ADVISORY COUNCIL ON UNEMPLOYMENT COMPENSATION:

BACKGROUND PAPERS

VOLUME III
JANUARY 1996

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PREFACE

In November 1991, the Congress of the United States passed the Emergency Unemployment Compensation Act (P.L. 102-164). The act included a section that created the Advisory Council on Unemployment Compensation, which was charged with the task of evaluating "the unemployment compensation program, including the purpose, goals, countercyclical effectiveness, coverage, benefit adequacy, trust fund solvency, funding of State administrative costs, administrative efficiency, and any other aspects of the program and to make recommendations for improvement."

The Advisory Council is made up of eleven members who represent the interests of business, labor, state governments, and the public. Five of the members are appointed by the President, three members are appointed by the Senate, and three members are appointed by the House of Representatives.

The Advisory Council has generally approached its work by focusing its attention on broad, fundamental elements of the Unemployment Insurance system. During 1993, its first year of operation, the Council examined the need for reform in the Extended Benefits component of the Unemployment Insurance system. Its work during the second year focused primarily on those issues related to benefits, eligibility, financing, and coverage. During its third and final year of operation, the Council is considering issues generally related to program administration, including appeals and federal-state responsibilities, as well as issues such as nonmonetary eligibility and program data.

In carrying out its mandate to evaluate and analyze the Unemployment Insurance system, the Advisory Council has relied on a diverse collection of information sources. The Council receives regular briefing materials from its staff and has also held a series of public hearings across the country in order to allow interested individuals and organizations to present their views to the Council. In addition, the Council has planned a number of academic conferences to facilitate the exchange of ideas and the presentation of works of research on Unemployment Insurance. These forums include two economics research conferences, one held in August 1994, and another planned for August 1995, and a legal symposium, sponsored jointly with the University of Michigan Journal of Law Reform in March 1995.

These two volumes contain much of the research that has been undertaken to date on behalf of the Council, both by the Council's staff and outside researchers. Additional reasearch will be published later this year. The papers presented at the legal symposium will be published separately by the University of Michigan Journal of Law Reform in 1996.

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Do Employers Use the Free-Layoff Loophole in Unemployment Insurance?

Paul L. Burgess
ICPLB@ASUACAD.ASU.EDU

Stuart A. Low
AOSAL@ASUACAD.ASU.EDU

Department of Economics
Arizona State University
Tempe, AZ 85287-3806
(602) 965-3531
(602) 965-0748 (FAX)

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I. OVERVIEW*

Numerous studies have analyzed the effects of UI incentives on employer layoffs, especially including the role of experience rating. But work prior to Burgess and Low (1993) ignored a hidden incentive which encourages employers to quickly screen and potentially lay off new employees. This incentive results from the unintended consequences of how nearly all states assign charges to employer accounts for UI benefits paid to workers placed on layoff in the first few months after initial hire. Importantly, this loophole is available to all employers, no matter what their experience rating.

The "free layoff" loophole noted above occurs because nearly all states assign layoff charges that determine employer tax rates only to the base-period employers of a worker who draws UI benefits. In turn, the base period in nearly all states is defined as the first four of the last five completed calendar quarters for a worker who begins a benefit year. Basing employer charges on base periods defined in this way means that employers have at least one full calendar quarter during which they can hire, screen and terminate a new employee without incurring any UI charges for the layoff. This UI incentive makes "early" layoffs of new employees free to the layoff firm and tends to increase both temporary and permanent layoff probabilities for workers during their first few months with a new firm. Thus, the free layoff incentive tends to destabilize rather than stabilize employment, and it reinforces other incentives firms already have to quickly determine whether workers are likely candidates for long-term employment. Certainly some early

*We thank Laurie Bassie, Daniel Hamermesh and Phillip Levine for helpful comments on an earlier version of the paper.

layoffs will be due to poor employee/firm job matches and would occur without the free layoff incentive. The opportunity to utilize free layoffs strengthens this behavior and extends it to marginal workers who otherwise may have remained employed.

Firm responses to the free layoff incentive we define above can be best analyzed with matched firm-worker data in which both the layoff employer and the base-period employers for each layoff can be determined. The matching process required is a huge undertaking because UI quarterly wage records from employers must be matched with the employment and layoff history records of individual workers and claimants. No state routinely maintains such a matched data base. However, we have a unique, matched firm-worker data base from the State of Illinois that has the necessary data for a systematic sample of 608 firms and more than 74,000 workers of these firms over a 6-year period. From this data base, it is possible to identify individual worker separations, including the layoff firm for each separation and whether the layoff was free or chargeable to the firm for experience-rating purposes.

Our matched firm-worker data set makes it possible to determine whether the free layoff loophole is quantitatively important enough to be of concern to policy makers. We find that the loophole is a potentially important one -- over one-fourth of all layoffs and over one-fifth of UI benefits paid are free to the layoff firm in our data set. We also find that the use of free layoffs is systematically related to firm characteristics, especially the UI tax rate.

The remainder of the paper is organized as follows: Section II contains a brief description of the institutional environment that leads to free layoff incentives. The data set used is described and contrasted to alternative data sources in Section III. An overview of layoff measures, especially focusing on the use of free layoffs, is presented in Section IV. Our multivariate results indicate that free layoffs are systematically related to certain firm characteristics. We present these results and discuss their implications in Section V. The final section contains the summary and conclusions.

II. INSTITUTIONAL BACKGROUND

The payroll tax faced by employers in all states has two components: a fixed rate that is independent of the firm's layoff history and an "experience rated" component that reflects the firm's prior layoff history. Firms between the minimum and maximum tax rates are subject to increased future tax rates due to increased UI compensable layoffs. Tax rate caps and other features of the present tax system result in incomplete experience rating and cross subsidies among firms that may increase the incidence of temporary layoffs (Feldstein, 1978; Topel, 1983 and 1985).

While these prior studies have considered the impact of the level of the taxable wage base, minimum and maximum tax rates and the slope of the tax rate schedule, they implicitly have assumed that the "last" employer is charged for its layoffs. In fact, however, employer charging provisions typically provide yet another dimension of incomplete experience rating through the free layoff loophole we identify.

In 1987, Illinois and most other UI jurisdictions assigned layoff costs to a firm depending on the base-period wages paid by the firm to its workers collecting UI benefits, where the worker's base period is defined as the first four of the last five completed calendar quarters prior to the benefit year.[1] The typical employer charging procedure involves a lag quarter between the base period and the benefit year, where the benefit year generally begins at the date of the initial

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[1]Nine UI jurisdictions in 1987 defined the base period as the 52 weeks immediately preceding the worker's benefit year. A base period with a lag quarter still is used in most UI jurisdictions.

valid UI claim. The firm initiating a layoff thus has between one and two full quarters (depending upon the date of hire) to terminate a new employee as a free layoff because the firm is not a base-period employer of the worker.[2] The lag quarter provides the largest free layoff incentive, but the ultimate impact of UI charging provisions also will be influenced by how a state allocates charges among base period employers and how the total UI benefits paid to any firm's workers affect its future tax rate. The intricacies of this process are considered in detail in Burgess and Low (1993, Chapter V).

The lag quarter provision leading to a free layoff incentive allows the firm to terminate its employment relationship with a worker without incurring any current UI charge. This incentive would tend to raise temporary and permanent layoff probabilities for workers during their first one to two quarters with a new firm. It should be emphasized that, apart from this UI incentive, layoff and quit probabilities tend to be much higher during the first few months of employment for new hires because of typical worker-firm sorting and firm screening. Thus, the free layoff incentive reinforces other incentives firms already have to quickly determine whether a new worker is a successful firm/worker match and whether the worker's continued employment is desired.

III. DATA

The sample analyzed is virtually identical to that analyzed in the Burgess and Low (1993) study of employer layoffs. The firms included are a subsample of the firm sample originally provided by the Illinois Department of Employment (IDES) for a study of firm compliance with

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[2]In addition to a free layoff potential due to a lag quarter, a worker in the midst of a benefit year may be laid off during the remainder of that benefit year with no charge to the layoff employer. In this paper we do not distinguish between these two free layoff forms.

UI reporting provisions by Burgess and St. Louis (1990).[3] See the Appendix in Burgess and Low (1993) for a detailed explanation of the sampling and data verification procedures. Here we only mention those procedures and note some minor differences between the samples and data in this study and in Burgess and Low (1993).

The probability sample of 611 firms analyzed in Burgess and Low (1993) was chosen to represent the third quarter of 1987 population of Illinois (UI covered) firms after applying appropriate sample weights. Because large (small) firms are overrepresented (underrepresented) in the raw sample, firm sample weights must be applied to accurately represent the population of firms. As our analysis here is more detailed than that in Burgess and Low (1993), we eliminate three firms with incomplete information and adjust the firm sample weights to reflect this difference. Otherwise, the firm sample weights are identical to those reported in the Appendix of Burgess and Low (1993).

The UI covered layoff behavior of our sample of 608 firms can be determined through time only by obtaining and matching the UI records of the individual employees of each of these firms. This match process resulted in approximately 1.5 million firm-worker records, but tracking that many records for layoffs and employment for nearly six years would be a complex and costly process. Consequently, random sampling was used to reduce the number of workers tracked for larger firms. Specifically, a random sample of the workers in each firm with reported 1987.3 employment of 60 or more was selected, and all employees of firms that reported fewer than 60 workers for 1987.3 were included (see Burgess and Low, 1993, Appendix Table 3 for the exact

[3]We thank the Illinois Department of Employment Security, especially Joseph Wojcik and Norman Harelik, for providing the unique data set we analyze.

employee sampling weights). This sampling reduced the number of matched firm-worker records to 74,422 for the 608 firms we analyze.

Each of these workers then was followed for the period from 1985.1 through 1990.2 and a hierarchial data record was constructed. Information on each worker consists of quarterly wage reports from each employer and full information on each UI claim, including its timing, duration, benefits paid, layoff employer and base period employers.[4]

To identify the layoff experience of each of our sample firms, each sampled worker is observed over a 4-5 quarter window beginning in the calendar quarter of 1987 when the employee first appears on the firm's employment roster.[5] Any valid UI reported layoff for the worker is then found and we determine whether the layoff employer is one of the sample firms.[6] Aggregating the worker experiences for each firm allows us to calculate firm-specific overall layoff and free layoff incidences and rates for our sample firms. The weighting issues to consider for calculating population estimates from our sample data are detailed in the Appendix.

Prior analyses of UI incentive effects have utilized several types of data. Many rely on CPS data for individual workers (Topel, 1983 and 1985). Drawbacks of this data source are that

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[4]Some data deficiencies were discovered on the data tape initially provided for Burgess and Low (1993). IDES graciously agreed to provide a corrected tape for this study. The new tape allows us to calculate layoff activity more accurately.

[5]Choosing a fixed one year period for all workers would not accurately capture the firm's layoff experience. Workers were selected for the sample based on the calendar year 1987 employee roster. A fixed period would mean that employees added to the firm at different times would be subject to different layoff risks. Further, as firm data are reported quarterly, we do not know how long an individual worked for a firm during any calendar quarter. Our 4-5 quarter window ensures that at least one year is included for all workers.

[6]Disallowed claims are not considered. In addition, workers who are terminated or quit and never file UI claims cannot be identified in our data. This implies that our layoff measures reflect only UI eligible unemployment.

individual UI status and entitlements are unknown and must be estimated for individual workers; no firm-specific information is available so aggregate state averages are ascribed; and the random sample of workers in the CPS does not correctly represent the population of firms. Other studies use aggregate state data (Brechling, 1981; Brown, 1986). These studies may have difficulty determining firm-specific impacts from aggregate data. Finally, several studies have employed the Continuous Wage and Benefit history data (Anderson, 1990; Anderson and Meyer, 1993), but these data sets are based on a sample of individuals, which may not be representative of the population of firms.[7]

Our matched firm-worker data circumvents these prior difficulties because it is based on the correct population of firms and contains firm and worker specific information. The primary drawback of our data is that it only considers one state at one point in time. Thus, one must be cautious when generalizing the results nationally or to a different stage of the business cycle. Obviously, it would be desirable to have additional data to evaluate the impact of the free layoff loophole we identify. For example, it would be especially helpful to have data for states that changed their provisions from the use of a lag quarter to the use of the last employer for charging UI benefits, both prior to and subsequent to the change in charging provisions. Alternatively, one could compare employer layoffs in different states that used the lag quarter versus the last employer in charging UI benefits. Unfortunately, none of the data required for using additional

[7]In the CPS, workers from firms with higher layoff rates and/or larger firms have a higher probability of representation in the sample. In the longitudinal CWBH, this difficulty is exacerbated because the workers of such firms are more likely to appear in the sample in successive quarters.

employer in charging UI benefits. Unfortunately, none of the data required for using additional states is available for our analysis. Thus, our free layoff findings rely on data for just one state.[8]

IV. OVERVIEW OF LAYOFF MEASURES

Various layoff and related measures are reported in Table 1. Free layoffs are those for which the layoff firm was not a base period employer because of the lag quarter. In contrast, the layoff firm was a base period employer for chargeable layoffs. Overall layoffs consist of both free and chargeable layoffs. These results are reported for the full sample and two subgroups -- firms that had at least one free layoff and firms that had at least one chargeable layoff (note that the same firms can appear in each of these subgroups). As noted above and explained in more detail in the Appendix, these propensity estimates are weighted by firm sampling weights, whereas the rate estimates are weighted by both firm and employee sampling weights to obtain appropriate population values.

Layoff propensity measures, defined as the proportions of firms with at least one layoff during the period are reported in item A. of Table 1. These results show that just over one-third (34.2 percent) of the firm population had some UI layoff activity during the 4-5 quarter window. For the full population, 11.6 percent of the firms had at least one free layoff and 30.5 percent had at least one chargeable layoff (implying that 7.9 percent engaged in both free and chargeable layoffs). For the subgroup of firms with some layoff activity, the free and chargeable layoff propensities obviously increase considerably above those for the full population, from 11.6

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[8]While our matched firm-worker data set evidently is the only such data presently available for the United States, the U.S. Department of Labor recently issued a solicitation (DAA-RFP 95-30) to develop similar data sets for additional states. Upon completion, analysis of these new data sets will permit expansion and generalization of our findings.

percent to 33.8 percent for free layoffs and from 30.5 percent to 89.2 percent for chargeable layoffs.

The UI layoff rates for this matched firm-worker population are reported in item B. of Table 1. For the full population, the overall layoff rate for this 4-5 quarter window is an estimated 5.6 percent. This overall layoff rate is comprised of a free layoff rate of 1.5 percent and a chargeable layoff rate of 4.1 percent. The layoff rates for firms with any layoff activity obviously are higher than the comparable layoff rates for the full population. The overall, free and chargeable layoff rates for firms with layoff activity in each category are 7.7 percent, 2.0 percent and 5.7 percent, respectively. Viewing the set of firms with at least one free or chargeable layoff, we note the highest overall layoff rate of 9.4 percent is for the free layoff subset.[9] It is interesting to note that the chargeable layoff rate is nearly the same for the three subgroups with layoff activity, including firms with free layoffs, varying from 5.7 percent to 6.3 percent. In contrast, the free layoff rate of 3.5 percent for firms with at least one free layoff is nearly twice as large as the free layoff rates for the overall and chargeable layoff categories of 2.0 percent and 1.9 percent, respectively. In other words, firms with free layoffs have higher overall layoff rates than other firms because of much more use of free layoffs, not because of different chargeable layoff rates. These results suggest that free layoff incentives may increase overall firm layoff rates, rather than encouraging firms to substitute free for chargeable layoffs while maintaining the same overall layoff rate.

The importance of free layoffs is further illustrated by items C.-E. of Table 1. These results show that 26.3 percent of all layoffs, 11.9 percent of overall UI weeks and 21.4 percent of

[9]Recall that firms with both free and chargeable layoffs are included in both subsets.

overall UI benefits were free to the layoff firm.[10] That is, free layoffs are a substantial portion of all layoff activity, accounting for over one-fourth of all layoffs and over one-fifth of all benefits paid.

The above results show considerable layoff activity during the 4-5 quarter window for the population. A natural question is whether firm layoff behavior varies with firm characteristics. As background for the analysis of systematic firm layoff behavior in the next section, we present some mean layoff estimates for firms classified by UI tax rate, industry and annual employment in Table 2. The layoff propensity means are for the full population, whereas the layoff rate means are for firms with at least one layoff during the 4-5 quarter window. These results reveal substantial variation in both layoff propensities and rates for firms with different characteristics. Further, the percent of layoffs that are free to the layoff firm also varies considerably with these firm characteristics.

Some implications may be drawn from the population estimates of mean UI layoff propensities and rates in Tables 1 and 2. First, the free layoff propensity is quantitatively large, comprising over one-fourth of all layoffs and over one-fifth of all UI benefits. The inability under the present chargeback system to charge the layoff employer for this large proportion of layoffs serves to exacerbate the imperfect experience rating impacts considered in other literature. Second, overall layoff rates for firms that engage in free layoffs are higher than for firms that engage in chargeable layoffs, but most of this impact is reflected in higher free layoff rates, with little difference in chargeable rates for the two subsets of firms. Thus, the free layoff incentive

[10]The only other estimates of the percentage of free layoffs are provided by Anderson and Meyer (1993, pp. 583-84). Using CWBH data for six states, they find that free layoffs vary from 16.3 to 37.2 percent, with a simple average of 24.1 percent for the six states.

may augment the chargeable layoffs the firm plans to undertake. In any case, the free layoff incentive may cause firms to more rapidly screen new workers to avoid UI charges for unsuccessful firm/worker matches. Finally, there is substantial variation in mean layoff propensities and rates for firms with different characteristics. The manner in which firm characteristics influence layoff behavior is considered in greater detail in the next section.

V. MULTIVARIATE RESULTS

We consider two main questions in this section. First, do overall firm layoff propensities vary systematically with firm characteristics? Second, does the proportion of free to total layoffs vary systematically with firm characteristics among the set of firms that makes at least one layoff? Because of our emphasis on free layoffs, we address overall layoff propensities only briefly as background for more extensively considering the second question. As large firms are overrepresented in the sample, weighted results are used for the population estimates we report.

The weighted probit estimates for the overall UI layoff propensities of all sample firms are reported in Table 3. Independent variables include UI tax rates, industry and firm size categories comparable to those in Table 2.[11] We also include controls for the number of quarters the firm has been in business (QUART), average annual earnings per worker (AVGERN) and the percent of workers who were paid via 1099s (T1099RAT).

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[11]For the tax rate, two dummy variables (the maximum tax rate class is excluded) are used. In alternate specifications, a continuous tax rate variable or a continuous tax rate variable with minimum and maximum tax rate dummy variables are estimated. Results are qualitatively similar to those reported. For firm size, four dummy variables (the intermediate annual employment category from 51 through 250 workers is excluded) are included. In an alternate specification, a continuous annual employment variable replaces the dummy variable, but this specification does not converge. Manufacturing is the omitted industry.

The results in Table 3 show that both minimum tax rate firms and those between the minimum and the maximum have significantly lower layoff likelihoods than maximum tax rate firms. The industry results reveal that construction and manufacturing tend to have higher layoff propensities than most other industry groups. Finally, layoff propensities are significantly lower for smaller firms (those under 50 workers) than for larger firms and also for firms with higher average annual earnings per worker.

The above results show that overall layoff propensities vary with firm characteristics, but our central issue is the degree to which firms with layoffs systematically vary their ratio of free to total layoffs. Because this ratio varies from zero to one, with some clustering at each limit, weighted double-limit Tobit estimates for this ratio are presented in Table 4.[12] Prior to considering the results, it is useful to consider how the independent variables are expected to influence the free layoff proportion.

Some facts of the experience rating process must be noted to identify the expected impact of the tax rate on the ratio of free to total layoffs. First, each tax rate category encompasses a range of experience, implying that a firm may vary its experience somewhat without moving into a different tax rate class. Second, as a firm moves up the tax rate schedule (into a higher category) a given increase in the UI tax represents a successively lower percentage rise in tax rates. For example, a .5% increase in the UI tax rate represents a 62.5% tax increase for a firm in the minimum rate category, compared to an increase of only 7.5% for a firm presently in the 6.7% tax category.

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[12]The Tobit model faces limits of 0 if no free layoffs were made and 1 if all layoffs were free. A model in which the double limit Tobit was censored by the probability of layoff could be formulated. Unfortunately, lack of identification of the two processes makes this infeasible at this time.

Because marginal layoffs have little (or no) effect on the future tax rates of many maximum-tax firms, they have little (or no) UI incentive to avoid layoffs or to be concerned about the distinction between free and chargeable layoffs. As maximum tax firms are our excluded group, we must consider incentives to vary the proportion of free to total layoffs by firms at other tax rates relative to maximum tax firms. At the other extreme, some minimum tax firms may remain at this rate even if they engage in some chargeable layoffs. But most minimum tax firms in our sample would face a relatively large percentage increase in future rates if they engage in chargeable layoffs.[13] Minimum tax firms thus have a strong incentive to rely on more free and fewer chargeable layoffs whenever possible and are expected to have a higher proportion of free layoffs than maximum tax firms. Finally, firms in the intervening tax ranges also have an incentive to utilize a higher proportion of free layoffs relative to maximum tax rate firms for two reasons. First, firms in the intermediate tax categories might avoid an increase in future rates by increasing the ratio of free to total layoffs. Alternatively, a higher proportion of free layoffs might reduce future tax rates for such firms and, as explained above, the incentive is stronger for firms in the intermediate tax range than for firms in the maximum tax category because the percentage reduction in tax rates is larger for the former group.

The above considerations suggest that both minimum tax firms and firms between the minimum and maximum tax rate have an incentive to use a higher proportion of free layoffs than otherwise comparable maximum tax firms. However, offsetting influences make it difficult to

[13]In our data, average firm size for minimum tax firms is far less than the overall firm size. For small firms in the minimum tax category, a single chargeable layoff is likely to cause the future tax rate to increase above the minimum (by a relatively large percentage). Thus, in our data, few minimum tax firms have the ability to engage in chargeable layoffs and remain at the minimum tax rate.

determine whether free layoff incentives are stronger for minimum tax firms or firms in the intermediate tax categories. On the one hand, firms in the intermediate tax range might reduce future tax rates by using a larger proportion of free layoffs, but minimum tax firms cannot. That is, possibly reducing future tax rates provides a stronger incentive for using a larger proportion of free layoffs by mid-range tax firms than by minimum tax firms. In contrast, a given increase in UI tax rates (e.g., 0.5%) represents a larger percentage increase in rates for minimum tax firms than for firms already above the minimum. That is, avoiding an increase in future tax rates provides a stronger incentive for using a larger proportion of free layoffs by minimum tax firms than by firms already above the minimum. Thus, which incentive dominates -- possibly reducing future UI tax rates or avoiding an increase in current UI tax rates -- determines whether the free layoff incentive tends to be stronger for firms in the intermediate tax range or minimum tax firms. In short, this is an empirical issue.

Controlling for the tax rate, there is no a priori reason that firms in certain industries or of particular sizes would have differential incentives to engage in higher free layoff proportions, although we may speculate that certain industries will tend to have high layoff rates (such as construction) or low layoff rates (such as financial services). Firms that have been in existence for a longer period may be both more cognizant of free layoff incentives and face a greater potential impact on future tax rates due to chargeable layoffs (as new firms may still be facing a constant "introductory" tax rate for their first 3 years). On the other hand, new firms will tend to have a less stable work force and thus more workers who are potentially free layoffs. Overall, the impact of QUART on the proportion of free layoffs is uncertain. The effect of higher average worker earnings also has an ambiguous impact. On the one hand, potential UI costs for layoffs tend to be smaller for higher (relative to lower) earning workers, tending to increase the likelihood of layoffs for firms with higher AVGERN values, ceteris paribus (see Burgess and Low, 1993, pp. 26-7). On the other hand, firms with higher AVGERN values may have better firm-worker matches and greater specific capital invested in the average worker, tending to reduce layoff likelihoods for high vs. low AVGERN firms, ceteris paribus. Finally, there is evidence firms that tend to use a relatively large proportion of 1099 workers attempt to avoid UI taxes by underreporting UI taxable wages (Blakemore, Burgess, Low and St. Louis, forthcoming, 1996). Such firms also may be especially sensitive to the tax implications of their layoff decisions and attempt to utilize free rather than chargeable layoffs to minimize future tax costs whenever possible.

The free layoffs results in Table 4 illustrate that minimum tax rate firms have a significantly higher proportion of free to total layoffs relative to maximum tax rate firms. The estimated partial impact indicates that this effect is quite strong. For example, if a typical maximum tax rate firm with layoffs has an expected proportion of free layoffs of 0.18, the comparable minimum tax rate firm would have an expected proportion of free layoffs of 0.59. Firms in the intermediate tax categories have insignificantly different free layoff proportions than maximum tax rate firms, but significantly lower proportions of free layoffs than minimum tax firms. Thus, the only significant tax effect we find is that minimum tax rate firms use a much larger proportion of free layoffs than firms in either of the other two tax categories. This suggests the incentive to avoid future tax increases is especially strong for firms in the minimum tax category.

The industry dummy variables show that, relative to the excluded manufacturing category, only financial firms have significantly lower free layoff proportions. Thus, it does not appear that there is much difference among industries in the proportion of free layoffs, ceteris paribus. Particularly after controlling for tax rates, this result is not surprising.

The annual employment results show that, relative to our excluded group (51-250 workers), smaller firms have significantly lower free layoff proportions and larger firms have insignificantly higher proportions. These firm size results may partly reflect the fact that only 20 of the 608 sample firms had no potentially free workers, yet 19 of these firms are of size 1-10 and the remaining firm is size 11-50. Thus, large firms in our sample always have some opportunity to choose between free and charged layoffs, but some small firms can only utilize chargeable layoffs or may have insufficient potentially free layoffs to handle all the layoffs the firm requires.

Neither the number of quarters a firm has been in business nor the firm's average earnings has any significant impact on the free layoff proportion. However, firms that utilize a greater proportion of 1099's have a significantly higher proportion of free to total layoffs. This confirms our expectation that such firms, already prone to minimize their UI taxes by underreporting, may be more responsive to free layoff incentives. The partial impact noted in Table 4 illustrates that a 10% increase in the proportion of workers classified as independent contractors (from the mean value of 10.7%) would be expected to lead to a 4% increase in the proportion of free layoffs for firms engaging in layoff activity. It thus appears firms that tend to reduce their UI tax burden through underreporting are more likely to recognize and act upon the free layoff incentive.

SUMMARY AND CONCLUSION

Nearly all states assign layoff charges that determine employer tax rates only to the base-period employers of workers who draw UI benefits. In turn, the base period in most states is defined as the first four of the last five completed calendar quarters. This gives employers at least one and up to two calendar quarters to screen and terminate new employees without incurring any UI charges. Using data for one state, we find that this free-layoff incentive is quantitatively important, with free layoffs accounting for over one-fourth of all layoffs and over one-fifth of all

UI benefits. These free layoffs are available to firms in all experience rating categories and represent an important addition to the imperfect experience rating analyzed in many other studies. The sample means also reveal considerable variation in the use of free (and chargeable) layoffs by firms with different characteristics. Thus, free layoffs may result in unintended cross-subsidies among different types of firms.

The multivariate results confirm that the proportion of free to total layoffs varies with certain firm characteristics. Two findings are especially interesting. First, minimum tax firms that engage in layoffs have significantly higher free layoff proportions than firms in other tax categories. Second, firms with a high proportion of independent contractors respond to the free layoff incentive by engaging in a higher proportion of free layoffs. Thus, two groups we expect to be especially sensitive to free layoff incentives -- minimum tax rate firms and firms that report a larger proportion of workers as independent contractors -- respond by engaging in greater free layoff proportions.

The quantitative importance of free layoffs and the differential use of them by different types of firms suggest that the free layoff incentive is important for policy makers to consider, especially since it appears counter to the objectives of the UI system. An obvious option is to eliminate the use of a lag quarter in charging base period employers for tax rate purposes. For example, the period used for charging employers for layoffs could be the 52 weeks immediately preceding a layoff (consistent with the base period definition already used in some states). On the one hand, such a change would eliminate the free layoff loophole we identify. However, such a change also could have detrimental impacts if it reduces a firm's propensity to hire a worker. Without a lag quarter, a firm may avoid hiring a "marginal" worker who could be immediately chargeable to the firm if placed on layoff. With the lag quarter, the firm may recognize its ability to engage in a free layoff and thus undertake more hiring. Which of these impacts may dominate ideally would require data from a state that has switched either from or to the use of the lag quarter in assigning employer charges. Although Illinois no longer has a lag quarter involved in assigning employer charges, data both before and after this change are not available.

REFERENCES

Anderson, Patricia M. "Linear Adjustment Costs and Seasonal Labor Demand: Unemployment Insurance Experience Rating in Retail Trade." Princeton University Department of Economics Working Paper (November 1990).

_________ and Bruce D. Meyer. "Unemployment Insurance in the United States: Layoff Incentives and Cross Subsidies." *Journal of Labor Economics* (January, 1993): 570-595.

Blakemore, Arthur E., Paul L. Burgess, Stuart A. Low and Robert D. St. Louis. "An Analysis of Employer Tax Evasion in the Unemployment Insurance Program," *Journal of Labor Economics*, (forthcoming, April 1996).

Brechling, Frank. "Layoffs and Unemployment Insurance," in Sherwin Rosen, ed., *Studies in Labor Markets*, NBER Conference on Income and Wealth. Chicago, Ill.: University of Chicago Press (1981): 187-207.

Brown, Eleanor P. "Unemployment Insurance Taxes and Cyclical Layoff Incentives," *Journal of Labor Economics* (January 1986): 50-65.

Burgess, Paul L. and Stuart A. Low. "Unemployment Insurance and Employer Layoffs," Unemployment Insurance Occational Paper #93-1, U.S. Department of Labor, (1993).

Burgess, Paul L. and Robert D. St. Louis, "Employer Compliance with Illinois Unemployment Insurance Reporting Requirements." Chicago, Ill.: Illinois Department of Employment Security (February 1990).

Feldstein, Martin. "The Effect of Unemployment Insurance on Temporary Layoff Unemployment," *American Economic Review* (December 1978): 834-46.

Topel, Robert. "Unemployment and Unemployment Insurance." In Ronald G. Ehrenberg, (ed.), *Research in Labor Economics* 7 JAI Press (1985): 91-136.

________. "On Layoffs and Unemployment Insurance." *American Economic Review* (September 1983): 541-59.

APPENDIX

Layoff Measures, Firm and Employee Sample Weights

Both firms and employees were sampled so appropriate sample weights must be applied to obtain population estimates of any layoff rates. Because the weighting procedure involves two sample weights for some measures and only one sample weight for others, we provide some hypothetical data in Panel A of Appendix Table 1. Then, we define the formulas we use and show the calculations for the hypothetical data in Panel B of Appendix Table 1.

Consider the hypothetical data for firm b in Panel A. Its firm sample weight of 5 in column (2) shows that 1/5 of the firms in this size category were sampled. The 200 employees sampled in column (3) represent 1/10 of the employees in this firm, as shown by the employee sample weight of 10 in column (4). Notice that the correct weight for the *population of employees* for this firm in column (5) is 50, or the product of the firm and employee sample weights (compared to a weight from column (2) of 5 for this firm in the *population of firms)*. Columns (6)-(8) show that this firm had 40 layoffs in the sample for a layoff indicator of 1 and a sample layoff rate of 2 percent.

The formulas and calculations for layoff propensities and layoff rates for the hypothetical data in Panel A are shown in Panel B. Two of the three sample firms have at least one layoff for an average layoff propensity of 66.7 percent in the *unweighted* sample. However, these firms have much different firm sample weights. Accounting for these firm sample weights, the correctly weighted average layoff propensity for the population of 35 firms represented by the three firms in our example is only 42.9 percent. This simple example illustrates why the (weighted) layoff propensity estimate for the population of firms is the relevant measure, not the simple layoff propensity for the unweighted sample. Thus, we report only layoff propensities that are weighted for the population of firms.

The layoff rates differ between the raw sample and the population of firms for the same reason explained above for the differences in layoff propensities. Specifically, as shown in the bottom of Panel B the average layoff rate for the raw sample is 10 percent, but the (correctly weighted) average layoff rate for our hypothetical firm population actually is 5.7 percent. Although 5.7 percent is the correct average layoff rate for the population of firms, it also is possible to calculate a weighted layoff rate for the population of employees.[1] In fact, the layoff rate for the population of employees is the rate normally reported in prior work.

The weighted average layoff rate for the population of employees accounts for both firm sample weights and employee sample weights, as shown in the bottom portion of the last column of Panel B. In our example, the average layoff rate for the population of employees is 12.6 percent. Because prior work emphasizes layoff rates for employees, we report layoff rates for the population of employees in text Tables 1 and 2.

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[1]Notice that the layoff rate formula for the weighted population of firms does not include employee sample weights. If the employee sample weight for each firm were included, it would simply cancel out because it would appear in both the numerator and the denominator of each firm's layoff rate.

Appendix Table 1
Hypothetical Sample Data and Formulas for Layoff Propensity and Rates

A. Hypothetical Sample Data

(1) Firm Sampled	(2) Firm Sample Weight for i[th] firm = W_i^f (= weight for population of firms)	(3) Number of Employees Sampled for i[th] firm = E_i	(4) Employee Sample Weight for i[th] firm = W_i^e	(5) Weight for Employee Population for i[th] firm = $W_i^f W_i^e$	(6) Number of Layoffs in Sample for i[th] firm = L_i	(7) Sample Layoff Indicator for i[th] firm = l_i ($l_i = 1$ if $L_i > 0$; $l_i = 0$ if $L_i = 0$)	(8) Sample Layoff rate for i[th] firm $= P_i = \frac{L_i}{E_i}$
a	20	100	2	40	0	0	0
b	5	200	10	50	40	1	.2
c	10	300	3	30	30	1	.1

B. Formulas for Layoff Propensity and Rates (illustrated for above data)

Measure	*Firm Sample (Unweighted)*	*Population of Firms (Weighted for Firm Sample Weights)*	*Population of Employees (Weighted for Both Firm and Employee SampleWeights)*
Firm Layoff Propensity	$\frac{\sum_i l_i}{\sum_i 1} = \frac{0+1+1}{3} = .667$	$\frac{\sum_i l_i W_i^f}{\sum_i W_i^f} = \frac{0(20)+1(5)+1(10)}{20+5+10} = \frac{15}{35} = .429$	Not Applicable
Layoff Rate	$\frac{\sum_i P_i}{\sum_i i} = \frac{0+.2+.1}{3} = .1$	$\frac{\sum_i P_i W_i^f}{\sum_i W_i^f} = \frac{0(20)+.2(5)+.1(10)}{20+5+10} = \frac{1+1}{35} = .057$	$\frac{\sum_i L_i W_i^f W_i^e}{\sum_i E_i W_i^f W_i^e} = \frac{0(40)+40(50)+30(30)}{100(40)+200(50)+300(30)} = \frac{2{,}900}{23{,}000} = .126$

Table 1
Population Estimates of Mean UI Layoffs and Related Measures During 4-5 Quarter Window

	Full Population	Firms with at Least One:		
		Overall Layoff	Free Layoff	Chargeable Layoff
A. Proportion of Firms with One or More UI Layoffs (Layoff Propensity):				
1. Overall	.342	1.000	1.000	1.000
2. Free to Layoff Firm	.116	.338	1.000	.258
3. Chargeable to Layoff Firm	.305	.892	.682	1.000
B. UI Layoff Rates:				
1. Overall	.056	.077	.094	.082
2. Free to Layoff Firm	.015	.020	.035	.019
3. Chargeable to Layoff Firm	.041	.057	.059	.063
C. Percent of Overall Layoffs Free to Layoff Firm	----	26.3%	37.2%	23.5%
D. Percent of Overall UI Benefits Free to Layoff Firm	----	21.4%	32.9%	18.2%
E. Percent of Overall UI Weeks Free to Layoff Firm	----	11.9%	18.9%	9.6%
F. Firm Sample Size	608	394	207	363

Table 2
Population Estimates of Mean UI Layoff Propensity and Rates by Statutory UI Tax Rate, Industry and Firm Size

Group	Layoff Propensity for Full Population		Layoff Rates for Firms with at Least One Layoff		
	Overall	Sample Size	Overall	Percent of Layoffs Free to Layoff Firm	Sample Size
Full Population	.342	608	.077	26.3%	394
UI Tax Rate					
.8% (min)	.146	98	.066	41.5%	25
.81-1.99%	.431	97	.028	22.4%	65
2.00-3.49%	.379	110	.033	26.4%	78
3.50-4.99%	.377	126	.060	14.9%	84
5.00-7.29%	.472	73	.065	25.3%	55
7.3% (max)	.601	104	.180	29.1%	87
Industry					
Construction	.458	46	.243	37.0%	36
Manufacturing	.617	116	.075	20.3%	98
Trans/Commun/Pub Util	.260	33	.046	22.9%	23
Wholesale Trade	.361	57	.085	16.2%	34
Retail Trade	.217	129	.033	22.2%	65
Fin/Ins/Real Estate	.230	55	.038	14.4%	36
Services	.324	149	.067	28.2%	86
Other	.469	23	.083	16.4%	16
Annual Employment					
1-10	.204	72	.232	16.9%	16
11-50	.398	177	.112	19.2%	81
51-250	.713	209	.077	29.1%	152
251-1000	.944	81	.068	36.2%	78
1000+	.977	69	.037	13.2%	67

Table 3
Weighted Probit Estimates of Layoff Propensity for the Full Sample

	Coef	t-Ratio	Partial[a]
UI Tax Rate			
.8% (min)	-1.108*	-6.00	-.28
.81-7.29%	-.652*	-3.72	-.20
Industry			
Construction	.062	.22	.02
Trans/Commun/Pub Util	-.620**	-1.96	-.19
Wholesale Trade	-.292	-1.08	-.10
Retail Trade	-.935*	-3.61	-.25
Fin/Ins/Real Estate	-.539***	-1.62	-.17
Services	-.140	-.58	-.05
Other	.145	.39	.05
Annual Employment			
1-10	-1.439*	-6.98	-.31
11-50	-.868*	-4.41	-.24
251-1000	1.061	1.44	.40
1000+	1.697	.46	.56
QUART	-.007	-.55	-.00
AVGERN	-.018*	-2.68	-.01
T1099RAT	.516	1.29	.02
CONSTANT	1.82*	4.99	
Sample Size	608		

The omitted categories for the dummy variables are the maximum tax rate (7.30%), manufacturing and annual employment of 51-250.

[a]The partial impact is calculated starting at the weighted sample probability of .34. Then, the partial impact for each variable is calculated by determining the change in the expected probability that results from: assigning dummy variables a value of 1; changing QUART and AVGERN by 1 unit; and changing T1099RAT by .1 (10%).

*indicates statistical significance at the .01 level.
**indicates statistical significance at the .05 level.
***indicates statistical significance at the .10 level.

Table 4
Weighted Double Limit Tobit Estimates of the Proportion of Layoffs Free to the Layoff Firm for Firms With At Least One Layoff

	Coef	t-Ratio	Partial*
UI Tax Rate			
.8% (min)	.897*	3.64	.41
.81-7.29%	-.088	-.53	-.02
Industry			
Construction	.236	.99	.08
Trans/Commun/Pub Util	-.442	-1.18	-.09
Wholesale Trade	-.193	-.75	-.05
Retail Trade	-.325	-1.16	-.08
Fin/Ins/Real Estate	-.712***	-1.70	-.13
Services	.261	1.19	.09
Other	-.395	-1.09	-.09
Annual Employment			
1-10	-1.189*	5.01	-.16
11-50	-.559*	-3.25	-.11
251-1000	.192	.64	.06
1000+	.439	.42	.16
QUART	-.012	-.78	-.00
AVGERN	-.004	-.47	-.00
T1099RAT	1.295*	2.70	.04
CONSTANT	.057	.17	
SIGMA	1.024*	10.88	
Sample Size	394		

The omitted categories for the dummy variables are the maximum tax rate (7.30%), manufacturing and annual employment of 51-250.

*The partial impact is calculated starting at the weighted sample free layoff proportion of .18. Then, the partial impact for each variable is calculated by determining the change in the expected proportion that results from: assigning dummy variables a value of 1; changing QUART and AVGERN by 1 unit; and changing T1099RAT by .1 (10%).

*indicates statistical significance at the .01 level.
**indicates statistical significance at the .05 level.
***indicates statistical significance at the .10 level.

2/20/95

Optimal Unemployment Insurance

Carl Davidson and Stephen A. Woodbury

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Paper prepared for the Advisory Council on Unemployment Compensation, whose support is gratefully acknowledged. We are grateful to Richard Deibel and Ellen Maloney for help in preparing the manuscript. Davidson is Professor of Economics, Michigan State University, East Lansing, MI 48824; Woodbury is Professor of Economics, Michigan State University and senior economist, W.E. Upjohn Institute.

I. Introduction

Risk averse workers facing uncertain employment prospects prefer to insure against adverse economic conditions such as unemployment. If they could, they would purchase private unemployment insurance in order to finance consumption during jobless spells. In fact, if the insurance were actuarially fair, it is well known that the all risk averse workers would choose to fully insure so that consumption during unemployment would exactly equal consumption while employed. But, for a variety of reasons, insurance markets are incomplete, and private unemployment insurance cannot be purchased.

In the absence of private insurance markets, agents will try and save during periods of employment and dissave during jobless spells. It is unlikely, however, that workers would be able to save enough to completely smooth consumption across periods of employment and unemployment. In response to this problem, virtually every developed country provides public unemployment insurance (UI). In the United States, there is considerable empirical evidence that UI does what it was intended to do -- it allows workers to smooth consumption. For example, in a recent paper, Gruber (1994) estimates that without UI consumption would fall by 22% during unemployment, whereas it falls by only 7% with UI in place.

But UI has unintended effects as well. By now there is considerable evidence that UI increases the length of unemployment spells.[1] By providing unemployment insurance, the government reduces the opportunity cost of unemployment. This reduces search effort and increases both the length of unemployment spells and the equilibrium rate of unemployment.[2] In designing an optimal UI program, the positive and negative effects of UI must be weighed against one another.

There are two classic theoretical treatments of optimal UI -- Baily (1978) and Flemming (1978). Both take the same approach, considering the situation faced by a typical unemployed worker and solving for optimal search effort as a function of UI. Although the actual spell of unemployment is a random variable, its expected value varies inversely with search effort. Both authors solve the optimal insurance problem by choosing UI to maximize the expected lifetime utility of the representative worker. The papers differ in their treatments of leisure, savings, and the capital market. Nevertheless, both papers and the empirical work making use of their approach all seem to conclude that UI payments in the United States are too generous (see, for example, Gruber 1994 and O'Leary 1994).

The purpose of this paper is to extend the analysis offered by Baily and Flemming in two ways. First, in formulating their models, both authors assume that UI is offered indefinitely -- that

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[1] See Davidson and Woodbury (1995b) for a review and new evidence based on the reemployment bonus experiments.

[2] It is often argued, on the other hand, that UI makes workers choosier about the jobs they accept, and that this may improve the quality of job matches. This notion has persisted despite very little empirical evidence in support of it.

is, unemployed workers collect UI benefits in every period until they find a job. But few UI systems are set up to pay benefits indefinitely. In the United States, workers usually exhaust their UI benefits after 26 weeks of unemployment. The potential duration of benefits is longer in Canada -- where it is about 1 year -- and in most of Western Europe -- where it is 3 years or longer in several countries, and indefinite in Belgium (OECD 1991). Even in the countries where UI is offered for 3 years or longer, a significant number of workers remain unemployed long enough to exhaust their benefits. In section III, we show that taking into account the finite potential duration of benefits drastically alters the conclusions reached by Baily and Flemming. For example, Flemming finds that if lending and borrowing are ruled out, the optimal replacement rate is approximately .75. The optimal replacement rate is close to .75 in our model as well (it is actually around two-thirds), assuming that UI is offered indefinitely. However, if UI is offered for only 26 weeks, the optimal replacement rate rises to 1.

Also in section III, we solve for the optimal UI program assuming that it can be characterized by two instruments -- the level of UI benefits (or the replacement rate) and the potential duration of benefits. Surprisingly, we find that the optimal UI program is characterized by an infinite potential duration of benefits. The argument is as follows. Let x denote the level of benefits and let T denote the potential duration of UI. Suppose that we compare two UI programs (x_1, T_1) and (x_2, T_2) with $x_1 > x_2$ and

$T_1 < T_2$ so that the second program offers lower benefits but a longer potential duration of benefits. Suppose further that these two programs cost taxpayers the same amount to fund so that employed workers earn the same after-tax wage under the two programs. We find that all risk-averse unemployed workers prefer the second program in spite of the fact that benefits are lower. They prefer the second program because the reduction in the probability that they will exhaust their benefits more than offsets the reduction in their benefits. In the terminology of decision making under uncertainty, the second program is "less risky" than the first program and is therefore preferred by all risk averse agents. Since the optimal UI program offers workers benefits indefinitely while most State programs in the United States offer benefits for only 26 weeks, the model's results suggest that the current United States system may not be generous enough.

The second extension we offer concerns the composition of the pool of unemployed workers. Both Baily and Flemming assume that all unemployed workers are eligible for UI benefits. In reality, fewer than half of all unemployed workers in the United States are UI-eligible (Blank and Card 1991). We show that this fact has important implications for the optimal replacement rate. Briefly, there are two effects. First, since an increase in UI benefits reduces the search intensity of UI-eligible workers, UI-ineligibles gain as they face less competition for jobs. This positive spill-over effect of UI increases the optimal replacement rate. The second effect is more subtle and depends on the degree of substitutability in production between UI-eligible and ineligible workers. Since UI-ineligibles receive no UI benefits, they search harder than UI-eligible workers. If these two types of workers are close substitutes, then treating all workers as if they are UI-eligibles will overstate the reemployment prospects for UI-eligible workers. In this case, the presence of UI-ineligibles in the workforce increases the optimal replacement rate; that is, since UI-ineligibles make it harder for UI-eligibles to find reemployment, the government needs to increase the level of insurance it provides to UI-eligibles. On the other hand, if UI-ineligibles tend to be lower-skilled workers who are poor substitutes for UI-eligible workers, then treating all workers as if the were UI-eligible will understate the reemployment prospects of UI-eligible workers. In this case, the presence of UI-ineligibles in the workforce lowers the optimal replacement rate (i.e., less insurance is needed). When we combine the spill-over effect and the effect of substitutability between UI-eligibles and UI-ineligibles, we find that unless the degree of substitutability between UI-eligibles and UI-ineligibles is extremely low, the presence of UI-ineligibles raises the optimal replacement rate.

In summary, we emphasize the importance of extending the models of Baily and Flemming to incorporate two empirical features of the UI system -- that UI benefits are offered only for a finite length of time and that not all workers are eligible for UI benefits. When their models are extended to include these features, the optimal replacement rate rises. In fact, we find that for reasonable parameter values, our model suggests that average statutory UI benefits in the United States are too low and that the potential duration of benefits is too short.

The paper is divided into three additional sections. In section II, we introduce a model that is similar in spirit to those of Baily and Flemming in that it assumes that all unemployed workers are eligible for UI. However, our model differs from theirs in that we allow for a finite potential duration of benefits. Using this model, we show in section III.A that any program that eventually cuts off benefits is Pareto-Dominated by another program that offers more periods of coverage. Thus, any optimal program must include an infinite potential duration of benefits. In section III.B, we solve for the optimal replacement rate under a program in which benefits are offered indefinitely. In section III.C we calculate optimal replacement rates for sub-optimal programs -- that is, programs in which benefits are cut off after a certain length of time. In section IV.A we drop the assumption that all unemployed workers are eligible for UI, and show that when UI-ineligibles are added to the model the optimal replacement rate is likely to increase. In section IV.B we consider the effects of adding voluntary saving to the model. We reason that, although including savings would reduce the optimal replacement rate somewhat, it would not alter our conclusion that an infinite potential duration of benefits is optimal. Finally, in section V we discuss the omission of worker heterogeneity from the model and offer some conjectures as to how this omission might affect our results. We also summarize and discuss the applicability of the results.

II. Model and Approach

We follow Baily and Flemming by modeling the behavior of a representative unemployed worker who is searching for employment. This worker earns a wage of w while employed and collects UI benefits of x while unemployed provided that he has not exhausted his benefits. Benefits are provided by the government to all jobless workers who have been unemployed for no more than T periods. UI is funded by taxing all employed workers' incomes at a constant rate τ.

We assume that unemployed workers choose search effort (p) to maximize expected lifetime income and that all workers are infinitely lived.[3] Given total labor demand (F), search effort determines equilibrium steady-state unemployment (U).[4] The government's goal is to choose x and T to maximize aggregate expected lifetime income. Increases in x and/or T provide unemployed workers with additional insurance but these increases also lower optimal search effort and therefore increase equilibrium unemployment. The optimal government policy must balance these two opposing forces.

Formally, we use L to denote total labor supply and let J represent the total number of jobs held in the steady-state equilibrium. Then, since every worker is either employed or

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[3] We assume infinite life since it makes the model much more tractable. Flemming also makes this assumption while Baily uses a two-period model.

[4] Following Baily and Flemming, we do not model the firm and treat F and w as exogenous variables.

unemployed, we have:

(1) $L = J + U.$

For later use, we define U_t to be the equilibrium number of workers who have been unemployed for t periods ($t = 1,\ldots,T$) and let U_x represent the equilibrium number of unemployed workers who have exhausted their UI benefits. We then write total unemployment as:

(2) $U = \Sigma_t U_t + U_x.$

Turn next to the firms. For simplicity, we assume that each firm provides only one job opportunity.[5] Thus, F denotes both the total number of firms and the total number of jobs available at any time. Each job is either filled or vacant. If we let V denote the number of vacancies in a steady-state equilibrium, it follows that:

(3) $F = J + V.$

The remainder of the model is explained in three stages. First, we describe the dynamics of the labor market and derive the conditions that must hold in a steady-state equilibrium. These conditions guarantee that the unemployment rate and the composition

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[5] This assumption is commonly used in general equilibrium search models (see, for example, Diamond 1982 or Pissarides 1990). Alternatively, we could simply assume that each firm recruits for and fills each of its many vacancies separately.

of unemployment both remain constant over time. Second, we relate search intensity by unemployed workers to their reemployment probabilities. We then use these reemployment probabilities to derive the expected lifetime incomes of employed and unemployed workers. Finally, in stage three, we derive the optimal level of search effort for all unemployed workers.

To describe the dynamics of the labor market, let s denote the probability that an employment relationship will break up in any given period -- that is, the job turnover or separation rate. In addition, let m_t and m_x denote the reemployment probabilities for workers in their t^{th} period of search and for UI-exhaustees, respectively. For any given worker, there are T + 2 possible employment states -- U_1, U_2,...., U_T, U_x, and J. If employed (i.e., if in state J) the worker faces a probability s of losing her job and moving into state U_1. If unemployed for t periods (i.e., if in state U_t), the worker faces a probability of m_t of finding a job and moving into state J. With the remaining probability of $1 - m_t$ this worker remains unemployed and moves on to state U_{t+1}. Finally, UI exhaustees face a reemployment probability of m_x, in which case they move into state J. Otherwise, they remain in state U_x.

In a steady-state equilibrium the flows into and out of each state must be equal so that the unemployment rate and its composition do not change over time. Using the above notation, the flows into and out of state U_1 are equal if:

(4) $sJ = U_1$.

The flows into and out of state U_t (for $t = 2,\ldots,T$) are equal if:

$$(5) \quad (1 - m_{t-1})U_{t-1} = U_t.$$

Finally, the flows into and out of state U_x are equal if:

$$(6) \quad (1 - m_T)U_T = m_x U_x.$$

In each case, the flow into the state is given on the left-hand-side of the expression while the flow out of the state is given on the right-hand-side.

Turn next to the reemployment probabilities. Each unemployed worker chooses search effort to maximize expected lifetime income. We use p_t to denote the search effort of a worker who is in her t^{th} period of search, with p_x playing the same role for UI exhaustees. Search effort is best thought of as the number of firms a worker chooses to contact in each period of job search. (For workers who contact fewer than one firm on average, p_t could be thought of as the probability of contacting any firm.) Once a worker contacts a firm, she files an application for employment if the firms has a vacancy. Since there are F firms and V of them have vacancies, the probability of contacting a firm with a vacancy is V/F. Finally, once all applications have been filed, each firm with a vacancy fills that vacancy by choosing randomly from its pool of applicants. Thus, if N other workers apply to the firm, the probability of a given worker getting the job is 1/(N+1). Since each other worker either does or does not apply, N is a random variable with a Poisson distribution with parameter λ equal to the average number of applications filed at each firm. It is straightforward to show that this implies that the probability of getting a job offer conditional on having applied at a firm with a vacancy is $(1/\lambda)[1 - e^{-\lambda}]$. The reemployment probability for any given worker is then the product of these three terms -- the number of firms contacted, the probability that a given firm will have a vacancy, and the probability of getting the job conditional on having applied at a firm with a vacancy:

$$(7) \quad m_t = p_t(V/F)(1/\lambda)[1 - e^{-\lambda}] \qquad \text{for } t = 1,\ldots,T$$

$$(8) \quad m_x = p_x(V/F)(1/\lambda)[1 - e^{-\lambda}]$$

where

$$(9) \quad \lambda = \{\Sigma_t p_t U_t + p_x U_x\}/F.$$

These equations define the reemployment probabilities of workers as a function of search effort and the length of time that they have been unemployed (since m_t varies over time). Note that for any given worker, the search effort of other workers affects that worker's reemployment probability through λ.

Finally, to determine optimal search effort we must first define expected lifetime income for all workers. Let V_w denote the expected lifetime income for an employed worker and let V_t and V_x play the same role for unemployed workers in their t^{th} period of search and for UI-exhaustees, respectively. For an employed worker, current income is equal to the net wage, $w(1 - \tau)$ where τ is the marginal (and average) tax rate. Her future income depends upon her employment status -- with probability s she loses her job and can expect to earn V_1 in the future, and, with the remaining probability she keeps her job and continues to earn V_w in the future. Thus,

$$(10) \quad V_w = w(1 - \tau) + [sV_1 + (1 - s)V_w]/(1 + r).$$

Note that future income is discounted with r denoting the interest rate.

For unemployed workers, current income is equal to unemployment insurance (if benefits have not yet been exhausted) less search costs. We assume that the cost of search is given by c(p) where c is a convex function with c(0) = 0. Future income depends on future employment status -- with probability m_t the worker finds a job and can expect to earn V_w in the future, while with the remaining probability she remains unemployed and can expect to earn V_{t+1} in the future. Thus,

$$(11) \quad V_t = x - c(p_t) + [m_tV_w + (1 - m_t)V_{t+1}]/(1 + r) \qquad \text{for } t = 1,\ldots,T.$$

$$(12) \quad V_x = - c(p_x) + [m_xV_w + (1 - m_x)V_x]/(1 + r).$$

Unemployed workers choose search effort (p_t) to maximize expected lifetime income (V_t). Thus,

(13) $p_t = \arg\max V_t$ for $t = 1,...,T$

(14) $p_x = \arg\max V_x$.

This completes the description of the model. Structurally it is very similar to Flemming's model. However, Flemming assumed that UI benefits are offered indefinitely and therefore, in his model all unemployed workers are identical. One of our purposes is to relax the assumption of indefinite benefits. Our model allows us to capture the notion that unemployed workers who have been unemployed for a longer period of time will search harder as they begin to worry about exhausting their benefits. In addition, as we show below, once we take into account the fact that UI is <u>not</u> offered indefinitely, conclusions about optimal UI levels are altered drastically.

Before we turn to optimal policy, it is useful to first describe the structure of equilibrium and some of its comparative dynamic properties. It is straightforward to show that the structure of equilibrium is such that $V_w > V_1 > V_2 >...> V_T > V_x$. That is, expected lifetime income is highest for employed workers, lowest for unemployed workers who have exhausted their benefits, and decreasing in the number of weeks that a worker has been unemployed. Intuitively, workers in the early stages of a spell of unemployment have more weeks to find a job before they have to worry about losing their UI benefits. Because of this, workers who have recently become unemployed will not search as hard as those who have been unemployed for a longer period of time -- that is, optimal search effort will be increasing in the number of weeks of unsuccessful search ($p_1 < p_2 <\ldots< p_T < p_x$).

A decrease in UI benefits (x) or the potential duration of benefits (T) decreases the level of insurance offered unemployed workers and triggers an increase in search effort by all UI-eligible workers (and therefore lowers equilibrium unemployment). Either change results in a decrease in V_t for all t. But decreases in x and T have opposite effects on the probability of exhausting benefits. A decrease in x makes it less likely that a worker will exhaust her UI benefits before finding a job (since she searches harder). But a decrease in T makes it more likely that benefits will be exhausted since the time horizon over which benefits are offered has been shortened (this is true even though search effort increases as a result of the decrease in T).

One final feature of the model needs to be emphasized. Although we assume that agents act to maximize expected lifetime income (as opposed to utility), they are in fact risk averse. Risk aversion follows from the assumption that search costs are convex in search effort. Any increase in the wage or decrease in UI benefits triggers an increase in search effort; but since search costs are convex, optimal search effort is concave in w and x. This implies that expected lifetime income is concave in w and x, making the worker risk averse with respect to income. This is important because it implies that any policy change that reduces the risk associated with unemployment will be welfare enhancing.

III. Social Welfare and Optimal UI Benefits

In the context of the model outlined above, social welfare can be calculated by aggregating expected lifetime income across all workers. In a steady-state equilibrium there are J employed workers with expected lifetime incomes of V_w, U_t unemployed workers who are in their t^{th} period of search with expected lifetime incomes of V_t, and U_x unemployed workers who have exhausted their UI benefits with expected lifetime incomes of V_x. Aggregating yields Social Welfare (SW):

$$(15)\quad SW = JV_w + \Sigma_t U_t V_t + U_x V_x.$$

The government's problem is to choose x (the UI benefit level) and T (the potential duration of benefits) to maximize (15) with the tax rate, τ, set such that the government budget balances:

$$(16)\qquad Jw\tau = x(U - U_x).$$

As noted above, increases in x or T increase the level of insurance provided to unemployed workers but also increase equilibrium unemployment and require that τ increase in order to fund the expanded program.

A. Optimal Potential Duration of Benefits

The most straightforward way to determine the optimal UI program is to proceed in two steps. First, for any tax rate (τ), we consider the set of all tax neutral programs (so that workers' incomes are the same while employed under any of the programs) and determine which one leads to the highest expected lifetime income for unemployed workers. Two programs are defined to be tax neutral if they are funded by taxing income at the same rate. It follows that if two programs are tax neutral, workers' net income while employed will be the same under either program. Thus, if one program leads to a higher V_t for all t and a higher V_x, it must be superior to the other program. This is in fact the case -- if we consider two tax neutral programs, the program with the longer potential duration of benefits (higher T) and the lower level of benefits (lower x) will lead to larger values of V_t for all t and a larger value of V_x. Thus, for any given τ, the optimal program is characterized by $T = \infty$. Setting $T = \infty$ allows us to write the optimal program for any given τ as $x(\tau)$. In the second step, we then maximize social welfare over $x(\tau)$.

To see why it is optimal to set $T = \infty$, consider any program (x, T) where T is finite. Now, increase the potential duration of benefits (T) by one period and lower the weekly benefit amount (x) in a tax neutral manner. What are the affects of this change in policy? Since the change is tax neutral, net income while employed is unchanged. For the unemployed, there are both direct and indirect effects on current income. The direct effect is that benefits are lower in the first T periods of unemployment but benefits are now offered for an additional period. The indirect effect works through search effort. For reasons that will become clear shortly, the policy change reduces search effort in all periods of unemployment, thereby lowering search costs. Once we combine these effects, we are left with three cases to consider -- there are periods $t = 1,\ldots,T$ in which the worker is eligible to receive UI under either program, there is period T+1 in which the worker receives UI under the new program but not the old program, and there are periods $t = T+2,\ldots.$ in which the worker does not receive UI benefits under either program. In periods $1,\ldots,T$, the direct effect of lowering benefits swamps the cost savings from reduced search effort so that current income falls. In period T+1, the worker receives benefits under the new program, raising current income. Finally, in periods T+2 and on, there are no benefits to lower, so everything depends on the indirect effect -- since search costs are lower, current income is higher.

This impact of the policy change on current income is depicted in Figure 1. Current income for the employed is unchanged, it falls for unemployed workers in the first T periods of search, and it increases for all unemployed workers who have been searching at least T+1 periods. Thus, this policy change increases income in the most adverse states of unemployment and lowers it in the least adverse states of unemployment -- it smoothes income across possible states of unemployment. Since the unemployed are risk averse and since total UI benefits given to the unemployed are the same under the two programs, this raises the expected lifetime utility of all unemployed workers.[6]

In this model with homogeneous workers, increasing T and lowering x in a tax neutral manner makes all unemployed workers better off. Accordingly, their expected lifetime incomes rise (V_t increases for all t). This is why the policy change lowers search effort -- since expected lifetime income while unemployed rises, the opportunity cost of unemployment falls, triggering a decrease in search effort.

Extending the potential duration of benefits in a tax neutral way also increases the expected lifetime income for employed workers (V_w). To see why, consider (10) which defines V_w. Since the policy change is tax neutral, $w(1 - \tau)$ does not change. However, since the unemployed are better off, V_1 rises. Thus, V_w increases. It follows that the shift in policy makes all agents better off.

In summary, a tax neutral change in policy that increases the potential duration of benefits (T) and lowers the weekly benefit amount (x) smoothes the receipt of income over states of unemployment without lowering the total amount of income received by the unemployed. Since all risk averse agents wish to smooth consumption, this makes all agents better off.

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[6] In the terminology of decision making under uncertainty, the policy change results in a "Rothschild-Stiglitz decrease in risk" for unemployed workers (see Rothschild and Stiglitz, 1970).

B. Optimal Replacement Rates with Unlimited Benefit Duration

We next obtain the optimal UI replacement rate under the assumption that T -- the potential duration of UI benefits -- equals infinity. Setting T equal to infinity makes sense for two reasons. First, we found above that it is the optimal policy. Second, setting T to infinity simplifies the model greatly because it makes all unemployed workers behave in an identical fashion over the entire spell of unemployment. Since no worker is getting close to exhausting benefits, all earn the same present and future income and choose the same level of search effort. If the potential duration of benefits were limited, search intensity would vary over the spell of unemployment, rising as the exhaustion point neared. (In the next sub-section, we obtain the optimal replacement rate under limited potential duration of UI benefits.)

When T is set to infinity, equations (1) and (3) are unchanged, while (2) becomes unnecessary. In addition, since we no longer need to keep track of the composition of unemployment, the steady-state equations can be simplified. Equations (5) and (6) can be dropped while (4) needs to be modified. While the flow into unemployment is still sJ, the flow out of unemployment becomes (1 - m)U, where m represents the reemployment probability for any unemployed worker. Thus, the new steady-state condition becomes sJ = mU.

The probability of reemployment (m) also becomes simpler to define -- it is now defined by (7) with the t subscripts on m and p dropped. Equation (8) can be dropped, and the definition of λ

simplifies to $\lambda = pU/F$.

Turn next to expected lifetime income and search effort. Define V_u to be the expected lifetime income earned by all unemployed workers. Then, using the same logic as in section A, (10) and (11) can be written as:

$$V_w = w(1 - \tau) + [sV_u + (1 - s)V_w]/(1 + r) \quad \text{and,}$$

$$V_u = x - c(p) + [mV_w + (1 - m)V_u]/(1 + r).$$

Optimal search effort (p) is chosen to maximize V_u.

Finally, for the government, Social Welfare can now be written as $SW = JV_w + UV_u$ while the government budget constraint can be simplified to $Jw\tau = xU$. The government's goal is now to choose x to maximize SW subject to its budget constraint.

Although this model is far simpler than the one laid out in section A, it is still too complex to yield a closed form solution for the optimal value of x. Again following Baily and Flemming, we choose parameter values and solve the model explicitly for the optimal x. Assuming that our parameters are chosen wisely, this should give us some idea of the range in which the optimal level of benefits falls.

The parameters of the model include the separation rate (s), the interest rate (r), the wage (w), the total number of jobs available (F), the size of the labor force (L), and the search cost function (c(p)). We can obtain an estimate of s from the existing literature on labor market dynamics. Ehrenberg (1980) and Murphy and Topel (1987) both provide estimates of the number of jobs that break-up in each period. If we measure time in 2-week intervals, their work suggests that s lies in the range of .007 to .013. For the interest rate we set r = .008 which translates into an annual discount rate of approximately 20%. Since our previous work suggests that results from this model are not sensitive to changes in r over a fairly wide range, this is the only value for the interest rate that we consider.

For F and L we begin by noting that our model is homogeneous of degree zero in F and L so that we may set L = 100 without loss of generality. If we then vary F holding all other parameters fixed we can solve for the equilibrium unemployment and vacancy rates. Abraham's (1983) work suggests that the ratio of unemployment to vacancies (U/V) varies between 1.5 and 3 over the business cycle. Although the actual values of U and V depend on the other parameters, we find that to obtain such values for U/V in our model F must lie in range of 95 to 97.5.

The remaining parameters are the wage rate and the search cost function. For these values we turn to our previous work, which makes use of data and results from the Illinois Reemployment Bonus Experiment (Davidson and Woodbury 1993, 1995). In the Illinois Reemployment Bonus Experiment a randomly selected group of new claimants for UI were offered a $500 bonus for accepting a new job within 11 weeks of filing their initial claim. The average duration of unemployment for these bonus-offered workers was approximately .7 weeks less than the average unemployment duration of the randomly selected control group (Davidson and Woodbury 1991). In our previous work, we estimated the parameters of the search cost function that would be consistent with such behavioral results. That is, we assumed a specific functional form for c(p) and then solved for the parameters that would make the model's predictions match the outcome observed in the Illinois experiment. The functional form that we used was $c(p) = cp^z$, where z denotes the elasticity of search costs with respect to search effort. Our results indicated that for the average bi-weekly wage rate observed in Illinois ($511), the values of c and z that are consistent with the Illinois experimental results are c = 282 and z = 1.269.[7]

In summary, our reference case uses the following parameter values: s = .010, r = .008, L = 100, F = 96.25, w = 511, c = 282, and z = 1.269. Once we have solved for the optimal value for x in the reference case, we vary s and F over the ranges described above to test for the sensitivity of our results with respect to each.

Table 1 summarizes the results of solving the model with infinite potential duration of benefits for the optimal bi-weekly UI benefit and the optimal replacement rate. For our reference case the optimal replacement rate -- the ratio of bi-weekly UI

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[7] As we show elsewhere (Davidson and Woodbury 1995), the Illinois bonus impact suggests that a 10 percentage point increase in the UI replacement rate lengthens the expected duration of unemployment by .8 week, and that a 1 week increase in the potential duration of benefits lengthens the expected duration of unemployment by .2 week. These are in the upper-middle of the range of existing estimates of the disincentive effects of UI.

benefits to the bi-weekly wage -- is .66.[8] For other values of the separation rate (s) and total available jobs (F), the optimal replacement rate varies from a low of .60 to a high of .74. This range falls between the optimal replacement rate estimates obtained by Baily (around .50) and Flemming (.75 in a model without borrowing or lending).

We obtain higher optimal replacement rates when s is low. Intuitively, when s is low, separations occur infrequently and the equilibrium unemployment rate is relatively low. With high employment, the government can afford to provide more generous assistance to the relatively few who are unemployed without generating a large tax burden for the employed. Also, we obtain higher optimal replacement rates when F low.

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[8] Remarkably, this rate is identical to the rate suggested by Hamermesh (1977) in his classic study of UI.

C. Optimal Replacement Rates with Limited Benefit Duration

We have argued that the optimal UI program entails offering benefits to unemployed workers indefinitely. Moreover, with savings ruled out and an elasticity of search with respect to UI benefits that is in the upper-middle of the range of existing estimates,[9] we find that the optimal replacement rate is roughly two-thirds. This result accords fairly well with some of the results reported in Baily (1978) and Flemming (1978). Baily finds an optimal replacement rate of approximately .50 when the elasticity of search effort with respect to UI benefits is relatively low. But his optimal replacement rate falls below .50 when this elasticity is high, which is the case he considers most relevant. In the end, he suggests that replacement rates in the United States, which designed to be about .50, are too high. Flemming finds that the optimal replacement rate is roughly .75 when agents cannot borrow or lend (as in our model). But he argues that when savings are incorporated into the model, the optimal replacement rate falls below .50. Thus, both authors strongly suggest that the existing UI programs in the United States are too generous.

As emphasized earlier, both Baily and Flemming assume that the potential duration of UI benefits is infinite. Although we have

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[9] Again, Illinois bonus impact, which was used to calibrate our model, suggests that a 10 percentage point increase in the UI replacement rate lengthens the expected duration of unemployment by .8 week, and that a 1 week increase in the potential duration of benefits lengthens the expected duration of unemployment by .2 week.

argued that such a policy is optimal, in reality every country that offers UI places a limit on the number of weeks of benefits that a worker may collect. This raises the following question: What is the optimal replacement rate when T = 26 (as in the United States), or T = 52 (as in Canada), or T = 104 (as in some European countries)? To answer this question, we return to the model introduced in sub-section A, fix T, and then solve for the optimal replacement rate.

The relationship between the optimal replacement rate (x/w) and the potential duration of UI benefits (T) in our reference case is depicted in Figure 2. The Figure reveals a striking finding of this exercise: for T < 32, the optimal replacement rate is 1. As T increases, the optimal rate falls fairly slowly, reaching .67 for T = 104. As T continues to increase, the optimal rate approaches .66 asymptotically.

Our model therefore suggests that if benefits are limited to 26 weeks, as is usually the case in the United States, the government should fully replace the lost earnings of UI-eligible unemployed workers during that limited period. This result suggests that unemployment insurance in the United States is sub-optimal. Either the potential duration of benefits should be increased substantially, or, if the potential duration of benefits is to remain limited, the replacement rate should be increased substantially.

Our basic conclusion -- that the existing UI system in the United States is too stingy -- is opposite that of Baily and

Flemming mainly because Baily and Flemming assume that UI benefits are provided in perpetuity, whereas we have examined optimal UI benefits under finite benefit duration. It is easy to see that the optimal UI replacement rate could never approach 1 if UI benefits were offered in perpetuity -- if full income replacement were offered indefinitely to unemployed workers, the unemployed would have no incentive to become reemployed and the economy would shut down. On the other hand, if the government were to offer full income replacement for only a limited time (say, 26 weeks), the unemployed would begin searching around the time their benefits were exhausted. The unemployment rate would not explode and the economy would not shut down. With full income replacement for 26 weeks, the unemployment rate <u>would</u> increase (to around 10% in our reference case, compared with 7% with a replacement rate of .5), but there would be a substantial smoothing of income that would increase the utility of all risk averse agents.

In summary, the assumption that the potential duration of UI benefits is unlimited in both the Baily and Flemming models leads to a basic misinterpretation of their results. Only if the government follows the optimal policy of offering UI benefits indefinitely is the optimal replacement rate as low as the values of .5 and below that Baily and Flemming report. If the potential duration of UI benefits is limited, then the optimal replacement rate is significantly higher.

IV. Extensions

A. UI-Ineligibles

Another assumption made by Baily and Flemming is that all unemployed workers are eligible to collect UI benefits. In reality this is not the case. Workers with a weak attachment to the labor force, new labor force entrants, and labor force reentrants are typically not eligible to collect benefits while unemployed. Blank and Card (1991) estimate that in the United States no more than 45% of the unemployed are UI-eligible.

Consideration of UI-ineligibles in the model can change the optimal replacement rate for two reasons. First, an increase in the generosity of the UI system will have a spill-over effect on the welfare of UI-ineligible workers. In general, a more generous UI system reduces the search effort of UI-eligible jobless workers. This reduction in search effort makes it easier for UI-ineligibles to find jobs and increases their expected lifetime utility. Once we take this spill-over effect into account, the optimal replacement rate rises.

Second, when we explicitly account for the fact that not all workers are UI-eligible, the reemployment probability faced by UI-eligible workers changes. Whether their reemployment prospects are brightened or dimmed depends on how hard UI-ineligibles search and the degree of substitutability in production between UI-eligible and UI-ineligible workers. For example, suppose that UI-eligibles and UI-ineligibles are considered close substitutes by firms and that UI-ineligibles search harder than UI-eligibles (since they receive no UI benefits). In this case, adding UI-ineligibles to the model will lower the reemployment probabilities faced by UI-eligibles and increase the desirable level of insurance (i.e., the optimal replacement rate will rise).

On the other hand, suppose that UI-ineligibles are low-skilled workers who do not vie for the same jobs as UI-eligible workers. In this case, treating all workers as if they are UI-eligible will overstate the difficulty that UI-eligibles will have in finding a job (since, in reality, there will be fewer workers vying for the jobs UI-eligibles seek than the model predicts). Since the presence of UI-ineligibles in the model makes it easier for UI-eligibles to find jobs, the level of insurance that the government needs to provide to UI-eligibles falls (i.e., the optimal replacement rate falls).

To investigate the size of these effects we add UI-ineligibles to a model in which the potential duration of benefits is unlimited and solve for the optimal replacement rate. The fundamental equations of the model as follows:

(1′) $$L = J + U$$

(2′) $$U = U_e + U_i$$

(3′) $$F = J + V$$

(4′) $$sJq = m_i U_i$$

(5′) $sJ(1 - q) = m_e U_e$

(7′) $m_j = p_j(V/F)(1/\lambda)[1 - e^{-\lambda}]$ for $j = i,e$

(9′) $\lambda = \{p_e U_e + p_i U_i\}/F$

(10′) $V_{wj} = w(1 - \tau) + [sV_j + (1 - s)V_{wj}]/(1 + r)$ for $j = i,e$

(11′) $V_e = x - c(p_e) + [m_e V_{we} + (1 - m_e)V_e]/(1 + r)$

(12′) $V_i = - c(p_i) + [m_i V_{wi} + (1 - m_i)V_i]/(1 + r)$

(13′) $p_j = \arg\max V_j$ for $j = i,e$

The subscripts e and i refer to UI-eligible and UI-ineligible workers. Thus, U_e and U_i are the numbers of UI-eligible and UI-ineligible workers seeking jobs in the steady-state equilibrium. The only new parameter is q (in equations 4′ and 5′), which is the fraction of the unemployed who are UI-ineligible.

As before, (1′)-(3′) are simple accounting identities. Equations (4′) and (5′) are the new steady-state equations -- (4′) equates the flows into and out of state U_e (UI-eligible unemployment) while (5′) equates the flows into and out of state U_i (UI-ineligible unemployment). Equation (7′) defines the reemployment probabilities for unemployed workers. Equation (10′)-(12′) define expected lifetime income for employed and unemployed workers. Note that in each case, a separate definition is provided for UI-eligible and UI-ineligible workers. Finally, (13′) defines optimal search effort.

The government's problem is the same as before, except that Social Welfare must now include the expected lifetime income of UI-ineligible workers as well.

It is important to note that in the above model the only difference between UI-eligible and UI-ineligible workers is that the UI-eligibles receive benefits while unemployed. That is, in this model firms consider the two types of workers good substitutes in production, and in equilibrium UI-ineligibles search harder than UI-eligibles (since UI-ineligibles receive no benefits).

An alternative to assuming that UI-eligibles and UI-ineligibles are good substitutes who compete for the same jobs is to assume that they are poor substitutes. We accomplish this by assigning UI-ineligibles a low reemployment probability that is unaffected by the behavior of UI-eligibles. That is, we replace (13′) for j = i with:

$$(14) \qquad p_i = \beta$$

where β takes some low value. Assigning a low reemployment probability to UI-ineligibles captures the notion that UI-ineligibles do not compete for the same jobs as UI-eligibles -- that is, they are poor substitutes for UI-eligibles.

We solve the model under the two alternative assumptions about substitutability between UI-eligibles and UI-ineligible and compare the results. Table 2 shows the optimal replacement rate under various assumptions about turnover (s) and the total number of jobs available (F), and assuming that UI-ineligibles and UI-eligibles are close substitutes. The only new parameter in the model is q, the proportion of unemployed workers who are UI-ineligible. Based on Blank and Card (1991), we consider q = .6 the most likely case, but report the optimal replacement rate for other values of q for comparison.

Table 2 shows that accounting for the fact that some workers are ineligible for UI increases the optimal replacement rate. In our reference case the optimal replacement rate rises from .66 when there are no UI-ineligibles to .74 when 60% of the unemployed are UI-ineligible. The optimal replacement rate also increases with q for the other cases considered in Table 2. Thus, assuming that all workers are eligible for UI (as we did above and as Baily and Flemming did) tends to bias downward estimates of the optimal replacement rate.

The intuition behind this result was described above. If all workers are assumed to be UI-eligible, the model cannot take into account the positive spill-over effect of UI on UI-ineligibles (that is, UI benefits improve the well-being of UI-ineligibles). Also, the model will overstate the reemployment prospects of UI-eligibles unless UI-eligibles and UI-ineligibles are very poor substitutes in production. Accounting for either of these effects results in a higher optimal replacement rate.

Consider now the case in which UI-eligible and UI-ineligible workers are not close substitutes. To solve for the optimal replacement rate in this case we need to choose a value for β, the search effort of UI-ineligibles. As β falls, the reemployment prospects of UI-eligibles brighten and less insurance is needed -- that is, as β falls, the optimal replacement rate falls. If β is low enough, adding UI-ineligibles to the model could actually lower the optimal replacement rate. That is, the positive effect of a low β on UI-eligible reemployment probabilities could outweigh the spill-over effect of UI on the well-being of UI-ineligibles.

The question now is, how low a value of β would be needed to leave the optimal replacement rate equal to what it would be in a model in which all workers are UI-eligible? For each of the cases shown in Table 2, we solve the model for the value of β that For all of the cases we have checked, the result is that β would have to be approximately 15% of the value that it would have been in the first model -- that is, in order for the optimal replacement rate to remain constant when UI-ineligibles are added to the model, UI-ineligibles would have to face a reemployment probability that is roughly 85% lower than the reemployment probability they face in the model in which UI-eligibles and UI-ineligibles are close substitutes. Thus, the degree of substitutability between UI-eligibles and UI-ineligibles would have to be extremely low for the optimal replacement rate to fall when UI-ineligibles are added to the model.

B. Savings

In our model workers are not allowed to save. This biases our estimates of the optimal replacement rate upwards since agents cannot self-insure against unemployment by saving during periods of employment. Extending our model to allow for savings is not straightforward -- we would have to choose a specific form for the utility function, model the capital market, and recalibrate the model to obtain estimates of the search cost parameters. Fortunately, we can say something about the effect of extending our model to include saving without actually going through the exercise. First, it should be clear that our basic result -- that the optimal potential duration to UI benefits is infinite -- would continue to hold even in a model where workers could save. Unless capital markets were perfect, agents would never save enough while employed to fully smooth consumption across periods of unemployment.[10] Thus, the qualitative nature of Figure 1 would continue to hold with savings in the model -- the vertical axis can simply be relabeled "present consumption." Extending benefits in a tax neutral manner will lower present consumption in the "good" states of unemployment (when present consumption is relatively high) and increase it in the most adverse states. It follows that it will still be optimal to offer UI indefinitely.

Second, since it is optimal to offer UI benefits indefinitely, and since Baily and Flemming allowed for savings in their models,

[10] As noted in the introduction, the empirical evidence is clear on this issue -- consumption does fall during periods of unemployment.

we can refer to their work to gauge how our results might be altered by allowing workers to save. Consider first Baily's findings. In a two-period model in which agents can save in the first period of life, he finds that the optimal replacement rate falls between .33 and .50, depending on the elasticity of search effort with respect to UI.[11] If the elasticity is low, the optimal replacement rate is close to .50. If the elasticity is high, the optimal replacement rate falls to .33.[12] Our results suggest that if Baily were to include UI-ineligibles in his model, his optimal replacement rates would rise by about 8 to 10 percentage points. Thus, combining our results with Baily's suggests that if workers can save, the optimal replacement rate will lie somewhere between .40 and .60. This rate is optimal, however, only if the potential duration of UI benefits is infinite.

Consider next Flemming's results. Flemming develops a model with infinitely lived agents and allows for varying degrees of capital market imperfections. If agents cannot borrow or lend, his model yields an optimal replacement rate of around .70. If capital markets are perfect, the optimal replacement rate lies in the range of .10 to .20.[13] Our results suggest that adding UI-ineligibles to

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[11] Baily makes a reasonable assumption about the degree of risk aversion -- specifically, that all agents have the same constant value of absolute risk aversion, and that this value is one.

[12] See his Table 2, column 2, rows 2 and 3.

[13] See his Tables 1 and 3 under the column d^* (for the "optimal dole").

Flemming's model would boost these rates by about 8 to 10 percentage points, yielding a range of .25 to .80 for the optimal replacement rate. However, we can probably rule out the extreme values since they are based on extreme assumptions -- capital markets do exist, but they are not perfect. This leaves us with optimal replacement rates quite similar to those discussed in the previous paragraph -- that is, .40 to .60. Again, it is important to emphasize that these rates are optimal only if the potential duration of UI benefits is infinite.

We conclude that if workers are allowed to save during periods of employment, the optimal replacement rate falls to a level that is consistent with existing average statutory rates in the United States. Hence, the current level of UI benefits would appear to be about right <u>if the potential duration of benefits were infinite</u>. But the current potential duration of benefits -- 26 weeks in most states in nonrecessionary times -- appears to be too short.

V. Discussion, Caveats, and Conclusions

Our results suggest that the structure of the existing UI system in the United States is sub-optimal. Most existing state systems limit the potential duration of UI benefits to 6 months, whereas insurance considerations suggest that it would be better to provide an unlimited potential duration of benefits (see section III.A). Also, most states' UI systems pay replacement rates on the order of .5 to most workers. But only when the potential duration of benefits is unlimited are UI replacement rates even as low as two-thirds optimal (section III.B). When the potential duration of benefit is limited to 32 weeks or less, insurance considerations suggest that an optimal replacement rate of 1 would be optimal (section III.C).

A likely objection to the finding that an infinite potential duration of benefits is optimal is that, if benefits were inexhaustible, then workers would never return to work. It is true that increasing the potential duration of benefits would lead workers to remain unemployed longer and would lead to a higher unemployment rate. In our model, increasing the potential duration of UI benefits from 6 months to infinity with a UI replacement rate of .5 would raise the unemployment rate from 7% to 10% (see section III.C). Raising the replacement rate to 1 (from existing levels around .5) would, similarly, increase the length of unemployment spells and increase the unemployment rate. But a higher unemployment rate is not a shut-down of the economy -- workers would not collect UI benefits paying a replacement rate of .5 (or .67) forever. Moreover, the increase in the unemployment rate would result from voluntary behavior, not from economic hard times, and would connote an improvement in workers' well-being.[14]

The model used to derive these conclusions is set out in section II, and extends earlier work by Baily (1978) and Flemming (1978) in two ways. First, whereas Baily and Flemming assumed that UI benefits are offered indefinitely, we consider a UI system in which the potential duration of benefits is limited to 26 weeks, as in the United States. We find that the optimal UI replacement rate under such a system is 1, rather than .75 or less, as Baily and Flemming suggested (see sections III.C).

Second, we consider how the optimal UI replacement rate is affected by the presence of workers who are ineligible for UI (section IV.A). This is important because fewer than half of all unemployed workers in the United States are UI-eligible. Adding UI-ineligibles to the model has two effects. The first is a positive spill-over effect that increases the optimal UI replacement rate: Since UI benefits reduce the search intensity of UI-eligible workers, UI-ineligibles face less competition for jobs when UI benefits are higher. The second concerns the substitutability in production between UI-eligible and ineligible workers. If UI-eligibles and UI-ineligibles are substitutes, then the presence of UI-ineligibles makes it harder for UI-eligibles to

[14] Increased unemployment, when it is in part increased in leisure, is hardly a bad thing. This point is made in an unusually entertaining way by Landsburg (1993).

find reemployment. (UI-ineligibles presumably search harder than UI-ineligibles because they receive no UI benefits.) Ignoring the presence of UI-ineligibles leads to an overstatement of the reemployment prospects for UI-eligible workers, and the optimal UI replacement rate needs to be increase to compensate. In general, then, the presence of UI-ineligibles in the workforce increases the optimal replacement rate.[15]

In section IV.B we consider the effects of adding voluntary saving to the model. If workers are able to save, then the optimal replacement rate falls by about 10 percentage points (for example, from .6 tp .5). But allowing workers to save would not alter our conclusion that an infinite potential duration of benefits is optimal.

In the model developed in section II, we assume that UI-eligible workers are homogeneous, that the disincentive effects of UI benefits are in the upper-middle of the range of effects that have been estimated, and that workers are unable to save. [Is there any way of saying something about the degree of risk aversion?] We have argued that the results are not especially sensitive to the savings assumption -- in particular, the finding that the optimal duration of UI benefits is unlimited holds even if

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[15] We also consider the case in which UI-ineligibles are lower-skilled workers who are poor substitutes for UI-eligible workers. In this case, the presence of UI-ineligibles in the workforce lowers the optimal replacement rate (i.e., less insurance is needed). Nevertheless, unless the degree of substitutability between UI-eligibles and UI-ineligibles is extremely low, the presence of UI-ineligibles raises the optimal replacement rate on net.

savings are allowed (section IV.A). Also, we believe that the assumptions about the disincentive effects of UI are reasonable and well-informed. However, we have not investigated whether results are sensitive to the assumption of worker homogeneity.

Worker heterogeneity could be considered in a number of ways. One approach would be to suppose that some UI-eligible workers face a high probability of layoff with a low expected duration of unemployment (blue-collar production workers), while others might face a low probability of layoff with a longer expected duration of unemployment (white-collar non-production workers). Another approach might be to suppose that some UI-eligible workers are strongly attached to the labor force (as most appear to be), but that a significant minority are weakly attached to the labor force. Whether an unlimited potential duration of benefits would remain optimal in a model that accounts for one or both of these types of heterogeneity is an open question.

References

Abraham, Katherine G. "Structural/Frictional vs. Deficient Demand Unemployment: Some New Evidence." American Economic Review 73 (September 1983): 708-724.

Baily, Martin N. "Some Aspects of Optimal Unemployment Insurance." Journal of Public Economics 10 (1978): 379-402.

Blank, Rebecca M. and David E. Card. "Recent Trends in Insured and Uninsured Unemployment: Is there an Explanation?" Quarterly Journal of Economics 106 (November 1991): 1157-1189.

Davidson, Carl and Stephen A. Woodbury. "Effects of a Reemployment Bonus under Differing Benefit Entitlements, or, Why the Illinois Experiment Worked." Manuscript, Department of Economics, Michigan State University, and W.E. Upjohn Institute for Employment Research, March 1991.

Davidson, Carl and Stephen A. Woodbury. "The Displacement Effect of Reemployment Bonus Programs." Journal of Labor Economics 11 (October 1993): 575-605.

Davidson, Carl and Stephen A. Woodbury. "Further Lessons from the Reemployment Bonus Experiments." Manuscript, Department of Economics, Michigan State University, and W.E. Upjohn Institute for Employment Research, December 1994.

Davidson, Carl and Stephen A. Woodbury. "Unemployment Insurance and Unemployment: Implications of the Reemployment Bonus Experiments." Manuscript, Department of Economics, Michigan State University, and W.E. Upjohn Institute for Employment Research, February 1995.

Ehrenberg, Ronald G. "The Demographic Structure of Unemployment Rates and Labor Market Transition Probabilities." Research in Labor Economics 3 (1980): 214-291.

Flemming, J. S. "Aspects of Optimal Unemployment Insurance." Journal of Public Economics 10 (1978): 403-425.

Gruber, Jonathan S. "The Consumption Smoothing Benefits of Unemployment Insurance." Manuscript, Department of Economics, MIT, May 1994.

Hamermesh, Daniel S. Jobless Pay and the Economy. Baltimore, MD: Johns Hopkins University Press, 1977.

Landsburg, Steven E. "How Statistics Lie: Unemployment Can Be Good for You." In The Armchair Economist: Economics and Everyday Life. New York: The Free Press, 1993. Pp. 127-137.

Murphy, Kevin M. and Robert H. Topel. "The Evolution of Unemployment in the United States: 1968-1985." in NBER Macroeconomics Annual 1987, edited by Stanley Fischer. Cambridge, MA: MIT Press, 11-58.

O'Leary, Christopher J. "The Adequacy of Unemployment Insurance Benefits." Manuscript, W.E. Upjohn Institute for Employment Research, August 1994.

Organization for Economic Cooperation and Development (OECD). "Unemployment Benefit Rules and Labour Market Policy." OECD Employment Outlook (July 1991): 199-236.

Figure 1
The Optimal Potential Duration of UI Benefits Is Unlimited

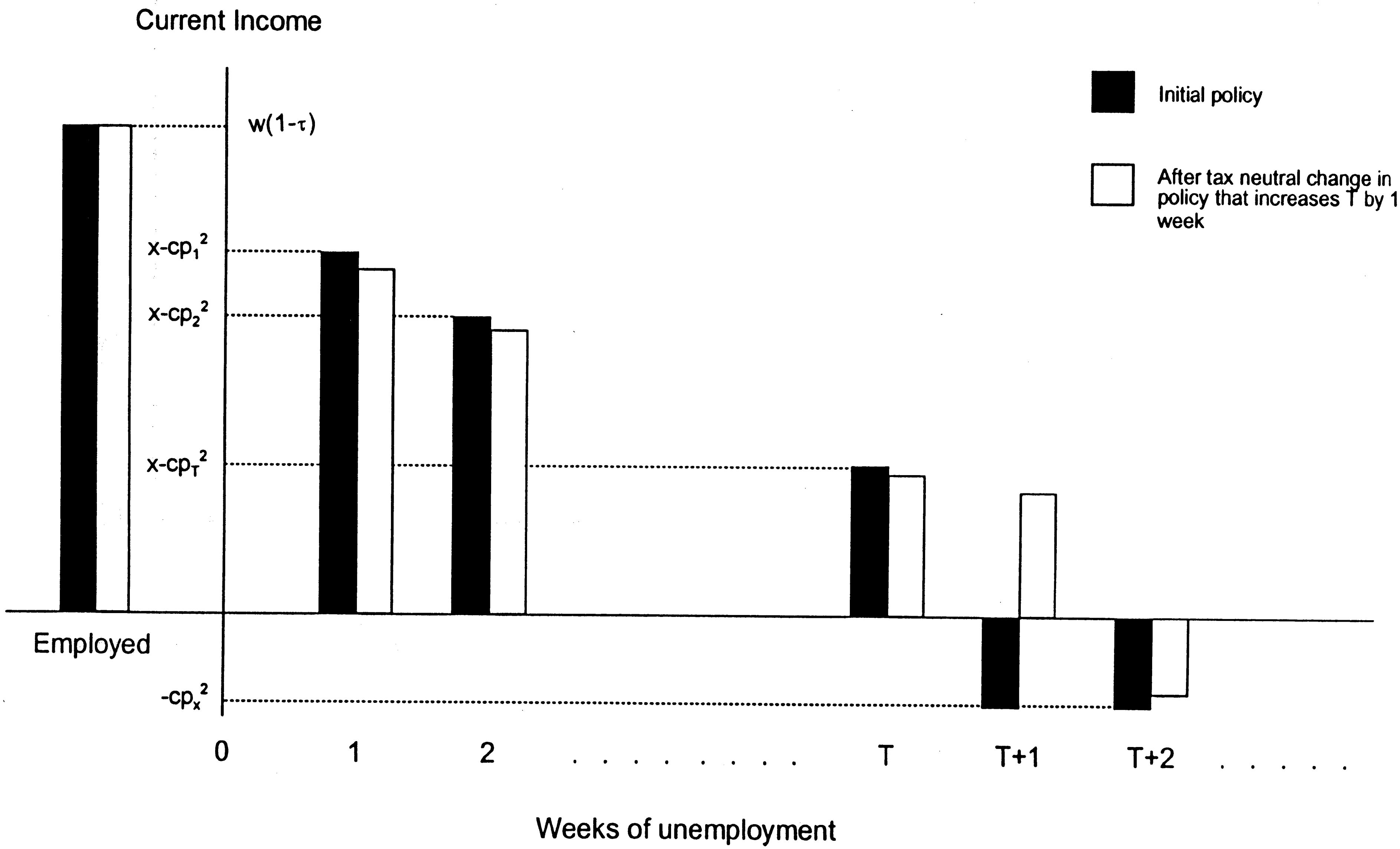

Figure 2

The Optimal UI Replacement Rate Increases as the Potential Duration of UI Benefits is Reduced

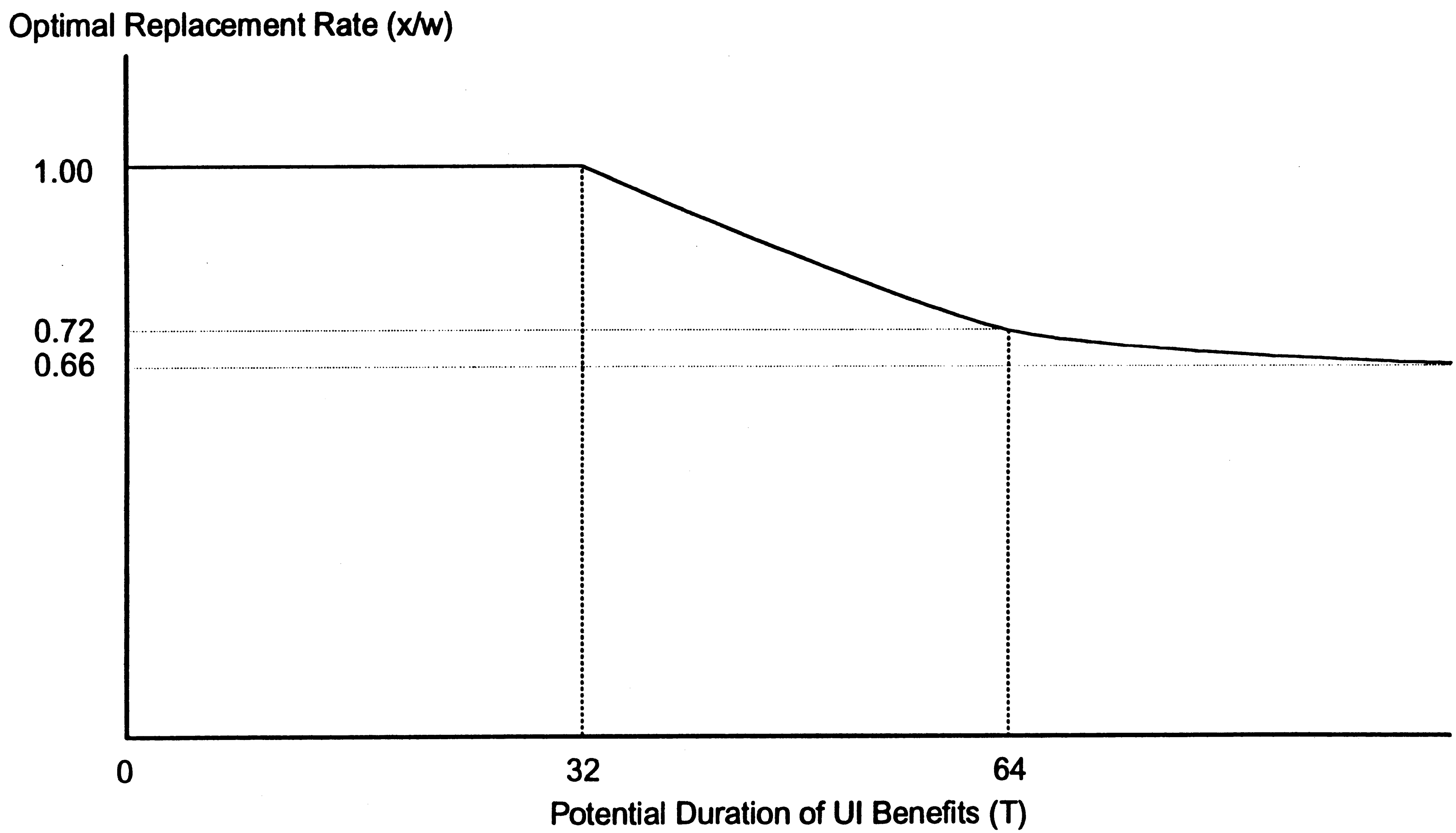

Table 1

Optimal Unemployment Insurance Benefits and Replacement Rates
under Various Assumptions,
Model with Infinite Potential Duration of UI Benefits

Assumptions	Optimal bi-weekly UI benefit (x)	Optimal replacement rate (x/w)
Reference case (s=.010, F=96.25)	335	.66
Low turnover (s=.007, F=96.25)	380	.74
High turnover (s=.013, F=96.25)	305	.60
Fewer total jobs available (s=.010, F=95)	356	.70
More total jobs available (s=.010, F=97.5)	317	.62

Note: Parameter values in the reference case are as follows: separation rate (s) = .010; total jobs available (F) = 96.25; labor force (L) = 100; bi-weekly interest rate = .008; bi-weekly reemployment wage = $500; search cost parameter (c) = 282; z = 1.269.

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Table 2

Optimal UI Replacement Rates When Some Workers Are Ineligible for UI, Various Assumptions, Model with Infinite Potential Duration of UI Benefits

	Proportion of unemployed workers ineligible for UI (q)				
	0	.15	.30	.45	.60
Reference case (s=.010, F=96.25)	.66	.67	.69	.72	.74
Low turnover (s=.007, F=96.25)	.74	.75	.77	.79	.81
High turnover (s=.013, F=96.25)	.60	.62	.64	.67	.70
Fewer total jobs available (s=.010, F=95)	.70	.71	.73	.75	.77
More total jobs available (s=.010, F=97.5)	.62	.64	.66	.69	.72

Notes: See Table 1. The results shown are from a model in which UI-eligibles and UI-ineligibles are good substitutes. Optimal replacement rates can fall below those shown in the table if UI-eligibles and UI-ineligibles are sufficiently poor substitutes.

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Further Aspects of Optimal Unemployment Insurance

by

Carl Davidson* and Stephen Woodbury**

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* Professor of Economics, Michigan State University.

** Professor of Economics, Michigan State University and Senior Economist, The W.E. Upjohn Institute for Employment Research

1. Introduction

Moral hazard is defined to be a situation in which "one party to a transaction may undertake certain actions that (a) affect the other party's valuation of the transaction but that (b) the second party cannot monitor/enforce perfectly" (Kreps, p. 577). Unemployment insurance is a classic example of moral hazard -- the government would like to provide a social safety net for those who are currently jobless but seeking reemployment. Unfortunately, the government cannot monitor perfectly the effort put forth by the unemployed to find new jobs. Thus, there is a tradeoff -- if the government provides too much insurance, the unemployed will not work hard enough to find new jobs, but, if too little insurance is provided the unemployed will bear too much risk. In devising an optimal unemployment insurance program, the government must find a way to provide adequate insurance without substantially reducing the incentive to seek employment.

The current UI program in the U.S. provides a benefit equal to roughly 50% of the wage earned on the previous job for one-half of a year after a worker loses her job. There are at least four relevant lines of literature that have been devoted to assessing whether this program is structured correctly and whether the current level of generousity is adequate. The purpose of this paper is to offer a brief critical review these literatures and to extend our previous work (Davidson and Woodbury 1995) on this issue. The paper divides into four additional sections. In section 2, we review three areas of the literature that deal explicitly with the issue of unemployment insurance. Section 3 provides a description of our model. Our previous results are reviewed and our new results are presented in Section 4. Finally, in section 5 we relate our results to the previous literature, compare them with insights that have been provided by the abstract literature on optimal insurance contracts, and discuss future extensions. We close the paper with a conjecture as to the structure of an optimal unemployment program that is radically different from the present system.

