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UNEMPLOYMENT INSURANCE, RECALL EXPECTATIONS AND UNEMPLOYMENT OUTCOMES

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ABSTRACT

This paper shows the importance of explicitly accounting for the possibility of recalls when analyzing the determinants of unemployment spell durations and the effects of unemployment insurance (UI) on unemployment outcomes in the United States. These issues are examined using a unique sample of UI recipients from Missouri and Pennsylvania covering unemployment spells in the 1979-1981 period. We find that those expecting recall who are not recalled tend to have quite long unemployment spells. Furthermore, ex-ante temporary layoff spells (the spells of individuals' who initially expect to be recalled) may account for over 60 percent of the unemployment of UI recipients and appear to account for much more unemployment than ex-post temporary layoff spells (spells actually ending in recall). We estimate a competing risks model in which the finding of a new job and recall are treated as alternate routes of leaving unemployment. Our results using this approach show that the recall and new job exit probabilities have quite different time patterns and are often affected in opposite ways by explanatory variables. We also find that the probability of leaving unemployment (both through recalls and new job finding) increases greatly around the time that UI benefits lapse.

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Temporary layoffs, where workers are laid off when demand declines and often rehired by their original employers, are an important feature of the U.S. labor market. Feldstein (1975) and Lilien (1980) conclude from examinations of establishment data on turnover that over seventy percent of workers laid off in U.S. manufacturing in the 1970's were subsequently rehired by their former employers. The layoff-rehire process also appears to be widespread outside of manufacturing. Data from the Panel Study of Income Dynamics for 1980 and 1981 indicate that about fifty percent of heads of households laid off from nonmanufacturing jobs have unemployment spells ending in recall.[1] Not only do many unemployment spells end in recall, but a large fraction of the total weeks of unemployment accumulated by some labor force groups (e.g. unemployment insurance recipients, manufacturing workers, and prime-age males) occur in temporary layoff spells. For example, Feldstein finds that forty percent of all weeks of unemployment for men age 45-59 in the 1966-71 period were accumulated by individuals who did not change employers.[2]

This paper shows the importance of explicitly accounting for the possibility of recalls when analyzing the composition of unemployment, the determinants of unemployment spell durations, and the effects of unemployment insurance (UI) on unemployment outcomes in the United States. These issues are examined using a unique

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[1]This data set is described and analyzed in Katz (1986).

[2]Temporary layoffs are also a substantial component of unemployment in Canada. Robertson (1988) finds for Canada in 1984 that approximately 35 percent of total weeks of unemployment were accounted for by persons who eventually returned to their former employers.

sample of UI recipients from Missouri and Pennsylvania covering unemployment spells in the 1979-1981 period. This data set combines Continuous Wage and Benefit History (CWBH) UI administrative records with information from a follow-up survey conducted approximately one year after individuals filed for UI benefits. The CWBH questions, which were asked at the time an individual filed for UI, include whether or not each person expected to be recalled. The follow-up survey determined whether each unemployment spell ended in recall, ended in the finding of a new job, or was censored at the survey date. This information allows us to determine the relationship between recall expectations and unemployment experience for UI recipients in the two states.

The first focus of our empirical work is on the composition of unemployment of UI recipients. In particular, we attempt to determine the fraction of unemployment spells and fraction of time spent unemployed accounted for by the layoff-recall process. Most previous research has concluded that temporary layoffs account for a large fraction of unemployment spells, but a much smaller fraction of total time spent unemployed.[3] Two distinct measures of temporary layoffs have been used to determine the fraction of total

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[3]See Feldstein (1975) and Lilien (1980). On the other hand, Clark and Summers (1979) argue that temporary layoffs do not account for a large share of total unemployment in the United States. The difference in the conclusions appears to come from the emphasis of the first papers on job-losers, manufacturing employees, and prime-age males, all of whom (like UI recipients) are more likely to be involved in temporary layoffs. Murphy and Topel (1987) present evidence for 1968 to 1985 on the fraction of ongoing unemployment spells in the Current Population Survey that are classified as temporary layoffs.

unemployment time accounted for by the layoff-recall process. The first captures the proportion of unemployment from spells involving no job change, while the second looks at the fraction of the unemployed at a point in time who expect to be recalled. These measures are likely to underestimate the total amount of unemployment affected by recall prospects. The first measure does not include the unemployment of those who waited for recall but were not eventually recalled. The second measure only partially includes people who expect to be recalled, since recall expectations are likely to fade as an unemployment spell continues. We provide evidence that those <u>expecting</u> recall who are <u>not</u> recalled tend to have quite long unemployment spells. Furthermore, ex-ante temporary layoff spells (the spells of individuals' who initially expect to be recalled) may account for much more unemployment than ex-post temporary layoff spells (those actually ending in recall).

The second part of our empirical work shows the value of explicitly treating the possibility of recall when analyzing unemployment spell durations. Models that allow recalls naturally lead to a competing risks specification of the duration of unemployment spell in which the finding of a new job and recall are alternate routes of leaving unemployment (Katz, 1986). This specification differs from the single risk approach typically used in most studies of unemployment spell durations. Our findings using the competing risks approach shows that the recall and new job exit probabilities have quite different patterns and are often affected in opposite ways by explanatory variables. We also find that the probability of leaving unemployment (both through recalls and new job finding) increases greatly around the time that UI benefits lapse.

The remainder of the paper is organized as follows. Section I discusses several theoretical models of the recall process and the impact of the potential duration of unemployment benefits on unemployment outcomes. These models both guide the empirical work and aid the interpretation of the results. Section II describes the Missouri-Pennsylvania unemployment insurance recipients data set. Section III analyzes the composition of unemployment, the search behavior of the unemployed, the distribution of unemployment spell durations, and post-unemployment wages using the Missouri-Pennsylvania data set. Section IV uses econometric duration models to empirically determine the impact of recall expectations, demographic characteristics, and UI variables on unemployment spell durations and the likelihood of a spell ending in recall. Section V concludes.

I. Theoretical Background

Recall Prospects and Unemployment Spell Durations

The duration of unemployment is typically analyzed using a standard job search model in which unemployed workers generate job offers by costly search. This approach leads to a single risk model of unemployment spell durations in which unemployment spells can only end through the finding of an acceptable new job. This formulation is less appropriate when analyzing the unemployment durations of workers on layoff with some possibility of recall. The prospect of recall affects the probability of leaving unemployment directly through the rate of actual recalls and indirectly by affecting worker search behavior. Katz (1986) extends the standard McCall (1970) model of job search to include an exogenous probability of recall.[4] Katz finds that under reasonable conditions better recall prospects reduce the new job finding rate by raising the reservation wage and reducing the likelihood of search. This suggests that workers who expect to be recalled may have extremely long unemployment spells if their expectations are not fulfilled.