2. The Literature

There are at least four relevant strands of literature that have investigated aspects of an optimal insurance program in the presence of moral hazard. The first three -- labor economics, macroeconomics, and public economics -- use similar approaches. They all adopt search models of the labor market in which unemployed workers choose search effort to maximize expected utility. More generous unemployment insurance increases the insurance offered to the unemployed, but also lowers optimal search effort, thereby triggering an increase in unemployment. Although the approaches are similar, these literatures seem to have developed, for the most part, independently. Thus, it is not surprising that they differ in the questions that are addressed, the complexity of the models, and the assumptions that are used to simplify the analysis. The purpose of this section is to provide a critical review of the contributions in each area.

In Section 5, we review work in a related area -- the abstract literature on optimal insurance contracts -- that does not directly deal with unemployment insurance. At that point, we discuss how the results from that related literature can be extended to provide insights concerning an optimal unemployment insurance program. We also combine the insights from the four literatures with own results to derive an unemployment insurance program that is fundamentally different from our current system, but which we believe makes more sense than the current one from an economic perspective.

A. Labor Economics

Perhaps the best known article on optimal unemployment insurance in the labor economics literature is Shavell and Weiss' 1979 paper in the Journal of Political Economy. This article addresses the following question -- given that the government is going to spend a *fixed* amount of money on unemployment compensation, how should the benefits be paid out to the unemployed? That is, how should benefits vary over the spell of unemployment? Note that this paper does not attempt to determine the optimal size of the program -- the generousity of the program is taken as given and fixed.

The authors consider a variety of models in order to indicate how different features of the model affect their results. Their basic approach is similar to that described above in that they use a search model of the labor market. However, in some of the cases that they consider they do not allow agents to alter search effort. This allows them characterize the optimal benefit path when moral hazard is not an issue. When they do allow search effort to vary, they assume that unemployed workers choose search effort to maximize expected lifetime utility and that greater search effort, while costly, increases the probability of finding employment. In all of their models unemployed workers are assumed to be identical. In addition, labor demand is not modeled and the wage rate is exogenous and independent of the UI program adopted. Finally, in each case, the benefit path over the spell of unemployment is chosen to maximize the expected lifetime utility of a representative unemployed worker.

Shavell and Weiss derive several results, depending on the assumptions of their model. For our purposes, there are three results that are important. The first result concerns the optimal benefit path when workers (a) cannot save and (b) cannot alter search effort so that they cannot affect their probability of reemployment. Thus, workers cannot self-insure and there are no moral hazard concerns. In this case, it is optimal to offer the same benefit rate in each period of unemployment. The logic is simple. Risk averse agents wish to smooth consumption across time. If agents cannot save, the only way provide a smooth path of consumption across the spell of unemployment is to make the benefit independent of the number of weeks a worker has been unemployed.

And, if agents cannot affect the probability of finding employment, there are negative side effects of such a UI program.

The second result concerns the optimal benefit path when agents can save but cannot affect their probability of reemployment. Thus, self-insurance is possible, but there are still no moral hazard issues to deal with. In this case, the optimal benefit rate is lowest in the initial stages of unemployment and rises over the spell of unemployment. As the spell lengthens, the benefit rate approaches an upper bound asympototically. Thus, benefits are offered indefinitely. The intuition for this result is straightforward. If agents can save while employed, then during the initial stages of unemployment they can smooth consumption by dissaving. However, as the spell of unemployment lengthens, savings are depleted, and the only way to maintian consumption is for the government to increase the benefit level. As before, if agents cannot affect their reemployment probabilities, then their are no negative side effects from this program.

Shavell and Weiss' last result describes the optimal benefit path when agents can affect their probability of reemployment but are unable to self-insure against employment risk. They show that due to moral hazard concerns, benefits should decline over the spell unemployment. The reduction in benefits induces workers to put forth effort to become reemployed. In the limit, the benefit converges to zero.

Unfortunately, Shavell and Weiss are unable to characterize the optimal benefit path when agents can save and can also affect their reemployment probabilities. However, the three results discussed above can be used to form a conjecture as to the optimal benefit path in this case. With savings, agents can maintain consumption in the early stages of unemployment without receiving benefits. Thus, providing high benefits in the early stages of unemployment would not be wise, since doing so would only serve to lower search effort and increase unemployment. As the spell lengthens and savings are depleted, the government must start to increase benefits in order to allow the unemployed to smooth consumption. However, increasing benefits too much or providing them for too long would have an adverse effect on search effort and unemployment. Thus, eventually the benefit rate must fall and converge to zero (a typical benefit path of this nature is depicted in Figure 1). It is important to emphasize that Shavell and Weiss' analysis provides no insight as to the optimal level of benefits or the point at which benefits should be cut-off, they are only concerned with the shape of the benefit path.

Several years after the publication of Shavell and Weiss, Hausman (1984) argued that it was possible to improve upon the type of UI program that they had advocated. He argued that by offering a large up front payment to newly unemployed workers followed by low (or zero) benefits during the spell of unemployment, the system would operate more efficiently. The reasoning behind this scheme is that the up front payment would provide the unemployed the funds necessary to smooth consumption while the low benefit payments during the spell of unemployment would provide a strong incentive to seek and accept reemployment. As in Shavell and Weiss, Hausman makes no attempt to determine the optimal size of the initial payment nor the optimal potential duration of benefits.

Both the Shavell and Weiss and Hausman analyses were largely theoretical. There have also been two important recent empirical investigations of the current U.S. program in the labor economics literature. In 1994 O'Leary used a consumer theory approach to estimate the optimal benefit path. His basic finding was that with the current U.S. program short spells of unemployment are over-compensated while long spells are under-compensated. Note that this result is similar to what one might conclude by comparing the current system with Figure 1.

In an even more recent paper, Hamermesh and Slesnick (1995) compare the well-being of UI recipeints with their counterparts who do not receive benefits. They conclude that since their welfare levels are similar, the current system provides the right level of insurance.

With the exception of O'Leary (1994), all of these papers attempt to analyze the UI system by focusing on its impact on the typical unemployed UI recipient. While this may seem reasonable at first, it ignores the costs of the program. If a more generous program increases the unemployment rate, it increases the tax burden on the employed for two reasons. First, it costs more to fund a more generous program. Second, with higher unemployment there are fewer employed workers to share the tax burden. Thus, it is important to investigate the impact of different programs on the unemployment rate -- which is something that these papers do not attempt to do. In short, these papers focus on the insurance aspects of unemployment insurance without paying adequate attention to the costs of the program.

B. Macroeconomics

Over the past five years it has become fashionable in macroeconomics to blame a large part of society's economic ills on unemployment insurance. It is argued that the disincentive effect of UI are so strong that they have lead to a significant increase in the unemployment rate throughout Europe (see, for example, Layard, Nickell and Jackman 1991). There have also been claims that the current U.S. unemployment insurance program generates a large welfare loss for the U.S economy (see, for example, Mortensen 1994).

In a recent book, Layard, Nickell and Jackman (1991) trace much of the recent European experience with unemployment to changes in UI programs in the European countries. They argue that the gradual increase in the "natural rate" of unemployment in several European countries can be explained by the increased generousity of their UI programs. In addition, they argue that much of the cross-country differences in unemployment can be attributed to differences in their UI programs. In fact, they estimate that approximately 91% of the variation in the 1983-88 unemployment rate averages across the major OECD industrial countries can be explained by nothing more than the variation in the generousity of labor market policies and the extent of collective bargaining coverage.

Based on their results, Layard et al suggest a variety of reforms to combat Europe's dual problems of high unemployment and long average duration of unemployment. For example, with respect to the U.K. they suggest reducing the unemployment benefit period, discarding policies that impose firing costs on firms, and instituting subsidies to offset recruiting and training costs incurred by firms.

The purpose of the Layard, Nickell and Jackman book is to provide estimates of the impact of various labor market policies on unemployment and to suggest reforms. However, the authors make no attempt to link the employment effects that they estimate to measures of economic welfare. Thus, it is difficult to assess whether or not European UI programs are welfare enhancing or debilitating. In addition, their analysis provides no guidance as to how the reforms they suggest would improve matters when compared to the present programs.

In two recent papers, Mortensen (1994) and Millard and Mortensen (1994) attempt to improve on the Layard et al approach by estimating the welfare effects of a variety of labor market policies including unemployment insurance. As opposed to the labor economics literature, they use a <u>general equilibrium</u> search model to carry out their analysis so as to capture the cost of UI through its impact on the aggregate unemployment rate. There are two primary reasons that UI generates economic costs (in addition to the tax burden it creates). First, as we have already discussed, more generous UI lowers the opportunity cost of unemployment resulting in lower search effort by the jobless. This increases the equilibrium rate of unemployment and reduces output. Second, since more generous UI makes the unemployed less likely to accept new jobs, the wage that firms must offer rises, making production less profitable. This decreases the total number of jobs available in the economy. This job destruction effect further lowers employment, production, and welfare. This latter effect is absent from all of the labor literature discussed in sub-section A since the authors do not employ equilibrium models nor do they model firm behavior.

For our purposes, the most important results from these papers concern the UI programs in the U.S. and the U.K.. To estimate the impact of these programs, the authors calibrate their model using data on labor market flows in the U.S. during the period covering 1983-1992 and estimates of key parameters that are obtained from the labor economics and macroeconomics literatures. Following Layard et al, they then recalibrate the model for the U.K. assuming that differences in the U.S. and U.K. unemployment experiences can be attributed to differences in their labor market policies and union coverage rates.

In both papers welfare is measured by aggregate income net of search, recruiting and training costs. With this measure, Mortensen (1994) estimates that a 50% reduction in the U.S.

replacement rate would reduce the equilibrium rate of unemployment by 1.48 percentage points and increase net output by slightly less than one percentage point. He also estimates that a 50% reduction in the potential duration of benefits would decrease the equilibrium rate of unemployment by .78 percentage points while increasing welfare by about .5 percentage points.

As for the U.K., Millard and Mortensen estimate that the welfare cost imposed on the U.K. by its current UI program is roughly equal to 1.7% of net output, a fairly large measure for dead weight loss. They also estimate that by limiting the benefit period to 2 quarters (as in the U.S.), the U.K. could increase welfare by more than one percentage point (and lower unemployment by over 2 percentage points). Moreover, if the firing costs currently imposed by the government were also eliminated (as suggested by Layard et al), Mortensen and Millard estimate that welfare in the U.K. would rise by as much as 3.5%.

It is easy to infer from these results that the current Ui programs in the U.S. and the U.K. impose significant welfare burdens on their economies. However, there is at least one serious drawback to these analyses. By using aggregate net income as their measure of welfare, the authors implicitly assume risk neutrality on the part of workers so that there is no need or desire for insurance of any kind. It follows that the positive aspects of UI -- the fact that it provides desired insurance against employment risk -- are given no weight in the welfare calculations. In contrast to the labor literature which focused on the insurance aspects of UI without measuring the economic costs of the program, these two papers focus on the costs of the program while ignoring the benefits it provides.

A recent paper by Wang and Williamson (1995) improves upon the Mortensen and Millard and Mortensen analyses by explicilty incorporating risk aversion into a general equilibrium model. In that paper, welfare is measured by summing the utilities of all the agents in the economy. Since each agent is risk averse, there is a desire for employment insurance, and, since a general equilibrium model is used, the authors are able to measure the impact of UI programs on aggregate unemployment. Thus, Wang and Williamson use an approach that measures both the benefits and costs of different UI programs. It is important to note, however, that this is not the only difference between the Millard/Mortensen and Wang/Williamson papers -- Wang and Williamson do not adopt a search framework, choosing to work instead in an abstract framework in which the process by which jobs are created and destroyed are not modelled. We discuss the importance of this difference in approach in Section 5.

The purpose of the Wang and Williamson paper is to derive the optimal unemployment insurance program assuming that the replacement rate can vary over the spell of unemployment and that the government can tax and/or subsidize transitions into various labor market states. Thus, they allow for extremely complex programs. In fact, the program that they find to be optimal is so complex that it is hard to imagine any government actually trying to implement it. In brief, they find that the replacement rate should vary non-monotonically with the spell of unemployment -- starting low and then rising before falling off eventually to zero. Thus, their optimal benefit path is similar to what we conjectured the optimal path would look like in the Shavell and Weiss analysis when agents can save and affect their reemployment probabilities (see Figure 1). In addition, they find that the government should subsidize transitions into employment (with, for example, a reemployment bonus).

Although Wang and Williamson use an approach that is quite different from ours (since they do not use a search model and do not include firms in their analysis) and although their optimal UI program is far more complex then any program that we allow the government to consider, their results share many of the important features of our optimal program. Therefore, in Section 5 we describe their results in greater detail and compare them with ours.

C. Public Economics

The two most heavily cited papers on optimal unemployment insurance appeared in the same 1978 issue of the Journal of Public Economics. These papers were written by Martin N. Baily and J.S. Flemming and were so similar in approach and conclusions that they were given almost identical titles. Both authors use a search model of the labor market in which unemployed agents choose search effort to maximize expected lifetime utility. Agents are risk averse, so that insuance is desired, and an equilibrium model is used in order to capture the impact of UI on unemployment. However, neither author explicitly models firm behavior so that neither paper is able to capture the job destruction effects of UI. This implies that all of the increase in unemployment from UI is due to its impact on search effort.

The papers differ in the time horizon that is considered (Baily uses a two-period model while Flemming uses an infinite horizon approach), the manner in which the capital market (and thus, savings) is handled, and the utility function that is used. Nevertheless, as we discuss below, they derive remarkably similar results.

Both authors have the same goal -- to determine the optimal replacement rate assuming that the rate remains constant over the spell of unemployment. The results are then compared to replacement rates offered in the U.S. and the U.K. in order to determine whether or not current UI programs are too generous. Briefly, Baily and Flemming both find that if agents cannot save then the optimal replacement rate lies in the 60%-70% range. This result is fairly robust, since it does not depend on the time horizon or the manner in which the authors calibrate their models. There is one exception -- this result does depend on the degree of risk aversion that is assumed. Baily assumes that the Arrow-Pratt measure of relative risk aversion is constant and equal to one, while Flemming assumes that the Arrow-Pratt measure of absolute risk aversion is constant and equal to one. For lower measures of risk aversion, they find lower optimal replacement rates.

When agents can save but capital markets are imperfect (so that workers can only partially self-insure), Baily and Flemming find that the optimal replacement rate falls by about 25-30 percentage points. Thus, they conclude that the optimal replacement rate is below 50% and that the current U.S. unemployment insurance program is too generous. Similar conclusions have been reached by Gruber (1994) who recently used Baily's framework to estimate the optimal replacement rate for the U.S..

In our earlier work, Davidson and Woodbury (1995b), we criticized Baily and Flemming for two of the assumptions that they used in their analysis -- both authors assume that all unemployed agents are eligible for UI benefits and that they receive such benefits for as long as they remain unemployed. In reality, less than 50% of the unemployed are eligible for UI benefits in the U.S. (Blank and Card 1991) while in the U.K. roughly 70% of the unemployed are eligible (Layard et al 1991). In addition, benefits are offered for only 26 weeks in the U.S. and are limited in almost every other country. In section 4, we review our earlier results which indicate that the conclusions reached by Bailey and Flemming are extremely fragile with respect to these two assumptions. We then go on to extend the Baily and Flemming analysis even further by explicitly modelling firm behavior and making the wage rate and the number of active firms endogenous. This allows us to capture the job destruction effects of UI and see exactly how this alters our results.

3. Our Model

In this section we provide a description of the model that we use to derive the optimal UI program. As we describe our model, we also point out the elements that are missing from each of the analyses described in Section 2. This should help clarify some of our criticisms of the earlier literature.

We follow the tradition in this literature by employing a search model of the labor market. In order to focus on the benefits and costs of UI we model the behavior of a representative unemployed worker who is searching for employment and desires employment insurance. This worker earns a wage of w while employed and collects UI benefits of x while unemployed provided that she has not exhausted her benefits. Benefits are provided by the government to jobless workers who have been unemployed for no more than T periods. Thus, at the outset we assume that all newly unemployed workers are eligible for UI. In the next section, we describe how the model is modified to take into account the fact that the actual UI take-up rate is below 100%.

In our model, UI is funded by taxing all employed workers' incomes at a constant rate τ. This assumption, common in the optimal UI literature, is used to capture the notion that in a competitive economy the incidence of a UI tax is likely to be borne by workers.

We assume that unemployed workers choose search effort (p) to maximize expected lifetime income and that all workers are infinitely lived. As for firms, we assume that each firm hires at most one worker and that new firms enter the labor market until the expected profit from creating a vacancy is zero. Once a firm with a vacancy and an unemployed worker meet, they negotiate the wage. Following a well-established tradition in the search literature, we assume that the negotiated wage splits the surplus created by the job evenly (this will be made precise below). Total labor demand (F) and search effort together determine equilibrium steady-state unemployment (U).

The government's goal is to choose x and T to maximize aggregate expected lifetime income. Increases in x and/or T provide unemployed workers with additional insurance but these increases also lower optimal search effort. In addition, since a more generous UI program reduces the opportunity cost of unemployment, it increases the wage rate and makes creating a vacancy less profitable. The reduction in search effort coupled with the destruction of job opportunties leads to an increase in equilibrium unemployment. The optimal government policy must balance these costs and benefits.

In terms of the literature reviewed above, our approach is very similar to that of Mortensen (1994) and Millard and Mortensen (1994), except that we assume risk aversion on the part of workers. Alternatively, our work could be viewed as an extension of Baily (1978) and Flemming (1978) in which we (a) make the potential duration of benefits variable, (b) take into account the fact that the UI take-up rate is below 100%, and (c) model labor demand so that the job destruction effects of UI are taken into account.

We describe the model in three steps. First, we show how to determine expected lifetime utility for all agents in the economy and use these measures to define welfare. We also show how these measures may be used to determine optimal search effort for unemployed workers. Second, we show how total labor demand and search effort can be combined to determine equilibrium unemployment. Finally, we introduce our model of firm behavior and show how total labor demand and the equilibrium wage are determined.

Before we begin, a few words about our notation are in order. Throughout the analysis we define variables such as search effort, expected lifetime utility, reemployment probabilities, et cetera that depend upon the employment status of the worker. In each case, we use sub-scripts on the variables to denote the employment status with w representing employed workers, t denoting unemployed workers in their t^{th} period of search, and x denoting unemployed workers who have exhausted their benefits. Thus, for example, if we use m to denote the reemployment probability, m_t would represent the reemployment probability for an unemployed workers in the t^{th} period of search while m_x would represent the reemployment probability for an unemployed worker who has exhausted her benefits.

A. Expetected Lifetime Utility, Search Effort, and Welfare

We use V_j to denote expected lifetime utility for a worker in employment state j (j = w if employed, t if unemployed for t periods, and x if unemployed and benefits have been exhausted). In addition, we use u( ) to represent the agents' common utility function. We assume that per period utility takes the form u(C) - c(p) with C denoting consumption, c(p) denoting the cost of search, and p denoting search effort (if unemployed). We assume that c(p) is a convex function and that c(0) = 0. We begin by assuming that agents cannot save so that in any given period consumption equals income. In Section 4 we discuss how relaxing this assumption affects our results.

For employed workers, current income consists of two components -- labor income, which is equal to the wage net of taxes, $w(1 - \tau)$, and non-labor income, which is equal to their share of the aggregate profits earned by the firms, θ_w. Thus, current utility is given by $u[w(1 - \tau) + \theta]$. Obviously, employed agents incur no search costs. To determine expected lifetime utility, we must also consider the worker's future prospects. Let s denote the probability that in any given period the worker will lose her job. Then, with probability (1 - s) the worker's expected future lifetime utility will continue to be V_w (since she remains employed). With the remaining probability of s the worker loses her job and her expected future lifetime utility falls to V_1. It follows that,

(1) $V_w = u[w(1-\tau)+\theta_w] + [sV_1+(1-s)V_w]/(1+r)$.

Note that future utility is discounted at rate (1+r) with r denoting the interest rate.

Turn next to the unemployed. For them, current income is equal to the sum of unemployment insurance (if benefits have not yet been exhausted) and profits. We use θ_u to denote a typical unemployed worker's share of aggregate profits. Future income depends on future employment status. We use m to deonte reemployment probabilities so that with probability m_t the worker finds a job and can expect to earn V_w in the future, while with the remaining probability she remains unemployed and can expect to earn V_{t+1} in the future. Thus,

$$(2) \quad V_t = u[x+\theta_u] - c(p_t) + [m_t V_w + (1-m_t)V_{t+1}]/(1+r) \qquad \text{for } t = 1,\ldots,T.$$

$$(3) \quad V_x = u[\theta_u] - c(p_x) + [m_x V_w + (1-m_x)V_x]/(1+r).$$

We are now in a position to define welfare (W). Let U_t represent the number of workers who have been unemployed for t periods and define U_x analogously for UI-exhaustees. Then, if we define J to be the total number of jobs held in the steady-state equilibrium and aggregate expected lifetime utility across all agents, we obtain

$$(4) \quad W = JV_w + U_x V_x + \Sigma_t U_t V_t.$$

Finally, since unemployed workers choose search effort (p) to maximize expected lifetime income (V) we have,

$$(5)\quad p_t = \arg\max V_t \qquad \text{for } t = 1,\ldots,T.$$

$$(6)\quad p_x = \arg\max V_x.$$

In maximizing expected lifetime utility, it is important to note that the reemployment probability (m) is an increasing function of search effort (p). We make the link between the two explicit in sub-section B below.

This completes the description of expected lifetime utility and the determination of search effort. At this point it is useful to note that if we were to stop here, we would have a model very similar to the one used by Shavell and Weiss (1979). In essence, their approach is to describe expected lifetime utility, assume that m is increasing in p, fix the total amount the government is going to spend on UI, and then choose a path of benefits (x_t for t = 1,2,....) to maximize V_1, the expected lifetime utility of a newly unemployed UI-eligible worker. As discussed above, this does not take into account the costs of the program nor does it tell us the optimal amount that the government should be spending on UI. In addition, it is not at all clear why Shavell and Weiss focus on the benefit path that maximizes V_1, since it seems clear that W is a more appropriate measure of welfare.

B. Determining Unemployment

In this sub-section we show how total labor demand (F) and search effort (p) can be combined to determine equilibrium unemployment. To do so, we first show how to determine steady-state unemployment once the reemployment probabilities have been determined. Second, we show how the reemployment probabilities vary with search effort, labor demand, and other features of the labor market.

Formally, we use L to denote total labor supply. Then, since every worker is either employed or unemployed, we have

(7) $L = J + U.$

In addition, given our definitions of U_t and U_x we can write total unemployment as

(8) $U = \Sigma_t U_t + U_x.$

Turn next to the firms. For simplicity, we assume that each firm provides only one job opportunity.[1] Thus, F denotes both the total number of firms and the total number of jobs available at any time. Each job is either filled or vacant. If we let V denote the number of vacancies in a steady-state equilibrium, it follows that

[1] This assumption is commonly used in general equilibrium search models (see, for example, Diamond 1982 or Pissarides 1990). Alternatively, we could simply assume that each firm recruits for and fills each of its many vacancies separately.

(9) $F = J + V$.

We are now in a position to describe the dynamics of the labor market and the conditions that must hold if we are in a steady-state equilibrium. These conditions guarantee that the unemployment rate and the composition of unemployment both remain constant over time. We begin by reminding the reader that s is defined to be the economy's separation rate -- that is, s denotes the probability that an employment relationship will dissolve in any given period. In addition, remember that reemployment probabilities are denoted by the m terms. Then, for any given worker, there are T + 2 possible employment states -- U_1, U_2,...., U_T, U_x, and J. If employed (i.e., if in state J) the worker faces a probability s of losing her job and moving into state U_1. If unemployed for t periods (i.e., if in state U_t), the worker faces a probability of m_t of finding a job and moving into state J. With the remaining probability of $1 - m_t$ this worker remains unemployed and moves on to state U_{t+1}. Finally, UI-eligible exhaustees face a reemployment probability of m_x, in which case they move into state J. Otherwise, they remain in state U_x.

In a steady-state equilibrium the flows into and out of each state must be equal so that the unemployment rate and its composition do not change over time. Using the above notation, the flows into and out of state U_1 are equal if

(10) $sJ = U_1$.

The flows into and out of state U_t (for t = 2,...,T) are equal if

(11) $(1-m_{t-1})U_{t-1} = U_t$

Finally, the flows into and out of state U_x are equal if

(12) $(1-m_T)U_T = m_xU_x$.

In each case, the flow into the state is given on the left-hand-side of the expression while the flow out of the state is given on the right-hand-side.

Equations (7)-(12) define the dynamics of the labor market given the reemployment probabilities and total labor demand. We must now explain how search effort translates into a reemployment probability for each unemployed worker. As described above, each unemployed worker chooses search effort (p) to maximize expected lifetime utility. Search effort is best thought of as the number of firms a worker chooses to contact in each period of job search. For workers who contact fewer than one firm on average, p_t could also be thought of as the probability of contacting any firm. Once a worker contacts a firm, she files an application for employment if the firm has a vacancy. Since there are F firms and V of them have vacancies, the probability of contacting a firm with a vacancy is V/F. Finally, once all applications have been filed, each firm with a vacancy fills it by choosing randomly from its pool of applicants. Thus, if N other workers apply to the firm, the probability of a given worker getting the job is $1/(N+1)$. Since each other worker either does or does not apply, N is a random variable with a Poisson distribution with parameter λ where λ is equal to the average number of applications filed at each firm. It is straightforward to show that this implies that the probability of getting a job offer conditional on having applied at a firm with a vacancy is $(1/\lambda)[1 - e^{-\lambda}]$. The employment probability for any given worker is then the product of these three terms -- the number of firms contacted, the probability that a given firm will have a vacancy, and the probability of getting the job conditional on having applied at a firm with a vacancy:

$$(13) \qquad m_t = p_t(V/F)(1/\lambda)[1-e^{-\lambda}] \qquad \text{for } t = 1,\ldots,T$$

$$(14) \qquad m_x = p_x(V/F)(1/\lambda)[1-e^{-\lambda}]$$

where

$$(15) \qquad \lambda = \{\Sigma_t p_t U_t + p_x U_x\}/F.$$

These equations define the employment probabilities of workers as a function of search effort and the length of time that they have been unemployed. Note that for any given worker, the search effort of other workers affects that worker's employment probability through λ.

Given the levels of search effort and expected lifetime utilities defined by (1)-(6), equations (7)-(15) can be solved for equilibrium unemployment (U), its composition (U_t for t = 1,...,T and U_x), and the reemployment probabilities (m_t for t = 1,...,T and m_x). If we were to stop developing the model at this point, treating F and w as exogenous, we would have a model almost identical to the one used by Flemming (1978). In fact, there would be only two real substantive differences between the models -- Flemming allows workers to save while employed while we do not and Flemming assumes that UI is offered indefinitely while we assume that it is only offered for T periods. As we mentioned above, we add a third distinction in the next section when we add UI-ineligible workers to the model.

C. Firms

To make the number of firm endogenous we assume that firms enter the market until the expected profit from doing so equals zero. When a firm enters the market, it creates a vacancy and starts to accept applications from unemployed workers to fill it. Once the vacancy is filled, the firm produces and sells output as long as its vacancy remains filled. If the firm loses its worker, it must restart the process of filling its vacancy.

We use Π_V to denote the expected lifetime profit for a firm that currently has a vacancy and use Π_J to represent the expected lifetime profit for a firm that has filled its vacancy. Thus, when a firm enters the market and creates a vacancy it can expect to earn Π_V in the future. Once it fills its vacancy, its expectations about future profits rise to Π_J. Firms enter until

(16) $\Pi_V = 0$.

To calculate Π_V and Π_J we folow the same procedure that was used to determine expected lifetime utilities -- we consider the current and future prospects of typical firms. Let q denote the probability of filling a vacancy, use R to denote the revenue earned by a firm that is producing, and let K represent the cost of maintaining a vacancy. Then, current profit for a firm with a vacancy is -K while current profit for a firm that is producing is R - w - K. Now consider their future prospects. A firm that has an opening fills it with probability q, in which case its expected lifetime profits rise to Π_J. With the remaining probability the vacancy remains open and the firm continues to expect to earn Π_V. Thus,

(17) $\Pi_V = -K + [q\Pi_J + (1-q)\Pi_V]/(1+r)$.

A firm that has already hired a worker keeps that worker with probability (1-s) and continues to earn Π_J. With the remaining probability, it loses its worker and sees its expected profits fall to Π_V. Thus,

(18) $\Pi_J = R - w - K + [s\Pi_V + (1-s)\Pi_J]/(1+r)$.

Note that, as before, future profits are discounted at rate (1+r).

The probability of filling a vacancy, q, depends on the number of firms competing for the unemployed (V), the number of unemployed workers (U) and the search effort of workers. In any given period the number of unemployed workers who find new jobs is equal to $\Sigma_t m_t U_t + m_x U_x$ while the number of vacancies that are filled is equal to qV. Since these values must be equal, we have

$$(19) \quad q = [\Sigma_t m_t U_t + m_x U_x]/V.$$

Note that the search effort of workers enters (19) through the reemployment probabilities.

The next step in developing our model is to use Π_V and Π_J to determine the profits that are distributed to workers in each period in the form of dividends (θ_w for the employed and θ_u for the unemployed). Since there are J jobs filled in equilibrium with each one generating Π_J in expected <u>lifetime</u> profits, aggregate expected lifetime profits are $J\Pi_J$. Thus, the aggregate <u>per period</u> profits are equal to $rJ\Pi_J/(1+r)$. These profits must be distributed to workers each period. We assume that these profits are distributed evenly to <u>employed</u> workers with the unemployed receiving nothing. It follows that $\theta_w = rJ\Pi_J/(1+r)J = r\Pi_J/(1+r)$ and $\theta_u = 0$. We make this assumption for the following reason. Suppose that the government were to reduce the generousity of the UI program, resulting in an increase in aggragte profits. If the unemployed were to receive a share of these profits, this increase in non-labor income could swamp the decrease in UI leaving the unemployed better-off. Since it is unlikely that the unemployed receive significant income from such non-labor sources, we assume that all profits go to the employed.

The final step in developing our model is to explain how the wage is determined. Following the general equilibrium search literature (see, for example, Diamond 1982 or Pissarides 1990), we assume that the firms and workers split the surplus created by the representative job evenly. For firms, when they fill a vacancy their expected profits rise from Π_V to Π_J. For an average worker, when they become employed their expected lifetime utility rises from V_u to V_w where V_u deontes the average expected lifetime utility for unemployed workers. That is,

$$(20)\quad V_u = [\Sigma_t U_t V_t + U_x V_x]/U.$$

It follows that the total surplus created by the average job when measured in dollars is $\Pi_J - \Pi_V + (V_w - V_u)MU_I$ where MU_I represents the workers marginal utility of income and allows us to transform the workers gain, $V_w - V_u$, which is measured in utility, into an appropriate dollar value. This surplus is split evenly between the firm and its employee if the wage solves

$$(21)\quad \Pi_J - \Pi_V = (V_w - V_u)MU_I.$$

In summary, when we model firms the number of firms demanding labor (F) is determined by (16) while the equilibrium wage is determined by (21).

The government's problem is to choose x (the UI benefit level) and T (the potential duration of benefits) to maximize welfare (W, as given in eq. 4) subject to the constraint that its budget balances. Since there are J employed workers each earning a wage of w, total tax revenue is equal to $Jw\tau$. In equilibrium, $U - U_x$ unemployed workers each receive benefits of x each period. Thus, the total cost of the program is $(U - U_x)x$. For the budget to balance it must be the case that

(22) $(U - U_x)x = Jw\tau.$

As noted above an increase in x or T increases the level of insurance provided to unemployed workers, but both increase equilibrium unemployment and require that τ increase in order to fund the expanded program.

This completes the description of our model. In structure it is very similar to that of Mortensen (1994) and Millard and Mortensen (1994). The major difference is in the manner in which welfare is measured -- while they use aggregate income net of search, recruiting, and training costs as their measure of welfare we use aggregate expected lifetime utility. These two measures are identical <u>if</u> agents are risk neutral. However, if the utility function is concave, so that agents are risk averse, the measures differ. As we argued above, we feel that it is important to assume risk aversion since this implies that there are positive benefits from providing unemployment insurance.

D. Properties of Equilibrium

Before we turn to optimal policy, it is useful to first describe the structure of equilibrium and some of its comparative dynamic properties. It is straightforward to show that in a steady-state equilibrium that $V_w > V_1 > V_2 > \ldots > V_T > V_x$. That is, expected lifetime income is highest for employed workers, lowest for unemployed workers who have exhausted their benefits, and decreasing in the number of weeks that a worker has been unemployed. Intuitively, workers in the early stages of a spell of unemployment have more weeks to find a job before they have to worry about losing their UI benefits. Because of this, workers who have recently become unemployed will not search as hard as those who have been unemployed for a longer period of time -- that is, optimal search effort will be increasing in the number of weeks of unsuccessful search ($p_1 < p_2 < \ldots < p_T < p_x$).

A decrease in UI benefits (x) or the potential duration of benefits (T) decreases the level of insurance offered unemployed workers and triggers an increase in search effort by all UI-eligible workers (and therefore lowers equilibrium unemployment). Either change results in a decrease in V_t for all t, but decreases in x and T have opposite effects on the probability of exhausting benefits. A decrease in x makes it less likely that a worker will exhaust her UI benefits before finding a job (since she searches harder). But a decrease in T makes it more likely that benefits will be exhausted since the time horizon over which benefits are offered has been shortened (this is true in spite of the fact that search effort increases as a result of the decrease in T). Of course, increases in x or T lead to the opposite effects.

Changes in the UI program also have important implications for firm behavior and labor demand. Since increases in either x or T reduce the cost of being unemployed, they make workers less willing to search for and/or accept jobs. This results in an increase in V_u and forces firms to increase the wage that they offer their new employees. This increase in the wage makes production less profitable and results in fewer firms and fewer job opportunities. This job destruction effect increases unemployment and lowers net output.

E. Calibration

In order to determine the optimal UI program we must choose values for the parameters of the model, solve for the equilibrium generated by each pair of policy parameters (x and T), and compare the levels of welfare achieved in the different equilibria. Assuming that we choose realistic values for the parameters, this exercise should give us some idea as to the ranges in which the optimal level of benefits and the optimal potential duration of benefits lie.

The parameters of the model are the separation rate (s), the interest rate (r), the size of the labor force (L), the search cost function ($c(p)$), the revenue earned by producing firms (R), the cost of maintaining a vacancy (K), and the utility function, $u(C)$. Since we are interested in varying the degree of risk aversion, we calibrate the model separately for a variety of different utility functions and compare the optimal programs that result.

We calibrate the model in two steps. First, we treat the model introduced in sub-sections A and B as if it were self-contained -- that is, we treat the number of firms (F) and the wage (w) as if they were parameters of the model. To calibrate this portion of the model we rely on data collected to analyze the Illinois Reemployment Bonus Experiment. Since we have discussed this calibration exercise in detail elsewhere (see, for example, Davidson and Woodbury 1993, 1994), we provide only a brief description of how we obtain estimates of the parameters of this abbreviated model. Briefly, this portion of the model is calibrated so that its predictions concerning the impact of a reemployment bonus offered to unemployed workers matches what was observed in Illinois. By treating F and w as fixed, we are implicitly assuming that the Illinois experiment had no wage or job creation/job destruction effects. In fact, the data does indicate that there were no wage effects from the reemployment bonus (Woodbury and Speigelman 1987) and, given that the program was temporary and limited in scope, it seems reasonable to assume that there were no significant changes in the number of firms seeking workers as a result of the bonus. Thus, we consider this approach appropriate.

In the second step, we expand the model (by adding sub-section C) so that F and w become endogenous. This adds two new parameters to the model -- R (the revenue earned by the firm when producing) and K (the cost of maintaining a vacancy). These values are then chosen so that the full model yields (a) a value for w that matches the data collected in Illinois, and (b) values for F that lie in the range predicted by the abbreviated model in the first stage of calibration.

Now, we begin with step one of the calibration. When considering the abbreviated model (as introduced in sub-sections A and B), the parameters of interest are the separation rate (s), the interest rate (r), the wage (w), the number of firms (F), the size of the labor force (L), and the search cost function (c(p)). We can obtain an estimate for s from the existing literature on labor market dynamics. Ehrenberg (1980) and Murphy and Topel (1987) both provide estimates of the number of jobs that break-up in each period. If we measure time in 2-week intervals, their work suggests that s lies in the range of .007 to .013. For the interest rate we set r = .008 which translates into an annual discount rate of approximately 20%. Since our previous work (Davidson and Woodbury 1993) suggests that results from this model are not sensitive to changes in r over a fairly wide range, this is the only value for the interest rate that we consider.

For F and L we begin by noting that our model is homogeneous of degree zero in F and L so that we may set L = 100 without loss of generality. If we then vary F holding all other parameters fixed we can solve for the equilibrium unemployment and vacancy rates. Abraham's (1983) work suggests that the ratio of unemployment to vacancies (U/V) varies between 1.5 and 3 over the business cycle. Although the actual values of U and V depend on the other parameters, we find that to obtain such values for U/V in our model with L = 100, F must lie in range of 95 to 97.5. Thus, in the second stage of the calibration, we must choose values for R and K such that F lies in the range 95-97.5.

The remaining parameters in sub-sections A and B are the wage rate and the search cost function. For these values we turn to the data and results from the Illinois Reemployment Bonus Experiment. In the Illinois Reemployment Bonus Experiment a rondomly selected group of new claimants for UI were offered a $500 bonus for accpeting a new job within 11 weeks of filing their initial claim. The average duration of unemployment for these bonus-offered workers was approximately .7 weeks less than the average unemployment duration of the randomly selected control group (Davidson and Woodbury 1991). In our previous work, we estimated the parameters of the search cost function that would be consistent with such behavioral results. That is, we assumed a specific functional form for c(p) and then solved for the parameters that would make the model's predictions match the outcome observed in the Illinois experiment. The functional form that we used was $c(p) = cp^z$, where z denotes the elasticity of search costs with respect to search effort. The values for c and z that make the model's predictions exactly match what occured in Illinois depends upon the utility function that is assumed. For example, if we assume that the utility function is linear in consumption, then our results indicated that for the average bi-weekly wage rate observed in Illinois ($511), the values of c and z that are consistent with the Illinois experimental results are $c = 338$ and $z = 1.23$. On the other hand, if the utility function takes on the form $u(C) = \ln(C)$, we find that the values of c and z that are consistent with the Illinois experimental results are $c = 2.05$ and $z = 1.38$.

Finally, turn to the second stage of calibration. In order to make F and w endogenous, we add the equations in sub-section C to the model. This adds only two new parameters, R and K. From the Illinois data we know that the average bi-weekly wage should be $511, and, from stage one of the calibration we know that F must lie in the range 95 to 97.5. Thus, we set x and T equal to their Illinois values -- x, the average bi-weekly UI benefit in Illinois is set equal to $242, and T, the potential duration of UI, in Illinois is set equal to 14 (since each period equals 2 weeks) -- and then we solve the model to determine what values of R and K would lead the model to predict that w = $511 and that F would fall in the range 95-97.5. Of course, the values of R and K depend upon the assumed functional form for the utility function. If the utility function is linear in consumption, then when $R = 724$ and $K = 2417$ the model predicts that $w = 511$ and $F = 96.25$. On the other hand, if $u(C) = \ln(c)$, then when $R = 1469$ and $K = 10863$ the model predicts that w = $511 and $F = 96.25$.

Once the calibration is complete, we set the parameters at the calibrated levels and solve for the welfare maximizing values of x and T. Once we have solved for the optimal values for x and T in one case, we vary the parameters over the ranges described above to test for the sensitivity of our results with respect to each parameter.

4. Results

In this section we begin by reviewing results from our earlier work, Davidson and Woodbury (1995), in which we solved for the optimal UI program in the abbreviated model outlined in sections 3.A and 3.B. These results are best thought of extensions of Baily (1978) and Flemming's (1978) work to an environment in which (a) the potential duration of benefits can vary and be controlled by the government, and (b) not all unemployed workers are eligible for UI. Next, we present new results concerning optimal UI when firm behavior is explicitly added to the model as in section 3.C. This allows us to examine how our initial results must be modified when the job destruction effects of more generous UI programs are taken into account. Finally, we extend our model once more in order to allow for worker heterogeniety and show how including workers with different labor market experiences in the model alters our results.

A. Optimal Potential Duration of Benefits without Job Destruction

The most surprising result from our earlier analysis is that in the abbreviated model the optimal potential duration of benefits is infinite -- that is, the government should offer UI benefits indefinitely to all unemployed UI eligible workers. Although there are some details omitted from the following reasoning[2], the crux of the argument is as follows. Agents facing employment risk would prefer a program that allows them to smooth consumption as much as possible across spells of unemployment. Thus, if given the choice between two UI programs that provide the same level of total benefits to the unemployed, agents would choose the program that does the best job of consumption smoothing. With this in mind, consider the following two UI programs -- the first program offers a benefit level of x for T periods while the second program offers a benefit level of x' for $T+1$ periods where $x' < x$ and is chosen so that the two programs provide the same level of total benefits to the unemployed. Thus, the first program offers higher benefits but for a shorter period of time. The key to the argument is to note that the second program allows for greater consumption smoothing -- in moving from the first program to the second program benefits are lowered during the least adverse states of unemployment (i.e., the initial phase) and increased in one of the most adverse states (period $T+1$ in which no benefits are offered in the first program) with total benefits provided remaining the same. In other words, by accepting slightly decreased benefits (and consumption) during the first T periods of unemployment, the unemployed can insure that benefits will not completely disappear for an additional period. Thus, all unemployed workers prefer the second program. Since this reasoning holds for all finite T, it follows that in an optimal UI

[2] See Davidson and Woodbury (1995b) for details.

program T must equal infinity.

This result has important implications for some of the work reviewed in section 2. Most importantly, this result implies that the conclusions reached by Baily and Flemming are misleading. Since both authors use models in which it is <u>assumed</u> that benefits are offered indefinitely and since, in their models it is indeed optimal to provide benefits indefinitely, the optimal replacement rates that they derive are correct -- without savings, the optimal replacement rate is in the 60-70% range, and, with savings but imperfect capital markets, the optimal replacement rate is in the 40-50% range[3]. However, these rates are optimal <u>only if they are offered indefinitely</u>. Thus, the conclusion that Baily and Flemming reach, that the U.S.'s 50% replacement rate is probably too high, is misguided, since the U.S. offers this rate for only 26 weeks. In fact, if we solve for the optimal replacement rate with T set exogenously at 26 weeks, we find that the optimal replacement rate is 1! It follows that if one ignores the job destruction effect of UI, the current U.S. unemployment insurance program is not generous enough.

It is important, however, not to place too much emphasis on this result. That is, we must remember the setting in which it was derived -- it was derived in a model in which the job destruction effects of UI were ignored. In fact, as we show below, when the job destruction effects are taken into account, this result no

[3] It is important to note that our abbreviated model yields almost identical predictions concerning optimal replacement rates.

longer holds. For this reason, we do not believe that an optimal UI program would indeed be characterized by an unlimited potential duration of benefits. However, what this result does indicate is that an optimal UI program is more likely to be characterized by low benefits and a long potential duration of benefits than a program with high benefits and a short potential duration of benefits (as in the U.S.). The intuition behind this result is clear -- programs with long potential durations of benefits lead to smoother consumption paths and therefore reduce the risk associated with unemployment more than programs with shorter potential durations.

B. Optimal Replacement Rates with UI-Ineligibles in the Model

Our second extension of the Baily and Flemming analyses was to explicitly take into account the fact that not all unemployed workers are eligible to collect UI. For example, for the U.S. Blank and Card (1991) report that over 50% of the unemployed are ineligible for UI and that of those who are eligible, only 75% bother to file for their benefits. Layard et al (1991) report that in the U.K. up to 30% of the unemployed are not eligible to collect UI benefits. This fact has important implications for the optimal replacement rate since more generous UI has positive spill-over effects on UI-ineligibles. The reasoning is as follows. If the government institutes a more generous UI program, UI-eligibles respond by searching less hard for employment. Assuming that UI-eligible and UI-ineligibles compete for some of the same jobs, this reduces the competition that UI-ineligibles face for those jobs and increases their reemployment probabilities. The existence of these positive spill-over effects implies that models that ignore the fact that not all unemployed workers are eligible to collect UI will underestimate the optimal replacement rate. Thus, Baily and Flemming's estimates of the optimal replacement rate are biased downwards.

To determine the optimal replacement rate when these positive spill-over effects are present, we extended the abbreviated model of sections 3.A and 3.B to allow for UI-ineligibility. Briefly, UI-ineligibles were modelled in exactly the same manner as other workers except that they were not allowed to collect UI while unemployed. For example, an equation almost exactly identical to (2) and (3) was used to define the expected lifetime utility for an unemployed UI-ineligible worker, and an equation almost identical to (1) was used to define the expected lifetime utility for an employed UI-ineligible worker. To be precise, let V_i represent the expected lifetime utility for an unemployed UI-ineligible worker, let V_{wi} denote the expected lifetime utility for an employed UI-ineligible worker, and use the sub-script i on all other variables to denote UI-ineligibility. Then, we can apply the same logic used to derive (1)-(3) to obtain:

(23) $V_{wi} = u[w(1-\tau)+\theta_w] + [sV_i+(1-s)V_{wi}]/(1+r)$

(24) $V_i = u[\theta_u] - c(p_i) + [m_iV_{wi}+(1-m_i)V_i]/(1+r)$.

Optimal search effort for UI-ineligibles is then the value of p_i that maximizes V_i:

$$(25)\ p_i = \arg\max V_i.$$

The remaining equations of the model can be modified in a similar fashion (interested readers are referred to Davidson and Woodbury 1995b for details) with only one new parameter added - the proportion of the unemployed who are ineligible for UI. Following Blank and Card (1991) we set this value equal to .6 for our basecase, and then vary it throughout the analysis from 0 to .6 to see how sensitive our results are to the value of the parameter.

We find, as expected, that including UI-ineligibles in the model does increase the optimal replacement rate. Depending upon the values of the other parameters (the interest rate, the separation rate, et cetera), we find that the positive spill-over effects of UI on UI-ineligibles increases the optimal replacement rate by 6 to 10 percentage points. Thus, if agents cannot save and the job destruction effects of UI are ignored, an optimal UI program offers a replacement rate in the 65-75% range indefinitely. If, on the other hand, agents can save but the job destruction effects of UI are ignored, then an optimal UI program entails the government offering replacement rates in the 45-55% range indefinitely.

This completes the description of our earlier results. Before moving on and discussing our new results, it is important to note that all of our previous results were derived assuming that utility is linear in consumption. If we had also assumed that search costs were linear in effort, this would have been equivalent to assuming risk neutrality and there would have been no demand for employment insurance. However, since we assumed that search costs were convex in effort, each individual's optimization problem is concave in the choice variable and thus, each agent is risk averse.

To see how increasing the degree of risk aversion affects these results, we have recently recalibrated the model for two different utility functions, namely $u[C] = \ln(C)$ and $u[C] = \sqrt{C}$, and rederived the optimal replacement rate in each case. The log utility function is characterized by constant Arrow-Pratt relative risk aversion equal to one and was chosen since it is identical to the one used by Baily (1978). The square root utility function is characterized by constant Arrow-Pratt relative risk aversion equal to 1/2 and was used since its measure of risk aversion falls mid-way between our other two extremes (the linear and log utility functions). Surprisingly, in this model without job destruction, we find that the degree of risk aversion does not make much difference -- optimal replacement rates rise by only about 5% when we go from the linear to the log utility function and only about 2% when we go from the linear to the square root utility function. The reason for this is that in recalibrating the model with the new utility functions, the values of the parameters change so that the model once agian yields predictions that are consistent with the Illnois data. For example, as we make the agents in the model more risk averse, the degree of convexity of the search cost function must also increase so that the model still yields the same predictions concerning a reemployment bonus. Since we recalibrate the model for each utility function so that the reemployment bonus impact is identical across the models, it is not surprising that the models yield similar predictions concerning UI.[4]

In summary, our earlier work focused on two shortcomings of the Baily and Flemming approaches -- the fact that they simply assumed that UI benefits would be offered indefinitely and the fact that they assumed that all agents are eligible for UI. We demonstarted that both of these assumptions bias their results in favor of less generous UI programs and led them to draw misleading conclusions. However, as we have emphasized above, these are not the only two shortcomings of the Baily and Flemming analyses -- they also ignored the impact of UI on firm behavior. In the next sub-section we discuss how extending the model to allow for the job destruction effects of UI forces us to further modify our conclusions concerning an optimal UI program.

C. Job Destruction and Risk Aversion

When firm behavior is endogenized, there are several additional effects of UI. First, if a more generous UI program is

[4] When we calibrate the model for the square root utility function we obtain the following values for the key parameters -- c = , z = , R = , K = .

offered, the average expected lifetime income for the unemployed (V_u) rises and this triggers an increase in the equilibrium wage. This higher wage lowers profit for producing firms (Π_J) and lowers the expected lifetime profit for a firm creating a vacancy (Π_V). This results in fewer firms (F) and fewer job opportunities. In terms of welfare, per period income for the employed could rise or fall (since the wage is increasing while non-labor income from firms is falling) while unemployment unambiguously rises due to the job destruction effect. Thus, in a model with endogenous labor demand the optimal UI program is likely to be <u>less</u> generous than the optimal UI program in a model in which firm behavior is ignored, and the size of the job destruction effect determines just how much less generous it will be.

Our results indicate that, regardless of the degree of risk aversion, the job destruction effect is large enough to overturn the result that it is optimal to offer UI benefits indefinitely. To see why, return to our earler argument concerning the potential duration of benefits. We argued that for any UI program in which T were finite there would exist another UI program with longer potential duration of benefits and lower benefits that would cost the same to finance and would be strictly prefered by all unemployed agents. Thus, it would always be possible to increase T and raise welfare. This argument no longer holds when labor demand is endogenous since increasing T in this manner reduces the number of job opportunities and increases unemployment. This negative effect of the decrease in job opportunities must be weighed against the positive impact of smoothing consumption to determine if the increase in T raises welfare. We find that for all levels of risk aversion, the job destruction effect of increasing T eventually outweighs the consumption smoothing effect of increasing T so that benefits should eventually be cut-off.

The point at which the government should stop providing benefits depends heavily on the degree of risk aversion. We consider three cases. In the first, we assume that utility is linear in consumption so that the degree of risk aversion is extremely low (the reader is reminded that in this case risk aversion enters through the convexity of the search cost function). This makes our model and approach very similar to that of Mortensen (1994) and Millard and Mortensen (1994). In fact, in this case, our model yields predictions that are almost identical to those in Mortensen (1994) -- we find that the current UI program in the U.S. generates a dead weight loss of roughly 1.2% of welfare.

The fact that we obtain results that are so similar to Mortensen (1994) in spite of the fact that our models are calibrated in very different manners using different data is comforting. In addition, the reader is reminded that the abbreviated model of sub-sections 3.A-3.B yielded results remarkably similar to those found in Baily (1978) and Flemming (1978). Thus, our model seems to be able to reproduce the existing results in the literature once the assumptions are altered to match the models used by previous authors. This is true in spite of the fact that virtually all of the previous models were calibrated in different ways using data from a wide variety of different sources.

Unlike Mortensen (1994), we go on to use our model to derive the optimal UI program when benefits are constant over the spell of unemployment. With this low level of risk aversion, we find that the optimal UI program entails no benefits at all! That is, when the degree of risk aversion is low, the job destruction effect of UI is large enough to out weigh the positive impact of even one unit of insurance. Clearly, this result depends upon the fact that when utility is linear in consumption the demand for employment insuarnce is relatively low.

In the second case that we consider we assume that U(C) = ln(C) so that the Arrow-Pratt measure of relative risk aversion is constant and equal to one. This is the utility function used by Baily (1978) and is probably the utility function that is most often used in the literature on decision making under uncertainty. With these preferences, we obtain very different results. First, in stark contrast to the results obtained with linear utility, we find that the current U.S. unemployment insurance system <u>increases</u> welfare above the level that would be achieved without publically provided UI. Moreover, the welfare gains are far from trivial -- our estimate is that welfare rises by 1.2%.

Our second set of results concern the optimal UI program. As before, the job destruction effect overturns the result that benefits should be offered indefinitely. However, in this case, the optimal value of T remains quite large -- 90 weeks -- so that benefits should be offered for almost two full years. Thus, the job destruction effect is not nearly as important when agents are reasonably risk averse. As for the optimal replacement rate, when agents cannot self-insure, the optimal replacement rate is 65%. With savings, this rate is likely to fall by roughly 20%. We conclude that with reasonable assumptions concerning risk aversion, the optimal UI program offers benefits slightly below 50% for almost two years. Our model predicts that instituting such a UI program would raise welfare above the level achieved with the current U.S. program by 5.5% of welfare -- a startlingly high measure for a potential welfare gain.

In the final case that we consider we assume that utility is equal to the square root of consumption. This utility function has a constant Arrow-Pratt measure of relative risk aversion equal to 1/2, so that it falls mid-way between our other two utility functions. With this utility function we find that the current UI program in the US is just about right -- the optimal program involves offering a replacement rate of 61% for 26 weeks. We also find that this optimal program increases welfare above the levels that would be achieved without a UI program by about 2%.

The differences in our three sets of results indicate that the assumptions made concerning risk aversion are crucial. Thus, it is important to determine which utility function represents the most reasonable assumption concerning risk aversion. To answer this question, there are two contradictory strands of literature that we may consult. First, there is the empirical literature on consumption behavior that attempts to directly estimate agents' degree of risk aversion (see, for example, Zeldes 1989). The work in this area seems to indicate that the best point estimate of the Arrow-Pratt measure of relative risk aversion is 2.

The other literature, which is theoretical, attepmts to infer the degree of risk aversion from observed behavior. For example, we can observe how agents adjust their investment portfolios as their wealth changes and we can build models of investment under uncertainty to explain such behavior. Most work in this area finds that the theories of choice under uncertainty are consistent with observed behavior only if the Arrow-Pratt measure of relative risk aversion is less than one.

The fact that these two literatures contradict one another is troubling and leaves us in an uncomfortable position. Our work indicates that if the Arrow-Pratt measure of relative risk aversion is close to (or above) one, then the current UI program in the US in not nearly generous enough. However, if the Arrow-Pratt measure of relative risk aversion is close to 1/2, then the current system is about right. If one chooses to believe the empirical literature on consumption (as we tend to do), then the former outcome is much more likely than the latter. Thus, we conclude that in the most general model with the most reasonable assumption concerning risk aversion, we find that the optimal UI program offers benefits that are close to the levels currently offered by most States in the U.S. but it offers those benefits for a considerably longer period of time -- almost two years. In other words, the current U.S. program does not offer sufficient employment insurance.

C. Heterogeniety

All of the previous work on UI, including our own, relies on the assumption that all agents are alike. In reality, however, workers are subject to a wide variety of labor market experiences. Some workers are never unemployed, others find jobs quickly, and some always face long spells of unemployment upon losing a job. In addition, some agents may attempt to take advantage of the UI system while others would never even consider filing for benefits much less exploit the system. This implies that agents will have different preferences concerning employment insurance based on their labor market histories and expectations. Moreover, the number of workers that attempt to exploit the system may depend upon the generousity of the program.

In order to take worker heterogeniety into account, we extend our model to allow for three different classes of workers. The first class represents the bulk of the labor force and is described by the model introduced above. These workers face employment risk, losing their jobs with probability s in each period, and actively search for a new job once unemployed.

The second class consists of workers who are never unemployed. We refer to this group as "professionals" and use ϕ to denote the proportion of the labor force that falls into this class. We also use L_p to denote the number of such workers and V_p to denote their expected lifetime utility. Since these workers are never unemployed, they earn w in each period of life, and thus, $V_p = u(w)(1+r)/r$. The total contribution of these workers to social welfare is therefore L_pV_p and adding professionals to the model is accomplished by adding this term to W as defined in equation (4).

The last class of workers consists of agents who try to take advantage of the system. We refer to such workers as "slouchers." We assume that these agents work only to become eligible for UI and that they live off of the dole as much as possible. We use L_s to denote the number of slouchers and use V_s to represent their expected lifetime utility. Thus, their contribution to social welfare (W) is equal to L_sV_s.

Presumably, the number of slouchers in the labor force will be a function of the generousity of the system -- a more generous UI program should result in more slouchers. To measure the generousity of the system, we introduce the following variable G:

$$G = \{u(x)/u(w)\}\{1 - (1/1+r)^{T+1}\}.$$

G measures the ratio of utility recieved by simply collecting benefits as opposed to working for wage w during one spell of unemployment that lasts T periods (the potential duration of benefits). Note that if x = 0 or T = 0, so that no UI is offered, G = 0. On the other hand, as the replacement rate approaches one and T approaches infinity, G approaches 1. Increases in G represent increases in the generousity of the UI program. We assume that α, the proportion of the labor force that are slouchers, is positively related to G. In particular, we assume that $\alpha = \eta G$.

To complete the extended model, we must describe the determination of η and V_s. Consider V_s first. We assume that, since these agents work as little as possible, they contribute less to social welfare than the average unemployed agent (who, after all, is at least seeking a job). Thus, since V_u is the average expected lifetime utility for unemployed workers, we set $V_s = \Omega V_u$ with $\Omega < 1$. We then vary Ω and see how this affects the optimal UI program.

For η, we solve the model under the assumption that the current US program is in effect (a 50% replacement rate offered for 26 weeks) and then vary η so that α ranges from 0 to .05. Thus, we consider values for η that imply that currently anywhere from 0% to 5% of the labor force is exploiting the system.

Our results for the square-root utility function are summarized in Tables 1 and 2 where we report the optimal UI program for various values of α, Ω, and ϕ. In each cell, the optimal Ui program is reported by first listing the optimal replacement rate and then listing the optimal potential duration of benefits. Table 1 shows how the optimal UI program varies with α and Ω when there are no professionals in the model (i.e., $\phi = 0$). If $\alpha = 0$, so that there are no slouchers in the model, the optimal program offers a 61% replacement rate for 26 weeks. As the number of slouchers increases, the generousity of the optimal program declines regardless of the value of Ω. This is hardly surprising -- with more slouchers in the economy the government needs to make the program less generous in order to discourage the expoitation of the system.

Table 1 also indicates that the generousity of the optimal program is decreasing in Ω, the parameter that measires the amount that slouchers contribute to social welfare. As Ω decreases, slouchers contribute less to social welfare and it becomes more important for the government to discourage slouching. Table 1 clearly indicates the importance of the actual values of α and Ω. If α is low or if Ω is close to one, then the optimal program is quite close to the optimal program in the model that ignores slouching. On the other hand, for large values of α and low values of Ω (e.g., α = .05 and Ω = .7), the optimal program is considerasbly less generous.

Table 2 report the optimal UI program when both slouchers and professionals are included in the model. These results are derived assuming that utility is eqaul to the square-root of consumption and that Ω = .8 (as in the middle row of Table 1). table 2 indicates that as the number of professionals increases the optimal program becomes more generous. The reasoning is as follows. Adding professionals to the model spreads out the tax burden that the UI system places on the employed and allows the government to afford a more generous system. As in Table 1, knowing the true value of ϕ is important -- for low values of α, the optimal UI program varies quite a bit with ϕ. For example, when α = 0 the optimal UI program when 10% of the work force is made up of professionals offers a replacement rate of 64% for 28 weeks. If, on the other hand, 30% of the work force are professionals, the optimal program offers a replacement rate slightly higher (68%) but for a much longer time (36 weeks).

5. Discussion

In this paper we have presented a general equilibrium search model of the labor market in order to determine the optimal UI program when (a) the government can control the optimal potential duration of benefits and (b) the replacement rate must remain constant over the spell of unemployment until benefits are exhausted. We believe that our approach is superior to those that have been used in the past for a number of reasons. First, with respect to the labor economics literature, we have used an equilibrium model that allows us to measure the costs of different UI programs through their impact on search effort, job creation and unemployment. With respect to the macroeconomics literature, we have assumed that workers are risk averse so that we can measure the welfare benefits of different UI programs through the insurance that they provide against employment risk. Finally, with respect to the literature in public economics, we have adopted their approach, but offered a richer model in that (a) we have allowed the potential duration of benefits to vary, (b) we have included UI-ineligibles in the model, and (c) we have modeled firm behavior so that we could measure the job destruction effects of UI.

Our basic finding is that current benefit levels offered by most States in the U.S. are about right, but that these benefits are not offered for a long enough period of time. Thus, we conclude that the current U.S. system in not generous enough.

Our finding that the optimal UI program is characterized by fairly a low replacement rate and a very long potential duration of benefits stands in stark contrast to most of the previous literature. However, we argue below that our results should have been expected, since they are consistent with the vast abstract literature on optimal insurance contracts in the presence of moral hazard. In the next sub-section we offer a brief review of this literature for two purposes. First, reviewing this literature allows us to view the UI issue from a different perspective -- one that makes the economic sense behind our results seem almost transparent. Second, the results in this literature suggest that there may be another slightly more complex UI program that is radically different from the current program and possibly superior to the one that we have proposed.

. The Optimal Insurance Literature

There are three issues that have been addressed in the abstract literature on optimal insurance contracts that have important implications for the design of an optimal UI program. The first issue concerns the design of an optimal insurance contract when the insured agent's behavior can effect the probability of a loss occuring (i.e., moral hazard is present). To investigate this issue, it is assumed that the agent's behavior cannot be observed by the insurance provider so that the contract must be structured in a manner that makes putting forth effort optimal for the agent. The key issue then is how to provide adequate insurance without reducing the agent's incentive to aviod the loss. Shavell (1979) is perhaps the best known work in this area.

The second issue concerns the optimal way to share risk between a risk neutral insurance provider and a risk averse agent when the total level of insurance coverage is fixed. Although the article actually addresses a host of other issues as well, Raviv (1979) provides the classic treatment of this issue.

The final issue concerns the design of insuarnce contracts in the presence of adverse selection -- a situation in which agents differ in a dimension that affects their need for insurance but is unobservable to the insurance providers. The main issue in this case is to devise insurance contracts that will lead agents to self-select into groups and therefore reveal their personal characteristics. The classic article in this area is Rothschild and Stiglitz (1976).

The remarkable thing about these three strands of literatyre is that in spite of the fact that they ask different questions, they all come up with the same answer -- in all three cases, the optimal insurance contract takes the form of a "deductible policy" in which coverage is not provided for losses below a certain level. The reasoning is as follows. When agents face uncertainty in income they would like to smooth income as much as possible by purchasing insurance. In fact, in the absence of moral hazard concerns, the optimal insurance contract in a competitive insurance market provides full coverage so that income is the same in all circumstances. However, when moral hazard is present, the market breaks down when full insurance is provided since, in that case, no agent would have any incentive to take care in order to avoid large losses. With no one taking care, large losses would occur and insurance providers would go broke compensating the insured. Thus, given that full insurance will not be provided, what type of insurance is best? To answer this, note that agents are most concerned about avoiding catastrophes -- that is, extremely large losses. It follows that the outcomes that they are most concerned about being insured against are the most adverse outcomes, and any optimal insurance contract will have to provide coverage in such cases. The insurance contract must also provide incentives to take care to avoid losses, and this is provided by not covering small losses -- there is a deductible that the insured agent must cover any time that a loss occurs. In summary, a deductible contract forces agents to cover all small losses and provides coverage against large losses. It is optimal since it provides coverage in the cases that agents are most concerned about and includes incentives for agents to put forth effort to avoid losses.

What are the implications for unemployment insurance? For unemployed workers, large losses occur when they suffer long spells of unemployment. Thus, an optimal UI program should provide compensation to those who have a particularly diificult time finding reemployment. This is why we find that a long potential duration of benefits is optimal. As for the deductible, we have ruled them out by requiring the replacement rate to remain constant over the spell of unemployment until benefits are exhausted. Therefore, the only way to force agents to search for employment is to keep the replacement rate relatively low. This explains why we find optimal replacement rates at or below the current rates offered in the U.S..

The results from the optimal insurance literature also imply that the current UI program in the U.S. is exactly the opposite of what it should be. By offering benefits for the first 26 weeks of unemployment, the government is covering <u>all</u> short spells of unemployment and thus, all small losses. In addition, by cutting off benefits after 26 weeks the government is not providing coverage in the most important cases -- ones in which agents suffer large losses due to long spells of unemployment.

Is there any way to design a UI program with a deductible? One simple way to do so would be to use a three stage program in which the government offers very low (or zero) benefits during the first stage of unemployment, followed by a higher replacement in the second stage that lasts for a considerable length of time, followed by no benefits in the final stage. For example, the replacement rate could be 25% for the first 26 weeks of unemployment, followed by 60% for weeks 27 through 90, followed by zero thereafter. Such a program would provide a strong incentive for unemployed workers to find rapid reemployment since they would be receiving very little from the government early on. However, in the unfortunate cases in which workers are unable to find new jobs quickly, the government would step in and provide help when it is most needed.

This type of program would also carry with it at least two additional benefits. First, it would end the government subsidization of temporary layoffs by firms. For quite some time economists have argued that since UI is not completely experienced rated, firms have an incentive to exploit the system by temporarily laying off workers and then recalling them as their benefits expire. Some authors have estimated that as many as 25-50% of all layoffs in the U.S. can be explained in this manner (see, for example, Anderson and Meyer 1995 or Topel). However, if laid-off workers receive little or no benefits during the initial satges of unemployment, they would have an incentive to move on and seek new jobs rather than wait for recall. And, if workers are unwilling to wait for the firm to recall them, then the firms will be less likely to lay them off initially.

The second benefit of such a program is that it would discourage those who attempt to exploit the system (slouchers). With a substantial waiting period before UI begins or low replacement rates during the initial stages of unemployment, agents who would like to live off of the dole would have to pay a substantial penalty in order collect the higher replacement rates that would be offered to the long term unemployed. Therefore, such a program should substantially reduce the number of slouchers in the system.

At this point it is useful to emphasize that previous results have hinted that such a "deductible" program might be more efficient than current UI programs. Figure 1, which shows the conjectured optimal benefit path when agents can save and affect their probability of reemployment has this flavor -- benefits are initially low to encourage search and then rise as savings are depleted in order to allow workers to smooth consumption. O'Leary's (1994) empirical results that short spells of unemployment are currently over compensated while long spells are under compensated is also consistent with this type of policy shift. Finally, Wang and Williamson (1995) have argued for a benefit path similar to the one depicted in Figure 1 along with reemployment bonuses as part of an optimal UI program.

The results of Wang and Williamson (1995) are especially worthy of review, given their similarity to ours. In their paper, they solve for the optimal benefit path <u>and</u> consumption stream when agents face randomness in employment. Thus, they allow the government to subsidize or tax movements into various labor market states (by choosing consumption) in addition to setting the benefit path. As noted above, they do not model firms and therefore do not capture the job destruction effects of UI. Nevertheless, their results have the same flavor as ours. Our Figures 2 and 3 are (slightly modified) reproductions of Figures 17 and 21 from their paper. Figure 2 shows the optimal benefit path as a function of the length of the spell of unemployment with one unit of time equally one quarter of a year. Note the non-monotionicity of the benefit path -- benefits are lower in the first quarter than the second quarter. In addition, note the generousity of the system -- the replacement rate remains above 50% for over 5 quarters!

Figure 3 shows consumption across the spell of <u>employment</u>. It is important to note that consumption in the first period after reemployment is much higher than it is in all subsequent periods -- there is a reemployment bonus. This bonus provides workers with an extra incentive to seek reemployment in the early stages of unemployment by rewarding those who find new jobs. Without such a bonus, the deductible that workers would have to pay in the first period of unemployment (as represented by the low replacement rate in the first quarter of unemployment) would be higher (so that the replacement rate in the first quarter would be lower).

References

Abraham, Katherine G. "Structural/Frictional vs. Deficient Demand Unemployment: Some New Evidence." American Economic Review 73 (September 1983): 708-724.

Baily, Martin N. "Some Aspects of Optimal Unemployment Insurance." Journal of Public Economics 10 (1978): 379-402.

Blank, Rebecca M. and David E. Card. "Recent Trends in Insured and Uninsured Unemployment: Is there an Explanation?" Quarterly Journal of Economics 106 (November 1991): 1157-1189.

Davidson, Carl and Stephen A. Woodbury. "Effects of a Reemployment Bonus under Differing Benefit Entitlements, or, Why the Illinois Experiment Worked." Manuscript, Department of Economics, Michigan State University, and W.E. Upjohn Institute for Employment Research, March 1991.

Davidson, Carl and Stephen A. Woodbury. "The Displacement Effect of Reemployment Bonus Programs." Journal of Labor Economics 11 (October 1993): 575-605.

Davidson, Carl and Stephen A. Woodbury. "Unemployment Insurance and Unemployment: Implications of the Reemployment Bonus Experiments." Manuscript, Department of Economics, Michigan State University, and W.E. Upjohn Institute for Employment Research, February 1995.

Davidson, Carl and Stephen A. Woodbury. "Optimal Unemployment Insurance." W.E. Upjohn Institute for Employment Research Working Paper 95-35, July 1995.

Diamond, Peter. "Wage Determination and Efficiency in Search Equilibrium." Review of Economic Studies 49 (April 1982): 217-228.

Ehrenberg, Ronald G. "The Demographic Structure of Unemployment Rates and Labor Market Transition Probabilities." Research in Labor Economics 3 (1980): 214-291.

Flemming, J. S. "Aspects of Optimal Unemployment Insurance." Journal of Public Economics 10 (1978): 403-425.

Gruber, Jonathan S. "The Consumption smoothing Benefits of Unemployment Insurance." Manuscript, Department of Economics, MIT, May 1994.

Hamermesh, Daniel S. Jobless Pay and the Economy. Baltimore, MD: Johns Hopkins University Press, 1977.

Hamermesh, Daniel and Daniel Slesnick.

Hausman, Jerry. "Tax Policy and Unemployment Effects on Labor Policy." in Removing Obstacles to Economic Growth (edited by Michael Wachter and Susan Wachter). Philadelphia: University of Pennsylvania Press, 1984, pp. 70-96.

Kreps, David. A Course in Microeconomic Theory. Princeton University Press, 1990.

Layard, Richard, Stephen Nickell and Richard Jackman. Unemployment: Macroeconomic Performance and the Labor Market, Oxford University Press, 1991.

Millard, Stephen and Dale Mortensen. "The Unemployment and Welfare Effects of Labor Market Policy: A Comparison of the U.S. and U.K." Northwestern University Working Paper, August 1994.

Mortensen, Dale. "Reducing the Supply-Side Disincentives to Job Creation." Northwestern University Working Paper, September 1994.

Murphy, Kevin M. and Robert H. Topel. "The Evolution of Unemployment in the United States: 1968-1985." in NBER Macroeconomics Annual 1987, edited by Stanley Fischer. Cambridge, MA: MIT Press, 11-58.

O'Leary, Christopher J. "The Adequacy of Unemployment Insurance Benefits." Manuscript, W.E. Upjohn Institute for Employment Research, August 1994.

Pissarides, Christopher. Equilibrium Unemployment Theory, Basil Blackwell, Inc., 1990.

Raviv, Arthur. "The Design of an Optimal Insurance Policy." American Economic Review 69 (1979): 84-96.

Rothschild, Michael and Joseph Stiglitz. "Equilibrium in Competitive Insurance Markets: An Essay on the Economics of Imperfect Information." Quarterly Journal of Economics 90 (1976): 629-649.

Shavell, Steven. "On Moral Hazard and Insurance." Quarterly Journal of Economics 93 (November 1979): 541-562.

Shavell, Steven and Laurence Weiss. "The Optimal Provision of Unemployment Insurance Benefits over Time." Journal of Political Economy 87 (December 1979): 1347-1362.

Wang, Cheng and Stephen Williamson. "Unemployment Insurance with Moral Hazard in a Dynamic Economy." University of Iowa Working Paper, April 1995.

Zeldes, Stephen. "Consumption and Liquidity Constraints: An Empirical Investigation." Journal of Political Economy, 97 (April 1989): 305-347.

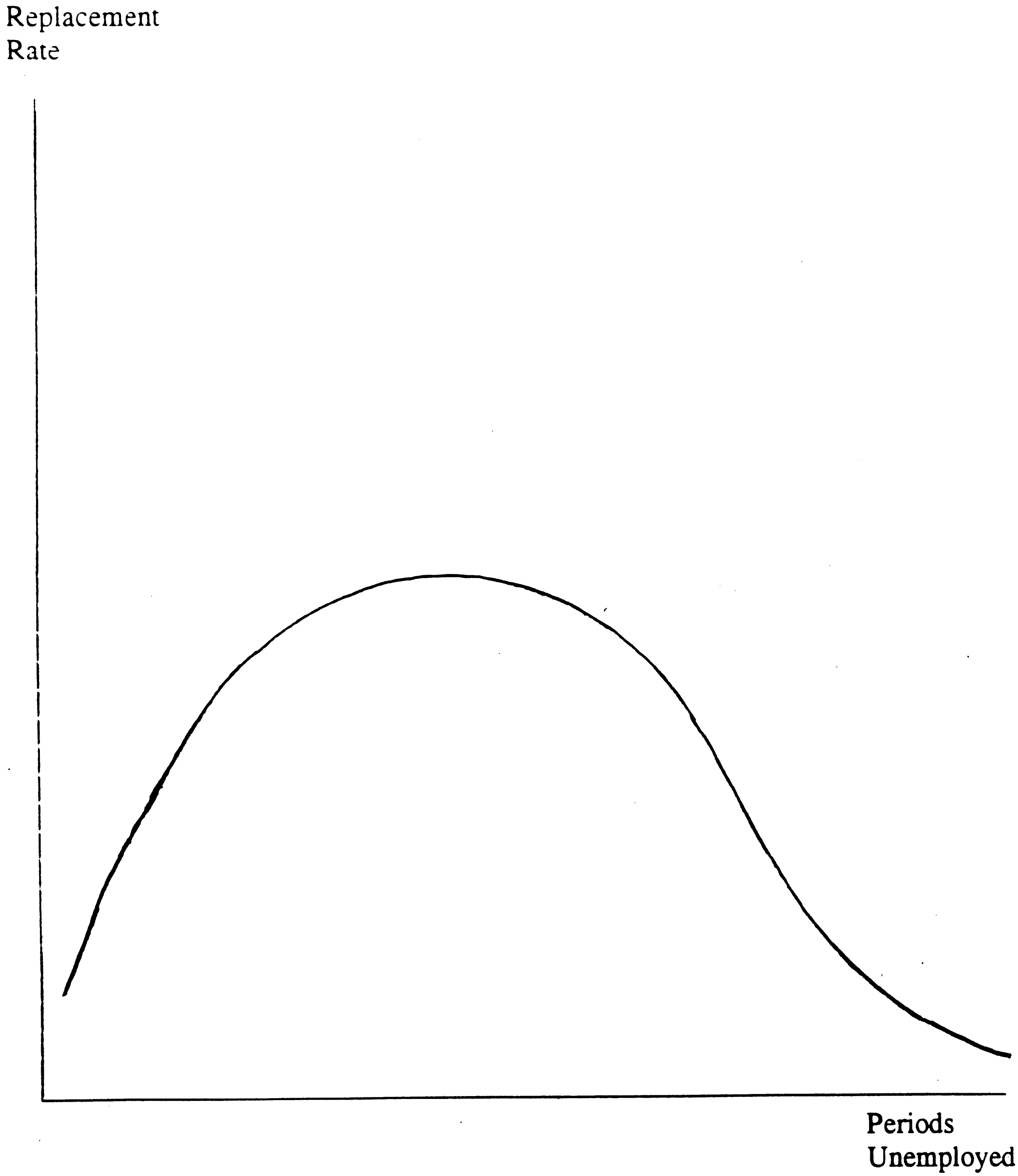


Figure 1

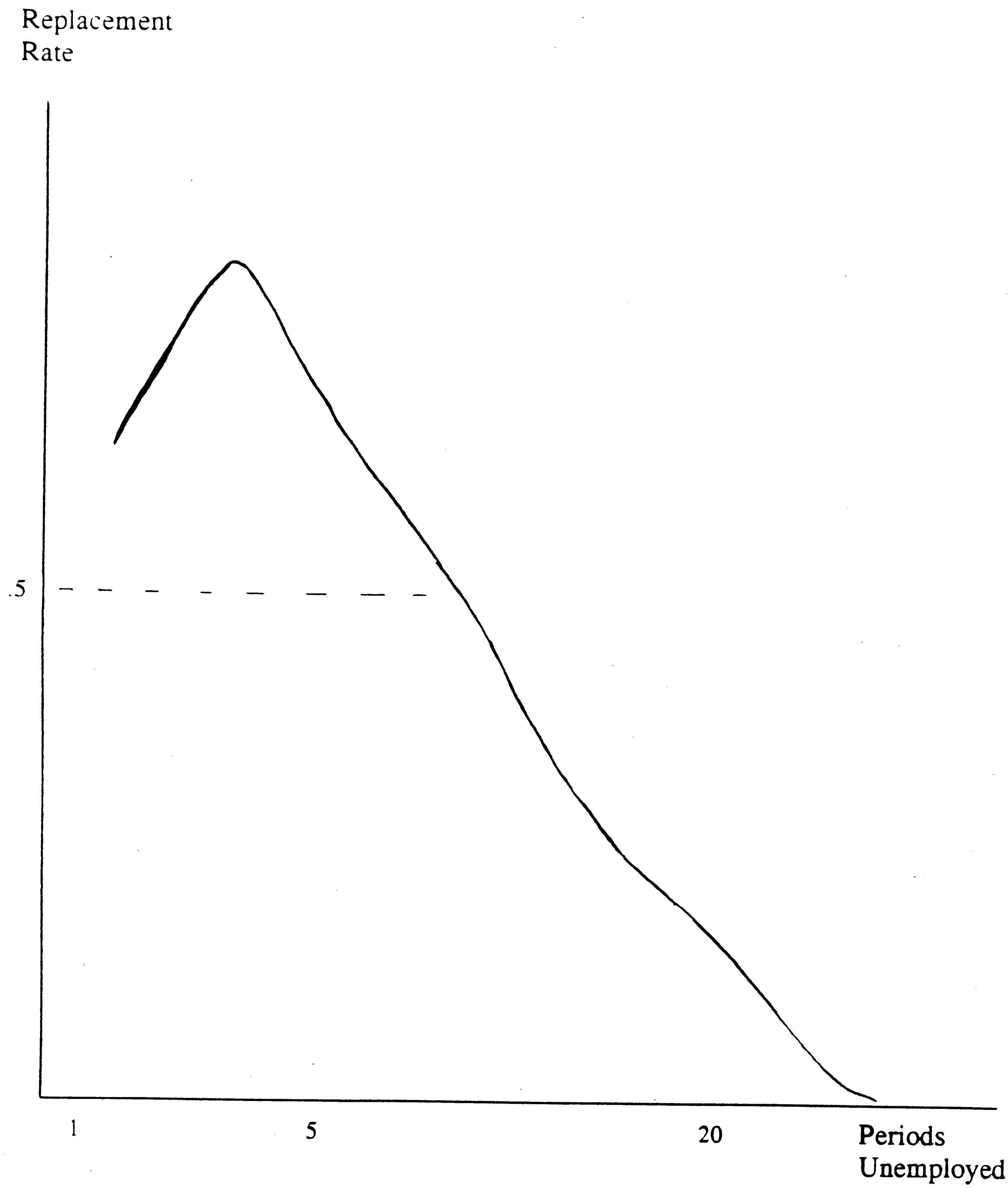


Figure 2

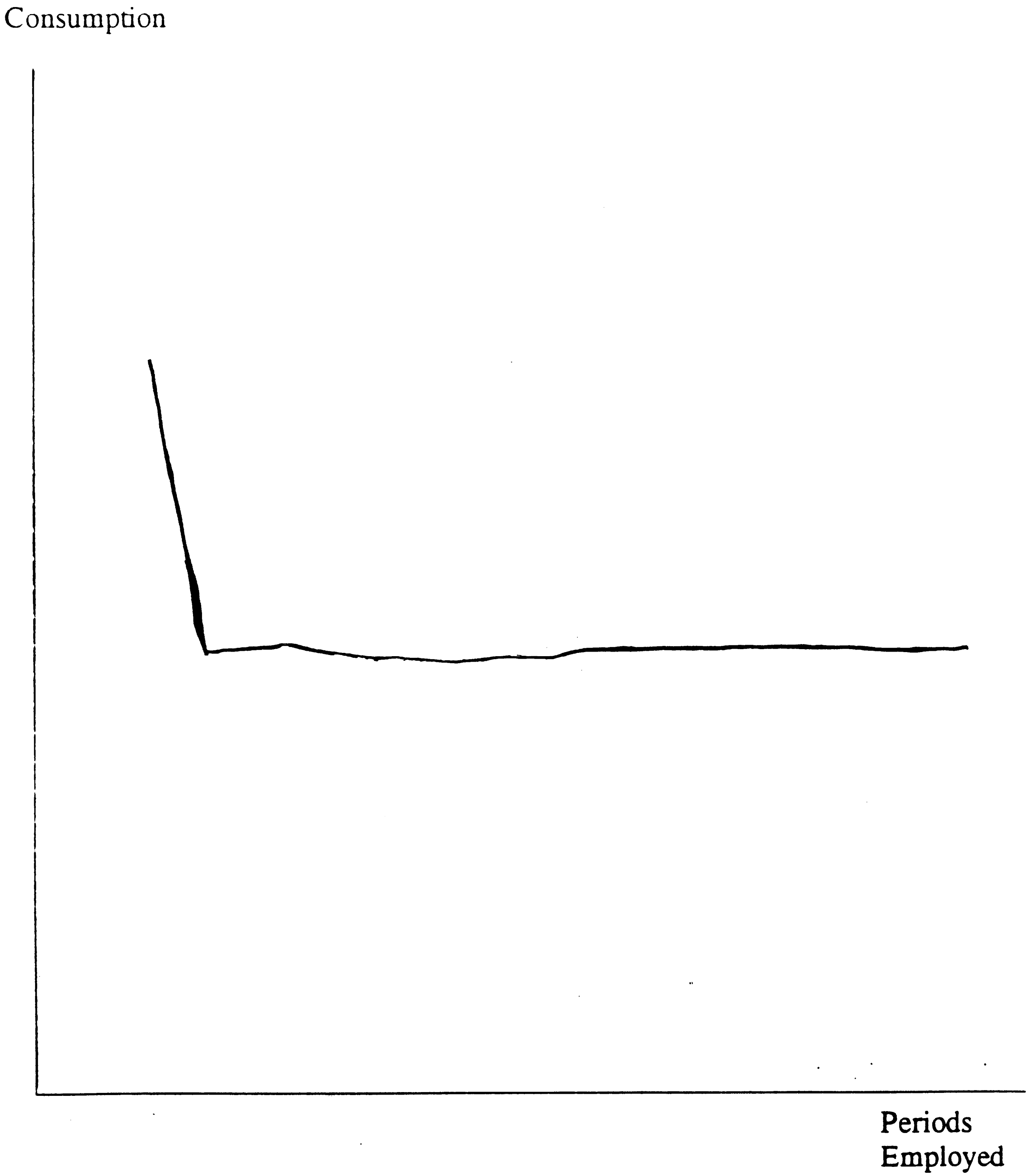


Figure 3

α	0	.01	.02	.05
Ω				
.9	(61%, 26)	(59%, 24)	(56%, 22)	(48%, 16)
.8	(61%, 26)	(60%, 22)	(55%, 20)	(42%, 14)
.7	(61%, 26)	(58%, 22)	(54%, 18)	(36%, 12)

Table 1

Optimal UI Programs with Slouchers but No Professionals

Square-root Utility

ϕ \ α	0	.01	.02	.05
0	(61%, 26)	(60%, 22)	(55%, 20)	(42%, 14)
.1	(64%, 28)	(60%, 26)	(57%, 22)	(45%, 14)
.2	(66%, 32)	(63%, 28)	(61%, 24)	(47%, 16)
.3	(68%, 36)	(67%, 32)	(64%, 28)	(51%, 18)

Table 2

Optimal UI Programs with Professionals and Slouchers

Square-root Utility

Ω = .8 in all cells

DRAFT FINAL REPORT
October 1995

UNEMPLOYMENT INSURANCE AND HOUSEHOLD WELFARE: Microeconomic Evidence 1980-93

Daniel S. Hamermesh and Daniel T. Slesnick*

*Edward Everett Hale Centennial Professor of Economics, University of Texas at Austin, and research associate, National Bureau of Economic Research; Professor of Economics, University of Texas at Austin. We thank Laurie Bassi, Paul Burgess, Carl Davidson, Stuart Low and Stephen Woodbury for helpful comments, and Jennifer Campbell for research assistance. Research underlying this study was carried out under Purchase Order No. M-5113-5-00-97-40 from the Advisory Council on Unemployment Compensation of the U.S. Department of Labor. The findings and conclusions expressed herein are solely those of the authors.

EXECUTIVE SUMMARY

This study examines the economic well-being of households that receive unemployment insurance (UI) benefits compared to those that do not. Well-being is measured by consumption flows that are derived from information on households' spending in the Consumer Expenditure Surveys from 1980-1993. For each quarter during this period we obtain the per-capita and equivalence-scale adjusted economic welfare of the two types of households. Adjusting for differences in the households' characteristics, we find:

> With UI benefits the average UI recipient household during this period had a level of economic well-being that was on average between 3 and 8 percent below that of otherwise identical households (depending on the welfare measure used).
>
> During a substantial part of this time the economic well-being of households that received UI benefits was at least that of other households.
>
> Unemployment benefits raised the well-being of recipient households by about 12 percent compared to where it would have been without benefits.
>
> There is no cyclical variation in the relative well-being of UI recipient households compared to others.

Taken together the findings in this study imply that during the 1980s and early 1990s states' UI programs did a satisfactory job of maintaining the well-being of UI recipients. Moreover, emergency programs enacted during recessions raised potential duration sufficiently to prevent the economic position of the average UI receipient from deteriorating. UI benefits may not be sufficient to maintain the consumption of unemployed Americans generally; but they do an adequate job in maintaining the well-being of households that receive them.

I. Introduction

How individuals spend their unemployment benefits and how well those benefits maintain households' incomes during times when a family member is unemployed are perhaps the crucial questions in the analysis of this particular form of social insurance. For example, the preamble to the House bill that became the Social Security Act of 1935 noted that its purpose was, "To alleviate the hazards of old age, unemployment...." The founders of these programs viewed maintaining consumption as central. Despite this, most research on unemployment insurance has dealt with other issues. This topic has, however, been examined carefully in a counting framework by, e.g., Burgess and Kingston (1978), Felder and Li (1980) and Browning (1995), all of whom examined spending changes in households during one member's compensated spell of unemployment. The issue has been placed in the context of the life-cycle theory of consumption by Hamermesh (1982), Dynarski and Sheffrin (1987) and Gruber (1994), who focused on the extent to which households containing unemployed workers depart from the spending path predicted by the theory.

The difficulty with this literature is that it has not been grounded in the formal consumer theory of the household. A huge body of research (e.g., as summarized by Deaton and Muellbauer, 1980) has demonstrated the fruitfulness of taking consumer theory seriously when studying households' expenditures and the well-being of household members. It is wrong to make comparisons across households without considering, for example, how to treat households of different sizes, how to compare the same household at different times and how to account for changes in the prices of goods bought by households distinguished by, for example, UI recipiency. In this study we therefore ground the analysis of the effects of UI benefits on recipients' well-being in the theory of the household, thus enabling us to answer the central question: How well do these benefits insure consumption streams against spells of unemployment?

The analysis focuses on the welfare effects of UI benefits, thus providing the first direct link of the literature on benefit adequacy to microeconomic theory. We base the evaluation of the effects of UI on its impact on consumption, the appropriate focus in light of the program's goals and in view of the fact that it is consumption, not income, that determines households' well-being. The comparison is not to the recipient household's pre-unemployment spending, but instead to otherwise identical households' economic welfare. The "bottom-line" result of the study is an upper bound on the difference in economic welfare between those households receiving UI benefits and other households.[1] We thus measure the extent to which UI benefits fail to protect households against income losses resulting from unemployment as compared to otherwise identical households that do not incur such losses.

We use representative quarterly cross-sections of households over a substantial period of time (1980-93). The estimates thus offer the first examination of how well unemployment insurance maintains consumption at different points of the business cycle. Such evidence speaks to the question of whether the emergency UI programs that have extended the potential duration of benefits have been sufficient to prevent cyclical declines in the well-being of households containing unemployed workers.

II. Measures of Economic Welfare

An essential first step in addressing the impact of unemployment insurance on the well-being of recipients is to develop a measure of individual welfare. Although many different measures are used in practice, most observers would agree that material well-being is fundamentally related to the goods consumed by individuals. The standard theoretical paradigm describes consumers as "rational" agents who choose that combination of goods and services that maximizes utility subject to the constraint of limited financial resources. Under this framework it is in principle possible to infer the level of individual welfare from the quantities consumed.

Specifically, assume that individuals maximize a (static) utility function subject to a budget constraint. The indirect utility function $V(\ )$ represents the maximum utility attainable for individual k with attributes $\mathbf{A_k}$ and total expenditure M_k facing prices $\mathbf{p}$:

$$V(\ ) = V(\mathbf{p}, M_k, \mathbf{A_k}) .$$

The indirect utility function represents the welfare derived from the consumption of goods and services but does not provide an intuitive, monetary measure of well-being. For that purpose we use the expenditure function $M(\ )$, which is the minimum level of total expenditure required to attain utility level V_k at prices $\mathbf{p}$:

$$M_k(\ ) = M(\mathbf{p}, V_k, \mathbf{A_k}) .$$

The expenditure function allows us to define a monetary measure of welfare using the indirect money metric utility function. This is defined to be the minimum expenditure required for individual k, who faces prices $\mathbf{p}^r$, to attain utility level $V(\mathbf{p}, M_k, \mathbf{A_k})$:

$$\mu(\mathbf{p}^r, \mathbf{A_k}; \mathbf{p}, M_k, \mathbf{A_k}) = M(\mathbf{p}^r, V(\mathbf{p}, M_k, \mathbf{A_k}), \mathbf{A_k}) . \tag{2.1}$$

The indirect money metric utility function provides a monetary measure of well-being that is ordinally equivalent to the individual's utility function. For fixed prices the money metric utility function increases if and only if utility and well-being rise.

The welfare function (2.1) provides a means of estimating the welfare of a household whose demographic composition does not change. Our problem, however, is that we wish to compare the relative levels of well-being of heterogeneous households. Is a UI-recipient household with four members and a expenditure level of \$10,000 as well off as a nonrecipient couple with \$5,000? To answer this more complicated question, the money metric utility function must be evaluated at a reference set of demographic characteristics $\mathbf{A}_r$. We can rewrite the welfare function as:

$$W_k = \mu(\mathbf{p}^r, \mathbf{A}_r; \mathbf{p}, M_k, \mathbf{A}_k)$$
$$= M_k/[P_k(\mathbf{p}, \mathbf{p}^r, V_k)\mathrm{m}_0(\mathbf{p}^r, V_k, \mathbf{A}_k)] \ . \tag{2.2}$$

In this alternative presentation of welfare P_k is a household-specific price index defined as the relative cost of attaining utility level V_k at prices $\mathbf{p}$ and $\mathbf{p}^r$ respectively:

$$P_k(\mathbf{p}, \mathbf{p}^r, V_k) = \mathrm{M}(\mathbf{p}, V_k, \mathbf{A}_k)/\mathrm{M}(\mathbf{p}^r, V_k, \mathbf{A}_k) \ .$$

The equivalence scale m_0 captures differences in needs across households and is the cost, relative to that needed by a reference household, of attaining a utility level V_k:

$$\mathrm{m}_0(\mathbf{p}^r, V_k, \mathbf{A}_k) = \mathrm{M}(\mathbf{p}^r, V_k, \mathbf{A}_k)/\mathrm{M}(\mathbf{p}^r, V_k, \mathbf{A}_r) \ .$$

The welfare measure in (2.2) can be interpreted as the level of per equivalent consumption. How does this differ from income-based estimates typically used to measure well-being? First and foremost, the welfare measure is based on the household's consumption rather than on an income concept such as earnings or pre-tax income. Within a single-period framework such as that described above consumption is clearly the appropriate concept, since utility is derived exclusively from the goods and services consumed.

Consumption is the preferred basis for a "snapshot" measure of welfare, but it is also a better proxy for lifetime income for reasons related to the permanent-income hypothesis. Assume that individuals choose stable paths of consumption through time based on their permanent incomes. Among those who have lower than usual annual incomes, permanent income is understated by current income. The reverse is true for individuals who have higher than usual income levels. In each case consumption is on average equal to permanent income, so that consumption approximates lifetime welfare more accurately than does current income.[2]

Our welfare index also incorporates household-specific price effects rather than measuring the effect of price changes through an aggregate price index such as the CPI. To see why this might be important, suppose that a poor household devotes a large fraction of its resources to food and housing compared to a household that is better off. If there are dramatic increases in the prices of food and housing, the cost of living facing the poor will have risen substantially relative to that facing the rest of the population. The usual deflation of income or expenditures by the CPI will fail to capture the differential impact of such a price change.

If two households have the same consumption, is it reasonable to conclude that they are equally well-off? In general, the answer depends on the composition of the two households and their relative needs. A single individual does not need the same level of consumption as a household with seven members. This aspect of welfare measurement is captured in (2.2) through the equivalence scales, which compare the relative cost of attaining a given welfare level to that facing a reference household.[3]

III. Data

A critical component to measuring the well-being of UI recipients using the consumer-theoretic model in Section II is a set of comprehensive data on consumption. Some studies of UI and spending have relied on measures of one or a few components of spending or on very ad hoc survey responses that are unlikely to generate data of the quality necessary to allow one to measure total consumption (e.g., Dynarski and Sheffrin, 1987; Browning, 1995).[4] We therefore rely on the only source of satisfactory spending data in the United States, the *Consumer Expenditure Surveys* (CEX), published by the Bureau of Labor Statistics. These surveys provide representative national samples that are used to calculate the expenditure weights that enter the monthly Consumer Price Index. Our sample includes quarterly observations beginning in the second quarter of 1980 and extending through the fourth quarter of 1993.

The basic observation in the CEX is the "consumer unit," defined either as a group of individuals who pool their monetary resources to make joint spending decisions, as individuals who are related by blood or legal arrangement, or as financially independent unrelated individuals. Financial independence is determined on the basis of whether the individual pays for at least two of the following three items: Food, housing and living expenses. The "reference" person, or "head" of the consumer unit, is the person who owns or rents the residence.[5] The coverage of the surveys can generally be described as the noninstitutionalized civilian population. Military personnel living off base are included, while those living on base are excluded. College students living outside their parents' homes are treated as separate consumer units. Travel expenditures abroad are part of the reported totals, but spending by U.S. citizens stationed overseas is not.

Since 1980 the surveys have been in a rotating panel format in which each consumer unit is interviewed for five consecutive quarters. Every quarter twenty percent of the households are dropped and replaced by a new group of consumer units. The first interview collects demographic information and a partial inventory of consumer durables. In the remaining four quarterly interviews detailed expenditure information covering the previous three months is collected. Each quarterly survey provides a sample of between 4000 and 6000 separate consumer units that is representative of the U.S. population.

The CEX reports on the out-of-pocket expenditures of the consumer unit. The BLS includes in the definition of total expenditure gifts and cash contributions to persons (and organizations) outside the household as well as spending on owner-occupied housing and durables. Contributions to pensions, retirement funds and Social Security are also included, while most in-kind transfers, such as employer-paid insurance and health care, Medicaid and government housing subsidies, are not. Food stamps, and meals and rent received as pay, are part of total expenditure to the extent that they are reported accurately.

We modify the BLS definition of consumption to correspond as closely as possible to the concept appropriate for the model described in Section II. We delete gifts and cash contributions (and discuss them separately later on), since altruism requires a conceptual formulation that differs from the basis of most studies of individual welfare. Pensions, retirement contributions and Social Security taxes are also deleted, since they are components of saving rather than consumption. Expenditures on owner-occupied housing and consumer durables are replaced by the appropriate rental equivalents, so that we obtain a flow of consumption of these items and avoid problems generated by, for example, lumpiness in purchases. For housing this is a straightforward exercise, since the estimated rental value of the home is reported in all but two of the surveys.[6] We calculate service flows from consumer durables using a method developed by Christensen and Jorgenson (1969) in which the services received correspond to the opportunity cost of holding the asset represented by the durable good.[7]

Income in the CEX is defined to be the combined income during the previous 12 months of all members of the consumer unit over 14 years of age. Questions related to income are asked only in the second and fifth interviews. If there is a new member in the consumer unit in the fifth interview, the earned income history is reconstructed for the previous quarters. The components of income included are: Wages and salaries, self-employment income, Social Security and other retirement income, interest, dividends, rental income, unemployment insurance, worker's compensation, veteran's benefits, public assistance, regular contributions for support, and other income. Unlike expenditures, no imputations are performed for nonresponses to the income questions. If the respondent did not provide values for major sources of income, such as wages and salaries, self-employment income or Social Security, the consumer unit was classified as an "incomplete income reporter."

For budgetary reasons only the urban population was sampled in 1982 and 1983. Since we want a continuous time series beginning in 1980, we delete all rural households from the CEX in the other years. Because very few households with a head over age 65 received UI benefits, such households were also excluded from the sample. A household was classified as a UI recipient if it reported receiving a positive level of benefits at some time during the last twelve months. In multiple-person households the CEX does not allow determining the identity of the recipient, and in no household can we determine the actual duration of the benefits.[8]

Throughout this study the accounting period is the most recent twelve months from the quarter indicated in the figures or the discussion. Each household can be in the sample from one to four quarters, depending on attrition. The focus on annual incomes, and benefits received during the past twelve months, is consistent with accounting periods that are standard in government programs. This does mean, however, that we cannot necessarily focus on the (possibly few) weeks when a household member might have been unemployed. That lacuna represents the trade-off one must make to obtain good data on spending flows over time intervals that are consistent with the ways in which economists usually measure consumption.

In the 55 quarters from 1980:2 to 1993:4 a total of 195,689 consumer units are in the sample. Each household is included in each quarter in which it appears in the CEX sample. The crucial point, however, is that in every quarter the sample is generally representative of households in the U.S. Of the households 12,999 (6.6 percent) are UI recipients, but the proportion fluctuates substantially over the fourteen years. Figure 1 presents this proportion in each quarter. Participation follows the business cycle fairly closely: An average of 9.2 percent of the sample received benefits around the business-cycle trough between 1982:2 and 1984:2, while only 4.7 percent did so during the business-cycle peak period between 1987:4 and 1989:4.

Table 1 compares the composition of the UI and non-UI populations. Most of the data reflect what has been observed in previous comparisons of UI recipients to other workers, but the focus on households is slightly novel. UI recipients have below-average total expenditures and income despite having more earners in the household.[9] The average age of the head of the household is identical for both recipients and nonrecipients, although the average size of the household is substantially larger among recipient households. Nonwhites and female-headed households are underrepresented among UI recipients, as are households living in the South. Perhaps most significant is the fact that UI households have, on average, less education than nonrecipients. While 53 percent of non-UI households have at least some college education, only 37 percent of UI households have similar educational levels; and only 13 percent of UI households

FIGURE 1- PROPORTION UI RECIPIENTS

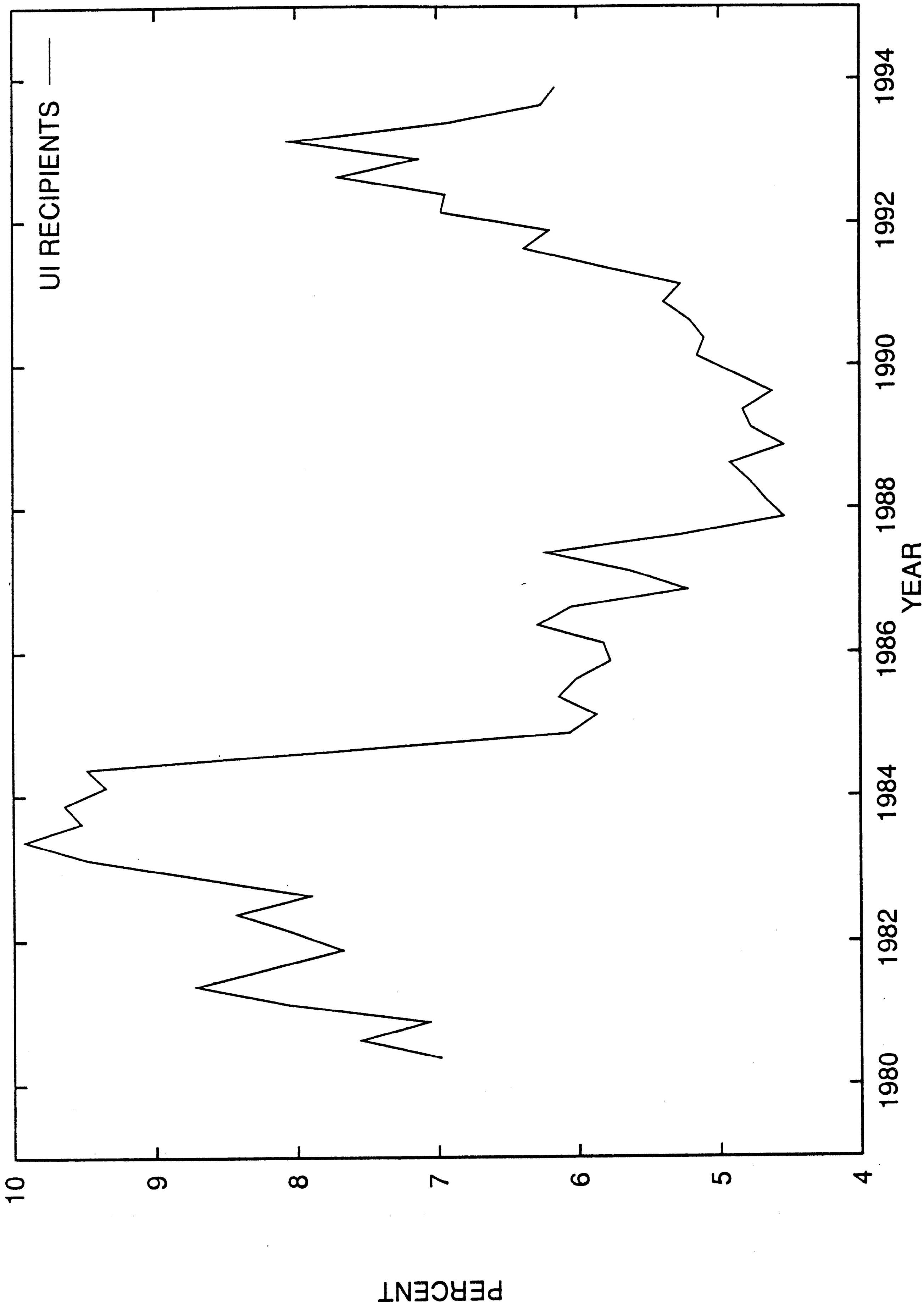


Table 1. Sample Means (standard deviations in parentheses)

	Overall	UI	Non-UI
Total expenditure	$22,918	$21,132	$23,045
	($15,013)	($12,191)	($15,186)
Before-tax income	$29,240	$26,632	$29,446
	($22,978)	($18,815)	($23,264)
Age of head	39.68	39.31	39.70
	(12.53)	(11.62)	(12.59)
Household size	2.79	3.14	2.76
	(1.59)	(1.65)	(1.59)
Number of earners	1.61	1.90	1.59
	(0.94)	(0.96)	(0.93)
Race of head			
White	0.84	0.86	0.84
Nonwhite	0.16	0.14	0.16
Gender of head: Male	0.75	0.82	0.74
Education of head			
No school	0.00	0.00	0.00
1 - 8 grade	0.06	0.08	0.06
9 - 11 grade	0.12	0.15	0.11
High school graduate	0.30	0.40	0.30
Some college	0.25	0.24	0.26
College graduate	0.14	0.08	0.14
Graduate school	0.12	0.05	0.13
Region			
Northeast	0.21	0.22	0.21
Midwest	0.26	0.30	0.26
South	0.29	0.22	0.29
West	0.24	0.26	0.24
Married	0.58	0.67	0.57
Earner Composition			
Head	0.38	0.29	0.39
Head & spouse	0.28	0.34	0.27
Head, spouse & others	0.09	0.13	0.09
Head & others	0.12	0.16	0.12
Spouse	0.02	0.02	0.02
Spouse & others	0.01	0.01	0.01
Others	0.02	0.02	0.02
No earners	0.07	0.03	0.08

are college graduates compared with 27 percent among non-UI households. Finally, UI-recipient households are more likely to have both spouses working in the past twelve months than are non-recipient households.

While the average income of UI recipients is substantially below that of nonrecipients, one would expect there to be differences across subgroups of the population. In the first two columns of Table 2 we examine the relative income and expenditure levels for groups classified by age, race, sex, education and region of residence. Younger recipients and those with lower educational attainment have higher incomes relative to comparable nonrecipients, while the average incomes are roughly the same for female-headed households. The income gap between recipients and nonrecipients is smaller among nonwhites than among whites. Among the four Census regions the gap is larger in the South, where we saw that those receiving UI are underrepresented.

Do the differences narrow when we examine total expenditure instead of income? In general the answer is yes. The fourth column of Table 2 indicates that the differences in average total expenditures between UI recipients and nonrecipients are closer in four of the five age categories than are differences in income. Only among households with a head age 55-64, where non-UI households are more likely to contain retirees and thus be saving less, is this not true. This is also found for whites, nonwhites, and male-headed households, and most educational categories.[10] In general, however, the qualitative results are the same whether one looks at total expenditure or income. Young, uneducated recipient households have higher total expenditures relative to nonrecipients, as do female-headed households. Comparing the four main Census regions, recipients living in the South have the lowest average total expenditure relative to the rest of the population. Households with both spouses working during the past twelve months have the

Table 2. **Household Income and Expenditures, Levels and Ratios for Recipient and Other Households**

	Income		Expenditures	
	UI	UI / Non-UI	UI	UI / Non-UI
Overall	$26,632	0.90	$21,132	0.92
Age of head				
16 - 24	$16,730	1.39	$13,997	1.22
25 - 34	$23,980	0.88	$18,516	0.92
35 - 44	$28,675	0.81	$23,430	0.86
45 - 54	$31,491	0.84	$24,467	0.86
55 - 64	$28,640	0.98	$22,753	0.95
Race of head				
White	$27,437	0.89	$21,688	0.91
Nonwhite	$21,688	0.95	$17,740	0.96
Gender of head				
Male	$28,838	0.85	$22,319	0.87
Female	$16,885	1.00	$15,825	1.02
Education of head				
No school	$15,817	0.99	$15,127	0.97
1 - 8 grade	$20,805	1.20	$17,785	1.09
9 - 11 grade	$22,496	1.21	$18,161	1.10
High school graduate	$25,981	0.97	$20,069	0.95
Some college	$28,013	1.01	$22,837	1.03
College graduate	$33,348	0.85	$25,906	0.91
Graduate school	$37,611	0.85	$29,151	0.90
Reigion of residence				
Northeast	$27,739	0.93	$22,277	0.94
Midwest	$26,572	0.92	$19,966	0.92
South	$24,167	0.84	$19,707	0.87
West	$27,795	0.90	$22,706	0.93
Earner Composition				
Head	$16,905	0.78	$15,647	0.87
Head & spouse	$30,847	0.77	$22,823	0.81
Head, spouse & others	$40,540	0.84	$29,169	0.85
Head & others	$28,405	0.91	$22,294	0.88
Spouse	$19,966	0.76	$19,104	0.76
Spouse & others	$29,512	0.92	$26,157	0.89
Others	$19,642	1.18	$20,403	1.05
No earners	$7,440	0.80	$11,849	0.91

lowest ratio of expenditures of UI recipients compared to others among households containing different combinations of earners.

IV. Welfare Comparisons of UI Recipients and Others

Under ideal circumstances the welfare function in (2.2) could be estimated by specifying a parametric form for the indirect utility function, fitting the implied demand functions to price and demand data and recovering the expenditure function by integrating backwards from the estimated demand functions (as in Hausman, 1981). Since the sample period is fairly short (14 years) and purchase prices are not reported in the CEX, it is very difficult to obtain precise estimates of price effects. These difficulties make this direct approach impractical here.

As a second-best alternative we approximate the welfare function using price indexes and equivalence scales that do not require a formal econometric model of demand. For each household we use a Paasche index defined as:

$$P_k^{-1}(\mathbf{p}, \mathbf{p}^r, V_k) = \sum_{n=1}^{N} s_{nk}[\, p_n^r / p_n \,] \,,$$

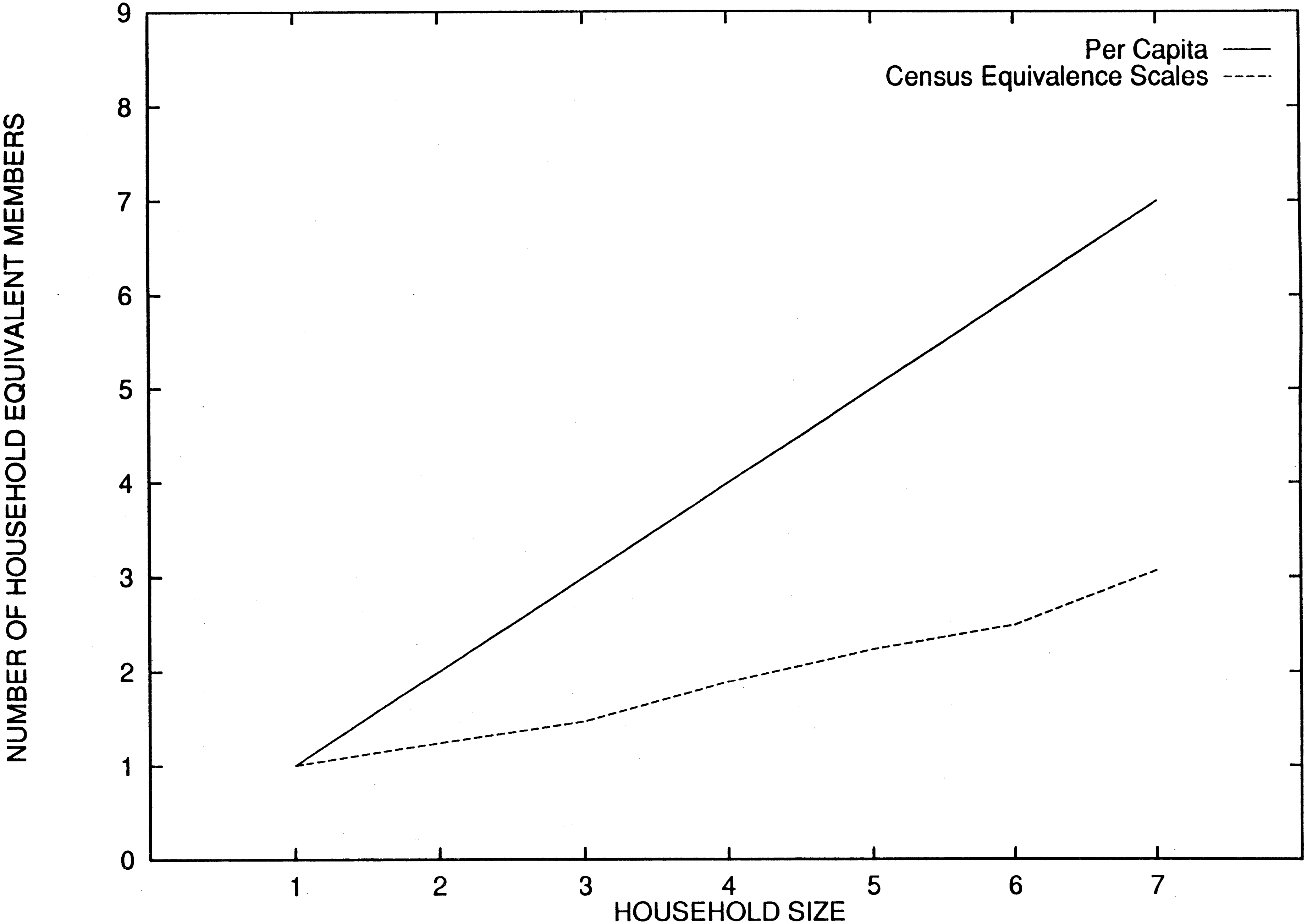
where s_{nk} is the budget share of good n for household k, and p_n is the price of the n'th commodity. This price index is calculated using the shares reported in the CEX for each household and the implicit price deflators of personal consumption expenditures in the National Income and Product Accounts. Since the Paasche index assumes that the weights s_{nk} are fixed, inflation will be underestimated relative to an index that incorporates households' adjustments in spending patterns that arise from changes in relative prices. Previous estimates of the "substitution bias" (for example, Braithwait, 1980, and Manser and MacDonald, 1988), however, indicate that it is very small.

The second critical component of the welfare function is the household equivalence scale, which is designed to measure the needs of heterogeneous households. Within this framework a larger household requires more resources to maintain the same level of well-being so that, other things equal, an increase in family size reduces welfare. The empirical issue is to determine the magnitude by which well-being decreases with each additional household member at a constant total consumption. A per-capita adjustment implicitly assumes that each additional family member, regardless of age or sex, increases the needs of the household by the same amount. While easy to implement, this approach is unlikely to provide an adequate representation of needs, since it presumes that the requirements of a child are the same as those of an adult. Economies of scale in consumption are also ignored.[11]

An alternative approach is to represent needs by an estimate of the number of adult-equivalent members in the household. One way of doing this is to base the equivalence scale on the nutritional intake required to avoid malnutrition. Adding up the costs of purchasing these nutrients for the different types of households provides the basis for comparing their relative needs. These types of scales are implicit in the poverty thresholds used by the Bureau of the Census to calculate the national poverty rate.

Figure 2 compares the per-capita adjustment to household spending with the equivalence scales used by the Census.[12] The two estimates represent polar extremes in terms of their implicit assumptions concerning economies of scale in consumption. The per-capita adjustment assumes that a family of four requires four times the expenditure of a single individual to attain the same level of well-being. The corresponding level for the Census equivalence scales is 1.89. For a family of seven or more individuals, the scale is only 3.09. While neither set of estimates is

FIGURE 2 -- HOUSEHOLD EQUIVALENCE SCALES



plausible, they provide reasonable upper and lower bounds on our estimates of welfare. The per-capita adjustment overdeflates consumption and underestimates the level of welfare, while the Census' equivalence scales do the opposite. In what follows we compare living standards of UI recipient households to those of nonrecipients using both types of adjustments.

The average logarithmic (percentage) differences in economic welfare between UI recipients and others are presented in the first column of Table 3 using per-capita consumption as an estimate of well-being. On average the welfare of UI recipients is 83.4 percent of nonrecipients'. There is, however, substantial variation across subgroups classified by age, race, sex, education and region of residence. UI households with a head age 16-24 have an average welfare level that is 3.6 percent below that of nonrecipients. The corresponding figure for those with a head age 45-54 is 21.1 percent. The gap between recipients and nonrecipients is narrower for nonwhites as well as for female-headed households. UI recipients with low educational attainment have average welfare levels much closer to those of nonrecipients compared to the gap among households with more education.

The second column of Table 3 repeats these tabulations using per-equivalent consumption as the welfare measure. Using this representation of well-being the discrepancy between UI recipients and others is narrower than when per-capita consumption is used. This primarily results from UI households being larger, but the effect is diminished because of the substantial economies of scale in consumption implicit in the equivalence scales. On average UI recipients have a welfare level that is 91.1 percent of nonrecipients' using this representation of well-being. However, the qualitative differences by demographic group that were shown in column (1) are preserved. The differences between UI recipients and nonrecipients are again smaller for

Table 3. **Welfare Difference - UI vs Non-UI (percent differences)**

	Total		**Unexplained**	
	W1	**W2**	**W1**	**W2**
Overall	-16.6	-8.9	-7.7	-3.6
Age of head				
16 - 24	-3.6	21.4	10.5	22.9
25 - 34	-17.5	-10.7	-9.7	-5.7
35 - 44	-17.9	-15.2	-7.4	-6.9
45 - 54	-21.1	-15.8	-10.3	-7.7
55 - 64	-19.6	-10.1	-10.9	-3.5
Race of head				
White	-18.3	-10.4	-8.0	-3.8
Nonwhite	-11.1	-3.8	-6.1	-2.8
Gender of head				
Male	-19.7	-13.8	-7.9	-4.2
Female	-9.1	0.0	-7.0	-1.1
Education of head				
No school	-15.3	-9.9	-12.6	-10.4
1 - 8 grade	-2.1	6.0	-2.2	3.4
9 - 11 grade	3.9	13.4	1.7	7.1
High school graduate	-12.0	-7.2	-9.8	-6.6
Some college	-5.9	2.7	-7.4	-3.0
College graduate	-19.6	-14.7	-18.1	-14.2
Graduate school	-13.3	-8.8	-13.4	-10.2
Reigion of residence				
Northeast	-14.9	-6.6	-7.7	-3.5
Midwest	-14.6	-6.5	-4.8	-1.2
South	-23.3	-15.2	-12.4	-8.1
West	-15.3	-8.7	-7.1	-2.8
Earner Composition				
Head	-16.6	-11.6	-7.4	-6.9
Head & spouse	-19.7	-18.4	-5.8	-5.9
Head, spouse & others	-18.7	-17.6	-4.9	-6.0
Head & others	-10.0	-9.6	-4.3	-5.0
Spouse	-25.4	-25.2	-13.8	-14.8
Spouse & others	-17.5	-14.5	-8.3	-7.2
Others	2.8	4.4	-1.7	-0.9
No earners	-19.4	-9.8	-11.0	-6.7

nonwhites, female-headed households and those where only the head is working. Young household heads who receive UI attain higher welfare levels relative to nonrecipients, as do those with some education but less than a high school diploma.

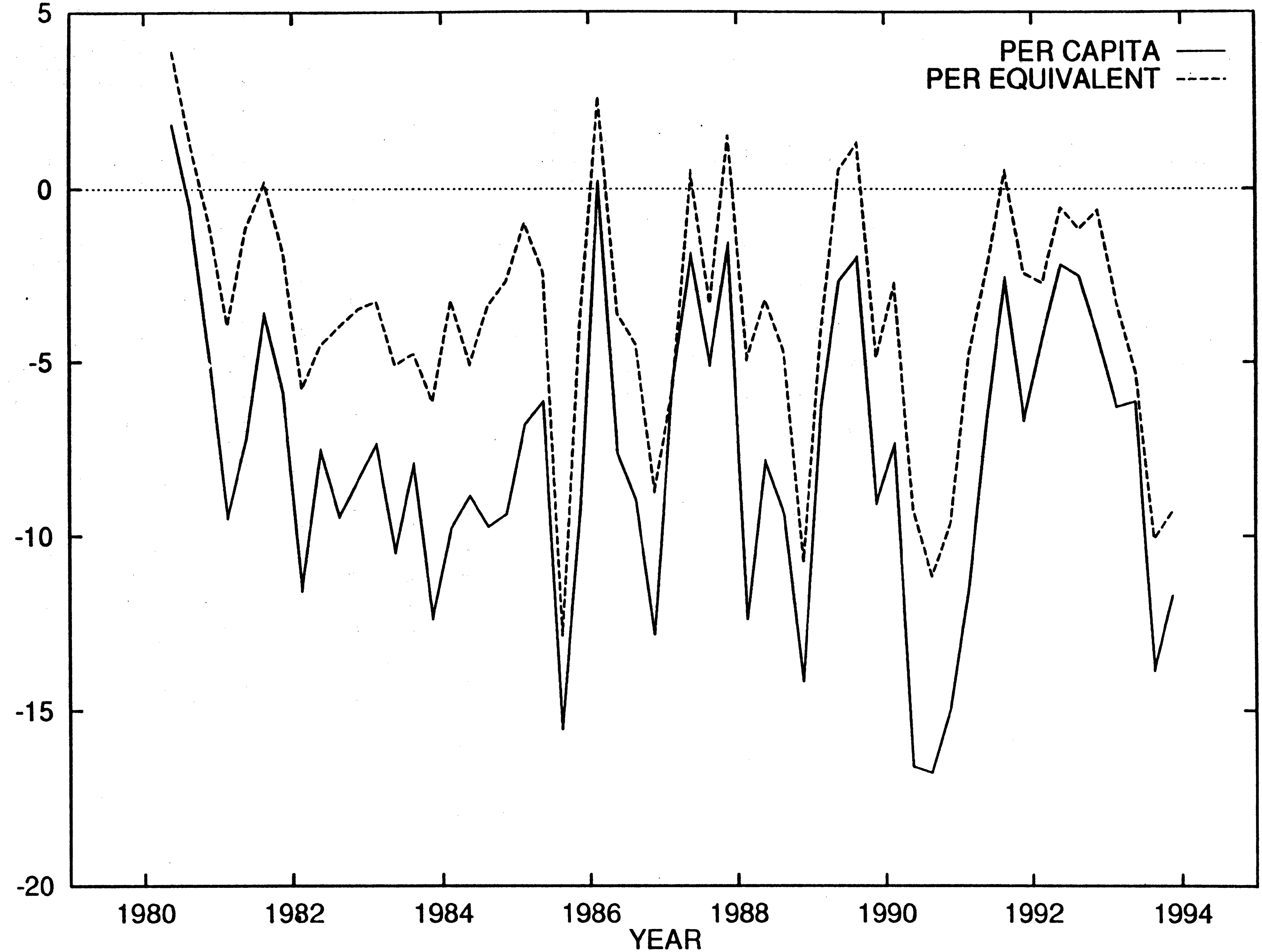
UI recipients are on average worse off than nonrecipients; but we saw in Table 1 that they also have different characteristics that could explain the lower welfare levels. To examine the role of compositional differences between the two groups we measure the average logarithmic differences in welfare levels accounting for differences in age, race, sex, education and region of residence.[13] Each figure in columns (3) and (4) thus adjusts for all the other demographic differences among households except the earner composition of the family and the particular characteristic listed in that section of the table.

Comparing columns (1) and (3), or (2) and (4), in Table 3, a large proportion of the differences between the average welfare level of the UI and non-UI samples can be explained by differences in their characteristics. The "unexplained" difference in the per-capita welfare measure averages -7.7 percent over the 55 quarters, although there is substantial variation from quarter to quarter. Of the variables included to create the adjusted differences, the differences in the human-capital characteristics alone explain most of the discrepancies in welfare levels. After accounting for age and education the unexplained differences only average -7.0 percent over the sample period using measure W_1. Adjusting the per-equivalent welfare measure W_2 yields qualitatively identical results, but the unexplained differences are even smaller. The fundamental result is that, once we account for differences in households' characteristics, a bit more than half of the differences in welfare between UI recipient households and others disappears.

Figure 3 shows how the unexplained differences vary over time. Clearly, there is substantial quarter-to-quarter variation in this measure. Indeed, using W_1 we find that in 9 quarters the adjusted welfare of recipient households is actually at least that of nonrecipient households.[14] The equations are estimated to account for seasonal changes, so there is no seasonal variation in these measures. They could, however, show cyclical variation, but they do not. Even though administrative data suggest that replacement rates fall in recessions (because higher-wage workers, whose benefits are limited by state maxima, constitute a larger fraction of recipients), the evidence here suggests that cyclical losses in well-being relative to those of nonrecipient households are small or nonexistent. This result suggests that the emergency programs in place during recessions in the 1980s sufficed to prevent UI recipient households' welfare from falling further below that of nonrecipient households.

While the central focus has and should be on differences between the average UI recipient household and other households, it is interesting to focus too on the impact of UI at the lower end of the distribution of income. We thus conclude this section by considering the role of UI benefits in relation to the poverty status of households. While the average welfare of recipients is substantially lower than that of nonrecipients, are they at greater risk of falling below the poverty line? To answer this question we arbitrarily choose the welfare of a family of four with $14,000 in 1993 as a "poverty threshold." Using per-equivalent consumption W_2 as the welfare measure, 9.6 percent of nonrecipient households are below the threshold compared to 9.3 percent of those receiving benefits. The poverty rates using the per-capita welfare measure W_1 are 9.6 percent and 10.4 percent respectively. Thus despite the fact that the mean of the distribution for recipients is roughly 10 to 15 percent below that of nonrecipients (depending on the welfare measure), those

FIGURE 3- UNEXPLAINED WELFARE DIFFERENCE UI VS. NONUI



receiving benefits are not at substantially greater risk of falling into poverty. Both because the UI population necessarily includes people with recent work histories and because UI benefits maintain the incomes of low-wage workers especially well due to states' benefit maxima, the chances of UI-recipient households being in poverty are essentially the same as those facing other households.

V. Expenditure Patterns

In the previous section we concentrated on the average well-being of UI recipients; but it is also important to examine how the difference between them and nonrecipients is reflected in their allocations of resources across different goods. Do those who receive benefits spend a larger fraction of their budgets on necessities? Do unemployment and the receipt of UI benefits result in a change in expenditure patterns? Does unemployment decrease spending on what one might view as discretionary or postponable items, for instance, purchases of durables or gifts to people outside the household? To answer these questions we examine the allocation of total expenditure across six broad categories of consumption:

1. Energy -- expenditures on electricity, natural gas, heating oil and gasoline.

2. Food -- expenditures on all food products, including tobacco and alcohol.

3. Consumer Goods -- expenditures on all other nondurable goods included in consumer expenditures.

4. Housing -- the service flow from owner-occupied and rental housing.

5. Durable Services -- the service flows from consumer durables such as cars and major appliances.

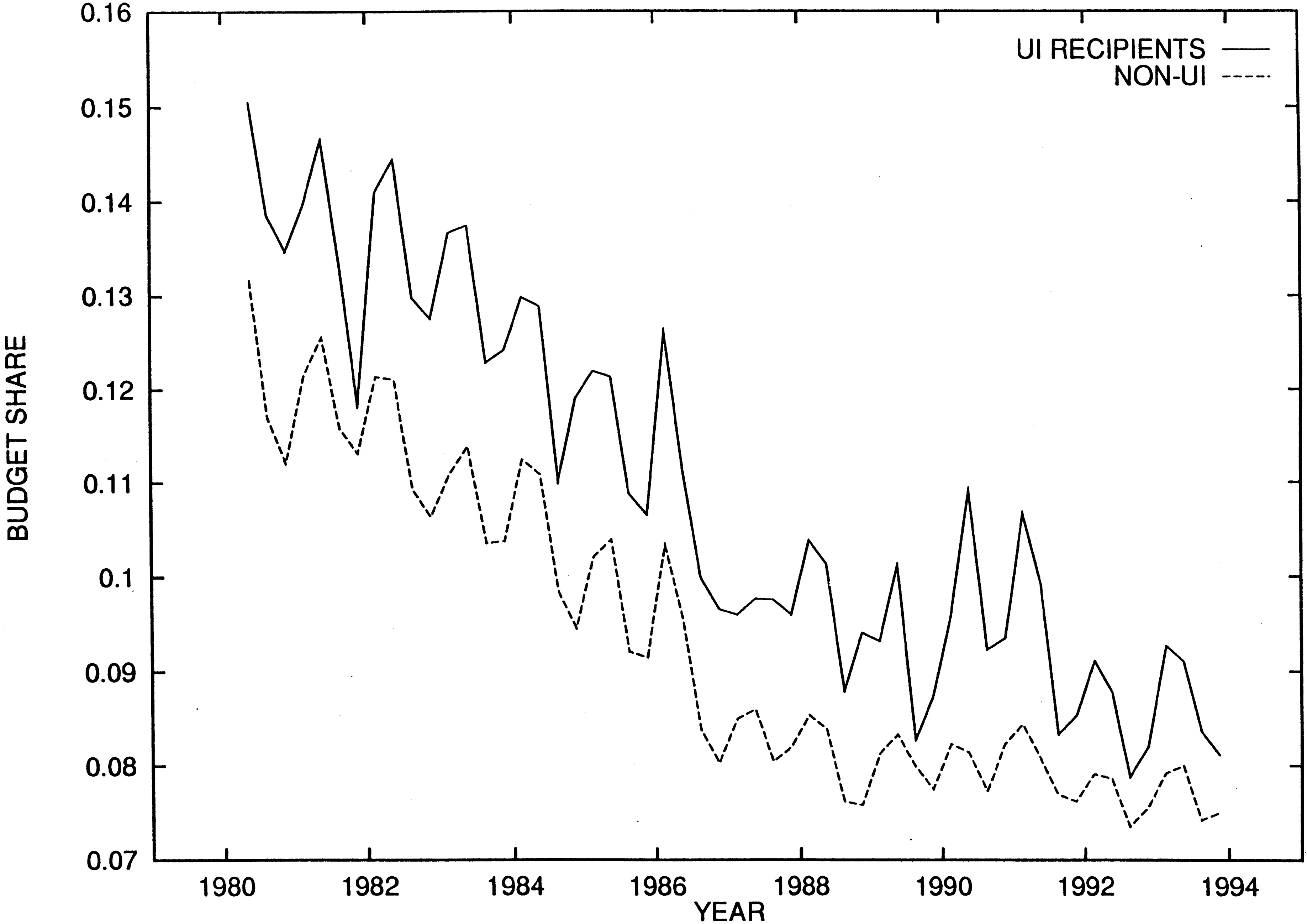
6. Consumer Services -- expenditures on consumer services, such as car repairs, medical care, entertainment and so on.

Figures 4-9 show the average shares of the expenditure groups over the period 1980:2-1993:4 for the UI and non-UI households. The trends in the average shares of all commodity groups are the same for the two sets of households. The share of spending on energy decreases over the sample period, as does the proportion of expenditures allocated to food. Spending on housing and, to a lesser extent, durables and consumer services rises over the sample period. Spending on consumer goods exhibits a classical seasonal pattern but comprises a small fraction of total expenditure and exhibits little trend over the 55 quarters.

The surprising fact is the absence of major differences in expenditure patterns between recipients and nonrecipients. Households receiving benefits spend larger proportions of their budgets on necessities such as energy and food, but the differences are small. Over the entire sample period the share of spending on energy by UI-recipient households is 1.6 percent higher, while the average food share is 1.8 percent higher. Expenditures on housing, consumer services and durable services comprise smaller fractions of total spending in almost every quarter, but again, the magnitudes of the differences are modest. Shares of spending on consumer goods are roughly the same for recipients and nonrecipients.

While average expenditure patterns are very similar between the two groups of households, there could be offsetting compositional effects. Specifically, UI households are typically larger, headed by white males and overrepresented in the Midwest and West, and these differences undoubtedly affect their allocation of total consumption across goods and services. To isolate differences in spending patterns between observationally equivalent recipients and nonrecipients we estimate Engel curves for each good. Budget shares for each of the six commodity groups are regressed against a quadratic in the logarithm of total expenditure and vectors of dummy variables

FIGURE 4- BUDGET SHARE OF ENERGY



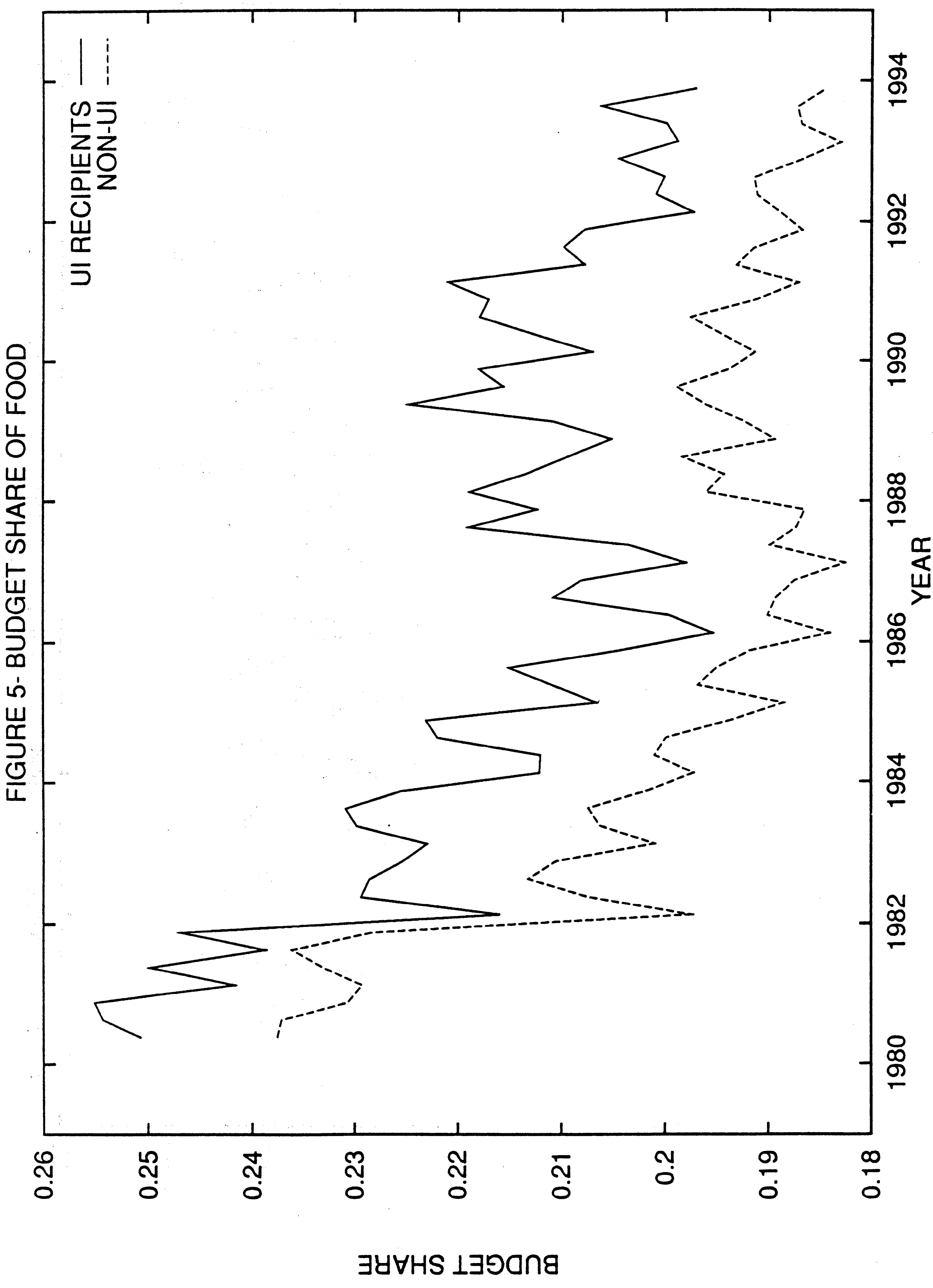


FIGURE 5- BUDGET SHARE OF FOOD
UI RECIPIENTS
NON-UI
BUDGET SHARE
0.26
0.25
0.24
0.23
0.22
0.21
0.2
0.19
0.18
YEAR
1980
1982
1984
1986
1988
1990
1992
1994

FIGURE 6- BUDGET SHARE OF CONSUMER GOODS

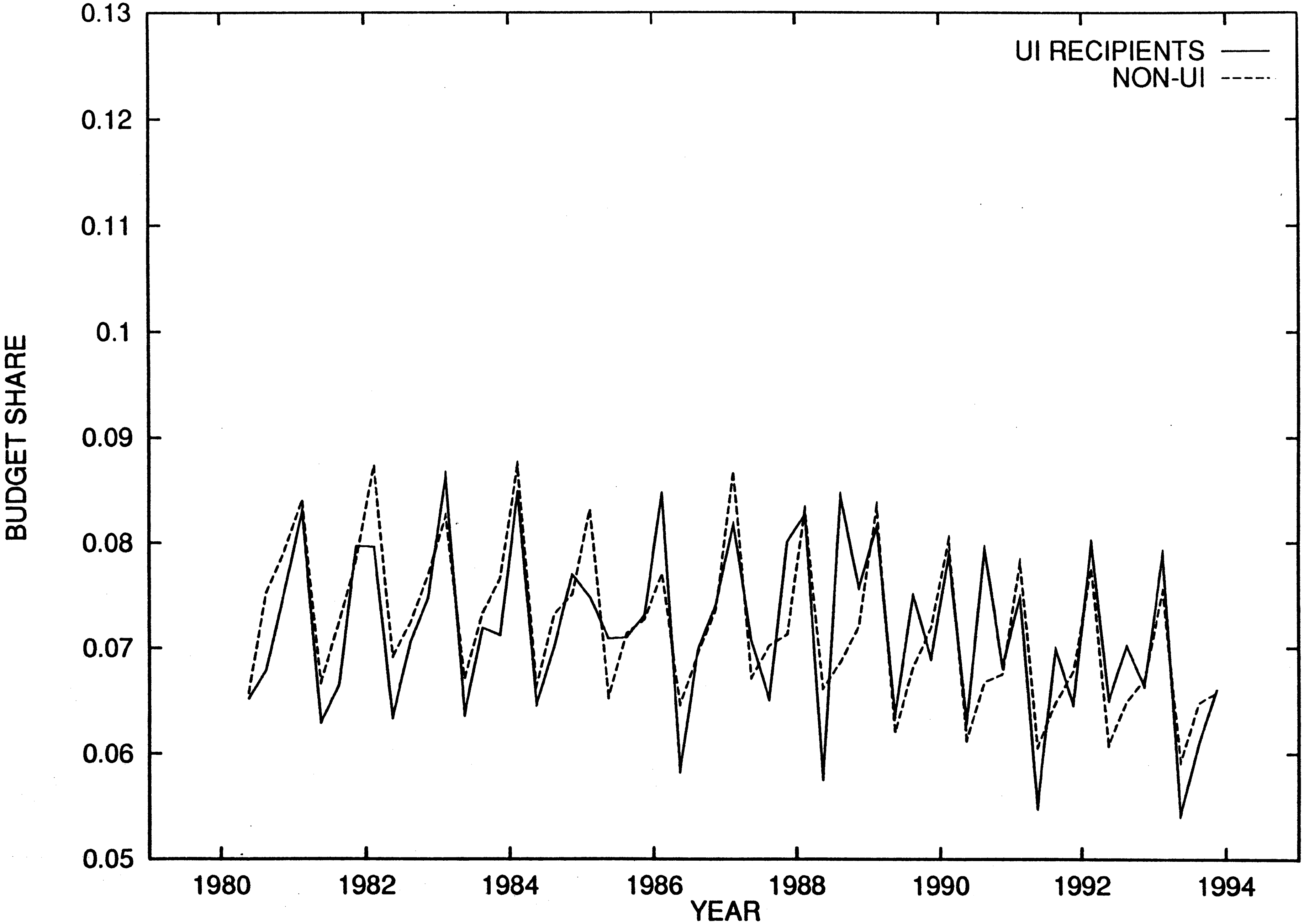


FIGURE 7- BUDGET SHARE OF HOUSING

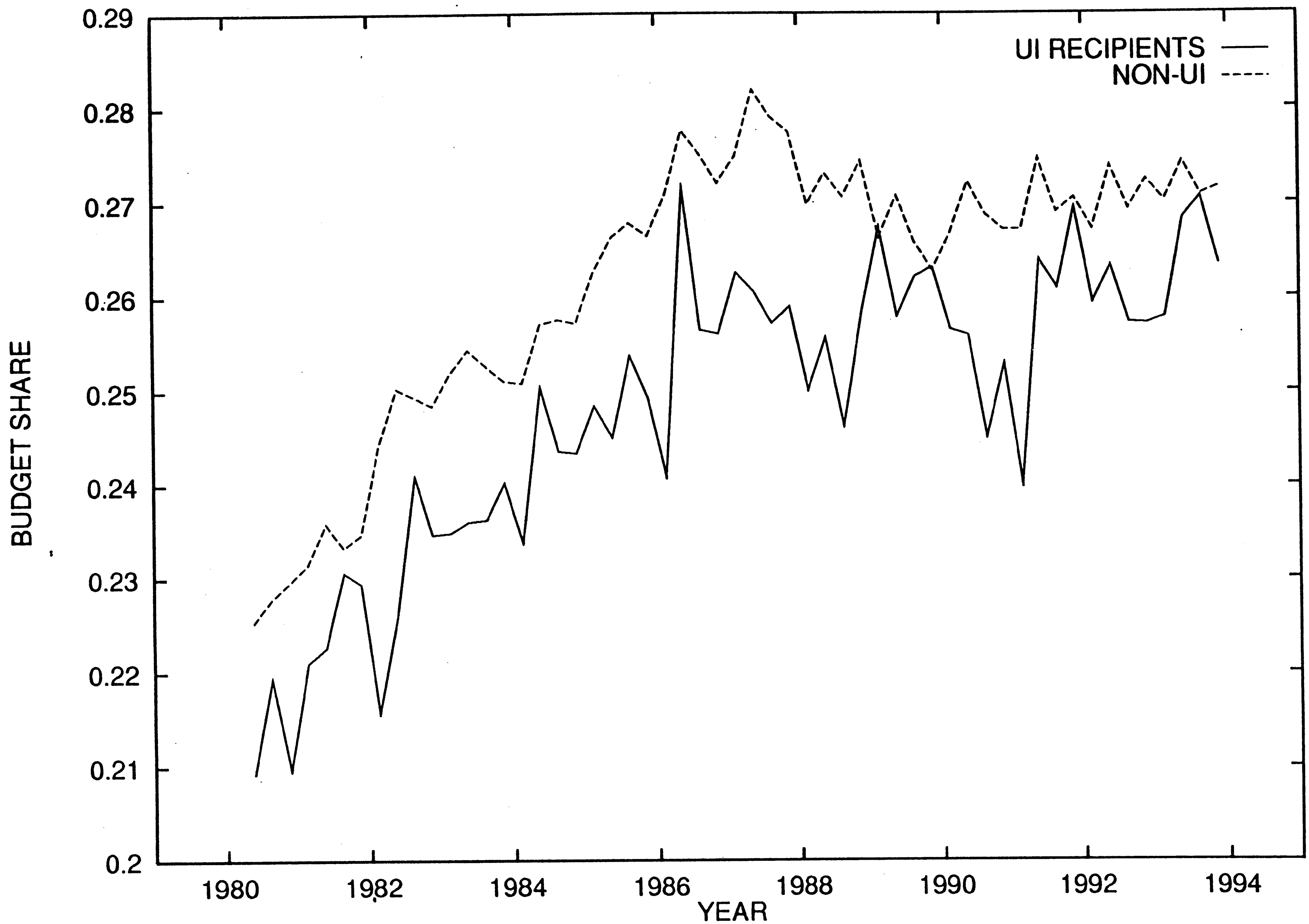


FIGURE 8- BUDGET SHARE OF DURABLES

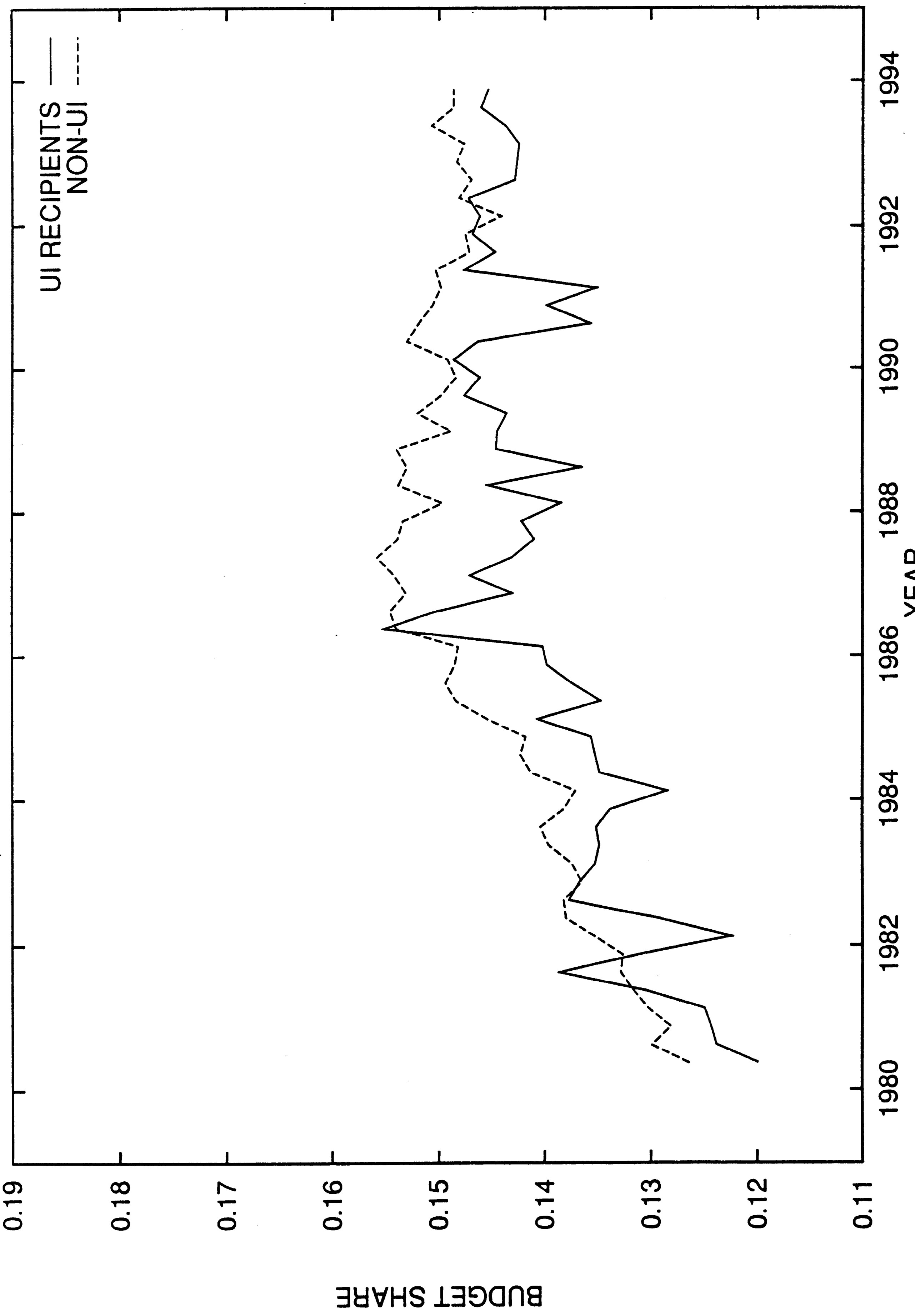
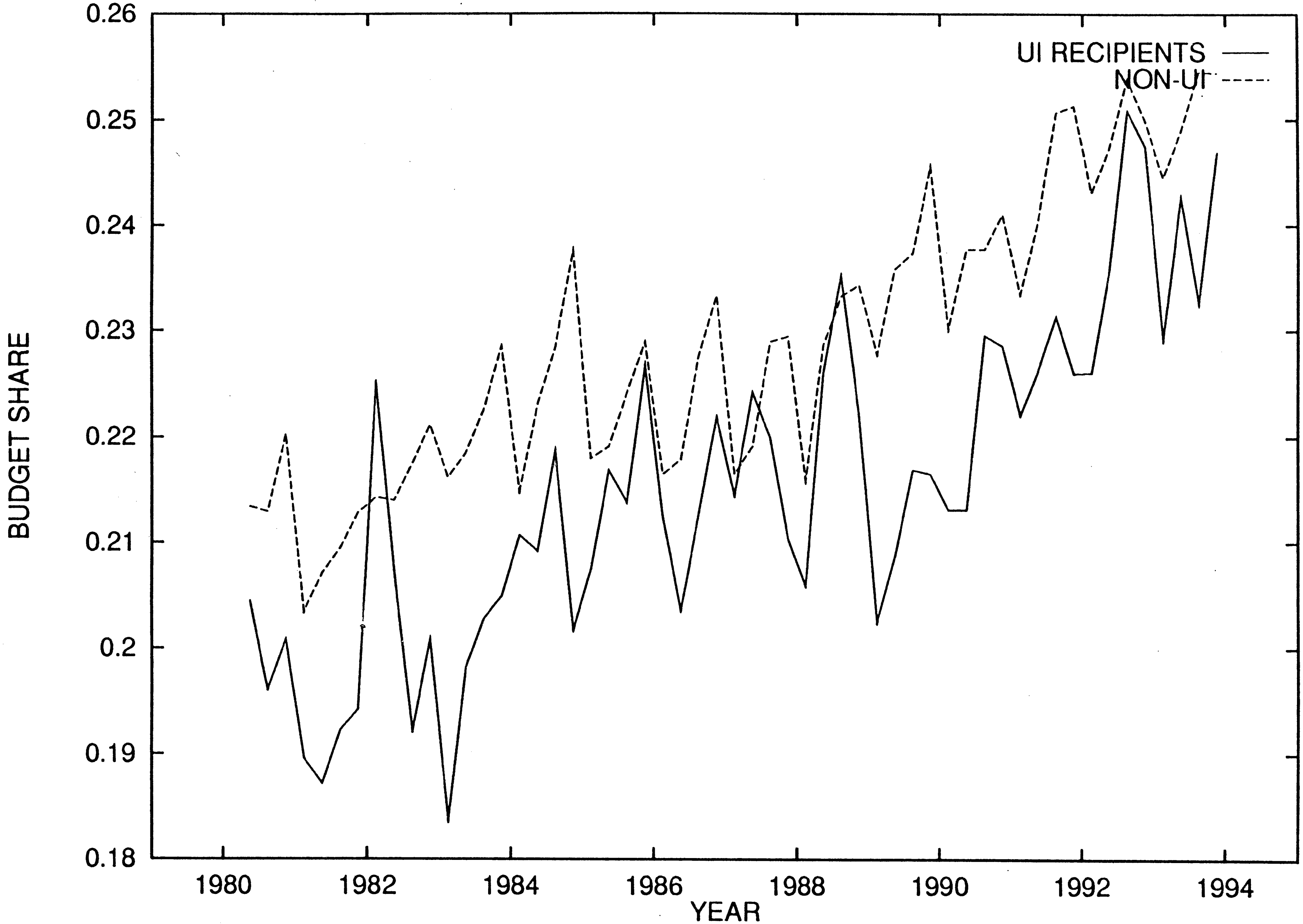


FIGURE 9- BUDGET SHARE OF CONSUMER SERVICES



for household size, age, region, race, sex, year, quarter and UI status. The results are presented in Table 4.

The central issue is the magnitude and sign of the coefficient on UI-recipient status. For every spending category except consumer goods that coefficient is statistically significant but very small. Relative to identical nonrecipients, households receiving UI benefits spend one percent more of their total budgets on energy and much less than one percent more on food and consumer goods. Recipients spend proportionately less on housing, durables and consumer services, although the magnitudes of the differences are well under one percent. The general conclusion is that the spending patterns of these two groups are very similar, especially after we account for differences in the demographic composition of the households.

While expenditure patterns do not differ substantially between UI recipients and others, it is possible that outlays on other items are influenced by whether or not the household has members who are unemployed. One might expect a priori that households receiving UI benefits are less likely to allocate expenditures toward what might be viewed as "postponable" items, such as purchases of durable goods, and less willing to provide gifts (either in-kind or in the form of cash) to individuals outside the household.[15] In Table 5 we present tobit estimates describing the variation in purchases of durables and gifts as determined by the logarithm of total expenditure and vectors of dummy variables representing household size, age of head, region of residence, race, sex, year, quarter and UI status.

The impact of UI status on both purchases of durables and the provision of gifts is negative and significantly different from zero, as expected. Once other demographic characteristics are accounted for, the level of total expenditure, a proxy for permanent income, has by far the largest

Table 4. Expenditure Share Equations, 1980:2 - 1993:4

	(1) Energy	(2) Food	(3) Consumer Goods	(4) Housing	(5) Durables Services	(6) Consumer Services
Log Expenditure	0.2114	-0.2976	-0.2298	0.4619	0.1847	-0.3306
	(0.0034)	(0.0057)	(0.0037)	(0.0055)	(0.0031)	(0.0064)
$(\text{Log Expenditure})^2$	-0.0122	0.0108	0.0125	-0.0224	-0.0085	0.0198
	(0.0002)	(0.0003)	(0.0002)	(0.0003)	(0.0002)	(0.0003)
UI Recipient	0.0109	0.0033	0.0006	-0.0071	-0.0040	-0.0038
	(0.0005)	(0.0009)	(0.0006)	(0.0009)	(0.0005)	(0.0010)
R^2	0.1664	0.2346	0.0743	0.1765	0.1794	0.0833

N = 195,689

Also included in the regessions were indicator variables for household size, age, race, sex, year, and quarter.

Table 5. Tobit Estimates of Impact of UI Receipts on Durable Purchase and Gifts

	(1) Durables	(2) Gifts
Log Expenditure	5519.39	1899.63
	(57.85)	(19.13)
UI Recipient	-498.57	-456.23
	(92.78)	(31.38)
χ^2	18840	21700

N = 195,689

Also included in the regressions were indicator variables for household size, age, race, sex, year, and quarter.

impact on the decision to purchase durables and a quite large effect on gift-giving. As with purchases of goods and services, these decisions do not differ much between the two populations.

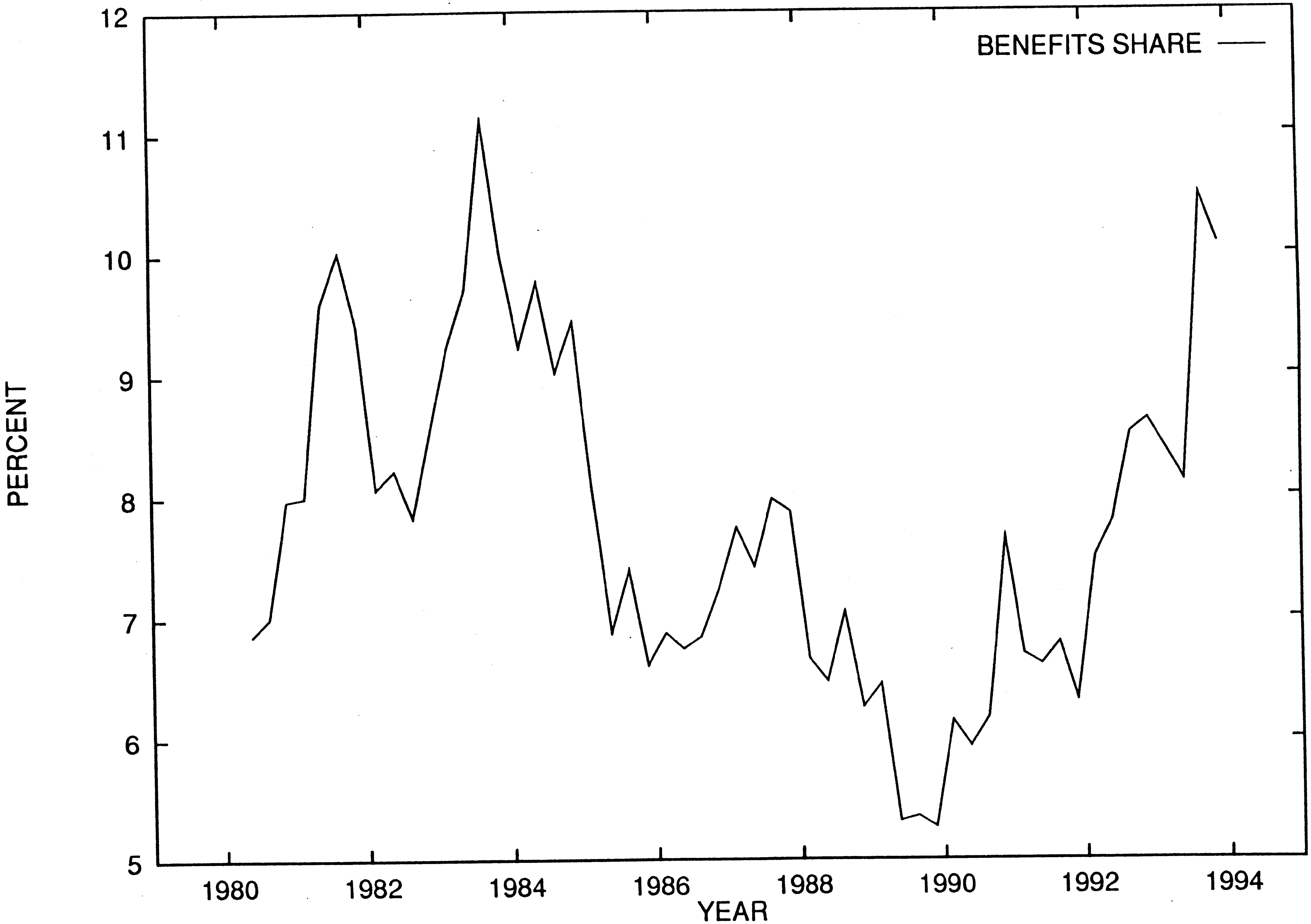
Taken together, this foray into analyzing disaggregated spending differences between UI recipient households and others corroborates the results of the previous section. Once we account for differences in their demographic characteristics and for the effect of family size and composition on well-being the two groups differ from each other much less than one might have expected. By inference, the level of UI benefits is sufficient to maintain recipients' spending patterns at what the behavior of otherwise identical households indicates they would have been if they had not experienced unemployment.

VI. UI Benefits and the Standard of Living

Thus far we have examined the welfare levels of recipients relative to those of nonrecipients, but we have not explicitly studied the role of UI benefit amounts in particular in maintaining the standard of living within the population of recipients. Average reported unemployment benefits for recipient households over the 55 quarters in our sample total $2,137, which constitutes 8.0 percent of the total annual income of these households. This fraction seems quite consistent with known facts: If average compensated duration is 13 weeks (one-fourth of a year), lost earnings constitute 50 percent of the household's flow of income, and gross replacement is 60 percent, we would observe that UI benefits add 7.5 percent to a receipient household's annual income.

There are substantial fluctuations in the importance of UI benefits over time. Figure 10 plots the average fraction of benefits in total income from 1980:2 through 1993:4. UI benefits are a more important source of income for recipient households in those quarters when the fraction

FIGURE 10- UI BENEFITS AS A PROPORTION OF RECIPIENT HOUSEHOLD INCOME



of households receiving benefits is highest. Over the two-year period 1982:2 through 1984:2 UI benefits averaged 9.3 percent of recipient households' incomes. From 1987:4 through 1989:4 the corresponding average was 6.3 percent. No doubt this is the result of the countercyclical pattern of the actual duration of benefit payments that unemployed workers receive.

To determine the impact of UI benefits on recipients' welfare we must assess their influence on recipients' consumption levels. This, in turn, depends fundamentally on the consistency of their behavior with the permanent-income hypothesis Rather than attempt to measure the extent to which consumers are able to borrow and lend to smooth consumption over time, we again bracket the magnitude of the effects of UI benefits by making two extreme assumptions:

> UI benefits would have the largest possible effect in maintaining expenditures if every recipient were completely constrained by liquidity, so that a change in current income would be fully reflected in an equal change in total expenditure. Among such individuals their well-being without the benefits is simply the observed level of total expenditure less the unemployment benefits received, deflated by the relevant price index and equivalence scale.
>
> UI benefits would have the smallest possible effect if every recipient were a "permanent-income consumer." In that case eliminating benefits would reduce recipients' lifetime wealth and therefore consumption by the annuitized value of this decrease.

We initially examine the effects of UI benefits under the extreme assumption that households are completely liquidity-constrained. Table 6 lists the two measures of household

Table 6. Ratios of Household Welfare, With / Without UI Benefits, UI Recipients Only

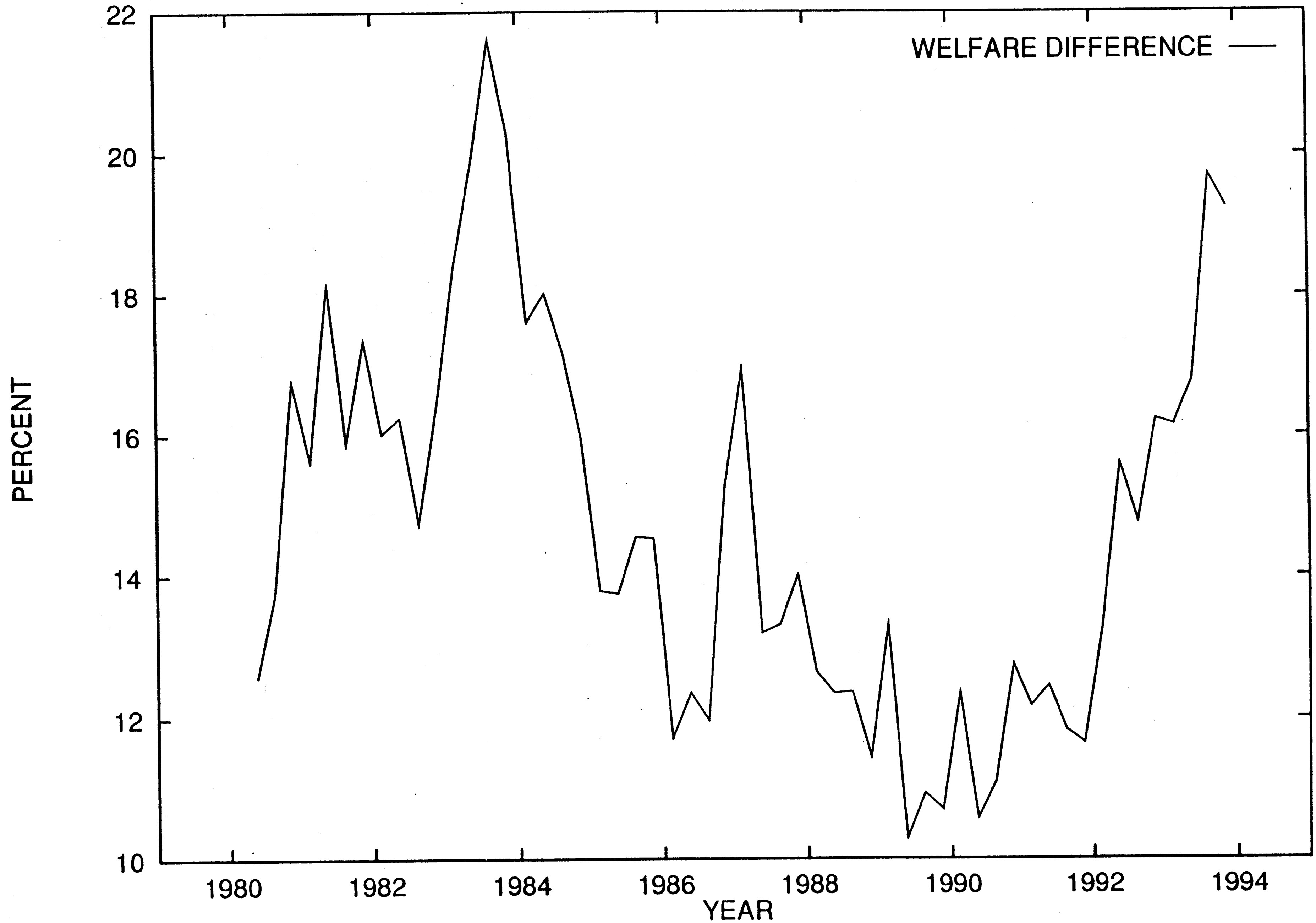
	(1) W1	(2) W2
Overall	1.12	1.12
Age of head		
16 - 24	1.14	1.14
25 - 34	1.13	1.13
35 - 44	1.12	1.12
45 - 54	1.11	1.10
55 - 64	1.12	1.11
Race of head		
White	1.12	1.11
Nonwhite	1.16	1.16
Gender of head		
Male	1.12	1.11
Female	1.14	1.13
Education of head		
No school	1.15	1.14
1 - 8 grade	1.14	1.14
9 - 11 grade	1.14	1.13
High school graduate	1.13	1.13
Some college	1.11	1.11
College graduate	1.11	1.10
Graduate school	1.09	1.09
Region of residence		
Northeast	1.14	1.13
Midwest	1.13	1.13
South	1.13	1.12
West	1.10	1.10
Earner Composition		
Head	1.15	1.15
Head & spouse	1.11	1.11
Head, spouse & others	1.07	1.07
Head & others	1.10	1.10
Spouse	1.17	1.17
Spouse & others	1.13	1.13
Others	1.17	1.17
No earners	1.44	1.44

welfare evaluated with and without benefits. For the entire UI sample over the 55 quarters the overall ratio of welfare with and without benefits is 1.12. Younger households gain more from receiving unemployment benefits relative to their older counterparts, however, as do nonwhite and female-headed households. The relative effects of UI are higher among less-educated households and those in the Northeast and the Midwest. These simple bivariate comparisons differ only slightly from comparisons based on multiple regressions that include all these determinants at once.

Has the impact of UI benefits on the welfare of recipients changed substantially over the sample period? Figure 11 shows the average proportional difference in the welfare of recipients with and without benefits during these nearly 14 years. The average increase in welfare resulting from the receipt of UI benefits is 14.7 percent. The gains are, however, positively related to the relative size of the program. Between 1982:2 and 1984:2, when the proportion of recipients was highest, the average welfare gain from benefits was 18.1 percent. Between 1987:4 and 1989:4, when only 4 percent of the sample received benefits, the average welfare gain among recipients was 12.0 percent.

All of these comparisons are based on the first polar case, that UI-recipient households are all constrained by liquidity. As such, the calculations present an upper bound on the proportional effect of UI benefits on the welfare of recipients relative to what it would have been in the absence of the program. A lower bound can be calculated in the second polar case by assuming that the typical recipient is no more likely to receive UI benefits in any particular year than the average labor-force participant (i.e., that UI recipiency is distributed randomly across the labor force). In this case UI benefits just raise permanent income and yield a flow of annual consumption equal

FIGURE 11- MAXIMUM PERCENTAGE IMPACT OF BENEFITS ON RECIPIENTS' WELFARE



to the amount of the benefits divided by the household's expected years of remaining lifetime. With benefits totalling only 8.0 percent of average income in recipient households; and with the average household head being only 39 years old, it is clear that, if households behave as permanent-income consumers, the fillip to consumption provided by UI benefits is very small. The lower-bound estimate of the welfare effects of UI benefits is nearly zero.

VII. Conclusion

This study has presented the first welfare-theoretic measures of the adequacy of UI benefits in the United States. The general result is very clear: Given the amount of benefits they are paid, households that receive benefits achieve nearly the same level of economic welfare as demographically identical households that do not receive benefits. By this economic criterion UI benefits are at least adequate to maintain the consumption of UI recipients. At the levels provided in the 1980s and early 1990s states' UI benefits and federal extended programs achieved the Act's original goal of "alleviating the hazards of ... unemployment."

This conclusion could be modified by additional analysis (which is not possible on any currently available set of data). On narrow grounds we may have overstated the welfare-improving effects of UI benefits by our inability to account for changes in the allocation of recipient households' labor between market and household production that is induced by their unemployment.[16] If, for example, one spouse's unemployment induces the other to enter the market, the welfare loss that seems to be overcome by the receipt of benefits is in fact not fully compensated. (Of course, the unemployed worker's leisure changes the welfare comparison in the opposite direction.) Also, we have not accounted for the effects of the burden of the UI tax on recipients and others. At first glance one would expect that accounting for the burden of the tax would strengthen our conclusion (since ipso facto recipients earn less than otherwise identical fully-employed workers); but the net effect of financing on the welfare comparisons that we have made would be extremely difficult to calculate.

On broader grounds our data do not allow us to determine whether spells of unemployment are being compensated at all, so that the group of nonrecipient households includes some households in which one member experiences a spell of unemployment during the twelve-month accounting period. This failure in the data makes UI benefits seem more adequate than they really are, since the income (and presumably consumption) of employed workers is understated by this classification problem. How important is this bias? If we assume (following Blank and Card, 1991) that fully two-thirds of unemployment spells are not compensated, and make the extreme assumption that the average income loss in those spells was the same as that lost by UI recipients, then the impact is still tiny. The average adjusted difference in welfare between UI recipients and households that experienced no unemployment during the year rises from the 7.7 percent difference (using the per-capita measure) that we noted in Section IV to no more than 8.2 percent.[17] Our central conclusion would thus hardly be changed if we could exclude unemployed nonrecipient households from the comparison group.

This central finding implies that benefit amounts and potential duration are adequate at their current levels. It says nothing about how well compensated spells of unemployment generally are. Indeed, the decline in the 1980s in the fraction of all unemployment that is compensated suggests that efforts to alter UI programs should be directed toward easing eligibility for benefits and ensuring that those who are eligible file for benefits in a timely fashion. Increasing benefits or potential duration for the relatively few who can currently obtain benefits would address a nonexistent problem: Assuming the goal of UI is to maintain well-being of recipients, the program already does that.

The results make it abundantly clear that programs that offered extended benefits during times of high unemployment, both temporary and triggered extensions, were of sufficient magnitude to prevent the welfare loss in recipient households relative to nonrecipient households from rising during the recessions of the early 1980s and of 1990-91. The constant legislative battles over these programs seem to have generated an outcome that was roughly consistent with a goal of leaving UI recipients relatively as well off during recessions as they are during other times in the business cycle.

REFERENCES

Orazio Attanasio and Guglielmo Weber, "Is Consumption Growth Consistent with Intertermporal Optimization? Evidence from the Consumer Expenditure Survey," National Bureau of Economic Research, Working Paper No. 4795, 1994.

Rebecca Blank and David Card, "Recent Trends in Insured and Uninsured Unemployment: Is There an Explanation?" Quarterly Journal of Economics, 106 (1991): 1157-89.

Steven Braithwait, "The Substitution Bias of the Laspeyres Price Index," American Economic Review, 70 (1980): 64-77.

Martin Browning, "Children and Household Economic Behavior," Journal of Economic Literature, 30 (1992): 1434-75.

----------------------, "UI, Income Distribution and Living Standards," Unpublished paper, McMaster University, 1995.

Paul Burgess and Jerry Kingston, "Benefit Adequacy Relative to Preunemployment Expenditure Levels," Proceedings of the Industrial Relations Research Association, 31 (1978): 39-45.

Laurits Christensen and Dale Jorgenson, "The Measurement of U.S. Real Capital Input, 1929-1967," Review of Income and Wealth, 15 (1969): 293-320.

Angus Deaton and John Muellbauer, Economics and Consumer Behavior. Cambridge: Cambridge University Press, 1980.

Mark Dynarski and Steven Sheffrin, "Consumption and Unemployment," Quarterly Journal of Economics, 102 (1987): 411-428.

Henry Felder and Jane-yu Li, "Adequacy of State Payments: An Assessment," in National Commission on Unemployment Compensation, Unemployment Compensation: Studies and Research, Volume 1. Washington: NCUC, 1980.

Jonathan Gruber, "The Consumption Smoothing Benefits of Unemployment Insurance," Working Paper No. 4750, National Bureau of Economic Research, 1994.

Daniel Hamermesh, "Social Insurance and Consumption: An Empirical Inquiry," American Economic Review, 72 (1982): 102-13.

Jerry Hausman, "Exact Consumer's Surplus and Deadweight Loss," American Economic Review, 71 (1981): 662-76.

Marilyn Manser and Richard McDonald, "An Analysis of Substitution Bias in Measuring Inflation, 1959-1985," Econometrica, 56 (1988): 909-30.

Christopher O'Leary, "The Adequacy of Unemployment Insurance Benefits," Unpublished paper, Advisory Council on Unemployment Compensation, 1994.

Daniel Slesnick, "Aggregate Consumption and Saving in the Postwar United States," Review of Economics and Statistics, 74 (1992): 585-97.

FOOTNOTES

1. This is an upper bound because it ignores any possible negative effects that the existence of public UI benefits might induce in households' precautionary saving.

2. Consumption also has a transitory component and therefore is itself a noisy measure of lifetime welfare. For the reasons described in the text, however, it is undoubtedly a more accurate measure than current income.

3. Arguments against the use of equivalence scales for welfare comparisons are summarized by Browning (1992).

4. As Attanasio and Weber (1994) show, basing studies on data that include only a few components of spending yields results on consumer behavior that are inconsistent with what is implied by complete measures of consumption.

5. To preserve consistency with earlier surveys, the head in consumer units with married couples is assumed to be the husband. Also, for ease of exposition we use the terms "household" and "consumer unit" interchangeably.

6. For 1980 and 1981 a hedonic regression is estimated using the 1984 survey and detailed information on the characteristics of the residence.

7. For details on how this method was applied to the CEX see Slesnick (1992).

8. The question posed is, "During the past 12 months, did you or any members of your consumer unit receive income from unemployment compensation? If so, how much was received from unemployment compensation?"

9. All tabulations of income exclude incomplete reporting of income and households with topcoded income.

10. Only 43UI-recipient households reported having had no schooling.

11. See Deaton and Muellbauer (1980) and, more recently, Browning (1992) for an exhaustive discussion of the issues involved in equivalence scales and their estimation.

12. The Census scales vary over dimensions other than household size. The equivalence scales in Figure 2 are for nonfarm households with a male head under age 65.

13. Annual and quarterly dummy variables are also incorporated in these regressions.

14. The absence of any cyclical variation in the difference in economic well-being between UI-recipient and other households is corroborated by respecifications of the regressions used to generate Figure 3 and Table 3. Adding the prime-age male unemployment rate for each quarter and an interaction of it with UI-recipient status yields t-statistics that are consistently well below 1 in absolute value.

15. The CEX includes information on both cash and in-kind gifts. Cash gifts are reported only in the household's fifth (final) appearance in the panel. Thus the totals in this category are likely to be underestimates.

16. O'Leary (1994) examines the effect of UI benefits on household well-being through their impact on each spouse's leisure only.

17. This means that 13.2 percent (twice the 6.6 percent that were recipient households) of all households (in the 93.4 percent of the sample that are nonrecipients) might have experienced a spell of unemployment.

With UI benefits accounting for 8.0 percent of UI recipients' incomes, and assuming sixty-percent replacement, we may have understated the incomes of households with no unemployment spells by 0.9 percent. Assuming, following Hamermesh (1982) that half of unemployed households are not liquidity-constrained, consumption-smoothing means that their expenditures are understated by perhaps only 0.5 percent. Given the shape of utility functions, their welfare is understated still less.

The Adequacy of Unemployment Insurance Benefits

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Prepared for:

Advisory Council on Unemployment Compensation
Employment and Training Administration
U.S. Department of Labor
200 Constitution Avenue, N.W.
Washington, DC 20210

Janet Norwood, Chair
Laurie Bassi, Executive Director

Prepared by:

W.E. Upjohn Institute for Employment Research
300 South Westnedge Avenue
Kalamazoo, Michigan 49007

Christopher J. O'Leary, Senior Economist

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1. INTRODUCTION

Since the inception of the federal-state unemployment insurance (UI) system nearly sixty years ago, there has been controversy about the adequacy of benefit payments.[1] Opinions have ranged from the view that UI does little more than subsidize leisure to the position that benefit levels grossly undercompensate for the physical and psychic hardships caused by unemployment.

The UI system was designed to be completely separate from relief programs, with eligibility determined by labor force attachment and benefit levels based on prior earnings experience. The benefit objectives of UI were recently set forth by the Advisory Council on Unemployment Compensation (1995, p. 8) in a statement of purpose for the UI system.

> The most important objective of the U.S. system of unemployment insurance is the provision of temporary, partial wage replacement as a matter of right to involuntarily unemployed individuals who have demonstrated a prior attachment to the labor force. This support should help meet the necessary expenses of these workers as they search for employment that takes advantage of their skills and experience.

In this statement the Advisory Council on Unemployment Compensation makes clear that the primary goal of UI is providing compensation for wage loss experienced as a result of involuntary unemployment. When making recommendations concerning benefit adequacy the Advisory Council on Unemployment Compensation (1995, p. 20) proposed:

> For eligible workers, each state should replace at least 50 percent of lost earnings over a six-month period, with a maximum weekly benefit amount equal to two-thirds of the state's average weekly wages.

---

[1]See Becker (1960) for an early history of public sentiment on several aspects of UI, and Curtin and Ponza (1980) for a summary of some more recent attitudes.

The Council's aim was to ensure one-half wage replacement *for a large number of beneficiaries*.

The most recent major effort to investigate the adequacy of UI was done in the 1970s by Paul Burgess and Jerry Kingston (1978a, 1978b) who conducted the Arizona Benefit Adequacy Study under the sponsorship of the U.S. Department of Labor. The methodology used by Burgess and Kingston closely paralleled that of earlier researchers.[2] The typical approach is to question a sample of UI recipients about their expenditures on a class of goods and services deemed "necessary" and compare the level of UI benefits to the level of these expenses.

Surveys of the type done by Burgess and Kingston, while extremely valuable, have proven to be quite expensive.[3] The high cost of gathering data has resulted in small sample sizes, but a more fundamental problem exists with the traditional approach. These studies presume that the analyst may determine which categories of expenditure are "necessary" or which items a household may least do without.

The problems of sample size and expenditure category selection, are both addressed in the present study by using a readily available large data set, the Current Population Survey (CPS) Annual Demographic File, and an agnostic approach to measuring unemployment compensation based on the economic theory of consumer-worker behavior. The methodology relies on a natural theoretical approach to estimating the upper limit on unemployment compensation--solve for the lump sum payment, which, when given to an unemployed individual, makes her indifferent between her current lot and her pre-unemployment one.

[2]Haber and Murray (1966) provide a summary of state studies done in the 1950s which used the same basic methodology later used by Blaustein and Mackin (1977) and Burgess and Kingston (1978a, 1978b).

[3] Becker (1961, p.23) noted that for the benefit adequacy studies done in the 1950s "[t]he time spent per interview averaged about three hours, with a range from one to fourteen hours, exclusive of the time spent in re-interviews of the more difficult cases."

Since UI is not intended to fully compensate the loss an individual experiences as a result of being unemployed, a financial inducement should remain for returning to work. Knowing the upper limit on the level of benefits is important for setting practical program guidelines.

In the next section a discussion of the accepted norms of benefit adequacy provides the framework for a review of the literature on assessing benefit adequacy. A simple theoretical approach to estimating the upper limit on unemployment compensation is given in Section 3 where explicit formulae for performing the computations are also given. In Section 4 the econometric methods to be used and the samples drawn from the 1992 CPS Annual Demographic File are discussed; basic labor supply results are also presented. Simulation results for a variety of household types, preference structures, and representative states are given in Section 6. The final section presents a summary of the new research findings, and considers program guidelines in light of the evidence presented.

2. STANDARDS OF BENEFIT ADEQUACY

In his classic monograph *The Adequacy of the Benefit Amount in Unemployment Insurance*, Father Joseph M. Becker (1961, p. 11) noted that; "A satisfactory norm of adequacy must have two elements--one positive, by which it can explain why benefits are as large as they are, and one negative, by which it can explain why they are no larger." Senator Paul Douglas (1932, p. 885) had earlier stated these principles in more substantive form. He suggested that "[t]here is a minimum of life which must be defended by the system of benefits," and that "[t]he amounts which the unemployed receive in benefits should always be appreciably less than what they would earn if employed [so that]...the temptation to shun work in order to draw the benefit will be greatly reduced" (Douglas 1932, p. 4). Douglas proposed that a balancing of these objectives might be achieved if unemployment benefits were to replace approximately one-half of lost wages for individuals who are unemployed and have demonstrated a significant attachment to the labor force.

While federal law has never specified the exact rate at which lost wages must be replaced under UI, every president since Eisenhower has reaffirmed the position that "payments to the great majority of the beneficiaries should equal at least half of regular earnings" (Becker 1980, p. 11). The Nixon administration specified that *great majority* should mean four-fifths of the nation's workforce (Becker 1980, p. 11). This criterion of benefit adequacy has come to be known as *one-half for four-fifths*.

2.1 The Wage Replacement Ratio: An Aggregate Criterion

While most states have benefit formula intended to replace approximately one-half of lost wages, the maximum on payments guarantees that many high wage workers will receive less than half their average lost earnings, and the minimum means that some low wage workers may receive more than half their average earnings. The data in Table 1 summarize the national historical experience on benefit adequacy using a very aggregate measure--the average wage replacement ratio (WRR). The national average WRR is defined by:

$$WRR = \frac{\sum_{i=1}^{n} WBA_i / n}{\sum_{j=1}^{m} WE_j / m}$$

where, WBA_i = the weekly benefit amount received by the ith UI recipient,

n = the number of UI recipients,

WE_j = the weekly earnings of the jth covered worker, and

m = the number of workers covered by UI.

In the first few years of UI, earnings of covered workers were unusually low, and the WRR was quite high. This is why there was little controversy about the adequacy of the weekly benefit amount until earnings rose rapidly after World War II. Figure 1 shows the declining trend of the WRR through the early 1950s. Since that time the WRR has ranged between thirty-two and thirty-seven percent, being approximately thirty-six percent in recent years.

Figure 1 which also shows a general upward trend in the WRR since about 1950. Controlling for the changing occupational mix of UI claimants, Hight (1980) arrived at lower bound estimates of 0.10 to 0.29 percent increase in the WRR per year over the period 1950-1977; and concluded that there has been some real gains in adequacy over the period. Table 2 lists the WRR for each state in 1994. While the national WRR was 36.05 percent in 1994, WRRs across the states ranged from a low of 26.8 percent in California to a high of 53.7 percent in Hawaii. A total of 18 states had WRRs greater than 40 percent in 1994.

Presumably the WRR is used as a rough gauge of benefit adequacy because the data needed to compute it is readily available. It is the main measure of benefit adequacy regularly reported by the U.S. Department of Labor.[4] However, the WRR as computed by the formula given above is a bit misleading. The denominator in the WRR considers wages for the entire population of covered workers, while the numerator considers only payments to

[4] It is reported by the U.S. Department of Labor quarterly in *UI Data Summary* and annually in updates to *UI Financial Data, ET Handbook No. 394.*

beneficiaries. Properly, we should examine benefit payments relative to lost earnings of beneficiaries.

Wayne Vroman (1980) who provided a comprehensive review of possible wage replacement rate computations called the series presented in Figure 1 and Tables 1 and 2 a *gross narrow wage replacement ratio* which is the one used historically. He also cited criticism that the measure underestimates the "true" replacement ratio because "unemployed workers receive lower wages than the average worker covered by the program."[5] Using unpublished micro data on the actual pre-unemployment earnings of beneficiaries from Illinois, Michigan, Pennsylvania, Texas, Washington, and Wisconsin for various periods during the 1980s, the Advisory Council on Unemployment Compensation (1995, p. 138) estimated that the gross narrow computation understates the true wage replacement rates by 25 to 30 percentage points.

The dramatic difference in wage replacement ratio estimates computed by the rather misleading gross narrow WRR formula and those produced using micro data on actual benefits and prior earnings convinced the Unemployment Compensation Advisory Council (1995, p. 21) to recommend that:

> The U.S. Department of Labor should calculate and report the actual replacement rate for individuals who receive Unemployment Insurance. This replacement rate should

[5] Vroman (1980, p. 170).

be calculated by dividing the weekly benefits paid to individuals by the average weekly earnings paid to those individuals prior to unemployment.

Vroman (1980, p. 170-72) reported that some researchers using micro data have arrived at very high net WRR figures. Feldstein (1974), who was concerned with the adverse incentive effects of UI, estimated that the net wage replacement ratio is often more than seventy percent. Munts and Garfinkel (1974) found replacement rates in Ohio in 1971-1972 to range from .38 to .89 for several distinct types of family units. Corson et al. (1977), determined the average ratio of benefits to lost wages in 1977 to be .66.

However, when broader measures of macro wage replacement which consider uncovered workers and non-compensated weeks are computed, replacement rates are much lower. For example Gramlich (1974), found that during the 1970-1971 recession for families headed by men, UI replaced only six to eight percent of lost earnings, and fourteen to eighteen percent for families headed by women. While the gross narrow WRR for 1971 was 0.363, Edgell and Wandner (1974) estimated the macro replacement rate for UI in the United States economy to be as low as 20 percent.

The wage replacement ratio estimates produced in the 1970s also varied because of differential treatment of taxes in the computations. This was a very important issue prior to the 1986 federal income tax changes which placed income received as unemployment compensation benefits in the same category for taxation as income from labor earnings.

2.2 Meeting Essential Expenditures: Support for the Standard

During the 1950s the U.S. Department of Labor financed a series of Unemployment Insurance benefit adequacy studies. The results of these studies have been summarized by Becker (1961), Lester (1962), and Haber and Murray (1966). Becker (1980), while discussing the principles which should underlie any proposal for a federal benefit standard, focused on the evidence from studies in Tampa, Fla. (1956), Anderson, S.C. (1957), Albany, N.Y. (1957), Portland, Ore. (1958), and St. Louis, Mo. (1958). These five similar studies were based on retrospective data on the income and expenditures of respondents during the period just prior to the survey date. Expenditures were divided into deferrable and non-deferrable categories. Spending on food, clothing, medical care, and housing constituted the non-deferrable group. Information was gathered on four household types. After examining these studies Becker (1980, p. 26) concluded that "[n]one of the states came close to the proposed goal of paying 80 percent of the beneficiaries half or more of their gross wage,...[and] [i]t is one of the weaknesses of the system that claimants without dependents' are treated much better than claimants with dependents." He suggested that benefit adequacy could be generally improved if benefit maximums were raised and programs for dependents allowances were expanded.

To give some examples from the 1950s studies, Table 3 presents a summary of the experience of those who fared best under the existing programs--households composed of a single beneficiary living alone. Becker (1961) found that benefits amounted to two-thirds or more of the income of unemployed single beneficiaries, more than 50 percent of family income for families with one wage earner, 40 percent for families with two wage earners and about 20 percent for families of secondary age earners. The 1950s studies demonstrated the usefulness of the one-half wage norm for assessing benefit adequacy. On the average, benefits that were half or more of the wage were sufficient to cover non-deferrable expenses for all claimant household types (Becker 1980, p. 13).

The deferrable/non-deferrable distinction used in the 1950s studies was expanded by Blaustein and Mackin (1977). They added expenditures made on a regular basis to repay outstanding debt to expenditures for food, clothing, medical care and housing, and labeled this "recurring" expenses. Using this concept as a basis for evaluating UI benefit adequacy they found that over two-thirds of the beneficiary households in South Carolina had adequate income in 1977. Nonetheless, they recommended increasing benefit maximums to improve adequacy.

Burgess and Kingston (1978a, 1978b) who conducted a detailed benefit adequacy study in Arizona, expanded the Blaustein-Mackin definition of recurring expenses to include expenditures on transportation, insurance, regular services, and regular support payments. They labeled this concept "necessary and obligated" expenses, and used it to assess benefit adequacy for seven recipient household types. The Arizona study revealed a wide disparity in terms of how closely benefits came to meeting the 10 necessary and obligated expenses for different categories of beneficiaries. As in the previous studies, the two most important factors, in addition to the weekly benefit amount, in determining the economic condition of the family during unemployment were the number of members to be supported and the number who were contributing to the support.

Burgess and Kingston found that benefits were most adequate for beneficiaries who had no other household members and lived with relatives--44 percent received a benefit equal to 100 percent or more of their share of the 10 expenses. The next most adequate category was husband and wife units in which both members worked. For 23.4 percent, the benefit amount represented 100 percent or more of expenses. Benefits were least adequate for beneficiaries in three or more person households in which the beneficiary was the only earner. For only 2.3 percent did the weekly benefit amount cover 100 percent or more of their expenses. For a majority of this category (56.1 percent), the benefit was half or less of the expenditures.

The low maximum weekly benefit amount was the principal reason for the disparity in the benefit-expense ratios among the different categories of Arizona beneficiaries studied. Sole wage earners, in households with two or more members including a spouse, generally had the highest wages and, consequently, were most often cut off by the maximum. For those beneficiaries, the weekly benefit amount--usually the $85 maximum--was less adequate than for any other category of beneficiary.

The Advisory Council on Unemployment Compensation (1995, p. 132) investigated whether states provided adequate UI benefits using data from the 1992 Consumer Expenditure Survey. Applying the narrow definition of necessary expenses used by Blaustein and Mackin (1977), the Advisory Council found that a majority of states provide UI compensation adequate to cover expenses for households with annual incomes in the $20,000 to $40,000 range. Very few state UI systems provided income replacement sufficient to meet the broader definition of Burgess and Kingston (1978).

Grossman (1973), Hamermesh (1982) and Gruber (1994a) have directly investigated how UI payments influence expenditure by unemployed workers. Grossman found that unemployed persons substitute leisure for market goods in an attempt to maintain customary consumption levels. Hamermesh concluded that UI benefits only partly help smooth consumption during periods of lost earnings due to unemployment, and that as much as half of the benefits received are spent as if "individuals were fully able to borrow or had sufficient savings to meet transitory losses of income without any disruption in their consumption spending."[6] Gruber estimated that in the absence of UI, average consumption expenditure by unemployed persons would fall by 22%, or more than three times the decline estimated in the presence of UI.[7]

[6] Hamermesh (1982, p. 110).

[7] Gruber (1994, p. 30).

The consumer expenditure studies also raise a question about the importance of UI in maintaining necessary expenditure. Gruber (1994b) investigated this question using two sources of microeconomic consumer expenditure data. Results based on both the Panel Study of Income Dynamics (PSID) and the Consumer Expenditure Survey (CES) "suggest that UI has a significant effect on consumption of the unemployed."[8] On a finer point evidence from the two data sets differed. The PSID results indicated UI is indispensable for unemployed workers trying to maintain necessary expenditure, while the CES suggested that other forms of consumption insurance, such as savings and earnings of other household members, are at least as important as UI benefits.

2.3 Optimal Unemployment Insurance: A Theoretical Approach

Baily (1978) and Flemming (1978) originated theoretical models of optimal unemployment insurance. The models are similar in that both attempt to solve for characteristics of the UI system which would maximize the expected lifetime utility of a representative worker. The UI program choice parameters for this problem are the wage replacement rate, and the potential duration of benefits. Both Baily and Flemming assume an infinite potential duration of benefits, and each determines that optimal replacement rates are in the range of those provided by the states. Baily (1978, p. 393) finds that:

[8] Gruber (1995, p. 31).

> [that if the] degree of relative risk aversion by workers [is] unity, and if workers do not prolong their duration of unemployment very much as a result of UI payments [i.e., if the elasticity of a spell of unemployment with respect to a change in the benefit amount is about 0.15] then if the benefit-wage ratio is 50% it is about right.

The elasticity of unemployment with respect to the benefit amount assumed by Baily (1978) is in line with estimates summarized in Chapter 7.

Flemming qualifies his statements with capital market considerations. He concludes that under perfect capital markets a replacement rate of 50% is too high, and "[i]f there is no lending or borrowing the optimal rates rise to about 75%."[9]

Davidson and Woodbury (1995, p. 1) examine optimal UI with "an equilibrium search and matching model calibrated using data from the reemployment bonus experiments and secondary sources." Like Baily and Flemming they find that if potential UI duration were infinite replacement rates should optimally be 50%. However, Davidson and Woodbury also estimate that if potential duration is limited to the standard 26 weeks, then the UI system should optimally replace all of lost earnings.

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[9] Flemming (1978, p. 403).

2.4 Econometric Approaches: Applying Theory

Burgess and Kingston (1980) investigated the possibility of evaluating benefit adequacy on the basis of readily available survey (Continuous Wage and Benefit History--CWBH) and claims data. They conclude, however, "that information on income and household composition must be supplemented with actual or estimated data on household expenditure patterns to predict individual benefit adequacy values with a reasonable degree of accuracy" (U.S. Department of Labor 1981, p. 43). Other writers have presented results which suggest a greater potential for econometric methods to yield reasonable estimates of adequate UI compensation.

Ashenfelter (1980), in the context of a household model where unemployment is treated as a rationing constraint, developed an approximation to a quantity which he refers to (Ashenfelter 1980, p.552) as the "lump-sum compensation required to restore the unemployed [rationed] worker's family to the welfare level of the fully employed family." This approximation is arrived at by taking a second-order Taylor Series approximation of the difference between the exogenous cost of achieving the unconstrained utility level in the presence of the ration and the cost of achieving the same level in the absence of any constraint, around the fully employed point. The result is "a conventional Harberger (1971) type triangle measure of welfare loss" (Ashenfelter 1980, p. 553), which is applied to aggregate time series data.

Hurd (1980), who examined the cost of unemployment to the unemployed, used a hybrid of approximation and direct methods to examine the experience of respondents to the 1967 Survey of Economic Opportunity. He estimated the parameters of a Taylor Series approximation of the substitution effect of a wage change on hours of work, integrated to find the compensated labor supply function, solved for the utility constant wage acceptance locus by inversion, and then determined the required lump sum compensation to constrained individuals by evaluating the area under this locus between the actual (constrained) and fully employed levels of labor supply.

O'Leary (1990) estimated the lump-sum compensation required to restore a single unemployed person with no dependents to the welfare level of a fully employed worker using a second-order Taylor Series approximation.[10] Results presented in Ashenfelter (1980), Hurd (1980), and O'Leary (1990) all suggest that the current UI practice of replacing one-half lost wages tends to overcompensate short spells of unemployment and undercompensate long spells.

2.5 A Consensus Standard

The norm of adequacy *one-half for four-fifths* is rooted in the common-sense recommendations of economists and politicians made over fifty years ago. The norm has

[10]This paper draws heavily on arguments and results presented in O'Leary (1986, 1990). Those previous studies of UI benefit adequacy focused on single workers without dependents.

been demonstrated to correspond roughly to the fundamental concern of satisfying needs of the unemployed, as well as being consistent with the fiscal integrity of the program. It is also appealing to policy makers and program managers because it is easy to apply. In the final section this norm is reviewed in light of full unemployment compensation estimates. The theoretical foundation for this exercise is laid in the next section.

3. A THEORETICAL APPROACH TO FULL UNEMPLOYMENT COMPENSATION

If the notion of a representative structure for individual preferences can be accepted, a method for directly evaluating the required compensation to an individual constrained in selling labor services is immediate. The method developed here is based on just such an assumption and is in the spirit of work by Rosen (1978), who examined the excess burden of income taxation, and Hurd and Pencavel (1981), who evaluated various wage subsidy programs.

3.1 Consumer Behavior with Employment Constraints

Satisfaction of each consumer-worker is represented as depending simply on the market resources at her command, Y, and the time available to enjoy these resources, L. It is assumed that each individual, given her exogenous non-labor income, I, and the rate at which she can transform labor services, H, into income, w, if unconstrained in the labor market, acts in a manner consistent with the problem:

$$\max_{L,Y} \{ U(L,Y); Y = wH + I \} = V(L(w,I), Y(w,I)). \tag{1}$$

She reaches an optimum where H(w,I) hours of work are supplied to the market and Y(w,I) goods are consumed in her residual discretionary time, L(w,I). Denoting T as the endowment of discretionary time, L(w,I) = T - H(w,I). In (1) V(w,I) is the indirect utility function, it represents the maximum level of satisfaction for given values of w and I.

While the above exposition is stated in terms of individual behavior, the question of appropriate unemployment compensation is best seen from the household perspective. To provide this viewpoint we follow the usual approach found in the economics literature as summarized by Shelly Lundberg (1988, p. 225):

> The most common empirical specification of family labor supply treats the work hours of married men as independent of the behavior or attributes of their wives and the husband's behavior, in turn, as exogenous with respect to the wife's work decision. Husband and wife maximize utility independently, with the wife treating husband's earnings as property income. This results in an asymmetric pair of labor supply functions [as stated below] with no cross equation restrictions.

Using the above notation and denoting subscripts for married males and females as m and f respectively, the following are labor supply equations for a household with a married couple where both partners work in the market: $H_m = H_m(w_m, I)$ and $H_f = H_f(w_f, I + w_m H_m)$. Under these assumptions the analysis of unemployment compensation to individuals in a household context may proceed using the model for individual consumer-worker behavior.

In many instances, the effective choice facing a consumer-worker is between working a standard day, week, or year or not working at all; in other cases an optimal wage-hour arrangement way be upset by an unexpected layoff. The analytic techniques required to investigate the effects of labor market constraints on consumer-worker behavior are formally similar to the methods used to evaluate the response to "straight rationing." Research on the effects of rationing began during World War II (see Rothbarth 1940-41, and Kaldor 1941) and has continued since (see Tobin-Houthakker 1950-51, Pollack 1971, and Neary and Roberts 1980).

Ashenfelter (1980) developed a model of household labor supply under rationing. This model has been applied by Blundell and Walker (1982), Deaton and Muellbauer (1981), Kneisner (1976), Parsons (1977), and Ransom (1987). Ham (1982) presented results based on a model of individual labor supply under rationing.

An individual faced with a binding constraint on the hours that he may sell in the labor market at, say, $\underline{H} < H(w,I) = T - L(w,I)$, achieves a utility level less than that attainable in the absence of the labor market constraint,

$$U(T - \underline{H}, w\underline{H} + I) < U(L(w, I), Y(w, I)), \tag{2}$$

or in terms of the indirect utility function,

$$\underline{V}(\underline{H}, w\underline{H} + I) < V(w, I). \tag{3}$$

Full unemployment compensation to an individual who is constrained in selling labor services is that lump sum grant, c, which solves:

$$U(T - \underline{H}, w\underline{H} + I + c) = U(L(w,I), Y(w, I)). \quad (4)$$

Stating this condition in terms of the indirect utility function,

$$\underline{V}(\underline{H}, w\underline{H} + I + c) = V(w, I), \quad (5)$$

where c is a Hicksian equivalent variation. It is the lump sum compensation required by an individual who is constrained in the labor market to make him as well off as if he were employed at equilibrium hours without any change in relative prices. Therefore

$$c = c(w, I, \underline{H}) \quad (6)$$

is the compensation he would need to forego an opportunity to be employed at equilibrium hours.

The concept of full compensation embodied in this approach may be easily understood by referring to the indifference curve analysis of Figure 2. An unconstrained individual, with preferences as represented by the map of indifference curves on Figure 2, would reach an unconstrained optimum equilibrium on U^0 at (L^0,Y^0). If, for some reason, market opportunities allow sales of only $\underline{H} = T - L^1$ hours of labor services, a lower level of utility is reached on U^1 at (L^1,Y^1). While there is a hardship experienced as a result of the associated earnings loss $(Y^0 - Y^1)$, the utility loss is partly compensated by an increase in leisure, and the income required to fully compensate the constrained individual $(\underline{Y} - Y^1)$ is less than the earnings loss.

3.1 Explicit Formulae for Computing Full Compensation

The approach to measuring full compensation proceeds from the estimation of a representative labor supply function. To compute an exact solution for full compensation utility function parameter estimates are required. For the model presented above, when the theoretical conditions required by neoclassical economic theory are satisfied utility function parameters can be recovered from estimation of a labor supply specification.

Deriving an explicit closed form solution to (6) is not always an easy matter. Two utility functions are used in this study, they are the familiar Stone-Geary (SG) which has been used widely in employment policy research, and the somewhat less familiar utility function derived by Hausman (1980) from the linear labor supply function. To crystalize the approach, the Stone-Geary case is now worked out in detail.