Katz (1985) also analyzes a model in which unemployed workers learn about their recall prospects in a Bayesian manner. He shows that the longer a worker is unemployed, all else held constant, the lower will be his or her subjective probability of recall. This result leads to a decreasing reservation wage and possibly increasing search intensity. Consequently, the new job finding rate for those who initially expect to be recalled should rise with unemployment duration (display positive duration dependence) under this scenario.

Furthermore, the statistical model of unemployment spell durations generated by the job search models extended to allow for recalls is a competing risks model in which unemployment spells can end either through recall or the finding of an acceptable new job.[5] The predictions of standard job search models for how variables

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[4]Burdett and Mortensen (1978) and Pissarides (1982) also analyze job search models that incorporate the possibility of recalls.

[5]See Kalbfleisch and Prentice (1980) for a detailed discussion of competing risks models.

affect the escape rate from unemployment really refer to the new job finding rate and these predictions need not hold for the overall escape rate from unemployment (the sum of the recall and new job finding rates). Information on whether spells ended through recall or the finding of a new job allows an econometrician to estimate a competing risks model. The competing risks specification has the advantage of permitting one to identify the distinct impact of variables on the recall rate and the new job finding rates.

The Duration of UI Benefits and the Duration of Unemployment Spells

The impact of finite duration UI on worker job search and employer recall behavior has been analyzed in several ways. Mortensen (1977) utilizes a standard dynamic search model with no recall possibility, a stationary known wage offer distribution, and a constant rate of job offers. As the remaining number of weeks of benefits decreases, the value of remaining unemployed also decreases. This drop causes search intensity to rise and the reservation wage to fall as an individual gets closer to when benefits lapse. These changes in behavior imply that the hazard rate (or escape rate) from unemployment rises till the date of benefits exhaustion and stays constant after the exhaustion date. On the other hand, if individuals can locate jobs and arrange not to begin work until their benefits run out, one could find a discrete increase in the escape rate near the point of benefits exhaustion.

An alternative approach is taken by Moffitt and Nicholson (1982) who use a static model where unemployed workers have preferences over income and unemployment. Some unemployment is valued because of its leisure component and because one can search while unemployed. At the time of job loss, individuals choose income and weeks of unemployment subject to a budget constraint. The budget constraint has a convex kink at the week of UI exhaustion because unemployment ceases to be subsidized. This kink implies that there may be a bunching of unemployment spell durations around the time benefits lapse.

Mortensen (1987) analyzes a joint wealth maximizing model of job separations with transitory demand changes facing firms and limited duration of unemployment benefits. The discrete change in the flow value of being unemployed when benefits are exhausted yields the prediction that many firms may recall laid-off workers around the benefit exhaustion point.

The theoretical models of unemployment spell durations surveyed indicate the importance of explicitly taking into account the recall process and the limited duration of unemployment benefits when analyzing the unemployment spells of UI recipients.

II. Data Description: The Missouri-Pennsylvania UI Recipients Sample

We use a unique data set to determine the fraction of unemployment of UI recipients due to ex-ante and ex-post layoffs and the relation of recall expectations to worker job search behavior. The data set consists of a sample of unemployment insurance recipients from Missouri and Pennsylvania beginning their unemployment insurance benefit years during the period October 1979 to March 1980.[6] The data set combines records collected by the Unemployment Insurance Service under the Continuous Wage and Benefit History (CWBH) system with information from special supplemental telephone interviews conducted in late 1980 and early 1981. The CWBH data include recall expectations, pre-UI weekly income, and demographic variables obtained from a survey administered when individuals filed for UI. Also included are administrative records on weekly UI benefits, the number of weeks of benefits for which an individual was eligible, and the timing and number of weeks collected. The follow-up telephone interviews ask when a job was found, the weekly wages on the job, whether the job was with the pre-UI employer, and additional information. The construction of the original data set is described by Corson and Nicholson (1983).[7]

The major advantage of this data set is that it provides information on whether individuals expected to be recalled at the time they filed for UI benefits, on how their initial unemployment spells ended, and on the level and length of unemployment insurance benefits available to each individual. Most previous work has used either CWBH data or survey data in isolation. Studies using only CWBH data (e.g. Moffitt, 1985; Meyer, 1988) miss the period prior to

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[6]In other words, the individuals in the sample all filed for unemployment insurance benefits during the October 1979 to March 1980 period. The initial unemployment spells for almost all of these individuals began during this time interval. The rare exceptions are those individuals with unemployment spells that began before October 1, 1979, but who did not file an unemployment insurance claim until after October 1, 1979.

[7]Corson and Hilton (1982) provide detailed documentation of the version of the data set we utilized to extract our sample.

filing for UI, the uncensored length of spells for those who exhaust benefits, and whether spells end through recall or the finding of a new job. Studies using micro survey data sets such as the Panel Study of Income Dynamics or National Longitudinal Survey (e.g. Dynarski and Sheffrin, 1987; Katz, 1986) tend to have poor information on the UI system parameters facing unemployed individuals and may have greater measurement error because of the retrospective nature of many of the questions. One disadvantage of the data set that we use is that it contains individuals from only two states over a short time period so that there is not a great deal of variation in the UI system parameters. A second disadvantage is that the data set does not allow us to examine the unemployment experiences of those who do not receive UI. This is an important issue because a large and increasing fraction of the unemployed in the U.S. are not UI recipients.[8]

The original Missouri-Pennsylvania telephone interview data set contains 2035 observations. Exclusions for missing demographic data and incomplete or inconsistent information on unemployment spells leaves a sample of 1499 observations.[9] Variable definitions and

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[8]Burtless (1983), Murphy and Topel (1987), and Kane (1988) document the decline in the fraction of the unemployed receiving UI in the U.S. and examine alternative explanations for this phenomenon.

[9]Observations were deleted for missing information on age, sex, marital status, education, the weekly UI benefit level, recall expectations, and on whether a definite recall date was given. Individuals who filed a UI claim but were disqualified before receiving benefits were deleted. 41 individuals in Missouri who received no weeks of full UI and apparently received partial UI while working part-time were also deleted. Finally, those individuals with missing UI claim dates, missing start dates of their initial

basic descriptive statistics for this sample are given in Table 1.

We focus most of our analysis on the initial spell of unemployment in the benefit year for each individual in the sample. Since the data set consists of a sampling of workers beginning unemployment insurance spells within a reasonably short interval in each state, the sample provides an approximation to a random sample drawn from the inflow of UI recipients into unemployment.[10] For an economy in a steady state, the distribution of first spells of unemployment of entrants into unemployment is the same as the distribution of the completed spells of a cross-section of the unemployed. This provides some justification for analyzing the characteristics of initial unemployment spells.[11] Our further justification is that the data set provides much better information on first spells then on total unemployment in the benefit year for

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unemployment spell, or irreconcilable inconsistencies among their claim dates, spell start dates, and spell end dates were deleted.