3.1.1 Full Compensation when Utility is Stone-Geary

The Linear Expenditure System is derived from the Stone-Geary utility function:

$$U(L,Y) = \alpha \ln(L - \gamma_1 - \Delta) + (1 - \alpha)\ln(Y - \gamma_2);\ \ 0 < \alpha < 1, \tag{7}$$

where the parameters α and $(1 - \alpha)$ are interpreted as marginal budget shares devoted to leisure and market goods, and γ_1 and γ_2 represent leisure and income origin translation parameters respectively, and $\Delta = \delta D$ with D the number of dependents and δ the effect of each dependent on the origin where leisure is defined. Maximizing (7) subject to the income, $Y = wH + I$, and time, $T = H + L$ constraints yields leisure demand,

$$L = \gamma_1 + (\alpha/w)((wT + I) - w(\gamma_1 + \Delta) - \gamma_2), \tag{8a}$$

or labor supply,

$$H = (T - \gamma_1) - (\alpha/w)(I + w(T - \gamma_1 - \Delta) - \gamma_2), \tag{8b}$$

and commodity demand,

$$Y = \gamma_2 + (1 - \alpha)((wT + I) - w(\gamma_1 + \Delta) - \gamma_2), \tag{8c}$$

functions. Given the adding up condition on neoclassical demand functions, the parameters of (8a) through (8c) can be determined by estimating the parameters of any one of the demand system equations. Denoting the estimated parameter values by the parameters themselves, substitution of (8a) and (8c) into the right-hand side of (4) yields the right-hand side of equation (5),

$$\left[T - \left[\frac{(1-\alpha)w(T-\gamma_1-\Delta)-\alpha(I-\gamma_2)}{w}\right]-\gamma_2\right] \times \left[(1-\alpha)\{w(T-\gamma_1-\Delta) + (I-\gamma_2)\}\right]^{1-\alpha} \tag{9}$$

the indirect Stone-Geary utility function. For this case the left-hand side of (5) is:

$$(T-\underline{H}-\gamma_1-\Delta)^{\alpha}\ (w\underline{H}+I+\tilde{c}-\gamma_2)^{(1-\alpha)} \tag{10}$$

Equating (9) and (10) and solving for $\tilde{c}$ yields:

$$\tilde{c} = \gamma_2-I-w\underline{H}+(1-\alpha)\left\{w(T-\gamma_1-\Delta)+(I-\gamma_2)\right\} \times \left[\frac{\alpha\{w(T-\gamma_1-\Delta)+(I-\gamma_2)\}}{w(T-\underline{H}-\gamma_1)}\right]^{\frac{\alpha}{(1-\alpha)}} \tag{11}$$

a closed form solution for full unemployment compensation when utility is Stone-Geary.

3.1.2 Full Compensation when Labor Supply is Linear

Hausman (1980) has shown that when labor supply takes the following linear form:

$$H = \alpha w + \delta I + Z\gamma, \tag{12}$$

with the variables H, w and I as defined above, Z representing socioeconomic variables such as the number of dependents and α, δ and γ being parameters to estimate, the indirect utility function satisfying neoclassical conditions is:

$$V(w, I) = e^{\delta w}\{I + (\alpha/\delta)w - (\alpha/\delta^2) + (s/\delta)\}, \tag{13}$$

where, $s = Z\gamma$, and the direct utility function consistent with the linear labor supply is:

$$U(L, Y) = e^{-[1 + \delta(Y + \check{s})(b - H)]}\{(H - b)/\delta\}, \tag{14}$$

where $b = \alpha/\delta$ and $\check{s} = (s/\delta) - (\alpha/\delta^2)$. In a different paper Hausman (1981) showed how these specifications may be used to compute exact welfare measures at the individual level. In the present case full compensation when labor supply is constrained to be $\underline{H} < H(w, I)$ is the Hicksian equivalent variation, c, which may be computed by the following formula:

$$c = \frac{\{\delta w + [1 + \delta(w\underline{H} + I + \check{s})(b - \underline{H})]\} + \ln\{(\delta I + bw\delta - b + s)/(\underline{H} - b)\}}{(-\delta(b - \underline{H}))} \tag{15}$$

4. SAMPLES, METHODS, AND BASIC ESTIMATION RESULTS

4.1 The Samples

The basic estimation was performed on samples from the 1992 Current Population Survey (CPS) Annual Demographic File. These data were collected in March of 1992, and describe respondent behavior during 1991.

This study ultimately examines what full UI compensation might be for workers in twelve different household situations. Six different categories of household member were examined in households with and without dependents. A total of 33,454 households were used for the basic estimations. This included:

11,739 households with married couples where *both* partners worked,
6,153 households with married couples where only the husband worked,
2,505 households with married couples where only the wife worked,
6,031 households with a single male working person, and
7,026 households with a single female working person.

Parameters of the preference structure were estimated for one person in each of the last four household types listed above, and for both partners among married couples where both worked. This results in workers who were in six different categories of household membership.

To arrive at this sample for analysis we eliminated households with earners aged less than 25 or more than 55 years, and examined only workers with positive earnings sometime during 1991. We also excluded households with more than two earners. Among households without a married couple we examined only those where there was one earner.

One of the most interesting aspects of household structure for our purposes is the dependents relationships. There was an average of 1.4 dependents in households with married couples where both partners worked and married couples with only the husband working. In households with a married couple and only the wife working the average was 0.9 dependents, while there was an average 0.2 dependents for single males and 0.7 dependents for single females. In addition to information on dependents, the mean values of annual hours worked, hourly wages, age, education, race, and urban residence status are presented in Table 4. Not surprisingly the samples show that workers in households with married couples and only one worker average about ten years older than households with married couples where both partners work, also the sample containing the largest fraction of black households are those where there is a single woman working.

The family non-labor income figure of $34,953, for wives in households where married spouses both work reflects the assumption that Shelly Lundberg (1988, p. 225) says is "the most common empirical specification"--labor income of husbands is regarded by working wives as part of exogenous income. The relative size of the means to the standard deviations of family non-labor income for the sub-samples indicates that for some households non-labor income is negative. This is because the CPS household non-labor income variable includes self-employment income and rental income, each of which may reasonably be negative in a given year.

Given the great diversity of American households, the sample selection restrictions are admittedly severe. However, even within the five narrowly defined household types examined there are many different dependent relationships so that the categories of household member multiply quickly. Assigning dependency relationships and non-labor income becomes quite complicated for other household structures. Not every possible combination can be examined; information yielded from examination of the household categories selected is rich and varied.

4.2 Estimation Methods

The parameter estimates which serve as the basis for compensation simulations are reported in Table 5. The equations estimated are similar in that each has a very small coefficient of determination. This is typical when estimating labor supply equations on cross-section data. While several omitted factors obviously explain the total variation in annual hours worked, every individual parameter in these equations is estimated with a high degree of statistical significance. Furthermore, these estimates are quite robust, being relatively invariant when other regressors were included. The parsimonious specifications were chosen for simplicity.

The labor supply specifications (8b) and (12) were each estimated on the six different samples of workers described above. The labor supply equations were estimated using ordinary least squares, correcting for the *division bias* problem involved in defining the hourly wage rate using the method proposed by Borjas (1980). In the labor supply regression equations the dependent variable, annual hours, is definitionally related to the important predictor, the hourly wage rate, since the latter is defined by dividing the former into annual earnings. To avoid the bias in parameter estimates which may result from division bias, first stage wage equations are run. Results of these estimations are reported in Table 6. All parameters in the wage equations were estimated with great precision, and overall the equations fit the data quite well. Wages were modeled as depending on age, education, race and urban residency status. These predictor variables were not later included in the hours equations so as to satisfy identification of the system and to avoid multicollinearity with the predicted wage.

All results in this study are based on empirical labor supply equations, which include only variables suggested by the theory which in this case includes the number of dependents. The number of dependents was incorporated into the two utility functions examined since dependency status is an important consideration in estimating UI benefit adequacy.

4.3 Basic Estimation Results

The direct utility function (14) derived by Hausman (1980) suggests no natural interpretations of the parameters estimated for the linear labor supply function and presented in Table 5. Interpretation of these results is limited to discussion of elasticities. On the other hand there are natural interpretations of the parameters of the Stone-Geary labor supply function reported in Table 7.

The budget share devoted to leisure is greatest for married males and single female workers. The complementary group of married women and single men, who have relatively lower valued market uses of time, have relatively higher minimum leisure requirements. Estimated minimum income requirements are large and negative for all groups. As mentioned earlier negative values are possible because the exogenous household income variable includes losses from self employment and rental property. The relative magnitudes of the estimated γ_2 across household types are reasonable. Working married males in one earner households have the highest subsistence income requirements, while married women in dual earner households have the lowest requirement.[11]

[11]Following the usual practice in the literature (Lundberg, 1988, p. 225) of adding husbands earnings to working wives non-labor income, as seen in Table 4, married women in dual earner households also have the greatest mean and standard deviation in exogenous non-labor income. This obviously drives the γ_2 estimate.

The labor supply equation estimates presented in Table 5 indicate that dependents increase hours of labor supplied by men, and decrease hours offered to the market by women. These results are given a finer interpretation in Table 7 where estimates for δ, the Stone-Geary utility function parameter indicating the minimum leisure required per dependent, are reported. The estimates indicate that an additional dependent reduces the minimum leisure required by a working woman with a working spouse by 119 hours per year, while increasing the minimum leisure required by married men whose spouse does not work by 123 hours per year.

Estimates of the structural Cournot (uncompensated) wage effect, income effect, substitution effect, and associated elasticities are presented in Table 8 for the Stone-Geary form and Table 9 for the Linear form. The labor supply estimation results are most easily reviewed in elasticity terms. For both the Stone-Geary and the Linear specifications, the elasticity estimates are consistent with the implication of consumer demand theory that the substitution effect on labor supply is positive. Furthermore, in each case leisure is found to be a normal good. Generally speaking, the Stone-Geary specification yields results more consistent with the received literature. The Linear form results in a relatively high labor supply elasticity for married men who are the sole earner in the household, this group is usually found to have the least elastic labor supply. Using the Stone-Geary specification, the labor supply elasticity estimate of married male sole earners falls to less than half that from the linear specification. Other estimates generated from the Stone-Geary model are also more in line with previous studies.

5. ESTIMATES OF FULL UNEMPLOYMENT COMPENSATION

In this section full compensation estimates based on the formulae (11) and (15) given in Section 3 are presented for various hypothetical degrees of labor market constraint. These figures are reported together with UI payment simulation results for four states having benefit computation provisions which span the variety of systems extant and an estimate of compensation which would result if one-half of lost wages were replaced--which is the standard norm of adequacy.

Under all state Unemployment Insurance (UI) laws, a claimant's benefit rights depend on four principal factors: "the amount of employment and wages required to qualify an individual for benefits, the period for earning such wages, the method of computing the weekly benefit amount, and the method of determining the length of time for which benefits may be paid."[12] Another factor which is an important determinant of benefits in 14 states is dependents allowances. While the level of wages and period of employment for qualification differ greatly across the states, there exist only four basic schemes for determining a UI claimant's weekly benefit amount. They are referred to as the Average-weekly-wage, High-quarter, Multi-quarter, and Annual-wage formulae.

Results of simple simulations, performed under the assumption of qualification for the maximum benefit payment period, are presented for state programs representative of each of the four benefit schemes: Michigan provisions are used to perform Average-weekly-wage simulations, Massachusetts laws provide the parameters to do High-quarter simulations, Illinois serves as an example of a Multi-quarter state, and Oregon's scheme is used to generate Annual-wage simulations. The particulars of the four categories of benefit rights provisions in each of these states are summarized in Table 10. The third section of the table highlights the distinguishing characteristics of the four different state benefit schemes. Under each scheme a formula is employed which yields a weekly benefit amount (WBA) which is equal to about one-half of lost gross wages. Under the Michigan plan seventy percent of the net AWW is paid; in Massachusetts a fraction between 1/21 and 1/26 of the HQ earnings is the WBA[13]; in Illinois 49 percent of earnings in the two highest quarters in the base period divided by 26; and in Oregon the WBA is 1.25 percent of annual income.

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[12] *Comparison of State UI Laws*, U.S. Department of Labor (1992, p. 3-1).

[13] The fraction 1/26 is used in the Massachusetts simulations because the statutory alternative of 1/52 of the highest two quarters yields the same WBA in our simulations since we use average quarterly earnings computed as annual earnings divided by four.

Tables 11-16 present simulation results for the four states, the one-half wage replacement rule, and for the two preference structures considered. Each table is divided into two parts, the left hand panel gives results for workers with no dependents, the right hand panel gives results for workers with two dependents. In each table the left most column lists the hypothetical number of weeks of unemployment (Weeks), which is allowed to range from one to thirty-one because among the simulation states the maximum entitled duration of regular benefits is 30 weeks in Massachusetts which also has a one week waiting period. The next four columns report the cumulative benefit payments which would be made to a qualified claimant in Michigan, Massachusetts, Illinois, and Oregon with the various weeks of unemployment, and sub-sample average gross hourly wages for the six categories of worker reported in Table 4. Column six reports a dollar amount which equals half of the total gross wages lost by a worker with the mean wage rate, and mean non-labor income. The seventh and eighth columns present the amount of "full" compensation implied by the closed form direct compensation formula for the Stone-Geary and Linear specifications respectively. The right panel in each table presents similar simulation results with the change that the hypothetical worker has two dependents instead of none.

In Michigan there is no waiting period before benefit payments begin. However, in Massachusetts, Illinois and Oregon the benefit payment is zero during the first full week of unemployment, with this waiting period acting as a form of coinsurance. The one-week waiting period was required in all but eleven states in 1991.[14] In all states, once benefit payments commence, total benefits increase in a linear fashion, with a fixed benefit amount being paid each week, until there is either a return to work or the claimant is no longer eligible. The one-half wage replacement rule results in a fixed benefit payment each week as well.

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[14]The other ten states without a waiting week in 1991 were: Alabama, Connecticut, Delaware, Iowa, Kentucky, Maryland, Nevada, New Hampshire, Virginia, and Wisconsin.

It is assumed in the simulations performed here that the stylized claimant considered qualifies for the maximum benefit period. In the absence of economic conditions which trigger extended benefits, 26 weeks is the maximum benefit duration under most UI programs.[15] As a consequence of the waiting period and the benefit maximums, the figures in the simulation tables for Illinois and Oregon are constant for weeks of unemployment beyond twenty-seven, for Michigan there is no change after 26 weeks because Michigan has a maximum entitlement of 26 weeks and no waiting week, cumulative compensation reaches a maximum in Massachusetts after 31 weeks. Just as the UI benefit totals increase in a linear fashion, so do the totals for one-half gross wage replacement (HALF).

In the simulations the generally accepted norm of benefit adequacy--*one-half wage replacement*--is met or slightly exceeded in all four states for workers with relatively low earnings. That is for the three categories of woman worker. The mean hourly wages across the three groups of woman worker were all approximately equal to $10.50, while the mean hourly wages for men were somewhat higher. The mean wages for both categories of married men, single earner ($16.47) and dual earner ($14.89) households, were too high to allow the average worker to qualify for half wage replacement in any of the states. However, single males who had mean hourly earnings of $13.24 would be provided with approximately half wage replacement when unemployed in either Michigan or Massachusetts. Naturally, in the simulations the waiting week delays wage replacement in Massachusetts, Illinois, and Oregon, but not in Michigan.

Differentiating each compensation formula with respect to hours, H, reveals that it is in general impossible to determine *a priori* how a change in hours of work affects utility based compensation. Comparing simulation results for "full" compensation from the theoretical formulae based on Stone-Geary and integrated Linear utility, with the figures for the actual benefit payments which would be forthcoming in the various states, the

[15]The exceptions (maximum duration in weeks) are: Louisiana (28), Massachusetts (30), Pennsylvania (30), Puerto Rico (20), Utah (36), Virginia (28), Washington, D.C. (34), and West Virginia (28).

general result is that current UI programs appear to overcompensate for wage loss during the first several weeks of unemployment and undercompensate for lengthy spells of unemployment.

The Stone-Geary form yields full compensation simulation estimates which nearly coincide with the one-half wage replacement rule for long durations of unemployment, but suggests that the states and the half-wage replacement formula is too generous in early weeks of unemployment.

Results based on the Linear form of labor supply generally accentuate the tendencies of the Stone-Geary simulation suggesting that compensation should be lower than the accepted norm in early weeks. However, for long durations of unemployment the Linear form suggests that compensation may safely be much higher than one-half wage replacement.

For a few categories of worker, simulation results based on the theoretical formulae have a surprising non-monotonic shape. For working husbands with non-working wives the pattern is exhibited for both the Stone-Geary and the Linear based formulae for men both with and without other dependents. For the Stone-Geary form the pattern is also apparent for single men with two dependents, and for the Linear form the pattern appears for married women workers with a working spouse and two dependents. In all of these cases the pattern is generally the same--full compensation in the first week of unemployment should be positive though not large, with cumulative full compensation declining for additional weeks of unemployment until it reaches zero in the early weeks of a spell and then rises thereafter. These results occur because of the non-linear form of the compensation formulae and the relative magnitude of the parameter estimates. The estimates suggest that the timing of benefit payments should be closely examined. Ignoring possible *entry effects* which may be created, the results suggest that the waiting period might be placed after the first weeks of compensation.

It is surprising that results on dependents allowances from the theoretical compensation formulae are not more consistent given that the dependents variable in the labor supply equations yielded the usual results found in the literature--independent of the household structure, because of strong income effects for men dependents tend to increase hours of market work for males, and perhaps because they more significantly raise the opportunity cost of working for women; dependents decrease hours of market work for females. For the Stone-Geary form adding dependents to the household lowers required full compensation for men and raises full compensation required for women, while precisely the opposite occurs for the Linear form with dependents lowering full compensation to women workers and raising full compensation.

Naturally, the conflicting simulation results across functional forms for dependents is due to the differing treatment of demographic variables in the compensation formulae. The result highlights the extreme sensitivity of the simulation results to the specifications. Taken together, the simulation results based on the theoretical specifications tend to be in the neighborhood of the standard norm of one-half wage replacement which is approximately what states provide for beneficiaries qualifying for less than the maximum weekly benefit amount. Results based on the Stone-Geary are slightly below and those based on the Linear form are somewhat above half wage replacement. Rather than contradict the standard norm of adequacy, these results tend to support the one-half wage replacement rule. If the theoretical simulation results raise any questions, they are about the best timing of payments.

6. SUMMARY, CONCLUSIONS, AND POLICY IMPLICATIONS.

Results from estimating explicit parameterizations of labor supply have been used to compute estimates for full unemployment compensation. The estimates generated were compared to hypothetical payments which would accrue under the unemployment insurance (UI) systems of representative states. Results on compensation amounts tend to support the accepted standard of UI benefit adequacy which calls for replacement of one-half of lost wages. While one-half wage replacement over the course of an average 15 week spell of unemployment appears to yield adequate and not excessive wage replacement, return to work incentives might be improved if the fixed nature of the weekly payment is examined. There may be ways to maintain or improve benefit adequacy while speeding return to work. This might be accomplished in part by a closer examination of partial benefit rules.

The direct compensation and state program simulations imply that current UI programs overcompensate for wage loss during short spells of unemployment, and under-compensate for lengthy spells. Overall, compensation is adequate in the present UI system, but the timing of payments should be more closely examined. Particular program features to consider are the length and timing of the waiting period.

Findings in this study concerning dependents allowances were extremely cloudy. The two different theoretical specifications produced opposite results. What the results suggested was that dependents affect required compensation to men and women in exactly opposite ways regardless of the household setting where the man or woman lives. It may require Solomon to craft a benefit policy which treats men and women differently in terms of dependents, and is still politically acceptable.

For the 12 different types of representative worker considered in this study, benefit simulations were performed for four representative states: Michigan, Massachusetts, Illinois, and Oregon. Among the 48 cases examined at least one-half of lost weekly earnings would be replaced during a week of unemployment in 24 of the cases. Clearly, each of the 48 cases is not equally likely to occur in practice. The four states studied differ greatly in size, and the probabilities of unemployment for each of the twelve types of household member differ as well. In the simulations one-half wage replacement is most likely to occur for women and single men, with dependents allowances greatly increasing the chance of one-half wage replacement. In 1993 single Americans were more than twice as likely to experience unemployment than were married people, and among women those with dependents were more likely to be unemployed.[16] This suggests that unemployment is a greater risk for those more likely to be adequately compensated by the UI system.

Since the 1950s a popular standard of unemployment insurance benefit adequacy is half wage replacement for eighty percent of the insured unemployed or *one-half for four-fifths*. Given that between the minimum and maximum weekly benefit amounts approximately one-half of lost wages are replaced, an important part of benefit adequacy concerns maximum benefit amount policy. Obviously raising the maximum weekly benefit amount (WBA) would allow one-half wage replacement to extend to more beneficiaries. In Table 17 we see that for the six worker types drawn from the 1992 Current Population Survey (CPS) and examined in this paper,[17] setting the maximum WBA at two-thirds the full sample average weekly wage (AWW) would extend one-half wage replacement to 77 percent of the population. The maximum WBA would need to be at about 71 percent of the AWW to allow *one-half for four-fifths*. Among working married women with husbands not working, the standard of adequacy would be reached with the maximum WBA at fifty percent of the AWW, while for working married men with wives not working setting the maximum at seventy-five percent of the AWW would still fall short of the adequacy standard. Clearly, earnings levels are different for the various categories of earners. Table 18 states what maximum WBA combined with fifty percent wage replacement below the maximum would yield *one-half for four-fifths* for each of the six categories of worker considered in this study.

Raising the maximum WBA is not a simple matter, adjustments of this parameter should always be considered in the larger context of UI trust fund adequacy. As Vroman (1990, p. 114) points out "symmetric treatment...of taxes and benefits...helps to reduce the risk of insolvency." It is generally believed that if the maximum weekly benefit amount is

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[16] Data from Table 1 in the 1993 *Geographic Profile of Employment and Unemployment* and Table 626 in the 1994 *Statistical Abstract of the United States.*

[17] There are six different worker types when the two alternative dependents possibilities are ignored.

set at two-thirds of the state average weekly wage, one-half wage replacement will be achieved for eighty percent of beneficiaries. Regarding maximum benefit amount policy Minnesota and Oklahoma should be studied as models. In Oklahoma, for example, the maximum weekly benefit amount is adjusted annually to a percentage between 60 and 67 percent of the state average weekly wage depending on the state UI trust fund balance.

While it was mentioned in this paper when reviewing earlier research but not analyzed, benefit adequacy also concerns those with low levels of prior earnings. Because "necessary and obligated" expenditures amount to a larger share of earnings for low income people, one-half wage replacement may be inadequate for this group. Programs which tie the minimum weekly benefit amount (WBA) to the maximum WBA amount should be closely examined. Kansas, where the minimum WBA is set at 25 percent of the maximum WBA, offers a useful approach to minimum WBA policy.

REFERENCES

Ashenfelter, O. (1980), "Unemployment as Disequilibrium in a Model of Aggregate Labor Supply", *Econometrica*, Vol. 48 (April): 547-64.

Baily, M.N. (1978), "Some Aspects of Optimal Unemployment Insurance," *Journal of Public Economics*, Vol. 10 (December): 379-402.

Becker, J.M. (1960), "Twenty-Five Years of Unemployment Insurance," *Political Science Quarterly*, Vol. 75 (December): 481-99.

Becker, J.M. (1961), *The Adequacy of the Benefit Amount in Unemployment Insurance*, W.E. Upjohn Institute for Employment Research, Kalamazoo, Michigan.

Becker, J.M. (1980), *Unemployment Benefits: Should There Be a Compulsory Federal Standard?*, American Enterprise Institute for Public Policy Research, Washington.

Blaustein, S.J. (1993), *Unemployment Insurance in the United States: The First Half Century*, Kalamazoo: W.E. Upjohn Institute for Employment Research.

Blaustein, S.J, and P.A. Mackin (1977), *Job Loss, Family Living Standards, and the Adequacy of Weekly Unemployment Benefits*, Unemployment Insurance Service, U.S. Department of Labor.

Blundell, R.W., and I. Walker (1982), "Modelling the Joint Determination of Household Labor Supplies and Commodity Demands," *Economic Journal*, Vol. 92 (April).

Borjas, George J. (1980), "The Role of Division Bias," *The Journal of Human Resources*, Vol. 15, No. 3 (Summer): 409-23.

Burgess, P., J. Kingston and C. Walters (1978a) *The Adequacy of Unemployment Insurance Benefits: An Analysis of Weekly Benefits Relative to Preunemployment Expenditure Levels*, U.S. Department of Labor, ETA.

Burgess, P., J. Kingston and C. Walters (1978b), *The Adequacy of Unemployment Insurance Benefits: An Analysis of Adjustments Undertaken Through Thirteen and Twenty-five Weeks of Unemployment*, U.S. Department of Labor, ETA.

Burgess, P., J. Kingston, R. St. Louis, and J. Sloane (1980), *Benefit Adequacy and UI Program Costs: Simulations with Alternative Weekly Benefit Formulas*, U.S. Department of Labor, Employment and Training Administration.

Burtless, G. (1990), "Unemployment Insurance and Labor Supply: A Survey," in W. Lee Hansen and James F. Byers, eds., *Unemployment Insurance: The Second Half-Century*, The University of Wisconsin Press, Madison, Wisconsin.

Classen, K.P. (1977), "The Effect of Unemployment Insurance on the Duration of Unemployment and Subsequent Earnings," *Industrial and Labor Relations Review*, Vol. 30 (July): 438-44.

Corson, W.S., et al. (1977), "A Study of Recipients of Federal Supplemental Benefits and Special Unemployment Assistance," Mathematica Policy Research, Inc., Princeton.

Crosslin, R.L. and W.W. Ross (1980), "Achieving Wage Replacement Goals," in *Unemployment Compensation: Studies and Research, Volume 1*, National Commission on Unemployment Compensation: 65-76.

Deaton, A., and J. Muellbauer (1980), *Economics and Consumer Behavior*, Cambridge University Press, Cambridge, England.

Douglas, P.H. (1932), *Standards of Unemployment Insurance*, The University of Chicago Press, Chicago, Illinois.

Edgell, David and Stephen A. Wandner (1974), "Unemployment Insurance: Its Economic Performance, " *Monthly Labor Review*, Volume 97, Number 4 (April): 33-40.

Ehrenberg, R.A. and R.L. Oaxaca (1976), "Unemployment Insurance, Duration of Unemployment, and Subsequent Wage Gain," *American Economic Review*, Vol. 66 (December): 754-66.

Feldstein, M.S. (1974), "Unemployment Compensation: Adverse Incentives and Distributional Anomalies," *National Tax Journal*, Vol. 27 (June): 231-44.

Flemming, J.S. (1978), "Aspects of Optimal Unemployment Insurance Search, Leisure, Savings and Capital Market Imperfections," *Journal of Public Economics* 10 (December): 503-425.

Gramlich, E.M. (1974), "The Distributional Effects of Higher Unemployment," *Brookings Papers on Economic Activity* 2: 293-341.

Grossman, Michael (1973), "Unemployment and Consumption," *American Economic Review*, Volume 63 (May): 208-213.

Gruber, Jonathan (1994a), "The Consumption Smoothing Benefits of Unemployment Insurance," paper prepared for the *Advisory Council on Unemployment Compensation* U.S. Department of Labor (August).

Gruber, Jonathan (1994b), "Unemployment Insurance, Consumption Smoothing, and Private Insurance: Evidence form the PSID and CEX" paper prepared for the *Advisory Council on Unemployment Compensation* U.S. Department of Labor (November).

Haber, W., and M.G. Murray (1966), *Unemployment Insurance in the American Economy: An Historical Review and Analysis*, Richard D. Irwin, Inc., Homewood, Illinois.

Ham, J.C. (1982), "Estimation of a Labor Supply Model with Censoring Due to Unemployment and Underemployment," *Review of Economic Studies*, Vol. 49 (July): 333-54.

Hamermesh, Daniel S. (1982), "Social Insurance and Consumption: An Empirical Inquiry," *American Economic Review*, Volume 72, Number 1 (March): 101-13.

Harberger, A.C. (1971), "Three Basic Postulates for Applied Welfare Economics: An Interpretive Essay," *Journal of Economic Literature* 9 (September): 785-97.

Hausman, J.A. (1981), "Exact Consumer's Surplus and Deadweight Loss," *American Economic Review*, Number 4 (71): 662-76.

Hausman, J.A. (1980), "The effect of wages, taxes, and fixed costs on women's labor force participation," *Journal of Public Economics*, Number 2 (14): 161-94.

Heckman, J.J. (1974), "Shadow Prices, Market Wages and Labor Supply," *Econometrica* Vol. 42 (March): 679-94.

Heckman, J.J. (1976), "The Common Structure of Statistical Models of Truncation, Sample Selection and Limited Dependent Variables and a Simple Estimator for Such Models," *Annals of Economic and Social Measurement*, Vol. 5 (Fall): 475-92.

Heckman, J.J. (1979), "Sample Selection Bias as a Specification Error," *Econometrica*, Vol. 47 (January): 153-61.

Height, J.E. (1980), "Trends in Unemployment Insurance Wage Replacement, 1950 to 1977," in *Unemployment Compensation: Studies and Research, Volume 1*, National Commission of Unemployment Compensation.

Hurd, M. (1980), "A Compensation Measure of the Cost of Unemployment to the Unemployed," *Quarterly Journal of Economics* 95 (September): 225-43.

Hurd, M., and J.H. Pencavel (1981), "A Utility Based Analysis of the Wage Subsidy Program," *Journal of Public Economics* 15 (January): 185-201.

Kaldor, N. (1941), "Rationing and the Cost of Living Index," *The Review of Economic Studies* 8 (June), 185-87.

Keeley, M.C. (1981), *Labor Supply and Public Policy,* Academic Press.

Killingsworth, M.R. (1982), *Labor Supply*, Cambridge University Press, London.

Kniesner, T.J. (1976), "An Indirect Test of Complementarity in a Family Labor Supply Model," *Econometrica* Vol. 44 (July): 651-69.

Lester, R.A. (1962), "The Economics of Unemployment Compensation," Industrial Relations Section, Princeton University.

Lopez, Eduard A. (1987), "Constitutional Background to the Social Security Act of 1935," *Social Security Bulletin*, Vol. 50, No. 1: 5-11.

Lundburg, Shelly (1988), "Labor Supply of Husbands and Wives: A Simultaneous Equations Approach," *Review of Economics and Statistics*, 70, 2 (May): 224-35.

Munts, R. and I. Garfinkel (1974), "The Work Disincentive Effects of Unemployment Insurance (September), Upjohn Institute, Kalamazoo, MI.

National Commission on Unemployment Compensation (1980), *Unemployment Compensation: Final Report*, Washington, DC: U.S. Government Printing Office.

Neary, J.P., and K.W.S. Roberts (1980), "The Theory of Household Behavior Under Rationing," *European Economic Review* 13 (January), 25-42.

O'Leary, C.J. (1990), *An Econometric Analysis of Unemployment Insurance Benefit Adequacy*, Staff Working Paper 90-05, Kalamazoo: W.E. Upjohn Institute.

O'Leary, C.J. (1986), *An Econometric Analysis of Unemployment Insurance Benefit Adequacy*, Ann Arbor: University Microforms.

Papier, W. (1974), "Standards for Improving Maximum Unemployment Insurance Benefits," *Industrial and Labor Relations Review* 27, No. 3 (April): 376-90.

Parsons, D.O. (1977), "Health, Family Structure and Labor Supply," *American Economic Review* 67 (September): 703-12.

Pollak, R.A. (1969), "Conditional Demand Functions and Consumption Theory," *The Quarterly Journal of Economics* 83, 60-78.

Ransom, M.R. (1987), "An Empirical Model of Discrete and Continuous Choice in Family Labor Supply," *Review of Economics and Statistics* 69 (August): 465-72.

Rosen, H.S. (1978), "The Measurement of Excess Burden with Explicit Utility Functions," *Journal of Political Economy* 54 86, no. 2, part 2 (April): 5121-35.

Rothbarth, E. (1940-41), "The Measurement of Changes in Real Income Under Conditions of Rationing," *The Review of Economic Studies* 8 (Winter): 100-107.

Spiegelman, Robert G., Christopher J. O'Leary, and Kenneth J. Kline (1992), *The Washington Reemployment Bonus Experiment: Final Report*, Unemployment Insurance Occasional Paper 92-6, Washington, DC: U.S. Department of Labor.

Tobin, J. and H. Houthakker, (1950-1951), "The Effects of Rationing on Demand Elasticities," *The Review of Economic Studies*, 18, 140-153.

U.S. Department of Commerce (1994), *Statistical Abstract of the United States*, Bureau of the Census, Economics and Statistics Administration.

U.S. Department of Labor (1992), Employment and Training Administration, *Unemployment Insurance Financial Data*, ET Handbook 394, Washington, D.C.

U.S. Department of Labor (1992), *Comparison of State Unemployment Insurance Laws*, Employment and Training Administration, Unemployment Insurance Service.

U.S. Department of Labor (1993), *Geographic Profile of Employment and Unemployment*, Bureau of Labor Statistics, Bulletin 2446.

Vroman, Wayne (1980), "State Replacement Rates in 1980," in *Unemployment Compensation: Studies and Research, Volume 1,* compiled by the National Commission on Unemployment Compensation, Washington, DC: U.S. Government Printing Office.

Vroman, W. (1990), *Unemployment Insurance Trust Fund Adequacy in the 1990s*, W. E. Upjohn Institute for Employment Research, Kalamazoo, MI.

Figure 1: Aggregate Average Wage Replacement Ratio (WRR) in the United States, 1938-1994.

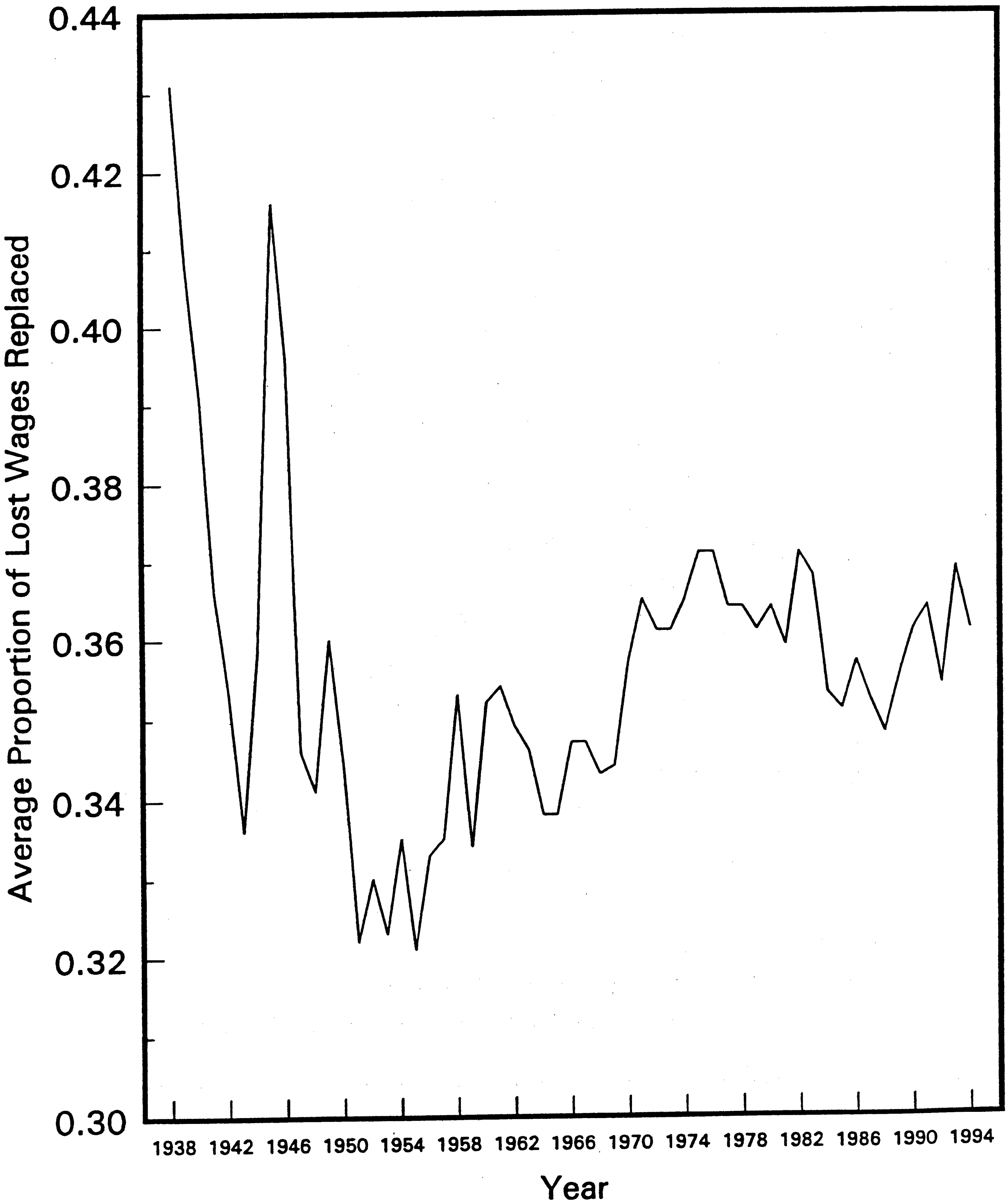

Figure 2. A graphic representation of full unemployment compensation as a Hicksian equivalent variation.

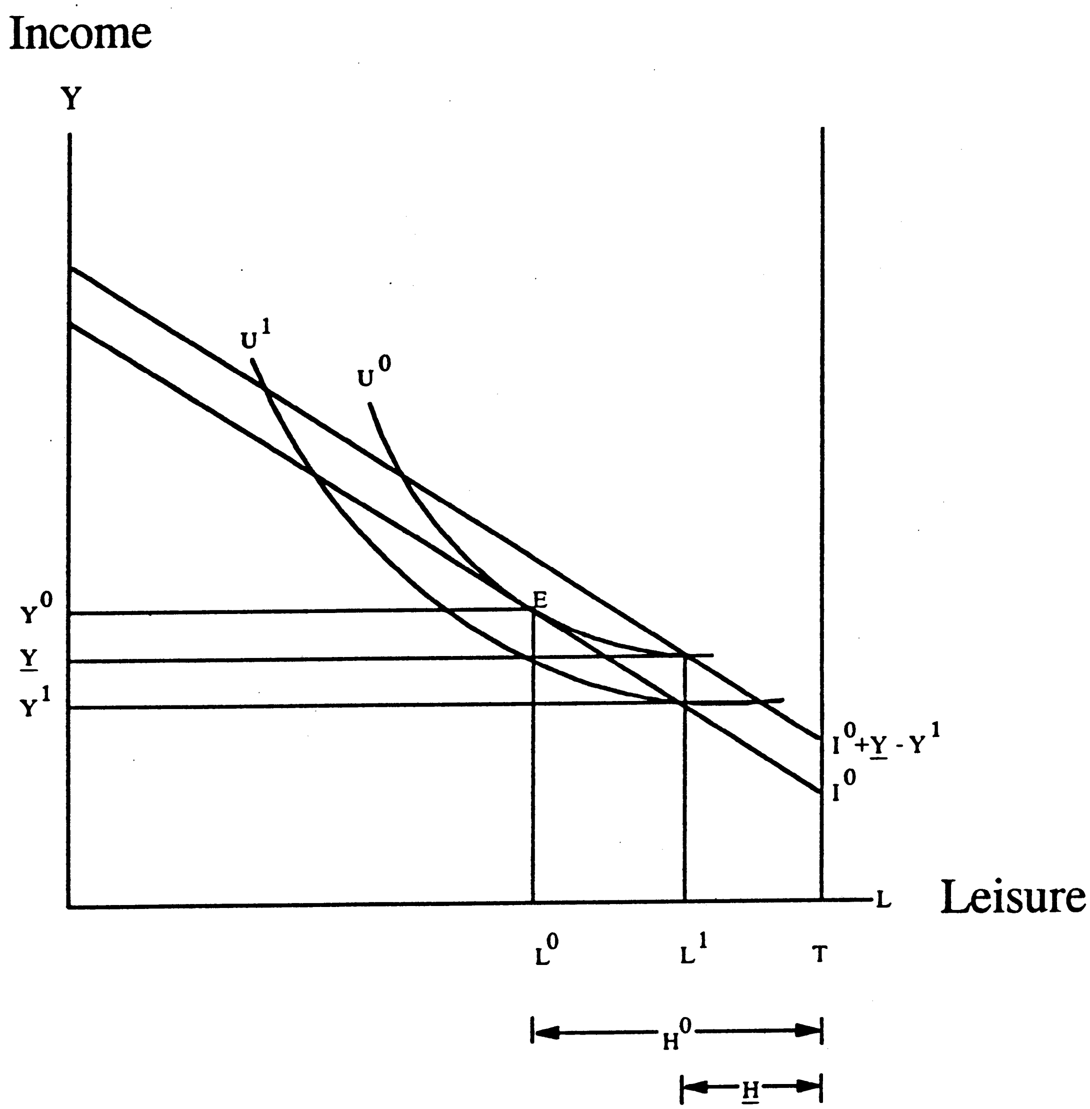

Table 1

Average UI Weekly Benefit Amount (WBA), and Wage Replacement Ratio (WRR) in the United States, 1938-1993.[a]

Year	WBA	WRR	Year	WBA	WRR
1938	10.94	0.431	1966	39.76	0.347
1939	10.66	0.408	1967	41.25	0.347
1940	10.56	0.391	1968	43.43	0.343
1941	11.06	0.366	1969	46.17	0.344
1942	12.66	0.353	1970	50.31	0.357
1943	13.84	0.336	1971	54.35	0.365
1944	15.90	0.359	1972	55.82	0.361
1945	18.77	0.416	1973	59.00	0.361
1946	18.50	0.396	1974	64.25	0.365
1947	17.83	0.346	1975	70.23	0.371
1948	19.03	0.341	1976	75.16	0.371
1949	20.48	0.360	1977	78.71	0.364
1950	20.76	0.344	1978	83.67	0.364
1951	21.09	0.322	1979	89.68	0.361
1952	22.79	0.330	1980	98.95	0.364
1953	23.58	0.323	1981	106.61	0.359
1954	24.93	0.335	1982	119.34	0.371
1955	25.04	0.321	1983	123.59	0.368
1956	27.02	0.333	1984	123.47	0.353
1957	28.17	0.335	1985	128.23	0.351
1958	30.54	0.353	1986	135.72	0.357
1959	30.40	0.334	1987	139.74	0.352
1960	32.87	0.352	1988	144.91	0.348
1961	33.80	0.354	1989	151.76	0.355
1962	34.56	0.349	1990	161.56	0.361
1963	35.28	0.346	1991	169.88	0.364
1964	35.96	0.338	1992	173.64	0.354
1965	37.19	0.338	1993	179.69	0.369
			1994	181.53	0.361

[a] Source: UI Financial Data, ET Handbook No. 394, United States Department of Labor, Employment and Training Administration (1992). Figures for 1993 and 1994 averaged from the four quarterly issues of *UI Data Summary*, United States Department of Labor, Employment and Training Administration (1993, 1994).

Table 2

State Wage Replacement Ratio (WRR), 1994
State Maximum Weekly Benefit Amount (MWBA), Jan 1993
as a Fraction of State Average Weekly Wage (AWW), 1992
and Any Statutory Rule for MWBA as a Fraction of AWW

State	WRR	MWBA	MWBA/AWW	Statutory Rule
Alabama	0.312	165	0.394	
Alaska	0.278	212	0.370	
Arizona	0.333	185	0.423	
Arkansas	0.423	212	0.564	66 2/3
California	0.268	230	0.421	
Colorado	0.405	250	0.526	55
Connecticut	0.330	306	0.487	60
Delaware	0.339	245	0.474	
District of Columbia	0.316	335	0.500	11
Florida	0.372	250	0.576	
Georgia	0.33[illegible]	185	0.393	
Hawaii	0.537	322	0.685	70
Idaho	0.414	223	0.573	60
Illinois	0.360	227	0.423	49.5
Indiana	0.360	140	0.310	
Iowa	0.447	200	0.505	53
Kansas	0.435	239	0.575	60
Kentucky	0.379	217	0.525	55
Louisiana	0.274	181	0.418	66 2/3
Maine	0.384	198	0.490	52
Maryland	0.345	223	0.445	
Massachusetts	0.410	312	0.548	57.5
Michigan	0.379	293	0.554	58
Minnesota	0.441	279	0.578	50-60%
Mississippi	0.343	165	0.453	
Missouri	0.329	175	0.390	
Montana	0.413	209	0.579	60
Nebraska	0.363	154	0.404	
Nevada	0.383	217	0.472	50
New Hampshire	0.312	188	0.395	
New Jersey	0.393	325	0.526	56 2/3
New Mexico	0.372	191	0.495	50
New York	0.321	300	0.472	
North Carolina	0.419	267	0.637	66 2/3
North Dakota	0.435	212	0.596	60
Ohio	0.389	228	0.486	
Oklahoma	0.407	229	0.559	60-66 2/3
Oregon	0.396	271	0.615	64
Pennsylvania	0.419	317	0.651	66 2/3
Puerto Rico	0.320	133	0.485	50
Rhode Island	0.465	294	0.653	67
South Carolina	0.370	191	0.474	66 2/3
South Dakota	0.397	154	0.467	50
Tennessee	0.334	170	0.393	
Texas	0.378	245	0.504	
Utah	0.438	240	0.584	60
Vermont	0.370	199	0.469	
Virginia	0.360	208	0.447	
Virgin Islands	0.484	203	0.494	50
Washington	0.412	273	0.572	70
West Virginia	0.393	270	0.643	66 2/3
Wisconsin	0.417	240	0.553	
Wyoming	0.416	200	0.500	55

Table 3

Experience of Single UI Beneficiaries
Selected from Five Benefit Adequacy Surveys, 1956-1958.[a]

Survey	PCTGW[b]	PCTNW[c]	PCTND[d]	PCTMAX[e]	WRR[f]	SWRR[g]
Tampa	28	65	95	21	.46	.31
Anderson	51	84	118	37	.56	.36
Albany	51	72	114	46	.54	.34
Portland	52	79	118	42	.58	.39
St. Louis	34	58	106	49	.48	.33

[a]Source: Becker (1980), Table 1, pp. 11-12.

[b]PCTGW: Percent of beneficiaries whose benefits were half or more of their gross wage.

[c]PCTNW: Percent of beneficiaries whose benefits were half or more of their net wage.

[d]PCTND: Average benefit as a percent of average non-deferrable expenditures.

[e]PCTMAX: Percent of beneficiaries who received the maximum benefit amount.

[f]WRR: Ratio of average weekly benefit amount in state to average weekly net wage of recipients.

[g]SWRR: Ratio of average weekly benefit amount in state to average weekly wage in state covered employment.

Table 4
Means of Characteristics of the Samples Selected from the
1992 Current Population Survey Annual Demographic File by Household Type
(standard deviations in parentheses)

	Household Type					
	Married Both Working		Married One Working		Single	
Characteristics	Husbands	Wives	Husbands	Wives	Males	Females
Annual Hours Worked	2,165 (589)	1,650 (705)	2,053 (731)	1,560 (755)	2,015 (670)	1,887 (653)
Hourly Wage	14.89 (9.12)	10.64 (10.07)	16.47 (18.67)	10.50 (12.25)	13.24 (28.65)	10.51 (6.87)
Family Non-labor Income	2,867 (7,225)	34,953 (21,730)	6,361 (13,992)	10,297 (15,881)	1,862 (7,372)	2,394 (5,521)
Number of Dependents	1.4 (1.2)	1.4 (1.2)	1.4 (1.4)	0.9 (1.1)	0.2 (0.6)	0.7 (1.0)
Age in Years	38.8 (7.7)	36.8 (7.3)	44.4 (13.6)	46.7 (12.8)	36.0 (8.1)	37.1 (8.2)
Education (Proportion in Category)						
8 years or less	0.03 (0.17)	0.02 (0.15)	0.09 (0.29)	0.05 (0.22)	0.05 (0.22)	0.03 (0.18)
9 to 12 years	0.07 (0.25)	0.06 (0.23)	0.11 (0.31)	0.10 (0.30)	0.08 (0.27)	0.08 (0.27)
High School grad	0.34 (0.47)	0.37 (0.48)	0.34 (0.47)	0.41 (0.49)	0.33 (0.47)	0.33 (0.47)
Some college	0.19 (0.39)	0.19 (0.39)	0.15 (0.36)	0.16 (0.37)	0.19 (0.39)	0.20 (0.40)
Associates degree	0.08 (0.27)	0.09 (0.29)	0.05 (0.22)	0.07 (0.26)	0.07 (0.25)	0.08 (0.28)
Bachelors degree	0.19 (0.39)	0.19 (0.40)	0.15 (0.36)	0.13 (0.34)	0.19 (0.39)	0.18 (0.39)
Advanced degree	0.11 (0.32)	0.08 (0.27)	0.11 (0.32)	0.07 (0.26)	0.09 (0.29)	0.09 (0.29)
Race (Proportion in Category)						
White	0.89 (0.31)	0.89 (0.31)	0.92 (0.28)	0.91 (0.29)	0.84 (0.36)	0.80 (0.40)
Black	0.07 (0.25)	0.07 (0.25)	0.04 (0.20)	0.06 (0.24)	0.11 (0.32)	0.17 (0.37)
Other	0.04 (0.19)	0.04 (0.20)	0.04 (0.20)	0.03 (0.18)	0.04 (0.20)	0.04 (0.19)
Urban Resident (Proportion in Category)	0.74 (0.44)	0.74 (0.44)	0.76 (0.43)	0.65 (0.48)	0.80 (0.40)	0.81 (0.39)
Sample Size	11,739	11,739	6,153	2,505	6,031	7,026

Table 5
Labor Supply Equation Regression Results by Household Type
(standard errors in parentheses)

	Household Type					
	Married Both Working		Married One Working		Single	
Independent Variable	Husbands	Wives	Husbands	Wives	Males	Females
Stone-Geary Form						
Intercept	2,503.69 (20.00)	2,132.16 (24.58)	2,105.40 (17.66)	1,887.51 (50.89)	2,360.79 (28.62)	2,343.38 (23.80)
Family Non-Labor Income / Predicted Hourly Wage ($I/\hat{w}$)	-0.14 (0.01)	-0.04 (0.00)	-0.26 (0.01)	-0.08 (0.01)	-0.13 (0.02)	-0.25 (0.02)
Reciprocal of Predicted Hourly Wage ($1/\hat{w}$)	-4,530.18 (259.08)	-1,936.67 (212.82)	-1,223.32 (160.99)	-1,932.67 (442.00)	-4,100.35 (332.10)	-3,631.54 (221.14)
Number of Dependents	9.45 (4.51)	-114.74 (5.32)	90.97 (6.65)	-44.27 (13.38)	43.09 (15.33)	-31.16 (7.60)
R^2	0.033	0.061	0.114	0.031	0.032	0.086
Linear Form						
Intercept	1,773.26 (23.18)	1,614.94 (24.68)	1,585.72 (27.65)	1,345.47 (47.85)	1,610.23 (33.19)	1,484.71 (29.00)
Predicted Hourly Wage ($\hat{w}$)	26.83 (1.43)	30.21 (2.18)	25.38 (1.47)	29.53 (4.18)	30.90 (2.42)	44.92 (2.55)
Family Non-labor Income (I)	-0.01 (0.00)	-0.004 (0.000)	-0.01 (0.00)	-0.01 (0.00)	-0.01 (0.00)	-0.02 (0.00)
Number of Dependents	8.17 (4.50)	-114.98 (5.31)	99.26 (6.62)	-39.34 (13.32)	43.59 (15.37)	-36.84 (7.63)
R^2	0.031	0.062	0.128	0.029	0.029	0.075
Sample Size	11,739	11,739	6,153	2,505	6,031	7,026

Table 6
Wage Equation Regression Results by Household Type
(standard errors in parentheses)

	Household Type					
	Married Both Working		Married One Working		Single	
Independent Variable	Husbands	Wives	Husbands	Wives	Males	Females
Intercept	-1.54	2.11	-0.72	2.96	0.62	-0.02
	(0.61)	(0.75)	(1.18)	(1.49)	(2.56)	(0.57)
Age in Years	0.21	0.06	0.17	0.03	0.15	0.10
	(0.01)	(0.01)	(0.02)	(0.02)	(0.05)	(0.01)
Education (Omit 8 years or less)						
9 to 12 years	2.64	1.06	2.35	1.45	2.23	1.38
	(0.53)	(0.68)	(1.02)	(1.29)	(2.11)	(0.48)
High School grad	4.48	2.69	5.79	2.66	4.61	3.18
	(0.45)	(0.59)	(0.84)	(1.11)	(1.78)	(0.42)
Some college	5.77	3.98	7.96	3.92	6.45	4.70
	(0.47)	(0.61)	(0.95)	(1.21)	(1.86)	(0.43)
Associates degree	6.86	6.35	10.13	5.82	6.71	5.71
	(0.51)	(0.64)	(1.27)	(1.38)	(2.20)	(0.48)
Bachelors degree	9.25	7.24	13.25	9.19	10.40	8.05
	(0.47)	(0.61)	(0.94)	(1.24)	(1.86)	(0.44)
Advanced degree	12.15	11.47	19.08	12.39	11.63	10.42
	(0.49)	(0.65)	(1.01)	(1.37)	(2.04)	(0.47)
Race (Omit White)						
Black	-1.72	-0.48	-3.40	-1.45	-1.98	-0.70
	(0.30)	(0.36)	(1.12)	(0.99)	(1.18)	(0.20)
Other	-0.71	-0.25	-1.57	-0.44	-1.15	-0.17
	(0.40)	(0.44)	(1.12)	(1.34)	(1.83)	(0.38)
Urban Resident	2.81	2.15	2.56	2.60	1.50	2.32
	(0.18)	(0.20)	(0.53)	(0.49)	(0.93)	(0.19)
R^2	0.176	0.095	0.109	0.082	0.014	0.197
Sample Size	11,739	11,739	6,153	2,505	6,031	7,026

Table 7
Stone-Geary Utility Function Parameter Estimates by Household Type
(standard errors in parentheses)

	Household Type					
	Married Both Working		Married One Working		Single	
Parameter	Husbands	Wives	Husbands	Wives	Males	Females
α	0.139 (0.012)	0.037 (0.003)	0.262 (0.013)	0.082 (0.011)	0.133 (0.018)	0.246 (0.016)
γ_1	5,854.7 (51.2)	6,547.0 (29.2)	5,907.1 (61.9)	6,704.1 (65.0)	6,038.5 (69.1)	5,653.6 (74.1)
γ_2	-32,774.2 (3,416.9)	-53,028.0 (7,908.7)	-4,668.7 (654.7)	-23,601.0 (6,495.7)	-30,933.5 (4,796.8)	-14,786.5 (1,314.7)
δ	11.0 (5.2)	-119.1 (5.5)	123.3 (8.8)	-48.2 (14.7)	49.7 (17.7)	-41.3 (9.9)
R^2	0.033	0.052	0.114	0.031	0.032	0.086
Mean Number of Dependents	1.4	1.4	1.4	0.9	0.2	0.7
Sample Size	11,739	11,739	6,153	2,505	6,031	7,026

α - Share of full budget devoted to leisure.

γ_1 - Minimum leisure required before utility is defined.

γ_2 - Minimum income required before utility is defined.

δ - Minimum leisure required per dependent.

$\Delta = \delta D$ - Minimum leisure (non-market time) required for dependents.

Table 8
Partial Effect and Elasticity Estimates of Labor Supply Implied by the Stone-Geary Utility Function for Various Household Types
(standard errors in parentheses)

	Household Type					
	Married Both Working		Married One Working		Single	
Effect	Husbands	Wives	Husbands	Wives	Males	Females
	Stone-Geary Form					
$(\partial H/\partial w)$[a]	22.222 (1.188)	28.380 (2.059)	10.652 (0.664)	25.187 (4.055)	24.788 (1.913)	38.228 (2.026)
$(\partial H/\partial I)$[b]	-0.00[illegible] (0.001)	-0.003 (0.000)	-0.016 (0.001)	-0.008 (0.001)	-0.010 (0.001)	-0.023 (0.001)
S[c]	42.321 (2.350)	34.041 (2.322)	43.306 (1.975)	37.358 (4.627)	44.953 (3.573)	82.332 (3.697)
$(\eta_{H.w})$[d]	0.153 (0.008)	0.183 (0.013)	0.085 (0.005)	0.169 (0.027)	0.163 (0.013)	0.213 (0.011)
$(\eta_{H.I})$[e]	-0.012 (0.001)	-0.073 (0.006)	-0.049 (0.002)	-0.051 (0.007)	-0.009 (0.001)	-0.030 (0.002)
$(\eta^{S}_{H.w})$[f]	0.291 (0.015)	0.220 (0.012)	0.348 (0.014)	0.251 (0.028)	0.295 (0.022)	0.458 (0.019)
$H = H(\bar{w}, \bar{I}, \hat{B})$	21,655	1,650	2,053	1,560	2,014	1,887
Mean Hourly Wage (w)	14.89	10.64	16.47	10.50	13.24	10.51
Mean Household Non-labor Income (I)	2,867	34,953	6,361	10,297	1,862	2,394
Sample Size	11,739	11,739	6,153	2,505	6,031	7,026

[a] $(\partial H/\partial w)$ = Cournot wage effect

[b] $(\partial H/\partial I)$ = pure income effect

[c] S = substitution effect = $(\partial H/\partial w) - H(\partial H/\partial I)$

[d] $(\eta_{H.w})$ = wage elasticity = $(\partial H/\partial w)(w/H)$

[e] $(\eta_{H.I})$ = income elasticity = $(\partial H/\partial I)(I/H)$

[f] $(\eta^{s}_{H.w})$ = substitution elasticity

$= (\eta_{H.w}) - (\bar{w}H/\bar{I})(\eta_{H.I})$

Table 9
Partial Effect and Elasticity Estimates of Labor Supply Implied by the Linear Labor Supply Function for Various Household Types

	Household Type					
	Married Both Working		Married One Working		Single	
Effect	Husbands	Wives	Husbands	Wives	Males	Females
Linear Labor Supply						
$(\partial H/\partial w)$[a]	26.827	30.205	25.380	29.532	30.898	44.923
	(1.430)	(2.180)	(1.474)	(4.185)	(2.416)	(2.545)
$(\partial H/\partial I)$[b]	-0.007	-0.003	-0.014	-0.006	-0.006	-0.019
	(0.001)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
S[c]	41.418	36.044	54.040	38.637	43.938	80.909
	(2.368)	(2.410)	(2.260)	(4.672)	(3.600)	(3.927)
$(\eta_{H.w})$[d]	0.185	0.195	0.204	0.199	0.203	0.250
	(0.010)	(0.014)	(0.012)	(0.028)	(0.016)	(0.014)
$(\eta_{H.I})$[e]	-0.009	-0.075	-0.043	-0.039	-0.006	-0.024
	(0.001)	(0.007)	(0.002)	(0.007)	(0.001)	(0.002)
$(\eta^{S}_{H.w})$[f]	-0.285	0.233	0.434	0.260	0.289	0.451
	(0.016)	(0.016)	(0.018)	(0.031)	(0.024)	(0.022)
$H = H(\bar{w}, \bar{I}, \hat{B})$	2,165	1,650	2,053	1,560	2,014	1,887
Mean Hourly Wage (w)	14.89	10.64	16.47	10.50	13.24	10.51
Mean Household Non-labor Income (I)	2,867	34,953	6,361	10,297	1,862	2,394
Sample Size	11,739	11,739	6,153	2,505	6,031	7,026

[a] $(\partial H/\partial w)$ = Cournot wage effect

[b] $(\partial H/\partial I)$ = pure income effect

[c] S = substitution effect = $(\partial H/\partial w) - H(\partial H/\partial I)$

[d] $(\eta_{H.w})$ = wage elasticity = $(\partial H/\partial w)(w/H)$

[e] $(\eta_{H.I})$ = income elasticity = $(\partial H/\partial I)(I/H)$

[f] $(\eta^{s}_{H.w})$ = substitution elasticity

$= (\eta_{H.w}) - (\overline{w}H/\overline{I})(\eta_{H.I})$

Table 10
Benefit Rights Provisions in the State UI Laws
of Michigan, Massachusetts, Illinois, and Oregon for the year 1991[a]

	Michigan Average-weekly-wage (MI)	Massachusetts High-quarter (MA)	Illinois Multi-quarter (IL)	Oregon Annual-wage (OR)
Base Period (BP)	52 weeks preceding BY	52 weeks preceding BY	1st 4 of last 5 quarters	1st 4 of last 5 quarters
Earnings to Qualify	20x30x min. wage	30x WBA	$1,600 in BP	$1,000 in BP
Employment to Qualify	20 weeks in BP	NS	NS	18 weeks in BP
Weekly Benefit Amount (WBA)	0.7 x Net AWW	1/21 to 1/26 of HQ earnings + dependant's allowance	49% of 2 highest quarters / 26	0.0125 x AWW
Min-Max WBA	$59 - $276	($14,$21) - ($282, $423)	$51 - ($206 - $270)	$57 - $247
To Qualify for Max WBA	$10,840 in BP	$7,332 in HQ	$10,881 in BP	$19,760 in BP
Entitled Benefit Duration:				
Max:				
Weeks	26	30	26	26
Dollars	$7,176	$8,460-$12,690	$5,356-$7,020	$6,422
Min:				
Weeks	15	10	26	5
Dollars	$ 825	$ 432	$1,600	$ 333
Provisions for Dependents:				
	Number of dependents is taken into account in after-tax weekly wage calculation.	$25 per dependent, up to $141.	5% of 2 highest quarters divided by 26.	NS

[a] Source: U.S. Department of Labor (1992), "Comparison of State Unemployment Insurance Laws", Manpower Administration, Unemployment Insurance Service, January.

BY: Benefit Year
WBA: Weekly Benefit Amount
BP: Base Period
NS: Not specified in the particular state law.
HQ: High Quarter

Table 11
Unemployment Compensation Simulation Estimates in Dollars
for Married Men with a Working Spouse[a]

	No Dependents							Two Dependents						
Weeks	Michigan	Massachusetts	Illinois	Oregon	Half	Stone-Geary	Linear	Michigan	Massachusetts	Illinois	Oregon	Half	Stone-Geary	Linear
1	276	0	0	0	327	30	13	276	0	0	0	327	13	29
2	552	282	206	247	653	100	72	552	376	270	247	653	66	105
3	828	564	412	494	980	205	178	828	752	540	494	980	156	229
4	1,104	846	618	741	1,306	341	333	1,104	1,128	810	741	1,306	280	401
5	1,380	1,128	824	988	1,633	506	536	1,380	1,504	1,080	988	1,633	434	622
6	1,656	1,410	1,030	1,235	1,959	696	788	1,656	1,880	1,350	1,235	1,959	614	892
7	1,932	1,692	1,236	1,482	2,286	910	1,090	1,932	2,256	1,620	1,482	2,286	818	1,211
8	2,208	1,974	1,442	1,729	2,612	1,145	1,441	2,208	2,632	1,890	1,729	2,612	1,045	1,579
9	2,484	2,256	1,648	1,976	2,939	1,399	1,842	2,484	3,008	2,160	1,976	2,939	1,292	1,998
10	2,760	2,538	1,854	2,223	3,265	1,671	2,294	2,760	3,384	2,430	2,223	3,265	1,557	2,467
11	3,036	2,820	2,060	2,470	3,592	1,959	2,796	3,036	3,760	2,700	2,470	3,592	1,839	2,987
12	3,312	3,102	2,266	2,717	3,918	2,263	3,350	3,312	4,136	2,970	2,717	3,918	2,137	3,558
13	3,588	3,384	2,472	2,964	4,245	2,580	3,955	3,588	4,512	3,240	2,964	4,245	2,450	4,181
14	3,864	3,666	2,678	3,211	4,571	2,911	4,612	3,864	4,888	3,510	3,211	4,571	2,776	4,855
15	4,140	3,948	2,884	3,458	4,898	3,254	5,322	4,140	5,264	3,780	3,458	4,898	3,114	5,582
16	4,416	4,230	3,090	3,705	5,224	3,608	6,084	4,416	5,640	4,050	3,705	5,224	3,464	6,362
17	4,692	4,512	3,296	3,952	5,551	3,972	6,900	4,692	6,016	4,320	3,952	5,551	3,825	7,195
18	4,968	4,794	3,502	4,199	5,877	4,347	7,769	4,968	6,392	4,590	4,199	5,877	4,196	8,082
19	5,244	5,076	3,708	4,446	6,204	4,731	8,693	5,244	6,768	4,860	4,446	6,204	4,576	9,023
20	5,520	5,358	3,914	4,693	6,530	5,123	9,671	5,520	7,144	5,130	4,693	6,530	4,966	10,019
21	5,796	5,640	4,120	4,940	6,857	5,524	10,705	5,796	7,520	5,400	4,940	6,857	5,364	11,070
22	6,072	5,922	4,326	5,187	7,183	5,933	11,793	6,072	7,896	5,670	5,187	7,183	5,770	12,176
23	6,348	6,204	4,532	5,434	7,510	6,349	12,938	6,348	8,272	5,940	5,434	7,510	6,184	13,338
24	6,624	6,486	4,738	5,681	7,836	6,772	14,140	6,624	8,648	6,210	5,681	7,836	6,604	14,557
25	6,900	6,768	4,944	5,928	8,163	7,201	15,398	6,900	9,024	6,480	5,928	8,163	7,032	15,833
26	7,176	7,050	5,150	6,175	8,489	7,637	16,714	7,176	9,400	6,750	6,175	8,489	7,466	17,166
27	7,176	7,332	5,356	6,422	8,816	8,079	18,087	7,176	9,776	7,020	6,422	8,816	7,906	18,557
28	7,176	7,614	5,356	6,422	9,142	8,526	19,520	7,176	10,152	7,020	6,422	9,142	8,351	20,007
29	7,176	7,896	5,356	6,422	9,469	8,979	21,011	7,176	10,528	7,020	6,422	9,469	8,802	21,516
30	7,176	8,178	5,356	6,422	9,795	9,437	22,562	7,176	10,904	7,020	6,422	9,795	9,259	23,084
31	7,176	8,460	5,356	6,422	10,122	9,900	24,173	7,176	11,280	7,020	6,422	10,122	9,720	24,713

[a] These results were computed using mean values of the hourly wage rate (w=$14.89) and family non-labor income (I=$2,867) from the sample of 11,739 husbands with working wives.

Weeks = Number of weeks unemployed in the year.
Michigan = Compensation payable in Michigan, an Average Weekly Wage State.
Massachusetts = Compensation payable in Massachusetts a High-quarter State.
Illinois = Compensation payable in Illinois, a Multi-quarter State.
Oregon = Compensation payable in Oregon, an Annual Wage State.
Half = Half of lost wages.
Stone-Geary = Full compensation at the means given the Stone-Geary Utility Function.
Linear = Full compensation at the means given the Linear Labor Supply Function.

Table 13
Unemployment Compensation Simulation Estimates in Dollars
for Married Men with a Non-working Spouse[a]

	No Dependents							Two Dependents						
Weeks	Michigan	Massachusetts	Illinois	Oregon	Half	Stone-Geary	Linear	Michigan	Massachusetts	Illinois	Oregon	Half	Stone-Geary	Linear
1	276	0	0	0	347	83	86	276	0	0	0	347	1,181	98
2	552	282	206	247	693	17	26	552	396	270	247	693	785	197
3	828	564	412	494	1,040	0	1	828	792	540	494	1,040	485	330
4	1,104	846	618	741	1,386	27	11	1,104	1,188	810	741	1,386	266	499
5	1,380	1,128	824	988	1,733	93	56	1,380	1,584	1,080	988	1,733	118	702
6	1,656	1,410	1,030	1,235	2,079	194	136	1,656	1,980	1,350	1,235	2,079	32	941
7	1,932	1,692	1,236	1,482	2,426	327	253	1,932	2,376	1,620	1,482	2,426	0	1,217
8	2,208	1,974	1,442	1,729	2,772	489	407	2,208	2,772	1,890	1,729	2,772	17	1,529
9	2,484	2,256	1,648	1,976	3,119	679	597	2,484	3,168	2,160	1,976	3,119	77	1,878
10	2,760	2,538	1,854	2,223	3,465	893	825	2,760	3,564	2,430	2,223	3,465	175	2,264
11	3,036	2,820	2,060	2,470	3,812	1,129	1,091	3,036	3,960	2,700	2,470	3,812	308	2,689
12	3,312	3,102	2,266	2,717	4,158	1,386	1,395	3,312	4,356	2,970	2,717	4,158	473	3,152
13	3,588	3,384	2,472	2,964	4,505	1,663	1,739	3,588	4,752	3,240	2,964	4,505	666	3,654
14	3,864	3,666	2,678	3,211	4,851	1,957	2,122	3,864	5,148	3,510	3,211	4,851	885	4,195
15	4,140	3,948	2,884	3,458	5,198	2,268	2,544	4,140	5,544	3,780	3,458	5,198	1,128	4,776
16	4,416	4,230	3,090	3,705	5,544	2,594	3,007	4,416	5,940	4,050	3,705	5,544	1,392	5,398
17	4,692	4,512	3,296	3,952	5,891	2,935	3,512	4,692	6,336	4,320	3,952	5,891	1,677	6,061
18	4,968	4,794	3,502	4,199	6,237	3,290	4,057	4,968	6,732	4,590	4,199	6,237	1,980	6,765
19	5,244	5,076	3,708	4,446	6,584	3,657	4,645	5,244	7,128	4,860	4,446	6,584	2,300	7,511
20	5,520	5,358	3,914	4,693	6,930	4,036	5,276	5,520	7,524	5,130	4,693	6,930	2,636	8,301
21	5,796	5,640	4,120	4,940	7,277	4,426	5,950	5,796	7,920	5,400	4,940	7,277	2,987	9,133
22	6,072	5,922	4,326	5,187	7,623	4,826	6,668	6,072	8,316	5,670	5,187	7,623	3,352	10,009
23	6,348	6,204	4,532	5,434	7,970	5,237	7,430	6,348	8,712	5,940	5,434	7,970	3,730	10,931
24	6,624	6,486	4,738	5,681	8,316	5,656	8,238	6,624	9,108	6,210	5,681	8,316	4,120	11,897
25	6,900	6,768	4,944	5,928	8,663	6,085	9,092	6,900	9,504	6,480	5,928	8,663	4,521	12,909
26	7,176	7,050	5,150	6,175	9,009	6,522	9,992	7,176	9,900	6,750	6,175	9,009	4,932	13,968
27	7,176	7,332	5,356	6,422	9,356	6,967	10,940	7,176	10,296	7,020	6,422	9,356	5,354	15,074
28	7,176	7,614	5,356	6,422	9,702	7,420	11,936	7,176	10,692	7,020	6,422	9,702	5,785	16,229
29	7,176	7,896	5,356	6,422	10,049	7,880	12,980	7,176	11,088	7,020	6,422	10,049	6,225	17,432
30	7,176	8,178	5,356	6,422	10,395	8,346	14,075	7,176	11,484	7,020	6,422	10,395	6,673	18,685
31	7,176	8,460	5,356	6,422	10,742	8,819	15,219	7,176	11,880	7,020	6,422	10,742	7,129	19,988

[a] These results were computed using mean values of the hourly wage rate (w=$16.47) and family non-labor income (I=$6,361) from the sample of 6,153 husbands with non-working wives.

Weeks = Number of weeks unemployed in the year.
Michigan = Compensation payable in Michigan, an Average Weekly Wage State.
Massachusetts = Compensation payable in Massachusetts a High-quarter State.
Illinois = Compensation payable in Illinois, a Multi-quarter State.
Oregon = Compensation payable in Oregon, an Annual Wage State.
Half = Half of lost wages.
Stone-Geary = Full compensation at the means given the Stone-Geary Utility Function.
Linear = Full compensation at the means given the Linear Labor Supply Function.

Table 14
Unemployment Compensation Simulation Estimates in Dollars
for Married Women with a Non-working Spouse[a]

	No Dependents							Two Dependents						
Weeks	Michigan	Massachusetts	Illinois	Oregon	Half	Stone-Geary	Linear	Michigan	Massachusetts	Illinois	Oregon	Half	Stone-Geary	Linear
1	182	0	0	0	174	108	62	193	0	0	0	173	347	1
2	364	173	170	198	347	188	138	386	223	222	198	347	468	8
3	546	346	340	396	521	286	245	579	446	444	396	521	602	45
4	728	519	510	594	694	400	383	772	669	666	594	694	748	113
5	910	692	680	792	868	528	551	965	892	888	792	868	904	212
6	1,092	865	850	990	1041	669	750	1,158	1,115	1,110	990	1,041	1,070	342
7	1,274	1,038	1,020	1,188	1,215	820	981	1,351	1,338	1,332	1,188	1,215	1,245	503
8	1,456	1,211	1,190	1,386	1,388	982	1,243	1,544	1,561	1,554	1,386	1,388	1,428	695
9	1,638	1,384	1,360	1,584	1,562	1,154	1,536	1,737	1,784	1,776	1,584	1,562	1,618	919
10	1,820	1,557	1,530	1,782	1,735	1,334	1,861	1,930	2,007	1,998	1,782	1,735	1,815	1,174
11	2,002	1,730	1,700	1,980	1,909	1,521	2,218	2,123	2,230	2,220	1,980	1,909	2,018	1,461
12	2,184	1,903	1,870	2,178	2,082	1,716	2,607	2,316	2,453	2,442	2,178	2,082	2,227	1,780
13	2,366	2,076	2,040	2,376	2,256	1,917	3,028	2,509	2,676	2,664	2,376	2,256	2,442	2,132
14	2,548	2,249	2,210	2,574	2,429	2,124	3,481	2,702	2,899	2,886	2,574	2,429	2,661	2,516
15	2,730	2,422	2,380	2,772	2,603	2,338	3,967	2,895	3,122	3,108	2,772	2,603	2,886	2,932
16	2,912	2,595	2,550	2,970	2,776	2,559	4,486	3,088	3,345	3,330	2,970	2,776	3,115	3,381
17	3,094	2,768	2,720	3,168	2,950	2,779	5,037	3,281	3,568	3,552	3,168	2,950	3,348	3,863
18	3,276	2,941	2,890	3,366	3,123	3,007	5,622	3,474	3,791	3,774	3,366	3,123	3,585	4,378
19	3,458	3,114	3,060	3,564	3,297	3,239	6,240	3,667	4,014	3,996	3,564	3,297	3,826	4,926
20	3,640	3,287	3,230	3,762	3,470	3,475	6,891	3,860	4,237	4,218	3,762	3,470	4,070	5,508
21	3,822	3,460	3,400	3,960	3,644	3,715	7,576	4,053	4,460	4,440	3,960	3,644	4,317	6,124
22	4,004	3,633	3,570	4,158	3,817	3,959	8,295	4,246	4,683	4,662	4,158	3,817	4,568	6,773
23	4,186	3,806	3,740	4,356	3,991	4,206	9,048	4,439	4,906	4,884	4,356	3,991	4,821	7,456
24	4,368	3,979	3,910	4,554	4,164	4,456	9,835	4,632	5,129	5,106	4,554	4,164	5,078	8,174
25	4,550	4,152	4,080	4,752	4,338	4,709	10,657	4,825	5,352	5,328	4,752	4,338	5,337	8,926
26	4,732	4,325	4,250	4,950	4,511	4,965	11,513	5,018	5,575	5,550	4,950	4,511	5,598	9,713
27	4,732	4,498	4,420	5,148	4,685	5,224	12,404	5,018	5,798	5,772	5,148	4,685	5,862	10,534
28	4,732	4,671	4,420	5,148	4,858	5,485	13,330	5,018	6,021	5,772	5,148	4,858	6,128	11,391
29	4,732	4,844	4,420	5,148	5,032	5,749	14,292	5,018	6,244	5,772	5,148	5,032	6,397	12,283
30	4,732	5,017	4,420	5,148	5,205	6,015	15,289	5,018	6,467	5,772	5,148	5,205	6,667	13,210
31	4,732	5,190	4,420	5,148	5,379	6,284	16,322	5,018	6,690	5,772	5,148	5,379	6,940	14,174

[a] These results were computed using mean values of the hourly wage rate (w=\$10.50) and family non-labor income (I=\$10,297) from the sample of 2,505 wives with non-working husbands.

Weeks = Number of weeks unemployed in the year.
Michigan = Compensation payable in Michigan, an Average Weekly Wage State.
Massachusetts = Compensation payable in Massachusetts a High-quarter State.
Illinois = Compensation payable in Illinois, a Multi-quarter State.
Oregon = Compensation payable in Oregon, an Annual Wage State.
Half = Half of lost wages.
Stone-Geary = Full compensation at the means given the Stone-Geary Utility Function.
Linear = Full compensation at the means given the Linear Labor Supply Function.

Table 15
Unemployment Compensation Simulation Estimates in Dollars for Single Men[a]

	No Dependents							Two Dependents						
Weeks	Michigan	Massachusetts	Illinois	Oregon	Half	Stone-Geary	Linear	Michigan	Massachusetts	Illinois	Oregon	Half	Stone-Geary	Linear
1	273	0	0	0	274	38	14	276	0	0	0	274	8	170
2	546	273	206	247	547	107	70	552	323	270	247	547	3	310
3	819	546	412	494	821	205	167	828	646	540	494	821	38	491
4	1,092	819	618	741	1,094	330	307	1,104	969	810	741	1,094	107	715
5	1,365	1,092	824	988	1,368	480	489	1,380	1,292	1,080	988	1,368	206	982
6	1,638	1,365	1,030	1,235	1,641	650	714	1,656	1,615	1,350	1,235	1,641	333	1,291
7	1,911	1,638	1,236	1,482	1,915	841	982	1,932	1,938	1,620	1,482	1,915	484	1,643
8	2,184	1,911	1,442	1,729	2,188	1,049	1,293	2,208	2,261	1,890	1,729	2,188	657	2,039
9	2,457	2,184	1,648	1,976	2,462	1,274	1,649	2,484	2,584	2,160	1,976	2,462	849	2,478
10	2,730	2,457	1,854	2,223	2,735	1,514	2,048	2,760	2,907	2,430	2,223	2,735	1,060	2,961
11	3,003	2,730	2,060	2,470	3,009	1,767	2,491	3,036	3,230	2,700	2,470	3,009	1,287	3,489
12	3,276	3,003	2,266	2,717	3,282	2,034	2,979	3,312	3,553	2,970	2,717	3,282	1,529	4,061
13	3,549	3,276	2,472	2,964	3,556	2,312	3,511	3,588	3,876	3,240	2,964	3,556	1,786	4,678
14	3,822	3,549	2,678	3,211	3,829	2,601	4,089	3,864	4,199	3,510	3,211	3,829	2,055	5,340
15	4,095	3,822	2,884	3,458	4,103	2,900	4,712	4,140	4,522	3,780	3,458	4,103	2,335	6,047
16	4,368	4,095	3,090	3,705	4,376	3,209	5,382	4,416	4,845	4,050	3,705	4,376	2,627	6,801
17	4,641	4,368	3,296	3,952	4,650	3,527	6,097	4,692	5,168	4,320	3,952	4,650	2,929	7,600
18	4,914	4,641	3,502	4,199	4,923	3,853	6,858	4,968	5,491	4,590	4,199	4,923	3,241	8,446
19	5,187	4,914	3,708	4,446	5,197	4,187	7,666	5,244	5,814	4,860	4,446	5,197	3,561	9,338
20	5,460	5,187	3,914	4,693	5,470	4,529	8,522	5,520	6,137	5,130	4,693	5,470	3,890	10,278
21	5,733	5,460	4,120	4,940	5,744	4,877	9,425	5,796	6,460	5,400	4,940	5,744	4,227	11,265
22	6,006	5,733	4,326	5,187	6,017	5,232	10,375	6,072	6,783	5,670	5,187	6,017	4,571	12,300
23	6,279	6,006	4,532	5,434	6,291	5,593	11,374	6,348	7,106	5,940	5,434	6,291	4,922	13,383
24	6,552	6,279	4,738	5,681	6,564	5,960	12,421	6,624	7,429	6,210	5,681	6,564	5,279	14,514
25	6,825	6,552	4,944	5,928	6,838	6,333	13,517	6,900	7,752	6,480	5,928	6,838	5,643	15,694
26	7,098	6,825	5,150	6,175	7,111	6,711	14,662	7,176	8,075	6,750	6,175	7,111	6,012	16,924
27	7,098	7,098	5,356	6,422	7,385	7,093	15,857	7,176	8,398	7,020	6,422	7,385	6,387	18,203
28	7,098	7,371	5,356	6,422	7,658	7,481	17,102	7,176	8,721	7,020	6,422	7,658	6,768	19,531
29	7,098	7,644	5,356	6,422	7,932	7,873	18,397	7,176	9,044	7,020	6,422	7,932	7,153	20,911
30	7,098	7,917	5,356	6,422	8,205	8,269	19,742	7,176	9,367	7,020	6,422	8,205	7,543	22,341
31	7,098	8,190	5,356	6,422	8,479	8,670	21,139	7,176	9,690	7,020	6,422	8,479	7,937	23,822

[a] These results were computed using mean values of the hourly wage rate (w=$13.24) and family non-labor income (I=$1,862) from the sample of 6,031 single males.

Weeks = Number of weeks unemployed in the year.
Michigan = Compensation payable in Michigan, an Average Weekly Wage State.
Massachusetts = Compensation payable in Massachusetts a High-quarter State.
Illinois = Compensation payable in Illinois, a Multi-quarter State.
Oregon = Compensation payable in Oregon, an Annual Wage State.
Half = Half of lost wages.
Stone-Geary = Full compensation at the means given the Stone-Geary Utility Function.
Linear = Full compensation at the means given the Linear Labor Supply Function.

Table 16
Unemployment Compensation Simulation Estimates in Dollars
for Single Women[a]

	No Dependents							Two Dependents						
Weeks	Michigan	Massachusetts	Illinois	Oregon	Half	Stone-Geary	Linear	Michigan	Massachusetts	Illinois	Oregon	Half	Stone-Geary	Linear
1	219	0	0	0	210	51	25	230	0	0	0	210	132	1
2	438	210	206	247	420	99	67	460	260	269	247	420	202	6
3	657	420	412	494	630	163	128	690	520	538	494	630	285	31
4	876	630	618	741	840	240	210	920	780	807	741	840	381	76
5	1,095	840	824	988	1,050	330	311	1,150	1,040	1,076	988	1,050	488	141
6	1,314	1,050	1,030	1,235	1,260	432	433	1,380	1,300	1,345	1,235	1,260	605	227
7	1,533	1,260	1,236	1,482	1,470	545	576	1,610	1,560	1,614	1,482	1,470	733	334
8	1,752	1,470	1,442	1,729	1,680	669	740	1,840	1,820	1,883	1,729	1,680	870	461
9	1,971	1,680	1,648	1,976	1,890	802	925	2,070	2,080	2,152	1,976	1,890	1,016	610
10	2,190	1,890	1,854	2,223	2,100	945	1,131	2,300	2,340	2,421	2,223	2,100	1,171	779
11	2,409	2,100	2,060	2,470	2,310	1,096	1,359	2,530	2,600	2,690	2,470	2,310	1,334	971
12	2,628	2,310	2,266	2,717	2,520	1,256	1,608	2,760	2,860	2,959	2,717	2,520	1,504	1,184
13	2,847	2,520	2,472	2,964	2,730	1,424	1,879	2,990	3,120	3,228	2,964	2,730	1,681	1,419
14	3,066	2,730	2,678	3,211	2,940	1,600	2,173	3,220	3,380	3,497	3,211	2,940	1,866	1,676
15	3,285	2,940	2,884	3,458	3,150	1,782	2,489	3,450	3,640	3,766	3,458	3,150	2,057	1,956
16	3,504	3,150	3,090	3,705	3,360	1,971	2,828	3,680	3,900	4,035	3,705	3,360	2,254	2,259
17	3,723	3,360	3,296	3,952	3,570	2,166	3,190	3,910	4,160	4,304	3,952	3,570	2,457	2,584
18	3,942	3,570	3,502	4,199	3,780	2,368	3,575	4,140	4,420	4,573	4,199	3,780	2,665	2,933
19	4,161	3,780	3,708	4,446	3,990	2,575	3,983	4,370	4,680	4,842	4,446	3,990	2,879	3,305
20	4,380	3,990	3,914	4,693	4,200	2,788	4,415	4,600	4,940	5,111	4,693	4,200	3,098	3,701
21	4,599	4,200	4,120	4,940	4,410	3,006	4,871	4,830	5,200	5,380	4,940	4,410	3,322	4,121
22	4,818	4,410	4,326	5,187	4,620	3,230	5,352	5,060	5,460	5,649	5,187	4,620	3,551	4,565
23	5,037	4,620	4,532	5,434	4,830	3,458	5,857	5,290	5,720	5,918	5,434	4,830	3,784	5,034
24	5,256	4,830	4,738	5,681	5,040	3,690	6,387	5,520	5,980	6,187	5,681	5,040	4,022	5,527
25	5,475	5,040	4,944	5,928	5,250	3,927	6,942	5,750	6,240	6,456	5,928	5,250	4,264	6,046
26	5,694	5,250	5,150	6,175	5,460	4,169	7,523	5,980	6,500	6,725	6,175	5,460	4,509	6,591
27	5,694	5,460	5,356	6,422	5,670	4,414	8,129	5,980	6,760	6,994	6,422	5,670	4,759	7,161
28	5,694	5,670	5,356	6,422	5,880	4,663	8,762	5,980	7,020	6,994	6,422	5,880	5,012	7,757
29	5,694	5,880	5,356	6,422	6,090	4,917	9,421	5,980	7,280	6,994	6,422	6,090	5,269	8,380
30	5,694	6,090	5,356	6,422	6,300	5,173	10,107	5,980	7,540	6,994	6,422	6,300	5,529	9,029
31	5,694	6,300	5,356	6,422	6,510	5,433	10,820	5,980	7,800	6,994	6,422	6,510	5,793	9,706

[a] These results were computed using mean values of the hourly wage rate (w=$10.51) and family non-labor income (I=$2,394) from the sample of 7,026 single females.

Weeks = Number of weeks unemployed in the year.
Michigan = Compensation payable in Michigan, an Average Weekly Wage State.
Massachusetts = Compensation payable in Massachusetts a High-quarter State.
Illinois = Compensation payable in Illinois, a Multi-quarter State.
Oregon = Compensation payable in Oregon, an Annual Wage State.
Half = Half of lost wages.
Stone-Geary = Full compensation at the means given the Stone-Geary Utility Function.
Linear = Full compensation at the means given the Linear Labor Supply Function.

Table 17
Proportion of Sub-sample with One-half Earnings Replaced when the Maximum Weekly Benefit Amount (WBA) is Set at Various Fractions of the Full-sample Average Weekly Wage (AWW=$519)

Fraction of AWW	Married Both Working Husbands	Married Both Working Wives	Married One Working Husbands	Married One Working Wives	Single Males	Single Females	Total
0.50	0.41	0.76	0.44	0.80	0.57	0.72	0.60
0.55	0.47	0.81	0.49	0.84	0.63	0.77	0.65
0.60	0.55	0.86	0.55	0.88	0.69	0.82	0.71
0.67	0.64	0.90	0.61	0.91	0.77	0.87	0.77
0.70	0.66	0.91	0.63	0.91	0.78	0.88	0.79
0.75	0.73	0.93	0.68	0.94	0.83	0.91	0.83
AWW (std. dev.)	653 (396)	378 (274)	693 (538)	347 (301)	547 (802)	420 (287)	519 (470)
Sample Size	11,739	11,739	6,153	2,505	6,031	7,026	45,193

Table 18
Maximum Weekly Benefit Amount (WBA) Required for Each Sub-sample to Yield *One-half for Four-Fifths*

Married Both Working Husbands	Married Both Working Wives	Married One Working Husbands	Married One Working Wives	Single Males	Single Females	Total
442	279	500	260	375	302	375

THE ROLE OF UNEMPLOYMENT INSURANCE IN ADDRESSING STRUCTURAL UNEMPLOYMENT: LESSONS FROM OTHER NATIONS

Duane E. Leigh*

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*Professor of Economics, Washington State University. The financial support of the Advisory Council on Unemployment Compensation, U.S. Department of Labor is gratefully acknowledged. Dan Hamermesh and Ernie Stromsdorfer supplied helpful comments on the initial draft of this study. Orley Ashenfelter and Robert LaLonde served as discussants when the paper was presented at the ACUC research conference held in Burlington, VT on August 17 and 18, 1995, and their suggestions are greatly appreciated.

ABSTRACT

Most industrialized nations, including the United States, are currently faced with rising long-term unemployment associated with increased structural unemployment as a fraction of total unemployment. The U.S. leads the world in funding demonstration projects designed to test the effectiveness of active labor market policies in combating structural unemployment. However, we have much less experience than many other nations with the actual implementation of these policies on a nationwide basis. This paper examines unemployment compensation system reforms recently implemented in Britain, Australia, and Canada as responses to rising long-term unemployment. Broadly speaking, these reforms are intended to enhance work incentives of unemployment compensation recipients and to strengthen the linkage between unemployment benefits and reemployment services. A number of lessons are drawn from the experiences of these nations that might be usefully considered in reforming our unemployment insurance system.

INTRODUCTION

A recent U.S. Department of Labor (1993) report documents a disturbing upward trend in structural unemployment, where structural unemployment refers to a permanent loss of the old job and difficulty in finding a new job. Job finding difficulty arises, in turn, because of the mismatch between skill requirements of vacant jobs and skills possessed by the unemployed. Structural unemployment is generally thought to be an unavoidable outcome of a dynamic economic system buffeted by demand-side shifts associated with new technologies, global markets, corporate restructurings, and the downsizing of the military. To the extent that the structurally unemployed possess substantial prelayoff labor market experience, as is often the case, they are also termed displaced workers.

Evidence supporting the existence of an upward trend in structural unemployment includes the observation that since the 1950s, long-term unemployment (those out of work for six months or longer) has risen at a faster pace than overall unemployment (see Table 1). For example, the average unemployment rate for the 1990-93 period is only slightly higher than the average rate for the decade of the 1970s, but the proportion of unemployment consisting of long-term unemployment has jumped by nearly half (from 11.0 percent to 16.0 percent). In 1992, moreover, some 76 percent of unemployed job losers did not expect to be recalled to their old jobs (USDOL 1993). This 76-percent level is reported to be the highest proportion of job losers not on temporary layoff ever recorded since these data become available in 1967.

The implication drawn from this evidence by USDOL analysts (1993) and many other observers is that traditional short-term income support provided through the unemployment insurance (UI) system is not sufficient to meet the challenge of rising structural unemployment. The argument is that UI income support is well-suited to workers on temporary layoff during a cyclical downturn and to those who can readily find new jobs on their own. At the same time, it is not a sufficient policy response to the needs of the structurally unemployed whose reemployment prospects are not expected to substantially improve with an upswing in the economy. Using the Organization for Economic Cooperation and Development (OECD) distinction between "active" and "passive" labor market policies, the USDOL report goes on to recommend increased funding for an active labor market policy designed to speed up reemployment and improve the long-run earnings potential of displaced workers. The expectation is that if active labor market programs are successful in expediting the reemployment of the displaced, fewer dollars will need to be spent on passive labor market programs like unemployment insurance.

Most of the suggested components of this "reemployment" response to rising structural unemployment were subsequently incorporated in the Reemployment Act of 1994. Major provisions of this proposed legislation are the following:

- A comprehensive reemployment system designed to assist all jobless workers, regardless of the cause of their job loss.[1]
- One-stop career centers designed to make it easy for the unemployed to access all available reemployment services at one location.
- A nationwide labor market information network.
- Early identification of displaced workers -- termed "profiling" -- so that they can promptly be referred to reemployment services.
- Long-term retraining services including income support provided to workers seeking to upgrade their skills to match the requirements of vacant jobs in growing industries. Displaced workers enrolled in an approved training program who had worked at least three years for a previous employer would qualify for up to one year of income support

beyond the usual six-month maximum. Workers with one or two years of job tenure would qualify for up to six months of additional support.

- Changes in the UI system to expedite the return to work including reemployment bonuses to claimants who find jobs quickly and measures to promote self-employment.[2]

The Reemployment Act failed to receive Congressional approval during the 1994 session. However, a policy response to the needs of displaced workers is not an issue that is likely to go away. In formulating new policies to assist displaced workers, the United States is in the fortunate position of being able to draw on the recent experience of other industrialized nations. Three nations -- Britain, Australia, and Canada -- are emphasized in the analysis. As described in the next section, each of these nations faced a more severe problem of long-term unemployment than confronts U.S. policymakers today, and each recently took action to make its unemployment compensation system more directly responsive to the needs of the structurally unemployed. The second of these points deserves emphasis. While the U.S. leads the world in funding demonstration projects to test the effectiveness of labor market policy innovations, we have less experience than most other nations with the actual implementation of such policies. As is widely noted in the evaluation literature, effects of a permanent program may be quite different from those of a comparable experiment.

The final section of the study draws lessons from the experiences of these three nations that might be usefully considered in discussing possible UI reforms in the context of a broader reemployment system.

UNEMPLOYMENT COMPENSATION REFORMS IN OTHER NATIONS

The English-speaking nations of Britain, Australia, and Canada are like the U.S. in that custom and national labor market policy make it much more likely that employers will respond to demand shocks by laying off workers than by adjusting compensation or hours of work. For their part, jobless workers in these countries are usually expected to prepare for and locate vacant jobs with relatively little government assistance. In contrast, most Western European countries support a comprehensive and stable institutional structure through which a wide variety of employment and training services are supplied. Sweden is a leading example of a Western European nation that has invested heavily in active labor market programs delivered through an established employment and training system (see Leigh 1995: Chap. 2). These nations typically also impose advance notice and severance pay requirements on employers who dismiss workers and encourage hours adjustments in lieu of layoffs.

An important way in which the English-speaking countries mentioned differ from one another is in the resources devoted to income support programs. The unemployment compensation (UC) system of the United States is undoubtedly the least generous of the four. Table 2 raises the important distinction in UC systems between unemployment insurance and unemployment assistance (UA). UI benefits are paid for a limited time period to workers who are involuntarily unemployed and actively seeking new employment. Eligibility for UI benefits depends on a past record of insured employment, and the amount of benefits received often hinges on past earnings. These conditions imply that many individuals may be unemployed but not eligible for UI benefits. Indeed, an important issue facing U.S. policymakers is the downward trend over the postwar period in the fraction of the unemployed receiving UI benefits (see Blank and Card 1991).

As distinct from unemployment insurance, unemployment assistance is typically of indefinite duration with eligibility for benefits independent of employment history. Moreover, the amount of UA benefits is determined by the level of income and assets of other household members via a means-test. Since unemployment assistance is not conditional on past employment history and may be of indefinite duration, the coverage of UA programs is likely to be greater than the coverage of UI. Table 2 points out that Australia offers a UA program to its workers, while in Britain UI is combined with a UA program to provide eligible workers with income support of unlimited length. Canada and the U.S. both offer unemployed workers a UI program only, but the potential duration of benefits is much longer in Canada. In addition, Canada, unlike the U.S., makes UI benefits available to voluntary job leavers.

Table 3 shows using OECD data for the second half of the 1980s that Britain, Australia, and, to a lesser extent, Canada all faced a more severe problem of long-term unemployment than did the U.S. In each case, policymakers took steps during the late 1980s or early 1990s to reform their UC systems to address the problem of rising long-term unemployment.

Britain

In 1911 the U.K. became the first highly industrialized country to institute a national compulsory unemployment insurance program. Blaustein and Craig (1977: 224) describe that during the 1920s and early 1930s high levels of unemployment meant that benefit outlays far surpassed the resources generated by the UI program's contributory financing structure. A reform carried out during the mid-1930s placed a limit on benefit duration beyond which workers who were still unemployed and could meet a means-test were eligible for further income support from a new unemployment assistance program. This program became known as Income Support.

Britain's public employment service historically was responsible for both disbursing unemployment benefits and providing labor exchange services including applicant screening and job placement. Placement service offices -- called Jobcentres -- were later separated from unemployment benefit offices in 1973, and in 1982 the legal requirement that benefit claimants register for job placement was dropped. Following a period of minimal intervention during the first half of the 1980s, British policymakers began in 1986 to implement programs emphasizing the provision of reemployment assistance services to the long-term unemployed. These initiatives are discussed below under the heading of the "scheduling and programming strategy." This programmatic emphasis was strengthened after 1989 by a multi-year program to bring about a gradual merging of the Jobcentres and Benefit Offices networks. The OECD (1993: 19) notes that by the mid-1990s, the network of merged Jobcentres and Benefit Offices is expected to include between 1,000 and 1,200 "integrated" offices offering one-stop access to employment services. Beginning in 1988, the provision of adult training services was made the responsibility of a national network of Training and Enterprise Councils (TECs). The TEC network represents an effort to decentralize the delivery of training services to the local labor market level and increase the role of the local business community in designing and operating training programs.

Unemployment Benefits

As noted in Table 2, unemployed British workers are eligible for up to 12 months of unemployment insurance benefits. Britain's UI system is unusual in that benefits are unrelated to previous earnings. Reubens (1989) suggests that academic research indicating that higher earnings replacement ratios tend to prolong unemployment led to the abolishment in 1982 of earnings-related UI benefits, leaving only a basic flat rate benefit. Storey and Neisner (1992: Table

3) report for the early 1990s that UI benefits are fixed at £34.71 weekly for a single person and £56.12 weekly for a person with a dependent spouse. Using December 31, 1991 exchange rates, these benefit levels convert to about $65 weekly for a single person and about $105 weekly for a person with a dependent spouse. The earnings replacement ratio is therefore quite low, except for persons with low previous earnings.

It is interesting to note that the academic research referred to by Reubens actually indicates that only quite large cuts in benefits would have much of an impact on length of unemployment spells. For example, a well-known study by Lancaster and Nickell (1980) estimates the elasticity of unemployment spells with respect to the earnings replacement ratio to be about 0.6. This means that a 10 percent rise in benefits would be associated with an increase in length of unemployment of one week, given an average unemployment duration of 17 weeks. Using a larger micro data set, a more recent study by Narendranathan, Nickell, and Stern (1985) concludes that this elasticity is even smaller, lying somewhere in the 0.28 to 0.36 interval.[3]

Table 2 also points out that unemployed workers who have exhausted their UI benefits or who failed to qualify for UI because of a lack of previous employment experience are eligible for the unemployment assistance program Income Support. Assuming no other sources of income, Income Support benefits are fixed at £39.65 ($74.17) weekly for a single person and £62.25 ($116.45) weekly for a couple, plus allowances for children (Storey and Neisner 1992: 26). Income Support claimants are also eligible for a housing subsidy under the Housing Benefit program. Reubens (1989) notes that between 1973 and 1983, the balance of unemployment compensation expenditures shifted from UI to Income Support benefits. During the same period, the proportion of the unemployed who received neither UI nor Income Support benefits shrank from almost 25 percent to less than 13 percent. A recent estimate suggests that three-quarters of all unemployed workers receive only Income Support benefits (see OECD 1992: 141).

Because Income Support benefits are related to family size, the OECD (1993: 66) raises the issue that Income Support plus Housing Benefit payments might well exceed after-tax earnings from low-wage jobs.[4] The potential for long-term dependency on British safety net programs has been widely noted. For example, Murray (1990) observes that among young male Britons, in particular, a belief in the "right" to full unemployment benefits appears to have developed without any sense that this right hinges on a past record of employment and on a reciprocal obligation to actively seek employment. The "official" British government perspective is expressed in an important White Paper presented to Parliament in December of 1988. In that document, the Department of Employment (1988: 55) writes that

> . . . there is evidence that a significant minority of benefit claimants are not actively looking for work. Some are claiming benefit fraudulently while working at least part-time in the black [or underground] economy. Others seem to have grown accustomed to living on benefit and have largely given up looking for work, despite the high level of job vacancies which are increasingly available throughout the country. Others believe, mistakenly, that they might be financially worse off taking a job or are reluctant to travel daily more than a short distance to where jobs are available.

The government's response to the problem of long-term dependency took basically three forms: (1) increase the return to work even at low wages through the Family Credit program, (2) encourage the long-term unemployed to reestablish contact with the labor market through the scheduling and programming strategy, and (3) increase the resources devoted to combating fraud. The first of these responses will be briefly described, while the second and third will be considered at greater length.

The Family Credit is designed to make most Income Support claimants financially better off working, even at a low wage, than they would be remaining unemployed. The program was initially restricted to individuals working over 24 hours a week on average; in 1992, this restriction was reduced to 16 hours per week. Workers meeting the weekly hours restriction receive a tax-free Family Credit benefit in addition to their labor market earnings, but the benefit is withdrawn at an implicit marginal tax rate of 70 percent applied to net earnings. Since the tax rate is high and the breakeven level of earnings is low, Family Credit payments are most likely to have a sizable impact on family income in cases where wages are low or one wage-earner is supporting a large family. The OECD (1993: 66) estimates that only about half of potentially eligible workers claim the Family Credit, and since 1989 the government has funded advertising campaigns to increase the program's takeup rate.

The Scheduling and Programming Strategy

To revitalize job search activity among the long-term unemployed, the Restart program was launched in 1986 at Jobcentres throughout the nation. Restart basically links a mandatory sequence of personal interviews to the offer of various reemployment services including more proactive, relative to traditional labor exchange services, job search assistance (JSA) and retraining programs. A Restart interview was initially required of individuals who had been unemployed for more than 12 months. The OECD (1993: 85) suggests that it's early success led to the extension of Restart to those unemployed for more than six months. Interviews are currently scheduled every six months for the duration of the unemployment spell (see Table 4).

In 1988 the role of the New Client Adviser was introduced to provide early access to JSA services for newly unemployed workers. For many years previously, new benefit claims typically involved only processing the necessary paperwork. The New Client Adviser supplies new claimants with a verbal presentation of various reemployment assistance options immediately rather than after months of unemployment. As indicated in Table 4, the initial interview with a New Client Adviser also results in a mutually agreed upon back-to-work plan which lays out a strategy for the claimant to follow in seeking reemployment. If the unemployment spell continues, subsequent interviews are conducted by a Claimant Adviser who is an expert on in-work benefits and the local job market. During the first six months of unemployment, the unemployed are viewed as being fundamentally job-ready but in need of assistance in sharpening job search skills.

This attitude changes after six months when an unemployed individual becomes eligible for the Restart program. Restart is intended for unemployed workers who may be beginning to doubt the value of continuing to actively engage in job search. As noted, the program begins with an interview with a Claimant Adviser. These interviews are repeated every six months with the same Claimant Adviser so that a good working relationship can be developed between adviser and client. Depending on the needs of the client, the adviser can recommend a variety of services including a slot in a TEC-operated adult skill training program, a place in a Jobclub, access to the Job Interview Guarantee program, a temporary public sector job under the Employment Action program, financial assistance to set up a small business, and a slot in a Restart Course. Eligible unemployed workers undergoing skill training qualify for continued income support payments plus a training allowance of £10 per week. Among the other available programs, two of the more important are Jobclubs and Restart Courses.

Job finding clubs are basically peer-support groups of jobless workers who meet to share their experiences in looking for work. In Britain, Jobclubs offer workers unemployed for at least six months two weeks of four half-day sessions per week of structured training on how to look for jobs, write more effective job applications, and come across well in interviews. Following these two weeks, clients are given access to "resource areas" containing telephones and sources of labor market information. At this stage, there is more support from other club members and less from the Jobclub leader. Jobclub participation is limited to four to six months to prevent the clubs from becoming used more for their social aspects than for their job search function. The OECD (1993: 86) reports that there are about 1,000 Jobclubs in Britain, with an average entry of about 150 unemployed persons per club per year. About 200 of the Jobclubs are operated by Employment Service staff, with the other 800 operated by outside suppliers on two-year contracts.

Attendance at a Restart Course is mandatory after two years of receiving unemployment benefits for those claimants who have turned down the opportunity to participate in other programs. Restart Courses are directed at those long-term unemployed persons who have settled into the mind-set that they are unlikely to ever get another job and have adjusted to life on the dole and a low standard of living. In addition to providing information on the range of available job search assistance options, Restart Courses involve small groups of about 12 individuals in exercises intended to overcome barriers to work, sort out career options, motivate the resumption of active job search, and produce an individualized reemployment plan. Courses are typically five days in length.

Fraudulent Claims

As might be expected, the most common form of fraud in Britain is continuing to claim Income Support benefits while receiving undeclared earnings which would, if declared, reduce or eliminate these benefits. Atkinson and Micklewright (1989) document for the 1979-88 period that the credible threat of benefit disqualification increased due to steps taken to tighten monitoring of benefit claims and job search activity. As of 1989, the OECD (1993: 88) reports that more than 10 percent of claimants sent a letter of invitation to attend a Restart interview stopped claiming benefits rather than attend the interview. More recently, experience with compulsory Restart Courses indicates that about nine percent of those scheduled for the courses stopped claiming benefits. While these dropout rates could be reflecting a successful reemployment experience, it is more likely that many who dropped their claim to benefits have access to other sources of income, in which case their Income Support claim could well be fraudulent. Supporting this conclusion is U.S. evidence for the Charleston, South Carolina (see Corson, Long, and Nicholson 1985) and Washington state (see Johnson and Klepinger 1994) UI reform experiments. This evidence suggests that strengthened work search reporting requirements and mandatory attendance at a job search workshop reduced UI payments. But the saving in UI expenditures was achieved by raising the costs of remaining on UI rather than by enhancing claimants' job search skills.

There are also individuals illegally claiming benefits who are undeterred by job search requirements. To combat this form of benefit fraud, the British Employment Service as of 1991-92 maintained a staff of about 1,000 fraud inspectors (in comparison to about 20,000 staff members in other benefit administration tasks). The OECD (1993: 72) reports that in recent years, inspectors investigated 300,000 to 400,000 cases per year leading to 65,000 to 85,000 withdrawals of claims and 3,500 to 4,500 prosecutions. Given that withdrawn claims lead to savings in Income Support payments and Housing Benefits and to increased tax receipts, it seems likely that fraud detection more than pays for itself.

Australia

Created in 1944, Australia's unemployment compensation system provides unemployed workers with benefits that vary by income, age, marital status, number of children, amount of rent, and location of residence. Benefits continue for an unlimited period of time. Although it is an unemployment assistance program (see Table 2), Australia's UC system was designed to meet the traditional goal of maintaining the incomes of workers temporarily unemployed during cyclical downturns in the economy. Long-term unemployment was largely unknown, and the system remained essentially unchanged for its first 40 years.

The event that led to a major revision of the UC system was the surge in structural unemployment during the 1982-83 recession (see DSS and DEET 1993: 13-15). Left to their own devices, many unemployed Australians laid off during the recession appeared to lack the skills and motivation to take advantage of job opportunities created by the subsequent economic recovery. The consequence was a period beginning in the mid-1980s and continuing into the early 1990s in which Australian policymakers were confronted with the dilemma of labor shortages co-existing with rising long-term unemployment. Table 3 indicates that in 1987 Australia faced a long-term unemployment problem that was less severe than Britain's but more severe than Canada's, while Australia's national unemployment rate was lower than that of either nation.

The initial policy response to rising long-term unemployment was the NEWSTART strategy implemented on a national basis in June 1989. Underlying NEWSTART is the principle of reciprocal obligation. That is, continued receipt of unemployment benefits by the long-term unemployed is made conditional on claimants' willingness to take advantage of available job placement and retraining assistance, with the ultimate objective of finding employment or improving employability. Monitoring of the work test was enhanced under NEWSTART by an upgraded interview process administered jointly by the Department of Social Security (DSS) and the Department of Employment, Education and Training (DEET). DSS pays out unemployment benefits, while DEET administers employment services delivered through the network of Commonwealth Employment Service (CES) offices. The DSS/DEET joint interview carried out with the long-term unemployed involved a range of assessment and counseling services leading up to a referral to a job vacancy or to a slot in a suitable labor market program. Client failure to attend a scheduled joint interview or to willingly follow up on a suggested referral could lead to a judgment that the work test had been failed, followed by suspension or termination of benefits.

The Newstart Strategy

In July 1991, the Newstart program replaced NEWSTART as a strategy to assist jobless workers, both the short-term and the long-term unemployed, to obtain reemployment. While NEWSTART was imposed on the existing UC program, Newstart substitutes two new mechanisms for providing unemployment benefits -- the Job Search Allowance paid to workers older than age 18 for the first 12 months of unemployment (and to eligible persons under 18 years of age) and the Newstart Allowance paid beyond 12 months of unemployment. Primary objectives of the Newstart program are to (1) prevent long-term unemployment by early intervention, (2) differentiate the services provided to the short-term and long-term unemployed, and (3) strengthen the reciprocal obligation of benefit recipients introduced by NEWSTART. Benefit levels are unchanged under Newstart. Table 5 outlines the timing of DSS and CES interventions in the Newstart program.

The Job Search Allowance. Newstart is designed to encourage unemployed workers to actively engage in job search from the time they first apply for unemployment benefits. As indicated in Table 5, eligibility for Job Search Allowance benefits is conditional on meeting requirements, known as "activity testing," to ensure that the unemployed are genuinely seeking employment. For the job-ready, activity testing involves contacting employers about possible vacancies, applying for suitable jobs, and indicating a willingness to accept suitable offers. Benefit recipients must provide the CES with a list of employers contacted on a bi-weekly basis.

Unemployed workers who are determined not to be immediately job-ready may meet the activity testing requirement by making themselves available for a job search assistance or skills training course. Job Search Allowance payments continue while clients are enrolled in an approved training program, and program participants receive a training allowance to help defray out-of-pocket training costs. Failure to meet the activity test means a temporary suspension of the Job Search Allowance or its permanent cancellation. All clients are subject to a DSS entitlement review with possible referral to CES employment services after three months of unemployment. After six months, the Commonwealth Employment Service carries out a second interview to assess claimants' employability and eligibility for labor market programs.

The Newstart Allowance. Continued receipt of unemployment benefits beyond a period of 12 months is not automatic. After 12 months, the unemployed must apply for the Newstart Allowance if they wish to continue receiving income support. At the time they apply, all applicants must review their situation with a CES staff member and develop a reemployment plan known as a Newstart Plan and Activity Agreement. The idea behind the Activity Agreement is first to diagnose those problems preventing the applicant from competing successfully in the labor market. Then the agreement lays out an action plan to succeed in meeting specified reemployment goals. Activities that may be proposed by the CES and agreed to by the client include job search training, vocational skills training, paid work experience, and measures designed to eliminate or reduce labor market disadvantages such as medical treatment or rehabilitation. As with the Job Search Allowance, participants in an approved training course receive the Newstart Allowance plus a training allowance.

The Newstart Activity Agreement takes the form of a concise, three-page written document. On the third page of the agreement, each Newstart Allowance recipient must indicate his or her understanding of and agreement to the following statement:

> My Newstart Allowance will continue to be paid only if I take reasonable steps to keep to this Agreement, and to my responsibilities set out in my copy of the Newstart Allowance claim form.

Also as part of the agreement, Newstart Allowance recipients are required to maintain a written record of their job search activity which is reviewed on a regular basis by a CES staff member (typically, the same CES staff member to encourage a more personal relationship between the CES and clients). Employer cooperation is sought in monitoring the work search effort of recipients. Employer Contact Certificates issued to job seekers by the CES are one form of evidence of active job search. When appropriate, employers may be requested to verify the job seeker's attempts to find work by completing the back of these certificates. Employers doubting the sincerity of a job seeker are encouraged to contact the CES with their concerns so that action can be taken.

An issue discussed earlier in connection with the British Family Credit program was how to make income support recipients financially better off working than they would be remaining unemployed. Under the Newstart program, long-term unemployed Australians are eligible for a variety of inducements to encourage reemployment. To help meet the costs of returning to work, the long-term unemployed are eligible to receive an employment entry payment of A$100 when they notify the Department of Social Security of their reemployment.[5] At least partial payment of the Newstart Allowance is usually also paid to help tide over the newly reemployed until their first pay check. Continuing help to low-income workers with dependent children is available through the Family Allowance Supplement (FAS). Under FAS, individuals taking low-wage jobs may continue to receive a DSS payment for children and rent assistance in addition to their earnings. Finally, individuals who accept a short-term job are eligible for immediate resumption of the Newstart Allowance when their job ends.

Evaluation Evidence

In 1994 an interim evaluation report on Newstart was released by the Australian government (see Sakkara et al. 1994). The report is based on surveys of CES staff and clients, a post-implementation review of Newstart, and net impact studies of four specific labor market programs. It should be noted that a worsening of labor market conditions occurred between 1991 and 1993. The large increase in the proportion of long-term and very long-term unemployed (12 months and 24 months of joblessness, respectively) during this interval taxed the capacity of Newstart and reduced the labor market opportunities available to CES clients upon leaving the program.

Some of the main findings of the evaluation report include the following:

- CES assistance to the long-term unemployed under Newstart increased at a faster rate than growth in the number of the long-term unemployed. This suggests that Newstart is meeting its objective of making assistance to the long-term unemployed a higher priority.

- Newstart assistance is distributed unevenly across client groups. Younger clients are more likely to receive job referrals and JOBSTART placements in a subsidized job, while slots in other labor market programs including job training and job search assistance activities tend to go to prime-aged jobseekers. The older unemployed and the very long-term unemployed receive less assistance than other groups.
- Newstart's concept is viewed favorably by staff.
- Clients report generally positive attitudes toward their Activity Agreements.
- The increased number of long-term unemployed during the observation period limited the ability of CES staff to implement the Newstart objective of providing individually-tailored assistance and may have led to inadequate activity testing.
- Breakdowns in DSS/DEET coordination led to difficulties in providing uniform delivery of services and consistent quality in client contacts.

The main conclusion reached in the evaluation is that more intensive and expensive labor market services such as JOBSTART should be reserved for the long-term unemployed. Except for those judged to be at high risk of joining the long-term unemployed, assistance to the short-term unemployed should be restricted to self-service activities and job club programs.

Canada

The passage in 1940 of the Unemployment Insurance Act established Canada's UI program with the primary objective of providing insurance against risk of temporary income loss due to unemployment. Coverage was restricted to workers with a strong previous labor market attachment. As noted earlier, included among covered workers are those who quit their previous job. Also in 1940, the National Employment Service was created to provide labor exchange services delivered through a national network of Canada Employment Centres (CECs). The National Employment Service is also responsible for enforcing the work search requirements of the UI Act.

The 1971 Reforms

Between 1940 and 1971 only gradual changes were made in the UI program, mainly in the direction of extending coverage to previously excluded groups. Much more dramatic changes expanding and liberalizing the system occurred in amendments to the UI Act passed in 1971. These changes were the following:

1. A substantial increase in coverage, so that virtually all wage and salary workers came under the UI Act.
2. An increase in the earnings replacement ratio to 67 percent of pre-tax earnings (with 75 percent provided to claimants with dependents who had low earnings or prolonged unemployment).
3. `An increase in the maximum benefit, with the maximum benefit subsequently indexed annually to changes in average wages.
4. A reduction in the minimum period of previous employment required to qualify for benefits from 30 weeks to just 8 weeks, regardless of economic conditions.
5. Introduction of a benefit structure raising duration of regular benefits to a maximum of 43 weeks depending on claimants' previous work experience. Regular benefits could be followed by extended benefits the duration of which depends on the national unemployment rate and the number of percentage points by which regional unemployment rates exceed the national rate.
6. Making benefits taxable.

7. Introduction of new benefits in cases of earnings interruptions due to sickness, maternity, and retirement.

Green and Riddell (1993) point out that the introduction of regional extended benefits represented a shift in objectives away from a pure insurance program toward concern with altering the regional distribution of income. But they also suggest that the most important of these changes in terms of their labor market impacts are the reduction in the minimum qualifying period and the increase in the maximum duration of benefits. A short minimum qualifying period continues to be a controversial feature of the Canadian system. As recently as 1993, the Wall Street Journal (1993: A8) writes:

> [U]nemployment insurance is so generous that in some parts of the country it has become a way of life: three months of work and nine months on the dole. In the little province of Prince Edward Island, the New Democratic Party lays off its leader for three months each year so he can collect federal unemployment benefits and save the party about $3,100.

As suggested in this quotation, there is a high economic payoff to quite brief periods of employment, particularly in high unemployment regions like Prince Edward island.

The 1971 amendments were followed by a rapid expansion in UI expenditures throughout the 1970s and continuing into the 1980s. Expenditures on UI benefits grew from about C$700 million in 1970 to almost C$2 billion in 1972 and almost C$12 billion in 1983. By the second half of the 1980s, Table 6 indicates that the ratio of Canadian passive program expenditures to Gross Domestic Product (GDP) was closing in on the ratio for Britain, a country well known for the number of its citizens living more or less permanently on the dole. Indeed, the OECD (1994: Table 20) reports for the 1991-92 period that public expenditures on UC programs as a percentage of GDP is 2.25 percent for Canada as compared to 1.35 percent for the U.K. The ratio shown for Canada in Table 6 is considerably higher than that for Australia and much higher than the ratio for the U.S. (Sweden is included in the table as an example of a Western European nation with a high ratio of total labor market expenditures to GDP but with a quite different mix of active to passive programs compared to the English-speaking countries.)

Table 7 compares Canada and the United States on important characteristics of their respective UI systems. The table shows that unemployed Canadians enjoy a higher earnings replacement ratio and receive UI benefits for a longer period than unemployed Americans. But the most important consideration explaining Canada's greater UI expenditures is the near certainly that an unemployed Canadian will receive UI benefits, whereas just over one-quarter of unemployed Americans are UI recipients. One factor involved in explaining the difference in UI receipt is that, as noted, workers who quit their previous job are covered in Canada but not in the U.S. Also important is the decline in insured unemployment in the U.S. since 1982 (see Blank and Card 1991).

Following the 1971 amendments, rapidly rising costs led to a number of proposals for UI reform. In 1985 the Macdonald Commission suggested returning UI to its original social insurance objective while recommending the creation of a comprehensive negative income tax plan to supplement incomes of the poor, especially of the working poor. The Forget Commission in 1986 also recommended focusing the UI program on insurance objectives and developing a separate comprehensive income security program. Large federal government deficits during the second half of the 1980s made it difficult to implement the more comprehensive program of income support recommended by the two commissions in their UI reform proposals. Instead, the Canadian government embarked on an alternative strategy of shifting resources away from passive income support programs to active labor market policies including job search assistance, skills training programs, and employment subsidies. This general strategy has been strongly recommended by the OECD throughout the 1980s and into the 1990s (see, for example, OECD 1990).

The 1989 Labour Force Development Strategy

In 1989 the Canadian government released a White Paper describing problems of structural unemployment and perceived skill deficiencies among Canadian workers and outlining the components of a new policy initiative (see EIC 1989). Called the Labour Force Development Strategy (LFDS), the new initiative has the purpose of tightening up on UI eligibility while, at the same time, allowing a reallocation of UI resources from income support to active labor market policies. Major changes in the UI program include the following.

1. Increased eligibility requirements. While retaining the UI program's sensitivity to differences in regional unemployment rates, a new UI benefit schedule increases the minimum weeks of employment required to become eligible for UI benefits to 10 to 20 weeks of employment depending on the region's unemployment rate. (A 1981 amendment raised the work experience requirement from the 8 weeks established in 1971 to 10 to 14 weeks.)

2. Reduction in the maximum duration of benefits. The maximum period of UI entitlement is reduced from 50 weeks except in those regions suffering very high unemployment rates. For example, claimants who lived in regions with an unemployment rate of 11 percent and who had worked for at least 30 weeks were entitled to just 42 weeks rather than 50 weeks of UI benefits.

3. Penalties for voluntary job leavers without just cause. Workers who quit their jobs without just cause are still eligible to receive UI benefits, but they face a seven- to 12-week delay before benefits commence, and the duration of their benefits is reduced. In addition, the before-tax replacement ratio for insurable

earnings is cut to 50 percent from 60 percent. There is no penalty imposed on those who leave their employment for just cause, such as hazardous working conditions, following a spouse to a new location, and sexual harassment.

4. Strengthening penalties imposed for fraud. Penalties are significantly strengthened for fraudulent use of the UI program by claimants and employers. As noted earlier in connection with the British Income Support program, the most common type of claimant fraud involves failure to report earned income while collecting UI benefits.

Several empirical studies based on Canadian data are worth noting in connection with the 1989 reforms. Evidence on the effect of shortening the UI entitlement period (reform 2) is provided by Ham and Rea (1987). Between 1976 and 1979, the Canadian government made two statutory changes in the UI system likely to affect unemployment duration. The first of these reduced the replacement ratio for claimants with dependents from 75 percent to 67 percent of insurable pre-tax earnings, and the replacement ratio for all other claimants was lowered from 67 percent to 60 percent. Second, changes in regulations determining weeks of benefit entitlement reduced the average period of entitlement, particularly in regions with low unemployment rates. Using longitudinal data covering the 1975-80 period, Ham and Rea find that the period of UI entitlement has a statistically significant effect on unemployment duration, even for those who do not ultimately exhaust their benefits. Holding age constant, a typical point estimate indicates that a decrease in initial entitlement of one week reduces expected unemployment duration by 0.33 weeks (relative to a mean expected duration of 20.8 weeks).

Further evidence has recently become available relating to the effect of increased eligibility requirements (reform 1). Green and Riddell (1994) examine a natural experiment occurring in 1989 in which the variable eligibility requirement prevailing in the maximum entitlement regions of Canada (regions with an unemployment rate of at least 11.5 percent in 1989-90) was replaced by a fixed eligibility requirement of 14 weeks of employment. In these regions, 10 weeks of employment in 1989 provided UI claimants with up to 42 weeks of benefits, whereas in 1990, 14 weeks of employment were required to qualify for benefits. The authors find that the "spike" in the job-leaving rate observed at 10 weeks in 1989 increases to 14 or more weeks in 1990.

Continuing the focus on eligibility requirement, Lemieux and MacLeod (1994) find, using longitudinal data for the 1972-92 period, that UI plays different roles for different groups of workers.[6] For low frequency users, UI remains essentially a pure insurance system that protects against the risk of layoff during an economic downturn. For high frequency users, however, UI looks increasingly like a permanent income support program that has little to do with layoff risk. Controlling for business cycle effects, they conclude that rising long-term unemployment in Canada is the consequence of high frequency users accounting for an increasing share of UI spells relative to low frequency users. Providing support for this conclusion is Corak's (1993) evidence that insured spells of unemployment increase in duration for each additional UI claim filed. This form of "occurrence dependence" is hypothesized to arise because each new claim reduces the fixed costs of gaining information about how to benefit from the UI system. That is, as unemployed workers gain experience with UI, an increasingly large fraction of them become part-year workers.

With the enactment of the LFDS, Employment and Immigration Canada (EIC 1989: 11) projected an annual UI program saving of about C$1.3 billion. This represents about 10 percent of the UI program's total annual expenditures beginning in 1990. Of the C$1.3 billion saving, approximately C$500 million was projected to be spent on improved UI benefits for maternity, sickness, and parental leave, as well as for workers over the age of 65. The remaining C$800 million was to be dedicated to active labor market programs designed to raise skills and enhance job placement and job development services. In June 1989, Bill C-21 to amend the UI Act was introduced with the intent of providing a higher ceiling (15 percent) on UI account funds that could be earmarked for programs designed (in the language of the legislation) "to provide an earlier return to employment by improving the employability of UI claimants through training." This bill was passed in November 1990. The OECD (1994: 74) reports that funding for UI-sponsored training increased from under C$300 million in 1989 to C$2.2 billion in 1993.

The trend toward reducing UI generosity continued into the 1990s. In 1993, Bill C-113 made the following legislative changes: (1) disentitlement for benefits of voluntary quitters without just cause, and (2) a reduction in the benefit replacement ratio for all claimants from 60 percent to 57 percent of insurable earnings. Concerning the first change, an analysis by Crossley and Kuhn (1995) indicates little effect of the voluntary quit provision on the UI takeup rate of job quitters, probably because the legislation includes a long list of reasons for justified quits. Regarding the second, Jones (1995) tentatively concludes that the Bill C-113 benefit rate reduction has negligible effects on search inputs, reemployment probability, and unemployment duration. Still more recently, a 1994 amendment to the UI Act once again raised the required number of weeks of employment to qualify for UI benefits, this time to 12 to 20 weeks, depending on the region's unemployment rate, from the earlier 10 to 20 weeks.

Evidence on the Impact of Training

Among the three English-speaking countries, Canada is the only one for which there is available evaluation evidence addressing the question of whether government training programs available to UI recipients actually enhance their reemployment prospects. A brief description of each of the five programs evaluated follows.

Feepayer clients must qualify for UI and must have been out of school for more than two years. Feepayers are eligible to enroll in any training program recommended by their CEC counselor including, since 1991, basic skills training.

DIR clients either took training (often in the evening) that did not interfere with their job search or did not inform their local CEC office that they were enrolled in training while on UI. In the latter case, clients were either disqualified from UI or agreed to continue job search and to accept reasonable employment offers while in training.

Job Entry clients include women reentering the work force after an absence of at least three years and out-of-school youths with little labor market experience. Priority for the program is given to high school dropouts.

Job Development clients must have suffered long-term unemployment defined as being unemployed for at least 24 out of the previous 30 weeks.

Skill Shortages clients are workers recommended by their CEC counselors for training in skills designated at the national level as being in short supply. Eligible applicants are clients determined not to be job-ready but who did not meet the criteria for other programs.

In 1991 a restructuring of Canadian employment and training programs was carried out in which the Job Development, Job Entry, and Skill Shortages programs incorporated in the existing Canadian Jobs Strategy (CJS) were spun off to the new Employability Improvement Program (EIP).[7] Added to the EIP, in addition, is the Feepayer option which was previously outside the CJS. The Feepayer option permits UI-eligible workers to receive income support while exempting them from job search requirements during their training. However, Feepayer option participants themselves, or a third party, were required to pay for training costs. Following 1989, as noted, the Labour Force Development Strategy made UI funds available to pay these costs. The policy shift toward decentralized decision-making incorporated in EIP led to increased use of the Feepayer option as CEC counselors gained greater flexibility to work with UI clients in developing individually tailored reemployment plans.[8]

The evaluation study by Park, Riddell, and Power (1993) is primarily based on longitudinal data collected for samples of UI trainees and non-trainees selected from all claimants who received unemployment benefits between January 1, 1988 and June 30, 1991. The training provided to UI claimants participating in all five programs is classroom training. However, the average length of training courses differed significantly by program, with Feepayer program participants enrolled in the longest courses (33.5 weeks on average) and Skill Shortages participants enrolled in the shortest courses (19.2 weeks).

Availability of longitudinal data collected for trainees and non-trainees makes it possible to use a "difference-in-differences" econometric methodology to control for the selection bias likely to be present because unobserved factors that are fixed over time influence program selection as well as earnings. Using this methodology, Table 8 displays estimated net impacts measured in terms of 1991 earnings for the 1988 and 1989 cohorts of UI claimants. For each cohort, estimates are calculated for base year earnings measured both two years and one year prior to initial receipt of UI benefits. The impact of UI-sponsored training is seen to depend strongly on the particular program studied and on the cohort. Across both cohorts, only the Job Entry program results in consistently positive and statistically significant impacts on earnings. In addition, the Skill Shortages program records very large and statistically significant estimated impacts for the earlier cohort and distinctly positive but not significant estimates for the later cohort. It is interesting to note that both the Job Entry and Skill Shortages programs provided trainees a much greater opportunity to combine on-the-job training with classroom training than was possible in the other three programs (see Park, Riddell, and Power 1993: 29). Greater access to OJT would be expected to increase the short-term net impacts estimated for training programs.

For the other three programs, net impact estimates are positive and statistically significant only for the 1988 cohort of Feepayer program participants. Given the length of Feepayer training programs, it is possible that the negligible estimates obtained for the 1989 cohort may be the consequence of program participants having very little chance by the time earnings are measured in 1991 to recoup a return on their training investment. All of the estimates shown in Table 8 are based on a simple comparison of average earnings. Controlling for residence and a limited number of personal characteristics in a regression framework, program net impact estimates display the same general patterns seen in the table, although the individual estimates tend to be smaller in magnitude.

LESSONS FOR THE U.S.

The unemployment compensation systems in place in Britain, Australia, and Canada each originated as a policy response providing income support to families whose head suffered a temporary loss of employment. Britain's system was instituted in 1911, and the systems in Canada and Australia began in the early and mid-1940s, respectively. Over time the UC systems of each country were expanded to furnish income support to the long-term unemployed. Recognizing that their evolving UC systems exceeded the original goal of maintaining incomes of the cyclically unemployed and were adversely affecting the incentive to work, policymakers in these nations implemented reforms during the late 1980s and early 1990s designed to couple stronger work incentives with enhanced reemployment services.

As indicated in the introductory section, the interest of U.S. policymakers in unemployment insurance reform is largely in response to a perceived increase in structural unemployment, and especially to the adjustment assistance needs of displaced workers. Rising long-term unemployment associated with industrial restructuring is a world-wide phenomenon, and in Australia the rise in structural unemployment in the early 1980s provided the motivation for initiatives undertaken in 1989 and 1991. In Britain and Canada, however, structural unemployment was not the only motivating factor leading to unemployment compensation reform. British policymakers faced the additional challenge of altering ingrained lifestyles of dependence on the public dole. In Canada, policymakers were also obliged to deal with the fiscal consequences of earlier reforms liberalizing UI and causing its evolution toward an income support system for marginal workers.

The complex issues faced, in particular, by British and Canadian policymakers point to the difficulty in making a substantial dent in long-term unemployment with marginal or even more radical UC reforms. At the same time, the unprecedented recent rise in Sweden's unemployment rate emphasizes that even innovative and well-funded active labor market programs for retraining and employing the jobless cannot prevent unemployment from rising sharply in a severe recession.[9] With these cautionary remarks in mind, this concluding section outlines six lessons that may be instructive to U.S.

policymakers. The first two suggest a stay-the-course attitude regarding current UI policy. The remaining four indicate possible reforms, some of which are already well under way, which would enhance the role of UI in a reemployment system designed to better serve the needs of the structurally unemployed.

1. Caution in extending the length of UI benefits. Contributing to the severity of long-term unemployment in the three nations studied relative to the U.S. are their more generous unemployment compensation systems, in particular, the much longer maximum duration of benefits. Evidence provided for Canada (see Ham and Rea 1987 and Green and Riddell 1994) and Germany (see Hunt 1995) suggests a positive relationship between duration of potential UC benefits and length of unemployment spells. Such a relationship also exists for American workers. For example, Katz and Meyer (1990) estimate in a well-known study that an one-week increase in potential benefits lengthens the average duration of UI claimants' unemployment by 0.16 to 0.20 weeks.

The two-tier reemployment programs implemented in Britain (Restart) and Australia (Newstart) attest to the difficulty of reintegrating jobless workers into the work force once the duration of their unemployment has stretched into the long-term category. In Canada, the increase in length of potential UI benefits legislated in 1971 appears to have played an important role in the encouragement of intermittent employment and the subsequent growth of UI expenditures. Among other restrictive provisions, the Labour Force Development Strategy of 1989 reduced the maximum duration of UI benefits available to Canadians. Rising long-term unemployment associated with structural change has the natural consequence of increased political pressure to lengthen the potential duration of UI. An implication for U.S. policymakers is that such pressure to raise the six-month maximum of regular UI benefits and to liberalize the federal-state extended benefits program should be resisted.

2. The importance of requiring considerable prior employment in determining benefit eligibility. Not only did the 1971 amendments to the UI Act increase the maximum length of UI benefits, they also substantially reduced the minimum period of prior employment required for Canadian workers to qualify for UI. Seeking to slow the resulting growth of UI outlays, a key provision of the 1989 legislation boosted the minimum number of weeks of employment required for UI eligibility in most regions of the country. Subsequent amendments to the UI Act further increased work requirements. However, writing as recently as 1994, the OECD (1994: 98-99) comments that

> [The] evidence . . . suggests that UI benefits encourage marginal attachment to the labour force by promoting unstable and seasonal industries through the subsidisation of temporary layoffs. In sum, the net effect of the provision of UI benefits may imply a reverse situation where benefit provision leads to unemployment.

The Canadian experience with reducing prior employment requirements suggests that U.S. policymakers are on the right track in holding firm to existing requirements, even in the face of a long-term downward trend in the fraction of unemployed workers receiving UI benefits.[10]

3. Claimants should be contacted early and monitored closely. To prevent short-term unemployment spells from stretching into long-term spells, the British scheduling and programming strategy implemented in 1986 offers new unemployment compensation claimants a range of more proactive employment services that go beyond the traditional labor exchange function of public employment agencies. In particular, new claimants are required to agree to a back-to-work plan which is followed up with the offer of a variety of other job search assistance services during the first six months of unemployment. A similar design emphasizing early intervention and personalized reemployment plans is found in Australia's Newstart and the Canadian Employability Improvement Program. Both initiatives were implemented in 1991. Newstart is

especially noteworthy for its model emphasizing to income support recipients from the first week of their unemployment the reciprocal obligation to actively engage in job search.

Often termed "profiling," early intervention with offers of reemployment assistance is not a concept that is new to U.S. policymakers. In fact, a key objective of the New Jersey UI Reemployment Demonstration initiated in July 1986 was to assess the feasibility of identifying early in the claim period UI claimants who, in the absence of intervention, were likely to face prolonged spells of unemployment and exhaust UI benefits (see Corson, et al. 1989). As of 1994, Runner (1995) reports that 18 states had amended their UI laws to require as a condition of UI eligibility that claimants must participate in job search assistance programs. This condition applies to individuals determined through state profiling systems to be likely to exhaust regular UI benefits. The profiling initiative should continue to be encouraged at the federal level.[11]

A monetary incentive that might be proposed to speed up the reemployment process is the reemployment bonus. Of the three countries studied, only the A$100 employment entry payment available in Australia at all resembles a reemployment bonus scheme. However, considerable evaluation evidence is available from four U.S. reemployment bonus experiments. This evidence is not entirely favorable. In particular, estimated effects on weeks of UI receipt are about the same as those achieved, but at much lower cost, in demonstration projects designed to test the effectiveness of more closely monitoring work search behavior.

4. Opening up Employment Service assistance to UI claimants. All three countries studied provide UC claimants with full access to public employment agency services. As noted, these services include not only traditional labor exchange services but also more proactive JSA services and skill training

programs. Evidence for this country on the effectiveness of job search assistance is available from a series of four federally-funded displaced worker demonstration projects implemented during the 1980s. This evidence indicates quite clearly that JSA speeds up the reemployment of displaced workers (see Leigh 1995: Chap. 3 and Meyer 1995). The demonstrations also indicate that expediting workers' return to work does not significantly decrease the quality of the job match (as realized in post-layoff earnings) upon reemployment.

In contrast to the open access of UC claimants to public employment agency services in Britain, Australia, and Canada, the U.S. Employment Service has over time been obliged by Congressional mandates to target its services to welfare recipients and other economically disadvantaged groups. Baily, Burtless, and Litan (1993: 136-39) argue that this obligation to the economically disadvantaged leads to a process of adverse selection. Adverse selection arises because as ES clients became younger, poorer, and less skilled due to government regulations, employers searching for good workers looked elsewhere. But as more attractive employers shied away from listing their job vacancies with Job Service offices, average workers had less reason to apply for agency services, further reducing the average quality of workers served by the ES.

Giving all UI claimants access to Employment Service resources via the profiling process and working more closely with employers in job development efforts can be anticipated to speed up the reemployment of the jobless and increase the proportion of job vacancies that employers choose to list. Displaced workers are highly likely to qualify for these reemployment services because of their previous labor force attachment.

5. <u>A need for greater coordination between the Employment Service and UI systems</u>. Implementation of a worker profiling system points to the need for

greater ES/UI coordination. Such coordination is an essential element of the Australian Newstart initiative which is designed to more closely link continued eligibility for UI benefits to active job search. Similarly, the British government is moving ahead on a plan initiated in 1989 to merge its Jobcentres and Benefit Offices networks into a nationwide system of integrated offices offering convenient one-stop access to employment services for UC recipients.

For the U.S., evidence produced from the evaluation of the Charleston, South Carolina job search demonstration project indicates that ES/UI coordination resulting in closer work test monitoring and enhanced employment services to UI claimants is cost-effective (see Corson, Long, and Nicholson 1985). The strengthening of linkages between UI and ES and the then-existing Job Training Partnership Act (JTPA) system providing limited skill retraining programs also appears to have played a key role in the success of the New Jersey UI Reemployment Demonstration (see Corson et al. 1989).

The introductory section described that an important provision in the proposed Reemployment Act of 1994 authorized the creation of one-stop career centers designed to make it easy for the unemployed to access all available reemployment services at one location. The Reemployment Act failed to obtain Congressional approval, but Balducchi, Johnson, and Gritz (1995) note that Congress has recently taken action to appropriate limited funds for a national system of One-Stop Career Centers under existing Wagner-Peyser Act authority (the law establishing the ES in 1933). Closer coordination between the ES and UI systems, possibly implemented in one-stop career centers, is a policy recommendation that should continue to receive emphasis. Not only would closer ES/UI coordination be likely to improve claimants' access to reemployment services through the profiling process, but it would also increase the Employment Service's ability to monitor the UI work test.

One additional related issue ties in to the ES's traditional role of referring displaced workers to retraining options. Going back as far as the 1960s, political considerations often required the coupling of legislation relaxing trade barriers with funding for retraining programs targeted to workers likely to be displaced from their jobs by the legislation. The consequence is our current fragmented system of categorical programs. The Reemployment Act would have consolidated these programs into a comprehensive reemployment system. With its defeat, a proposal put forward by President Clinton in his 1995 state of the union address is to scrap the present system, using the funds freed up to finance essentially a voucher program which would allow UI claimants determined to be in need of retraining by their ES counselors to pay for training programs of their own choice. This proposal warrants serious consideration, and the Canadian Feepayer option providing funding for individualized training programs provides an example of how such a program might work.

6. Extend UI benefits to displaced workers engaged in long-term education and training programs. Britain and Australia provide income support payments to unemployed workers engaged in an approved retraining program. Similarly, the Feepayer option allows UI-eligible Canadian workers to receive income support while being exempted from job search requirements for the duration of their training. All three countries also help UI claimants meet the direct costs of retraining programs.

At present, American workers enrolled in approved training programs may receive a work test exemption in all states allowing them to receive up to six months of income support while undergoing retraining. Longer term assistance of up to 12 months of additional income support is available to only the limited number of trade-displaced workers who qualify for Trade Adjustment Assistance (TAA) retraining. As a consequence, only a small fraction of displaced workers who might benefit from training can be expected to possess the necessary financial resources to support themselves and their families while being retrained, particularly if the training period lasts more than a few months.

A widely-cited study by Kane and Rouse (1995) provides evidence that returns to community college programs are substantial and about the same per year of credits as those estimated for four-year colleges and universities. Community colleges are the primary subcontractors for federal- and state-funded adult training programs. This evidence for longer-term education and training contrasts favorably with the generally discouraging findings obtained for short-term classroom training programs implemented in the displaced worker demonstration projects of the 1980s (see Leigh 1995: Chap. 3).

The proposed Reemployment Act of 1994 would have authorized UI benefits for up to a year beyond the usual six months for claimants enrolled in training or educational activities. Given the cost and duration of most retraining programs, UI claimants may need to be carefully screened to limit access to only those who require retraining to qualify for jobs in other industries or to meet the higher skill requirements of jobs in their current industries. Subject to this restriction, a provision extending UI benefits to claimants undergoing longer-term training should be considered for inclusion in future legislation.

FOOTNOTES

1. This would replace the current multiplicity of categorical displaced worker programs including Economic Dislocation and Worker Adjustment Assistance (EDWAA), the main program currently supplying reemployment assistance to the displaced, and Trade Adjustment Assistance (TAA), which provides income assistance and retraining to trade-displaced workers.

2. Legislation passed in December 1993 implementing the North American Free Trade Agreement (NAFTA) permits states to establish self-employment assistance programs through which selected UI claimants may continue to receive regular UI payments while engaged full time in establishing their own small businesses.

3. Estimates of the relationship between duration of benefits and length of unemployment spells are not available for British data. However, evidence on this relationship for Germany's combined UI/UA program has recently been presented by Hunt (1995). Beginning in 1985, a series of statutory changes extended the duration of UI benefits for older workers with considerable work experience. Although most of these workers would have been eligible for UA assistance once their UI benefits ended, UI benefits are preferred by German workers to UA because of UI's higher replacement ratio and the fact that UA benefits are means-tested. Hunt's results for workers in the 44-48 and 49-57 age brackets indicate that extended UI benefits increase length of unemployment spells, especially for those who ultimately dropped out of the labor force. Her interpretation of these findings is that many of the unemployed who left the labor force were not really looking for work, but described themselves as unemployed while receiving UI benefits, and then dropped out of the labor force when their benefits were exhausted.

4. Horowitz (1991: A9) offers this example from the northern England port city of Hartlepool:

> All Britons have access to free medical care. An unemployed couple with two children receives about $160 a week, plus free rent, free school meals, free medicine, even "cold weather" payments if the temperature falls below 32 degrees for seven days in a row. . . . Because of the benefits structure, an unemployed parent of two in Hartlepool must find a job paying almost $300 a week to improve on the dole. The few jobs available here typically pay $150 to 250.

5. Atkinson and Micklewright (1991: 1719) suggest that the A$100 employment entry payment may be viewed as a reemployment bonus, albeit a reemployment bonus with a long qualifying period (since it is available only to the long-term unemployed) and a modest bonus payment. Underlying the reemployment bonus concept in general is the notion that a bonus payment for finding a new job quickly helps overcome the disincentive of UC claimants to engage actively in job search. For the United States, Meyer (1995) reviews evidence gathered from a series of random-assignment experiments on the effectiveness of reemployment bonuses in expediting reemployment. Estimated effects of reemployment bonus plans on weeks of UI payments appear to be about the same as those achieved, but at much lower cost, by more closely monitoring work search behavior in the Charleston and Washington state job search experiments referred to earlier in the text.

6. Green and Riddell (1994) and Lemieux and MacLeod (1994) are two of a series of studies commissioned in the spring of 1993 by Human Resource Development Canada as part of a major evaluation of the Canadian UI system. Two additional studies in this series that are mentioned later in the text are Crossley and Kuhn (1994) and Jones (1995).

7. Created in 1985, the Canadian Jobs Strategy was an umbrella program including five major component programs designed to assist the following client groups: (1) the long-term unemployed, (2) women and youth who face difficulty

in making the transition to full participation in the labor market, (3) employed workers whose job security is threatened by technological or market changes, (4) employers needing workers possessing skills in short supply, and (5) communities with chronic high unemployment or faced with major layoffs.

8. Prior to the EIP's implementation in 1991, the preferred approach was the Direct Purchase option under which workers eligible for CJS programs were assigned to program slots purchased from training institutions by the government.

9. Sorrentino (1995) describes that for many years, Swedish retraining and wage-subsidy and job-creation programs helped keep unemployment at a very low level by international standards. However, as joblessness rose to record postwar highs during the early 1990s, increased participation in labor market programs could no longer hold down unemployment as it had in earlier, milder recessions.

10. In the United States, monetary eligibility for UI benefits is primarily earnings based. Nevertheless, a recent discussion by the Advisory Council on Unemployment Compensation (1995: Chap. 7) makes it clear that for low-wage workers, eligibility depends directly on previous employment duration. In particular, a minimum-wage worker who was previously working full-time and full-year meets monetary eligibility requirements in every state. However, a minimum-wage worker previously working half-time for half of the year would be eligible for UI benefits in just 37 states. That is, low-wage, part-time workers must have worked more hours to qualify for benefits than higher-wage workers. In contrast to monetary eligibility requirements, nonmonetary requirements are intended to ensure that UI claimants are available for and actively seeking work, and are either involuntarily unemployed or voluntarily left their jobs for good cause.

11. Legislation passed in 1993 extending regular UI benefits also requires the U.S. Secretary of Labor to establish a program for encouraging the implementation of a system for profiling new UI claimants.

REFERENCES

Advisory Council on Unemployment Compensation. 1995. Unemployment Insurance in the United States: Benefits, Financing, and Coverage. Washington, D.C. (February).

Atkinson, Anthony B., and John Mickelwright. 1989. "Turning the Screw: Benefits for the Unemployed 1979-88." In Andrew Dilnot and Ian Walker, eds., The Economics of Social Security. Oxford: Oxford University Press. Pp. 17-51.

_______ and _______. 1991. "Unemployment Compensation and Labor Market Transitions: A Critical Review." Journal of Economic Literature 29 (December): 1679-1727.

Baily, Martin N., Gary Burtless, and Robert E. Litan. 1993. Growth with Equity: Economic Policymaking for the Next Century. Washington: The Brookings Institution.

Balducchi, David, Terry R. Johnson, and R. Mark Gritz. 1995. "The Function of a Labor Exchange System." Presented at the W. E. Upjohn Institute and U.S. Department of Labor conference, "Unemployment Insurance in the United States: Analysis of Policy Issues," Kalamazoo, MI, June 15-17, 1995.

Blank, Rebecca M., and David E. Card. 1991. "Recent Trends in Insured and Uninsured Unemployment: Is There an Explanation?" Quarterly Journal of Economics 106 (November): 1157-89.

Blaustein, Saul J., and Isabel Craig. 1977. An International Review of Unemployment Insurance Schemes. Kalamazoo, MI: W. E. Upjohn Institute for Employment Research.

Corak, Miles 1993. "Is Unemployment Insurance Addictive? Evidence from the Benefit Durations of Repeat Users." Industrial and labor Relations Review 47 (October): 62-72.

Corson, Walter, David Long, and Walter Nicholson. 1985. "Evaluation of the Charleston Claimant Placement and Work Test Demonstration." Unemployment Insurance Occasional Paper 85-2, U.S. Department of Labor.

Corson, Walter, Shari Dunstan, Paul Decker, and Anne Gordon. 1989. "New Jersey Unemployment Insurance Reemployment Demonstration Project." Unemployment Insurance Occasional Paper 89-3, U.S. Department of Labor.

Crossley, Tom, and Peter Kuhn. 1995. "Effects of Bill C-113 Disentitlement on UI and Welfare Take-up Rates. HRDC Evaluation Brief No. 17.

Department of Employment. 1988. Employment for the 1990s. London (December).

Departments of Social Security (DSS) and Employment, Education and Training (DEET). 1993. Helping the Long-Term Unemployed: Report of the Initial NEWSTART Strategy. Commonwealth of Australia (February).

Employment and Immigration Canada (EIC). 1989. "Success in the Works: A Policy Paper." Ottawa.

Employment Department Group. 1991. "Just the Job." England (September).

Green, David A., and W. Craig Riddell. 1993. "The Economic Effects of Unemployment Insurance in Canada: An Empirical Analysis of UI Disentitlement." Journal of Labor Economics 11 (January, Part 2): S96-S147.

_______, and _______. 1994. "Qualifying for Unemployment Insurance." HRDC Evaluation Brief No. 1.

Ham, John C., and Samuel A. Rea. 1987. "Unemployment Insurance and Male Unemployment Duration in Canada." Journal of Labor Economics 5 (July): 325-53.

Horwitz, Tony. 1991. "No Expectations: Working Class Culture Erodes Britain's Rank In a Unified Europe." Wall Street Journal (February 11): A1, A9.

Hunt, Jennifer. 1995. "The Effect of Unemployment Compensation on Unemployment Duration in Germany." Journal of Labor Economics 13 (January): 88-120.

Johnson, Terry R., and Daniel H. Klepinger. 1994. "Experimental Evidence on Unemployment Insurance Work-Search Policies." Journal of Human Resources 19 (Summer): 695-717.

Jones, Stephen. 1995. "Effects of Benefit Rate Reduction and Changes in Entitlement (Bill C-113) on Unemployment, Job Search Behaviour and New Job Quality." HRDC Evauation Brief No. 20.

Kane, Thomas J., and Cecilia E. Rouse. 1995. "Labor-Market Returns to Two- and Four-Year College." American Economic Review 85 (June): 600-14.

Katz, Lawrence, and Bruce Meyer. 1990. "The Impact of the Potential Duration of Unemployment Benefits on the Duration of Unemployment." Journal of Public Economics 41 (February): 45-72.

Lancaster, Tony, and Stephen Nickell. 1980. "The Analysis of Re-employment Probabilities." Journal of the Royal Statistical Society 143, series A: 141-65.

Leigh, Duane E. 1995. Assisting Workers Displaced by Structural Change: An International Perspective. Kalamazoo, MI: W. E. Upjohn Institute for Employment Research.

Lemieux, Thomas, and Bentley MacLeod. 1994. "Learning Effects and Unemployment Insurance." HRDC Evaluation Brief No. 4.

Meyer, Bruce D. 1995. "Lessons from the U.S. Unemployment Insurance Experiments." Journal of Economic Literature 33 (March): 91-131.

Murray, Charles. 1990. "The British Underclass." The Public Interest (Spring): 4-28.

Narendranathan, W., Stephen Nickell, and J. Stern. 1985. "Unemployment Benefits Revisited." Economic Journal 95 (June): 307-29.

Organization for Economic Cooperation and Development (OECD). 1989. OECD Employment Outlook. Paris (July).

_______. 1990. Labour Market Policies for the 1990s. Paris.

_______. 1992. OECD Employment Outlook. Paris.

_______. 1993. The Public Employment Service In Japan, Norway, Spain and the United Kingdom. Paris.

_______. 1994. OECD Economic Surveys: Canada. Paris.

Park, Norman, W. Craig Riddell, and Robert Power. 1993. "An Evaluation of UI-Sponsored Training." Employment and Immigration Canada (August).

Reubens, Beatrice G. 1989. "Unemployment Insurance in the United States and Europe, 1973-83." Monthly Labor Review 112 (April): 22-31.

Runner, Diana. 1995. "Changes in Unemployment Insurance Legislation in 1994." Monthly Labor Review 118 (January): 60-64.

Sakkara, Mary-Anne, Stephen Beyers, Karen Buchanan, Philip Gatenby, and Kathy Patston. 1994. Newstart Evaluation: Interim Report on Effectiveness. Canberra: Australian Government Publishing Service.

Sorrentino, Constance. 1995. "International Unemployment Indicators, 1983-93." Monthly Labor Review 118 (August): 24-30.

Storey, James R., and Jennifer Neisner. 1992. "Unemployment Compensation in the Group of Seven Nations: An International Comparison." Report prepared by the Congressional Research Service at the request of the Committees on Finance, U.S. Senate, and Ways and Means, U.S. House of Representatives. Washington: U.S. Government Printing Office.

U.S. Department of Labor (USDOL). 1993. "The Changing Labor Market and the Need for a Reemployment Response." Washington, D.C. (December).

Wall Street Journal. 1993. "Day of Reckoning: After Years of Bliss, Canada Faces Trouble from Excess Spending." (August 27): A1, A8.

Table 1. Trends in Long-Term Unemployment in the U.S.

Time period	Average unemployment	Long-term unemployment as % of total unemployment[a]
1950s	4.5%	9.4
1960s	4.8	10.5
1970s	6.2	11.0
1980s	7.3	15.0
1990-93[b]	6.6	16.0

Source: USDOL (1993: Table 1).

[a]Long-term unemployment defined as 6 months or longer.

[b]Data through September 1993.

Table 2. Unemployment Compensation Programs in Selected English-Speaking Countries

Country	Type of program	Duration of benefits	Fuller description
U.K.	UI and UA	indefinite	UI for up to 12 months (duration depending on recent record of employment). Possibility of UA instead of, during, or after UI.
Australia	UA only	indefinite	Eligible workers must be available for work and actively seeking employment. Benefits are paid for as long as an individual is qualified.
Canada	UI only	50 weeks	Up to 50 weeks depending on regional unemployment rate and employment history.
U.S.	UI only	6 months	Up to 6 months of regular UI benefits. 3 additional months of extended benefits sometimes available in high unemployment states.

Source: U.K., Canada, and U.S.: Atkinson and Micklewright (1991: Tables 2 and 3); Australia: Storey and Neisner (1992: 31).

Table 3. Unemployment Rates and Incidence of Long-Term Unemployment, Selected English-Speaking Countries

Country	Standardized unemployment rate, 1985-88	Long-term unemployment, 1987[a]	
		6+ months	12+ months
U.K.	10.2%	62.8	45.2
Canada	9.1	24.1	9.4
Australia	7.9	48.7	28.7
U.S.	6.4	14.0	8.1

Source: OECD (1989: Tables 1.5 and M).

[a]As a percentage of total unemployment.

Table 4. Schedule of British Employment Service Contacts in the Scheduling and Programming Strategy, by Duration of Unemployment

Duration of unemployment	Employment Service Action
Day 1	Basic check of entitlement; benefit forms issued.
First week	Benefit claim taken; explanation of the "benefit contract" (i.e., claimant's legal rights and obligations); agreement on a back-to-work plan.
Week 13	Review of the back-to-work plan; claimants may attend a 2-day course on job search techniques or a 2-day workshop on alternative careers.
Week 26	First Restart interview: review benefit contract; develop new back-to-work plan; access to reemployment programs including Jobclubs, Employment Training, and Job Interview Guarantee.
Week 52	Second Restart interview: same actions as week 26.
Week 78	Third Restart interview: same actions as week 26, plus warning that attendance at a Restart Course could be mandatory.
Week 104	2-year Restart interview: same actions as week 26, plus mandatory attendance at a 1-week Restart Course. Involves overcoming perceived barriers to employment, assessment of skills and experience, enhancement of job search skills, and designing a plan of action leading back to work.

Source: Employment Department Group (1991) and OECD (1993: Table 19).

Table 5. Schedule of Department of Social Security and Commonwealth Employment Service Interventions in the Australian Newstart Program, by Duration of Unemployment

Duration of unemployment	DSS and CES Actions
First week	Registration for the Job Search Allowance with the DSS; explanation of "activity testing" requirements by the CES.
Three months	First Job Search Allowance interview. Entitlement review carried out by the DSS with referal to the CES for further reemployment assistance.
Six months	Second Job Search Allowance interview. CES assesses claimant's employment prospects and advises on eligibility for labor market programs.
12 months	Interview and application for the Newstart Allowance; agreement between claimant and the CES on a Newstart Plan and Activity Agreement.
Between 12 and 36 months	Regular reviews by the CES of the Activity Agreement (on average every 6 months), plus DSS entitlement reviews every 12 months.
36 months	All Newstart Allowance recipients interviewed jointly by the CES and DSS to reassess clients' job prospects, review entitlement, and determine if some other forms of assistance (including alternate income support) are appropriate.

Source: Sakkara et al. (1994: 16-21).

Table 6. Public Expenditures on Active and Passive Labor Market Programs as a Percentage of GDP, Selected Countries, Annual Averages for 1985-88

Country	Active[a]	Passive[b]	Total
U.K.	0.80	1.89	2.69
Australia	0.36	1.20	1.56
Canada	0.58	1.74	2.32
U.S.[c]	0.26	0.51	0.77
Sweden	1.96	0.80	2.76

Source: OECD (1989: Table A.1)

[a] Includes employment services and administration, adult labor market training, special youth measures, direct job creation and employment subsidies, and measures for the disabled.

[b] Includes unemployment compensation and early retirement for labor market reasons.

[c] For 1986-88.

Table 7. Characteristics of the UI Systems in Canada and the U.S., 1988

Characteristic	Canada	U.S.
Weeks of UI/weeks of unemployment	99.8%	27.0%
Average weekly UI amount/average weekly earnings[a]	43.7%	34.8%
Average duration in weeks of UI claims	17.9	13.7

Source: Green and Riddell (1993: Table 5).

[a]Insured workers only for the U.S.

Table 8. Impact of Canadian UI-Sponsored Training Programs, 1991 Earnings, by Cohort of UI Recipients and Pre-Training Base Year

Training program	1988 cohort		1989 cohort	
	1986	1987	1987	1988
Feepayer	4816***	3494***	-20	906
DIR	1981	2181	-1018	1013
Job Development	1531	2221*	-215	1291
Job Entry	4461***	4935***	4054**	6296***
Skill Shortages	6188***	5429***	1965	1784

Source: Park, Riddell, and Power (1993: Table 5.12).

Note: Estimates are based on comparisons of average earnings.

***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

The Sensitivity of Unemployment Compensation Expenditures to Variations in the Key Program Parameters: Evidence From Five OECD Countries

Pravin K. Trivedi*
Department of Economics
Indiana University
Bloomington, IN 47405, U.S.A.

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Contents

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[1]For a detailed description and analysis of the US system see *Advisory Council on Unemployment Compensation, Report and Recommendations,* February 1994.

Abstract

This paper presents a framework for evaluating the role of policy changes in the evolution of unemployment compensation (UC) expenditures. Within this framework UC expenditures are related to average weekly (or annual) benefits, initial or first time payments and average duration on unemployment compensation. These three variables in turn are modelled in terms factors which include policy variables that reflect changes in the scope, coverage and generosity of the UC system in each country. The modelling framework allows one to study the channels through which UC policy works. The framework is implemented econometrically using annual data for the 1960-93 period from five O.E.C.D. countries. The emphasis is on medium and long term cointegrating relations rather than short term dynamics. The econometric analysis helps to identify the reasons why the aggregate sensitivity of UC expenditures to cyclical variations in unemployment differs significantly in the five countries. Estimates are provided of the impact of past policy changes on the inflows into UC benefit pool and on the average of duration of compensation.

1. Introduction

Unemployment compensation (UC) has been studied from many different aspects - financing methods, impact on labor force participation and labor supply, and relationship to other income maintenance and labor market programs. The focus of the present paper is on the budgetary cost of UC systems which has been an important policy issue in many OECD countries due to high and persistent unemployment during the 1980s. See Figures 1.1 to 1.5. The policy issues involve not only the economics of passive versus active labor market policies, but also an evaluation of the net budgetary impact of numerous piecemeal changes to the coverage, generosity and eligibility of various UC programs.

This paper uses aggregate time series analysis to model the cost of UC in five OECD countries - the United States, Canada, the United Kingdom, Germany and France. This choice was motivated by the following considerations. While the UC expenditures are strongly related to the level of unemployment, their growth has been strongly influenced by policy decisions of the 1960s and 1970s, when the UC systems in Canada and several European countries were greatly expanded in terms of their scope and generosity by measures intended to expand coverage and to protect the real value of the benefits in the face of inflation that followed the oil price shocks of the 1970s. Later policy changes, for example, in the mid- and late 1970s in Canada, and the early 1980s in the United Kingdom, attempted to curtail or reverse the earlier changes, and, in the case of some countries, to modify them significantly, in order to control the budgetary costs of the UC system. In the United States, where the UC system is State based, the changes were more piecemeal and less long term, compared with Canada and Europe. The time series data therefore provide an opportunity to model and to compare the impact of policy changes on UC systems in different institutional environments.

Cost considerations have a key role in determining the feasibility of changes to an existing program of unemployment compensation. Reubens (1989) provided a comparative study of unemployment insurance in the United States and five European countries based on 1973-83 data, and attempted to highlight factors that contributed to the high and rising costs of UC expenditures. Her study did not attempt to econometrically model the data. It is useful to know how the different components of cost contribute to the total cost within the extant program, and how they themselves respond to changes in policy parameters. This requires a quantitative model. It is clear, especially from the experience of the recent recessions, that the cost of UC programs has increased sharply in many countries; yet the role different contributory factors to this cost blow-out at the aggregate level is not clearly understood. This consideration motivates the construction of a small econometric models for describing the evolution of different components of aggregate unemployment compensation. The objective of this paper is to develop and apply a simple econometric method of making international comparisons of the factors that contribute to the expenditure on UC programs. Such country specific econometric models would help summarize the major features of different UC systems in terms of certain key parameters.

Cross country comparisons of UC programs are complicated by differences in the mix of income maintenance and unemployment insurance elements, and by the way in which UC systems integrate with coexisting income maintenance programs. The US and the Canadian UC systems are based on the insurance

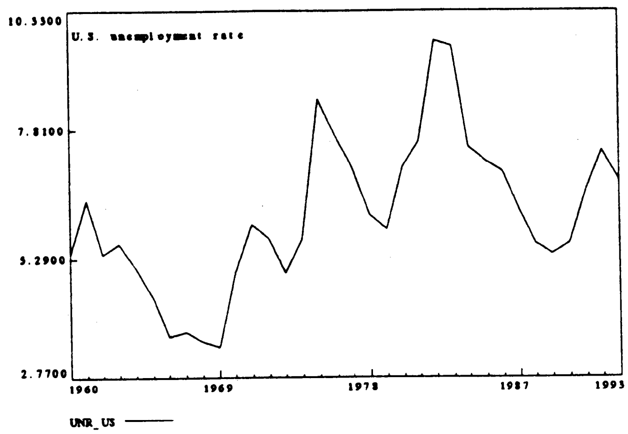

Figure 1.1: Unemployment Rate - USA

principle, and have significant regional variation, whereas the three European systems have had insurance and income maintenance and social security elements simultaneously present, or integrated to varying degrees. These features preclude a mechanical extrapolation of the experience of one country or time period to different economic environments. However, it seems possible to use econometric models as a framework for measuring the sensitivity of the direct cost of the UC program to movements in unemployment and to policy shifts.

The paper is organized as follows. Section 2 provides an overview of the UC systems in the five countries, with emphasis on changes whose impact I wish to measure. Section 3 lays out the econometric methodology and the framework which are empirically implemented in Section 4. Section 5 summarizes the main findings and concludes with qualifying remarks.

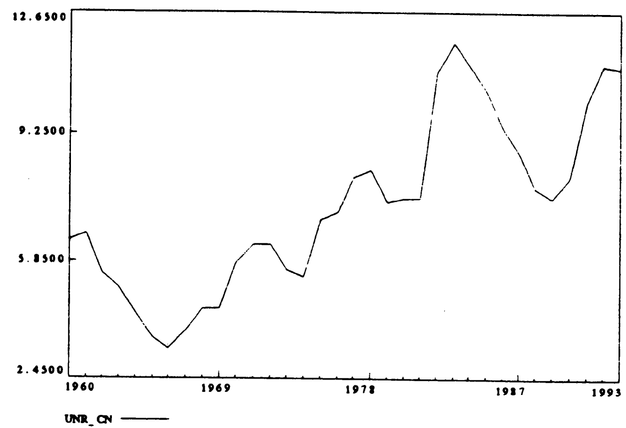

Figure 1.2: Unemployment Rate - Canada

2. The Unemployment Insurance and Assistance Systems in Five OECD Countries

A brief comparison between the UC systems of the five countries is the starting point. The US system operates under State rules and hence varies more than the remaining four. To finance the cost of benefits payroll taxes are used in all five countries. In the United Kingdom, France and Germany, unemployment insurance (UI) is supplemented by means (income) tested unemployment assistance (UA), the latter being financed by general revenues. UI is a contributory system, financed usually be payroll taxes, with benefit related to the past work and earnings history, and providing benefits for a finite duration and not related to

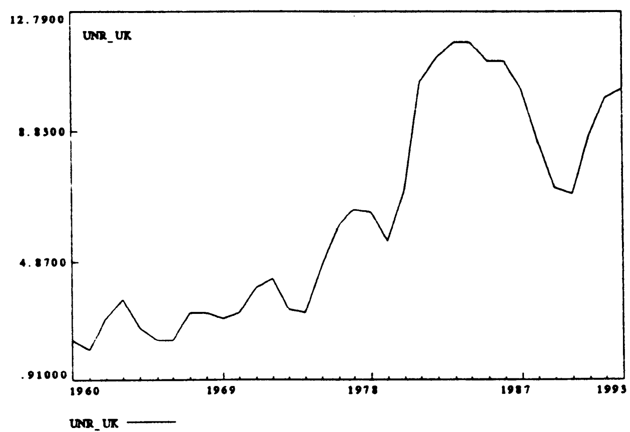

Figure 1.3: Unemployment Rate - UK

family circumstances. UA is non-contributory, with means tested benefits usually independent of employment history and extending over an indefinite period. The UI system in the United States is experience rated on employers. Finally neither the United States nor Canada have means tested UA programs. Tables 1 and 2 provide a comparison of the benefit formulas in two years, 1987 and 1977.[2]

Most countries supplement the UI spending by a variety of other (often means

[2] An excellent example of an attempt at an international comparison of unemployment compensation in the Group of Seven Nations is contained in a 1992 report prepared by the Congressional Research Service. Table 3 (p. 21) and Table 4 (p. 24) of the report provide information on the determinants of unemployment benefit amounts including such "parameters" as Age, Work history, Wage history, Maximum benefit duration and so forth; such parameters define the operating characteristics of a given unemployment benefit program. Appendix B provides a tabular summary of the major events in the development of unemployment compensation. Appendix C provides the data on the public expenditures for unemployment compensation programs, expressed as a proportion of GDP.

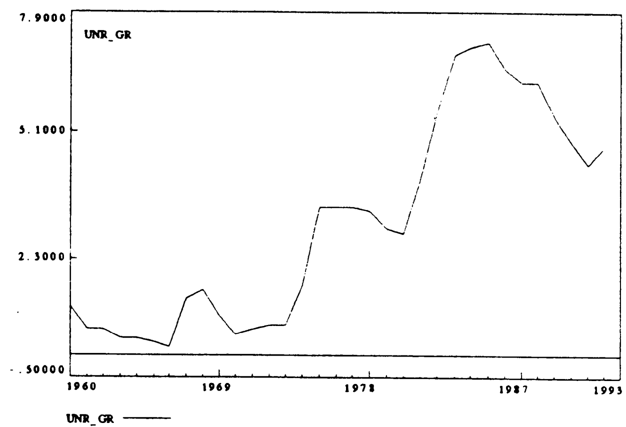

Figure 1.4: Unemployment Rate - Germany

tested) income maintenance, assistance and/or adjustment programs (UA). An immediate consequence is that to make international comparisons of the overall expenditure on UC, some specific indicator is required. A commonly used measure is the ratio of all UC expenditures, UI plus UA, to GDP.[3] The first row for each country in Table 5 shows these figures for selected fiscal years during the period 1970-1990. Canada has the highest percentage figure and the United States has the lowest. The variations across time reflects both unemployment experience and discretionary policy changes. Since the time path of unemployment differs between countries, a crude adjustment may be made by dividing the percentages in Table 5 by the average standardized unemployment rate for that year. The adjusted figure is shown in row 2 for each country. Again Canada and Germany

[3]See *OECD Employment Outlook*, 1992.

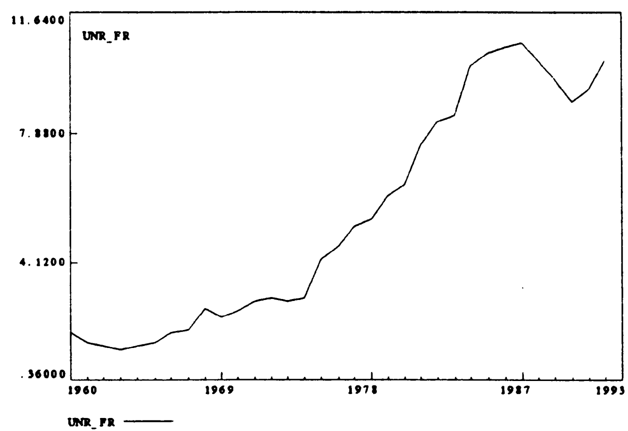

Figure 1.5: Unemployment Rate - France

have the highest average adjusted expenditure and the United States has the lowest. In all countries, and especially in Canada and Germany, the expenditure levels have declined from the high levels of the 1975-80 period.

Figures 2.1 to 2.7 show either the ratio of total UC (UI plus UA) to GDP, or just UI to GDP, the distinction being important for the European countries.

2.1. The Basic System and the Discretionary Changes in UC Policy

The UC systems tend to be complex. In the United States this is partly a consequence of the State basis of the system. In Canada also there is significant regional variation. In all five countries examined here this is in part because of the coexistence of a multiplicity of income assistance programs and because of the large number of parameters that affect the eligibility, duration and level of

UC. Tables 3 and 4 provide comparative information on the generosity of the UC systems in 1989 and 1979, respectively. A comparison of the two tables indicates several important changes that have occurred during the 1980s.[4] The main focus of this paper is on the changes to the UC system in each country and their impact on the components of UC expenditure. We begin with a summary of changes to UC system in each country.[5]

2.1.1. United States[6]

The UC system was designed to compensate for job loss due to normal business cycle fluctuations. It is State based; the benefits are financed by payroll taxes levied by the States on employers only. The States determine eligibility and level of benefits which vary considerably. Since 1986 the benefits have been taxable. There is no means tested benefit integrated with UC. Table 3 provides information on the representative eligibility and benefit rates. Unfortunately, there is a loss of information when no indicators are given to reflect the substantial variation across the States.

The maximum duration of the basic UI compensation is 26 weeks in most states.[7] Since 1970 extended benefits (EB) program has operated in .... states. It is triggered by the State's insured unemployment rate over 13 weeks exceeding 6 per cent. It provides additional benefits for up to 13 weeks. EB has been supplanted by a temporary emergency UC program funded entirely from Federal

[4]See "Unemplyment Benefit Rules and Labor Market Policy", chapter 7 in *OECD Employment Outlook,* July 1991.

[5]See Appendix A in the Congressional Research Service 1992 report for additional details.

[6]For a detailed description and analysis of the US system see *Advisory Council on Unemployment Compensation, Report and Recommendations,* February 1994.

[7]It is 30 weeks in 2 states and 18 weeks or less in two other states (Alabama and Washington).

payroll taxes. The duration of emergency benefits has varied from one program to the next, some times being as little as 13 weeks and on other occasions it has provided 26-33 weeks of added benefits, depending on the State unemployment rate. Finally the Trade Adjustment Assistance (TAA) program exists to help dislocated workers faced with long term unemployment who need to make a career transition by retraining. TAA benefits are paid after the basic UI/EB benefits expire. The total combined duration of all benefits is 52-78 weeks.

2.1.2. Canada

The UI system is financed by payroll taxes on employers and employees and covers all wage and salary earners except those providing less than 15 hours of work per week and earning less than C$136 per week. There is no waiting period for workers satisfying eligibility requirement. The duration of the benefit varies according to the number of insurable weeks of work and the regional unemployment rate, and may reach up to 50 weeks.

The sample period covers some major changes to the system. *Unemployment Insurance Act of 1971* was a major reform since it extended coverage and relaxed eligibility criteria, and introduced distinctions in the treatment of claimants with major and minor labor force attachments (Green and Riddell (1993)). The labor force covered by UI has risen from 61% in 1966 to 90% in 1975. This led to a very sharp increase in total UC expenditures which more than doubled between 1971 and 1977. During the period 1976-1979, there was a partial reversal of the liberalization of 1971-72. The elderly were disentitled from UI due to maximum age for UI entitlement being reduced from 70 to 65, the maximum disqualification period was increased from 3 to 6 weeks, and the benefit rates for claimants with dependents and low income were reduced. Between 1977-79 eligibility conditions were tightened, regional differentiation of the UI program was increased. The period of the 1980s was one of relative stability. Further changes in the generosity of the program occurred in 1990-92 after the passage of Bill C-21 in 1990 which reduced benefits in most regions and increased eligibility requirements.[8]

2.1.3. United Kingdom

The British UC system, like those in Germany and France, has UI and income tested UA components. Full UI benefits are available to those satisfying the minimum threshold for taxable earnings, with a reduced rate for those failing to do so. Micklewright (1989) has analyzed the consequences of the earnings threshold. Maximum benefit period for UI, after a 3-day waiting period is 52 weeks, but the period on UA for those qualified for it is indefinite. UI benefits are financed by payroll taxes, and the UA component is financed through general revenues.

The UK system has seen several major changes beginning with the introduction in 1966 of the Earnings Related Supplement (ERS) to supplement a flat rate UI benefit. This measure was aimed to produce a UA program with a guaranteed standard of living. Beginning 1980, and especially after 1982, there has been a significant tightening up of eligibility for UI. In 1982 ERS was abolished and benefits were taxed. The Social Security Act 1988 tightened contribution conditions for the receipt of UI. Those who failed to meet eligibility conditions could qualify for UA. This represented a significant shift away from UI to means tested UA. In

[8] Moorthy (1989) has argued that the generosity of the Canadian UI system partly accounts for the unemployment rate differential between Canada and the USA. Card and Riddell (1993) provide detailed comparisons between the generosity of the UI system in the USA and Canada.

1988 Supplementary Benefits were replaced by Income Support.

Atkinson and Micklewright (1988) have listed changes in the UC system according to whether they increased or decreased the generosity of the UC system. It is clear that as a consequence of the changes during the 1980s the proportion of the unemployed receiving UI has declined, with more unemployed individuals receiving the means- tested UA.

Table 6 shows the proportions of registered unemployed receiving UI and UA in Britain and Germany in 1988. This proportion is relatively low and is of the same order as the recipiency rate for UI in the United States (Blank and Card (1991), ACUC (1994)). Among the reasons for nonreceipt of UI in Britain and Germany are benefit exhaustion after a long spell, insufficient contribution history and waiting period. The importance of these factors can be expected to vary over a business cycle.

2.1.4. Germany

The German UC system covers all employees with earnings subject to social security tax. UI (*Arbeitlosengeld*) benefits are paid to unemployed with insured employment history of 360 days during the past 3 years. Benefits, which are not subject to income tax, are payable without a waiting period for a duration that varies with the length of covered work history and age. Maximum duration is 52 weeks. The means tested UA benefits (*Arbeitlosenhilfe*) are paid to the ineligible unemployed workers for another 52 weeks, and renewable for further periods. The rules differ somewhat for seasonal and part-time workers.

There have been few major changes to the German system. The eligibility criteria were made more stringent in 1977, but made less stringent in 1987 for some classes of benefit recipients, e.g. older workers with long and continuous contribution records. The UC system was also established for the former GDR during 1990-92 (which is outside the sample period for Germany used in regression analysis).

2.1.5. France

Before 1969 the UI system was private and contributory and received no government assistance. In 1979 a unified UC system with support from government revenues was proposed. This largely failed and was replaced in 1984 by the present system which has two components, UI and means-tested UA. Workers under age 60, excluding seasonal workers, are covered by UI. The "solidarity" UA program provides income support for UI benefit exhaustees and flat-rate one year benefit for certain categories of new labor force entrants.

Coverage of UI has increased from around 40% in the early 1970s to about 75% in mid-1980s. Finance is through payroll taxes on employers and employees. UI benefit has a fixed component determined by past work history, and a variable component which is related to previous earnings. There is no waiting period for benefits.

2.2. An overview

A chronology of important policy changes is given in Appendix A. It is difficult to give an accurate but brief description of the net effect of the policy changes on the benefit recipients. Table 7 summarizes these in terms of the percentage change in the real value of average benefits.

These results show that over the sample period there has been a negligible

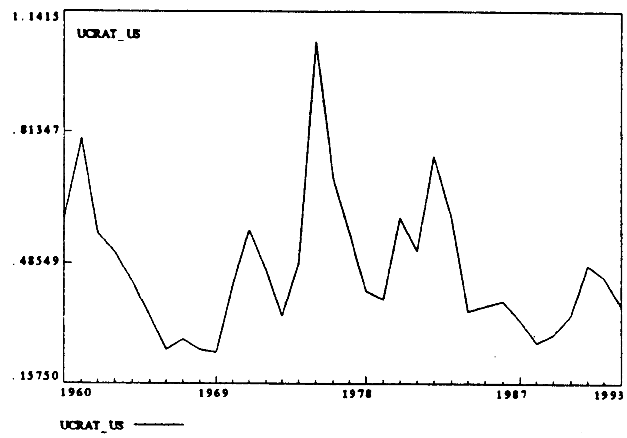

Figure 2.1: Ratio of total UC to GDP - USA

change in the real UC compensation in the United States. Over the period 1971-90, the United States registered a *negative* cumulated growth of around 4%; at the opposite end of the spectrum was Canada with cumulated growth of over 50%. Of the three European countries, the United Kingdom has the smallest growth (25%), while Germany and France have about 36% and 43% respectively. In all cases, the growth rate of the UC systems slowed during the 1980s. In the European countries there has been an attempt to strengthen the insurance element of the UI system, leaving the UA system to cope with issues of income maintenance. See Figures 2.1-2.7 which show either the UC (UI plus UA) or the UI expenditures as a proportion of GDP in each country.

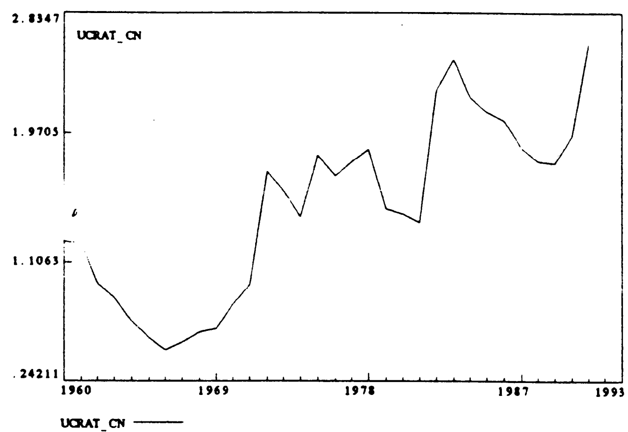

Figure 2.2: Ratio of Total UC to GDP - Canada

3. Econometric Methodology

The direct cost of a UC program depends upon the coverage, eligibility rules and benefit levels; the level, composition and duration of unemployment; the size of the labor force and coverage of the UC program; and the relative importance of the income maintenance and unemployment insurance elements in the program.

The following stylized identity isolates the key components of unemployment compensation:

$$UC_t = AWB_t \times NC_t \times AD_t \tag{3.1}$$

where UC is the total annual unemployment compensation, AWB is the actual average weekly UI or UA benefit, NC is the average weekly number of benefit recipients (sometimes referred to as the number of first time payments, and to be

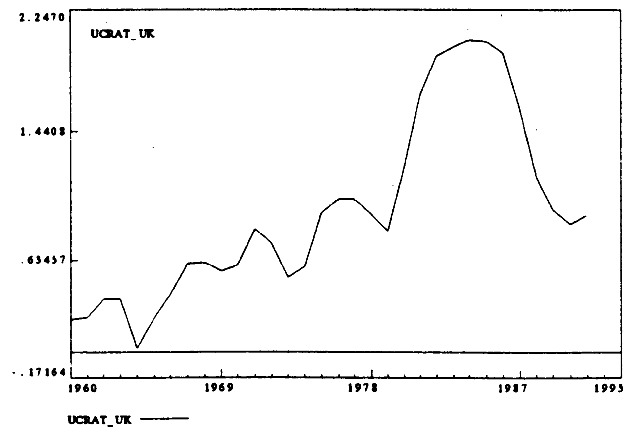

Figure 2.3: Ratio of Total UC to GDP - UK

distinguished from the stock of benefit recipients at a point in time), and AD is average duration (in weeks) of a spell on unemployment compensation.[9]

An alternative identity is

$$UC_t = UCPBR_t \times NBR_t \tag{3.2}$$

where $UCPBR$ is the annual UC per benefit recipient and NBR is the average annual number of benefit recipients. This identity does not use information on average weekly flows and average duration and hence may be useful when the latter information is not available. In practice the information on $UCPBR$ has to be derived by combining accounting information on $UCPBR$ and occasionally

[9] In practice this is calculated as the total number of weeks compensated (NWC) divided by the average duration of unemployment compensation (AD).

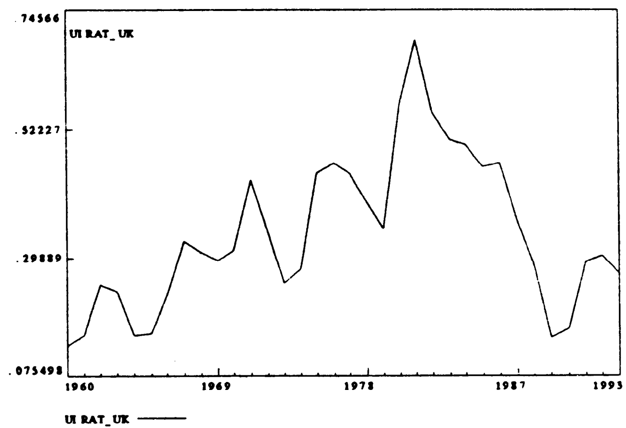

Figure 2.4: Ratio of UI Expenditure to GDP - UK

collected information on NBR. For example the information on NBR may be based on a count in one week or month rather than an appropriately calculated average.

The goal of the paper is to develop and estimate econometric specifications for all components on the right hand side in terms of its determinants, especially including the policy variables. For the above identity to hold exactly, the UC system should be homogeneous, and the data should be constructed from a consistent set of accounting conventions. In practice there is more than one class of benefit recipient and several classes of payments, both within UI and UA categories. Further the administrative records may not be held on a consistent basis.[10]

[10]For example, data may only permit an econometric analysis based on claims figures for one week in the year rather than the average for the year. The published AWB may be calculated

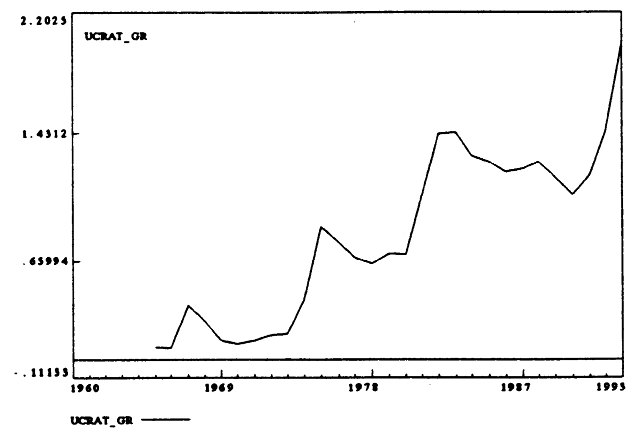

Figure 2.5: Ratio of Total UC to GDP - Germany

In such cases the above identity may be replaced by a quasi-identity with an error term (to account for measurement errors):

$$UC_t = \alpha_0 AWB_t^{\alpha_1} NC_t^{\alpha_2} AD_t^{\alpha_3} \varepsilon_t \tag{3.3}$$

The parameters $\alpha_0, \alpha_1, \alpha_2, \alpha_3$ measure the elasticity of UC with respect to AWB, NC, and AD. Their deviation from unity may be interpreted as a measure of the extent of inconsistency in the data. The above equation may be interpreted as an equilibrium (long-term) equation. Given delays in processing of the claims for UC, accounting conventions for treating and classifying claims, and measurement errors, the above relation will not hold exactly at any point in time. As a consequence all four parameters estimated from actual data can be expected to be

for different classes of payments while the data for NC or AD may be aggregated.

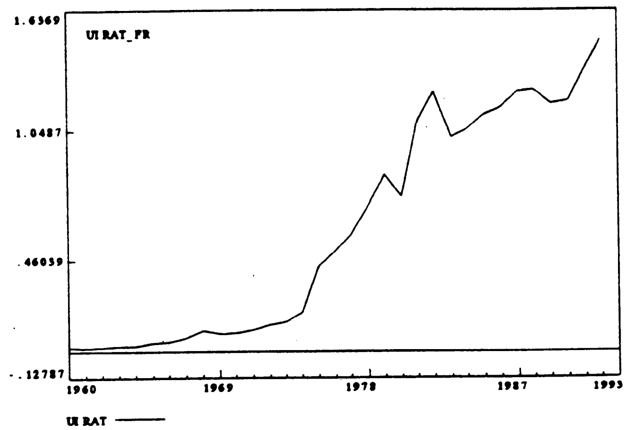

Figure 2.6: Ratio of UI Expenditure to GDP - Germany

significantly different from unity.

This paper will attempt to model the time series behavior of the three components on the right hand side of the above identity. In choosing an appropriate econometric methodology for doing so, it is relevant to consider the properties of these time series. Since UC_t and AWB_t will be measured in nominal units of national currency, they will be ***nonstationary*** time series, more precisely, I(1) (integrated to order 1) time series. The variable NC_t will also be an I(1) series since, with the secular growth of the labor force, one can expect that the number of unemployed persons to rise in a trend like fashion. Finally, AD_t is likely to have a *stationary* (I(0)) time series representations because, firstly, it will be subject to an upper bound due to the restriction on maximum weeks of benefit entitlement, and secondly, because a priori we expect the duration of payment

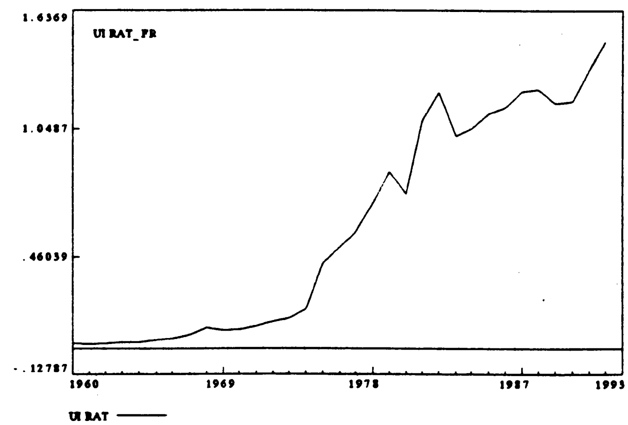

Figure 2.7: Ratio of UI Expenditures to GDP - France

over long periods of time to vary largely with the business cycle. In choosing a suitable estimator for the AWB and NC equation the I(1) property will affect the choice of the econometric estimator.

3.1. A Preliminary Decomposition of the UC Expenditure Growth

I begin by using the identity (3.1) to decompose the growth in UC expenditures by periods. Allowing for data inconsistencies and measurement errors, (3.1) implies the following growth rate equation:

$$\Delta \log\ UC_t = \Delta \log\ AWB_t + \Delta \log\ NC_t + \Delta \log\ AD_t + \ error$$

which indicates the contributions of the three factors to the percentage growth in UC expenditures. The first component will be generally positive as AWB in nominal terms almost always increase over time. The other two components will have strong cyclical elements and may be positive or negative over particular periods. If the UC expenditure growth is dominated by cyclical variations, one expects the second and third components to account for the bulk of the movements. During periods in which benefit eligibility is liberalized and/or the insurance coverage is expanding the first two components will contribute more to the total growth.

Table 8 shows the breakdown for a selected number of four year periods for the three countries for which the data will support such a calculation. The contrast between the United States on the one hand and Canada and Germany on the other is very marked. Though the growth of AWB is a significant factor in all three countries, it is orders of magnitude smaller in the United States. In all three cases, and especially in the United States, inflows and average benefits appear to be more important than average duration of compensation in accounting for the overall growth of UC. Over specific periods such as 1980-83, the duration component was substantial in all three countries. Finally, the contribution of inflows has been substantially *negative* during the 1980s in the United States, whereas in Canada and Germany this component also shrank over this period.

3.2. A Model for Benefits (AWB)

The hypothesis that AWB and AWE will be cointegrated is the starting point; that is, in the long run AWB will move in line with the average weekly earnings (AWE) according to the log-linear equation:

$$\log\ AWB_t = \beta_0 + \beta_1 \log\ AWE_t. \tag{3.4}$$

This hypothesis implies that the earnings replacement ratio moves around a long run (equilibrium) value. Because of the focus on the long term, I shall neglect lags in this relationship even though the statistical fit of the regression equation may be improved by including variables with short lags. The presence of lagged AWE will not change the implied long run hypothesis.

The impact of UC policy changes will enter through the parameters (β_0, β_1). Some policy interventions may be interpreted as attempts to change the equilibrium earnings replacement rate, and hence may be modelled by specifying that (β_0, β_1) depend upon such policy variables. A relatively generous UI or UA policy will have a higher intercept and a larger slope. For a relatively 'neutral' policy setting we should expect $\beta_1 = 1$. Interpreting $\frac{AWB}{AWE}$ as the average earnings replacement ratio, β_1 equal to (greater than) 1 implies a constant (growing) earnings replacement rate.

Policy settings which affect the parameters of the UI system also affect the long run by either changing the replacement rate. For example, changes in UI or UA that relate benefits to recent earnings will work through changes in the slope parameter β_1, whereas changes that affect eligibility, waiting periods or duration of benefits will work through changes in the intercept β_0. Temporary changes can be regarded as finite life shocks to the intercept, and could enter the equation as dummy variables. Assuming one shift dummy variable of each type, say Z_{1t} and Z_{2t}, yields the regression equation

$$\log\ AWB_t = \beta_0 + \delta_1 Z_{1t} + \beta_1 \log\ AWE_t + \beta_1 \delta_2 Z_{2t} \log\ AWE_t. \qquad (3.5)$$

This equation is consistent with a changing long run replacement ratio.

Policy changes may be either of one-shot type, such as the US Temporary Emergency Benefit Program, or lasting longer such as permanent changes in eligibility conditions. In principle, one should expect such policies to affect both the intercept and the slope. In practice, a sample period that is long in the relevant sense is required to estimate the policy induced changes in the parameters.

The single equation least squares (OLS) estimator of a cointegrating relationship is known to be biased in small samples. Many estimators have been proposed. In this paper the fully modified Phillips-Hansen (FMPH) estimator will be used where it is feasible. (Phillips and Hansen (1990), Banerjee et al (1993, chapter 7)). All equations were estimated by OLS as well as the fully modified Phillips-Hansen (FMPH) method with Bartlett weights. In small samples the FMPH estimator may not yield theoretically valid estimates of the variances. In those cases where the FMPH procedure did not produce theoretically acceptable estimates of variances, the OLS estimator was used.

3.2.1. Results

Variants of (3.4) were estimated for the United States and for Canada. For the three European countries consistent time series on AWB could not be constructed, so the variable $UIPBR$ was used in stead of AWB, where $UIPBR$ stands for UI benefits per recipient, calculated as the ratio of total annual UI expenditure to the average number of benefit recipients during the year. A similar measure can also be constructed for UA beneficiaries.

Intercept shift dummies seemed most appropriate for the United States, France, and Germany. Slope and intercept dummies were considered in the case of United Kingdom and Canada. The results are given in Table 9 and are discussed below.

- The coefficient of log AWE, which reflects the variation in the nominal earnings replacement rate, exceeds unity in four cases and has the largest value (1.3) for Canada; and the smallest for the United Kingdom (0.7-0.8).

Although the policy changes in the late 1970s made the conditions for the receipt of UI in Canada relatively more stringent, the benefit itself rose on average faster than did nominal average earnings. Only in the case of the United Kingdom is there evidence that the policy changes of the 1980s led to a reduction (of about 12%) in the coefficient of log AWE.

- There is evidence that UI policy changes had significant impact on AWB. The liberalization of the Canadian UI system in 1971, the TAA legislation in the United States, and the introduction of the supplementary benefits (SB) in the United Kingdom had raised the AWB significantly. The abolition of SB and the earnings related supplement (ERS) in the United Kingdom, the application of a generally more stringent UI benefit policy after 1982, and the introduction of Income Support in 1988, and the not only lowered the intercept in the AWB equation but also lowered its slope in relation to AWE.

- There is only weak evidence for the effect of other policy measures, such as the tightening up of eligibility conditions in the United States (1981-82), in the U.K. (1988), Germany (1969 and 1980) and France (1984-85), and the disentitlement of the elderly in Canada (1976-79). This is not entirely surprising because such policies were probably aimed at restricting the inflow into the pool of UI benefit recipients, rather than towards reducing the level of average benefits.

3.3. A Model for First Payments (NC)

The equation for the average flow of benefit recipients is as follows:

$$\log\ NC = \gamma_0 + \gamma_1(\log\ IUN) + \gamma_2 GINDEX + \gamma_3 Z + error, \qquad (3.6)$$

where IUN denotes the number of covered unemployed, $GINDEX$ denotes an index of UI system generosity (proxied by the ratio $\frac{AWB}{AWE}$) and Z denotes the set of policy interventions. It is expected that IUN is the most important variable in this relationship. The distinction between covered and total unemployment is important if the coverage ratio is changing through time. If the actual unemployment is used in place of covered unemployment the coefficient will reflect the effects of recipiency rate for UI. For the United States, the data on IUN are easily available and were used, but for the remaining four countries data on *actual* unemployment was used, as a proxy.

The coefficient γ_1 measures the sensitivity of inflows to changes in the level of unemployment. This depends upon the generosity of the UI system in the sense that short waiting periods and easy eligibility cause the coefficient to be larger. Following earlier reasoning the coefficient γ_1 may be treated as a function of policy variables Z. This approach is relatively more promising for the United Kingdom where the important policy changes occurred in the early 1980s which leaves enough observations for reliable differentiation of the effect of the policy.

The coefficient γ_1 also depends upon the behavior of the insured unemployed. A lower recipiency (claim) rate for benefits will cause the coefficient to be smaller. To see this note that $\frac{\partial NC}{\partial IUN} = \frac{\partial NC}{\partial RECRAT} \cdot \frac{\partial RECRAT}{\partial IUN}$. The first term on the right is positive. It is known that RECRAT has been falling in the United States (Blank and Card (1991), ACUC (1994)), while the coverage rate (IUN) has been rising, which implies that there is a negative component in the response. To capture this effect I specify the coefficient γ_1 to be linear in $RECRAT$. When substituted into the equation (3.5) this results in an additional interactive variable, log ($IUN \times RECRAT$), whose coefficient should be negative. This factor was found to be extremely significant in explaining the inflow into the UI beneficiary pool.

Two alternative measures for the number of claims (NC) are used. The first (and theoretically preferred) measure is the average weekly number of initial (first time) benefit payments, denoted NFP. This information is available for the United States, Canada and Germany, but not for the United Kingdom and France. In the latter case we use the number of UI benefit recipients, denoted NBR. This variable is less than satisfactory since it refers to a particular month and is not an average measure which is required to satisfy the basic cost identity.

3.3.1. Results

The estimated equations, shown in Table 10, support the following observations.

- As expected inflows are highly cyclical and strongly related to total unemployment. The γ_1 coefficient is positive and is larger for the European countries than for the United States. Specification tests also confirm that for the United States the coefficient is negatively related to the recipiency rate. In the case of the United Kingdom also there is some evidence of variation in this coefficient. For Germany the coefficient seems to have become smaller - it is estimated to be 0.70 after 1980 compared with 0.8 before 1980. In both cases the results may also not be very robust. The equation for France shows the largest γ_1 and that for Canada the smallest. The latter finding is a little surprising.

- There is clear evidence that the inflows were significantly affected by policy changes in all countries. In Canada the inflow was significantly increased by 1971 liberalization, decreased by the 1976 disentitlement of the elderly, and increased by the 1990 legislation. The last is a surprising result. In the United States the inflows were decreased by extended benefits, and increased slightly by the TAA policy, but unaffected by the temporary emergency benefit legislation. In the United Kingdom supplementary benefits raised the inflows substantially, but the removal of earnings related supplement reduced them slightly. In Germany there was one policy measure (1969) that temporarily raised the inflows and one (1980) which reduced them, with the latter effect being nearly three times larger than the former. In France two policy interventions seemed to have roughly equal and opposite effects.

- The direct impact of the $GINDEX$ is found to be positive but insignificant in most cases, but France is an exception. Unfortunately the sign of the coefficient is negative which is a priori implausible. Past empirical studies have generally shown average duration to be more sensitive than inflows to the changes in the average replacement ratio.

3.4. A Model for Average Duration of Benefit (AD)

In modeling the average duration on unemployment compensation the main focus is on UI, not UA. The specified equation is:

$$\log\ AD = \delta_0 + \delta_1 UNR + \delta_2 UNRSQ + \delta_3 GINDEX + \delta_4 Z + error. \tag{3.7}$$

One expects average duration (AD) to be strongly positively related to the aggregate unemployment rate (UNR). Since in all countries there is a maximum duration on UI, the positive relation should not be monotone; hence UNR^2 ($UNRSQ$) is used as an additional explanatory variable. A priori we expect $\delta_1 > 0$ and $\delta_2 < 0$. Maximum AD is obtained when $UNR = -\frac{1}{2}\frac{\delta_1}{\delta_2}$. The generosity index ($GINDEX$) and policy actions affecting eligibility and coverage of the UI system (generically denoted as Z) are added to the regression.

3.4.1. Results

The estimated equations are shown in Table 11. For the United States two equations are reported, the first for the period 1960-93 and the second based on a the shorter time period 1968-89, the latter to facilitate a comparison with the equation for Canada for the same time period. The comparison is based on mutually consistent data for the replacement ratio ($GINDEX$) from Green and Riddell (1993). Beyond this there is only limited scope for international comparisons because data constraints prevent the estimation of duration equations for the United Kingdom and France. For Germany the data were not of the same reliability as Canada and the United States.

- As expected UNR is highly significant in all equations. Canada and the US have coefficients which are almost the same size. The coefficient is nearly twice as large for Germany. The coefficient of $UNRSQ$ appeared to be negligible for Canada, but that of ΔUNR indicates that in the short run a rising rate of unemployment shortens the average duration.
- For Canada there is some evidence that AD on UI benefits was affected by the earnings replacement ratio. The coefficient of $GINDEX$ is positive and

statistically significant. The US equation shows a statistically significant impact of $GINDEX$ on AD, when the longer sample period is used, but this effect is not evident over the shorter time period. Some will interpret this as a confirmation that the relatively greater generosity of the Canadian UI system raises the level of search unemployment (Moorthy (1989)), especially in the 1970s and 1980s.

- The policy dummy variables that were so important in explaining inflows are found to have a smaller impact on duration. For example, of the two major changes in eligibility conditions, only the disentitlement of the elderly in 1976 has a measurable impact - it reduced the duration on UI, which is plausible since the elderly may have been overrepresented in among long durations. The fact that the dramatic changes of 1971 did not seem to have raised AD significantly (there is a small positive coefficient) could be explained partly by the fact that the $GINDEX$ variable already absorbs a part of the effect of policy changes in that period. It also seems plausible that the impact of variations in coverage and eligibility criteria are felt largely through inflows.

- Two other US policy interventions appear to have had an impact. The extended benefits program raised the average duration in the 1970s and the tightening of the eligibility criteria lowered it in the early 1980s.

4. Equations for UC Expenditures

As was noted earlier, there are wide differences in the ratio of UC expenditures to GDP between the five countries. The preceding section attempts to provide an explanation of how these differences arise. The explanation is complex because it involves the *interaction* between the individual country UC systems and the country's unemployment experience. In this section I shall report an econometric attempt at *summarizing* the mechanisms that lead to the inter-country differences in the $\frac{UC}{GDP}$ ratios. These equations may be regarded as a reduced (form) description of the more detailed models given above. For this purpose UC may be defined purely in terms of UI or may additionally include UA also.

The $\frac{UC}{GDP}$ equations reported below have the following form:

$$\frac{UC}{GDP} = \phi_0 + \phi_1 UNR + \phi_2 \Delta UNR + \phi_3 Z_1 + \phi_4 Z_2 + error \tag{4.1}$$

where Z_1 refers to the policy interventions mentioned earlier and Z_2 refers to other country specific factors such as the UI coverage ratio or the benefit recipiency rate.

Notice that $GINDEX$ does not appear as a separate variable; its impact can be observed through intercountry differences in the ϕ_1 coefficient, provided that the variations in $GINDEX$ are not very large over the sample period. If, hypothetically, the UC system in the five countries were very similar, then the intercountry differences in ϕ_1 would reflect purely the differences in the cyclical sensitivity of the $\frac{UC}{GDP}$ ratio. The dummy variables Z_1 will account for policy induced intercept shifts.

Table 12 contains the results based on UI expenditures in all five countries and Table 13 contains the regressions for aggregate UC expenditures in the three European countries.

The following are notable features of the equations:

- In the UI expenditures equations, the greater propensity of the UC system to react to changes in the unemployment rate is reflected in larger values of the ϕ_1 coefficient. This proneness reflects the joint effects of unemployment on inflows, initial payments and average duration. Canada and Germany show the greatest sensitivity to variations in UNR. A one percentage point increase leads to an increase in the $\frac{UI}{GDP}$ ratio of between $\frac{1}{4}$ and $\frac{1}{3}$ percentage point. This is roughly $1\frac{1}{2}$-2 times as large as in France, 2-$2\frac{1}{2}$ times as large as in the United States, and 3-4 times as large as in the United Kingdom (focussing only on the direct effect of the change).