[10]While the entire sampling frame covers UI spells beginning from October 1979 to March 1980, the vast majority (95 percent) of the individuals from Missouri have benefit year begin dates in from November 1979 to January 1980 and the vast majority (92 percent) from Pennsylvania have benefit year begin dates from January to March 1980.

[11]One problem with the data set is that the steady state assumption is likely to be violated. First, since the sample includes individuals with spells starting in the fourth and first quarters, many seasonally unemployed workers are likely to be included. This is especially likely in Pennsylvania since most of the sample has unemployment spells beginning in the first quarter. The likely importance of seasonal unemployment means one must be somewhat cautious in drawing inferences concerning the distribution of unemployment over the full year from this sample. Second, fewer weeks of UI benefit eligibility, if any, are likely to be remaining for second or third unemployment spells in a benefit year. Thus, the incentives during the first unemployment spells examined here are different from those in a random sample of unemployment spells of UI recipients.

each individual.

We have developed several different measures of unemployment spell durations. IUSR measures the unemployment spell starting from the UI claim date which is available from administrative records, and FSPELL measures the spell from the respondent's self-reported spell start date. These two measures can be computed for both Pennsylvania and Missouri. PAYSPELL is an alternative measure that more fully utilizes administrative records on the actual number of weeks of benefits received, but it can be computed only for individuals from Missouri.[12] All three unemployment spell measures lead to similar conclusions concerning the fraction of unemployment due to either spells ending in recall or individuals who expected recall. The PAYSPELL measure provides much more accurate information for analyzing the distribution of unemployment spell durations.

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[12]PAYSPELL is defined as weeks from UI first payment date until reemployment (or until the interview date if the first unemployment spell is still in progress at the interview date). We utilized administrative records rather than respondent retrospective information whenever possible in constructing PAYSPELL. CWBH administrative information on the first payment date and on the weeks of benefits received in the benefit year were only available for Missouri. For individuals who had a single compensated unemployment spell in the benefit year and who gained reemployment before benefits were exhausted, PAYSPELL can be computed from CWBH administrative records and equals the weeks of benefits received in the benefit year. Since the available CWBH data does not disaggregate the total number of weeks of benefits received in a benefit year into individual spells of compensated unemployment, we were forced to use respondent retrospective information on weeks of benefits received for individuals with multiple compensated unemployment spells in the benefit year. In this case, PAYSPELL equals the survey respondent's self-reported weeks of benefits received during his or her initial unemployment spell. For individuals who exhausted their benefits during their initial unemployment spell, PAYSPELL is given by weeks from the UI first payment date (from CWBH records) until the self-reported reemployment date (or until the interview date if the spell is censored).

The descriptive statistics in Table 1 indicate that the prospect of recall was relevant for a large majority of the UI recipients in the sample. When asked soon after their unemployment spells began, seventy-five percent expected to be recalled and eighteen percent had a definite recall date from their employer.[13] Fifty-seven percent of the individuals in the sample had initial unemployment spells ending in recall. The mean unemployment spell duration is about 15 weeks when measured from the claim date and about 18 weeks when measured from the end of the pre-UI job.[14]

Unemployment insurance recipients in Missouri and Pennsylvania largely consisted of blue collar occupations and workers previously employed in construction and manufacturing. The importance of recalls varied substantially across industries. Sixty-six percent of the workers laid-off from construction, mining, and manufacturing had spells ending in recall as opposed to thirty-seven percent of the workers from transportation, trade, services and administration.

The unemployment spell durations are substantially longer on average for the Missouri sample. The mean spell length using IUSR is approximately 17 weeks for Missouri and approximately 13 weeks for Pennsylvania. The fraction of spells ending in recall is 64 percent

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[13] The recall expectations information arises from a claimant survey questionnaire which clearly indicates that the information is confidential and only for statistical and research purposes. The information is not utilized to determine claimant job search requirements or benefits eligibility.

[14] These means are underestimates of the true mean duration of completed spells since only incomplete spell durations are available for the 8 percent of the spells censored at the interview date.

in Pennsylvania and only 51 percent in Missouri.

The rules concerning the level and duration of UI benefits were much more generous in Pennsylvania than in Missouri during the period of our sample.[15] In particular, the maximum weekly benefit available was $105 in Missouri and $170 in Pennsylvania in 1980. In fact, the mean weekly benefit amount of $125 for individuals from Pennsylvania in our sample is greater than the maximum possible benefit in Missouri. Pre-UI earnings were similar in the two states leading to a much higher replacement rate in Pennsylvania. Regular UI benefits in Pennsylvania had a uniform duration of 30 weeks, while Missouri had a maximum potential duration of 26 weeks with variation in the potential duration that depended on base period and high quarter earnings. The Missouri sample provides substantial variation in the potential length of benefits, while the Pennsylvania sample provides almost none. Extended benefits were triggered in February 1980 in Pennsylvania and in May 1980 in Missouri. The extensions raised the potential length of benefits to 39 weeks in Pennsylvania and increased the potential length by fifty percent in Missouri.

The extent to which firms' UI payroll taxes depended on their previous layoff rates also differed greatly in the two states. A firm's future payroll taxes increased with layoffs until the firm's unemployment rate reached 6.3 percent in Missouri as compared to 3.6 percent in Pennsylvania. On the other hand, Topel (1985) calculates

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[15] Topel (1985) provides detailed information on the characteristics of the UI systems for Missouri, Pennsylvania and several other states circa 1980. Corson and Nicholson (1983) provide further information on the rules concerning the determination of UI benefit levels and potential durations in Missouri and Pennsylvania.

that the marginal effect of layoffs on future taxes was much greater in Pennsylvania than in Missouri when unemployment was below the maximum.

These UI system characteristics indicate that firms in highly cyclical or seasonal industries in Pennsylvania were unlikely to be experience-rated on the margin and that the replacement rate was much higher in Pennsylvania. These factors help explain why a larger fraction of UI recipients in Pennsylvania were involved in short layoff spells ending in recall to their original employer. 25 percent of recipients in Pennsylvania had definite recall dates and 64 percent were recalled. The corresponding figures for Missouri were 12 and 51 percent.

Accuracy of Survey Responses

The combination of administrative records and survey data available in the Missouri-Pennsylvania data set provides a unique opportunity to explore the accuracy of survey information on weekly benefit levels, weeks of benefit receipt, and unemployment spell durations. The data set allows us to compare accurate administrative records with the survey responses of the UI recipients. We find that the sample members provide quite accurate information on the level of UI benefits they received and quite poor information on the weeks of benefits received and the dates of the beginning and ending of their unemployment spells. 67.5 percent of the 1408 individuals in our sample that provided information on the level of weekly benefits reported exactly (to the dollar) the benefit level indicated by administrative records. 85 percent of the sample were within $10 of the true amount. The mean self-report was slightly downward biased ($102 reported vs. $105 actual) and the variance of the reporting error divided by the variance of the true value was a fairly small .26.

On the other hand, very few individuals in the sample reported weeks of benefit receipt the same as indicated by their CWBH records. Only 15 percent of the 561 individuals in Missouri with a single spell of unemployment in the benefit year reported weeks of benefit receipt equivalent to the number provided by administrative records, 35 percent have deviations from CWBH records of over 4 weeks.[16] The mean absolute difference between weeks reported by respondents and CWBH records is 4.5 weeks. Many inconsistencies in reported dates are apparent in the sample (e.g. reported end date of pre-UI job after UI claim date or UI first payment date available from administrative records). It appears that people may remember salient dollar amounts far better than the timing of events such as the start and end dates of unemployment spells.

III. Recall Expectations and Unemployment Outcomes: Some Evidence

In this section, we analyze the fraction of the unemployment of UI recipients in Missouri and Pennsylvania that can be accounted for

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[16] Consistent comparisons of self-reported weeks and administrative records can only be done for individuals from Missouri with a single spell of unemployment. Administrative records on weeks of benefit receipt are not available for Pennsylvania. Individuals with multiple spells from Missouri may include weeks of benefit receipt in latter spells outside the benefit year covered in our CWBH data set.

by ex-ante and ex-post layoffs, the relation between recall expectations and job search behavior, and the importance of taking into account the possibility of recall when analyzing the determinants of unemployment spell durations.

Recall Expectations, Job Search and the Composition of Unemployment

The usual method of assessing the contribution of temporary layoffs to unemployment uses an ex-post concept of temporary layoffs (Feldstein, 1975; Lilien, 1980). Ex-post temporary layoffs are unemployment spells that end through rehire to the original employer. This concept is appropriate if one is interested in the amount of unemployment that does not involve a change in employers.

This concept is not the correct one for assessing the contribution of the temporary layoff process to total unemployment. The ex-post approach does not take into account the fact that some workers who expect to be recalled at the time of layoff are not recalled or find other jobs before being recalled. Workers expecting recall, whose expectations are not met, may have quite long unemployment spells since they are unlikely to search intensively for a new job as long as they regard the probability of recall to a valuable old job as high. If these workers receive UI benefits, they may be willing to wait as long as the benefits last before searching for another job. Imperfectly experience-rated firms may have an incentive to encourage workers in whom they have invested to wait for recall. Other employers may be unwilling to incur the initial fixed costs of hiring and training workers with reasonable prospects of recall to a more attractive job. These factors suggest an ex- ante temporary layoff concept as the proper measure of the amount of unemployment affected by the layoff-recall process. We define ex-ante temporary layoff unemployment as the unemployment arising in spells in which the individual expected to be recalled at the time of layoff. The recall expectations information in our Missouri-Pennsylvania data set allows us to compare this unemployment concept with the usual ex-post temporary layoff approach.

Table 2 presents the distribution of first unemployment spells and weeks of first spell unemployment by spell outcome, recall expectations, and definite recall status for our entire sample using the IUSR unemployment concept. Since it is unlikely that many of the long censored spells ended in recall, it appears reasonable to conclude that about 57 percent of the unemployment spells and 32 percent of the weeks of unemployment of UI recipients in our two states are accounted for by ex-post temporary layoffs.[17] The typical spell ending in recall was substantially shorter than those ending in the finding of a new job. Less than 10 percent of unemployment is accounted for by spells in which individuals had a definite recall date. On the other hand, almost 64 percent of unemployment is accounted for by ex-ante temporary layoffs. Table 3 yields qualitatively similar findings for Missouri alone using the PAYSPELL unemployment spell measure which more fully uses available

[17] The share of unemployment accounted for by temporary layoffs is likely to be overstated in this sample relative to a random sample of unemployment spells over the calendar year since most of the spells started in the peak period for temporary seasonal layoffs (December, January, and February).

administrative information than the IUSR measure.

Table 4 provides more detailed information on the relation between recall expectations and unemployment outcomes. 72 percent of those who expected to be recalled and 13 percent of those who did not expect to be recalled had spells actually ending in recall. An interesting finding from this table is that <u>although the vast majority of those who expected to be recalled were recalled, more than fifty percent of the total unemployment of those who expected to be recalled is accounted for by the minority who were not recalled</u>. UI recipients who ex ante expected to be recalled and ex post were not recalled tend to have quite long unemployment spells. While this group accounts for only 21 percent of the entire sample, it accounts for approximately 34 percent of first spell unemployment. Since many of the individuals of this group had spells censored at the interview date, 34 percent may be an underestimate.[18]

One plausible reason why those who expect be recalled but are not tend to have long unemployment spells is that they may rationally decide to wait for recall and not search very intensively for a new job. (They may also have a difficult time gaining new jobs since employers will be reluctant to hire those likely to return to their old jobs.) Table 5 provides some information on the search behavior of the UI recipients in our sample. 59 percent of the UI recipients claimed to have looked for work at the time they were laid off. The average searcher spent 12 hours a week looking for work. Those that

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[18]On the other hand, many of the censored spells may involve individuals who have dropped out of the labor force.

expected to be recalled were substantially less likely to search than those who did not expect to be recalled and they searched many fewer hours on average as well. This result is consistent with the finding of Barron and Mellow (1979) that those who classify themselves as being on "temporary layoff" in the Current Population Survey spend less time searching than do other individuals who classify themselves as unemployed. Low search intensity may play a role in the low rate of new job finding of those who expect to be recalled.

One may be interested in the distribution of total weeks of unemployment in a benefit year rather than just first spell unemployment. If some groups have proportionally more unemployment in second and third spells, examining only the first spell would give a distorted picture of the distribution of unemployment. This bias would occur if past unemployment was either an inoculation against future unemployment or a cause for greater difficulty in finding and keeping a job. While the data set does not allow the construction of a good measure of total unemployment weeks in the benefit year, it does provide information on total weeks of compensated unemployment (weeks of UI benefit receipt) for the Missouri sample. This measure is also directly relevant for evaluating the fraction of UI benefits accruing to individuals involved in temporary layoffs.