- An interesting result for the United States is that the UI benefit recipiency rate ($RECRAT$) is a useful explanator of the $\frac{UI}{GDP}$ ratio. $RECRAT$ has shown a long term decline, falling from around 0.85 in the early 1960s to around 0.4 in the early 1990s.[11] Since the variable has a well determined coefficient of 0.7, it appears to contribute significantly towards lowering the unemployment sensitivity of $\frac{UI}{GDP}$ ratio in the United States.

- Consistent with the results of earlier sections, the policy dummy variables explain significant variation in the ratio. Especially interesting is the downward shift in the United Kingdom intercept due to the abolition of supplementary benefits ($D1$) and earnings related supplement ($D2$), and the advent of a more stringent UI system in the post-1982 period ($SHIFT$). All these have served to lower the $\frac{UI}{GDP}$ ratio. The pattern of results for Canada and France is consistent with those discussed earlier.

- The reductions in UI expenditure may be offset if the unemployed shift to the UA benefits. The results in Table 13 show that the $SHIFT$ dummy is not a significant variable in the total UC equation for United Kingdom

[11] See ACUC, Report and Recommendations 1994.

. As expected, the sensitivity of aggregate UC to UNR is also higher than that of UI expenditures alone both in United Kingdom and France, though in Germany it is of the same order of magnitude.

5. Summary and concluding remarks

This article has attempted to econometrically model three components of UC expenditures, and $\frac{UC}{GDP}$ ratio, using annual data from the early 1960s to the early 1990s for five OECD countries. It complements previous comparative research on UC systems which, while providing valuable data and insights, stopped short of econometric modeling; see Burtless (1987, 1990), Card and Riddell (1993), Green and Riddell (1993), Reubens (1989). Schmidt et al (1992), and the U.S. Congress (1992). Such models provide a useful way of summarizing the net impact of numerous intercountry differences on UC expenditures. They also provide a method of evaluating the impact of policy interventions on the evolution of the different components of expenditures. Their limitation results from the shortness of the time series used in estimating regressions.

The results show that the sensitivity of the unemployment insurance component of the UC to variations in unemployment is the greatest in Canada and Germany, and the least in the United States and United Kingdom. The UC expenditure indicator, after correcting for differences in the unemployment experience, was the highest in Germany and Canada and the lowest in the United States. Relative to the other four countries, Germany appears to have the most generous eligibility and waiting rules for UI benefits. When both UI and UA are aggregated, the relative differentials in the impact of unemployment on UC between the three European countries is narrower, yet again the German UC expenditures appear to have the highest sensitivity to unemployment.

The source of greater sensitivity UC expenditures to unemployment could potentially lie in greater cyclical sensitivity of all three components of expenditures. In both Germany and Canada, the elasticity of weekly benefits to weekly earnings exceeds one by a significant margin; that is, on average the benefits increased faster than weekly earnings over this period. By contrast, this elasticity was less than one in the United Kingdom. This factor will generate significant differences in the cost structure of the UC system. Because of differences in the quantity and quality of data, the sensitivity of inflows of benefit recipients to variations in the unemployment rate are more difficult to compare, though the intercountry differences seem less striking in this case. Finally, the elasticity of average duration on benefits with respect to the unemployment rate varies between .06 and .12. There is not much robust evidence that differences in the generosity of the UC system, as measured by the earnings replacement rate, can account for the variations in average duration, which seems a very cyclical variable in all countries. Future research could usefully concentrate on other measures of the generosity of the UC system as a source of absolute intercountry differences. To summarize: large differences in the UC expenditures seem to arise partly from the differential cyclical sensitivity of UC expenditures across countries, and partly from the differences in the inherent cost structures of the UC systems. It is worth emphasizing that over the sample period of this paper significant differences in the cost structures existed even if the attention were restricted to the UI component of the UC system.

Finally, consider the role of policy in the growth of the UC expenditures. The paper has produced quite strong evidence that all three components of the UC expenditures responded to policy interventions in all five countries. Some of the policies were aimed at greater liberalization of the UC system, while others, especially those in the 1980s, were in the opposite direction. For example, there is clear evidence that the weekly benefits as well as initial claims in Canada rose sharply between 1971-76 in response to liberalization of the UC system. In response to the disentitlement of the elderly in 1976 the initial claims for benefits declined. Similarly, for the United States average benefits and initial claims have responded to emergency benefit legislation much as expected. The example of the United Kingdom shows that some of the most important policy interventions that affect the UC expenditures have come through fundamental changes that reduce the responsiveness of UI benefits to changes in the earnings. To the extent that such policies follow the US model by accentuating the insurance element of the UC system, as appears to have happened in the United Kingdom, Canada and France to varying degrees, the differences between the US and other countries are now relatively smaller than in the early 1980s.

References

[1] Advisory Council on Unemployment Compensation, *Report and Recommendations* February 1994.

[2] Advisory Council on Unemployment Compensation, *Unemployment Insurance in the United States: Benefits, Financing and Coverage*, February 1995.

[3] Atkinson, A. B. and Micklewright, John. (1985) *Unemployment benefits and unemployment duration: A study of men in the United Kingdom in the 1970s.* London School of Economics and Political Science, Suntory-Toyota International Centre for Economics and Related Disciplines, 1985, 282pp.

[4] Atkinson, A. B. and Micklewright, John. (1991) "Unemployment Insurance and Labor Market Transitions: A Critical Review", *Journal of Economic Literature,* 29(4) , 1679

[5] Atkinson, A. B. and Micklewright, John. (1989) "Turning the Screw: Benefits for the Unemployed 1979-88", Chapter 1 in *The Economics of Social Security,* edited A. W. Dilnot and I. Walker, Oxford: Oxford University Press.

[6] Banerjee, A., J. Dolado, J. W. Galbraith, and D. F. Hendry. (1993) *Co-integration, Error-Correction, and the Econometric Analysis of Non-stationary Data.* Oxford: Oxford University Press.

[7] Blank, R. M., and D. E. Card (1991) "Recent Trends in Insured and Uninsured Unemployment", *Quarterly Journal of Economics,* 1157-1189.

[8] Burtless, Gary T., (1990) Unemployment insurance and labor supply : a survey Washington, D.C. : Brookings Institution, 1990.

[9] Burtless, Gary T., (1987) "Jobless Pay and High European Unemployment", in Schultze, C. L. and Lawrence, R. Z., (eds) *Barriers to Economic Growth: A Transatlantic View,* Washington: Brookings Institution, pp. 105-162.

[10] Card, D. and W. Craig Riddell (1993) "A Comparative Analysis of Unemployment in the U.S. and Canada", in *Small Differences That Matter,* edited David Card and Richard B. Freeman, Chicago and London: The University of Chicago Press.

[11] Disney, R. and Webb, S. (1990) "Why Social Security Expenditure in the 1980s Has Risen Faster than Expected: The Role of Unemployment", *Fiscal Studies;* 11(1*),* 1-20.

[12] Emerson, M. (1988) *What Model for Europe?,* Cambridge: The MIT Press.

[13] Fortin, Bernard and Lanoie, Paul (1992) "Substitution between unemployment insurance and workers' compensation: An analysis applied to the risk of workplace accidents", *Journal of Public Economics,* 49(3), 287

[14] Green, David A. and Riddell, W. Craig (1993) "The Economic Effects of Unemployment Insurance in Canada: An Empirical Analysis of UI Disentitlement", *Journal of Labor Economics,* 11, (1), S96-S147.

[15] Layard, R., S. J. Nickell, and R. Jackman (1991). *Unemployment: Macroeconomic Performance and the Labour Market.* Oxford : Oxford University Press.

[16] Meyer, Bruce D. (1992) Policy Lessons From the U.S. Unemployment Insurance Experiments. Cambridge, MA : National Bureau of Economic Research.

[17] Micklewright, J. (1991) "The Reform of Unemployment Compensation: Choices for East and West", *European Economic Review,* 35(2 / 3), 419-429

[18] Micklewright, J. (1989) "The Strange Case of British Earnings Related Unemployment Benefit", *Journal of Social Policy,* 18(4), 527.

[19] Micklewright, J. (1989) "Why Do Less Than a quarter of the Unemployed in Britain Receive Unemployment Insurance?", EUI Working Paper ECO #90-17, Florence: European University Institute.

[20] Micklewright, J. (1985) "On Earnings-related Unemployment Benefits and Their Relation to Earnings", *Economic Journal,* 95(377), 133-45.

[21] Moorthy, V. (1989) "Unemployment in Canada and the United States: The Role of Unemployment Insurance Benefits" *Federal Reserve Bank of New York Quarterly Review,* 14(4), Winter 1989-90, 48-61

[22] OECD (1984) *High Unemployment: A Challenge for Income Support Policies.* Paris: OECD

[23] OECD (1992), "Unemployment Benefit Rules and Labour Market Policy", chapter 7 in *Employment Outlook,* 199-234, Paris: OECD.

[24] Phillips, P. C. B. and B. E. Hansen (1990) "Statistical Inference in Instrumental Variables Regression with I(1) Processes", *Review of Economic Studies,* 57, 473-95.

[25] Reubens, Beatrice G. (1989) "Unemployment insurance in the U.S. and Europe, 1973-83". *Monthly Labor Review,* April 1989.

[26] Roberti, P. (1984) "Major Policy Issues in Unemployment Compensation", *High Unemployment: A Challenge for Income Support Policies.* Paris: OECD.

[27] Schmidt, G., B. Reissert, and G. Bruche (1992) *Unemployment Insurance and Labor Market Policy: An International Comparison of Financing Systems,* Detroit: Wayne State University Press.

[28] United States. Congress. House. Committee on Ways and Means. Subcommittee on Human Resources. *Reforming the unemployment compensation system* : hearing before the subcommittee on Human Resources of the Committee on Ways and Means, House of Representatives, One Hundred First Congress, first session, May 24, 1989. Published: Washington : U.S. G.P.O., 1990. Description: v, 240 p. : ill. ; 24 cm.

[29] United States Congress. *Unemployment Compensation in the Group of Seven Nations : An International Comparison.* Washington : U.S. G.P.O. : 1992.

[30] United States Department of Health and Human Services. *Social Security Programs Throughout the World*, various years.

Table 1: Comparison of Unemployment Benefit Formula - 1987

Country	Relation to	UI benefit formula varies in relation to			
	past wages	Age	Work history	Region	Dependents
United States	IP: none except	-	-	X	-
	in some states				
	E: 0.8% ,max $7000/yr				
Canada	IP: 2.35% of earnings	X	X	-	-
	E: 3.29% of payroll				
	Max: $530/wk				
	Min: $106/wk				
U.K.	Varies	-	-	-	X
Germany	IP: 2.15% of earnings	-	X	-	-
	E: 2.15%				
France	IP: 2.31% of earnings	X	X	-	-
	E: 4.27% of payroll				
	Max: FF462240/mon				

Source: US DHHS: Social Security Programs Throughout the World
Key: IP: insured person; E: Employer. X indicates the presence of a feature.

Table 2: Comparison of Unemployment Benefit Formula - 1977

Country	Relation to	UI benefit formula varies in relation to			
	past wages	Age	Work history	Region	Dependents
USA	IP: none except	-	-	X	-
	in some states				
	E: 0.4%, max $4200/yr				
Canada	IP: 1.4% of earnings	X	X	X	-
	E: 1.96% of payroll				
	Max: $185 /wk				
U.K.	Varies	-	X	-	X
Germany	IP: 1.5% of earnings	-	X	-	-
	E: 1.5% of payroll				
	Max: DM3,400/mo				
	Min: DM				
France	IP: 2.31% of earnings	X	X	X	-
	E: 4.27% of payroll				
	Max: FF462240/mon				

Table 3: Comparison of Generosity of UI Systems - 1989

Country	Replacement Ratio 1985 UI	Min weeks to qualify/ Reference Period	Min waiting period	Max weeks of benefit for minimally qualified	Coverage %
USA	50	20wks/1 yr	1 week	26 wk	34
Canada	60	10-14 wks/1 yr	2 days	10-42 wk	92
		(27 wk/ 1yr)		(50 wk)	
U.K.	36	11wk+/yr	3 days	52 wk	71
Germany	63	3yr/4 yr	none	52 wk	61
France	57	52 wk/104 wk	none	30 mon	41-47

Source: OECD, Employment Outlook, 1991.

Table 4: Comparison of Generosity of UI Systems - 1979

Country	Replacement Ratio		Min weeks to qualify/ Reference Period	Min waiting period	Max weeks of benefit for minimally qualified	Coverage %
	Year 1	Year 2				
	UI	UA				
USA	37	0	14-20wks/1 yr	1 week	26-39 wk	34
			Varies			
Canada	38		10-14/1 yr	1 week	10-42 wk	89
U.K.	48	40	26-50wk/1 yr	3 days	52 wk or indefinite	70
Germany	66	56	26-50 wk	none	52 wk or indefinite	74
France	67	33-70	3 mo/3 yr	none	21-45 mon	65
			180 hr/3 mo			

Table 5: Expenditure Indicators for UC

		1970	1975	1980	1985	1990
USA	1	0.42	1.18	0.62	0.61	0.60
	2	0.09	0.14	0.09	0.08	0.11
Canada	1	1.67	2.76	2.32	1.87	1.57[a]
	2	0.29	0.40	0.31	0.17	0.21
U.K.	1	0.47	0.70	0.94	2.01	0.90
	2	0.15	0.15	0.13	0.18	0.13
Germany	1	0.40	1.49	1.12	1.41	1.14
	2	0.80	0.44	0.40	0.20	0.22
France	1	0.32	0.76	1.20	1.20	1.27[a]
	2	0.13	0.19	0.23	0.12	0.13

Notes: The first row for each country gives UC expenditures as a percentage of GDP and the second row standardizes the figure by dividing it by the average (OECD standardized) unemployment rate;

a : refers to 1989.

Table 6: Proportions of registered unemployment stock in receipt of UI and UA in 1988

	% Britain	% Germany
UI	26	40
UA	56	23
Neither UI nor UA	18	37

Source - Micklewright (1991), p. 423.

Table 7 : Percentage change in Real Average Weekly UI Benefit (AWB) or UI per benefit recipient (UIPBR)

	Period					
	1961-65	1966-70	1971-75	1976-80	1981-85	1986-90
USA	1.71	5.63	-2.30	-4.05	0.43	1.85
Canada	-5.31	17.02	41.08	-5.22	9.17	6.81
UK[(2)]	25.45	-2.34	5.37	7.54	-0.50	12.74
Germany[(3)]	-	26.31	11.01	26.74	−7.1[(4)]	5.63
France	-	-	56.13[(5)]	11.50	-17.41	-7.22

Notes:
(1) The GDP deflator is used except for Germany.
(2): The number of benefit recipients are estimates for one month.
(3): There is a break in the series in 1981.
(4): Refers to 1983-85.
(5): Refers to 1972-75.
Sources:
United States - *Economic Report of the President*; ET Handbook Financial Data.
Canada - *Canada Yearbook*, and Roberti (1984)
U.K. - *Annual Abstract of Statistics* and Roberti (1984)
Germany - Roberti (1984) and *Arbeitmarkt* (Amtliche Nachrichten der Bundenstalt für Arbeit)
France - *Annuaire Statistique De La France.*

Table 8: Decomposition of UC Growth

United States: Total UC				
4 years ending	Δuc	Δawb	Δnc	Δad
1971	101	27	27	23
1975	100	26	48	8
1979	-58	24	-21	-18
1983	76	32	12	29
1987	-32	12	-29	-18
1991	59	19	31	5

Canada: Total UC				
4 years ending	Δuc	Δawb	Δnc	Δad
1971	93	44	27	67
1975	126	78	19	5
1979	24	25	-9	12
1983	93	34	28	28
1987	3	21	-6	-7
1991	53	25	19	5

Germany: Total UI				
4 years ending	Δuc	Δawb	Δnc	Δad
1971	-	-	-	-
1975	219	45	138	39
1979	-4	41	-29	-16
1983	83	0	-10	54
1987	-11	6	-3	-13
1991	44	47	-20	5

Notes: Lower case letters prefixed by Δ denote percentage changes over the relevant period.

Table 9: The AWB and UIPBR equations

Variable	USA	Canada	UK	Germany	France
Intercept	-1.33	-2.73	3.22	2.19	2.48
	(23.42)	(11.81)	(25.77)	(5.80)	(9.55)
log AWE	1.09	1.30	-	1.18	1.10
	(98.48)	(33.44)	-	(17.93)	(29.41)
log AWE1	-	-	.80	-	-
			(18.77)		
log AWE2	-	-	.70	-	-
			(6.38)		
D1	-.01	0.21	0.42	-.034	-
	(.69)	(5.48)	(5.98)	(.22)	-
D2	-.029	0.043	-	.15	.090
	(1.17)	(1.29)	-	(.96)	(1.15)
D3	.05	0.044	-	-	-
	(3.63)	(1.03)	-		
D4	-	-	0.62	-	-
			(1.02)		
Estimator	FMPH	FMPH	OLS	FMPH	OLS
R^2	.995	.975	.986	.947	.980
DW	.92	1.57	2.24	.85	1.55
Period	1961-93	1970-91	1960-93	1966-93	1972-92

Notes for Table 9:

Here and in later tables absolute 't-ratios' are shown in parenthesis below each coefficient. .

United States: D1 denotes the extended benefit (EB) dummy, equals 1 in 1971-72, 1974-77, and is zero otherwise. D2 denotes eligibility dummy which equals 1 from 1981-82. D3 denotes Temporary Emergency Benefit dummy which equals 1 from 1975 to 1977, 1983-85, 1991-92, and is zero otherwise. D4 denotes TAA assistance dummy which equals 1 from 1974 to 82, and is zero otherwise.

Canada: D1 is the 1971 reform dummy which equals 1 from 1971-76, and is zero otherwise. D2 is the 1976 elderly disentitlement dummy which equals 1 from 1977-79 and is zero otherwise. D3 is the 1990 reform dummy which equals 1 from 1990 to 93, and is zero otherwise.

United Kingdom : D1 denotes the Supplementary Benefits dummy which equals 1 from 1966 to 81. D2 denotes the Earnings Related Supplement dummy which 1 from 1982-88. D3 denotes the 1982 regime shift dummy which equals 1 from 1982 onwards and is zero otherwise.

$\log AWE1 = (1\text{-}D3) \times \log AWE$; $\log AWE2 = D3 \times \log AWE$.

Germany: $D1 = 1$ in 1969, zero otherwise; $D2 = 1$ in 1980, zero otherwise.

France: D1 equals 1 from 1967 to 1969, zero otherwise. D2 equals 1 from 1979 to 1980, zero otherwise. D3 equals 1 from 1984-85, zero otherwise.

Table 10: The NFP and NBR equations

Variable	USA	Canada	UK[1]	Germany	France[2]
Intercept	0.303	4.66	-2.56	2.26	-2.86
	(1.76)	(61.5)	(3.49)	(20.7)	(6.55)
log UN	.743	.480	.911	.824*	1.45
	(20.1)	(33.5)	(15.23)	(45.76)	(22.13)
GINDEX	-.605	.121	-	-	-1.87
	(1.38)	(1.20)			(3.14)
log UN2			1.178	.683*	-
			(12.21)	(46.99)	
D1	-.161	.055	2.25	.26	-
	(7.63)	(4.78)	(2.65)	(5.09)	-
D2	-	-.038	-.022	-.80	.272
		(3.45)	(1.45)	(16.30)	(2.88)
D3	.063	.066		-	-.283
	(2.67)	(4.76)			(3.00)
	-				
Estimator	FMPH	FMPH	FMPH	FMPH	FMPH
R^2	.93	.95	.962	.80	.94
DW	2.01	2.01	1.19	.87	1.29
Period	1963-91	1968-91	1967-90	1968-92	

Notes:
1,2: The dependent variable is the number of UI benefit recipients.

Table 11: The AD equations

Variable	USA		Canada	Germany	Germany	Germany
	(1)	(2)		(UI)	(UI)	(UA)
Intercept	1.75	2.36	1.92	2.09	2.28	2.26
	(14.70)	(33.26)	(5.09)	(16.20)	(36.22)	(46.31)
UNR	0.094	.055	.055	.109	.243	.12
	(3.71)	(9.22)	(9.34)	(16.16)	(2.45)	(6.80)
UNRSQ	-.0028	-	-	-	-.022	-
	(1.41)		-	-	(.95)	
ΔUNR	-		-.031	-	-.118	.039
			(2.77)		(4.11)	(1.47)
GINDEX	0.96	-.0028	.005	.24	.286	-.27
	(3.48)	(1.63)	(2.50)	(0.89)	(1.82)	(1.51)
D1	-.05	.045	.044	-.066	-	-.022
	(1.79)	(1.77)	(.83)	(.82)		(.34)
D2	-	-.064	-.209	-.109	-	
		(2.19)	(2.91)	(1.61)		
R^2	.911	.902	.912	.937	.935	.887
DW	1.63	1.93	2.25	1.14	1.11	2.18
Period	1960-93	1968-89	1968-89	1965-90	1965-80	1965-80

Table 12: $\frac{UC}{GDP}$ Regressions based on UI Expenditures

Variable	USA	Canada	UK	Germany	France
Intercept	-.53	-.43	.061	-.05	-.20
	(7.15)	(4.47)	(2.33)	(1.89)	(5.53)
UNR	.092	.255	.047	.271	.15
	(12.78)	(21.03)	(7.57)	(11.86)	(25.11)
ΔUNR	.039	-.036	.034	.110	-.013
	(4.11)	(1.28)	(2.95)	(5.38)	(1.22)
UNRSQ				-.027	-
				(8.34)	
D1	-	-.22	.081	-	-.048
		(1.54)	(3.13)		(.75)
D2	-.088	.72	.016	-	.082
	(2.25)	(4.72)	(2.60)		(1.10)
D3	.065	.43	-	-	-.21
	(2.98)	(3.96)			(2.66)
RECRAT	.718	-	-	-	-
	(9.48)	-			
SHIFT			-.21	.26	-
			(2.95)	(4.49)	
R^2	.941	.962	.903	.977	.970
DW	1.72	1.57	1.58	1.60	1.13
Period	1962-93	1965-91	1961-93	1961-92	1961-92

Notes:

United States - RECRAT is the UI benefit recipiency ratio defined as the insured unemployment rate divided by aggregate unemployment rate.

United Kingdom - SHIFT factor refers to dummy variable which takes the value 1 after 1982 and is zero otherwise.

Germany - SHIFT factor refers to dummy variable which takes the value 1 after 1980 and is zero otherwise.

France - SHIFT factor refers to dummy variable which takes the value 1 after 1984 and is zero otherwise.

Table 13: $\frac{UC}{GDP}$ Regressions based on Total UC Expenditures

Variable	UK	Germany	France
Intercept	-.23	-.03	-.50
	(2.63)	(.83)	(9.18)
UNR	.179	.285	.30
	(8.63)	(10.41)	(2.60)
ΔUNR	.004	.080	-.02
	(.12)	(3.54)	(.36)
D1	.187	-	
	(2.19)		
D2	.029	-	
	(1.25)		
SHIFT	-.08	.32	-.30
	(.50)	(5.03)	(2.59)
R^2	.944	.977	.980
DW	1.05	1.59	.62
Period	1961-91	1965-92	1961-88

Appendix A
Chronology of Major Changes to the Unemployment Insurance and Assistance System

United States

Year	Change
1935	*Social Security Act of 1935* established a Federal-State program to . provide temporary fianacial assistance
1958	*Temporary Unemployment Compensation* program enacted $\frac{1}{2}$ of regular benefit entitlement up to 13 weeks financed by a temporary Federal program of loans to states.
1961	*Temporary Extended Unemployment Compensation* program enacted.
1970	Permanent Federal-State extended benefits (EB) program established.
1971	*Emergency Unemployment Compensation Act* passed to further extend $\frac{1}{2}$ of benefits for up to 13 weeks.
1974	Federal supplemental benefits passed to provide up to 26 weeks of benefits financed through FUTA taxes and general revenues.
1981	Eligibility for EB tigthened.
1982	*Federal Supplemental Compensation* passed, providing benefits which varied by State
1986	UI benefits made fully taxable.
1991	*Emergency Unemployment Compensation* passed to provide additional weeks of benefits to those exhaustees.
1992	EUC program amended to provide additional weeks of benefits.

CANADA

Year	Change
1940	Establishment of compulsory UI program through the *Unemployment Insurance Act of 1940*
1955	The 1940 UI Act repealed and replaced by *Unemployment Insurance Act of 1955.* Increased coverage, benefit rates and duration of benefits and eased eligibility conditions.
1971	*Unemployment Insurance Act of 1971* included dramatic changes to universal coverage, easier eligibility, distinctions in the treatment of claimants with major and minor labor force attachments.
1976	UI disentitlement of persons of persons between 65 and 70.
1977-79	Tightening up of eligibility and reduction in the maximum weeks of benefits for minimally qualified. Increases in the regional differentiation of UI program
1990-92	Bill C-21 passed in 1990. Reduced maximum benefit period in most regions; extension of coverage to workers over 65; provision of a multitier special benefit structure. Substitution of employer and employee taxes in place of general revenues for the financing of UI benefits.

Sources: Congressional Research Service (1992); Green and Riddell (1993).

United Kingdom

Year	Change
1946	Enactment of *National Insurance and Industrial Injuries Scheme* - a unified system of social insurance.
1948	Introduction of National Insurance and National Assistance
1966	UA program merged into a general means-tested supplementary benefit systemwith a guranteed minimum standard of living. Maximum duration of NI benefit set at 12 months. Earnings Related Supplement (ERS) introduced to supplement flat rate UI benefit up to 6 months. ERS replaces National Assistance.
1973	Distinction between ordinary and long term rates for Supplementary. Benefits introduced
1980	Start of a phased reduction and abolition of the ERS. Changes in the linked spells rule.
1982	Abolition of the ERS. Taxation of benefits. Flat rate system plus supplementary benefit. Beginning of more stringent administration
1986	Extension of the disqualification period.
1988	The Social Security Bill 1988 tightens the contribution conditions of NI benefit.

Sources: Atkinson and Micklewright (1985); Atkinson and Micklewright (1989); Congressional Research Service (1992).

GERMANY

Year	Change
1952	Establishment of institutions to administer UI program.
1969	UI benefits and means tested UA benefits program financed by general revenue.
1977	Slight tighttening up of eligibility.
1990-92	UC system established for the former GDR consisting of flat rate allowance plus redundancy.

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USING THE UNEMPLOYMENT INSURANCE SYSTEM TO TARGET SERVICES TO DISLOCATED WORKERS

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Authors:

Walter Corson
Paul T. Decker

Submitted to:

Advisory Council on Unemployment Compensation
U.S. Department of Labor
200 Constitution Avenue, N.W.
Suite S-4206
Washington, DC 20210

Project Officer:
Laurie Bassi

Submitted by:

Mathematica Policy Research, Inc.
P.O. Box 2393
Princeton, N.J. 08543-2393
(609) 799-3535

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The Unemployment Insurance (UI) system provides limited duration income support for workers who have lost their jobs. This temporary income support is sufficient for workers who will be recalled to their previous employer and those who find a new job quickly. However, workers who are permanently separated from their previous employers and who can expect to have difficulty becoming reemployed may need more active employment services, such as job search assistance or skill retraining, as well as income support.

Policymakers have recently expressed concern that UI claimants receive little in terms of additional employment services, even though the need for services appears to be growing. The UI system has generally provided such services to UI claimants through referrals to the Employment Service (ES), but referral mechanisms have often been weak, and few services have been received by claimants (Richardson et al. 1989). UI claimants may also be eligible to participate in training offered through programs operating under the authority of the Economic Dislocated Worker Adjustment Assistance (EDWAA) Act. However, these services have not been closely coordinated with UI and ES, so few UI claimants have actually participated in the available training opportunities. In addition to the historical lack of services received by UI claimants, recent increases in the proportion of job losers who are on permanent layoff has led to the perception that the need for services has escalated. As a result, there have been calls to strengthen the referral mechanism and to increase the likelihood that UI claimants receive reemployment services (see, for example, the recent publication, "The Changing Labor Market and the Need for a Reemployment Response" (DOL 1993)).

Recent legislation (the Unemployment Compensation Amendments of 1993) has addressed this issue by requiring states to implement systems, called Worker Profiling and Reemployment Service (WPRS) systems, to identify permanently dislocated claimants expected to experience long spells of unemployment and to provide them with mandatory job search assistance services. One of the primary objectives of the WPRS initiative is to coordinate the provision of services from different employment and training agencies so that claimants will have better information about the opportunities for participating in training and other reemployment services.

In this paper, we look at the changing nature of the unemployed population in the U.S. and consider how UI policy has changed to address the needs of this population through the UI system. First, we examine data on the dislocated worker population and address the overlap between dislocated workers and UI claimants. One objective of this study is to investigate whether dislocated workers represent a larger share of the unemployed population now than they did in the past. Next, we examine the receipt of reemployment services by claimants and summarize evidence from recent UI demonstrations on the effectiveness of these services. We also describe the current policy initiatives that are intended to expand the services available to UI claimants. Finally, we consider the methods developed by DOL to target scarce services to claimants who are likely to exhaust their benefits, and we evaluate whether and the extent to which these methods are effective in directing services to claimants who need the services and can benefit from the services. We conclude that these methods are effective.

DISLOCATED WORKERS AND UI

The reemployment problems of permanently laid-off workers have received national attention since the early 1980s. The number of these workers, who have been called "dislocated" or "displaced," is sizeable. For many, the labor market experiences following layoff include long spells of unemployment and a reduction in wages after reemployment.

Since 1984, the Bureau of Labor Statistics of the U.S. Department of Labor (DOL) has identified and tracked dislocated workers through biannual supplements to the Current Population Survey (CPS). In this survey, workers who report "having lost or left a job because of a plant closing, an employer going out of business, a layoff from which they were not recalled, or other similar reason" are classified as dislocated. The 1994 survey showed that about 5.5 million workers were dislocated in the two year period, 1991-1992. Nearly half of this group had been employed in their jobs for three or more years (Gardner 1995).

In an earlier analysis of these data on dislocated workers, the Congressional Budget Office (CBO 1993) found that about 2 million were dislocated each year during the 1980s. Although the numbers were higher than average during the recession of the early 1980s,[1] substantial numbers were dislocated in all years, including those in which the unemployment rate was relatively low. The CBO study also found that workers in goods-producing industries--agriculture, mining, construction, and manufacturing--and in blue-collar occupations were at greater risk of dislocation than workers in service-producing industries and in white-collar occupations. However, substantial fractions of the dislocated worker population were employed in service-producing industries and white-collar occupations. Moreover, differences in the risk of dislocation for these groups narrowed during the 1980s, a trend that continued in the early 1990s (Gardner 1995).

The CBO study also showed that many dislocated workers have long spells of unemployment and reductions in wages after reemployment. One to three years after losing their jobs, half were either not working or had new jobs with weekly earnings of less than 80 percent of their prelayoff earnings. The workers with the largest losses had the least education, were the oldest, and had the longest tenure with the previous employer. Furthermore, dislocated workers who held a job at the time of the survey had endured relatively long jobless spells--the average duration was just under 20 weeks.

Additional studies of dislocated workers based on individual-level data sets have also demonstrated that worker dislocation is costly. Topel (1993) cites three studies that estimated wage losses of 10 to 30 percent as a result of dislocation,[2] depending on the point of observation--that is, dislocated workers who became reemployed earned about 10 to 30 percent less than they earned in their pre-dislocation job. Even five years after their job loss, the wages of dislocated workers in these studies were still about 15 percent

[1]As indicated in the previous paragraph the annual number of dislocated workers was also higher during the recession of the early 1990s than the annual number in the 1980s.

[2]The three studies are Topel (1990, 1991) and Ruhm (1991).

lower than their predislocation levels. The large loss in wages, together with the relatively long jobless spells experienced by dislocated workers, implies that the total cost of dislocation is high. This is confirmed by estimates based on a sample of dislocated workers in Pennsylvania (Jacobson, Lalonde, and Sullivan 1993). Total discounted earnings losses for these workers over the six years after their job loss were equal to an average of $41,000 per worker.

Many dislocated workers enter the UI system. Furthermore, many UI recipients can be classified as dislocated workers. The CBO study found that 70 percent of dislocated workers who were jobless for at least five weeks reported receiving UI benefits. In addition, more than half of the dislocated workers who received UI reported exhausting their benefits. Data from a study of UI recipients in 1988 show that more than half of the UI population had no recall expectations at the time they entered the UI system, and about 36 percent could be characterized as dislocated according to a definition similar to that used in the CBO survey (Corson and Dynarski 1990). Not surprisingly, these figures were higher among UI exhaustees--67 percent had no recall expectations, and 52 percent could be classified as dislocated.

Dislocated workers who enter the UI system, like dislocated workers in general, have longer-than-average spells of unemployment and a greater likelihood of wage reductions than other claimants. Corson and Dynarski (1990) used their sample of UI claimants from 1988 to compare employment and UI benefit outcomes of dislocated and nondislocated workers.[3] They found that dislocated workers, particularly those with substantial job tenure, had lower reemployment rates, longer spells of unemployment, higher UI exhaustion rates, and a lower ratio of post-UI to pre-UI weekly wages than other claimants. For example, only 81 percent of the dislocated workers with 3 or more years of job tenure had become reemployed during the first 20 months after their initial claim, compared with 92 percent of the nondislocated workers.

---

[3]Corson and Dynarski use the BLS definition of dislocated workers, which includes workers who lose their jobs because their plants close, their employer went out of business, or they were laid off and not recalled.

Data from a demonstration program in New Jersey, in which claimants were followed for six years, showed that claimants targeted for demonstration services--permanently separated claimants with three or more years of job tenure--experienced large reductions in annual earnings relative to their UI base period earnings throughout the six-year period (Corson and Haimson 1995). This drop in earnings was considerably larger than that experienced by other claimants. Even claimants who became reemployed had substantial earnings losses; average earnings for employed individuals did not reach pre-UI levels until the fourth year after the initial claim. By the sixth year, annual average earnings for employed individuals exceeded the base period average by $1,889, but this 10.5 percent increase in nominal earnings did not keep pace with inflation (the consumer price index for the Northeast rose approximately 34 percent in this period) or with the average weekly earnings of manufacturing workers in New Jersey (average weekly earnings rose by approximately 25 percent in this period).

UI claimants who exhaust their benefits also have especially high earnings losses. Findings based on a sample of UI exhaustees from manufacturing drawn from 10 states for an evaluation of the Trade Adjustment Assistance program (Corson et al. 1993) illustrate these losses, at least for manufacturing workers. These findings show that the costs of dislocation among UI exhaustees from manufacturing, as measured by earnings losses, were about $35,000 (undiscounted) over the first three years after the initial UI claim. Furthermore, since average earnings were still relatively low among the sample at three years after the initial claim, we can conclude that the full earnings losses would be significantly larger if we were able to expand the post-layoff period of observation.

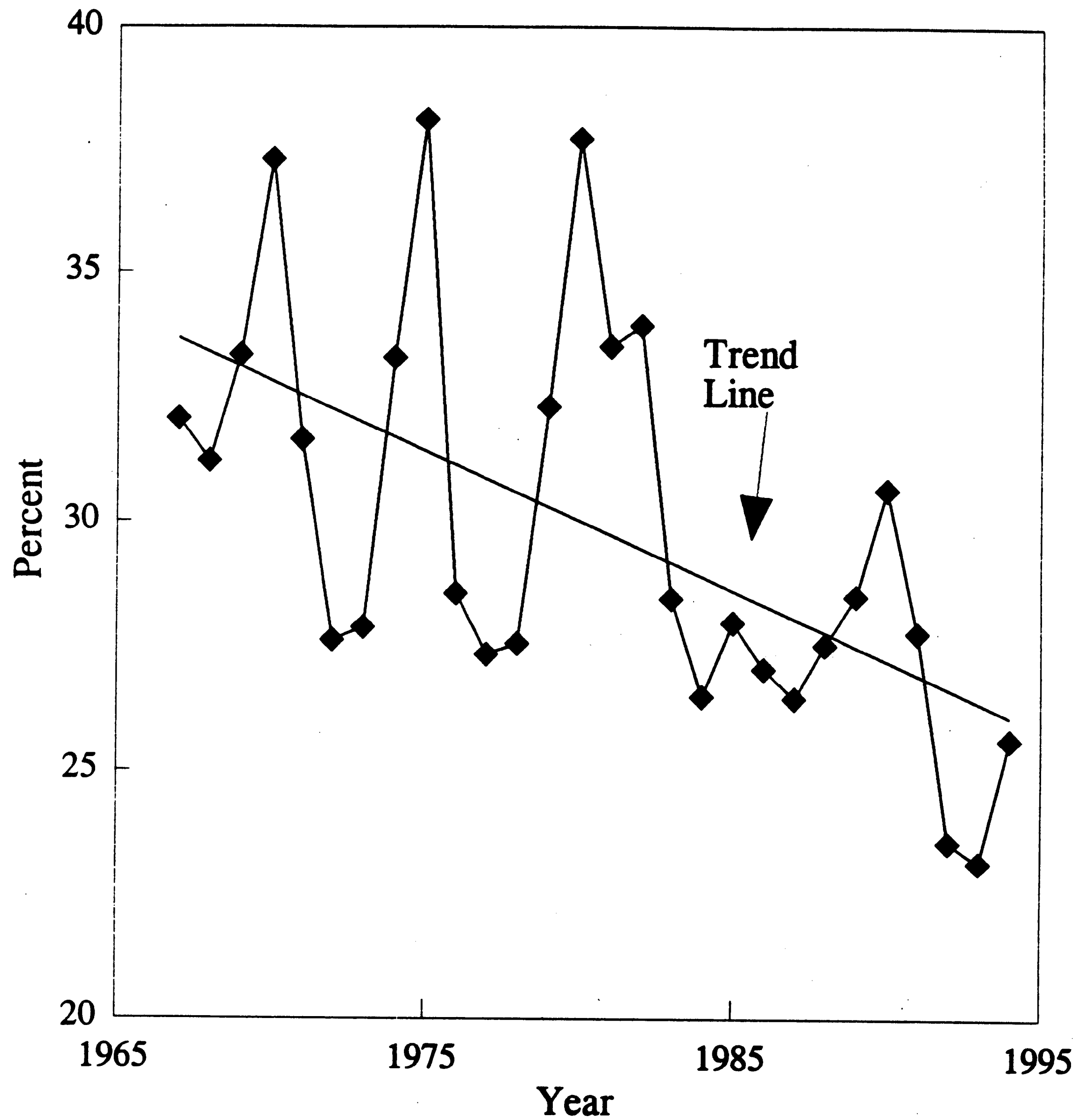
TRENDS IN UNEMPLOYMENT AND DISLOCATION

Trends in three unemployment measures suggest that an increasing proportion of the unemployed population is made up of dislocated workers. These measures include the proportion of unemployed workers on temporary layoff, the proportion of unemployed workers with long unemployment spells, and the proportion of UI claimants who exhaust their benefits. The trend in the proportion of job losers from the CPS who report that they are on temporary layoff is shown in Figure 1. Although the series has been relatively volatile between 1967 and 1994, the long-run trend is clearly a decline in the proportion of job losers on temporary layoff and hence, a corresponding increase in the proportion on permanent layoff. The trend line in Figure 1 implies that the proportion of job losers on temporary layoff declined by nearly three-tenths of a percentage point per year over the observation period. This downward trend is statistically significant at the 99 percent confidence level. Further evidence on the relative decline in temporary layoffs is provided by comparing the average proportion of temporary layoffs early in the observation period with the proportion later in the observation period. The average annual proportion over the first 10 years of the series is 32 percent, compared with about 27 percent over the last 10 years of the series.

At the same time that the share of temporary layoffs has declined, the proportion of unemployed workers who are unemployed for 15 or more weeks has increased. Figure 2 shows the data on unemployment, which are drawn from the CPS, between 1950 and 1994 and the estimated trend in the unemployment data over this period. The trend line shows that an increasing proportion of unemployed workers have stayed unemployed for at least 15 weeks. The highest rate of long-term unemployment shown in Figure 2 occurred not in recent years, but during the recession of the early 1980s, when the proportion of long-term unemployed reached 39 percent in 1983. Nevertheless, the general trend since 1950 has been for long-term unemployment to become more prevalent. The estimated trend suggests that the proportion of unemployed workers who were unemployed for at least 15 weeks increased annually by a quarter of a percentage point over the observation period, and this estimated trend is statistically significant at the 99 percent confidence level. The trend is further illustrated in Table 1, which shows that the average proportion of unemployed workers who are unemployed for at least 15 weeks has been greater in each decade than in the previous decade.

The findings on UI benefit exhaustion parallel those on long-term unemployment. The benefit exhaustion rate, which is shown in Figure 3, applies only to unemployed workers who file for and begin

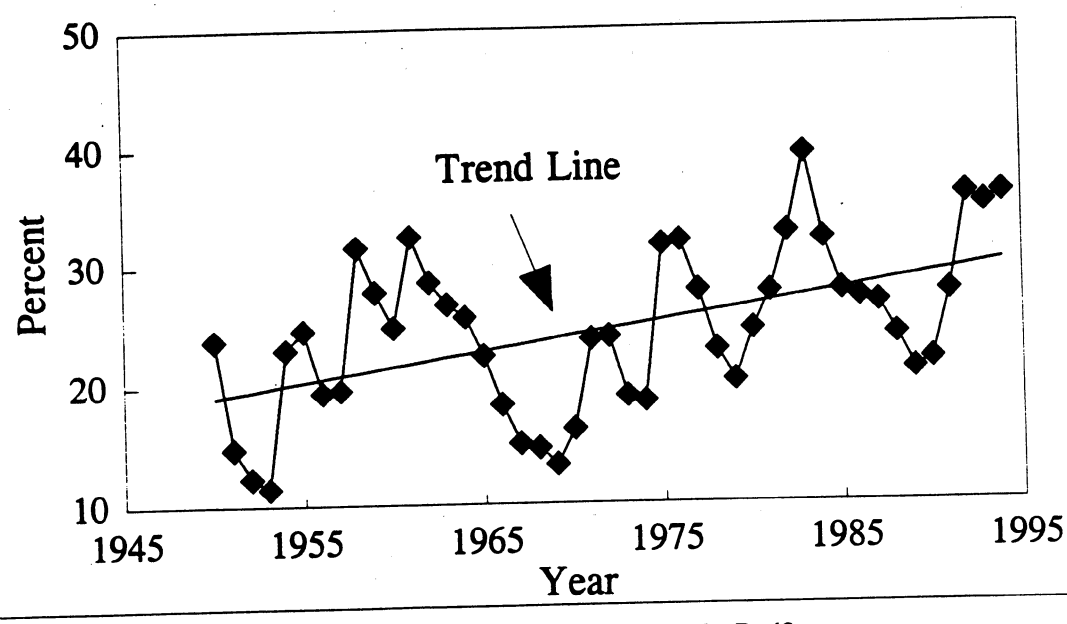
FIGURE 1
PERCENT OF JOB LOSERS ON TEMPORARY LAYOFF



SOURCE: *Economic Report of the President, Table B-42.*

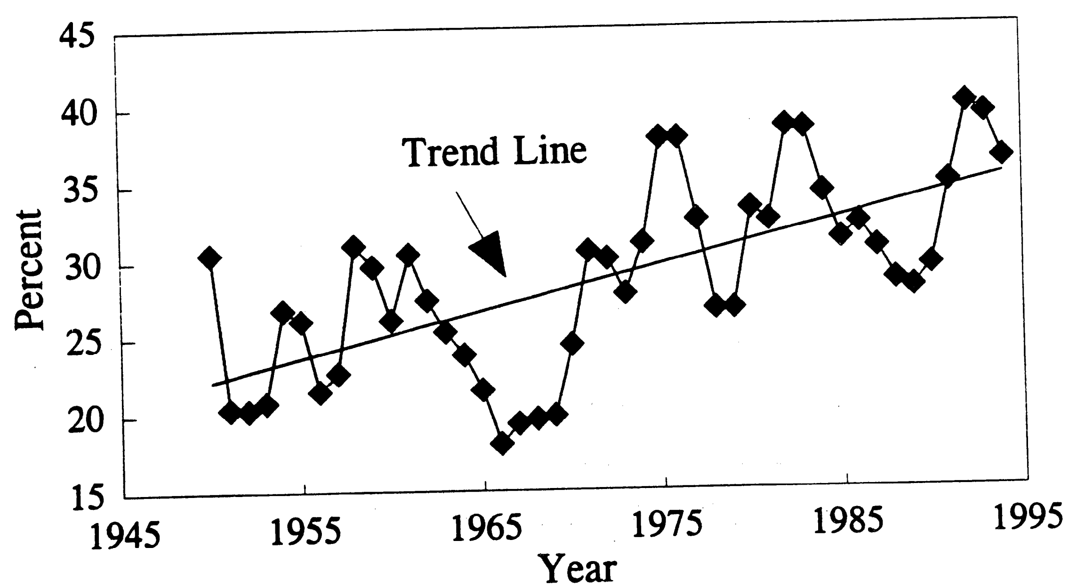
NOTE: Job losers on temporary layoffs are individuals who were laid off and are expecting to be recalled.

FIGURE 2
PERCENT UNEMPLOYED FIFTEEN WEEKS OR MORE



SOURCE: *Economic Report of the President, Table B-42.*

FIGURE 3
UI EXHAUSTION RATE



SOURCE: *Unemployment Insurance Financial Data, ETA Handbook 394, U.S. Department of Labor.*

TABLE 1

RATES OF LONG-TERM UNEMPLOYMENT
AND UI EXHAUSTION, BY DECADE

Time Period	Long-Term Unemployment (As a percentage of all unemployed)	UI Exhaustion Rate (Percent)
1950 - 59	20.82	24.97
1960 - 69	22.20	23.12
1970 - 79	23.55	30.51
1980 - 89	28.27	33.72
1990 - 94	31.18	35.92

SOURCES: *Economic Report of the President 1994*, Table B-42 and *Unemployment Insurance Financial Data*, ETA Handbook 394, U.S. Department of Labor.

NOTE: Long-term unemployment refers to workers unemployed for 15 or more weeks.

to collect UI benefits. Although the trend in benefit exhaustion may respond to changes in the type of workers filing for UI, it is still a useful indicator of the reemployment difficulties of unemployed workers who are receiving UI benefits. As expected, the pattern of the exhaustion rate over time is similar to the pattern of the long-term unemployed measure over time. Estimates of the long-term trend in benefit exhaustion suggest that the exhaustion rate increased annually by an average of three-tenths of one percentage point between 1950 and 1994. The increase has been relatively steady. Table 1 shows that since 1950, the average exhaustion rate in each decade has generally been greater than the rate in the previous decade.

Trends in the three measures shown in Figures 1, 2, and 3 show that an increasing number of job losers do not expect to return to work at their previous employer, that unemployed workers are increasingly more likely to remain unemployed for at least 15 weeks, and that UI claimants are increasingly more likely to exhaust their benefits. These trends suggest that unemployed workers are more likely now than in the past to face long unemployment spells with uncertain reemployment prospects, and, accordingly, more of them could be characterized as dislocated workers.

TRENDS IN RECALL EXPECTATIONS OF UI CLAIMANTS

The positive trend in the UI exhaustion rate suggests that UI claimants are more likely than in the past to have a long spell of unemployment. Two data sets provide further information about the likelihood that UI claimants find it difficult to become reemployed. These data sets are based on representative samples of UI claimants from several states for two different years during the 1980s. The Current Wage and Benefit History (CWBH) data set contains, for samples of claimants in 14 states, information from state administrative records and a survey administered at the initial claim. While these data were collected for several years in the early 1980s, we report on data from 1981, the year in which the data are the most complete. The second data set, the survey of UI exhaustees and nonexhaustees, contains information for samples of exhaustees and nonexhaustees in 20 states. These samples were selected to provide nationally representative estimates of exhaustees and nonexhaustees in 1988 (see Corson and Dynarski 1990).

Table 2 presents data on the recall expectations of UI recipients for the seven states represented in both data sources. It includes information on the number of recipients with general expectations of being recalled to their previous employer and the percentage of recipients who have a definite recall date. Of course, these data represent only the recall expectations at the time of the job loss rather than the actual outcomes. However, Corson and Dynarski (1990) suggest that recall expectations of UI claimants tend to be relatively accurate.

In nearly all of the states shown in Table 2, the proportion of claimants who either expected to be recalled to their previous employer or who had a definite recall date declined during the 1980s, suggesting that the proportion of claimants who were dislocated increased over the 1980s. For example, in New York, the proportion of claimants who expected to be recalled to their previous employer declined from 56.4 percent in 1981 to 42.4 percent in 1988. The decline in New York claimants who had a definite recall date was even larger--from 42.6 percent to 10.1 percent. Two exceptions to the general downward trend in recall expectations are also shown in Table 2. In North Carolina, the percentage of claimants who said they expected to be recalled remained approximately constant between 1981 and 1988, and in Missouri, the percentage of claimants with a definite recall date increased slightly from 11.8 percent to 15 percent over the same period.

Two issues should be considered in interpreting the trends in percentages in Table 2. First, the measures shown in the table are probably sensitive to the business cycle and this may affect the validity of our comparisons of recall rates at the two points of observation. The observation points were chosen because of data availability rather than because they represented similar points in the business cycle. In fact, 1981 was during a recession and 1988 was a non-recessionary period. It is therefore difficult to determine whether the differences between the recall rates in 1981 and 1988 are due to short-term

TABLE 2

UI CLAIMANTS' RECALL EXPECTATIONS IN SELECTED STATES

	Percent Expecting Recall		Percent with a Definite Recall Date	
State	1981	1988	1981	1988
Georgia	98.8	50.1	52.7	30.2
Louisiana	70.7	32.2	14.2	2.0
Missouri	60.0	43.4	11.8	15.0
New York	56.4	42.4	42.6	10.1
North Carolina	65.6	65.9	52.2	36.1
Pennsylvania	70.0	59.4	27.3	26.0
Wisconsin	74.4	67.4	35.7	32.7

SOURCE: Data for 1981 come from tabulations of Continuous Wage and Benefit History data prepared by the Unemployment Insurance Service, U.S. Department of Labor. Data for 1988 were tabulated from a survey of UI exhaustees and nonexhaustees conducted by Mathematica Policy Research. For a description of the survey see Corson and Dynarski (1990).

fluctuations in economic conditions or to a long-term trend in recall expectations. The available data provide no simple way to evaluate the importance of this issue. The second issue has to do with trends in the composition of the UI recipient population. The changes in recall expectations may reflect the type of unemployed workers who file for UI rather than changing conditions faced by the UI population. However, it may not be necessary to disentangle these two sources of a decline in recall expectations. Regardless of the source, it is clear that fewer claimants expect to return to their previous employer. Consequently, it may be more appropriate or more efficient than in the past to use the UI program to target reemployment services to dislocated workers because these workers now make up a larger proportion of the claimant population.

REEMPLOYMENT SERVICES RECEIVED BY UI CLAIMANTS

Although increasing numbers of UI claimants appear to be dislocated workers, relatively few UI claimants receive substantive reemployment services. For example, a recent study of long-term recipients found that just 6 percent were receiving job search assistance that was more intensive than simple ES work registration (Richardson et al. 1989). Rates of service receipt reported in the 1988 survey of UI recipients were considerably higher (55 percent of recipients and 64 percent of exhaustees said they received some services), but a substantial number did not receive services, and few received any intensive services such as assessment, counseling, or job search workshops (Corson and Dynarski 1990).[4]

Dislocated workers may also receive reemployment services and training through several other programs targeted explicitly to them. The main program, EDWAA, which operates as Title III of the Job Training Partnership Act (JTPA), provides funding for states to provide training and related services to dislocated workers.[5] As part of EDWAA, states also conduct rapid response activities, to inform

[4]Less than one-third of the individuals going to the Employment Service received intensive services.

[5]EDWAA uses a relatively broad definition of dislocated workers. Workers are eligible for EDWAA if they have been laid-off or received a notice of termination, are UI eligible, and are unlikely to return to

dislocated workers of available services as soon as a plant closing or mass layoff is announced. Funding under this program has grown rapidly in recent years, from under a half billion dollars in 1990 to more than one billion dollars in 1995. A number of other much smaller programs also provide services to specific groups of dislocated workers. For example, the Trade Adjustment Assistance program provides extended unemployment compensation and additional employment services for workers who lose their jobs because of international trade. The overall number of workers served by EDWAA and other dislocated worker programs is, however, relatively small. The 1988 UI study data suggest that 5 percent of UI recipients and under 10 percent of exhaustees receive any services from these programs (Corson and Dynarski 1990).

EFFECTIVENESS OF SERVICES FOR DISLOCATED UI CLAIMANTS

The results of several recent demonstrations suggest that providing mandatory reemployment services for dislocated UI claimants is likely to speed their reemployment and that the resulting savings in UI expenditures are sufficient to pay for the services. These demonstrations, which were conducted in a number of states (New Jersey, South Carolina, and Washington among others) used the UI system to identify dislocated workers early in their unemployment spells and to refer them to reemployment services. In addition, the referrals to services were mandatory in the sense that workers who did not report for services could have their UI benefits denied.[6]

The most extensive of these demonstrations was conducted in 10 offices in New Jersey in 1986. Services were targeted to dislocated workers determined to be eligible through a series of screens that excluded workers who (1) did not receive a UI first payment within five weeks after their initial claim, (2) were collecting partial UI benefits, (3) were younger than 25 years of age, (4) were employed for fewer

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their previous industry or occupation; if they have been laid off or received a notice of termination as a result of a plant closing or substantial layoff; or if they are long-term unemployed individuals with limited opportunities for reemployment in their occupation.

[6]See Meyer (1995) for a review of the UI demonstrations.

than three years on their last job, (5) had a specific recall date from their employer, or (6) were usually hired through union hiring-hall arrangements. As a whole, these screens excluded approximately 73 percent of all workers who received a first payment from UI during the sample period.

Eligible claimants were then randomly assigned to one of three treatment groups--job search assistance only, job search assistance combined with training or relocation assistance, and job search assistance combined with a cash bonus for early reemployment--and referred to ES and EDWAA for services. The main component of each treatment was an initial set of mandatory job search services--orientation; testing; a one-week, half-day job-search workshop; and a one-on-one counseling/assessment session--delivered over approximately three weeks early in claimants' unemployment spells. Claimants who did not report for these services were referred back to the UI system for a review of their continued eligibility for UI. The demonstration was successful in increasing the level of service receipt and in forging links between the ES and UI to monitor participation.

UI and labor market outcomes of claimants in the three treatment groups were compared to those of a randomly selected control group of claimants who received only regular services. This comparison showed that each treatment in the New Jersey demonstration had a statistically significant effect on reducing the collection of UI benefits and on raising subsequent employment and earnings (Corson et al. 1989 and Corson and Haimson 1995). UI benefits were reduced in both the initial benefit year and in subsequent years. The total benefits of the treatments also exceeded their total costs from the perspectives of both society and the participants. From the perspective of government alone, however, only the job search and reemployment bonus treatments were unambiguously beneficial. There was no clear evidence that providing training or relocation assistance in addition to job search assistance was cost-effective.

An alternative way of changing the incentives to find a job for UI claimants is to offer a financial reward for rapid reemployment. The reemployment bonuses tested in New Jersey and in three other demonstrations--in Illinois, Pennsylvania, and Washington--did just that by offering a cash payment to claimants who found a job within a specified period. In contrast to the New Jersey demonstration, the bonus demonstrations were not explicitly targeted to dislocated workers, but to a broader group of permanently separated claimants. These demonstrations showed that the offer of a reemployment bonus prompted claimants to find jobs more quickly, and as a result, UI benefit payments were reduced. In Illinois, these reductions were large enough to offset the costs of offering the bonuses, but in New Jersey, Pennsylvania, and Washington, the amount of bonus payments plus the administrative costs associated with making the offers generally exceeded the estimated savings for UI payments generated by the offers. These findings suggest that reemployment bonuses offered to a general population of UI claimants are unlikely to be a cost-effective way to speed reemployment, at least from the standpoint of the UI system.

CURRENT POLICY INITIATIVES

The UI system is a logical avenue for identifying workers who might benefit from reemployment services because the majority of dislocated workers collect UI benefits, and they usually begin collecting early in their unemployment spells. Other mechanisms for identifying dislocated workers (such as the rapid response program outreach efforts under EDWAA) are important but more limited than a UI-based approach because they tend to focus on specific groups of dislocated workers (such as those from plant closings or mass layoffs, in the case of EDWAA). Identifying workers early in their unemployment spells has several advantages. By beginning the adjustment process earlier, claimants can use UI benefits as income support during training if training is necessary. For workers who do not need training, the risk of exhausting UI benefits can be lessened. Because many dislocated workers collect UI benefits for a substantial period of time, potential program savings from more rapid reemployment can also be achieved.

This reasoning, combined with evidence from the New Jersey demonstration that long-term UI recipients can be identified early in their unemployment spells, resulted in the Unemployment Compensation Amendments of 1993, which require state UI programs to *profile* claimants as they enter the UI system so that dislocated workers can be identified. DOL's subsequent interpretations of this requirement provide guidance on how states should implement profiling.[7] Specifically, states are encouraged to adapt an approach developed by DOL (1994), in which a two-step process is used to identify dislocated workers and to allocate scarce resources to a subset of these workers. In the first step, claimants who are permanently separated from their previous employer are identified; in the second, a probability of exhaustion is estimated for each claimant on the basis of education, job tenure, industry, occupation, and other variables. Those with the highest probabilities of exhaustion are considered the target group. States that do not have sufficient data to estimate such models are expected to use a fixed set of screens to identify dislocated workers (as in the New Jersey demonstration), but they are encouraged to develop profiling models as more data become available.

Identifying dislocated workers is the first step in helping them become reemployed; allowing them greater access to reemployment services is the second step. For this reason, the worker profiling legislation requires state UI systems to refer profiled claimants to reemployment services. Referred claimants are expected to participate in reemployment services as a condition of eligibility for UI unless they have already done so or have a justifiable cause for failure to participate.

To operationalize these requirements, states are expected to establish agreements between the UI system and service providers (the ES or EDWAA programs) so that profiled claimants can be referred to a service provider and receive services. Service providers in each locality hold initial orientation sessions with claimants, followed by assessment sessions in which a service plan is developed for each claimant. Participation in the reemployment services identified in the plans is a condition for continued UI eligibility. In addition to orientation and assessment, reemployment services include counseling, job search assistance (such as job search workshops), referrals to jobs and job placement, and other similar services, but they do not include training or education. However, claimants can be referred to training or educational

[7]Unemployment Insurance Program Letter No. 45-93, Field Memorandum No. 35-94, and other documents in DOL (1994).

services; if they participate, they do not have to participate in other reemployment services, although participation in training or education is not mandatory. So that UI can monitor compliance with the requirement to participate in reemployment services, states are expected to develop feedback mechanisms to provide UI with information about the participation of referred claimants.

States are currently implementing profiling and reemployment services systems. In late 1994, Delaware, Florida, Kentucky Maryland, New Jersey, and Oregon began profiling and referring claimants to the service system. Systems in the remaining states are expected to become operational during 1995.

Comparable legislation permitting states to offer reemployment bonuses to UI claimants has not been enacted, though the proposed Reemployment Act of 1994 would have allowed states to offer reemployment bonuses as part of their regular UI systems. Under this act, bonuses would have been targeted to claimants identified by state profiling systems as likely to exhaust UI.

EVIDENCE ON TARGETING OF SERVICES

Data from the UI demonstrations can be used to investigate the implications of using the two-step, regression-based approach developed by DOL to target services to claimants for whom the probability of benefit exhaustion is expected to be high. We focus our analysis on two of the UI demonstrations--the New Jersey Unemployment Insurance Reemployment Demonstration and the Pennsylvania Reemployment Bonus Demonstration. Our approach is to simulate the use of the DOL profiling model using the data from these two demonstrations. Then we estimate the impact of the treatments in each of the demonstrations on UI receipt among those claimants targeted and those claimants not targeted according to the profiling model.

Targeting Claimants

In estimating and applying the DOL profiling model, we use the same two-step approach proposed by DOL. For both demonstrations, we start with the full sample of claimants, including those both eligible and ineligible for the demonstration. We then apply the first step of the DOL profiling model by excluding all claimants who either were not permanently separated or were determined demonstration-ineligible for some nonemployment reason.[8] In the second step, we estimate a statistical model of the probability of benefit exhaustion using the control group members who were not screened out in the first step. The statistical models that we estimate are similar to the one developed by DOL for its simulation of the second step in the two-step profiling system. We estimate two models--one for each demonstration.[9] Because of differences in data availability, we cannot specify exactly the same model for the two demonstrations. Although both models include explanatory variables in the five general categories shown below, the particular variables in the two models differ. The explanatory variables in the models include the following:

- ***Education.*** The New Jersey model includes three binary education indicators: (1) no high school diploma, (2) some college, and (3) college degree or more. The Pennsylvania model includes two indicators: one for college degree or more and one for no high school diploma.

- ***Job Tenure.*** The New Jersey model includes a single binary variable, pre-UI job tenure of three years or more. The Pennsylvania model includes the three job-tenure variables used by DOL in its model: pre-UI job tenure of 3 to 5 years, 6 to 9 years, and 10 or more years.

- ***Occupation.*** The New Jersey model contains no variables for occupation. The Pennsylvania model includes a series of binary indicators that correspond to broad occupational categories.

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[8]The resulting sample includes all demonstration-eligible claimants in the Pennsylvania demonstration and all demonstration-eligible claimants plus claimants with fewer than three years of job tenure in the New Jersey demonstration.

[9]The New Jersey demonstration estimates are based on the full sample and data drawn from administrative records. The Pennsylvania demonstration estimates are based on the sample that was interviewed one year after their initial claim and data drawn from both administrative records and the interviews. Different data sources and samples were used for the two demonstrations because differences in data availability in the demonstrations.

- ***Industry.*** In the New Jersey model, we include industry employment growth rates based on 9 broad industrial classifications (see footnote to Table 3). The Pennsylvania model simply includes a series of binary indicators that correspond to 10 broad industrial categories, which are nearly identical to those used to create the New Jersey industry employment growth variable.

- ***Unemployment Rate.*** Both models include unemployment rate data. The New Jersey data refer to unemployment rates in six substate areas in 1986. The Pennsylvania data refer to local unemployment rates in each demonstration site in 1989.

Most of the estimated coefficients in both models have the expected signs (see Tables 3 and 4). In both models, UI exhaustion is positively related to tenure and the unemployment rate. The unemployment rate is statistically significant in both models, and the tenure variables are significant in the Pennsylvania model but not in the New Jersey model. Exhaustion, for the most part, is negatively related to higher levels of education, but only the college graduate indicator in the New Jersey model is statistically significant. One unanticipated result of the New Jersey model estimates is the negative sign on the high school dropout coefficient, suggesting that high school dropouts are less likely to exhaust UI than high school graduates. This coefficient, however, is small and not statistically significant.

Industry and occupation of the pre-UI job also appear to affect the probability of benefit exhaustion. In the New Jersey model, the variable for industrial employment change has a small and insignificant effect on exhaustion. But the simple industry and occupation indicators in the Pennsylvania model show that the probability of exhausting benefits varies among claimants from different occupations and industries. With regard to industries, claimants from mining and finance, insurance, and real estate tend to be less likely to exhaust, while those from wholesale trade tend to be more likely to exhaust. However, none of the industry UI exhaustion rates is significantly different than the rate for the excluded group (retail trade). With regard to occupations, claimants from professional or clerical occupations have the highest exhaustion rates, and these claimants are significantly more likely to exhaust than are claimants from the excluded occupational group (service). Exhaustion is lowest for claimants from forestry, fishing, and farming occupations.

TABLE 3

LOGIT ESTIMATION OF PROBABILITY OF UI EXHAUSTION,
NEW JERSEY UI REEMPLOYMENT DEMONSTRATION

Independent Variable	Mean of Independent Variable	Coefficient	Standard Error	Change in Probability per Unit Change of Independent Variable (Percentage Points)[c]
No High School Diploma	.334	-.110	.086	-2.7
High School Diploma	.412	b	b	b
Some College	.140	-.060	.112	-1.5
College Degree	.114	-.298 **	.123	-7.2
Tenure Less than 3 Years	.262	b	b	b
Tenure 3 or More Years	.738	.070	.083	1.7
Industrial Employment Change	.832	-.008	.007	-.2
Local Unemployment Rate	5.474	.124 ***	.027	3.0
Constant	—	-.870 ***	.171	—

NOTE: Sample includes 2,252 control group members and 901 noneligibles.

The unweighted mean value of the dependent variable (the probability of UI exhaustion) is .445.

The -2 log likelihood is 33.34 with a p value of .0001.

[a] Dependent variable is assigned value of 1 if exhausts UI; 0 otherwise.

[b] Omitted category for dummy variables.

[c] Evaluated at mean of independent variable.

* Significantly different from zero at the 90 percent confidence level, two-tailed test.
** Significantly different from zero at the 95 percent confidence level, two-tailed test.
*** Significantly different from zero at the 99 percent confidence level, two-tailed test.

TABLE 4

LINEAR PROBABILITY ESTIMATES OF UI EXHAUSTION,
PENNSYLVANIA REEMPLOYMENT BONUS DEMONSTRATION

Independent Variable	Mean	Coefficient	Standard Error	Change in Probability Per Unit Change of Independent Variable (Percentage Points)
Unemployment Rate (Percent)	4.86	.017*	.009	1.7
Job Tenure				
3 to 6 years	.155	.046	.033	4.6
6 to 10 years	.087	.022	.042	2.2
10 years or more	.152	.079**	.033	7.9
Education				
High School Dropout	.152	.004	.033	0.4
College Graduate	.105	-.047	.045	-4.7
Industry				
Agriculture	.023	-.082	.122	-8.2
Mining	.005	-.159	.174	-15.9
Construction	.120	-.015	.061	-1.5
Nondurable Manufacturing	.135	-.047	.051	-4.7
Durable Manufacturing	.208	.034	.048	3.4
Transportation	.068	-.051	.059	-5.1
Wholesale Trade	.037	.064	.068	6.4
Finance/Insurance/Real Estate	.036	-.103	.070	-10.3
Services	.198	.031	.042	3.1
Government	.017	-.088	.093	-8.8
Occupation				
Professional	.094	.150***	.058	15.0
Technical	.017	.042	.094	4.2
Sales	.083	.080	.054	8.0
Clerical	.183	.182***	.045	18.2
Forestry/ Fishing/Farming	.022	-.069	.123	-6.9
Mechanical and Repair	.043	.107	.065	10.7
Construction	.072	.015	.072	1.5
Precision Production	.025	.102	.084	10.2
Machine Operator	.172	.022	.053	2.2
Transportation	.070	-.009	.061	-0.9
Handler	.095	.038	.057	3.8

*Significantly different than zero at the 90 percent confidence level in a two-tailed test.
**Significantly different than zero at the 95 percent confidence level in a two-tailed test.
***Significantly different than zero at the 99 percent confidence level in a two-tailed test.

Although few of the individual coefficients in either model are statistically significant, the overall models are clearly useful in identifying claimants with relatively high probabilities of UI exhaustion. As shown in Table 5, just the initial screens used in the first step of the profiling model separate workers into two groups that differ substantially in their actual rates of benefit exhaustion. In the New Jersey demonstration, claimants screened out in the first step have an exhaustion rate of 29.7 percent, compared with an exhaustion rate of 44.1 for the remaining claimants. The Pennsylvania data are similarly affected in the first step of the profiling model. Claimants who are screened out have an 10.8 percent probability of exhaustion, compared with 28.8 percent for the remaining claimants.[10]

The regression model further identifies a group of workers with a relatively high likelihood of exhaustion. Table 5 shows the actual exhaustion rates for permanently separated claimants who have predicted probabilities above and below the 70th percentile of the exhaustion probabilities. For the New Jersey sample, the exhaustion rate is 52.5 percent for claimants above the 70th percentile, compared with 40.5 percent for claimants below the 70th percentile. For the Pennsylvania sample, the exhaustion rate is 41.4 percent for the claimants above the 70th percentile, compared with 23.4 percent for those below the 70th percentile.

Impacts of Services on Targeted and Nontargeted Claimants

The final step in our analysis is to determine whether the profiling system could effectively target reemployment services and bonuses to claimants who are most likely to benefit from services. To answer this question, we estimate the impact of the New Jersey and Pennsylvania demonstration treatments, assuming that services would be offered only to individuals with probability of exhaustion above the 70th

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[10]Exhaustion rates shown in Table 5 for claimants screened out in the first step are much lower in Pennsylvania than in New Jersey. This difference arises partly because the exhaustion rate in general was higher among New Jersey claimants than among Pennsylvania claimants. In addition, the Pennsylvania demonstration excluded claimants who expected to be recalled only if they had a specific recall date within 60 days of their initial claim, while the New Jersey demonstration excluded claimants with any specific recall date.

TABLE 5

BENEFIT EXHAUSTION RATES FOR GROUPS TARGETED AND NOT TARGETED BY CLAIMANT PROFILING (PERCENT)

Claimant Group	New Jersey Demonstration	Pennsylvania Demonstration
Claimants not Permanently Separated and other Ineligibles	29.7	10.8
Permanently Separated Claimants	44.1	28.8
Below 70th percentile	40.5	23.4
Above 70th percentile	52.5	41.4

percentile. Those with predicted probabilities above the 70th percentile were designated as "targeted claimants," while those with predicted probabilities below this level were designated as "nontargeted claimants."[11] We estimate the impact differences in each model by including the variable indicating whether or not a claimant is targeted and a set of variables that represent the interaction of each of the treatment indicators with this variable.

The estimated impacts of both demonstrations tend to be larger for the targeted claimants than for the nontargeted claimants. The impacts of the three New Jersey treatments on UI receipt are about twice as large, on average, for the targeted claimants than for the nontargeted claimants (Table 6). The difference is relatively small for the JSA-only treatment but is larger for the other two treatments. The pattern is similar in the Pennsylvania demonstration--impacts of the treatments tend to be larger for the targeted group than for the nontargeted group. For example, the effect of the combined treatments is to reduce average UI receipt by 1.3 weeks among targeted claimants and by 0.6 weeks for nontargeted claimants. Despite the consistent differences in estimated impacts between the targeted and nontargeted groups in the two demonstrations, none of the differences is statistically significant. This is not surprising, since splitting the sample into targeted and nontargeted groups lowers the effective sample sizes used to generate the impact estimates.

These findings suggest that the DOL profiling model can help to target services to a group of claimants who are likely to benefit from the services. Although the difference in estimates for the two groups are not statistically significant, the estimates are consistently larger for the targeted claimants than for the nontargeted claimants. These estimates also imply that targeting of services based on the profiling model is a relatively efficient way to provide services. The profiling model directs services to a specific group of dislocated claimants who can benefit more from them than a random group of dislocated claimants, thereby generating relatively large savings in UI receipt for the given level of expenditures on services.

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[11]The choice of the 70th percentile as a cutoff for targeting services to claimants accounts for the estimated costs of serving targeted claimants and the resources available to serve them.

TABLE 6

IMPACTS ON UI WEEKS PAID IN BENEFIT YEAR
TARGETED AND NONTARGETED WORKERS
(Standard Errors in Parentheses)

Treatment	Nontargeted	Targeted	Difference
New Jersey UI Reemployment Demonstration			
JSA Only	-0.391 (0.322)	-0.583 (0.433)	-0.192 (0.540)
JSA Plus Training or Relocation	-0.309 (0.293)	-0.736* (0.389)	-0.427 (0.487)
JSA Plus Reemployment Bonus	-0.763** (0.322)	-1.297*** (0.430)	-0.534 (0.537)
Pennsylvania Reemployment Bonus Demonstration			
Low bonus/short qualification	-0.021 (0.830)	0.606 (1.364)	0.627 (1.570)
Low bonus/long qualification	-0.444 (0.515)	-1.315 (0.811)	-0.871 (0.945)
High bonus/short qualification	-0.036 (0.564)	-0.677 (0.901)	-0.641 (1.046)
High bonus/long qualification	-0.784 (0.652)	-1.350 (1.077)	-0.566 (1.238)
Declining bonus	-0.887 (0.575)	-2.089** (0.892)	-1.202 (1.044)
Combined bonuses	-0.555 (0.352)	-1.267** (0 .563)	-0.712 (0.653)

*Significantly different than zero at the 90 percent confidence level.
**Significantly different than zero at the 95 percent confidence level.
***Significantly different than zero at the 99 percent confidence level.

CONCLUSION

The evidence presented in this paper demonstrates that a growing share of unemployed workers in the U.S. face significant barriers to reemployment and the possibility of long unemployment spells. Because many of these workers file for UI at the time of their job separation, UI has been proposed as an avenue through which workers needing reemployment services can be identified and assisted. Recent legislation has moved in this direction by requiring states to develop systems to identify UI claimants who are likely to be unemployed for a long time and to provide these claimants with services. DOL has developed a model that states can use to profile claimants and identify the long-term unemployed. Several states have now implemented their own worker profiling systems based on the DOL model to target reemployment services to long-term unemployed workers.

Our simulations of the use of the DOL profiling model based on data from two of the recent UI policy demonstrations confirm that the profiling model helps to identify claimants who are likely to be unemployed long enough to exhaust their benefits. Those claimants who are predicted by the model to have a high probability of benefit exhaustion do turn out to exhaust their benefits at a higher rate than other claimants. The simulations also suggest that the DOL profiling model effectively targets services to claimants who are likely to benefit from the services. All of the services tested in the demonstrations appear to have had a greater impact on claimants targeted by the model than on other claimants, although the differences are not statistically significant because of the relatively small sample sizes.

These simulation findings suggest that appropriate targeting of reemployment services may be important in determining whether services are cost-effective. In the New Jersey UI Reemployment Demonstration, job search assistance services were originally targeted to a group of claimants who were expected to have long unemployment spells. Our findings suggest that this may have contributed to the finding that these services were cost-effective. In contrast, the Pennsylvania Reemployment Bonus Demonstration targeted bonus offers to a somewhat broader claimant population, and the bonus offers were not found to be cost-effective. Our simulations suggest that if the profiling model were used to target bonus offers to claimants with high exhaustion probabilities, the bonus offers might be cost-effective. However, targeting bonus offers to particular claimants may not be politically feasible if it elicits a strong negative reaction from claimants who do not receive bonus offers.

REFERENCES

Congressional Budget Office. "Displaced Workers: Trends in the 1980s and Implications for the Future." Washington, DC: Congress of the United States, Congressional Budget Office, 1993.

Corson, Walter, Paul Decker, Shari Dunstan, and Anne Gordon. "The New Jersey Unemployment Insurance Reemployment Demonstration Project: Final Evaluation Report." Unemployment Insurance Occasional Paper 89-3. Washington, DC: U.S. Department of Labor, Employment and Training Administration, 1989.

Corson, Walter, Paul Decker, Phillip Gleason, and Walter Nicholson. "International Trade and Worker Dislocation: Evaluation of the Trade Adjustment Assistance Program." Princeton, NJ: Mathematica Policy Research, 1993.

Corson, Walter and Mark Dynarski. "A Study of Unemployment Insurance Recipients and Exhaustees: Findings from a National Survey." Unemployment Insurance Occasional Paper 90-2. Washington, DC: U.S. Department of Labor, Employment and Training Administration, 1990.

Corson, Walter and Joshua Haimson. "The New Jersey Unemployment Insurance Reemployment Demonstration Project: Six-Year Followup and Summary Report." Unemployment Insurance Occasional Paper 95-2. Washington, DC: U.S. Department of Labor, Employment and Training Administration, 1995.

Gardner, Jennifer M. "Worker Displacement: A Decade of Change." *Monthly Labor Review*, vol. 118, no. 4, April 1995, pp .48-57.

Jacobson, Louis, Robert J, LaLonde, and Daniel G. Sullivan. "The Costs of Worker Dislocation." Kalamazoo, MI: The W.E. Upjohn Institute, 1993.

Meyer, Bruce D. "Lessons from the U.S. Unemployment Insurance Experiments." *The Journal of Economic Literature*. vol. 23, March 1995, pp.91-131.

Richardson, Philip et al. "Referral of Long-Term Unemployment Insurance Claimants to Reemployment Services." Unemployment Insurance Occasional Paper 89-2. Washington, DC: U.S. Department of Labor, Employment and Training Administration, 1989.

Ruhm, Christopher J. "Are Workers Permanently Scarred by Job Displacements?" *American Economic Review*, vol. 81, March 1991, pp. 319-324.

Topel, Robert. "Specific Capital and Unemployment: Measuring the Costs and Consequences of Worker Displacement." *Carnegie-Rochester Series on Public Policy*, vol.33, 1990, pp. 181-214.

Topel, Robert. "Specific Capital, Mobility, and Wages: Wages Rise with Job Seniority." *Journal of Political Economy*, vol 99, February 1991, pp. 145-176.

Topel, Robert. "What Have We Learned from Empirical Studies of Unemployment and Turnover?" Paper presented at the ASSA meetings, Anaheim, California, January 1993.

U.S. Department of Labor. "The Changing Labor Market and the Need for a Reemployment Response." Washington, DC: U.S. Department of Labor, December 1993.

U.S. Department of Labor. "The Worker Profiling and Reemployment Services System: Legislation, Implementation Process, and Research Findings." Unemployment Insurance Occasional Paper 94-4. Washington, DC: U.S. Department of Labor, Employment and Training Administration, 1994.

Draft: Comments Welcome

Repeat Use of Unemployment Insurance

by Brian P. McCall*
Industrial Relation Center
University of Minnesota
271 19th Avenue South
Minneapolis, MN 55455

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I. Introduction

Much of the early research on unemployment insurance focused on the effect of weekly benefit amounts and/or benefit durations on unemployment durations (See Atkinson and Micklewright (1991) and Devine and Kiefer (1991)). The question of what determines whether or not an individual files a unemployment insurance claim received relatively little attention. In the United States, at least, this question merits attention since a substantial fraction of unemployed individuals fail to file a unemployment insurance claim even when they are eligible for benefits (Blank and Card (1991), Corson and Nicholson (1988), and McCall(1995a)). Recent research has begun to remedy this oversight (See Anderson and Meyer (1994), Blank and Card (1991), Budd and McCall (1995), McCall (1995a), and Vroman (1991) for example).

There has been very little recent research, however, on the extent of repeat use of the unemployment insurance system. Much of the current research has focused on the extent of repeat use in Canada's unemployment insurance system. For example, Corak analyzed, conditional on first use, the determinants of repeat use in a given time interval (Corak (1992,1993a)) and the differences in the duration of insured unemployment between first claimants and repeat users (Corak (1993b)).

Using a unique data set which track nearly 700,000 men over a twenty year period Lemieux and MacLeod (1995) analyzed the extent of repeat use in Canada and, in particular, looked at the extent to which learning the unemployment insurance system accounts for repeat use. One drawback with the data is its limited demographic information and the fact that demographic information is available only for those who have received unemployment insurance benefits at least once. Nevertheless, Lemieux and MacLeod (1995) provide evidence that is consistent with a learning effect.

The purpose of this paper is to analyze the extent of repeat use of unemployment insurance in the United States. Specifically, the question of whether, among the unemployed who qualify for unemployment insurance benefits, past recipients are more likely to file a claim during a layoff will be analyzed. The extent to which this difference is due to differences in the observable and unobservable individual characteristics or state dependence (Heckman (1978, 1981a,b and 1991)) will also be examined. Such state dependence may occur if there are fixed costs to learning the unemployment insurance system or if using the unemployment insurance system in and of itself changes an individual's preferences for benefit receipt (e.g. lowers the stigma they attach to receiving unemployment insurance benefits). If state dependence is important, then a large shock to take-up rates via an increase in the local or national unemployment rate, for example, may cause higher takeup rates in the future.

This research employs data from both the National Longitudinal Survey of Youth and the Survey of Income and Program Participation. The next section shows that, as in Canada, repeat use of the unemployment insurance system in the United States is quite extensive. Even for a group of youths ranging from ages 26 to 34, nearly two-thirds of their use of the unemployment insurance system in a given calendar year involves repeat use. Repeat users are more likely to be male, white and come from families with lower parental education. There is also empirical evidence of a seasonal component to repeat unemployment insurance benefit receipt. The hazard rate for next usage has a "spike" that occurs precisely twelve months after the start of the previous spell of unemployment insurance benefit receipt.

To analyze the role learning or state dependence in repeat use of the unemployment insurance system, Section III employs instrumental variable techniques to analyze whether receipt of unemployment insurance during the first job separation in which an individual qualifies for unemployment insurance benefits increases the chances of benefit receipt in the next job separation in which he or she qualifies for unemployment insurance benefits. Estimates from the National Longitudinal Survey of Youth and a sample of youths from the Survey of Income and Program Participation find significant state dependence. Point estimates suggest that previous receipt of unemployment insurance benefits increase the probability of receipt in a job separation in which the individual qualifies for benefits from between 0.17 to 0.37.

II. The Extent of Repeat Use of the Unemployment System

This study uses data from the National Longitudinal Survey of Youth (NLSY) and the 1990 wave of the Survey of Income and Program Participation (SIPP). The NLSY is a longitudinal data set comprising individuals selected from the population of American youth aged 14-21 in January 1979 (see Center for Human Resource Research, 1993). Using this data, a history of unemployment insurance (UI) system use for the calendar years 1978 through 1991 was constructed using the non-military subsample which includes 11,000 youths.

The SIPP data was provided to me by Fu Associates and was constructed by them under contract for the Advisory Council on Unemployment Compensation (see Bassi et al. (1995) for more details on the data construction). The SIPP panel is a nationally representative sample of the non-institutionalized population. There are eight interviews that take place very four months. For this data all unemployment spells involving a job loss were identified during the thirty-two month observation period of the 1990 wave. Unemployment insurance receipt was then determined for each spell. The sample consists of 18,150 spells involving 11,407 individuals.

Table 1 presents the extent of UI use for the NLSY and SIPP. Since the NLSY's period of observation is over five times as long as the SIPP's, it is not surprising the extent of repeat use is considerably more. Among those using the UI system at least once, 46% have used the system again in the NLSY as compared to 11% for the SIPP. What is perhaps more surprising is that among those in the NLSY who have used the system at least once, over 8% have used the UI system five or more times and that these individuals account for 28% of all UI receipt during the sample period.

Table II presents the extent of UI use for the NLSY by calendar year. Since this is a sample of youth, more are entering the labor market over time. On the other hand, attrition from the NLSY is also occurring over time. Nevertheless, the raw counts presented in the first column of Table II exhibit a cyclical pattern with peak usage of the UI system occurring during the 1982-1983 and 1991 recessions. This is similar to what is found in Canada (See Lemieux and MacLeod (1995)).

As is shown in the second column of Table II, the fraction of first-time users falls over the sample period. During 1991, 65 percent of all UI recipients in the NLSY were repeat users. Moreover, 17 percent of all UI recipients in 1991 had received UI four or more times in the past.

Some characteristics of UI recipients by frequency of use are presented in Table 3. In the both the NLSY and the SIPP, UI recipients are under-represented by both women and African Americans. Moreover, this under-representation grows as the frequency of UI receipt increases. There is also some evidence that Hispanics are over-represented among those who frequently use the UI system.

In the NLSY, the average AFQT scores of UI recipients is nearly 4.5 percentiles lower than the sample as a whole. This AFQT differential remains fairly constant as the frequency of UI receipt increases. The average years of completed schooling of the father and mother of an individual who has received UI at least once during the observation period is approximately 0.5 years less, than the NLSY sample as a whole. For those who have received UI five or more times, this deficit in parental education increases to 0.8 years for fathers and 1.3 years for mothers.

To investigate the rate at which first-time UI recipients reuse the system, Figures 1 through 5 present Kaplan-Meier hazard estimates of the duration between first and second UI receipt (See Kalbfleisch and Prentice (1980)) broken down by gender.[1] Figures 1 and 2 present monthly hazard estimates for the NSLY and SIPP, respectively, when duration is measured as the time between the end of the first UI spell and the beginning of the second UI spell. For those who have used the UI system only once during the observation period, the duration is measured as the time between the end of the first spell and the end of the observation period. These spells are then treated as incomplete spells in the hazard estimations (See Kalbfleisch

[1]Since the data are discrete, the hazard for second UI receipt at time k refers to the probability of receiving UI a second time k months after first receipt conditional on not receiving UI a second time in the first k-1 months after first receipt.

and Prentice (1980)).

There is some indication from Figure 1 that the hazard for second receipt of UI is higher for male than female first-time recipients. These hazard estimates imply that 25 percent of all male first-time recipients will receive UI again within one year after first receipt while only 20 percent of all first-time female recipients will receive UI again within one year after first receipt.

As Figure 2 shows, there are no clear gender differences in the hazard for second UI receipt in the SIPP. For both men and women, 20 percent of those who receive UI once during the observation period will receive UI again within one year. The different findings in the NLSY and SIPP may reflect the fact that the SIPP is a nationally representative sample while the NLSY focuses on youths.

A long term perspective on the rate of second use of the UI system is provided in Figure 3 which presents Kaplan-Meier estimates of the hazard rate for second UI receipt when the NLSY data are grouped into years. This figure clearly shows that the hazard for second UI receipt declines over time. Moreover, the hazard rate for second UI receipt is higher for men than women for eight years after for first receipt.

To examine the extent of seasonal use of UI, Figures 4 and 5 present monthly hazard estimates of second UI receipt, for the NLSY and SIPP samples, respectively, when the duration between first and second receipt is determined by the time between the beginning of the first UI spell and the beginning of the second UI spell. The spike in the hazard at 12 months observed in Figure 4 and 5 is consistent with notion of seasonal use of the UI system by some individuals.

To investigate further the determinants of the duration until second receipt of UI among those who have received UI at least once during the observation period, a Cox regression was estimated (See Kalbfleisch and Prentice (1980)). Columns (1) and (2) of Table 4 present the results of these estimations for the NLSY and SIPP samples respectively. In the NLSY, the hazard for second receipt of UI is significantly lower for women and Hispanics. An increase in the father's education also significantly decreases the hazard for second receipt of UI.

For the SIPP, the coefficient for the linear age term is significantly positive while the coefficient for the quadratic age term is significantly negative. The coefficient estimates imply that the hazard rate for second receipt of UI increases up to age 58 and decreases thereafter.

Although the evidence presented above shows that repeat use of UI is extensive in the United States and that the rates of use of the UI system differ among various demographic groups, it remains unclear whether part of the propensity to reuse the system is due to an individual "learning" the system. That is does use of the UI system in of itself increase the probability of future use? This question is addressed next.

III. The Impact of Previous Receipt of UI on the Probability of Current Receipt

The question of whether past receipt of UI affects the probability of current receipt will be investigated by constructing a sequence of job separations in which an individual qualifies for UI benefits and by determining whether the probability of UI receipt in the current job separation is affected by UI receipt in past job separations. Specifically, this paper will focus on individual's first two qualifying job separations in the sample period and whether UI receipt during the first affects the probability of UI receipt during the second.

NLSY

The NLSY are ideal data to analyze this question. Since the respondents are first observed at a young age, the initial conditions problem is mitigated (See Heckman (1981b)). Also, the NLSY has followed these youths for more than a dozen years so that the fraction of individuals with multiple job separations in that qualify for UI benefits is large.

To select a sample, for each survey year, beginning with the 1980 survey, every individual who reported a layoff or (beginning with the 1984 survey) reported that they lost their job because of a plant closing or temporary job was identified.[2] Only these categories of job displacement were included because of the nonmonetary UI eligibility requirements (e.g. not quit) and the desire to minimize the number of ineligibles in the sample. Moving forward in time from the previous interview, the first involuntary job separation (as defined above) within each survey year was then identified. Information on the individual's previous job (i.e. the lost job) was identified and the duration of the joblessness spell calculated. Whether an individual received UI benefits was then determined by seeing whether the individual reported receiving UI benefits at any time during the period of joblessness. Subsequent layoffs within the same 12 month interview period were not analyzed. This was done to minimize the amount of involuntary job separations occurring within the same benefit period and, hence, better address the question of long-term state dependence.[3]

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[2]Layoffs from the first survey year, 1979, were not used due to the lack of earnings information needed to assess financial eligibility.

[3]If an individual becomes employed and subsequently loses a part-time job within the same benefit period they may have actually never stopped receiving UI benefits since most states allow individuals to continue receiving at least some benefits while working part-time.

Unfortunately, the NLSY does not ask respondents about their eligibility status or the level of UI benefits they are eligible to receive. Thus, it is necessary to estimate eligibility and impute benefit levels.[4] To estimate eligibility, information on state UI systems contained in the U.S. Department of Labor (various issues) was used to convert state minimum earnings requirements into a base period minimum earnings amount (the base period is the year immediately preceding the layoff).[5] Then, an individual was included in the sample only if the individual's calculated base period earnings exceeded this state minimum amount.[6] If a state had a base period minimum weeks or hours requirement, data on usual hours of work, weeks of tenure in the lost job, and hours of work in the preceding calendar year were used to determine eligibility. The amount of weekly benefits an individual was eligible to receive was imputed using information on state benefit formulas contained in the U.S. Department of Labor (various issues) and reported weekly earnings in the lost job. Finally, data provided by the Advisory Council on Unemployment Compensation on supplementary UI benefits was used

[4]The lack of information on eligibility and benefit levels is not limited to the NLSY (See Blank and Card (1991), Levine (1993), McCall (1995a), and Portugal and Addison (1991), for example).

[5] If a state had both high quarter earnings requirements and a base period earnings requirement, the more stringent of the two requirements was used. This was done in order to minimize the number of ineligibles in the sample.

[6]Since actual base period earnings may include jobs prior to the lost job, the individual's base period earnings were calculated in two ways. First, base period earnings were calculated from the pervious calendar year. Second, base period earnings were calculated using hourly wage, usual hours of work in the lost job, and the minimum of 52 weeks or reported weeks of tenure in the lost job. The results reported below use the maximum of these two base period measures to determine eligibility. However, using the minimum of the two measures does not change the results.

to determine the number of weeks of extended benefits available at the time of the layoff.

The sample consists of 6497 job separations which qualify for UI benefits. A total of 1497 individuals had two or more of these separations. The first columns of Table 5 present means of selected variables. Comparing the sample means with those in the first column of Table 3 shows that women are over-represented in the sample of qualifying job separations. This may simply reflect the fact that the labor participation rate of men is higher than that of women. Although slight, Hispanics are also over-represented in the sample of qualifying job separations.

The takeup rate in the sample of 0.22 is significantly lower than the 0.83 calculated by Blank and Card (1991) and the 0.65 calculated by McCall (1995a). However, Blank and Card (1991) only have data on household heads who are more likely to pick up UI benefits (See McCall (1995a)). Also the regression results in Blank and Card (1991) and McCall (1995a) imply that younger workers are less likely to pick up benefits and the sample mean for the NLSY sample is 23.7 years. Moreover, the Current Population Survey's Displaced Workers Supplements (DWS) used by McCall (1995) ask questions only about the longest held job that was lost in the previous five years. As was shown in McCall (1995b), layoffs in the NLSY involve nearly six less months of job tenure, on average, than a comparable sample of youth layoffs from the DWS. That tenure is an important determinant of takeup can be seen by noting that when NLSY sample is restricted to youths with one or more years of tenure in the lost job, the takeup rate increases from 0.22 to 0.33

Table 2 presents regression results for the determinants of UI receipt in the second job separation using the sample of individuals with at least two qualifying job separations.

Columns (1) through (3) present the results for all individuals. Although the NLSY begins following individuals at a young age, some individuals are already twenty-two by 1979. A substantial fraction of these older individuals may have had a past job separation in which they qualified for UI benefits. To mitigate the initial conditions problem further, columns (4) through (6) present regression results for a sample of individuals who were seventeen years old or younger at the beginning of 1979.

Columns (1) and (4) of Table 2 present regression results when only an indicator variable for UI receipt in the first job separation is used as a predictor variable. Columns (2) and (4) present regression results when independent variables for demographic, UI system, and lost job characteristics are also included. These variables are evaluated at the time of the second job separation. To account for the potential endogeneity of the indicator variable for UI receipt in the first job separation, columns (3) and (5) present two-stage least squares estimates which use demographic, UI system, and lost job characteristics evaluated at the time of the first job separation as instruments for the indicator variable of UI receipt during the first job separation. Since the dependent variable, UI receipt during the second job separation, is a dichotomous variable, the errors are heteroskedastic (Goldberger (1964)). Thus, standard errors that are robust to arbitrary forms of heteroskedasticity (White(1980)) are computed.

As the first column shows, the probability of UI receipt increases by 0.15 when an individual has received UI in the first job separation.[7] Adding other explanatory variables diminishes the effect to 0.09 but it remains significantly different from zero at the 5 percent

[7]The takeup rate during the first layoff was 0.21.

significance level.

Instrumenting for UI receipt in the first job separation increases the coefficient estimate to 0.17. Although the standard errors also increase, the 2SLS estimates is significantly different from zero at the 5% level. A Hausman (1978) test for endogeneity fails to reject the null hypothesis of no endogeneity at the 10% significance level ($\chi^2 = 1.10$) while the instruments are not rejected on the basis of an over-identification test ($\chi^2 = 29.18$). Moreover, the instruments (i.e. those exogenous variables included in the first-stage but not second-stage estimations) are jointly significant at the 0.1% in the first-stage estimations (see column (1) of Table A1 in the appendix). Thus the instruments appear to be reasonable, although some caution is warranted given the recent results on the small sample properties of the IV estimator (See Nelson and Startz (1990a,b)).

The results for the sample of youths born after December 31, 1960 mirror those of the sample which includes all birth years. Again, the 2SLS estimates produce a larger coefficient estimate for the indicator variable of UI receipt in the previous job separation, the Hausman test fails to reject the null of no endogeneity at the 10% significance level ($\chi^2 = 1.38$), the instruments are not rejected on the basis of an over-identification test ($\chi^2 = 21.23$) and the instruments are jointly significant at the 0.1% level in the first-stage estimations (see column (2) of Table A1 in the appendix).

Although not reported, estimates from a bivariate probit model yield similar conclusions for both samples.[8] Evaluating the other regressors at their sample means, the

[8]The estimates from a censored bivariate probit which also includes the job separations of individuals who experience only one qualifying job separation yield similar results.

bivariate probit estimates imply that UI receipt in the first job separation increases the probability of UI receipt in the next job separation in which an individual qualifies for UI benefits by 0.12 for the sample consisting of all birth years and 0.19, for the sample consisting of individuals born after 1960. Moreover, the estimated correlation coefficient in the bivariate probit model while negative is insignificantly different from zero for both samples.

SIPP

Recall that this data consists of 18,150 unemployment spells that were identified during the thirty-two month observation period of the 1990 wave. Whether an individual satisfied the monetary requirements for eligibility was determined from information on state UI eligibility requirements and data on previous earnings history contained in the SIPP. In addition, the amount of weekly benefits an individual was eligible to receive was imputed using information on state benefit formulas and an individual's previous earnings history contained in the SIPP.

Among the 18,150 unemployment spells, 42 percent satisfied the monetary conditions for eligibility.[9] The SIPP also contains information about non-monetary conditions for eligibility, although this is somewhat limited. Among those who satisfied the monetary

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Although job separations occur across different survey years, some may still occur within the same benefit period. Estimates derived from a sample which further restricts job separations to be at least one year apart yields similar results.

[9]Some job separations failed to satisfy the monetary eligibility conditions simply because they occurred near the start of the observation period and, hence, lacked sufficient information to assess eligibility.

conditions for eligibility only 61 percent had valid responses to questions involving reason for job loss. The fact that the analysis below focuses on individuals with at least two job separations increases the severity of this problem. For this reason two samples were constructed. The first sample (SIPP1) determines eligibility using only monetary criteria. This is similar to the Continuous Wage and Benefit History (CWBH) data used by Anderson and Meyer (1993) to analyze UI takeup decisions which also lacks information on non-monetary eligibility. The second sample (SIPP2), which is considerably smaller, determines eligibility using both monetary and non-monetary criteria.[10]

The sample consists of 7366 job separations which satisfy the monetary conditions for UI eligibility and 2900 job separations which satisfied both the monetary and non-monetary conditions for eligibility. A total of 1497 (454) individuals in the SIPP1 (SIPP2) sample had two or more job separations which qualified for UI benefits during the thirty-two month observation period.

A shortcoming of the SIPP in terms of its suitability for addressing the issue of state dependence of UI receipt is its relatively short period of observation. First, many of the individuals who didn't receive UI benefits in their first job separation during the observation period may have received UI benefits in the past. Second, the short period of observation makes it necessary to consider all job separation during the observation period which qualify

[10]Non-monetary conditions for eligibility were satisfied if the individual reported losing there job because of a loss of temporary job, a layoff, or a discharge. Among those who satisfied the monetary conditions for eligibility, 17, 35, and 22 percent of those who lost a job because it was temporary, because of a layoff, and because of a discharge, respectively, received UI benefits. By comparison, only 6 percent of those who qualify for UI benefits on the basis of monetary conditions but who reported quitting their job received benefits

for UI benefits not just those occurring across different benefit years. As noted above, a drawback of including contiguous job separations occurring within the same benefit year is that some individuals may have continued receiving UI benefits during the intervening job if it was part-time. In this case UI receipt across the two job separations may reflect the same spell of UI receipt.

Summary statistics for the SIPP1 and SIPP2 samples are presented in columns (2) and (4) of Table 5. In order to better compare the results between the SIPP and the NLSY, and to limit the initial conditions problem at least to some extent, samples of individuals 30 years and younger at the time of the first eligible job loss are also analyzed. Summary statistics for these two samples (SIPP1-Y and SIPP2-Y) are presented in columns (3) and (5) of Table 5.

Relative to the full sample of job separations reported in Table 3, women and blacks are under-represented in the sample of separations that satisfy the monetary conditions for eligibility. This under-representation increases when the sample is further restricted to those who satisfy the non-monetary conditions for eligibility. Hispanics, on the other hand, are slightly over-represented in both the SIPP1 and SIPP2 samples.

When all job separations satisfying the monetary eligibility conditions are included, the sample takeup rate is 0.22 which is similar to that found by Anderson and Meyer (1993) with the CWBH. When only involuntary job separations the satisfying monetary eligibility conditions are included, the take-up rate increases to 0.30. For the sample of young adults, the take up rate is 0.14 when all separations satisfying the monetary eligibility conditions are included and 0.22 when the sample is further restricted to involuntary separations. This latter take-up rate is similar to that found in the NLSY.

As with the NLSY, only individuals with at least two job separations satisfying the eligibility conditions for UI receipt are analyzed in the SIPP. Again, only the first two qualifying job separations for this group of individuals are used in the statistical analysis. Tables 7 and 8 present the regression results for the SIPP1 and SIPP2 samples. The first three columns of Table 7 and 8 present regression results for the samples which include all ages while the last three columns present results for the thirty and under sample.

As is seen in columns (4) through (6) of Tables 7 and 8, the results for the SIPP1-Y and SIPP2-Y samples are similar to those found with the NLSY with regard to the effect of previous UI receipt on the probability of current UI receipt. Again, instrumenting for UI receipt in the first job separation increases its coefficient estimate, a Hausman test for endogeneity fails to reject the null of no endogeneity (χ^2 = 3.65 and χ^2 = 0.42 for the SIPP1-Y and SIPP2-Y samples, respectively), the instruments pass the over-identification test (χ^2 = 19.25 and χ^2 = 16.52 for the SIPP1-Y and SIPP2-Y samples, respectively), and the instruments are jointly significant at the 0.01 % level and the 3% level for the SIPP1-Y and SIPP2-Y samples (see Table A2 of the appendix), respectively, in the first stage estimates. While the 2SLS coefficient estimate for the UI1 variable is significantly different from zero in the SIPP1-Y sample it is imprecisely estimated in the SIPP2-Y sample.[11]

The results from the samples which include all ages, reported in columns (1) through (3) of Tables 7 and 8 are similar to the results for individuals thirty and under, a Hausman test for endogeneity fails to reject the null of no endogeneity (χ^2 = 0.04 and χ^2 = 0.25 for the

[11]Again bivariate probit estimates yield similar conclusions.