The distribution of total compensated unemployment in the benefit year by outcome of the first spell and first spell recall expectations is presented in Table 6. Individuals whose first spell ended in recall account for almost 41 percent of the total weeks of compensated unemployment. This percentage is substantially larger than their share of total weeks of first spell unemployment. This difference arises because those recalled are more likely to have multiple spells of UI receipt in a year and because weeks of unemployment after UI exhaustion are not included. A reasonable conclusion from this table is that a large fraction (maybe 40 percent or more) of the weeks of compensated unemployment in Missouri in this period were accounted for by ex-post temporary layoff spells. This finding is quite similar to that of Robertson (1988) for Canada. Robertson finds that 44 percent of total UI weeks in Canada in 1984 were accounted for by ex-post temporary layoffs. Thus, a substantial proportion of insured unemployment in both the U.S. and Canada appears to be related to the layoff-recall process.

Reemployment Earnings

An important element in the evaluation of the success of a UI program is the effect of UI on the wages of reemployed workers. Table 7 provides information on the post-UI job earnings relative to pre-UI job earnings of those individuals in the Missouri-Pennsylvania sample reemployed by the interview date.[19] Those with unemployment spells ending in recall appear to go back to their old jobs since their post-UI hourly earnings are quite similar to their pre-UI

[19] Pre-UI earnings are from information provided by respondents at the time that they made their UI claims. Post-UI earnings are from the follow-up survey. The choice of deflator (Average Hourly Earnings vs. CPI) affects conclusions about the magnitude of earnings changes. The earnings losses are substantially larger when the CPI is used as the deflator. On the other hand, the choice of deflator does not substantively affect any conclusions concerning relative earnings changes of any of the groups compared.

hourly earnings. On the other hand, the usual weekly hours of those rehired by their previous employers do decline by about 4.5 percent on average. The reduced hours of those recalled suggest may relate to the cyclical downturn that gained force by the middle of 1980.

Individuals with spells ending through the finding of new jobs typically experienced substantial earnings declines. In particular, the hourly earnings of those who expected to be recalled but were not fell by 15 percent on average, while new job finders who did not expect to be recalled experienced 11 percent earnings losses on average. Table 7 also illustrates that individuals who exhausted their benefits experienced the largest earnings declines by a substantial margin. Their hourly earnings declined by 30 percent on average and their weekly earnings declined even further. The large losses of exhaustees suggest that reservation wages are likely to fall substantially and that the new job finding rate is likely to increase substantially as benefits run out. An alternative explanation for the low relative reemployment earnings of those with long spells is heterogeneity in reemployment prospects. Workers with low job offer arrival rates are likely to have both low reservation wages and low escape rates from unemployment for many plausible wage offer distributions (Mortensen, 1986).

The Distribution of Unemployment Spell Durations

The pattern of initial unemployment spell durations in our Missouri sample of UI recipients using the PAYSPELL unemployment spell concept is illustrated in Table 8 and Figures 1 and 2. We focus our duration analysis on the Missouri sample since more information to construct accurate spell durations is available for this sample than for Pennsylvania. Table 8 gives the Kaplan-Meier empirical hazards for the PAYSPELL data. The overall empirical hazard for a given week is the fraction of spells ongoing at the start of that week which end during the week.[20] The recall and new job empirical hazards are analogously defined as the fraction of spells ongoing at the start of the week which end during the week through recall and through the finding of a new job respectively. The total hazard basically trends downward except for a rise at 12 and 16 weeks and a valley at around 32 weeks. The overall hazard masks the quite distinct patterns in the recall and new job hazards. The recall hazard drops sharply over time except for spikes at 12 and 16 weeks and becomes quite low after about 25 weeks. The new job hazard starts out quite low and increases on average until about 28 weeks.[21] These basic differences in the recall and new job finding hazards are quite similar to those found for UI recipients in a national sample of household heads from the PSID analyzed by Katz (1986). The upward sloping new job hazard provides some support for the UI exhaustion effects emphasized by Mortensen (1977) and the

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[20]More formally, the Kaplan-Meier empirical hazard for week t (H_t), is the number of failures during the week (D_t), divided by the size of the risk set at the beginning of the week. The size of the risk set at the beginning of week t (R_t), is the number of people whose spells have not ended or been censored at the beginning of week t. Thus, $H_t = D_t/R_t$.

[21]A pronounced even-odd effect, where the hazard tends to be higher in even weeks, is also evident in Figures 1 and 2. A possible explanation for this anomaly is that the cards used to claim benefits in Missouri are mailed two at a time to potential recipients.

impact of changing recall expectations on job search behavior discussed by Katz. Direct evidence on exhaustion effects is somewhat masked in Figures 1 and 2 because of the fair amount of variation in potential durations contained in the Missouri sample.

Figures 3 and 4 provide a direct look at possible effects of finite length UI benefits on spell durations. The figures present time until exhaustion empirical hazards analogous to the usual Kaplan-Meier estimators. The time axis is time until benefits lapse rather than time since a spell began. The data behind these plots are reported in Table 9. There is a large spike in the hazard at the week of benefits exhaustion.[22] This spike is apparent for both the new job and recall hazards. The new job finding rate remains relatively high after exhaustion, while the recall rate becomes minuscule after exhaustion. This suggests that workers may stop waiting for recall and start taking new jobs as their benefits run out. In fact, when we look only at workers who indicated when their spells began that they expected to be recalled, the new job finding rate is extremely low early in spells and there is a prolonged sharp increase in the new job escape rate from four weeks before exhaustion through three weeks after exhaustion.

The recall spike around exhaustion in figure 4 provides some support for the Mortensen (1987) joint wealth maximizing model of the layoff-recall process. The model predicts many recalls occurring when

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[22]The spike in the hazard function at the week of benefits exhaustion is not primarily a phenomenon related to hiring halls and seasonal fluctuations in the construction industry. Only 4 of the 26 individuals with spells ending in the UI exhaustion week were construction workers.

the flow value of being unemployed drops discretely as benefits run out. In this case, recalls may make sense even if demand has not recovered. Additionally, a rotating system of layoffs and recalls may make sense when benefits are of limited duration.

The exhaustion spikes support the findings on a PSID sample of Katz (1986) and on a CWBH sample by Moffitt (1985) and Meyer (1988). Katz also found that spikes in the hazard near likely exhaustion points (26 and 39 weeks) were not apparent for non-UI recipients. The absence of similar spikes for non-UI recipients provides strong support for the view that the exhaustion spikes for UI recipients are strongly related to the finite length of UI benefits.

IV. Formal Duration Models for the Missouri Sample of UI Recipients

In this section, we analyze the impact of recall expectations, individual and pre-UI job characteristics, and UI system variables on the total, recall, and new job exit rates from unemployment for the Missouri sample of UI recipients.

Model Specification

The exit rates from unemployment are analyzed using formal hazard model techniques. Hazard models have several advantages over other techniques for analyzing unemployment spell data. Unemployment spells are positive random variables which are often censored (9.3 percent are censored in our sample). Many important explanatory variables (e.g. weeks until benefits exhaustion, local labor market conditions, etc.) change values during an unemployment spell. The entire time path of time-varying explanatory variables and the possible censoring of the dependent variable are easily incorporated in a hazard model.

We use a proportional hazards model estimator that allows for time-varying explanatory variables and which nonparametrically estimates the change in the hazard over time. This semiparametric approach is analyzed in detail in Meyer (1986). The estimates are the parameters of a continuous time hazard model and thus retain a clear interpretation. Nonparametrically estimating the change in the hazard over time eliminates the need to impose a potentially restrictive functional form that has no theoretical justification. If an incorrect functional form were assumed, all of the parameter estimates from the model would be inconsistent. This danger is avoided by nonparametrically estimating the baseline hazard.

Formally, we parameterize the overall hazard rate from unemployment for individual i at time t, $\lambda_i(t)$, using the proportional hazards form.

Let T_i be the length of individual i's unemployment spell. Then

$$\lambda_i(t) = \lim_{h \to 0^+} \frac{\text{prob}[t+h>T_i \geq t \mid T_i \geq t]}{h}$$

$$= \lambda_o(t)\exp\{z_i(t)'\beta\},$$

where

$\lambda_o(t)$ is the baseline hazard at time t, which is unknown,

$z_i(t)$ is a vector of time dependent explanatory variables for individual i, and

β is a vector of parameters which is unknown.

The probability of a spell lasting until t+1 given that is has lasted until t is easily written as a function of the hazard:

$$(1) \qquad P[T_i \geq t+1 \mid T_i \geq t] = \exp\left[- \int_t^{t+1} \lambda_i(u)du \right] .$$

Assuming that $z_i(t)$ is constant between t and t+1, equation (1) can be rewritten as

$$(2) \qquad P[T_i \geq t+1 \mid T_i \geq t] = \exp\left[- \exp(z_i(t)'\beta + \gamma(t)) \right]$$

where

$$(3) \qquad \gamma(t) = \ln\left(\int_t^{t+1} \lambda_o(u)du \right).$$

The log-likelihood for a sample of N individuals can be written as a function of terms such as (2)[23]:

$$(4)\ L(\gamma,\beta) = \sum_{i=1}^{N} \left(d_i \cdot \ln[1-\exp(-\exp[\gamma(k_i)+z_i(k_i)'\beta])] - \sum_{t=1}^{k_i-1} \exp[\gamma(t)+z_i(t)\beta] \right)$$

where k_i = the time a spell ends or is censored, and

d_i = 1 if the spell ends before the survey date and

0 if the spell is censored.

This approach assumes that censoring does not provide any information about T_i beyond that available in the covariates.

We utilize an analogous methodology to estimate the recall and new job hazards within a competing risks model framework. The recall and new job hazards are specified analogously to the total hazard above. In the estimation of the recall hazard, spells ending in the finding of a new job are treated as censored (d_i = 0) at the date of new job finding. Spells ending in recall are analogously treated as

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[23] See Meyer (1986) for a discussion of the derivation of the likelihood function and the properties of this estimator.

censored at the recall date in the estimation of the new job hazard.

The effects of unemployment insurance on the hazard rates are measured using functions of the benefits level and the time until benefits lapse. The level of weekly UI benefits is included as a time varying covariate whose impact is allowed to vary depending on whether the individual is still receiving benefits or has exhausted benefits. Also included are time until benefit exhaustion dummy variables for five intervals covering both weeks before and after benefits have expired. These variables are designated UI 6-10 through UI ≤-1. Each of these time-varying exhaustion dummies takes on the value of 1 in its designated interval and takes on the value of 0 in all other periods. For example, UI 6-10 takes on the value 1 when the individual is 6 to 10 weeks until exhaustion, UI 0 takes on the value of 1 in the week of benefits exhaustion, and UI ≤-1 takes on the value of 1 when the individual is one week or more after exhaustion. Those 11 or more weeks before exhaustion are the comparison group, the group corresponding to the omitted dummy variable.

Results for the Missouri UI Recipient Sample

Semiparametric hazard model estimates of the total, recall and new job hazards for the Missouri sample using the PAYSPELL unemployment spell variable are presented in Table 10.[24] Initial recall expectations have a strong effect on the hazards, raising the

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[24]The sample size falls to 756 in the hazard model estimates since 52 individuals in the original Missouri sample have missing pre-UI job tenure data.

recall and reducing the new job hazards substantially. Using the estimates in Table 10, those expecting recall have a recall hazard that is almost ten times as high as those who do not expect to be recalled. Furthermore, those expecting recall have a new job hazard which is almost forty percent lower. The large negative coefficient on expect recall in the new job hazard indicates that workers who expect to be recalled and are not, tend to have much longer unemployment spells than observationally equivalent workers who realized they were permanently displaced at the time of layoff. This result is consistent with the finding of Katz (1986) that individuals with unemployment spells initiated by plant closings have higher new job finding rates than those with unemployment spells initiated by layoffs. The expect recall and definite recall variables also have strong effects on the total hazard in the estimates presented in Table 10. Those that have a definite recall date (and necessarily expect recall) have a total hazard which is over twice as high as those not expecting recall. A definite recall date also increases further the recall hazard by a factor of 1.7, but has no significant effect on the new job hazard.[25]

Increases in pre-UI job tenure, a possible measure of firm

[25]The industry dummy variable coefficients are fairly small and statistically insignificant when expect recall and definite recall date are included in the hazard model estimates. The recall rate is substantially higher in nondurable goods and the new job finding rate is substantially lower in durable goods than in other industries in the specifications presented in Table 10. When the expect recall and definite recall date dummies are excluded, the industry dummy variables have much larger and statistically significant effects with construction, durable goods, and nondurable goods industries having significantly higher recall rates and significantly lower new job finding rates than other industries.

specific human capital or job match quality, is associated with a significantly increased recall hazard and decreased new job hazard. Older workers appear to have longer spells because of both lower recall and new job finding rates after controlling for tenure. The total hazard estimates mask many large differences between the effects of the covariates on the recall and new job finding hazards.

The large and significant increases in the recall and new job hazards apparent in Figure 4 at the week of benefits exhaustion are strongly confirmed in the more sophisticated hazard model estimates. Higher UI benefits are associated with higher recall rates and lower new job finding rates. The UI benefit coefficients in the new job hazard appear reasonable; higher benefits greatly depress the new job finding rate, and this effect disappears after benefits are exhausted. The positive and significant coefficient in the recall hazard is a puzzle. High UI benefits may be linked to the short-term temporary layoff sector of the Missouri economy. The effect of UI and the pre-UI wage on the total hazard are of opposite sign from the findings of most studies, although they are not statistically significant. The odd UI benefit coefficient estimates may arise because the variation in benefits in Missouri has a peculiar form: over 71 percent of the sample received either exactly \$105 or exactly \$85 in benefits. Variation in benefits across states and points in time is usually available in studies finding that UI benefits increase unemployment spell durations (e.g. Ehrenberg and Oaxaca, 1976; Meyer, 1988).