SIPP1 and SIPP2 samples, respectively), the instruments pass the over-identification test (χ^2 = 18.41 and χ^2 = 14.89 for the SIPP1 and SIPP2 samples, respectively), and the instruments are jointly significant at the 0.01 % level and the 0. 3% level for the SIPP1-Y and SIPP2-Y samples, respectively, in the first stage estimates. The 2SLS coefficient estimates for the UI1 variable, however, are less than the OLS coefficient estimates and are not significantly different from zero for both the SIPP1 and SIPP2 samples.

To summarize, the two-stage least squares estimates from both NLSY, SIPP1-Y, and SIPP2-Y samples suggest that there is significant state dependence (See Heckman (1978, 19801a,b)) in UI receipt. That is, previous UI receipt, in and of itself increases the probability of UI receipt during a layoff. Point estimates from the NLSY suggest that the probability of UI receipt may be between 0.17 and 0.37 higher, all else equal, for those who have received UI in the past, although the two-stage least square estimates are admittedly imprecise.

The results from the SIPP1 and SIPP2 samples are less conclusive. This may stem, however, from the short time horizon of the SIPP and the fact that these samples include older individuals. For these older individuals UI receipt during their first qualifying job separation in the observation period may be a poor indicator of past UI receipt.

IV. Summary

This paper analyzed the extent repeat use of UI system in the United States. The data from the NLSY and the SIPP suggest that the extent of repeat use is considerable. For example, among 26 to 34 year olds using the UI system in the United States during 1992, nearly two-thirds are repeat users and 17 percent have used the UI system four or more times in the past.

There is also some evidence that using the UI system in and of itself increases the probability of future use. Although the coefficients are imprecisely estimated they suggest that, all else equal, individuals are nearly twice as likely to use the UI system during an involuntary job separation if they had received UI in the previous involuntary job separation than if they had not. This is consistent with the results of Lemieux and MacLeod [1995] who find that some of the repeat use observed in Canada's UI system can be explained by individual's "learning" the system.

One aspect of repeat use which this study ignores is that receipt of UI may affect the type of job an individual accepts and, hence, the likelihood of a future eligible layoff. That is an individual who finds that he or she "likes" UI may choose types of work such as a seasonal job that heavily rely on the UI system.

References

Anderson, P. and Meyer, B. (1993), "UI Benefits and Takeup Rates," unpublished manuscript.

Atkinson, A. and Micklewright, J. (1991), "Unemployment Compensation and Labor Market Transitions: A Critical Review," Journal of Economic Literature, 29, 1679-1727.

Bassi, L., Chasanov, A., Cubanski, E., Grundman, S., and McMurrer, D. (1995) "The Evolution of Unemployment Insurance," mimeo.

Blank, R. and Card, D, (1991), "Recent Trends in Insured and Uninsured Unemployment: Is There an Explanation?" Quarterly Journal of Economics, CVI, 1157-1189.

Budd, J. and McCall, B. (1995), " The Effect of Unions on the receipt of UI Benefits," University of Minnesota, mimeo.

Corak. M. (1992), "Repeat Users of the UI Program," Canadian Economic Observer, Statistics Canada, Catalogue No 11-010, 3.1-3.25.

Corak, M. (1993a), "UI Once Again: The Incidence of Repeat Participation in the Canadian UI Program", Canadian Public Policy, 19, 162-176.

Corak M. (1993b), "Is UI Addictive? Evidence from the Duration of Repeat Users," Industrial and Labor Relations Review, 47, 62-72.

Corson, W. and Nicholson, W. (1988), "An Examination of Declining UI Claims During the 1980's," Washington, D.C.: U.S. Department of Labor, UI Occasional Paper 88-3.

Devine, T. and Kiefer, N. (1991), Empirical Labor Economics: A Search Approach, Oxford University Press: New York,

Goldberger, A. (1964), Econometric Theory, New York, NY: Wiley.

Hausman, J. (1978), "Specification Tests in Econometrics," Econometrica, 46, 1251-1271.

Heckman, J. (1978), "Simple Statistical Models for Discrete Panel data Developed and Applied to Test the Hypothesis of True State Dependence Against the Hypothesis of Spurious State Dependence," Annales de INSEE, 30, 227-269.

Heckman, J. (1981a), "State Dependence and Heterogeneity," in S. Rosen, ed., Conference on Labor Markets, University of Chicago Press: Chicago, IL.

Heckman, J. (1981b), "The Incidental Parameters Problem and the Problem of Initial Conditions in Estimating a Discrete Time-Discrete Data Stochastic Process," in Structural Analysis of Discrete Data with Econometric Applications, eds. C. Manski and D. McFadden, Cambridge: MIT Press.

Heckman, J. (1991), "Identifying the Hand of the Past," American Economic Review 81, 75-79.

Kalbfleisch, J. and Prentice, R. (1980), The Statistical Analysis of Failure Time Data, John Wiley & Sons: New York, NY.

Lancaster, T. (1990): The Econometric Analysis of Transition Data, Cambridge: Cambridge University Press.

Lemieux, T. and MacLeod, B. (1995) "State Dependence and Unemployment Insurance," Human Resources Development Canada, mimeo.

Levine, P. (1993), "Spillover Effects Between the Insured and Uninsured Unemployed," Industrial and Labor Relations Review, 47, 73-86.

McCall, B. (1995a), "The Impact of UI Benefit Levels on Recipiency," Journal of Business and Economic Statistics, forthcoming.

McCall, B. (1995b), "The Incentive Effects of Partial UI Rules," to appear in Research in Employment Policies Vol. 1, Laurie J. Bassi and Stephen A. Woodbury eds., JAI Press.

Nelson, C. and Startz, R. (1990a), "Some Further Results on the Exact Small Sample Properties of the Instrumental Variable Estimator," Econometrica, 58, 967-976.

Nelson, C. and Startz, R. (1990b),"The Distribution of the Instrumental Variable Estimator and its t-Ratio When the Instrument is a Poor One," Journal of Business, 63, s125-s140.

Portugal, P. and Addison, J. (1990), "Problems of Sample Construction in Studies of the Effects of UI on Unemployment Duration," Industrial and Labor Relations Review, 43, 463-477.

U.S. Department of Labor (various issues), "Significant Provisions of State UI Laws," Employment and Training Administration.

Vroman, Wayne (1991), "The Decline UI Claims Activity in the 1980s," Washington, D.C.: U.S. Department of Labor, UI Occasional Paper 91-2.

White, H. (1980), "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity," Econometrica, 48, 817-838.

Table 1
Extent of Use of UI System

Number of Times	NLSY (1978-1992)	SIPP (1989-1992)
0	0.657	0.858
1	0.184	0.125
2	0.077	0.016
3	0.036	0.001
4	0.018	0.000
5	0.011	-
6	0.007	-
7	0.004	-
8 or more	0.007	-

Source: See Text

Table 2
Use of UI System by Calendar Year: NLSY

Year	Total	Fraction of first time users
1978	198	1.00
1979	325	0.81
1980	564	0.70
1981	632	0.60
1982	881	0.56
1983	812	0.42
1984	696	0.39
1985	706	0.44
1986	603	0.41
1987	471	0.39
1988	432	0.38
1989	452	0.36
1990	462	0.34
1991	594	0.35

Source: See Text.

Table 3. Characteristics of UI System Users

	NLSY					SIPP		
		Times Received UI					Times Received UI	
	Full Sample	1 or more	2 or more	3 or more	4 or more	Full Sample	1 or more	2 or more
Age[a]	18.39	18.67	18.95	19.11	19.17	31.38	37.37	39.74
Female	0.51	0.43	0.38	0.35	0.30	0.48	0.38	0.39
African American	0.26	0.26	0.24	0.21	0.18	0.13	0.10	0.07
Hispanic	0.17	0.17	0.17	0.18	0.20	0.14	0.15	0.19
Years of Schooling[b]	12.87	12.37	12.10	11.94	11.80	11.76	12.10	11.56
AFQT	39.54	35.95	35.06	35.27	35.10	-	-	-
Mother's Education	10.90	10.30	9.97	9.77	9.62	-	-	-
Fathers Education	10.81	10.41	10.18	10.04	9.99	-	-	-
Sample Size	11,000	3,778	1,759	909	509	11,407	1,625	195

Source: See Text

[a]Age in the NLSY refers to age in 1979. Age in the SIPP refers to age in 1990.

[b]Years of Schooling in the NLSY refers to years of completed schooling in 1992. Years of schooling in the SIPP refers to years of completed schooling in 1990.

Table 4. Cox Regression Estimates
Waiting Time Between First and Second Use of UI System

	NLSY	SIPP
Age[a]	-0.005 (0.015)	0.030** (0.008)
Age squared/100	-	-0.026** (0.009)
Female	-0.216** (0.056)	-0.126 (0.150)
African American	-0.042 (0.078)	-0.286 (0.285)
Hispanic	-0.210** (0.098)	0.182 (0.211)
Years of Completed Schooling	-0.030 (0.015)	-0.037 (0.026)
AFQT	-0.000 (0.001)	-
Mother's Education	-0.015 (0.010)	-
Fathers Education	-0.024** (0.012)	-
Year Effects	Yes	yes
Region Effects	Yes	yes
Sample Size	2764	1441

Source: See Text
[a] Age, years of completed schooling, region of residence are measured at the time of first UI receipt.

Table 5. Summary Statistics
Job Separations Qualifying for UI Receipt
(standard deviations in parentheses)

	NLSY	SIPP			
		SIPP1	SIPP1-Y	SIPP2	SIPP2-Y
Variable	(1)	(2)	(3)	(4)	(5)
Received UI	0.216	0.215	0.142	0.300	0.218
Age[a]	23.682 (3.729)	33.548 (11.948)	23.436 (3.786)	34.218 (11.669)	23.830 (3.713)
Lost Job Blue-collar	0.705	0.613	0.606	0.603	0.608
Lost Job Unionized	0.191	0.16	0.090	0.143	0.097
African American	0.256	0.107	0.108	0.096	0.098
Hispanic	0.183	0.147	0.139	0.156	0.139
Female	0.368	0.431	0.418	0.379	0.368
Married	0.282	0.463	0.288	0.492	0.306
Years of schooling[a]	12.011 (2.151)	12.090 (2.740)	12.164 (2.455)	12.170 (2.809)	12.278 (2.426)
Sample Size	6497	7366	3511	2900	1287

Source: See Text
[a]Calculated at the time of layoff.

Table 6. Linear Probability Models of UI Receipt: NLSY Sample[a]
(standard errors in parentheses)

	Variable	All Ages OLS (1)	OLS (2)	2SLS (3)	Born After 1960 OLS (4)	OLS (5)	2SLS (6)
1.	Intercept	0.190** (0.012)	-0.279* (0.160)	-0.240 (0.168)	0.151** (0.014)	-0.271 (0.246)	-0.221 (0.256)
2.	Received UI in Previous Layoff	0.154** (0.028)	0.087** (0.028)	0.166** (0.082)	0.146** (0.039)	0.092** (0.039)	0.193** (0.095)
3.	Weekly Benefit Amount (Hundreds of 1979 Dollars)	-	0.091 (0.062)	0.092 (0.062)	-	0.059 (0.080)	0.056 (0.080)
4.	Weeks of Extended Benefits	-	-0.006** (0.002)	-0.006** (0.002)	-	-0.001 (0.003)	-0.002 (0.003)
5.	Waiting Week	-	0.009 (0.032)	0.013 (0.032)	-	0.040 (0.038)	0.046 (0.039)
6.	Years Tenure in Lost Job	-	0.029** (0.008)	0.027** (0.009)	-	0.038** (0.013)	0.036** (0.013)
7.	Weekly Earnings Lost Job (Hundreds of 1979 Dollars)	-	0.000 (0.015)	0.000 (0.015)	-	0.006 (0.018)	0.006 (0.018)
8.	Lost Part-time Job	-	-0.084** (0.028)	-0.082** (0.028)	-	-0.096** (0.033)	-0.090** (0.033)
9.	Lost Blue-collar Job	-	0.014 (0.028)	0.012 (0.028)	-	0.011 (0.034)	0.010 (0.034)
10.	Lost Unionized Job	-	0.045 (0.030)	0.039 (0.030)	-	0.012 (0.037)	0.004 (0.038)
11.	Local Unemployment Rate	-	0.007 (0.004)	0.006 (0.004)	-	0.011** (0.005)	0.011** (0.005)
12.	African American	-	-0.024 (0.027)	-0.022 (0.027)	-	-0.006 (0.033)	-0.004 (0.034)
13.	Hispanic	-	0.033 (0.033)	0.035 (0.033)	-	0.007 (0.042)	0.008 (0.043)
14.	Female	-	0.034 (0.025)	0.032 (0.033)	-	0.014 (0.029)	0.015 (0.029)
15.	Married	-	0.029 (0.027)	0.025 (0.027)	-	0.034 (0.035)	0.029 (0.036)
16.	Children present	-	0.004 (0.026)	0.002 (0.027)	-	-0.006 (0.033)	-0.008 (0.034)

Table 6. Linear Probability Models of UI Receipt: NLSY Sample[a]
(standard errors in parentheses)

	Variable	All Ages			Born After 1960		
		OLS	OLS	2SLS	OLS	OLS	2SLS
		(1)	(2)	(3)	(4)	(5)	(6)
17.	Age	-	0.016**	0.014**	-	0.002	-0.002
			(0.005)	(0.006)		(0.012)	(0.013)
18.	Years of schooling	-	0.000	0.000	-	-0.000	0.000
			(0.005)	(0.005)		(0.007)	(0.007)
19.	Live in an SMSA	-	-0.013	-0.015	-	0.008	0.007
			(0.028)	(0.028)		(0.034)	(0.037)
20.	Year Effects	no	yes	yes	no	yes	yes
21.	Industry Effects	no	yes	yes	no	yes	yes
22.	Region Effects	no	yes	yes	no	yes	yes
	Sample Size	1497	-	-	824	-	-

Source: See Text
Notes: [a]Dependent Variable: UI Recipient (1 if received UI). Standard errors are robust to arbitrary forms of heteroskedasticity. Standard errors are in parentheses.

* Statistically significant at the 0.10 level (two-tailed test).

* Statistically significant at the 0.05 level (two-tailed test).

Table 7. Linear Probability Models of UI Receipt: SIPP1 Sample[a]
(Standard errors in parentheses)

		All Ages			Thirty and Under		
		OLS	OLS	2SLS	OLS	OLS	2SLS
	Variable	(1)	(2)	(3)	(4)	(5)	(6)
1.	Intercept	0.203**	-0.270*	-0.268*	0.160**	-0.375	-0.523
		(0.011)	(0.150)	(0.150)	(0.015)	(0.530)	(0.539)
2.	Received UI in Previous Separation	0.177**	0.116**	0.097	0.157**	0.078	0.374**
		(0.032)	(0.034)	(0.010)	(0.053)	(0.055)	(0.166)
3.	Weekly Benefit Amount (Hundreds of Dollars)	-	0.083**	0.085**	-	0.091*	0.052
			(0.026)	(0.026)		(0.047)	(0.052)
4.	Earnings-Quarter 1 (Thousands of Dollars)	-	0.014**	0.014**	-	0.012	0.018
			(0.006)	(0.006)		(0.010)	(0.011)
5.	Earnings-Quarter 2 (Thousands of Dollars)	-	-0.018**	-0.019**	-	-0.017	-0.013
			(0.006)	(0.006)		(0.012)	(0.013)
6.	Earnings-Quarter 3 (Thousands of Dollars)	-	-0.011	-0.011	-	0.006	-0.004
			(0.007)	(0.007)		(0.011)	(0.012)
7.	Earnings-Quarter 4 (Thousands of Dollars)	-	0.004	0.004	-	0.006	0.002
			(0.006)	(0.007)		(0.012)	(0.012)
8.	Weeks Worked Last Year	-	0.001	0.001	-	0.002	0.004**
			(0.001)	(0.001)		(0.001)	(0.002)
9.	Lost Blue-collar Job	-	-0.031	-0.031	-	-0011	-0.027
			(0.028)	(0.028)		(0.034)	(0.035)
10.	Lost Unionized Job	-	-0.037	-0.031	-	-0.043	-0.073
			(0.032)	(0.032)		(0.057)	(0.062)
11.	State Unemployment Rate	-	0.016	0.016	-	0.030**	0.031**
			(0.010)	(0.010)		(0.013)	(0.012)
12.	African American	-	-0.018	-0.018	-	-0.020	-0.020
			(0.037)	(0.037)		(0.046)	(0.047)
13.	Hispanic	-	-0.012	-0.012	-	-0.030	-0.028
			(0.032)	(0.032)		(0.042)	(0.043)
14.	Female	-	0.006	0.006	-	-0.011	-0.032
			(0.025)	(0.025)		(0.031)	(0.032)
15.	Married	-	0.021	0.022	-	0.075**	0.053
			(0.026)	(0.027)		(0.037)	(0.041)
16.	Number of Kids	-	0.004	0.004	-	0.005	0.011
			(0.010)	(0.010)		(0.013)	(0.014)

Table 7. Linear Probability Models of UI Receipt: SIPP1 Sample[a]
(Standard errors in parentheses)

	Variable	All Ages			Thirty and Under		
		OLS	OLS	2SLS	OLS	OLS	2SLS
		(1)	(2)	(3)	(4)	(5)	(6)
17.	Age	-	0.014**	0.014**	-	0.013	0.023
			(0.006)	(0.006)		(0.045)	(0.046)
18.	Age-Squared/100	-	-0.013*	-0.014*	-	-0.011	-0.037
			(0.008)	(0.008)		(0.094)	(0.096)
19.	Years of schooling	-	0.003	0.003	-	-0.002	-0.003
			(0.005)	(0.005)		(0.007)	(0.008)
20.	Year Effects	no	yes	yes	no	yes	yes
21.	Industry Effects	no	yes	yes	no	yes	yes
22.	Region Effects	no	yes	yes	no	yes	yes
	Sample Size	1497	-	-	712		

Source: See Text
Notes: [a]Dependent Variable: UI Recipient (1 if received UI). Standard errors are robust to arbitrary forms of heteroskedasticity. Standard errors are in parentheses.

* Statistically significant at the 0.10 level (two-tailed test).

* Statistically significant at the 0.05 level (two-tailed test).

Table 8. Linear Probability Models of UI Receipt: SIPP2 Sample[a]
(standard errors in parentheses)

		All Ages			Thirty and Under		
		OLS	OLS	2SLS	OLS	OLS	2SLS
	Variable	(1)	(2)	(3)	(4)	(5)	(6)
1.	Intercept	0.246** (0.023)	-0.338 (0.329)	-0.297 (0.336)	0.225** (0.032)	-0.831 (1.199)	-0.848 (0.539)
2.	Received UI in Previous Separation	0.174** (0.054)	0.132** (0.061)	0.082 (0.109)	0.226** (0.095)	0.116 (0.104)	0.217 (0.186)
3.	Weekly Benefit Amount (Hundreds of Dollars)	-	0.111** (0.052)	0.117** (0.052)	-	0.034 (0.093)	-0.043 (0.093)
4.	Earnings-Quarter 1 (Thousands of Dollars)	-	0.022** (0.010)	0.021** (0.010)	-	0.034* (0.019)	0.036* (0.020)
5.	Earnings-Quarter 2 (Thousands of Dollars)	-	-0.025** (0.012)	-0.025** (0.012)	-	-0.005 (0.026)	-0.005 (0.026)
6.	Earnings-Quarter 3 (Thousands of Dollars)	-	-0.016 (0.012)	-0.016 (0.012)	-	0.037 (0.028)	-0.041 (0.028)
7.	Earnings-Quarter 4 (Thousands of Dollars)	-	0.021* (0.012)	0.022* (0.012)	-	0.024 (0.024)	0.019 (0.024)
8.	Weeks Worked Last Year	-	0.002 (0.002)	0.002 (0.002)	-	-0.002 (0.004)	-0.001 (0.004)
9.	Lost Blue-collar Job	-	-0.022 (0.057)	-0.022 (0.057)	-	0.042 (0.072)	0.040 (0.071)
10.	Lost Unionized Job	-	0.004 (0.065)	0.008 (0.064)	-	0.056 (0.103)	0.038 (0.011)
11.	State Unemployment Rate	-	0.036* (0.020)	0.036* (0.020)	-	0.051** (0.026)	0.053** (0.026)
12.	African American	-	0.033 (0.079)	0.035 (0.079)	-	-0.067 (0.102)	-0.057 (0.104)
13.	Hispanic	-	-0.024 (0.059)	-0.023 (0.060)	-	0.002 (0.092)	0.001 (0.090)
14.	Female	-	0.072 (0.054)	0.075 (0.054)	-	0.008 (0.069)	0.006 (0.090)
15.	Married	-	0.047 (0.053)	0.047 (0.053)	-	0.108 (0.074)	0.107 (0.075)
16.	Number of Kids	-	0.006 (0.021)	0.006 (0.021)	-	0.043 (0.030)	0.043 (0.030)

Table 8. Linear Probability Models of UI Receipt: SIPP2 Sample[a]
(standard errors in parentheses)

	Variable	All Ages			Thirty and Under		
		OLS	OLS	2SLS	OLS	OLS	2SLS
		(1)	(2)	(3)	(4)	(5)	(6)
17.	Age	-	0.008 (0.013)	0.007 (0.013)	-	0.074 (0.107)	0.070 (0.107)
18.	Age-Squared/100	-	-0.010 (0.017)	-0.009 (0.017)	-	-0.134 (0.217)	-0.130 (0.218)
19.	Years of schooling	-	0.006 (0.009)	0.006 (0.009)	-	-0.013 (0.016)	-0.015 (0.016)
20.	Year Effects	no	yes	yes	no	yes	yes
21.	Industry Effects	no	yes	yes	no	yes	yes
22.	Region Effects	no	yes	yes	no	yes	yes
	Sample Size	454	-	-	204		

Source: See Text
Notes: [a]Dependent Variable: UI Recipient (1 if received UI). Standard errors are robust to arbitrary forms of heteroskedasticity. Standard errors are in parentheses.

* Statistically significant at the 0.10 level (two-tailed test).

* Statistically significant at the 0.05 level (two-tailed test).

Appendix

Table A1. 1st Stage Estimates for 2SLS: NLSY Sample[a]
(standard errors in parentheses)

	Variable[b]	All Ages (1)	Born After 1960 (2)
1.	Intercept	-0.354*	-0.226
		(0.186)	(0.314)
2.	Weekly Benefit Amount (Hundreds of 1979 Dollars)	0.161**	0.167
		(0.072)	(0.101)
3.	Weeks of Extended Benefits	0.003	0.004
		(0.002)	(0.003)
4.	Waiting Week	-0.036	-0.028
		(0.052)	(0.064)
5.	Years Tenure in Lost Job	0.041**	0.044**
		(0.010)	(0.015)
6.	Weekly Earnings Lost Job (Hundreds of 1979 Dollars)	0.013	0.042
		(0.020)	(0.032)
7.	Lost Part-time Job	-0.037	-0.018
		(0.032)	(0.039)
8.	Lost Blue-collar Job	-0.076**	0.066
		(0.033)	(0.041)
9.	Lost Unionized Job	0.003	0.033
		(0.029)	(0.037)
10.	Local Unemployment Rate	0.005	0.002
		(0.005)	(0.006)
11.	Married	0.080**	0.069
		(0.035)	(0.048)
12.	Children present	-0.032	-0.047
		(0.041)	(0.055)
13.	Age	0.006	0.006
		(0.011)	(0.021)
14.	Years of schooling	0.031**	0.052**
		(0.015)	(0.019)
15.	Live in an SMSA	-0.039	-0.063
		(0.042)	(0.055)
16.	Year Effects	yes	yes
17.	Industry Effects	yes	yes
18.	Region Effects	yes	yes

Table A1. 1st Stage Estimates for 2SLS: NLSY Sample[a]
(standard errors in parentheses)

	All Ages	Born After 1960
Variable[b]	(1)	(2)
F-statistic[c]	3.765**	2.777**
Sample Size	1497	824

Source: See Text

Notes: [a] Dependent Variable: UI Recipient (1 if received UI) in the first eligible layoff. While all exogenous variables are included in the regressions, only coefficient estimates for those exogenous variables excluded from the second stage estimates are reported. Standard errors are in parentheses.

* Statistically significant at the 0.10 level (two-tailed test).

* Statistically significant at the 0.05 level (two-tailed test).

[b] All variables are measured at the time of the first qualifying job separation.

[c] Test of joint significance of variables excluded from second stage.

Table A2. 1st Stage Estimates for 2SLS: SIPP Sample[a]

	Variable[b]	SIPP1 All Ages (1)	SIPP1 Thirty and Under (2)	SIPP2 All Ages (3)	SIPP2 Thirty and Under (4)
1.	Intercept	-0.317** (0.148)	0.480 (0.483)	0.812** (0.298)	0.271 (1.114)
2.	Weekly Benefit Amount (Hundreds of Dollars)	0.000 (0.032)	-0.079* (0.046)	-0.055 (0.068)	-0.215** (0.107)
3.	Earnings-Quarter 1 (Thousands of Dollars)	0.030** (0.007)	0.038** (0.011)	0.036** (0.011)	0.051** (0.026)
4.	Earnings-Quarter 2 (Thousands of Dollars)	0.008 (0.007)	0.000 (0.008)	0.017 (0.013)	0.007 (0.027)
5.	Earnings-Quarter 3 (Thousands of Dollars)	-0.005 (0.010)	0.024* (0.007)	-0.011 (0.018)	0.017 (0.038)
6.	Earnings-Quarter 4 (Thousands of Dollars)	-0.001 (0.009)	-0.008 (0.013)	0.007 (0.015)	-0.055* (0.038)
7.	Weeks Worked Last Year	0.005** (0.001)	0.002 (0.001)	0.008** (0.002)	0.011** (0.003)
8.	Lost Blue-collar Job	-0.059* (0.031)	-0.033 (0.035)	0.032 (0.060)	0.007 (0.078)
9.	Lost Unionized Job	0.051 (0.035)	0.032 (0.046)	0.025 (0.072)	0.335** (0.123)
10.	State Unemployment Rate	0.001 (0.011)	0.004 (0.014)	-0.003 (0.022)	0.015 (0.035)
11.	Married	0.025 (0.046)	-0.026 (0.055)	0.052 (0.085)	0.030 (0.104)
12.	Number of Kids	0.023 (0.019)	0.022 (0.022)	0.024 (0.033)	0.020 (0.045)
13.	Age	0.053 (0.045)	0.148 (0.119)	0.073 (0.083)	0.155 (0.279)
14.	Age-Squared/100	-0.079 (0.062)	-0.309 (0.252)	-0.179 (0.116)	-0.597 (0.590)
15.	Years of schooling	-0.002 (0.015)	-0.000 (0.014)	-0.042 (0.030)	0.007 (0.041)
16.	Year Effects	yes	yes	yes	yes
17.	Industry Effects	yes	yes	yes	yes

Table A2. 1st Stage Estimates for 2SLS: SIPP Sample[a]

		SIPP1		SIPP2	
		All Ages	Thirty and Under	All Ages	Thirty and Under
	Variable[b]	(1)	(2)	(3)	(4)
18.	Region Effects	yes	yes	yes	yes
	F-statistic	4.027**	1.778**	3.623**	1.636**
	Sample Size	1497	712	454	204

Source: See Text

Notes: [a] Dependent Variable: UI Recipient (1 if received UI) in the first eligible layoff. While all exogenous variables are included in the regressions, only coefficient estimates for those exogenous variables excluded from the second stage estimates are reported. Standard errors are in parentheses.

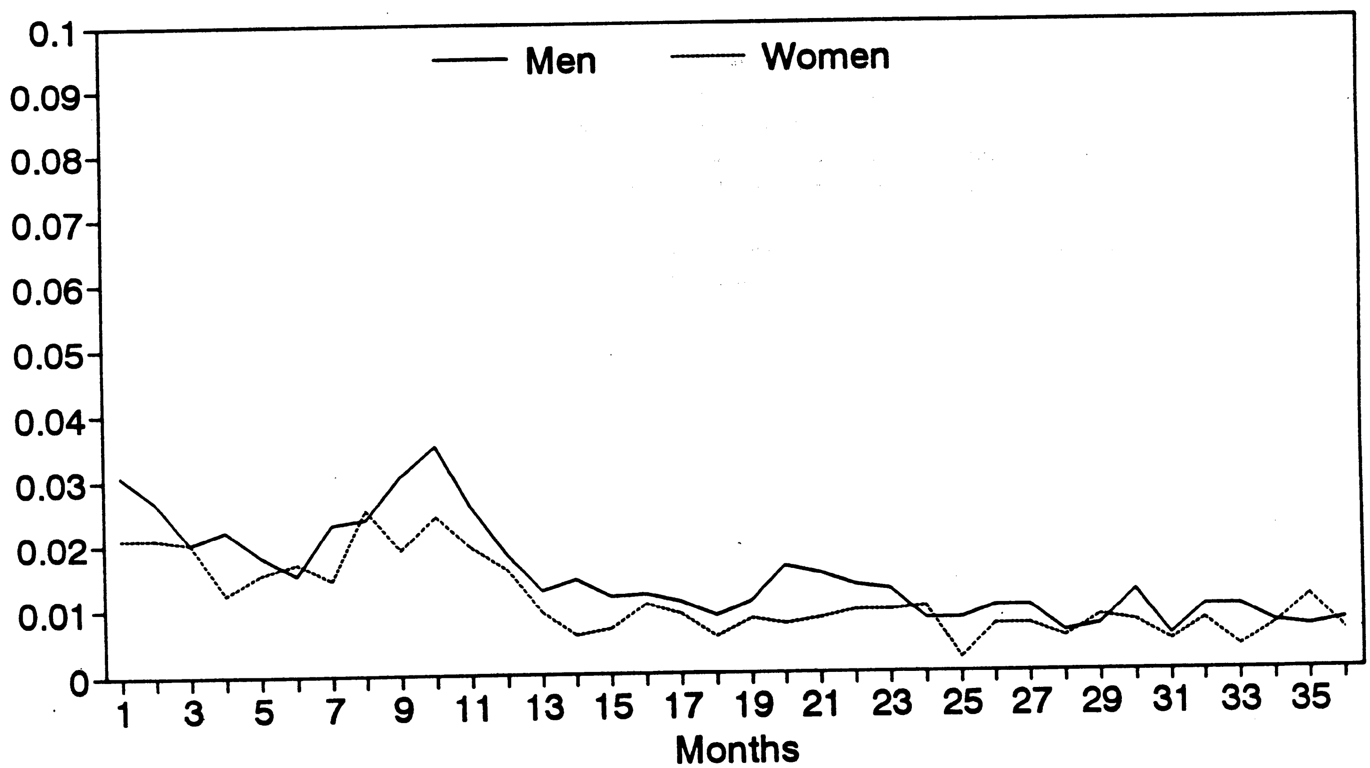
* Statistically significant at the 0.10 level (two-tailed test).

* Statistically significant at the 0.05 level (two-tailed test).

[b] All variables are measured at the time of the first qualifying job separation.

[c] Test of joint significance of variables excluded from second stage.

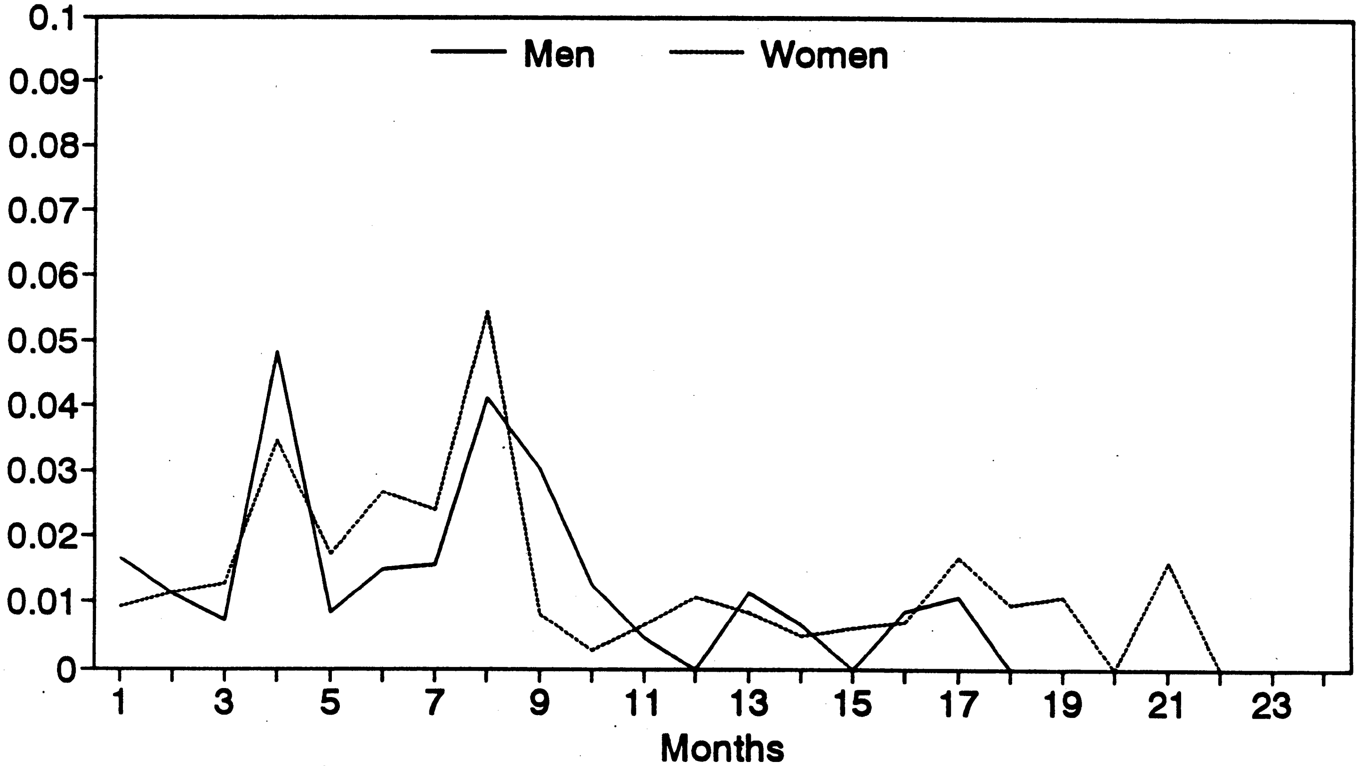
FIGURE 1
Repeat Use of UI Hazard: K-M Estimates



Source: National Longitudinal Survey of Youth

Notes: Spells are measured from the end of the first UI use to the start of the second UI use.

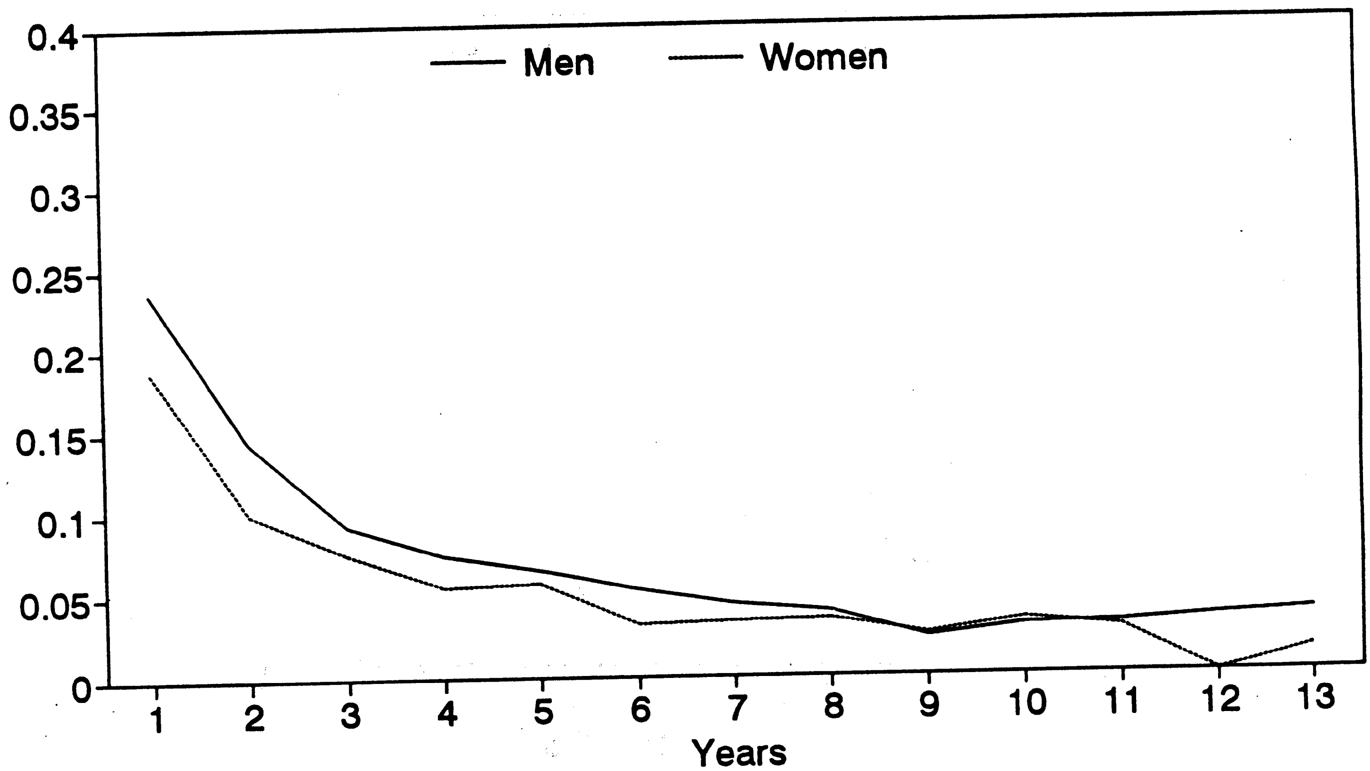
FIGURE 2
Repeat Use of UI Hazard: K-M Estimates



Source: 1990 Survey of Income and Program Participation

Notes: Spells are measured from the end of the first UI use to the start of the second UI use.

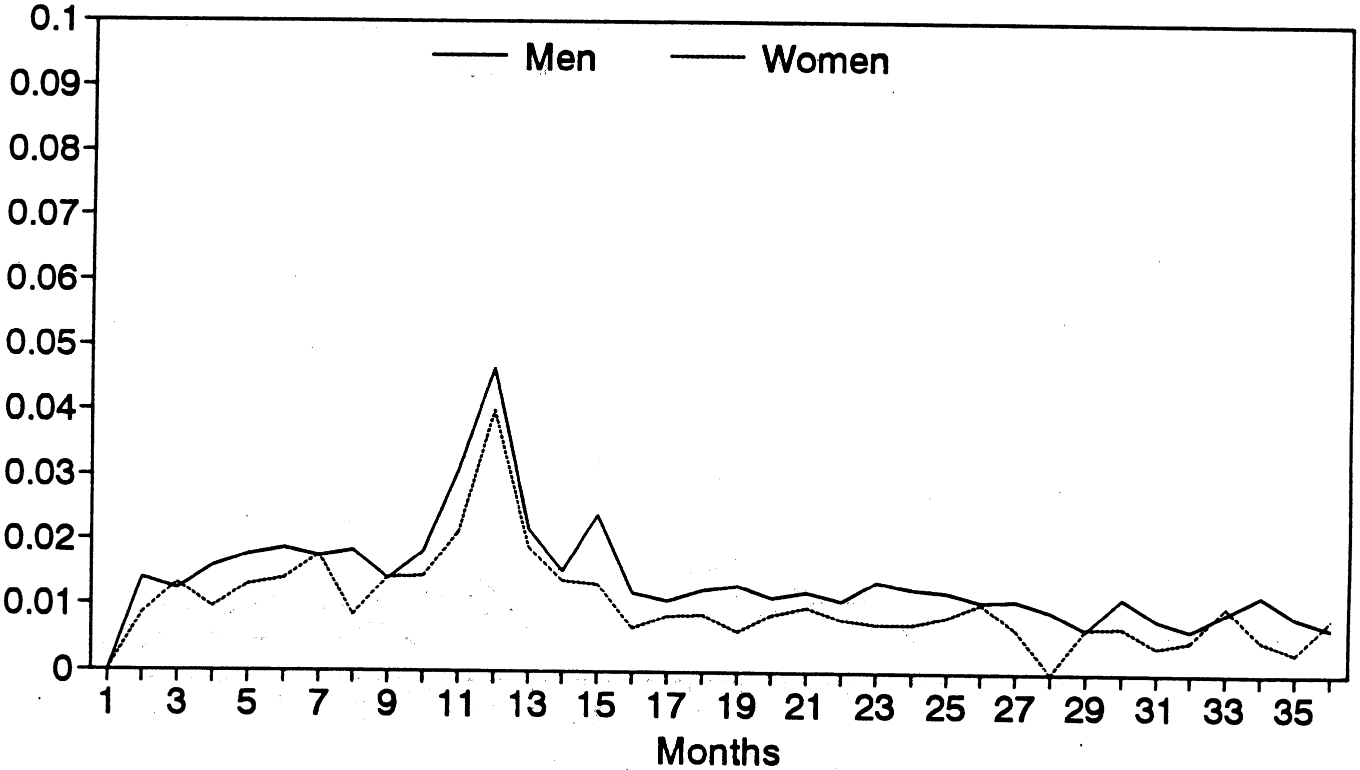
FIGURE 3
Repeat Use of UI Hazard: K-M Estimates



Source: National Longitudinal Survey of Youth

Notes: Spells are measured from the end of the first UI use to the start of the second UI use.

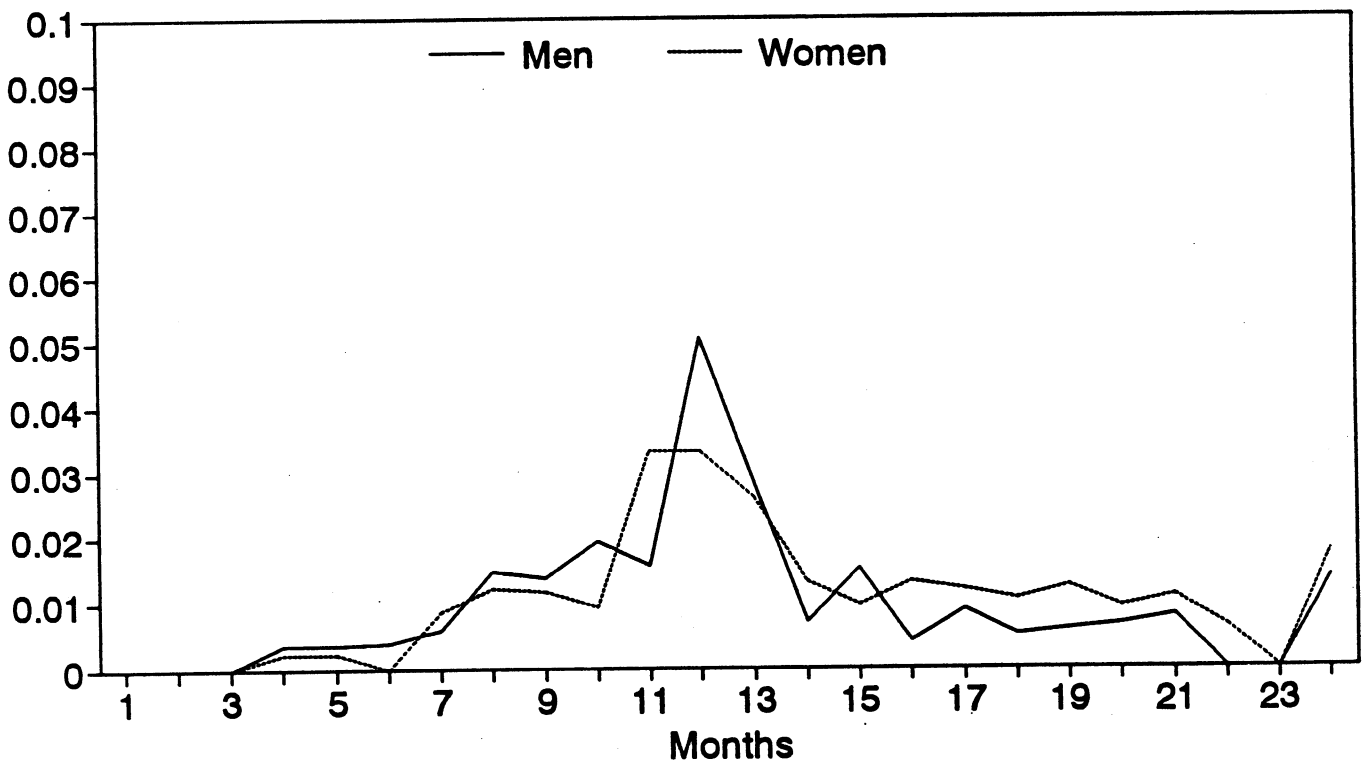
FIGURE 4
Repeat Use of UI Hazard: K-M Estimates



Source: National Longitudinal Survey of Youth

Notes: Spells are measured from the start of the first UI use to the start of the second UI use.

FIGURE 5
Repeat Use of UI Hazard: K-M Estimates



Source: 1990 Survey of Income and Program Participation

Notes: Spells are measured from the start of the first UI use to the start of the second UI use.

Emergency Extensions of Unemployment Insurance:
A Critical Review and Some New Empirical Findings

Stephen A. Woodbury

Michigan State University and
W.E. Upjohn Institute for Employment Research

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Starting with the recession of 1958, the potential duration of unemployment benefits that are regularly provided by states has been extended by six separate temporary Federal programs. These temporary extensions have been controversial because they raise a variety of questions about the optimal potential duration of unemployment insurance (UI): Should the potential duration of benefits be linked to labor market conditions, and if so, how should the link between potential duration and labor market conditions be made? That is, how should extended benefits be activated or "triggered"? Should the same eligibility conditions apply to extended benefits as to regular benefits? Should extended benefits be financed differently than regular benefits?

These questions about temporary benefit extensions have become increasingly controversial since 1970, when Congress passed a permanent "stand–by" Extended Benefits program. In principle, the EB program activates automatically when unemployment durations rise in the wake of a recession. But, especially since the "triggers" that activate EB were changed in the early 1980s, the EB program has activated only infrequently. For example, EB activated in only 10 states during the recession of the early–1990s, and never activated in some of the largest states that were hard–hit by the recession, such as California, New Jersey, New York, and Pennsylvania. As a result, four of the six temporary extensions have occurred since implementation of the permanent EB program, and each successive temporary extension has been increasingly complicated.

This paper provides background and a review of some of the analytical issues that arise in making policy on the potential duration of UI benefits. Section I reviews briefly the history of Federal extensions of unemployment insurance and illustrates the increasing complexity of each successive emergency benefit extension. Section II provides a brief review of how the potential duration of benefits is determined under regular state programs, since it is important to understand how potential durations are set in state programs as a background to accessing extended benefit programs.

The main analytical issue surrounding emergency extensions of unemployment benefits is whether (or to what extent) they create a disincentive for workers to seek reemployment, and hence lengthen spells of unemployment. Accordingly, the next three sections of the report focus on various aspects of this issue. Section III reviews the theoretical issues that arise in estimating the impact of extending the potential duration of UI benefits, and discusses the main class of model that has been relied on in grounding those empirical estimates.

Section IV provides a critical review of the empirical techniques that have been used to obtain estimates of the disincentive effects of increasing the potential duration of unemployment, pointing up the strengths and weaknesses of the various techniques. Section IV also provides a summary of the empirical estimates that have been obtained in past studies

Section V offers some new estimates of the impact of increased potential benefit duration on the duration of unemployment, using two data sets. In one of these data sets -- from Washington State in 1988-89 -- variation in the potential duration of benefits exists because the state provides greater potential duration of benefits to workers with stronger work histories. In the other data set -- from Illinois in 1984-85 -- variation in the potential duration of benefits exists because an emergency benefit extension program expired. This expiration allows a before-after comparison that yields an estimate of the impact of the emergency extension on the duration of unemployment. These are the two types of variation in potential duration of benefits that have been the basis of most estimates of the disincentives of benefit extensions. The main purpose of section V is to explore the sensitivity of the estimates to changes in model specification, estimating technique, and source of variation in potential benefit duration. The main question addressed is, How robust (or how fragile) are estimates of the disincentive effects of extending the potential duration of unemployment benefits?

Finally, section VI summarizes the main points and suggests some directions for future research. The main conclusion, perhaps, is that most existing research has avoided the difficult issue of testing the sensitivity of econometric estimates, tending instead to put forward one or another rather fragile set of estimates as representing the truth. In the process, many of the issues that are central to designing sensible and defensible extended benefit programs have been side-stepped, and research has focused instead on econometric issues that, while technically interesting, may yield only a modest return to policy.

I. A Brief History of Federal Extended Benefit Programs

Currently, the maximum potential duration of unemployment benefits provided to job losers by regular state programs is 26 weeks in all states except Massachusetts and Washington (where the maximum potential duration is 30 weeks). In 10 states, the potential duration of benefits is 26 weeks for all claimants who qualify for any benefits (Illinois and New York are the only large states that provide such "uniform" potential duration of benefits). In all other states, the potential duration of benefits varies with a claimant's work experience in the base period — roughly the year preceding the claim for benefits. The ways in which "variable" potential duration states compute the potential duration of benefits are described in section II below. The regular state programs is sometimes referred to as the "first tier" of the UI system.

Table 1 provides a summary of the main features of the six Federal programs that have temporarily extended the potential duration of unemployment benefits beyond the duration provided by state programs. The permanent stand-by Extended Benefit program is also summarized there. The stand-by EB program is often referred to as the "second tier" of the UI system, and emergency extensions are collectively referred to as the "third tier."

The first two Federal emergency benefit extensions — Temporary Unemployment Compensation (TUC) and Temporary Extended Unemployment Compensation (TEUC) — were enacted in 1958 and 1961. They were similar in that each lasted slightly over a year and extended the potential duration of benefits to workers who exhausted their regular state benefits by 50%, up to a maximum of 13 weeks. They differed, however, in that TUC was a voluntary program financed by interest–free loans to 17 participating states. TEUC, on the other had, was mandatory and was financed through increases in the Federal Unemployment Tax.

Nearly ten years after the first two emergency extended benefit programs, Congress enacted the permanent stand–by Extended Benefit program (EB). EB was modeled on TUC and TEUC in that it extends benefits to claimants who exhaust their regular state benefits by an amount equal to one–half their regular benefit duration, up to 13 weeks. Also, the weekly benefit amount is the same as the weekly benefit amount under the regular state program. EB is financed half–and–half by Federal and state revenues. It was originally activated either nationally by a "trigger" based on the national insured unemployment rate, or on a state–specific basis by state–level insured unemployment rates. The Federal trigger activated the program whenever the national insured unemployment rate reached 4 percent for a three-month period; the state trigger activated the program whenever a state's insured unemployment rate reached 4 percent for 13 consecutive weeks, and was at least 20 percent above the average insured unemployment rate of the corresponding 13-week periods in the two previous years.

As shown in Table 1, in 1980 and 1981, the national trigger was dropped and the state–level trigger was raised from an insured unemployment rate of 4 percent to a rate of 5 percent. Both changes made it less likely that EB would be activated in a recession. Also, the amendments of the early-1980s made eligibility for EB more restrictive -- the program now requires that workers have at least 20 weeks of work (or the equivalent) in the base period to qualify for EB. Combined with falling insured unemployment rates, which have resulted mainly from decreased participation in unemployment insurance, the changes of 1981 led to a situation in which EB was nearly defunct by the time of the recession of the early 1990's. As already noted, EB was activated in only 10 states during that recession and failed to activate in several states where many observers felt labor market conditions were bad enough to warrant it. In response to the failure of EB to be activated widely during the early–90's recession, Congress passed legislation allowing states to adopt an alternative trigger based on the total unemployment rate (TUR) in 1993, although few states have adopted the alternative trigger (see below).

States were allowed to adopt EB as early as October 1970, and required to do so no later than January 1972. But even before EB became available in all states, Congress adopted the Emergency Unemployment Compensation Act (sometimes called "Temporary Compensation" or "TC"), which provided up to 13 weeks of extended benefits to claimants who either exhausted EB or exhausted regular benefits in states where EB was not available. Temporary Compensation was activated by special triggers that differed from the stand–by EB triggers. It was financed from Federal Unemployment Tax revenues in the Extended Unemployment Compensation Account (EUCA). The program, which originally was set to run from January 1972 until September 1972, was extended through March 1973.

During the severe recession of mid–1970s, the national EB trigger activated the Extended Benefits in all states, permitting workers to receive up to 26 weeks regular unemployment benefits followed by up to 13 weeks of EB. Nevertheless, the recession was so severe that Congress enacted another emergency extension in January 1975 -- Federal Supplemental Benefits (FSB), which provided up to 13 additional weeks of benefits to those who exhausted regular benefits and EB.

In March 1975, the FSB program was extended and made more generous by providing yet another 13 weeks of benefits. As a result of this and further extensions of FSB, a claimant could receive up to 65 weeks of unemployment benefits for the period March 1975 through March 1977 — 26 weeks of regular state benefits, 13 weeks of EB, and 26 weeks of FSB.

In April, 1977, FSB was extended again (through January 1978), but the potential duration of benefits was reduced to 13 weeks from May 1977 through the end of the program. This extension also added special federal disqualifications for refusal of suitable work and failure to actively seek work, defined suitable work for the FSB program, and added special penalty and repayment provisions for fraudulent acts on the part of both claimants and employers. This was the first time such disqualifications had been imposed as part of an emergency extension.

As already noted, Congress eliminated the national trigger for EB in 1980, and increased the rate of insured unemployment needed to activate EB in a recession. In addition, Congress changed the definition of insured unemployment to omit EB claimants from the computation, and imposed special eligibility and disqualifying conditions on EB claimants. All of these changes reflected a changed attitude toward extended benefits, one that suggested an intent by the new Reagan Administration and Congress to reduce the cost of domestic programs. These changes clearly did reduce the cost of EB -- indeed, they very nearly disabled the program. But ironically, the parade of emergency unemployment benefit extensions continued.

In 1982, Congress enacted Federal Supplemental Compensation (FSC) as part of the Tax Equity and Fiscal Responsibility Act of 1982. FSC was different from previous emergency extended benefit programs in that the number of weeks payable in each state varied according to different criteria at different times. In fact, FSC went through four "phases," each of which provided different potential benefit durations for each state depending on the state's labor market conditions (see Table 1, under "potential duration of extended benefits provided"). Under Phase II, a UI claimant in a high unemployment state could be eligible for up to 55 weeks of benefits -- 26 from the regular state program, 13 from EB (assuming the state had triggered on), and 16 from FSC.

Potential durations were somewhat shorter under Phases III and IV of FSC, but the interstate differences in potential benefit durations remained. Under FSC, then, there was more tinkering (or, more charitably, greater effort to fine-tune the program) than under previous emergency extensions in two senses. First, the idea that emergency extensions should provide different potential benefit durations to different states was wholly new -- even the stand-by EB program has never done this. Second, four phases of FSC led to frequent changes in potential benefit duration and created administrative difficulties for the states. Both of these aspects of FSC began to call into question the roll of emergency extensions and seemed to be an admission that the stand-by EB program was already defunct.

The most recent emergency extension of unemployment benefits, Emergency Unemployment Compensation (EUC), was enacted in November 1991 after months of foot-dragging by the Bush Administration, which had vetoed several earlier emergency extensions. EUC was the most complicated emergency benefit extension of all: it went through *five* phases, provided different potential durations across states at a given time, and different potential durations within a state over time (see Table 1). The potential duration of benefits within a state could change either because of Congressional fiat (that is, a change from one phase to the next), or because a state changed its classification as either high-unemployment or low-unemployment. By all accounts, EUC was a UI administrator's nightmare. In Pennsylvania, for example, the potential duration of benefits changed *nine* times between November 1991 when EUC became effective and February 1994 when Phase V of EUC terminated. Five of these changes resulted from enactment of EUC or a change from one phase to another, and four resulted because Pennsylvania was reclassified as high- or low-unemployment. At one point, Congress let EUC lapse, but subsequently resuscitated it, and during the hiatus, state administrators were left hanging.

During Congressional debate on whether to extend EUC, Republicans in Congress argued that if Congress continued its pattern of enacting emergency extensions whenever the economy went into recession, then there would be no incentive for the states to switch to the new alternative EB trigger, based on the total unemployment rate (TUR). The old insured unemployment rate trigger, as already discussed, has been ineffective since the early-1980s and rarely moves a state onto EB, whereas the alternative TUR trigger would be more effective. But states naturally prefer to have the Federal government step in and provide emergency benefits, since financing of emergency benefits is wholly Federal, rather than 50-50 state-Federal as with EB. As long as the states can argue that EB is not providing adequate benefit durations, they can reasonably urge Congress to enact emergency extensions. And as long as Congress accommodates the states in enacting emergency extensions, the states have no incentive to switch to the alternative TUR trigger, which would be more effective but would also result in greater benefit payments from the state UI trust funds.

A cynic might argue that Congress really does not want the stand-by EB program to work effectively -- that members would prefer to step in and enact an emergency program whenever the economy slumps. An emergency program shows that Congress has "done something" in an economic downturn and offers the politicians a concrete program to point to when they stand for reelection. Such a cynical view is not wholly unrealistic. Congress could require the states to switch to the alternative TUR trigger, but it has not done so.

The future of the EB program and emergency extensions is highly unclear at this time. Congress seems to pay attention to the Unemployment Insurance System only when there is a recession, so the role of politics would seem to be more important than the role of policy analysis in determining the future of extended benefits. It needs to be noted that very little effort has been devoted to understanding what is (or would be) the socially optimal potential duration of benefits, or to analyzing the extent to which that optimal potential duration should change with changing labor market conditions. These gaps, convincingly addressed, could have an impact on policy and the future direction of unemployment insurance in this country.

II. How States Determine the Potential Duration of Benefits

From the beginning of the UI program in the United States, the generally accepted goal has been to provide a limited number of weeks of benefits, payable only long enough to tide an unemployed worker and household over a temporary spell of unemployment. Consensus on the meaning of "temporary" has changed -- from 15 weeks, which was the most common potential duration at the beginning of the program in 1935, to 26 weeks, which is the maximum in all but two states today.

The apparent consensus that 26 weeks is a reasonable duration of benefits masks considerable variation among the states in how the duration of benefits is determined. Some states provide the same duration of benefits to all eligible claimants, whereas others vary benefit duration with the extent of a claimant's past employment or wages. Accordingly, there are substantial differences among the states in the amount of prior work or wages required to qualify for different benefit durations.

Table 2 provides a summary of the methods used by the states to determine the potential duration of benefits. As can be seen in the first two columns, nine states currently provide the same potential duration of benefits to all who meet the minimum qualifying requirement (that is, the minimum and maximum potential durations are the same). These are usually referred to as uniform duration states. The number of states providing uniform duration has fallen over the years, as Blaustein (1993, Table 10.7, p. 304) has shown.

The other 44 states vary potential duration according to each claimant's past employment or earnings. These states use one of two methods to compute potential duration. In 6 states -- Florida, Michigan, New Jersey, Ohio, Oklahoma, and Pennsylvania -- potential duration is an increasing function of the number of "credit" weeks worked (or wages, in the case of Oklahoma) in the base period, up to the maximum 26 weeks. A credit week is a week in which earnings equaled or exceeded some specified minimum, so that,

(2.1) $$D_{pot} = \min[\, f(\text{credit weeks}), 26],$$

where D_{pot} denotes the potential duration of UI benefits and f is a function increasing in credit weeks. For example, in Florida, a credit week is a week in which a worker earned at least $20, and potential duration equals one-half the number of credit weeks.[1] It follows that, in order to be eligible for the maximum potential duration of 26 weeks of benefits, a worker must have 52 credit weeks; that is, the worker must have worked in every week of the base period.

In 38 states, the potential duration of benefits depends on the ratio of a claimants' base-period earnings to high-quarter earnings, up to the maximum 26 weeks. If we let BPE denote base period earnings and HQE denote high-quarter earnings, then,

(2.2) $$D_{pot} = \min[\, f(BPE/HQE), 26],$$

where f denotes a function increasing in BPE/HQE. Note that BPE/HQE ranges from 1 for a worker all of whose base period earnings were earned in a single quarter (BPE = HQE for such a worker) to 4 for a worker who had identical earnings in all four quarter (BPE = 4[HQE]). The idea here is that a worker with stable earnings throughout the base period will have a higher BPE/HQE and hence a higher potential duration of UI benefits.

In 5 states, the relationship between BPE/HQE and potential duration is explicit. For example, in North Carolina, potential duration is simply 8.67 times BPE/HQE (up to 26 weeks), so that a UI-eligible worker with BPE/HQE of 3 or greater is eligible for the maximum potential duration of 26 weeks of benefits.

In 33 states, however, the relationship between BPE/HQE and potential

[1] In Florida, a worker must have at least 20 weeks in which there were earnings of at least $20 to be eligible for any UI benefits, so the minimum potential duration of UI benefits is 10 weeks.

duration is masked by the formula used to calculate potential duration. In these states, potential duration is calculated as some fraction a of base period earnings divided by the weekly benefit amount (WBA), up to the maximum:

(2.3) $$D_{pot} = \min [\, a(BPE)/WBA; 26].$$

The parameter a limits the total UI benefits paid to a worker in the benefit year to some fraction of base period earnings. In 18 states, a = 1/3, and in the other 15, a ranges between .25 and .6. What needs to be noted is that in all of these states the weekly benefit amount is computed in turn as a fraction of high-quarter earnings (or in some cases, average earnings in the two highest quarters of the base period) up to some maximum:

(2.4) $$WBA = \min [\, b(HQE), WBA_{max}].$$

Typically, b is 1/25 (.04), so that the weekly benefit amount equals one-half of average weekly earnings in the high quarter. [The parameter b ranges from 1/26 (.038) to 1/20 (.05) in these 33 states.] Substituting the WBA equation (2.4) into the potential duration function (3) yields:

(2.5a) $$D_{pot} = a(BPE)/b(HQE), \qquad \text{if } WBA < WBA_{max},$$

or

(2.5b) $$D_{pot} = a(BPE)/WBA_{max}, \qquad \text{if } WBA = WBA_{max}.$$

It follows that for eligible claimants whose WBA is less than the state's maximum,

(2.6) $$D_{pot} = g(BPE/HQE),$$

where

(2.7) $$g = a/b,$$

so the dependence of potential duration on BPE/HQE is clear for claimants whose WBA is below the maximum. For claimants whose WBA is at the maximum, potential duration will still depend on the relationship between base period and high-quarter earnings. For example, a worker who obtains the maximum WBA as a result of high earnings in just one quarter may have potential duration below the maximum 26 (or 30) weeks, since base period earnings will be low relative to the weekly benefit amount for such a worker.

The parameter g can be usefully interpreted as an index of a state's duration generosity. Specifically, it gives the increase the number of weeks of potential duration that result from a unit increase in BPE/HQE. In Table 2, g has been computed for all 53 "states" (that is, UI jurisdictions). For states that do explicitly use the parameters a or b in computing the potential duration of benefits, an implied g has been calculated numerically.

Table 2, also displays the minimum base period earnings and high-quarter earnings that an eligible claimant would need in order to receive the state's maximum potential duration of benefits.

An examination of g and the minimum earnings required for maximum potential duration in Table 2 shows that the variations in states' duration provisions are significant. Claimants with similar base-period work experience qualify for quite different potential durations depending on the state in which they reside, and the requirements for 26 weeks of regular benefits vary dramatically among the states. For example, to qualify for 26 weeks of regular benefits requires as little as \$130 in the base period (and \$33 to \$105 in the high-quarter) in Hawaii to as much as \$18,757 in the base period (and \$4,800 in the high quarter) in Indiana.

Variable duration reflects the notion that workers "earn" their rights to benefits by working, and that each week of benefits is earned by a given number of weeks of employment or earnings. The widespread use of variable duration also reflects two further concerns: first, that uniform duration is more expensive than variable duration, and second, that uniform duration can generate a high ratio of total benefits paid to base period earnings, which could in turn lead to strong work disincentives.[2] These issues are explored further below.

[2] See Advisory Council on Unemployment Compensation (1995, p. 129) for a discussion replacement rates based on the administrative records of six states. Although the ACUC discussion is based on a different definition of the replacement rate (the ratio of weekly benefits to average base period earnings) than the definition used in the text, it does suggest that benefit durations in excess of base period employment durations would imply strong work disincentives.

III. Theoretical Issues

Most estimates of the effects of both benefit duration and benefit amounts on the duration of joblessness have been based on one or another model of job search (see Mortensen 1986 for a review)[3]. The job search models provide a theoretical link between the duration of joblessness, on the one hand, and job–search intensity, individual characteristics, and labor market conditions, on the other. It is useful to review a general job-search model as a prelude to the empirical work that is reviewed and developed below.

Let T denote the week in which a UI recipient returns to work, and let P_t denote the probability that a UI recipient returns to work in week t, given that she has not already returned to work by then; that is, $P_t = Pr[T{=}t|T{\geq}t]$. Then P_t can be expressed as the product of (a) the probability of receiving a job offer in week t and (b) the probability of accepting that job offer, given that an offer has been made. The probability of receiving an job offer in week t (J_t) depends on the intensity of the worker's job search (i) and a vector of characteristics of the worker that determine the demand for the worker's labor (c); that is, $J_t = J_t(i,c)$. The probability of offer acceptance (A_t) depends on whether the offered wage (w^o) equals or exceeds the worker's reservation wage (w^r); that is, $A_t = Pr[w^o{\geq}w^r|J_t]$. Hence, the probability of finding reemployment during week t (given that reemployment has not already occurred) can be expressed as:

(3.1) $P_t = J_t(i,c)\ A_t.$

[3] The income-leisure model developed by Moffitt and Nicholson (1982) is also appealing because of its link to well-known principles of consumer theory, but it has been the basis of relatively little empirical work. One reason may be that the income-leisure model implicitly views unemployment as compensated leisure, and views the combination of income and leisure (that is, unemployment) as chosen by the worker subject to the constraints posed by the available wage rate and the UI system. Since the extent to which unemployment is voluntary is itself an important question, a model that assumes that unemployment is voluntary may be rather uninformative.

P_t has been defined in discrete time above. In the limit, as the time interval over which reemployment is measured approaches zero, P_t becomes an instantaneous rate of reemployment, or hazard rate, h(t). The hazard rate is linked to unemployment duration in the following way. If t has cumulative distribution F(t), and frequency distribution f(t), then $h(t) = f(t)/[1 - F(t)] = f(t)/S(t)$, where S(t) is the so–called survivor function, or the probability of being unemployed to time t (Lancaster 1979). The survivor function can also be expressed in terms of the hazard: $S(t) = f(t)/h(t)$. Thus, the hazard rate is inversely related to the survival probability, and any factor that increases the hazard rate should decrease expected unemployment duration.

Equation (3.1) highlights the fact that longer spells of joblessness can result from less–intense job search, from individual characteristics (c) that imply lower demand for a worker's services, or from a lower probability of job–offer acceptance. From the point of view of work disincentives, the probability of offer acceptance, A_t, and the intensity of job search, i, are central, since they depend on the generosity and potential duration of UI benefits.

IV. Potential Duration of Unemployment Benefits and the Duration of Joblessness: Models and Existing Estimates

Since the mid–1970s, the most–researched question about the Unemployment Insurance (UI) system has been whether and to what degree higher weekly UI–benefit amounts lengthen UI recipients' jobless spells. But an equally important and under–researched question is how the potential duration of those benefits influences the length of jobless spells. The latter question is important for two reasons. First, as already discussed, Congress has legislated six temporary or emergency extended UI–benefit programs since the 1950s, making the potential duration of UI benefits a highly variable aspect of the UI system. Second, as will become clear, econometric problems make inferences about the influence of UI–benefit extensions on the expected length of UI recipients' jobless spells especially tenuous.

This section offers a summary of the evidence on the disincentives effects of extending the potential duration of benefits by an additional week. Rather than simply provide a range of estimates, though, an effort is made to provide some insight into the quality of the existing evidence. Subsection A begins with a discussion of the problems that arise in using censored data, which has generally been used in obtaining estimates of the disincentives of unemployment insurance. Subsection B then reviews several models that have been used to infer the effects of extended benefits -- a simple linear duration model, a parametric jobless duration model that accounts for censoring of the dependent variable and non–normality of the error term, and a semi–parametric model of the conditional probability (or hazard) of returning to work. Subsection B also summarizes the estimates that have been derived from each of the models.

A. Censoring Problems

Most analyses of the effects of potential benefit duration on jobless duration have used claims and benefits data from UI administrative files. These data are extremely rich: For example, they usually contain demographic data on claimants, the dates of their UI claims, and the amount and timing of benefits received. But claims and benefits data from UI administrative files are usually deficient in that they exclude any information on the subsequent earnings of claimants. Hence, they fail to offer data on actual spells of unemployment. Rather, they indicate only the duration of *insured* unemployment experienced by a claimant.

In some data sets -- including both the Illinois and Washington State data used in section V, this deficiency can be overcome to some extent by using data from Unemployment Insurance Wage Records, which contain information on the earnings histories of workers both before and after their spell of insured unemployment. By matching a claimant's Wage Records to his or her claims and benefits data, it is possible to determine whether a spell of insured unemployment was followed by a period of earnings. If the observed spell of insured unemployment was followed by a period of earnings, then it can be inferred that the insured spell and the actual spell of joblessness were the same. On the other hand, if the insured spell was not followed by a spell of earnings, the insured spell must be considered a censored or truncated measure of the actual spell of joblessness.

It was rare for early studies of the disincentive effects of UI to make use of Wage Records, as will be done in Section V below. As a result, existing research using administrative data has necessarily taken a different approach to drawing inferences about actual spells of joblessness from observations on insured unemployment spells. First, because the number of weeks of unemployment that can be observed in administrative data is limited by the potential duration of UI benefits, it has been assumed that any spell of insured unemployment that is at the maximum potential is an incomplete spell. For example, a worker eligible for 26 weeks of state regular benefits who is observed receiving 26 weeks of UI benefits would be considered to have had a jobless spell of greater than 26 weeks. Second, and conversely, a workers eligible for 26 weeks of benefits who is observed receiving less than 26 weeks of UI benefits would be considered to have had a jobless spell of exactly the observed length.

Neither of these assumptions is necessarily correct, as can be seen in Table 3, which uses a random sample of administrative data on male UI recipients in Illinois during 1984-85 to illustrate four cases, labeled A through D.[4] Cases A and B are those of workers who received the maximum potential weeks of UI benefits—that is, exhausted their benefits. It is possible for such workers to return to work immediately after receiving their last benefit payment (Case A), or to continue to be out of covered employment (Case B, which implies either continuing to seek employment or dropping out of the labor force after receiving the last benefit payment).[5] The usual assumption is that all workers who exhaust benefits continue without covered employment, as in Case B. But the right–most column of Table 3 shows that this assumption is incorrect for nearly 40 percent of the workers who exhausted their benefits in this sample. That is, 283 of the 717 workers who exhausted their benefits returned to work immediately (or very shortly) after receiving their last benefit payment.

It is also possible to misclassify a worker who did not exhaust his or her benefits. Cases C and D in Table 3 are for workers who received fewer than the potential weeks of benefit payments. Again, the usual assumption is that all such

[4] The data are for men in the control group of the Illinois Reemployment Bonus experiment. These data are described further in section V, where they are used further.

[5] Obtaining uncovered (usually underground) employment and moving out of state are additional possibilities.

workers returned to work immediately after they stopped receiving benefits, as in Case C. But the right–most column shows that 405 (or 28 percent) of the 1,445 workers who ended their benefits before exhausting did not return to covered employment.[6]

The problems of using censored data to infer the effects of extended UI benefits on expected unemployment duration can also be illustrated using the Illinois data that underlie Table 3. About one-half of the sample used in Table 3 was drawn before expiration of Phase IV of the Federal Supplemental Compensation (FSC) program in late 1984, and the other half was drawn after FSC expired. Phase IV of FSC in Illinois provided an additional 12 weeks of potential benefits to initial claimants who satisfied the usual state eligibility criteria plus a somewhat more stringent monetary eligibility criterion that was specific to FSC.[7] As a result, the workers sampled, all of whom were eligible for regular state benefits, can be divided into four categories: (a) those who were eligible for FSC because they met the additional monetary eligibility criteria for FSC and filed their initial UI claim while FSC was still in effect; (b) those who were monetarily eligible for FSC, but claimed benefits too late to actually receive FSC; (c) those who were monetarily ineligible for FSC and filed their initial claim before FSC expired; and (d) those who were monetarily ineligible for FSC, and filed their initial claim after FSC expired.

Table 4 shows the mean insured unemployment duration for each of these four groups. The expiration of FSC appears to offer a natural experiment. The mean unemployment duration of workers eligible for FSC (21.4 weeks) can be compared with the mean unemployment duration of workers monetarily eligible but temporally ineligible because they filed after FSC expired (17.9 weeks). As a quasi–control, the mean unemployment duration of workers who were monetarily *in*eligible but who filed

[6] Most likely, these workers either dropped out of the labor force or took uncovered employment, although it is possible that they stopped participating in UI and continued to seek employment. There is no way of distinguishing between the two possibilities in the administrative data.

[7] That is, not all UI claimants who were monetarily eligible for regular state benefits were also monetarily eligible for FSC.

for benefits while FSC was still in effect (16.5 weeks) can be compared with the mean unemployment duration of workers who were neither monetarily nor temporally eligible (20.1 weeks).

Two comparisons are shown in the bottom row of Table 4 (labeled "Difference"). The difference between the two groups of monetarily eligible workers, 3.5 weeks, suggests that FSC prolonged unemployment spells significantly. Moreover, the difference between the two groups of monetarily ineligible workers, –3.5 (with a large standard error), suggests that there was no underlying macroeconomic or other reason for expecting unemployment spells to be longer after the expiration of FSC. The conclusion would seem to be that workers eligible for FSC tended to take over three weeks longer to return to work than did workers who were not eligible for FSC.

Such an inference would clearly be wrong, though, because the claims and benefits data make it impossible to observe more than 26 weeks of unemployment among FSC-ineligibles, whereas we can observe up to 38 weeks of unemployment among FSC-eligibles. The truncation or censoring of unemployment spells at the maximum potential duration leads to a situation in which the two group means cannot be compared. To take Moffitt's (1985a) extreme example, every worker in each of the two groups might have an actual spell of joblessness of 30 weeks, but we would observe an average of 26 weeks for the first group (because the data are censored at 26 weeks) and 30 weeks for the second (because censoring occurs only at 38 weeks). It may still be the case that FSC tended to lengthen jobless spells, but the data in Table 4 cannot be used to make such an inference.

In the presence of censoring, quasi–experimental comparisons like those presented in Table 4 fail to yield reliable estimates of the effects of extended benefits on jobless duration. Hence, other methods of inference must be considered.

B. Models of Unemployment Duration and Reemployment Hazard

Estimates of how potential benefit duration affects the expected duration of joblessness have progressed through three stages. This section outlines the approach represented by each of these stages and summarizes past studies that have used each of the methods. Table 5 provides a synopsis of the various studies that are referred to.

1. *Linear Models of Insured Unemployment Duration*. The earliest empirical work on the effects of potential duration on expected jobless duration simply regressed the duration of insured unemployment in weeks (D), or the natural logarithm of weeks of unemployment on appropriate explanatory variables ($x_1, ..., x_K$), including measures of the replacement ratio and potential duration of benefits:

(4.1) $D = a_0 + a_1x_1 + ... + a_Kx_K + u,$

where u is assumed to be a normally distributed disturbance term. The coefficients of x_1 through x_K provide an estimate of the relationship between the explanatory variables and weeks of insured unemployment. Studies taking this approach include Ehrenberg and Oaxaca (1976), and Holen (1977), among others; however, only Holen estimated the effect of additional weeks of benefit entitlement. Her estimates suggest that a 1-week increase in the potential duration of benefits increases unemployment duration by about .8 week -- a very high estimate.

It seems unlikely that such estimates can be relied on for convincing assessments of the behavioral impact of an additional week of benefit entitlement on the duration of unemployment. The model used takes no account of censoring in the data, so that any measured impact of longer benefit entitlement could simply be the result of the ability to observe more weeks of unemployment for workers whose benefit entitlement is longer, as discussed in section A above. The estimates provided by such studies do provide accurate *descriptive* evidence on extended benefits. That is, they they do give an unbiased and consistent estimate of the average weeks of extra benefit payments that are paid to workers who receive an additional week of benefit entitlement. But this descriptive estimate cannot be used to infer how an increase in the potential duration of benefits would change the behavior of workers. As a result the descriptive estimate cannot be used to predict how unemployment durations would increase if benefits were extended.

2. *Parametric Models of Time to Reemployment.* The problem with applying Ordinary Least Squares (OLS) to equation (4.1) is that the error them u in the equation is not normal, as OLS requires. There are two reasons for this. First, as already discussed, D is a censored measure of actual jobless duration, since each worker is eligible for a specified maximum number of weeks of benefits. As a result, the distribution of D is truncated at the maximum benefit duration. Realization of this problem lead to some studies that assumed that the underlying distribution of jobless spells is normal, and assumed in turn that the distribution of u in equation (4.1) is truncated normal. For example, Classen (1979) and Newton and Rosen (1979) both used Tobit analysis—which assumes that u has the truncated normal distribution — to correct for the truncation of the dependent variable. Classen's estimates suggest that an additional week of potential benefit duration leads to at most an additional 0.12 week of insured unemployment, whereas Newton and Rosen's estimates suggest an additional .6 week (see Table 5).

The second reason for questioning the assumption that u in equation (4.1) is normal is that the empirical frequency distribution of weeks of insured unemployment in most data is not bell-shaped, as the normality assumption requires. Rather, it shows one spike at zero weeks of unemployment, and falling frequencies for greater unemployment durations, until a spike appears where censoring occurs (that is, at maximum benefit duration.) Except for the spike at the censoring point, the empirical distribution looks much like an inverse exponential. This latter problem can be solved in a jobless–duration equation like (4.1) by making an appropriate assumption about the distribution of u, and estimating equation (4.1) under that alternative distributional assumption. The Weibull distribution has been widely assumed in studies of jobless duration because it provides an approximation to the empirical distribution of jobless duration that appears to be valid (Lancaster 1979). (The exponential distribution is a special case of the Weibull. Whereas the exponential restricts the conditional probability that a UI recipient will become reemployed (P_t, or the hazard rate) to be constant over the spell of unemployment, the Weibull allows for the possibility that a UI recipient's probability of reemployment rises or falls over the spell. The greater generality of the Weibull distribution makes it the preferred choice.)

Several studies have imposed a more appropriate distributional assumption on u in equation (4.1) to examine the effects of potential benefit duration on the length of unemployment spells. These studies have obtained estimates suggesting that an additional week of benefit entitlement increases the duration of unemployment by about .2 to .4 week. For example, Katz and Ochs (1980) estimate that an additional week of benefit entitlement increases the duration of unemployment by .17-.23 week. Moffitt (1985b) estimates an additional .45 week for men and an additional .28 week for women using a 15-state sample. Using Georgia data, he obtains an estimate of .17 week for men and .37 for women. Solon (1985) also uses Georgia data and estimates that an additional week of potential benefit duration leads to 0.36 additional weeks of insured unemployment.

Such estimates are far more convincing than those based on Ordinary Least Squares, in that they take account of the censoring of data and make a defensible assumption about the distribution of unemployment spells. Nevertheless, they have been criticized for imposing any distributional assumption, and because they are unable to take account of factors that change during a spell of unemployment and that may influence the ultimate duration of unemployment

3. *Semi–Parametric Hazard Models.* The two problems just mentioned cannot be handled in either of the duration modeling frameworks discussed to this point. They are important enough to discuss in somewhat more detail. The first is that the duration models force an assumption about the distribution of the error term u in equation (4.1). Incorrect distributional assumptions may yield misleading inferences about the effects of extended benefits. For example, the Weibull seems a good approximation to the empirical distribution of jobless spells as long as it is true that the spike in the empirical distribution in the week following benefit exhaustion results from censored data. But if the distribution of jobless spells shows a true spike in the week following benefit exhaustion—that is, if workers tend to put off finding taking a job until just after their benefits terminate—then the Weibull is a poor choice. Ideally, one would like to impose no distributional assumption at all.

The second problem is that some variables may change during a worker's spell of joblessness. For example, the number of weeks until exhaustion of benefits can be thought of as a variable that decreases weekly. There is no way of understanding the effects of such "time–varying" explanatory variables in a duration model.

To analyze the effects of time–varying explanatory variables and to avoid any assumptions about the distribution of jobless spells requires reconceptualizing the duration problem as a problem of rate of escape from joblessness. In other words, rather than regress some measure of duration on various explanatory variables, one could regress a dummy variable (R_t) equal to one if a worker escaped from unemployment in week t (zero otherwise) on various explanatory variables, some of which are time–invariant ($x_1, ..., x_K$), and others which are time–varying ($z_1(t), ..., z_N(t)$).

Since the dependent variable is a measure of a worker's probability of escaping joblessness in week t (given that the worker was "at risk" of escaping joblessness at the beginning of the week), it is appropriate to interpret the dependent variable as a hazard rate, and to call it h(t). Hence, this model can be written:

(4.2) $R_t = h(t) = b_0 + b_1x_1 + \ldots + b_Kx_K + c_1z_1(t) + \ldots + c_Nz_N(t) + e.$

If e is assumed normally distributed, then this a linear probability model, which is useful in exploratory work.[8]

Each coefficient in equation (4.2) represents the change in the probability of reemployment that results from a unit change in the independent variable. For example, if x_1 were age in years, then b_1 would show the change in probability of reemployment associated with an additional year of age. This change in probability would be assumed constant over the spell of unemployment (unless age were interacted with a time–varying explanatory variable). Note that a positive coefficient indicates a higher probability of returning to employment, and hence a shorter duration of unemployment.

Whereas the unit of observation in the various duration models represented by equation (1) is the *claimant*, the unit of observation in a hazard model is the *claimant–week*. The transformation of claimant records into claimant–week records is illustrated in Table 6. Panel A shows records for three claimants, who experienced 4, 38, and 0 weeks of insured unemployment. The first and third claimant became reemployed after receiving their last UI benefit payment, whereas the second remained jobless. Also, the first and third were ineligible for FSC, whereas the second was eligible.

The second panel of Table 6 shows the claimant–week records that are

[8] If e is assumed to have the logistic distribution, then we have a logit model, which is preferred to a linear probability model because it yields, consistent, and efficient coefficient estimates. In practice, logit estimates of equation (4.2) are virtually identical in statistical significance and quantitative response to changes in explanatory variables to linear probability estimates.

generated by the three claimant records in Panel A. The first claimant contributes a total of 6 observations to the claimant–week data set—one for the waiting week, one for each week in which UI benefits were received, and one more for the week following the spell of insured unemployment, since this worker became reemployed. The dependent variable in the hazard analysis, reemployment, is zero in all weeks except the last, in which reemployment occurred. The second claimant contributes 39 observations to the claimant–week data set—one for the waiting week, and one for each week in which UI benefits were received. The reemployment variable is zero for all of these observations, and since this claimant did not find reemployment after exhausting his UI benefits, there is no fortieth observation following the spell of insured unemployment in which the reemployment variable equals one. Note that, when claimant records are transformed into claimant–week records, each claimant contributes exactly as much information as is known about him or her to the analysis of reemployment probability (Allison 1982).

Hazard models such as (4.2) start from the pioneering work of Cox (1972), and are often referred to as "semiparametric" because they implicitly make no assumption about the distribution of u in the duration equation (4.1). Studies that have estimated hazard models such as (4.2) that also provide estimates of the effects of increases in potential benefit duration on jobless duration include Moffitt (1985a, 1985b), Ham and Rea (1987), Grossman (1989), and Katz and Meyer (1990). The estimates provided by these studies are wide-ranging: The estimates in Moffitt (1985a, 1985b) and Katz and Meyer (1990) suggest that a one–week addition to potential duration leads to an increase in unemployment duration of only .15 to .2 week. Ham and Rea's (1987) estimate of .26-.35 week is somewhat higher. Grossman's (1989) estimate of .9 week, derived from Phase IV of FSC, is the highest of all.

V. How Robust Are the Estimates? Some Exploratory Findings

The estimates of the impact of extending UI benefits reviewed above are based on different estimating techniques, various data sources, and various specifications of the incentives (or disincentives) facing UI claimants. The estimates vary widely, from virtually no impact of extending the potential duration of benefits to an increase in unemployment duration of .9 week for each additional week of benefit eligibility. This is a disturbingly wide range of estimates. In order to get a better understanding of how robust (or how fragile) these estimates are, this section reports results from two data sets that are typical of the data used in the studies reviewed above. The first are data from the Washington Reemployment Bonus experiment, which was conducted during 1988-89, and was evaluated using administrative data from the Washington State Unemployment Insurance system (see Spiegelman, O'Leary,and Kline 1991 for a full description). The advantage of using the Washington data is that Washington is a variable duration state -- that is, the potential duration of UI benefits depends on the earnings history of workers during their base period, and varies from a low of 10 weeks to a high of 30 weeks.[9] Since many estimates of how the potential duration of benefits affects unemployment come from data in which the main source of variation in potential duration occurs within-state (that is, under a given "regime" of potential duration), it seems useful to explore the extent to which various model specifications yield different findings in such a setting. The workers examined below are the 9,982 UI–eligible claimants who filed valid claims and were assigned either to the control group or to one of three treatments that offered low bonuses and had on impact on behavior.

[9] The distribution of potential durations is highly skewed in Washington: Less than 1 percent of eligible claimants have potential duration of 16 or fewer weeks; about 15 percent are eligible for 17 to 21 weeks (about 3 percent each for 17, 18, 19, 20, and 21 weeks); about 28 percent are eligible for 22 to 28 weeks (about 4 percent each for 23, 24, 25, 26, 27, and 28 weeks); about 5 percent are eligible for 29 weeks, and 51 percent are eligible for 30 weeks.

The second data set used is from the Illinois Reemployment Bonus experiment, which was conducted during 1984-85, and like the Washington experiment, was evaluated using administrative records of the State of Illinois Unemployment Insurance system (see Woodbury and Spiegelman 1987 for a complete description). The advantage of the Illinois data is that they span the expiration of one of the the emergency UI benefit extensions -- the Federal Supplemental Compensation program (FSC), which expired about half-way into the enrollment period of the Illinois bonus experiment. In Illinois, Phase IV of FSC provided 12 weeks of Federal benefit eligibility on top of the 26 weeks of regular state benefit eligibility that Illinois provides (Illinois is a uniform duration state). Consequently, the Illinois data permit one to compare the jobless spells of workers who were eligible for FSC (that is, were eligible for a total of 38 weeks of benefits, and knew this at the time they filed their initial claim) with the spells of workers who were eligible for only 26 weeks of benefits, but who would have been eligible for an additional 12 weeks if they had become unemployed and filed for benefits only one to six weeks earlier.

It is important to remark that, although FSC was terminated by Congress because national labor market conditions had improved following the severe recession of the early 1980s, it did not "trigger off" as the standing Extended Benefits program would have done. That is, whereas EB would have triggered off when specific conditions in Illinois had improved, FSC ended by Congressional fiat and in response to the impression that labor market conditions nationwide no longer required a Federal emergency benefit program. Accordingly, the expiration of FSC provides a "natural experiment" and an alternative method of making inferences about the independent contribution of potential benefit duration to the length of a jobless spell -- one that might be more convincing than inferences from data where potential duration is correlated with work history (as in the Washington data). The workers examined below are the 7,443 UI–eligible claimants who filed valid claims and were assigned to either the control or Employer experimental groups of the experiment.[10]

Both the Washington and the Illinois data are typical of the data used in past studies of the effects of extended benefits on the duration of unemployment because they are administrative data—that is, data gathered and maintained by agencies responsible for administering the UI program. But both the Washington and the Illinois data include both earnings history data, which allow an improved classification of jobless spells as complete or censored.[11] Most of the studies described and reviewed in section IV had to impose what appear to be erroneous assumptions about whether a spell of insured unemployment represents a complete or censored spell of joblessness. The consequences of this issue were discussed above in section IV.A.

A. Potential Duration and Unemployment in a Variable Duration State: Washington

1. *Estimates from Duration Models.* Table 7A displays the results of estimating various specifications of model (4.1) using the Washington State data. In each case the dependent variable is the natural log of the weeks of benefits paid to the claimant during the benefit year. Hence, the models are flawed in that they fail to account for censoring of the dependent variable and impose the assumption of log-normality on the dependent variable. Although the latter is an improvement over assuming normality, assuming a Weibull or inverse exponential distribution would be an

[10] The Employer bonus experiment had no measured impact on behavior. Hence, including it in the analysis provides a way of increasing the sample size. As it turns out, the basic results would be the same if only the Control group were examined.

[11] Because these are administrative data, there is no way of distinguishing unemployment from out-of-the labor force status for workers who have no earnings after the spell of insured unemployment ends. Accordingly, I refer to the duration of joblessness (meaning either unemployment or out-of-labor force status) and the probability of return to work.

improvement (this is done below).[12]

The point of the estimates displayed in Table 7A is to explore the sensitivity of estimates of the effects of the potential duration of benefits to various specifications of the incentives facing UI claimants. Columns 1 through 4 use two different measures of the replacement rate, along with base period earnings, to characterize the disincentives to reemployment faced by UI recipients. Columns 5 through 7 use a combination of the weekly benefit amount and base period earnings to characterize those disincentives, and columns 8 through 10 use a specification suggested by Welsh (1977) and implemented by Classen (1979). Welsh's suggestion was to include earnings in the two high-quarters of the base period (the amount from which the weekly benefit amount is calculated) and the amount by which earnings in the two high quarters exceed the amount that would give a claimant the maximum weekly benefit amount. His argument is that there is no independent information contained in the weekly benefit amount, base period earnings, or the replacement rate that is not contained in these two variables.

Two other aspects of the models estimated need to be noted. First, a "recall" variable is included, which equals one if the claimant was reemployed by the same employer after the spell of insured unemployment as before. Also, the recall variable is interacted with the potential duration of benefits, D_{pot}, so that differences in the impact of potential benefit duration that might arise between workers on temporary layoff and others can be estimated. Second, a measure of earnings variability -- the standard deviation of each worker's four quarterly earnings amounts -- is added to some specifications (or substituted for D_{pot} in some cases). Since less earnings variability is

[12] Note that a variety of control variables, which are of secondary importance for present purposes, have also been included in the model: age (4 dummies), gender, ethnicity (4 dummies), the number of referrals received by the claimant from the Employment Service, the number of employers that the claimant worked for during the base period, geographic location (20 dummies),and industry of employment before job loss (10 dummies).

precisely what leads to longer potential duration of benefits -- as shown in section II above -- some control for earnings variability is required in order to ensure that the potential duration variable does not merely reflect greater earnings stability during the base period.

Table 7A's estimates of the impact of an additional week of benefit eligibility on the duration of unemployment range from a low of .7 percent (specification 8) to a high of 2.8 percent for workers who are not recalled to their pre-layoff employer (specifications 1 and 3).[13] Since the sample mean of the dependent variable is 16.13 weeks, this range amounts to an increase in the duration of unemployment of between .11 and .45 week as a result of an additional week of benefit eligibility. This spans the range of a large number of the estimates summarized in Table 5. A rather striking implication of this result is that a fairly wide range of estimates of the disincentive effects of an additional week of benefits can be obtained simply by manipulating the way benefit levels are entered in an estimating equation.

Note that controlling for earnings variability in the base period does not reduce the estimated impact of an additional week of benefit eligibility (columns 3, 4, 7, and 10). In fact, just the opposite is true. The highest estimates of the impact of an additional week of benefit eligibility come from the specifications that include the replacement rate that is based on base period earnings (columns 1 and 3). If we discount this estimate, we can narrow the range of estimates from between .11 and .45 week to between .11 and .29 week.

The estimates in Table 7A are all based on a flawed estimating method, as already noted. The estimates displayed in Table 7B are intended to show the extent to which different estimating techniques yield different answers to questions about the impact of increased potential duration of benefits. Columns 1 and 2 display linear OLS

[13] Since the specification is semi-logarithmic, the coefficients can be interpreted as the percentage change in the dependent variable induced by a unit change in the respective independent variable.

estimates of two insured unemployment duration models -- the first controls for the weekly benefit amount and base period earnings (as did specifications 5 through 7 in Table 7A); the second is the Welch-Classen specification described above (specifications 8 through 10 in Table 7A). Columns 3 and 4 in Table 7A display the same two models, but this time estimated with the natural log of the weeks of benefits paid as the dependent variable. (These estimates are identical to those already displayed in columns 7 and 10 of Table 7A.) Finally, columns 5 and 6 show the results of estimating the two models with a correction for censoring of the dependent variable and with the assumption that the underlying distribution of unemployment spells in characterized by the Weibull distribution. This is the parametric model of time to reemployment that was discussed in section IV.

The estimates displayed in Table 7B suggest that the differences between the two specifications are minimal (compare column 1 with column 2, column 3 with column 4, and column 5 with column 6), but that the different estimating techniques give rather different answers about the impact of an additional week of benefit eligibility. The linear OLS estimates suggest that an additional week of benefit eligibility adds about .28 week to the duration of unemployment. The semi-log estimates suggest that an additional week of benefit eligibility adds about .20 week to the duration of unemployment. But the Weibull models, which are the most defensible, suggest that an additional week of benefit eligibility adds about .45 week to the duration of unemployment. If we take the Weibull models' estimates as "true," and if these results are representative of the estimates reported in the literature, then they suggest that models that fail to account for censoring and that make questionable assumptions about the distribution of unemployment spells may understate the disincentive effects of additional weeks of unemployment eligibility.

Although the models displayed in Table 7B include a control for the variability of earnings, the question remains whether the potential duration of unemployment (D_{pot}) is really just a proxy for some aspect of a worker's experience before being laid off and claiming benefits. This is more than possible, since a worker's potential benefit duration is inversely related to the variability of his or her earnings during the base period. That base period experience experience may in turn reflect some unobserved characteristics of the worker that would influence his or her unemployment duration. If so, then the estimated coefficient of D_{pot} in a duration equation may capture not the disincentive effect of longer potential duration of benefits, but rather some unobserved characteristic.

One way of gaining some insight into this issue is to take data on UI claimants in a state where the potential duration of benefits is in fact uniform, and simulate a potential duration of benefits for each claimant as if he or she were in a variable duration state. This can be done using one of the potential duration formulas discussed in section II. If such a *simulated* potential duration variable, when included in the duration equation, were to yield results similar to those obtained above for Washington State, then it would reduce confidence in estimates obtained from variable duration states.

Table 8 displays the results of such an exercise. Illinois is a uniform duration state -- all eligible claimants have a 26-week potential duration of benefits. Using the potential duration formula from Washington State, a simulated potential duration of benefits was created for each of the workers in two subsamples of Illinois UI claimants -- those eligible and those ineligible for FSC -- and included in Weibull duration models similar to those displayed in Table 7B.

The main result of the estimates shown in Table 8 is that the simulated D_{pot} variable is a weak predictor of the duration of unemployment. This is true whether or not the earnings variability measure is included in the estimating equation. If anything, the simulated D_{pot} variable is negatively related to unemployment duration, suggesting that workers who would be eligible for more weeks of benefits under the Washington State formula could be expected to have shorter spells of unemployment. (This makes sense, since workers with stable work histories would be expected to have a stronger attachment to the labor force, and to return to work relatively quickly.) The inference to be drawn is that the estimated coefficients of potential duration in the Washington State models (in Table 7B, for example) should perhaps be taken seriously as estimates of the disincentive effects of increasing the potential duration of UI benefits. That is, Table 8's results make it more difficult to dismiss estimates of the disincentive effects of increased potential benefit duration from states where potential duration varies only with work history during the base period.

2. *Estimates from Hazard Models.* Table 9 displays unadjusted estimates of the conditional probability of reemployment--or discrete reemployment hazards--for each of three groups of UI recipients in Washington State: workers who were eligible for 19 to 21 weeks of benefits; those eligible for 24 to 26 weeks; and those eligible for 30 weeks.[14] These hazards, which are based on the well-known Kaplan-Meier estimator, are descriptive in that they do not adjust for observable differences among these three groups of workers.

Because UI claimants in Washington State are certified for two weeks of benefits at a time, time until exhaustion of benefits is measured in two-week intervals.[15] The unadjusted reemployment hazards in Table 9 are computed by dividing the

[14] Workers eligible for 19, 20, and 21 weeks of benefits are aggregated in order to yield a group of workers large enough to give a hazard function in which some confidence can be placed. Similarly for workers eligible for 24, 25, and 26 weeks. The results reported below are not appreciable changed if workers are not aggregated in this way.

[15] Some information is lost by using discrete two-week time periods. Any worker who ended his or her spell of unemployment an even number of weeks before exhaustion (including zero) and gained reemployment in the following week is counted as gaining reemployment one week too late. The importance of this information loss is lessened by the fact that, as Harris (1987) found, only about half as many workers receive an even number of weeks of benefits as receive odd number of weeks, mainly because of the system of certifying for two weeks of benefits at a time. The increased simplicity that results from using two-week periods outweighs the information loss that may result.

number of workers who became reemployed during two-week period t prior to benefit exhaustion by the number of workers who were unemployed at the beginning of period t. This latter group--the so-called risk set, or the number of workers "at risk" of reemployment--is shown in the columns labeled Risk Set, and the unadjusted reemployment probability is shown in the columns labeled Hazard.

Consider the 877 workers who began their spell of unemployment with potential benefit duration of 19 to 21 weeks. Since 78 of these workers were reemployed by the end of weeks 19 and 20 before exhausting their benefits, the reemployment hazard for this period is 0.0889. Other reemployment hazards are computed similarly. Note that the risk set in period t-1 does not generally equal the risk set in period t minus the number of workers who gained reemployment by the end of period t. For example, there were 877 workers eligible for 19-21 weeks of benefits at the beginning of pre-exhaustion weeks 19 and 20, and 78 of these found reemployment before pre-exhaustion weeks 17 and 18. But the risk set in period 18 is 775, which is less than 877 minus 78. This occurs because 24 workers left the labor force (that is, stopped searching for work and collecting UI benefits) during pre-exhaustion weeks 19 and 20.[16]

The general time-pattern of the hazards shown in Table 9 is similar for the three groups: All three hazard functions have an early spike, then fall gradually to a flat segment with hazards in the neighborhood of .02 to .03., and finally show a large spike at the time of benefit exhaustion (week 0). Note that there are noticeable upturns in the hazards just before the exhaustion of benefits.

Although the hazards for the three groups shown in Table 9 are similar in a general way, closer comparison of the three hazard functions shows some differences. Mainly, the workers who are eligible for 30 weeks of benefits appear to have higher

[16] It is also possible that these workers left Washington State, in which case it is impossible be know their labor force status. In either case, treating these cases as incomplete or censored spells of joblessness is appropriate.

reemployment hazards early in their unemployment spells than do workers eligible for 19 to 21 weeks or 24 to 26 weeks. This has implications for the unemployment durations experienced by the three groups.

The bottom rows of Table 9 shows two estimates of the expected duration of unemployment that is implied by each of the three unadjusted hazard functions. Estimator 1 of the expected durations is defined as:

(5.1) $d_1 = 2\,(f_1 1 + f_2 2 + f_3 3 + \ldots + f_t t + \ldots)$

where

(5.2) $f_t = (1\text{-}h_1)(1\text{-}h_2) \ldots (1\text{-}h_{t-1})(h_t)$.

Equation (5.2) gives the unconditional probability of experiencing t two-week periods of unemployment (calculated as the product of the probabilities of not finding a job in each of the first t-1 periods, times h_t, the conditional probability of finding a job in period t).