We further examine the time pattern of the baseline hazards from these models now that we have controlled for observable differences across individuals. After including explanatory variables, the time pattern of the hazards is captured by the $\gamma(t)$'s, the baseline hazard parameters defined in equation (3). These parameters confirm the patterns seen in Figures 1 and 2. A total hazard which falls with unemployment duration masks the combination of an upward sloping new job hazard and a downward sloping recall hazard. A test of these patterns which confirms the visual impression was performed using GLS regressions of the baseline hazard parameters on the length of spell. As a summary of the data, we used the specification $\gamma(t) = a + b \cdot \ln(t) + \epsilon$. This specification roughly corresponds to a Weibull baseline hazard. These regressions yield a positive coefficient on ln(t) for the new job hazard, and a negative coefficient on ln(t) for the total hazard and the recall hazard, all of which are significant at the 5 percent level. These results clearly show the value of the competing risks specification which allows the disentangling of the two effects which produce the total hazard. Furthermore, the finding that the new job escape rate rises with spell duration, even after controlling for the remaining potential duration of UI benefits, suggests that falling reservation wages from declining assets and changing recall expectations may play an important role in the reemployment process of laid-off workers.[26]

[26]Although uncontrolled heterogeneity biases estimates of the overall hazard towards spurious findings of negative duration dependence, a bias in the opposite direction is possible for an individual escape route hazard in a competing risks framework. If uncontrolled factors that raise the recall hazard also lower the new job hazard, then one can in theory find spurious positive duration dependence in the new job hazard. Han and Hausman (1986) have

A potential problem with the estimates in Table 10 is that it is likely that some individual attributes which affect the hazard rate are omitted from the list of covariates. If unobserved heterogeneity is present, but not allowed for in the estimation, the coefficient estimates will be biased (Lancaster 1979, 1985). Table 11 reports estimates which allow for individual specific omitted attributes under the assumption that a gamma distribution is a reasonable approximation to the distribution of heterogeneity in the population.[27] The estimates in Table 11 are very similar to those in Table 10 except for the rescaling effect suggested by Lancaster (1985). Lancaster finds that the omission of heterogeneity biases parameter estimates towards zero, even though elasticities of mean duration with respect to covariates may not change.

Specifications were also tried which included several additional covariates: a dummy variable set equal to 1 if the individual engaged in job search at the time of job loss, the time-varying state unemployment rate, and five occupation dummy variables. None of these additions noticeably changed the key findings. The state

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developed an estimator to handle correlated, unobserved heterogeneity in a competing risks model. They implement their estimator on the PSID layoff unemployment spell data set developed by Katz (1986) and find essentially zero correlation among the unobserved heterogeneity factors in the new job and recall hazards.

[27] Whether or not the gamma distribution is sufficiently flexible is a subject of debate. Heckman and Singer (1984) argue that it is not, but their results come from an example where a fairly restrictive parametric form for the baseline hazard is assumed. Ridder and Verbakel (1983) offer some evidence that the gamma distribution does fairly well. Some preliminary Monte Carlo experiments by one of the current authors indicates that coefficient estimates are relatively insensitive to the distribution of heterogeneity when the baseline hazard is estimated nonparametrically.

unemployment rate and occupation dummies were always insignificant. The behavior of the search variable again illustrates the usefulness of the competing risks approach. In the total hazard the search variable comes in negative and highly significant, implying that those who search initially are reemployed less quickly. However, this may arise because initial search acts as a further proxy for the likelihood of recall. Those who strongly expect to be recalled may not search and may also be recalled quickly. The recall and new job hazard estimates provide some support for this interpretation. The search variable has a large negative value in the recall hazard, but is small and insignificant in the new job hazard.

Overall, the lack of variation in the UI parameters within Missouri suggests the need to look at a data set covering more states and a longer time period to more accurately determine the impact of the length and level of UI benefits on spell durations.[28] The results with the Missouri sample do indicate that the recall process plays a major part in determining the duration of unemployment spells of UI recipients and the increase in unemployment escape rate around when benefits lapse.

We also examined hazard model estimates for the Pennsylvania subsample, though we were less confident about the accuracy of the spell lengths since their derivation relied more heavily on survey responses rather than administrative records. The benefit level and

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[28]We are currently beginning an analysis of a large data set that is better suited to the estimation of UI effects. The data set covers 9 states over a six year period and includes most of the variables available in the Missouri-Pennsylvania data set plus other information.

pre-UI wage coefficients in the total hazard had the signs found in previous studies (e.g. Classen, 1979), but they were not significantly different from zero. The signs and significance of the expect recall and definite recall date variables were very similar to those found for Missouri. The definite recall date coefficients tended to be larger than for Missouri. This is not surprising given the greater use of definite recall dates in Pennsylvania as seen in Table 1. Definite recall had the expected negative sign in the new job hazard and the effects of tenure on the previous job were similar in the two states.

Finally, several authors, including Hamermesh (1977) have suggested a subtle reason why studies which use weeks compensated by UI as the dependent variable might yield biased benefit coefficients. They suggest that higher benefits might induce people to claim UI more promptly, so that a larger fraction of an unemployment spell of a given length would be spent receiving UI. This effect might lead to the finding that higher benefits cause longer compensated spells even when there is no effect on the total length of unemployment. This effect is of potential importance in our Missouri sample where the mean number of weeks from loss of job till UI claim is 3.6 weeks and the standard deviation is 4.3 weeks. This hypothesis was tested by estimating hazard models where the dependent variable is the time from loss of job to the UI claim date. We used a set of control variables like that used for the unemployment spell specifications. The Hamermesh hypothesis would require a large positive coefficient on the benefit level, but the estimated coefficient was close to zero, negative and insignificant. This result provides some support for the reliability of studies which use weeks compensated as the dependent variable.[29]

V. Conclusion

This paper has examined the extent to which the unemployment of UI recipients in the U.S. can be attributed to alternative concepts of temporary layoffs and the impact of the potential duration of UI benefits on the distribution of unemployment spell durations of UI recipients.