Estimator 2 of the expected durations is defined as:

(5.3) $d_2 = \{(U_1)(h_1)(1) + (U_2)(h_2)(2) + (U_3)(h_3)(3) + \ldots + (U_t)(h_t)(t) + \ldots\} / U_1$

where U_t denotes the number of workers in the risk set at the beginning of period t before benefit exhaustion.

These alternative estimators make different assumptions about whether the labor market is in equilibrium. Estimator 1 assumes that the market is in equilibrium and tends to yield higher estimates of the expected duration of unemployment. Estimator 2 does not make such an assumption and is, in effect, a "mechanical" way of estimating the expected duration. In both estimators, it is assumed that workers who have exhausted benefits have a constant reemployment hazard equal to h_e (the hazard in the period in which benefits were exhausted) in perpetuity; for example, h_e = 0.3485 for workers eligible for 19 to 21 weeks of benefits in Table 9.

Table 9 shows that the expected duration for workers eligible for 19 to 21 weeks of benefits was between 14.5 weeks (estimator 1) and 16.6 weeks (estimator 2). For workers eligible for 24 to 26 weeks of benefits, expected duration was between 14.1 weeks and 18.0 weeks; and for workers eligible for 30 weeks of benefits, expected duration was between 14.0 weeks and 18.4 weeks.

These estimated expected durations of unemployment can be used to obtain a direct estimate of how additional weeks of potential benefit duration can be expected to affect the duration of a worker's unemployment. Estimator 1 suggests that there may be some impact of extending benefits -- the 5-week increase in potential duration between 19-21 weeks and 24-26 weeks is associated with an increase of about 1.5 in the duration of unemployment, or .29 week per additional week of benefit eligibility. However, the 5-week increase in potential duration between 24-26 weeks and 30 weeks is associated with only a very small increase in unemployment duration -- about .33 week, or .07 week per additional week of benefit eligibility. By the same sort of reasoning, estimator 2 suggests no impact -- or possibly even a negative impact -- of extending benefits.

An important caveat regarding the estimates shown in Table 9 is that they are not adjusted for observable characteristics of the workers that may in turn be associated with unemployment duration. This is important, since the apparent negative impact of additional weeks of benefits implied by estimator 2 could simply result from the fact that workers who are eligible for more weeks of benefits are more likely to return to work quickly, rather than the separate effect of additional weeks of benefits. It is possible to adjust for worker characteristics, but this is left for future work.

B. The Impact of Federal Supplemental Compensation (FSC) in Illinois

As already noted, the Illinois data that were used above span the expiration of one of the the emergency UI benefit extensions -- the Federal Supplemental Compensation program (FSC). Workers eligible for Phase IV of FSC in Illinois had a total potential duration of benefits of 38 weeks -- 26 weeks of regular state benefit, plus 12 weeks of FSC. In this section, the natural experiment presented by the expiration of FSC is used to obtain estimates of the additional 12 weeks of benefit eligibility. It is important to repeat that FSC did not "trigger off" as the standing Extended Benefits program would have done. Rather, FSC was ended by Congressional fiat in response to the impression that labor market conditions nationwide -- not just in Illinois -- no longer required a Federal emergency benefit program.

Table 10 displays estimates of four Weibull duration models of unemployment; that is, parametric models of time to reemployment that estimate equation (4.1) under the assumption that the disturbance term u has the Weibull distribution. The interpretation of the Weibull model's coefficients is straightforward: Each coefficient gives the approximate proportional change in unemployment duration that is attributable to a unit change in the explanatory variable. As already discussed, these are the most defensible of the duration models that are available, although the hazard models presented below arguably yield more convincing estimates.

The four models shown in Table 10 use different specifications of the incentives facing UI recipients. Specifications 1 and 2 include the replacement rate (based either on base period earnings or high-quarter earnings), specification 3 includes the weekly benefit amount and the base period earnings, and specification 3 uses the Welch-Classen variables described above. The main finding of the results in Table 10 is that the impact of FSC-eligibility on unemployment duration is insensitive to these variations in specification. The impact of the additional 12 weeks of potential benefit duration provided by FSC is estimated to be a 19 percent increase in unemployment duration.[17] This translates into an increase of 4 weeks in unemployment duration, or .34 week per additional week of benefit eligibility. This estimate is higher than most of the estimates obtained using the Washington State data -- that is, the estimates that were based on variability across individuals that resulted from differing base period work histories.

Table 11 displays estimates of the conditional reemployment probabilities for two groups of workers in Illinois: those who were eligible for FSC and those who were ineligible for FSC. These hazards are constructed in an identical manner to those presented in Table 9, which examined UI recipients in Washington State. That is, they show the conditional probability of reemployment at various times before exhaustion of benefits. The patterns of the two hazard functions shown in Table 11 are similar to each other (and also to the Washington State hazards). Both have an early spike, then fall gradually to a flat segment, turn up just before exhaustion of benefits, and then show a large spike at the time of benefit exhaustion (week 0).

There are also some differences between the hazards for FSC-eligibles and FSC-ineligibles. Mainly, the FSC-eligibles appear to have higher reemployment hazards early in their unemployment spells (weeks 38 through 30 before exhaustion), but lower reemployment hazards later (weeks 26 through 4 prior to exhaustion).

The implications of the hazard functions for unemployment duration are displayed at the bottom of Table 11. As in Table 9, two estimates of the expected duration of unemployment are shown. Estimator 1 suggests that the FSC-eligible workers experienced about 2.8 more weeks of unemployment than did the FSC-ineligibles, or about .23 week per additional week of benefit eligibility. Estimator 2, on the other hand, suggests that the FSC-eligible workers experienced 1.3 weeks *less*

[17] The estimate is obtained by dividing the coefficient of FSC-eligibility by the Weibull shape parameter.

unemployment than did the FSC-ineligibles, or about .11 week less for each additional week of benefit eligibility. This rather unlikely result could suggest a need to adjust for observable characteristics of the workers, which could differ between the FSC-eligible and FSC-ineligible workers.

VI. Summary and Conclusions

The main goals of this paper have been to review the sources of variation in the potential duration of unemployment benefits (sections I and II), to review critically existing estimates of the extent to which increasing potential benefit duration affects the duration of unemployment (section IV), and to explore the effect of variation in potential duration on the expected duration of unemployment of UI recipients (section V). In particular, data from two states were examined in an effort to understand the extent to which estimates of the effects of extended benefits are sensitive to differences in model specification, estimating technique, and the source of variation in potential benefit duration.

Two conclusions can be drawn from the discussion of sources of variation in potential benefit duration and the review of existing estimates in sections I, II, and IV. First, differences in the effects of emergency benefit extensions, the Extended Benefits program, and within-state variation in states have variable benefit duration have never been systematically analyzed. This is an important omission because the impact of an emergency extension may be quite different from adding to the potential duration of benefits in a variable duration state. The existing estimates, summarized in Table 5, have not sorted out the extent to which differences in estimated impacts result from differences in the underlying source of variation in potential benefit duration.

Second, few of the existing studies have examined the extent to which their estimates are sensitive to estimating technique and model specification. This has made it difficult to know whether the behavioral impacts being estimated are real or simply an accident of the data. Section IV reviewed some of the problems inherent in estimating the disincentive effects of increasing the potential duration of benefits, and it seems fair to say that they are unusually daunting and make the estimates less convincing than most econometric estimates that are based on cross-sectional or panel data. Accordingly, the importance of sensitivity analysis seem especially important in regard to this question, but sensitivity analyses have rarely been pursued.

The exercises presented in section V can be viewed as an attempt to provide such a sensitivity analysis on two data sets. The main conclusions of this exercise are as follows. First, estimates of the impact of an additional week of potential benefit duration that are derived from a variable duration state such as Washington State are quite fragile, although they do suggest that there may be some increase in the duration of unemployment that results from increased potential duration. The largest estimates of the impact of increased potential duration derive from parametric models of the time to reemployment that account for censoring and assume the underlying distribution of unemployment spell conforms to the Weibull distribution. These models yield suggest that an added week of potential benefit duration adds .45 week to the duration of unemployment of workers who are not recalled to their pre-layoff employer. This estimate is not especially sensitive to model specification. Also, using a simulated potential duration variable in a uniform duration state (Illinois) yields results that increase confidence that there is a real impact of additional weeks of potential duration.

Interestingly, though, estimates of the impact of an additional week of potential benefit duration that are based on a reemployment hazard function suggest a smaller impact of increasing the potential duration of benefits by one week. Those estimates suggest that adding a week to the potential duration of benefits adds at most .29 week to the expected duration of unemployment, and may have no impact at all. Clearly, then, the estimates, although relatively insensitive to model specification, are quite sensitive to modeling technique.

Second, estimates of the impact of an additional week of potential benefit duration that come from the expiration of the Federal Supplemental Compensation program are also sensitive to estimating technique. A Weibull model of the duration of joblessness suggests that the availability of FSC increased the expected duration of a worker's jobless spell by roughly 4 weeks, or by .34 week per additional week of benefit eligibility. This estimate is actually somewhat lower than the estimate derived from Washington State using a Weibull model (that estimate was .45 week per additional week of benefit eligibility).

But a hazard model of the conditional probability of becoming reemployed again yields lower estimates of the influence of extended benefits on jobless duration. Specifically, the hazard estimates suggest that FSC eligibility increased the duration of unemployment by at most .23 week per additional week of eligibility, and may have had no impact at all.

Given the fragility of the estimates, it seems sensible to be rather modest in making claims about our knowledge of the disincentive effects of extended unemployment benefits. Although the evidence does suggests that additional weeks of benefits may increase the duration of unemployment, it is clear that the impact is difficult to estimate and especially sensitive to estimating technique. In particular, the findings presented suggest that two avenues of further work could be especially fruitful. First, examining the robustness of estimates from the hazard models would seem to be especially important, given that the hazard model is most appealing a priori and has the potential to provide results that are most convincing. Second, though, it seems important to explore further the variation in disincentives that arise from different sources of variation in the potential duration of benefits -- emergency benefit programs, the Extended Benefit program, or within-state variation in potential benefit duration. This will require additional data, particularly data that cover an extended period of time and a variety of potential duration "regimes." Only with additional work that includes careful sensitivity testing can we expect to improve understanding of the disincentives of extended benefits.

References

Advisory Council on Unemployment Compensation. *Report and Recommendations*. Washington, D.C.: Advisory Council on Unemployment Compensation, February 1994.

Advisory Council on Unemployment Compensation. *Unemployment Insurance in the United States: Benefits, Financing, Coverage*. Washington, D.C.: Advisory Council on Unemployment Compensation, 1995.

Akerlof, George A. and Brian G. M. Main. "Unemployment Spells and Unemployment Experience." *American Economic Review* 70 (December 1980): 885–893.

Allison, Paul D. "Discrete–Time Methods for the Analysis of Event Histories." In *Sociological Methodology 1982*, edited by Samuel Lienhardt. San Francisco: Jossey–Bass, 1982. Pp. 61–98.

Allison, Paul D. *Event History Analysis*. Beverly Hills: Sage Publications, 1984.

Blaustein, Saul J. with Wilbur J. Cohen and William Haber. *Unemployment Insurance in the United States*. Kalamazoo, Michigan: W.E. Upjohn Institute for Employment Research, 1983.

Burgess, Paul L. and Jerry L. Kingston. "Labor Market Experiences of Unemployment Insurance Exhaustees. Unemployment Insurance Occasional Paper 79–3. Washington, D.C.: U.S. Department of Labor, Education and Training Administration, Unemployment Insurance Service, 1979.

Clark, Kim B. and Lawrence H. Summers. "Labor Market Dynamics and Unemployment: A Reconsideration." *Brookings Papers on Economic Activity* 1979 (1): 13–60.

Classen, Kathleen P. "Unemployment Insurance and Job Search." In *Studies in the Economics of Search*, edited by S.A. Lippman and J.J. McCall. Amsterdam: North–Holland, 1979. Pp. 191–219.

Corson, Walter and Mark Dynarski. "A Study of Unemployment Insurance Recipients and Exhaustees: Findings from a National Survey." Unemployment Insurance Occasional Paper 90–3. Washington, D.C.: U.S. Department of Labor, Education and Training Administration, Unemployment Insurance Service, 1990.

Corson, Walter, Jean Grossman, and Walter Nicholson. "An Evaluation of the Federal Supplemental Compensation Program." Unemployment Insurance Occasional Paper 86–3. Washington, D.C.: U.S. Department of Labor, Education and Training Administration, Unemployment Insurance Service, 1986.

Corson, Walter, David Long, and Walter Nicholson. "An Analysis of the 1981–82 Changes in the Extended Benefit Program." Unemployment Insurance Occasional Paper 85–1. Washington, D.C.: U.S. Department of Labor, Education and Training Administration, Unemployment Insurance Service, 1985.

Corson, Walter, David Long, and Walter Nicholson. "Evaluation of the Charleston Claimant Placement and Work Test Demonstration." Unemployment Insurance Occasional Paper 85–2. Washington, D.C.: U.S. Department of Labor, Education and Training Administration, Unemployment Insurance Service, 1985.

Corson, Walter and Walter Nicholson. *The Federal Supplemental Benefits Program.* Kalamazoo, Michigan: W.E. Upjohn Institute for Employment Research, 1982.

Cox, D. R. "Regression Models and Life Tables." *Journal of the Royal Statistical Society* (Series B) 34 (Number 2, 1972): 187–220.

Davidson, Carl and Stephen A. Woodbury. "Optimal Unemployment Insurance." Paper prepared for the Advisory Council on Unemployment Compensation, April 1995.

Davidson, Carl and Stephen A. Woodbury. "Unemployment Insurance and Unemployment: Implications of the Reemployment Bonus Experiments." Paper prepared for the Advisory Council on Unemployment Compensation, February 1995.

Davidson, Carl and Stephen A. Woodbury. "The Displacement Effect of Reemployment Bonus Programs." *Journal of Labor Economics* 11 (October 1993): 575-605.

Davidson, Carl and Stephen A. Woodbury. "Effects of a Reemployment Bonus under Differing Benefit Entitlements, or, Why the Illinois Experiment Worked." Manuscript, Michigan State University and W.E. Upjohn Institute for Employment Research, March 1991.

Dynarski, Mark and Steven M. Sheffrin. "New Evidence on the Cyclical Behavior of Unemployment Durations." In *Unemployment and the Structure of Labor Markets,* edited by Kevin Lang and Jonathan S. Leonard. New York: Basil Blackwell, 1987. Pp. 164–185.

Ehrenberg, Ronald G. and Ronald L. Oaxaca. "Unemployment Insurance, Duration of Unemployment, and Subsequent Wage Gain." *American Economic Review* 66 (December 1976): 754–766.

Felder, Henry E. and Richard W. West. "The Federal Supplemental Benefits Program: National Experience and the Impact of P.L. 95–19." Washington, D.C.: U.S. Department of Labor, Education and Training Administration, Unemployment Insurance Service, 1978.

Grossman, Jean Baldwin. "The Work Disincentive Effect of Extended Unemployment Compensation: Recent Evidence." *Review of Economics and Statistics* 71 (February 1989): 159–164.

Haber, William and Merrill G. Murray. *Unemployment Insurance in the American Economy.* Homewood, Illinois: Richard D. Irwin, 1966.

Hall, Robert E. "Turnover in the Labor Force." *Brookings Papers on Economic Activity* 1972 (3): 709–756.

Ham, John C. and Samuel A. Rea, Jr. "Unemployment Insurance and Male Unemployment Duration in Canada." *Journal of Labor Economics* 5 (July 1987): 325–353.

Hamermesh, Daniel S. *Jobless Pay and the Economy.* Baltimore, Maryland: Johns Hopkins University Press, 1977.

Han, Aaron and Jerry A. Hausman. "Flexible Parametric Estimation of Duration and Competing Risk Models. *Journal of Applied Econometrics* 5 (January-March 1990): 1-28.

Hansen, W. Lee and James F. Byers. "Unemployment Compensation and Training: Can a Closer Link Be Forged?" In *Unemployment Compensation: The Second Half–Century,* edited by W. Lee Hansen and James F. Byers. Madison, Wisconsin: University of Wisconsin Press, 1990. Pp. 267–301.

Hight, Joseph. "Insured Unemployment Rates, Extended Benefits, and Unemployment Insurance Exhaustions." In *Proceedings of the Twenty–Eighth Annual Winter Meeting of the Industrial Relations Research Association,* edited by James L. Stern and Barbara D. Dennis. Madison, WI: IRRA, 1976. Pp. 242–249.

Holen, Arlene. "Effects of Unemployment Insurance Entitlement on Duration and Job Search Outcome." *Industrial and Labor Relations Review* 30 (July 1977): 445–450.

Jacobson, Louis S. "The Ability of Trade Adjustment Assistance's Wage-Subsidy and Generous Benefits to Offset Dislocated Workers' Earnings Losses." Manuscript, Westat, Inc., September 1994.

Jacobson, Louis, Robert LaLonde, and Daniel Sullivan. *The Costs of Worker Dislocation.* Kalamazoo, MI: W.E. Upjohn Institute, 1993a.

Jacobson, Louis, Robert LaLonde, and Daniel Sullivan. "Earnings Losses of Displace Workers." *American Economic Review* 83 (September 1993b): 685–709.

Jacobson, Louis, Robert LaLonde, and Daniel Sullivan. "Long–Term Earnings Losses of High–Seniority Displaced Workers." *Federal Reserve Bank of Chicago Economic Perspectives* 17 (November/December 1993c): 2–20.

Johnson, Terry R. and Daniel R. Klepinger. "Evaluation of the Impacts of the Washington Alternative Work Search Experiment." Unemployment Insurance Occasional Paper 91–4. Washington, D.C.: U.S. Department of Labor, Education and Training Administration, Unemployment Insurance Service, 1991.

Katz, Arnold and Jack Ochs. "Implications of Potential Duration Policies." In *Unemployment Compensation: Studies and Research*, Volume I. Washington, D.C.: National Commission on Unemployment Compenastion, July 1980. Pp. 111–126.

Katz, Lawrence F. and Bruce D. Meyer. "The Impact of the Potential Duration of Unemployment Benefits on the Durationof Unemployment." *Journal of Public Economics* 41 (February 1990): 45–72.

Kiefer, Nicholas M. "Economic Duration Data and Hazard Functions." *Journal of Economic Literature* 26 (June 1988): 646–679.

Lancaster, Tony. "Econometric Methods for the Duration of Unemployment." *Econometrica* 47 (July 1979): 939–956.

Leigh, Duane E. *Assisting Displaced Workers: Do the States Have a Better Idea?* Kalamazoo, Michigan: W.E. Upjohn Institute for Employment Research, 1989.

Leigh, Duane E. *Does Training Work for Displaced Workers?* Kalamazoo, Michigan: W.E. Upjohn Institute for Employment Research, 1990.

Mathematica Policy Research. "A Longitudinal Study of Unemployment Insurance Exhaustees: Waves I and II." Report prepared for the Employment and Training Administration, U.S. Department of Labor, January 1976.

Moffitt, Robert. "Unemployment Insurance and the Distribution of Unemployment Spells." *Journal of Econometrics* 28 (April 1985a): 85–101.

Moffitt, Robert. "The Effect of the Duration of Unemployment Benefits on Work Incentives: An Analysis of Four Data Sets." Unemployment Insurance Service Occasional Paper 85–4. Washington, D.C.: U.S. Department of Labor, Education and Training Administration, Unemployment Insurance Service, 1985b.

Moffitt, Robert and Walter Nicholson. "The Effect of Unemployment Insurance on Unemployment: The Case of Federal Supplemental Benefits." *Review of Economics and Statistics* 64 (February 1982): 1–11.

Mortensen, Dale T. "Job Search and Labor Market Analysis." In *Handbook of Labor*

Economics, Volume II, edited by Orley Ashenfelter and Richard Layard. Amsterdam: North–Holland, 1986. Pp. 849–919.

Murray, Merrill G. *The Duration of Unemployment Benefits*. Kalamazoo, Michigan: W.E. Upjohn Institute for Employment Research, 1974.

Newton, Floyd C. and Harvey S. Rosen. "Unemployment Insurance, Income Taxation, and Duration of Unemployment: Evidence from Georgia." *Southern Economic Journal* 45 (January 1979): 773–784.

Nicholson, Walter and Walter Corson. "A Longitudinal Study of Unemployment Insurance Exhaustees: Waves I and II." Report prepared for the Employment and Training Administration, U.S. Department of Labor, January 1976.

Perry, George. "Unemployment Flows in the U.S. Labor Market." *Brookings Papers on Economic Activity* 1972 (2): 245–278.

Rubin, Murray. *Federal–State Relations in Unemployment Insurance: A Balance of Power*. Kalamazoo, Michigan: W.E. Upjohn Institute for Employment Research, 1983.

Rubin, Murray. "State–Federal Relations in Unemployment Insurance." In *Unemployment Compensation: The Second Half–Century*, edited by W. Lee Hansen and James F. Byers. Madison, Wisconsin: University of Wisconsin Press, 1990. Pp. 207–243.

Solon, Gary. "Work Incentive Effects of Taxing Unemployment Benefits." *Econometrica* 53 (March 1985): 295–306.

Steinberg, Danny, and Frank A. Monforte. "Estimating the Effects of Job Search Assistance of the Unemployment Durations of Displaced Workers." In *Unemployment and the Structure of Labor Markets*, edited by Kevin Lang and Jonathan S. Leonard. New York: Basil Blackwell, 1987. Pp. 186–206.

Storey, James R. and Gene Falk. "Unemployment Compensation: The Emergency Unemployment Compensation (EUC) Program." Washington, D.C.: The Library of Congress, Congressional Research Service, 93–367 EPW, March 1993.

Sueyoshi, Glenn T. "Semiparametric Proportional Hazards Estimation of Competing Risks Models with Time-Varying Covariates." *Journal of Econometrics* 51 (January/February 1992): 25-58.

Woodbury, Stephen A. and Robert G. Spiegelman. "Bonuses to Workers and Employers to Reduce Unemployment: Randomized Trials in Illinois." *American Economic Review* 77 (September 1987): 513–530.

Woodbury, Stephen A. and Murray Rubin. "The Duration of Benefits." Draft prepared for a conference on "Unemployment Insurance in the United States: Analysis of Policy Issues," W.E. Upjohn Institute, Kalamazoo, Michigan, June 1995.

Table 1
Federal Extended Unemployment Benefit Programs,
1958 to 1995

Program and Enabling Legislation	Effective Dates and Extensions	Potential Duration of Extended Benefits Provided	Financing	Notes
Temporary Unemployment Compensation Act, P.L. 85-441	6/58 - 7/59	50% of regular state duration, up to 13 weeks.	Interest-free loans to 17 participating states	State participation voluntary.
Temporary Extended Unemployment Compensation Act (TEUC), P.L. 87-6	4/61 - 6/62	50% of regular state duration, up to 13 weeks.	Temporary increases in Federal Unemployment Tax (.4% in 1962, .25% in 1963)	
Extended Unemployment Compensation Act of 1970 (EB), P.L. 91-373, with major amendments in P.L. 96-364, P.L. 96-499, P.L. 97-35, P.L. 102-318	8/70 to present	50% of regular state duration, up to 13 weeks	One-half from Federal Unemployment Tax revenues paid to Extended Unemployment Compensation Account (EUCA); one-half from state UI reserves.	EB activated in a state by an insured unemployment rate (IUR) trigger, 8/70 to present; EB could be activated in all states by a national IUR trigger, 8/70-8/81. Starting 1980, EB denied to claimants refusing to seek or accept suitable work, and to claimants who had quit or been discharged. State triggers were made more restrictive, 8/81. Eligibility for EB made more restrictive, 8/81. States permitted to adopt a total unemployment rate (TUR) trigger, 3/93.
Emergency Unemployment Compensation Act, P.L. 92-224 and P.L. 92-329	1/72 - 9/72, extended to 3/73	50% of regular state durations, up to 13 weeks.	Extended Unemployment Compensation Account (EUCA)	State-level triggers (different from EB triggers) used to activate program.

Table 1
(Continued)

Program and Enabling Legislation	Effective Dates and Extensions	Potential Duration of Extended Benefits Provided	Financing	Notes
Federal Supplemental Benefits (FSB), P.L. 93-572, P.L. 94-12, P.L. 94-45, P.L. 95-19	1/75 - 12/76, extended to 1/78	50% of regular state duration, up to 13 weeks (1/75-2/75 and 5/77-1/78); additional 50% of regular state duration, up to 13 weeks provided 3/75-4/77 (that is, up to 26 weeks of FSC total).	Repayable advances to EUCA from general revenues; general revenues after 3/77	EB program was activated in all states, so total potential benefit duration was 65 for those exhausting EB between 3/75 and 4/77. State-level triggers applied starting 1/76. Uniform Federal eligibility and disqualification standards implemented 4/77 (P.L. 95-19).
Federal Supplemental Compensation (FSC), P.L. 97-248, P.L. 97-424, P.L. 98-21, P.L. 98-135	9/82 - 3/83, extended to 9/93 and 3/85	FSC-I (9/82-1/83): 50% of regular state duration, up to 6 to 10 weeks. FSC-II (1/83-3/83): 65% of regular state duration, up to 8 to 16 weeks. FSC-III (4/83-9/83): 55% of regular state duration, up to 8 to 14 weeks. FSC-IV (10/83-3/85): Same as FSC III, except entitlement did not vary once established.	General revenues	Potential duration varied with state's EB status and separate FSC triggers. Except in FSC-IV, potential duration would vary when EB and FSC status changed. FSC-I and FSC-II exhaustees could collect FSC-III benefits, but not FSC-IV benefits. EB eligibility criteria applied to all phases of FSC. Available regular state benefits and EB (if activated) had to be exhausted to receive FSC.

Table 1
(Continued)

Program and Enabling Legislation	Effective Dates and Extensions	Potential Duration of Extended Benefits Provided	Financing	Notes
Emergency Unemployment Compensation Act of 1991 (EUC), P.L. 102-164, P.L. 102-182, P.L. 102-244, P.L. 102-318, P.L. 103-6, P.L. 103-152	11/91 - 6/92, extended to 7/92, 3/93, 10/93, and 2/94	EUC-I (11/91-2/92): Lesser of 100% of regular benefits, or 13 or 20 weeks. EUC-II (2/92-7/92): Lesser of 130% of regular benefits, or 26 or 33 weeks. EUC-III (7/92-3/93): Lesser of 100% of regular benefits, or 20 or 26 weeks. EUC-IV (3/93-10/93): Lesser of 60% of regular benefits, or 10 or 15 weeks EUC-V (10/93-2/94): Lesser of 50% of regular benefits, or 7 or 13 weeks	Extended Unemployment Compensation Account (EUCA) until 7/92, general revenues thereafter	Potential duration determined at time of filing for EUC, and depended on state's classification as high- or low-unemployment. EUC entitlement could be increased if state moved from low to high status, or if program became more generous; EUC entitlement could not be decreased. Claimants exhausting benefits between 3/91 and 11/91 could receive benefits under "reach-back" provisions (but no retroactive benefits paid). EB eligibility criteria applied to all phases of ECU. Once EUC was exhausted, a claimant needed to regain regular UI eligibility to receive additional EUC.

Table 2
Potential Duration of Unemployment Insurance Benefits: Summary of State Practices

State	Potential Duration (weeks)		Minimum requirement for 26 weeks					
	Minimum	Maximum	Base Period Earnings ($)	High-Quarter Earnings ($)	*a*	*b*	*g*	State Minimum Weekly Benefit Amount
Alabama	15	26	1716	516	0.33	4.17%	7.91	22
Alaska	16	26	1000	250-286	1.31	17.60%	7.44	44
Arizona	12	26	3120	1000	0.33	4.00%	8.25	40
Arkansas	9	26	3588	897-1183	0.33	3.85%	8.57	46
California	14	26	2080	900-920	0.50	4.35%	11.49	40
Colorado	13	26	1950	488-649	0.33	3.85%	8.57	25
Connecticut	26	26	600	150	0.65	3.85%	16.88	15
Delaware	23	26	2184	966	0.50	4.35%	11.49	21
District of Columbia	20	26	2600	1300	0.50	3.85%	12.99	50
Florida	10	26	1040	260	0.25	3.85%	6.49	10
Georgia	9	26	3848	962	0.25	4.00%	6.25	37
Hawaii	26	26	130	32.5-105	1.00	4.76%	21.00	5
Idaho	10	26	3690	1144	0.31	3.85%	8.05	44
Illinois	26	26	1600	400-1160	0.83	3.77%	22.02	51
Indiana	14	26	18757	4800	0.28	5.00%	5.60	50
Iowa	11	26	2496	740	0.33	4.34%	7.60	32
Kansas	10	26	4914	1229-1482	0.33	4.25%	7.76	63
Kentucky	15	26	1857	750	0.33	4.74%	6.96	22

State	Potential Duration (weeks)		Minimum requirement for 26 weeks					State Minimum Weekly Benefit Amount
	Minimum	Maximum	Base Period Earnings ($)	High-Quarter Earnings ($)	*a*	*b*	*g*	
Louisiana	8	26	3081	800	0.27	4.00%	6.75	10
Maine	21	26	2730	683	0.33	4.55%	7.25	35
Maryland	26	26	900	576	0.72	4.17%	17.27	25
Massachusetts	10	30	2000	500	0.36	3.85%	9.35	14
Michigan	15	26	2100	525-781	0.52	5.38%	9.67	42
Minnesota	10	26	2999	1000	0.33	3.85%	8.57	38
Mississippi	13	26	2340	780	0.33	3.85%	8.57	30
Missouri	11	26	3510	1000	0.33	4.50%	7.33	45
Montana	8	26	4469	1117-1375	0.32	4.00%	8.00	55
Nebraska	20	26	1575	394-400	0.33	5.00%	6.60	20
Nevada	12	26	1248	400	0.33	4.00%	8.25	16
New Hampshire	26	26	2800	1200	0.30	4.40%	6.82	32
New Jersey	15	26	4375	1094-1623	0.45	4.62%	9.74	75
New Mexico	19	26	1777	1068	0.60	3.85%	15.58	41
New York	26	26	1600	400	0.65	3.85%	16.88	40
North Carolina	13	26	2603	651-868	0.33	3.85%	8.57	25
North Dakota	12	26	3572	1118	0.32	3.85%	8.31	43
Ohio	20	26	6864	1716	0.25	3.85%	6.49	66
Oklahoma	20	26			0.40	4.00%	10.00	16
Oregon	4	26	5304	1326-1360	0.33	5.00%	6.60	68

State	Potential Duration (weeks)		Minimum requirement for 26 weeks					
	Minimum	Maximum	Base Period Earnings ($)	High-Quarter Earnings ($)	*a*	*b*	*g*	State Minimum Weekly Benefit Amount
Pennsylvania	16	26	1357	900	0.69	4.00%	17.25	35
Puerto Rico	26	26	280	75	0.58	9.30%	6.24	7
Rhode Island	15	26	2961	890	0.36	4.62%	7.79	41
South Carolina	15	26	1560	540	0.33	3.85%	8.57	20
South Dakota	15	26	2183	728	0.33	3.85%	8.57	28
Tennessee	12	26	3120	780	0.25	3.85%	6.49	30
Texas	9	26	4044	1011-1050	0.27	4.00%	6.75	42
Utah	10	26	1800	450-486	0.27	3.85%	7.01	17
Vermont	26	26	1628	1163	0.42	4.44%	9.46	25
Virgin Islands	13	26	2574	858	0.33	3.85%	8.57	33
Virginia	12	26	6760	1625	0.25	4.00%	6.25	65
Washington	16	30	5694	1825	0.33	4.00%	8.25	73
West Virginia	26	26	2200	550-600	0.28	4.00%	7.00	24
Wisconsin	12	26	3250	1250	0.40	4.00%	10.00	50
Wyoming	12	26	3467	1000	0.30	4.00%	7.50	16

Source: *Comparison of State Unemployment Insurance Laws*, U.S. Department of Labor, Employment and Training Administration; and authors' calculations.

Notes: Parameter a is the maximum proportion of base period earnings that can be paid in UI benefit during a given benefit year (see equation 3 in the text).
Parameter b is the proportion of high-quarter earnings paid as the weekly benefit amount (see equation 4 in the text).
Parameter $g = a/b$ and is an index of the state's potential duration generosity.

Table 3

Classification of Workers by Weeks of
UI Benefits Claimed and
Subsequent Labor Force Status

Case	Number of Weeks of UI Benefits Claimed	Labor Force Status after Benefit Termination	Number Observed in Illinois Data (proportion)	
A	Maximum Potential	In Covered Employment	283	(0.13)
B	Maximum Potential	Out of Covered Employment	434	(0.20)
C	Fewer than Potential	In Covered Employment	1040	(0.48)
D	Fewer than Potential	Out of Covered Employment	405	(0.19)

Notes: Cases B and C are correctly characterized by usual censoring conventions; Cases A and D are misspecified by the usual conventions.

Table 4

Mean Insured Unemployment Durations for Men
by Monetary and Temporal Eligibility for
Federal Supplemental Compensation (FSC)

(Standard Errors in Parentheses)

Temporal Eligibility for FSC	Monetary Eligibility for FSC	
	Eligible	Ineligible
Eligible	21.387 (0.402) (N=1131)	16.532 (1.144) (N=79)
Ineligible	17.864 (0.318) (N=866)	20.058 (0.957) (N=86)
Difference	3.524 (0.513)	-3.527 (1.979)

Notes: In order to be monetarily eligible for FSC, a claimant needed to have total base period earnings equal to at least 1.5 times high-earnings quarter of the base period. To be Temporally Eligible for FSC, a claimant needed to file an initial claim for UI benefits before September 30, 1984. Insured unemployment duration refers to the total number of weeks of benefits (both state regular and FSC) received in the claimant's full benefit year.

Table 5

Selected Estimates of the Impact of Increased Potential Duration of UI Benefits

Study	Data	Change in weeks of unemployment from 1 added week of potential UI	Remarks
Holen (1977)	UI claimants in San Francisco, Boston, Phoenix, Seattle, Minneapolis, 1969-70	.77-.81	OLS linear duration estimates
Classen (1979)	UI claimants in Arizona and Pennsylvania, 1967-69	0-.12	Tobit duration estimates
Newton and Rosen (1979)	UI recipients in Georgia, 1974-76	.6	Tobit duration estimates
Katz and Ochs (1980)	Current Population Survey, individuals in 26 states, 1968-70 and 1973-77	.17-.23	Maximum likelihood duration estimates
Moffitt and Nicholson (1982)	Recipients of EB and FSC, 15 states, 1975-77	.1	Labor supply model, maximum likelihood estimates
Moffitt (1985a)	Continuous Wage and Benefit History, 1978-83	.15	UI exit rate estimates
Moffitt (1985b)	Continuous Wage and Benefit History, 1978-83:		UI exit rate estimates
	White men	.17	
	White women	.10	
	FSC and EB recipients in 15 states, 1975-78:		Maximum likelihood duration estimates
	Men	.45	
	Women	.28	
	UI recipients in Georgia, 1974-76:		Maximum likelihood duration estimates
	Men	.17	
	Women	.37	
Solon (1985)	UI claimants in Georgia, 1978-79	.36	Maximum likelihood duration estimates
Ham and Rea (1987)	Canadian men, 1975-80	.26-.35	UI exit rate estimates
Grossman (1989)	Continuous Wage and Benefit History, individuals in 3 states, 1981-84	.9	UI exit rate estimates of FSC impacts
Katz and Meyer (1990)	Continuous Wage and Benefit History, men in 12 states, 1978-83	.16-.20	UI exit rate estimates
Davidson and Woodbury (1995)	UI recipients in:		Translation of reemployment bonus impacts using equilibrium search model
	Illinois 1984-85	.2	
	Pennsylvania 1988-89	0-.2	
	Washington 1988-89	0-.2	

Table 6

Transformation of Data on Claimants into
Data on Claimant-Weeks

Panel A: Claimant Records

Claimant	Weeks of Insured Unemployment	Reemployed	Weekly Benefit	Eligible for FSC
1	4	1	$149	0
2	38	0	$161	1
3	0	1	$128	0

Panel B: Claimant-Week Records

Claimant	Weeks of Since Initial Claim	Reemployed	Weekly Benefit	Eligible for FSC
1	0	0	$149	0
1	1	0	$149	0
1	2	0	$149	0
1	3	0	$149	0
1	4	0	$149	0
1	5	1	$149	0
2	0	0	$161	1
2	1	0	$161	1
2	2	0	$161	1
2	3	0	$161	1
.	.	.	.	.
.	.	.	.	.
.	.	.	.	.
2	37	0	$161	1
2	38	0	$161	1
3	0	0	$128	0
3	1	1	$128	0

Notes: The claimant is the unit of observation in the alternative models of employment duration. The claimant-week is the unit of observation in the reemployment hazard models.

Table 7A

Potential Duration of Benefits, Benefits, and Unemployment Duration: Alternative Specifications for Washington State
[Dependent Variable: ln (weeks of benefits paid)]

Explanatory Variable	Mean (Std. Dev.)	1	2	3	4	5	6	7	8	9	10
D_{pot}	26.86 (4.17)	.028 (.004)	.013 (.003)	.028 (.004)	.018 (.003)	.009 (.003)	–	.013 (.003)	.007 (.003)	–	.012 (.003)
D_{pot} * Recall	2.93 (8.49)	-.060 (.007)	-.060 (.007)	-.059 (.007)	-.060 (.007)	-.058 (.007)	–	-.057 (.007)	-.058 (.007)	–	-.057 (.007)
Recall	.110 (.312)	.679 (.195)	.677 (.195)	.635 (.198)	.639 (.198)	.620 (.194)	-.975 (.039)	.569 (.198)	.625 (.194)	-.977 (.039)	.573 (.198)
Earnings Variability ($1,000s)[1]	1,401 (1,901)	–	–	.017 (.006)	.028 (.006)	–	.016 (.006)	.021 (.006)	–	.021 (.008)	.028 (.008)
Earnings Variability * Recall ($1,000s)[1]	154 (762)	–	–	.017 (.016)	.016 (.016)	–	.030 (.016)	.021 (.016)	–	.031 (.016)	.012 (.016)
Replacement Rate (Base period)	.618 (.208)	.400 (.109)	–	.375 (.109)	–	–	–	–	–	–	–
Replacement Rate (High quarter)	.464 (.104)	–	.627 (.165)	–	.849 (.172)	–	–	–	–	–	–
Base Period Earnings ($10,000s)[1]	15,594 (10,786)	.070 (.016)	.082 (.018)	.046 (.018)	.068 (.018)	-.040 (.014)	-.055 (.014)	-.066 (.016)	–	–	–
Weekly Benefit Amount ($100s)[1]	151.86 (52.12)	–	–	–	–	.272 (.030)	.294 (.029)	.276 (.030)	–	–	–
Two High-Quarters Earnings ($10,000s)[1]	9,768 (6,662)	–	–	–	–	–	-	–	.447 (.051)	.460 (.047)	.410 (.052)
Two High-Quarters >Maximum ($100s)[1]	80.42 (197.23)	–	–	–	–	–	–	–	-.122 (.015)	-.144 (.016)	-.134 (.016)
R^2 (adjusted)	–	.134	.134	.135	.140	.140	.136	.141	.139	.135	.140
F	–	35.4	35.4	34.2	34.5	37.2	35.8	36.0	36.9	35.5	35.7

Notes: Mean weeks of benefits paid in the sample = 16.13 (std. dev. = 10.87). Mean of the dependent variable (ln of week of benefits paid) = 2.41 (std. dev. = 1.03).
Estimates derived from a sample of 9,982 unemployment insurance claimants who filed valid claims in Washington State during 1988-89. All equations estimated include the following explanatory variables in addition to those displayed: age (4 dummies), gender, ethnicity (4 dummies), number of job referrals received from the Employment Service, number of employers during the base period, geographic location (20 dummies), and industry of employment before job loss (10 dummies).
[1] Scaling applies to regression coefficients only; not to descriptive statistics.

Table 7B

Potential Duration of Benefits, Benefits, and Unemployment Duration: Alternative Estimating Techniques for Washington State

Explanatory Variable	Mean (Std. Dev.)	1 OLS Linear	2 OLS Linear	3 OLS Semi-log	4 OLS Semi-log	5 Weibull	6 Weibull
D_{pot}	26.86 (4.17)	.284 (.034)	.274 (.033)	.013 (.003)	.012 (.003)	.027 (.005)	.026 (.005)
D_{pot} * Recall	2.93 (8.49)	-.623 (.077)	-.624 (.077)	-.057 (.007)	-.057 (.007)	-.049 (.008)	-.050 (.008)
Recall	.110 (.312)	8.217 (2.114)	8.227 (2.114)	.569 (.198)	.573 (.198)	-.357 (.227)	-.349 (.227)
Earnings Variability ($1,000s)[1]	1,401 (1,901)	.160 (.066)	.209 (.088)	.021 (.006)	.028 (.008)	.064 (.014)	.072 (.016)
Earnings Variability * Recall ($1,000s)[1]	154 (762)	.095 (.175)	.100 (.175)	.021 (.016)	.021 (.016)	-.051 (.020)	-.052 (.020)
Base Period Earnings ($10,000s)[1]	15,594 (10,786)	.481 (.167)	–	-.066 (.016)	–	-.067 (.022)	–
Weekly Benefit Amount ($100s)[1]	151.86 (52.12)	2.156 (0.319)	–	.276 (.030)	–	.225 (.046)	–
Two High-Quarters Earnings ($10,000s)[1]	9,768 (6,662)	–	3.390 (0.555)	–	.410 (.052)	–	.281 (.080)
Two High-Quarters >Maximum ($100s)[1]	80.42 (197.23)	–	-1.092 (0.170)	–	-.135 (.016)	–	-.098 (.024)
Weibull Shape parameter	–	–	–	–	–	.953 (.012)	.959 (.012)
R^2 (adjusted)	–	.124	.124	.141	.140	–	–
F	–	31.1	31.0	36.0	35.7	–	–
ln likelihood for Weibull	–	–	–	–	–	-9,975	-9,979

Notes: See notes to Table 7A. The dependent variable in the linear OLS models is the weeks of benefits paid. In the semi-log and Weibull models, the dependent variable is the natural logarithm of weeks of beneits paid.

Table 8

Simulated Potential Duration of Benefits, Benefits, and Unemployment Duration: Alternative Specifications for Illinois Using Weibull Duration Model

	FSC - Eligible Workers				FSC - Ineligible Workers			
Variable	Mean (Std. Dev.)	1	2	3	Mean (Std. Dev.)	4	5	6
Simulated D_{pot}	22.07 (6.67)	-.012 (.006)	--	-.011 (.007)	21.53 (7.00)	-.007 (.005)	--	-.007 (.006)
Simulated D_{pot} * Recall	5.15 (10.00)	.015 (.007)	--	.014 (.007)	6.88 (10.59)	.005 (.005)	--	.005 (.005)
Recall	.229 (.420)	-.388 (.165)	-.127 (.070)	-.450 (.172)	.310 (.463)	-.093 (.123)	-.052 (.055)	-.173 (.131)
Earnings Variability ($1,000s)[1]	954 (949)	--	.0003 (.022)	-.011 (.024)	1,017 (992)	--	.010 (.023)	.020 (.025)
Earnings Variability * Recall ($1,000s)[1]	228 (574)	--	.082 (.054)	.075 (.054)	341 (728)	--	.071 (.040)	.073 (.040)
Weekly Benefit Amount ($100s)[1]	121.48 (39.54)	.190 (.091)	.100 (.072)	.183 (.100)	121.01 (40.54)	.047 (.079)	-.025 (.063)	-.035 (.091)
Base Period Earnings ($10,000s)[1]	13,398 (9,069)	-.094 (.033)	-.117 (.029)	-.093 (.035)	13,135 (9,564)	-.058 (.027)	-.066 (.024)	-.053 (.028)
Weeks of Benefits Paid	22.28 (13.79)	--	--	--	19.09 (9.44)	--	--	--
ln (Weeks of Benefits Paid)	2.72 (1.09)	--	--	--	2.69 (.904)	--	--	--
Weibull Shape parameter	--	.962 (.017)	.963 (.017)	.962 (.017)	--	.718 (.015)	.718 (.015)	.718 (.015)
ln likelihood for Weibull	--	-5,845	-5,847	-5,844	--	-3,586	-3,586	-3,584

Notes: Estimates derived from samples of 4,367 FSC - Eligible UI claimants and 3,076 FSC - Ineligible UI claimants who filed valid claims in Illinois during 1984. All equations estimated include the following explanatory variables in addition to those displayed: age (3 dummies), gender, ethnicity (4 dummies), number of job referrals received from the Employment Service, number of employers during the base period, whether a dependants' allowance was received, the length of time between job loss and filing the UI claim, the labor market in which the worker was seeking a job (5 dummies), and industry of employment before job loss (10 dummies).

[1] Scaling applies to regression coefficients only; not to descriptive statistics.

Table 9

Conditional Reemployment Probabilities (Discrete Hazards), Washington State

	Potential Duration of Benefits					
	19-21 weeks		24-26 weeks		30 weeks	
Weeks until Exhaustion of Benefits	Risk Set	Adjusted Hazard	Risk Set	Adjusted Hazard	Risk Set	Adjusted Hazard
30	--	--	--	--	5,085	.0810
28	--	--	--	--	4,593	.0664
26	--	--	--	--	4,147	.0511
24	--	--	1,210	.0893	3,733	.0447
22	--	--	1,075	.0409	3,374	.0341
20	877	.0889	1,003	.0469	3,081	.0305
18	775	.0555	921	.0347	2,806	.0328
16	718	.0460	847	.0307	2,553	.0239
14	660	.0273	786	.0255	2,349	.0234
12	619	.0216	714	.0336	2,159	.0278
10	580	.0224	648	.0232	1,993	.0211
8	544	.0184	595	.0219	1,845	.0222
6	511	.0313	527	.0209	1,706	.0258
4	468	.0235	484	.0227	1,571	.0216
2	434	.0346	439	.0296	1,460	.0336
0	373	.3485	384	.2318	1,317	.3144
Expected duration of unemployment						
Estimator 1	16.55		18.02		18.35	
Estimator 2	14.46		14.06		13.95	

Table 10

Potential Duration of Benefits, Benefits, and Unemployment Duration: Alternative Specifications for Federal Supplemental Compensation (FSC) in Illinois Using Weibull Duration Model

Explanatory Variable	Mean (Std. Dev.)	1	2	3	4
FSC-eligible	.587 (.492)	.167 (.031)	.167 (.031)	.167 (.031)	.167 (.031)
FSC-eligible * Recall	.134 (.341)	-.045 (.060)	-.045 (.060)	-.045 (.060)	-.046 (.060)
Recall	.262 (.440)	.007 (.044)	.007 (.044)	.006 (.044)	.006 (.044)
Earnings Variability ($1,000s)[1]	980 (967)	.017 (.015)	.013 (.015)	.013 (.015)	.041 (.016)
Replacement Rate (Base period)[1]	.600 (.234)	-.020 (.093)	–	–	
Replacement Rate (High-quarter)[1]	.449 (.111)	–	-.103 (.197)	–	
Base Period Earnings ($10,000s)[1]	13,289 (9,277)	-.092 (.023)	-.096 (.022)	-.099 (.019)	–
Weekly Benefit Amount ($100s)[1]	121.29 (39.96)	–	–	.046 (.049)	–
Two High Quarter Earnings ($10,000s)[1]	8,059 (5,085)	–	–	–	-.105 (.070)
Two High Quarter > Maximum ($100s)[1]	59.79 (143.06)	–	–	–	-.025 (.022)
Weibull Shape parameter	–	.868 (.011)	.868 (.011)	.868 (.011)	.868 (.011)
ln likelihood for Weibull	–	-9,517	-9,517	-9,517	-9,518

Notes: Mean weeks of benefits paid in the sample = 20.96 (std. dev. = 12.28). Mean of the dependent variable (ln of weeks of benefits paid) = 2.71 (std. dev. = 1.02).
Estimates derived from a sample of 7,443 UI claimants who filed valid claims in Illinois during 1984. All equations estimated include the following explanatory variables in addition to those displayed: age (3 dummies), gender, ethnicity (4 dummies), number of job referrals received from the Employment Service, number of employers during the base period, whether a dependents' allowance was received, the length of time between job loss and filing the UI claim, the labor market in which the worker was seeking a job (5 dummies), and industry of employment before job loss (10 dummies).
[1] Scaling applies to regression coefficients only; not to descriptive statistics.

Table 11

Conditional Reemployment Probabilities (Discrete Hazards), Illinois

Weeks until Exhaustion of Benefits	Ineligible for FSC		Eligible for FSC	
	Risk Set	Hazard	Risk Set	Hazard
38	--	--	2,105	.0912
36	--	--	1,909	.0602
34	--	--	1,782	.0466
32	--	--	1,686	.0463
30	--	--	1,593	.0439
28	--	--	1,508	.0345
26	1,600	.0688	1,445	.0291
24	1,488	.0477	1,384	.0275
22	1,405	.0349	1,333	.0315
20	1,348	.0423	1,265	.0269
18	1,278	.0376	1,214	.0264
16	1,217	.0394	1,165	.0309
14	1,160	.0474	1,105	.0443
12	1,086	.0359	1,042	.1008
10	1,028	.0418	738	.0203
8	968	.0382	706	.0255
6	923	.0563	670	.0239
4	862	.0360	644	.0230
2	807	.0682	605	.0430
0	731	.3926	553	.3816

Expected Duration of Unemployment	Ineligible for FSC	Eligible for FSC
Estimator 1	21.29	24.08
Estimator 2	21.31	19.99

Unemployment Insurance and Unemployment:

Implications of the Reemployment Bonus Experiments

Carl Davidson and Stephen A. Woodbury

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Unemployment Insurance and Unemployment:
Implications of the Reemployment Bonus Experiments

Carl Davidson and Stephen A. Woodbury

February 1995

Abstract

We translate the results of the three reemployment bonus experiments that were conducted during the 1980s into (a) impacts of a 10-percentage point increase in the Unemployment Insurance (UI) replacement rate on the expected duration of unemployment; and (b) impacts of adding 1 week to the potential duration of UI benefits on the expected duration of unemployment. Our approach is to use an equilibrium search and matching model, calibrated using data from the bonus experiments and secondary sources. The results suggest that a 10-percentage point increase in the UI replacement rate increases the expected duration of unemployment by .3 to 1.1 week (a range consistent with, but only somewhat narrower than, the existing range of estimates), and that adding 1 week to the potential duration of UI benefits increases the expected duration of unemployment by .05 to .2 week (which is toward the low end of existing estimates).

I. Introduction

During the last 20 years, perhaps the most researched question about Unemployment Insurance (UI) has been whether and to what degree increases in the UI replacement rate lengthen UI recipients' jobless spells. Yet estimates of the effect of UI on unemployment have varied over a wide range. For example, Hamermesh (1977) provided an early (and much-quoted) "best estimate" that a 10-percentage point increase in the UI replacement rate (defined as the ratio of the weekly UI benefit to the average weekly wage) increases the duration of unemployment by about one-half week. But this "best estimate" was based on studies that found impacts ranging from approximately zero to over 1.5 weeks. More recent studies have continued to produce a wide range of estimates: Moffitt and Nicholson (1982) found that a 10-percentage point increase in the replacement rate increases unemployment duration by about 1 week; Solon (1985) found an impact of between one-half and one week; and Meyer (1990) found an impact on the order of 1.5 weeks. Clearly, this is a rather broad range -- one that gives relatively little guidance in deciding whether claims that UI is a serious deterrent to job search should be taken seriously and used as a basis for reforming the UI system.

A related issue that has received less attention in the literature concerns the degree to which extending the potential duration of UI benefits affects job search behavior. A number of studies have estimated the impact of adding 1 week to the potential duration of benefits on the expected duration of unemployment. All find evidence that an increase in potential duration reduces search effort and increases the average length of unemployment spells. But again the estimates vary widely: Newton and Rosen (1979) find that an additional week of potential benefits raises the expected duration of unemployment by .4 to .5 week; Moffitt and Nicholson (1982) find an impact of .1 week; Solon (1985) finds an impact of .3

week; Ham and Rea (1987) find an impact of .33 week; and Katz and Meyer (1990) find an impact of .16 to .2 week. The differences among these estimates may seem small, but they do matter. A typical UI benefit extension of 10 weeks would increase the expected duration of unemployment by as little as 1 week, or as much as 4 weeks, depending on which of these estimates is correct.

In the last decade, three social experiments have been performed in the United States that have the potential to narrow the range of estimates of how UI affects the behavior of unemployed workers (Woodbury and Spiegelman 1987; Corson, Decker, Dunstan, and Kerachsky 1992; Spiegelman, O'Leary, and Kline 1992). Each of the experiments tested one or more variants of the so-called reemployment bonus -- a cash bonus paid to UI recipients who find rapid reemployment and cut short their spell of insured unemployment. Because each of the three experiments randomly assigned UI claimants either to a control group or to a bonus-offered group, the experiments offer a potentially powerful way of discerning whether (and to what extent) UI benefits are a benign or nondistortionary income transfer.

The reemployment bonus experiments provide clear evidence that UI benefits are not merely a benign transfer. All three of the experiments found that bonus offers reduce the duration of insured unemployment of bonus-offered UI recipients. But unfortunately, there is no simple "back-of-the-envelope" way to translate the results of the reemployment bonus experiments into estimates of the effects of UI replacement rates or potential duration on unemployment duration. The essence of the reemployment bonus is the offer of cash for rapid reemployment, not a direct change in the weekly UI benefit amount or the potential duration of benefits.

Nevertheless, the reemployment bonus experiments provide convincing evidence of the effect of financial incentives on unemployment behavior. Accordingly, it is tempting to devise a way to use the observed reemployment bonus effects to infer how unemployment behavior would be altered by changes in either UI benefit amounts or potential duration.

In this paper, we attempt to solve the problem of translating the observed effects of the reemployment bonus into the effects that have been of most concern to economists and policy-makers -- that is, the effects of changes in the UI replacement rate and in the potential duration of UI benefits on unemployment duration. In section II, we offer a brief review of the three reemployment bonus experiments that have been carried out. In section III, we develop an equilibrium search/matching model that incorporates the relevant institutional characteristics of the UI system in the United States. In section IV, we describe how we make use of the results of the reemployment bonus experiments to infer key unobservable parameters of the model. Finally, in section V, we apply the model to our main question -- what are the implications of the results of the reemployment bonus experiments for the relationship between UI benefits and unemployment duration?

In addition to yielding estimates of how increasing the UI replacement rate and potential duration increase the expected duration of unemployment, the method we use yields estimates of spillover effects of the UI system. That is, UI has the potential to affect the behavior and employment outcomes of workers other than UI claimants -- for example, workers who are ineligible for UI (or "UI-ineligibles"). Because the model we use considers the effect of UI benefits on these other groups of workers -- as well as on UI claimants -- we can estimate the potentially important effect of changes in the replacement rate and potential duration of benefits on workers other than UI claimants. Such spillover effects have only rarely been considered -- Levine's work (1993) is a recent exception.

II. The Reemployment Bonus Experiments

Table 1 summarizes the three reemployment bonus experiments and their results. Central to each of the experiments is the notion of random assignment: One or more randomly chosen groups of new UI claimants were offered a reemployment bonus, and another randomly chosen group -- the control group -- received no special treatment. Random assignment was made in each of the reemployment bonus experiments by referring to the last two digits of the Social Security number. In the Illinois experiment, for example, UI claimants with a Social Security number ending in 00 through 33 were assigned to the control group, whereas those with a Social Security number ending in 34 through 66 were offered a bonus. The power of random assignment is that, if it is effectively carried out, then on average the observable and unobservable characteristics of workers in the experimental and control groups will be identical. Accordingly, the only difference between the experimental and control groups is that the experimental groups receives a "treatment," and comparing the outcomes of the control and experimental groups is sufficient to obtain an unbiased estimate of the effect of the experimental treatment on behavior.[1]

The earliest of the three reemployment bonus experiments, in Illinois, was also the simplest in that it had just one treatment -- a $500 bonus offered to new UI claimants who found a job within 11 weeks and held that job for four months. Note that the base period earnings of workers assigned to the treatment and control groups are essentially similar[2] (see column 3 of Table 1), suggesting that the random assignment was successful. (Similarly, the weekly benefit amounts of the treatment and control groups, shown in column 4, are the

[1]This is, of course, a rather rosy accounting of the virtues of social experimentation. For a discussion of the possible pitfalls of experimentation, see Spiegelman and Woodbury (1990).

[2]That is, we cannot reject the hypothesis that the difference between the base period earnings of the treatment and control groups is zero at conventional significance levels.

same.) Column 5 shows the duration of insured unemployment for the experimental and control groups -- specifically, the weeks of insured unemployment during the full year of eligibility following the initial claim for UI (the benefit year).[3] Comparing the weeks of insured unemployment received by the Illinois treatment group with the Illinois control group suggests that the $500 bonus reduced the duration of insured unemployment by .71 week (19.27 - 18.56 = .71).[4] This treatment effect, shown in column 6, suggests that financial incentives do influence the job search behavior of unemployed workers.

Further, it is clear from data on earnings after reemployment (column 7) that the jobs accepted by the bonus-offered workers did not pay significantly less than the jobs accepted by the control group. The implication is that the bonus-offered workers did not lower their reservation wages, cut short productive job search, and accept a poor job match (which would be evidenced by a lower wage) simply to qualify for the bonus. Rather, the results suggest that to qualify for the bonus, bonus-offered workers increased the intensity of their job search.

The two experiments that followed Illinois -- one in Pennsylvania and the other in Washington State -- each tested several bonus offers. These new treatments varied in two ways. First, the <u>length of the qualification period</u> -- the time within which a worker needed

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[3]We focus on the duration of insured unemployment during the full benefit year on the assumption that this best captures the overall impact of the bonus on a worker's propensity to become and remain reemployed.

[4]Note that this is the bonus effect for workers who were eligible for 26 weeks of state-regular UI benefits. About half of the workers enrolled in the Illinois experiment were eligible for an additional 12 weeks of Federal Supplemental Compensation (FSC). The bonus impact for these FSC-eligible workers appears to have been greater: for state-regular eligibles and FSC-eligibles <u>combined</u> the bonus impact was 1.13 weeks (Woodbury and Spiegelman 1987), and for FSC-eligibles alone the bonus impact was 1.8 weeks (Davidson and Woodbury 1991). In this paper, we restrict our attention to Illinois claimants who were eligible only for state-regular benefits. This makes the Illinois results comparable with the Pennsylvania and Washington results, in which UI claimants were eligible only for state-regular benefits.

to find reemployment in order to qualify for a bonus -- varied across the treatments. In Pennsylvania, qualification periods of 6 and 12 weeks were tried (see column 1 of Table 1). In Washington, qualification periods of 20% and 40% of a claimant's potential duration of UI plus 1 week -- or about 6 and 11 weeks -- were tried. Second, the <u>size of the bonus offer</u> varied across the treatments. In Pennsylvania, bonus offers equal to 3 times and 6 times the weekly UI benefit amount were tried -- that is, about $500 and $1,000 on average (see column 2 of Table 1). In Washington, bonus offers of 2 times, 4 times, and 6 times the weekly UI benefit amount were tried -- that is, roughly $300, $600, and $900 on average. In both Pennsylvania and Washington (as in Illinois), a worker had to hold the new job for 4 months in order to receive a bonus.

The results of the Pennsylvania and Washington experiments are not entirely consistent, but do suggest that larger bonus offers and longer qualification periods tend to reduce insured unemployment by more than smaller bonus offers and shorter qualification periods (see column 7 of Table 1). One inconsistency arises in Washington, where the <u>long</u> qualification period/<u>low</u> bonus treatment had a greater effect than would be expected based on the results of the other Washington treatments. (Alternatively, the <u>long</u> qualification period/<u>medium</u> bonus treatment could be viewed as anomalously low.) In Pennsylvania, the <u>short</u> qualification period/<u>low</u> bonus treatment had nearly the same effect as the <u>short</u> qualification period/<u>high</u> bonus treatment, which is also inconsistent.[5]

Comparing results across the three experiments poses additional puzzles. Mainly, the Illinois treatment effect of .71 week is larger than one would expect given the results of

[5]Note that is both cases, the formal statistical tests fail to reject the hypothesis that the treatments had no effect. We take the point estimates at face value, however, on the assumption that large enough samples would reveal effects of the magnitude reflected by the point estimates.

similar treatments in Pennsylvania (the long qualification period/low bonus treatment) and Washington (the long qualification period/medium bonus treatment), although this latter treatment, as already noted, is anomalously low in comparison with the other Washington results.

Although providing a full explanation of the differences among the various bonus offers is beyond the scope of this paper, the differences across experiments may be explained in part by reference to features of the experiments other than the bonus offers themselves. For example, base period earnings were highest in Washington State and lowest in Illinois, but UI benefit levels were highest in Pennsylvania. As a result, the mean replacement rate (defined simply as the ratio of the average weekly UI benefit to the average weekly base period earnings) was highest in Pennsylvania at about .6, and lower in both Illinois and Washington at about .5 (the replacement rates are not shown in Table 1). Also, workers' weekly earnings before and after the spell of insured unemployment were nearly identical in Illinois, but in Pennsylvania, post-unemployment earnings were somewhat lower than pre-unemployment earnings, whereas in Washington post-unemployment earnings were higher than pre-unemployment earnings. Finally, average spells of insured unemployment were considerably longer in Illinois (over 19 weeks on average for controls) than in either Pennsylvania or Washington (about 16 and 15 weeks respectively). Some of these differences could influence the experimental outcomes, and one advantage of the model developed next is that it accounts for these differences in translating the bonus effects into other behavioral impacts.

In summary, two results are consistent across all three bonus experiments. First, all three experiments provide evidence that financial incentives do influence job search behavior and the duration of unemployment, with larger bonus offers and longer qualification periods tending to induce workers to shorten their unemployment spells by more than smaller bonus offers and shorter qualification periods. Second, the shorter spells of insured unemployment that resulted from the bonus offers did not come at the expense of lower earnings after reemployment. This latter is arguably the most important finding of the reemployment bonus experiments because it suggests that bonus-offered workers shortened their spells of unemployment by increasing their search intensity rather than by lowering their reservation wages. By inference, the finding suggests that UI lengthens unemployment spells mainly by reducing job search intensity rather than by increasing the reservation wage.

III. The Model

Our goal is to translate the experimental results of the reemployment bonus experiments into impacts of UI generosity on unemployment duration. To do this, we employ a partial equilibrium matching model of the labor market in which unemployed workers search randomly among firms for employment. After workers have applied for a job, firms with vacancies randomly select workers from their pool of applicants. Each unemployed worker chooses his or her optimal search effort -- the number of firms to contact -- by equating the marginal benefit from increasing search effort with the associated marginal cost. To obtain results on the impact of changes in the UI replacement rate and potential duration of UI benefits, we solve the model for different replacement rates and potential benefit durations, and compare outcomes. The basic structure of our model is similar to the one used in Davidson and Woodbury (1993, 1995).

We require a model that is institutionally rich, yet tractable. Accordingly, we introduce three categories of unemployed workers. The first consists of workers who are ineligible for unemployment insurance. We refer to such workers as UI-ineligibles and use U_i to denote the number of such workers in the steady-state equilibrium. UI-ineligibles are generally workers with relatively weak attachments to the labor force -- new labor force entrants and reentrants -- and typically account for roughly 60% of the unemployed (Blank and Card 1991). We denote the proportion of the unemployed who are UI-ineligibles by q, and set it equal to .6.

The second category of unemployed workers consists of those who are eligible for UI, but do not bother to claim their benefits. We refer to these workers as UI-eligible non-claimants and use U_k to denote the number of them in the steady-state equilibrium. We include these workers in the model because the fraction of UI-eligible workers who claim benefits -- the UI take-up rate -- is less than 100%. If we denote the UI take-up rate by k, then the proportion of UI-eligible workers who fail to claim their benefits is 1 - k. Recent work by Blank and Card (1991) indicates that k falls in a range between .65 and .75. (We set k = .75 in our model. This choice is somewhat arbitrary, but the results are insensitive to changes in k in the range of .65 to .75.) Although there are probably a variety of reasons why workers who are eligible for benefits do not claim those benefits, we assume that they expect to be reemployed rapidly, so that the expected benefit from claiming their benefits falls short of the cost of doing so. This assumption is not crucial for what follows.

Finally, we refer to the remainder of the unemployed as UI-eligible claimants and use U_t to denote the number of such workers who are receiving benefits and are in their t^{th} period of search. We assume that these workers exhaust their eligibility after T periods of unemployment and use U_e to denote the number of UI-eligible claimants who have exhausted their benefits in the steady-state equilibrium.

We describe the model in four steps. First, we introduce three accounting identities that describe the distribution of the workforce (between employment and unemployment), the distribution of jobs (between employment and vacancies), and the distribution of unemployed workers (among UI-ineligibles, UI-eligible non-claimants, and UI-eligible claimants, both recipients and exhaustees). In the second step, we equate the flows into and out of each state of unemployment to yield a steady-state. Third, we demonstrate how search effort translates into reemployment probabilities. Finally, we define optimal search effort for each unemployed worker.

A. Identities

Since UI-claimants must be certified for benefits every two weeks, we measure time in two-week intervals. Let F denote the total number of jobs available, J represent the total number of available jobs that are filled, and V represent the number of job vacancies in the steady-state equilibrium. Then, since all jobs are either filled or vacant, we have:

(1) $$F = V + J.$$

In addition, since all workers in the labor force must be either employed or unemployed, we have:

(2) $$L = U + J$$

where L denotes the total number of workers and U represents total unemployment.

The final identity divides unemployed workers into UI-ineligibles, UI-eligible non-claimants, and UI-eligible claimants who are receiving benefits, and UI-eligible claimants who have exhausted their benefits:

(3) $$U = U_i + U_k + \Sigma_{t=1,T} U_t + U_e.$$

B. Steady-State Conditions

The second set of equations equates the flows into and out of each employment state. These equations must hold to insure that total unemployment and its composition remain constant in steady-state. Let s denote the rate of job separation or turnover -- that is, the probability that a randomly chosen employed worker will lose his or her job in any given period. Thus, sJ worker lose their job in each period. Of these, qsJ are UI-ineligible. It follows that the flow into state U_i is qsJ. To calculate the flow out of this state, let m_i denote the reemployment (or job match) probability for a typical UI-ineligible worker. Then the flow out of U_i is m_iU_i. Equating these flows yields the first steady-state condition:

$$(4) \qquad qsJ = m_iU_i$$

Applying the same logic to the class of UI-eligible non-claimants yields

$$(5) \qquad (1\text{-}q)(1\text{-}k)sJ = m_kU_k$$

where m_k denotes the reemployment probability for UI-eligible non-claimants.

Next, turn to the UI-eligible claimants. Let m_t denote the reemployment probability for a UI-eligible claimant in the t^{th} period of search and let m_e play the same role for UI-eligible claimants who have exhausted their benefits. Then, of the U_t workers in their t^{th} period of search, m_tU_t find jobs and the remaining $(1-m_t)U_t$ do not. Those who find jobs move to state J while those who do not move on to state U_{t+1}. It follows that all workers who begin the period in state U_t flow out of that state at the end of the period and that the flow into state U_{t+1} is given by $(m_t\text{-}1)U_{t\text{-}1}$. Equating these flows yields the following steady-state conditions

$$(6) \qquad (1\text{-}q)ksJ = U_1$$

$$(7) \qquad (1\text{-}m_{t\text{-}1})U_{t\text{-}1} = U_t \qquad\qquad \text{for } 2 \leq t \leq T$$

UI-eligible claimants who have exhausted their benefits leave state U_e if and only if they find a job, which happens with probability m_e. Entry into state U_e occurs if workers fail to find employment after T periods of search. Thus, the flows into and out of state U_e are equal if:

$$(8) \qquad (1-m_T)U_T = m_e U_e$$

If equations (4)-(8) hold, unemployment and its composition remain constant over time; that is, a steady-state exists.

C. Reemployment Probabilities

The probability that a searching worker finds a job is a function of his or her own search effort, the search effort of other workers, and the slackness or tightness of the labor market. We use p_t to denote the search effort of a UI-eligible claimant in the t^{th} period of search. The terms p_i, p_k, and p_e denote the search effort of UI-ineligible workers, UI-eligible non-claimants, and UI-eligible claimants who have exhausted their benefits, respectively. Each of the p terms gives the probability that a worker contacts a firm and applies for a job (or, if $p > 1$, the number of firms contacted by the worker) in any given period. Assuming that workers choose firms at random, the probability that any given firm has a vacancy is V/F. Thus, the probability that a worker contacts a firm with a vacancy is p(V/F). If we let λ denote the average number of applications filed per firm, then the probability that a worker gets a job <u>conditional</u> on applying to a firm with a vacancy is $(1 - e^{-\lambda})/\lambda$ (see Davidson and Woodbury 1993). Thus, the probability of a UI-eligible claimant in the t^{th} period of search finding a job is given by:

(9) $m_t = p_t(V/F)[(1 - e^{-\lambda})/\lambda]$ for $1 \leq t \leq T$.

where $\lambda = (1/F)[p_iU_i + p_kU_k + \Sigma_{t=1,T}p_tU_t + p_eU_e]$. The analogous conditions for the UI-ineligibles, UI-eligible non-claimants, and UI-exhaustees are:

(10) $m_i = p_i(V/T)[(1 - e^{-\lambda})/\lambda]$

(11) $m_k = p_k(V/F)[(1 - e^{-\lambda})/\lambda]$

(12) $m_e = p_e(V/F)[(1 - e^{-\lambda})/\lambda]$.

Note that the search effort of other workers enters into each workers reemployment probability through λ.

D. Optimal Search Effort

We assume that workers choose search effort to maximize expected lifetime income. Workers can increase the probability of reemployment by increasing search effort, but doing so is costly. We assume that the cost of search is a function of search effort, p, and specify the search cost function as cp^z, where c and z are search cost parameters. Note that z (> 1) denotes the elasticity of search costs with respect to search effort. We assume that the parameter c differs between UI-eligible and UI-ineligible workers (we refer to it as c for UI-eligibles and c_i for UI-ineligibles), but that z is the same for all.

To calculate expected lifetime income we must consider both the current and future prospects faced by the each worker. For example, let V_t denote the expected lifetime income of an unemployed UI-eligible claimant in the t^{th} period of search; let V_w denote the expected lifetime income of an employed UI-eligible claimant; let w represent the wage earned by such a worker when employed; and let x denote unemployment benefits. Then, an unemployed UI-eligible claimant in the t^{th} period of search earns $x - c(p_t)^z$ currently. With probability m_t this worker finds a job and can expect to earn V_w in the future. With the remaining probability, $1-m_t$, the worker remains unemployed and can expect to earn V_{t+1} in the future. Therefore,

$$(13) \quad V_t = x - c(p_t)^z + [m_t V_w + (1 - m_t)V_{t+1}]/(1 + r) \quad \text{for } 1 \leq t \leq T.$$

Note that future income is discounted, with r denoting the interest rate. An analogous condition describes the expected lifetime income of workers in every other state of unemployment. If we let V_e and V_i denote the expected lifetime incomes of a UI-claimant who has exhausted benefits and an unemployed UI-ineligible worker, then we have:

$$(14) \quad V_e = - c(p_e)^z + [m_e V_w + (1 - m_e)V_e]/(1 + r)$$

$$(15) \quad V_i = - c_i(p_i)^z + [m_i V_{wi} + (1 - m_i)V_i]/(1 + r)$$

Recall that V_w is the expected lifetime income of an employed UI-eligible worker, and let V_{wi} denote the expected lifetime income of an employed UI-ineligible worker. To calculate V_w and V_{wi} we follow the procedure outlined already. Current income equals the worker's wage, w (or w_i if UI-ineligible). With probability (1 - s) this worker keeps his job for another period and continues to earn V_w (or V_{wi} if UI-ineligible). With probability s the worker loses his job and has to search for new employment, resulting in a future income of V_1 (or V_i if UI-ineligible). Therefore,

$$(16) \quad V_w = w + [sV_1 + (1 - s)V_w]/(1 + r)$$

$$(17) \quad V_{wi} = w_i + [sV_i + (1 - s)V_{wi}]/(1 + r).$$

For each unemployed worker, search effort is chosen to maximize expected lifetime income. Therefore, we have the following equations defining optimal search effort for all but one possible state of unemployment

$$(18) \quad p_t = \arg\max V_t \qquad \text{for } 1 \leq t \leq T$$

$$(19) \quad p_{\bullet} = \arg\max V_{\bullet}$$

$$(20) \quad p_i = \arg\max V_i.$$

The one exception is made for UI-eligible non-claimants. Presumably, these workers do not claim UI benefits because they do not expect to be unemployed for a significant length of time -- that is, they expect to be able to find jobs relatively easily and with little effort. Therefore, we treat these workers differently, by assigning them a high reemployment probability and ignoring their search decision. Provided that their reemployment probability is set high enough (so that their expected duration of unemployment is roughly half the expected duration faced by UI-eligible claimants), our results are not sensitive to this assumption.

IV. Calibration

In order to solve the model, we must first set values of its parameters. We begin by dividing the model's parameters into three categories. First, we have parameters that have either been estimated directly in previous work or that can be inferred from estimates of other variables in previous work. These include the separation rate (s), total jobs available (F), the size of the labor force (L), the fraction of unemployed workers who are ineligible for UI benefits (q), the UI take-up rate (k), and the interest rate (r).

The second set consists of variables that are observable and can be taken directly from the data collected to analyze the three reemployment bonus experiments. Included in this set are wages or earnings (w and w_i) and unemployment benefits (x).

The third set consists of variables that have not been previously estimated and that cannot be observed directly -- the search cost parameters c, c_i, and z.

The reasoning used to obtain values of parameters in the first category is described in detail elsewhere (Davidson and Woodbury 1993, 1995). Here, we simply report the range of values considered and cite the sources used to support our choices. Recall that we measure time in 2-week intervals since UI claimants are typically certified for 2 weeks of benefits at a time. With that in mind, we use values of s (the bi-weekly separation rate) ranging from .006 to .014, with s = .010 considered to be the best estimate (Eherenberg 1980; Clark and Summers 1982; and Murphy and Topel 1987). Since the system of equations is homogeneous of degree zero in F and L, we can set L = 100 without loss of generality. We consider values for F ranging from 95 to 97.5, with F = 96.25 considered the best estimate (inferred from values of the ratio of unemployment to vacancies reported in Abraham 1983). As mentioned above, we set q (the fraction of unemployed workers who are ineligible for UI benefits) equal to .6 and k (the UI take-up rate) equal to .75 (Blank and Card 1991). For the interest rate, we consider values ranging from .002 to .020 with r = .008 considered the best estimate. Thus, we have a "reference case" in which s = .01, r = .008, and T = 96.25. As we show below, as long as these parameters remain in the ranges described above, our results are remarkably robust to changes in the parameter values.

The wages and UI benefit levels used for each bonus treatment are displayed in Table 1. (Since Table 1 reports values in weekly terms, we multiply by 2.) For each treatment, we use the average of the control group and the treatment group. For example, in Pennsylvania the bi-weekly UI benefit amount for the control group was approximately $328, and the bi-weekly UI benefit amount for the long qualification period/high bonus treatment group was $330. Therefore, when analyzing the Pennsylvania long qualification period/high bonus treatment, we set x = $329.[6]

We use the information we have on bonus effects to infer values of the search cost parameters. For a given set of search cost parameters, the model predicts an expected duration of unemployment for each class of worker and a bonus effect for UI-eligible claimants. Let D denote the expected duration of unemployment <u>predicted by the model</u> in the absence of the bonus for UI-eligible claimants, and let ΔD denote these workers' bonus-induced change in unemployment duration, again as <u>predicted</u> by the model. The <u>actual</u> values of D and ΔD for each experiment are reported in Table 1. For example, in Illinois D = 19.27 and ΔD = -.71. For each experiment, we choose search cost parameters such that

[6]The proper wage to use in implementing the model is arguably the reemployment wage (column 7 of Table 1) rather than the base period wage. As it turns out, the base period wage and the reemployment wage are close enough that use of either yields similar results.

the values of D and ΔD predicted by the model match those reported in Table 1.[7] This yields a vector of search cost parameters that makes the model's prediction as close as possible to the actual outcome of each experimental treatment.

To investigate the impact of varying x (the level of UI benefits), and T (the potential duration of UI benefits), we solve the model for a variety of x and T values and compare the outcomes. Increasing x or T decreases the opportunity cost of unemployment for UI-eligible claimants, resulting in a decrease in search effort. The decrease in search effort increases their duration of unemployment, decreases steady-state employment, and may increase the number of jobs held by other workers. By solving the model for different values of x and T, we can estimate the magnitude of these different impacts.

[7]The expected durations of unemployment and the treatment impacts shown in Table 1 are actually in terms of durations of insured unemployment and changes in insured unemployment. Our model, on the other hand, is stated in terms of spells of actual joblessness (not just insured unemployment). Accordingly, we have adjusted the treatment effects shown in Table 1 so as to reflect the expected durations of unemployment and changes in unemployment duration induced by each bonus offer. These adjustments are described in Davidson and Woodbury (1991).

V. Results

Table 2 displays the main results of our efforts to translate the reemployment bonus effects into (a) impacts of a 10-percentage point increase in the UI replacement rate on unemployment duration; and (b) impacts of adding 1 week to the potential duration of UI benefits.

Column 1 of Table 2 shows the estimated impact of a 10 percentage point increase in the UI replacement rate -- the ratio of weekly UI benefits to pre-unemployment wage -- on the expected duration of unemployment of UI-eligible claimants. The estimates cover a range of roughly .5 to 1.1 weeks. That is, a 10 percentage point increase in the UI replacement rate is predicted to increase the expected duration of UI-eligible workers by between .5 and 1.1 weeks. This range is in the middle of the existing empirical estimates reviewed in the introduction -- that range runs from a low of about zero to a high of over 1.5 weeks. However, most of the range lies above Hamermesh's (1977) "best estimate" of one-half week. The range we obtain is more in keeping with some of the more recent findings, such as Solon's (1985), whose estimates are in the range of .5 to 1 week.

The second column of Table 2 shows that increases in the UI replacement rate can be expected to shorten the unemployment spells of UI-ineligibles. That is, since an increase in the UI replacement rate reduces the search effort of UI claimants, the competition for jobs is reduced so that UI-ineligibles (whose search effort is essentially unchanged by the increase in the UI replacement rate) have a higher probability of getting a job offer when they apply for a job. The estimates suggest that a 10 percentage point increase in the UI replacement rate shortens the expected unemployment spells of UI-ineligibles by about half a day to a day (that is, on the order of one-tenth to two-tenths of a week). Although this is not a large effect, it does suggests that UI benefit increases reduce job competition and make it easier for UI-ineligibles to find work.

We have omitted three of the six Washington treatments from the results shown in Table 2 -- the short qualification period/low bonus treatment, the short qualification period/medium bonus treatment, and the long qualification period/medium bonus treatment. These three bonus treatments produced relatively small effects and thus generated small predicted impacts on the expected duration of unemployment. For all three, the estimated impact of a 10 percentage point increase in the replacement rate on D falls in the range of .3 to .4 week, and the impact on D_i falls in the range of -.1 to -.05.

Whether we should expand the range of estimates to include these Washington treatments is a judgement call. The treatment effects in question were so small that they are inconsistent with the findings from the other bonus offers in Washington as well as in Illinois and Pennsylvania. On the other hand, excluding findings of a small treatment effect in three cases out of a total of 11 is quite arbitrary -- we may be throwing away real information here. In any case, if we include these three Washington treatments in our range of estimates, then we would conclude that a 10 percentage point increase in the UI replacement rate is predicted to increase the expected duration of UI-eligible workers by between .3 and 1.1 weeks. This range is consistent with existing estimates of the disincentive effects of UI, but the extent to which it narrows that range is disappointing. Indeed, from the viewpoint of policy, a range of .3 to 1.1 weeks is hardly more informative than a range of 0 to 1.5 weeks.

The third column of Table 3 shows the estimated impact of a 1-week increase in the potential duration of UI benefits on the expected duration of unemployment for UI-eligible claimants. The estimates fall in the range of .1 to .2 week, implying that a 10-week benefit extension would increase the expected duration of unemployment by between 1 and 2 weeks.

These estimates are clearly toward the low end of the existing empirical estimates of the impact of extending benefits -- recall that those estimates fall in the range of .1 to .4 week.

The fourth column of Table 2 shows that extending the potential duration of benefits can be expected to make it slightly easier for UI-ineligibles to find jobs. Specifically, the estimates suggest that a 1-week increase in the potential duration of benefits shortens the expected duration of unemployment of UI-ineligibles by about one-quarter of a day. This is a very small effect, but it illustrates that job competition is reduced when UI benefits are extended.

The three omitted Washington treatments yield relatively small estimates of the impact of extending the potential duration of UI benefits. For those three treatments, the estimated impact of a 1-period increase in the potential duration of UI benefits on D is in the range of .05 to .08, and the impact on D_i is in the range of -.02 to -.01. If we include these three treatments, then our range of estimates widens, and we would conclude that a 1 week increase in the potential duration of benefits increases the expected duration of unemployment by between .05 and .2 week. The implication is that <u>a 10-week benefit extension would increase the expected duration of unemployment by between one-half and 2 weeks</u>. Again, this is clearly at the low end of existing estimates, and suggests that the disincentive effects UI extensions may be less than previously believed.

It is important to determine whether the impacts found above vary depending on the initial replacement rate or the initial potential duration of unemployment. That is, if the replacement rate were .1 or .9 to begin with (rather than .5 or .6), would the impacts that we estimate be different? Similarly, if the potential duration of benefits were 16 or 40 weeks to begin with (rather than 26 weeks), would the results differ? Table 3 shows that the estimated impacts do not vary much with the initial UI replacement rate or with the initial potential duration of benefits. For example, the results shown suggest that, depending on the initial replacement rate, a 10 percentage point increase in the replacement rate would lengthen unemployment by as little as .761 week or as much as .816 week (see the first and second columns of Table 3). This is a variation of only a quarter of a day (.055 week) in response to dramatic variation in the initial replacement rate. Similarly, the variation in response to a 1 week benefit extension that results from changing the initial potential duration of benefits is of little significance (see the right three columns of Table 3). Similar results were found for Pennsylvania and Washington. We conclude that the results are largely insensitive to the initial replacement rate or initial potential duration of benefits.

It is also important to examine the sensitivity of the results to variation in some of the key parameters that we have obtained from secondary sources. Table 4 shows how the main estimates vary with changes in the separation rate (s) and total available jobs (F). We show results for the reference case (s = .01, F = 96.25), for high and low values of s (.006 being low and .014 being high), and for high and low values of F (95 being low and 97.5 being high). The sensitivity analysis is shown for the Illinois treatment, for one treatment in Pennsylvania (the long qualification period/high bonus treatment), and for one Washington treatment (the short qualification period/high bonus treatment). (The Pennsylvania and Washington treatments selected each gave results that were in the middle of the range of their respective experiment.)

The main finding of the sensitivity analysis shown in Table 4 is that the results are generally quite insensitive to changes in F (total available jobs), but somewhat sensitive to change in s (the separation rate). Consider first the impact of a 10 percentage point increase in the replacement rate. The Pennsylvania reference case shown suggests that such an increase would lengthen unemployment by .627 week. For low F, the estimate is .629, and for high F, the estimate is .627. Hence, the results are robust to variation in F. (Similar results obtain for the Illinois and Washington cases shown.) But for low s, the estimate is .677, whereas for high s, it is .577. Thus, we have variation of about one-half day (.1 week) in response to varying s between .006 and .014. (Again, the Illinois and Washington cases shown give similar results.) Based on this finding, it might be wise to widen further (perhaps by .1 week on each side) the range discussed above for the impact of a 10 percentage point increase in the UI replacement rate. But doing so would not basically alter our conclusions.

Consider next the impact of a 1 week extension of the potential duration of UI benefits. The Pennsylvania reference case suggests that a 1 week extension would lengthen unemployment by .15 week. For low F, the estimate is .178, and for high F, the estimate is again .15. For low s, the estimate is .177, whereas for high s, it is .144. These variations -- about .03 week in each case, or less than a quarter of a day -- are probably too small to worry about (the Illinois and Washington cases shown give similar results.) However, they may suggest a need to broaden slightly the range discussed above for the impact of a 1 week extension of UI benefits.

In short, the results shown in Table 4 suggest that choosing different values of the s and F parameters might widen slightly the estimated ranges of UI impacts, but would not change our basic inferences.

VI. Discussion and Conclusions

Our main goal has been to translate the effects of reemployment bonus offers, as estimated in three separate UI field experiments, into estimates of the disincentive effects of UI that have been of most concern to economists and policy-makers -- that is, the effects of changes in the UI replacement rate and the potential duration of UI benefits on unemployment duration. An advantage of the findings presented here is that the logic of verification underlying them is quite different from that underlying earlier empirically-based findings on the incentive effects of the UI system. Yet the estimates presented clearly fall within the ranges of the earlier estimates, and arguably narrow those ranges.

We have four main findings. First, a 10 percentage point increase in the UI replacement rate can be expected to increase the unemployment duration of UI claimants by between .3 and 1.1 weeks (see Table 2 and the accompanying discussion). Existing empirical work offers a somewhat broader range than this, placing the expected increase in unemployment duration anywhere from zero to slightly over 1.5 weeks. The estimates presented here might be viewed as providing evidence that the range may be somewhat narrower -- but not much narrower -- than previously estimated. However, from a policy perspective, there is some question whether a range of .3 to 1.1 weeks is any more informative than a range of 0 to 1.5 weeks.

Second, we find that a 1 week increase in the potential duration of benefits increases the expected duration of unemployment by between .05 and .2 week (see again Table 2 and the accompanying discussion). The implication is that a 10-week benefit extension would increase the expected duration of unemployment by between one-half week and 2 weeks. This is clearly at the low end of existing estimates of the disincentive effects of UI benefit extensions.

Third, increases in the UI replacement rate and the potential duration of benefits reduce the job search intensity of UI claimants so that unemployed workers who are ineligible for UI face less competition for jobs. The result is shorter spells of unemployment for UI-ineligibles. We estimate that a 10 percentage point increase in the UI replacement rate shortens the expected unemployment spells of UI-ineligibles by about one-half day to a day. Also, a 1-week increase in the potential duration of benefits shortens the expected duration of unemployment of UI-ineligibles by about one-quarter of a day. These are admittedly very small effects, but they do illustrate that increasing the generosity of UI benefits reduces job competition and has benefits for workers who are ineligible for UI.

Fourth, we find that the disincentive effects of increases in the UI replacement rate and extensions of UI benefits are invariant to the initial replacement rate or the initial potential duration of benefits. That is, at least when they are averaged over fairly large groups, the disincentive effects of UI appear to be similar whether the initial replacement rate is high or low, and whether the initial potential duration of benefits is high or low (see Table 3 and the accompanying discussion).

Because these estimates are based on randomized trials, they are arguably free of many of the complicating and contaminating factors that plague nonexperimental estimates. Moreover, there is no particular reason to favor or disfavor any of the estimates, in that each arises from a similar experimental design that was implemented and monitored with some care. In that respect, it is striking that we find a range of estimates that is nearly as broad as that in the existing literature. Existing studies are based on various data and various econometric techniques, each of which might be expected to add variation to the range of estimates independent of variation in the actual behavior underlying those estimates. In other words, the results presented here suggest substantial variation in the behavior of unemployed workers, even when measurement of that behavior is averaged over rather large groups. At the least, the results suggest that it is unwise -- and perhaps futile -- to try to concoct summary "best estimates" of the disincentive effects of increasing the UI replacement rate or extending the potential duration of UI benefits.

References

Atkinson, A.B., J. Gomulka, J. Micklewright, and N. Rau. "Unemployment Benefit, Duration, and Incentives in Britain: How Robust Is the Evidence?" Journal of Public Economics 23 (1984): 3-26.

Atkinson, Anthony B. and John Micklewright. "Unemployment Compensation and Labour Market Transitions: A Critical Review." Journal of Economic Literature 29 (December 1991): 1679-1727.

Blank, Rebecca M. and David E. Card. "Recent Trends in Insured and Uninsured Unemployment: Is There an Explanation?" Quarterly Journal of Economics 106 (November 1991): 1157-1189.

Burtless, Gary. "Unemployment Insurance and Labor Supply: A Survey," in Unemployment Insurance: The Second Half Century, edited by W. Lee Hansen and James F. Byers. Madison, WI: University of Wisconsin Press, 1990. Pp. 69-107.

Classen, Kathleen P. "Unemployment Insurance and Job Search." In Studies in the Economics of Search, edited by S. A. Lippman and J. J. McCall. Amsterdam: North-Holland, 1979. Pp. 191-219.

Corson, Walter and Walter Nicholson. The Federal Supplemental Benefits Program. Kalamazoo, MI: W. E. Upjohn Institute for Employment Research, 1982.

Corson, Walter, Paul Decker, Shari Dunstan, and Stuart Kerachsky. Pennsylvania Reemployment Bonus Demonstration: Final Report. Unemployment Insurance Occasional Paper 92-1. Washington, D.C.: U.S. Department of Labor, Employment and Training Administration, 1992.

Davidson, Carl and Stephen A. Woodbury. "Effects of a Reemployment Bonus under Differing Benefit Entitlements, or, Why the Illinois Experiment Worked." Manuscript, Department of Economics, Michigan State University, and W. E. Upjohn Institute, March 1991.

Davidson, Carl and Stephen A. Woodbury. "The Displacement Effect of Reemployment Bonus Programs." Journal of Labor Economics 11 (October 1993): 575-605.

Davidson, Carl and Stephen A. Woodbury. "Wage Subsidies for Dislocated Workers." Paper prepared for the Advisory Council on Unemployment Compensation, January 1995.

Decker, Paul T. and Christopher J. O'Leary. An Analysis of Pooled Evidence for the Pennsylvania and Washington Reemployment Bonus Demonstrations. Unemployment Insurance Occasional Paper 92-7. Washington, D.C.: U.S. Department of Labor, Employment and Training Administration, 1992.

Ehrenberg, Ronald G. and Ronald L. Oaxaca. "Unemployment Insurance, Duration of Unemployment, and Subsequent Wage Gain." American Economic Review 66 (December 1976): 754-766.

Ham, John C. and Samuel A. Rea, Jr. "Unemployment Insurance and Male Unemployment Duration in Canada." Journal of Labor Economics 5 (July 1987): 325-353.

Hamermesh, Daniel S. Jobless Pay and the Economy. Baltimore, MD: Johns Hopkins University Press, 1977.

Hight, Joseph. "Insured Unemployment Rates, Extended Benefits, and Unemployment Insurance Exhaustions." In Proceedings of the Twenty-Eighth Annual Winter Meeting of the Industrial Relations Research Association, edited by James L. Stern and Barbara D. Dennis. Madison, WI: IRRA, 1976. Pp. 242-249.

Holen, Arlene. "Effects of Unemployment Insurance Entitlement on Duration and Job Search Outcome." Industrial and Labor Relations Review 30 (July 1977): 445-450.

Katz, Lawrence F. and Meyer, Bruce D. "The Impact of the Potential Duration of Unemployment Benefits on the Duration of Unemployment." Journal of Public Economics 41 (February 1990): 45-72.

Levine, Phillip B. "Spillover Effects between the Insured and Uninsured Unemployed." Industrial and Labor Relations Review 47 (October 1993): 73-86.

Meyer, Bruce D. "Unemployment Insurance and Unemployment Spells." Econometrica 58 (July 1990): 757-782.

Moffitt, Robert. "The Effect of the Duration of Unemployment Benefits on Work Incentives: An Analysis of Four Data Sets." Unemployment Insurance Service Occasional Paper 85-4, Employment and Training Administration, U.S. Department of Labor, 1985a.

Moffitt, Robert. "Unemployment Insurance and the Distribution of Unemployment Spells." Journal of Econometrics 28 (April 1985b): 85-101.

Moffitt, Robert and Walter Nicholson. "The Effect of Unemployment Insurance on Unemployment: The Case of Federal Supplemental Benefits." Review of Economics and Statistics 64 (February 1982): 1-11.

Newton, Floyd C. and Harvey S. Rosen. "Unemployment Insurance, Income Taxation, and Duration of Unemployment: Evidence from Georgia." Southern Economic Journal 45 (January 1979): 773-784.

Solon, Gary. "Work Incentive Effects of Taxing Unemployment Benefits." Econometrica 53 (March 1985): 295-306.

Spiegelman, Robert G., Christopher J. O'Leary, and Kenneth J. Kline. The Washington Reemployment Bonus Experiment: Final Report. Unemployment Insurance Occasional Paper 92-6. Washington, D.C.: U.S. Department of Labor, Employment and Training Administration, 1992.

Spiegelman, Robert G. and Stephen A. Woodbury. "Controlled Experiments and the Unemployment Insurance System," in Unemployment Insurance: The Second Half Century, edited by W. Lee Hansen and James F. Byers. Madison, WI: University of Wisconsin Press, 1990. Pp. 355-392.

Woodbury, Stephen A. and Robert G. Spiegelman. "Bonuses to Workers and Employers to Reduce Unemployment: Randomized Trials in Illinois." American Economic Review 77 (September 1987): 513-530.

Wunnava, Phanindra V. and Syed Ali Raza Mehdi. "The Effect of Unemployment Insurance on Unemployment Rate and Average Duration: Evidence from Pooled Cross-Sectional Time-Series Data." Applied Economics Letters 1 (1994): 114-118.

Table 1
Basic Features of the Reemployment Bonus Experiments: Means (with Standard Errors in Parentheses)

Experimental Treatment	Qualification Period (1)	Bonus Amount (2)	Weekly Base Period Earnings (3)	Weekly Benefit Amount (4)	Weeks of Insured Unemployment (Benefit Year) (5)	Treatment Effect in weeks (6)	Weekly Reemployment Earnings (7)
Illinois Experiment							
Treatment (FSC-ineligible)	11 (0)	$500 (0)	$261 (4.8)	$121 (1.0)	18.56 (.25)	-.71* (.34)	$240 (6.1)
Control (FSC-ineligible)	--	--	250 (4.5)	122 (1.0)	19.27 (.23)	--	247 (5.7)
Pennsylvania Experiment							
Short qualification period/low bonus	6 (0)	500 (5.6)	279 (5.6)	167 (1.9)	15.54 (.28)	-.40 (.34)	252 (6.7)
Short qualification period/high bonus	6 (0)	1003 (9.4)	276 (4.6)	167 (1.6)	15.44 (.25)	-.50* (.31)	267 (7.0)
Long qualification period/low bonus	12 (0)	499 (4.3)	281 (4.4)	166 (1.4)	15.51 (.22)	-.43 (.28)	268 (5.4)
Long qualification period/high bonus	12 (0)	991 (7.6)	267 (4.4)	165 (1.3)	15.02 (.19)	-.93* (.27)	273 (3.7)
Control	--	--	272 (3.5)	164 (1.2)	15.94 (.18)	--	263 (4.7)
Washington Experiment							
Short qualification period/low bonus	5.7 (.02)	302 (2.2)	296 (4.3)	151 (1.1)	15.14 (.23)	-.07 (.30)	317 (6.9)
Short qualification period/medium bonus	5.8 (.02)	610 (4.3)	302 (4.3)	153 (1.1)	15.05 (.22)	-.16 (.30)	324 (6.7)
Short qualification period/high bonus	5.7 (.02)	917 (7.8)	298 (5.1)	153 (1.3)	14.61 (.27)	-.61* (.34)	315 (7.3)
Long qualification period/low bonus	11.0 (.04)	308 (2.1)	302 (4.1)	154 (1.1)	14.71 (.23)	-.50* (.30)	321 (6.6)
Long qualification period/medium bonus	11.0 (.04)	612 (4.3)	306 (4.5)	153 (1.1)	15.08 (.23)	-.14 (.30)	319 (6.5)
Long qualification period/high bonus	11.1 (.04)	923 (8.0)	307 (5.4)	154 (1.3)	14.50 (.28)	-.72* (.34)	326 (8.8)
Control	--	--	296 (3.6)	151 (0.9)	15.21 (.20)	--	321 (6.5)

Notes: Authors' tabulations of the Illinois, Pennsylvania, and Washington Reemployment Bonus Public Use Data files.

* denotes rejection of the hypothesis that the treatment effect is zero using a 10-percent significance level.

Table 2

Change in Expected Duration of Unemployment in Response to Changes in the UI Replacement Rate and Potential Duration of UI Benefits, Reference Case

Experimental Treatment	Response of Expected Duration of Unemployment to:			
	10 percentage point increase in the replacement rate		1 week increase in potential duration of UI benefits	
	ΔD (UI-eligible claimants)	ΔD_i (UI-ineligible)	ΔD (UI-eligible claimants)	ΔD_i (UI-ineligibles)
Illinois (FSC-ineligible)	.815	-.203	.194	-.051
Pennsylvania				
Short qualification period/low bonus	.777	-.211	.184	-.052
Short qualification period/high bonus	.481	-.121	.114	-.030
Long qualification period/low bonus	.532	-.136	.125	-.033
Long qualification period/high bonus	.627	-.125	.150	-.041
Washington*				
Short qualification period/high bonus	.653	-.177	.128	-.036
Long qualification period/low bonus	1.113	-.348	.219	-.071
Long qualification period/high bonus	.575	-.152	.111	-.031

*For other Washington treatments, see text.

Notes: ΔD is the change in expected duration of unemployment of UI-eligible claimants (in weeks); ΔD_i is the change expected duration unemployment of UI-Ineligible workers (in weeks).

Table 3

Variation in Response to Changes in the UI Replacement Rate and Potential Duration of UI Benefits: Results Based on Illinois Experiment

Initial Replacement Rate	Response to a 10 percentage point increase in the replacement rate		Initial Potential Duration of UI Benefits	Response to a 1 week increase in potential duration of UI benefits	
	ΔD (UI-eligible claimants)	ΔD_i (UI-ineligible)		ΔD (UI-eligible claimants)	ΔD_i (UI-ineligible)
.000	.761	-.195	16	.180	-.049
.100	.778	-.198	18	.185	-.050
.200	.791	-.201	20	.190	-.050
.300	.803	-.202	22	.193	-.051
.400	.811	-.203	24	.194	-.051
.500	.816	-.203	26	.194	-.051
.600	.816	-.201	28	.194	-.050
.700	.811	-.198	30	.192	-.050
.800	.801	-.194	32	.191	-.049
.900	.784	-.188	34	.186	-.049
			36	.183	-.048
			38	.179	-.047
			40	.175	-.046

Notes: See Table 2.

Table 4

Sensitivity of Results to Changes in the Separation Rate (s) and Number of Jobs Available (F)

	Response of Expected Duration o Unemployment to:			
	10 percentage point increase in replacement rate		1 week increase in potential duration of UI benefits	
	ΔD (UI-eligible claimants)	ΔD_i (UI-ineligibles)	ΔD (UI-eligible claimants)	ΔD_i (UI-ineligibles)
Illinois (FSC-ineligible)				
s low	.879	-.303	.211	-.076
F low	.819	-.229	.215	-.050
Reference Case	.815	-.203	.194	-.051
s high	.747	-.157	.182	-.040
F high	.817	-.225	.195	-.046
Pennsylvania (Long qualification period/high bonus)				
s low	.677	-.156	.177	-.057
F low	.629	-.141	.178	-.040
Reference Case	.627	-.125	.150	-.041
s high	.577	-.101	.144	-.036
F high	.627	-.143	.150	-.039
Washington (short qualification period/high bonus)				
s low	.679	-.209	.143	-.048
F low	.653	-.193	.128	-.035
Reference Case	.653	-.177	.128	-.036
s high	.615	-.159	.124	-.032
F high	.655	-.197	.128	-.035

Notes: See Table 2.

* U.S. G.P.O.:1996-403-281:40538 VOL.III