We find that an understanding of the layoff-rehire process is critical to understanding the composition of unemployment (particularly insured unemployment) in the United States. Over 30 percent of the total weeks of unemployment of UI recipients in Missouri and Pennsylvania were attributable to unemployment spells ending in recall. Ex-ante temporary layoffs (those in which the individual initially expected to be recalled) may account for over 60 percent of the unemployment of UI recipients. Individuals who initially expect to be recalled search less intensely for new jobs than other UI recipients and tend to have extremely long unemployment spells if they are not actually rehired by their original employer. The recall rate is quite high at short spell durations and right around the point at which UI benefits lapse.

The potential duration of UI benefits appears to have a

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[29] Solon (1981) found a similar result in an examination of CWBH data for 3 states during the 1978-79 period.

substantial impact on the length of the unemployment spells of U.S. UI recipients. Our findings indicate that the probability of leaving unemployment (both through recalls and new job finding) increases greatly around the time that benefits are exhausted. Furthermore, some rough simulations based on hazard model estimates for twelve states in Katz and Meyer (1988) indicate that an increase in the potential duration of benefits of the size that naturally occurs when a state passes through an extended benefits trigger increases the mean weeks of compensated unemployment in a benefit year by 15 percent (2.6 weeks). This is a large effect given that most spells are completed well before regular benefits run out. In fact, the impact of the extended benefits trigger on the duration of UI recipient unemployment spells is estimated to be almost identical to the impact of a uniform 20 percent increase in the level of benefits. Alteratively, Moffitt and Nicholson (1982) using a labor supply estimation framework find that a one-week increase in the potential duration of benefits increases the length of an average spell by 0.1 week.

These findings suggests that a further examination of the impact of the potential length of UI benefits on unemployment could be quite useful. Most work on the effects of UI focuses on differences in replacement rates or experience rating provisions. Rules concerning potential benefit durations vary greatly across OECD countries with the typical potential duration ranging from 26 weeks in the United States, to two and a half years in Denmark, and to a virtually unlimited duration in Belgium (Emerson, 1988). Shorter maximum durations of UI and the much greater importance of the layoff-recall process may play a major role in the lower incidence of extremely long-term unemployment in the U.S. and Canada than in most European countries.

References

Barron, John M. and Wesley Mellow. "Search Efforts in the Labor Market." Journal of Human Resources 14 (1979): 389-404.

Burdett, Kenneth and Dale Mortensen. "Labor Supply Under Uncertainty." Research in Labor Economics 2 (1978): 109-57.

Burtless, Gary. "Why Is the Insured Unemployment Rate So Low?." Brookings Papers on Economic Activity (1983:1).

Classen, Kathleen P. "Unemployment Insurance and Job Search." In S.A. Lippmann and J.J. McCall, eds., Studies in the Economics of Search. Amsterdam: North Holland, 1970.

Clark, Kim B. and Lawrence H. Summers. "Labor Market Dynamics and Unemployment: A Reconsideration." Brookings Papers on Economic Activity (1979:1): 13-60.

Corson, Walter and Susan Hilton. "Unemployment Spells Public Use File Documentation," Mimeo, Mathematica Policy Research, February 1982.

Corson, Walter and Walter Nicholson. "An Analysis of UI Recipients' Unemployment Spells," Unemployment Insurance Occasional Paper 83-1, U.S. Department of Labor, Employment and Training Administration, 1983.

Dynarski, M. and S. Sheffrin. "New Evidence on the Cyclical Behavior of Unemployment Durations." in K. Lang and J. Leonard, eds., Unemployment and the Structure of Labor Markets. London: Basil Blackwell, 1987.

Ehrenberg, Ronald G. and Ronald L. Oaxaca. "Unemployment Insurance, Duration of Unemployment, and Subsequent Wage Growth." American Economic Review 66 (1976): 754-66.

Emerson, Michael. What Model for Europe?. Cambridge, MA.: MIT Press, 1988.

Feldstein, Martin S. "The Importance of Temporary Layoffs: An Empirical Analysis." Brookings Papers on Economic Activity (1975:3): 725-44.

Hamermesh, Daniel. Jobless Pay and the Economy. Baltimore: John Hopkins Press, 1977.

Han, Aaron and Jerry A. Hausman. "Semi-Parametric Estimation of Competing Risks Models." MIT Working Paper, November 1986.

Heckman, James J. and Burton Singer. "A Method for Minimizing the Impact of Distributional Assumptions in Econometric Models for

Duration Data." Econometrica 52 (1984): 271-320.

Kalbfleisch, John D. and Ross Prentice. The Statistical Analysis of Failure Time Data. New York: Wiley, 1980.

Kane, Thomas J. "What Happened to Unemployment Insurance? Administrative Reforms and Compositional Changes." Mimeo, Harvard University, Kennedy School of Government, February 1988.

Katz, Lawrence F. "Worker Mobility and Unemployment." Unpublished Ph.D. Dissertation, MIT, 1985.

Katz, Lawrence F. "Layoffs, Recall and the Duration of Unemployment," NBER Working Paper No. 1825, January 1986.

Katz, Lawrence F. and Bruce D. Meyer. "The Timing of Unemployment Insurance Payments: Theory and Evidence." Mimeo, May 1988.

Lancaster, Tony. "Econometric Models for the Duration of Unemployment." Econometrica 47 (1979): 939-56.

Lancaster, Tony. "Generalised Residuals and Heterogeneous Durations Models with Applications to the Weibull Model." Journal of Econometrics 28 (1985): 155-169.

Lilien, David M. "The Cyclical Importance of Temporary Layoffs." Review of Economics and Statistics 62 (1980): 24-31.

McCall, J.J. "Economics of Information and Job Search." Quarterly Journal of Economics 84 (1970): 113-26.

Meyer, Bruce D. "Semiparametric Estimation of Hazard Models." Mimeo, MIT, November 1986.

Meyer, Bruce D. "Unemployment Insurance and Unemployment Spells," NBER Working Paper No. 2546, March 1988.

Moffitt, Robert. "Unemployment Insurance and the Distribution of Unemployment Spells." Journal of Econometrics 28 (1985): 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 (1982): 1-11.

Mortensen, Dale T. "Unemployment Insurance and Job Search Outcomes." Industrial and Labor Relations Review 30 (1977): 595-512.

Mortensen, Dale T. "Job Search and Labor Market Analysis." In O. Ashenfelter and R. Layard, eds., Handbook of Labor Economics, vol. 2. New York: North Holland, 1986.

Mortensen, Dale T. "A Structural Model of UI Benefit Effects on the Incidence and Duration of Unemployment." Mimeo, Northwestern University, June 1987.

Murphy, Kevin M. and Robert Topel. "The Evolution of Unemployment in the United States: 1968-1985." NBER Macroeconomics Annual 2 (1987).

Pissarides, Christopher A. "Job Search and the Duration of Layoff Unemployment." Quarterly Journal of Economics 97 (1982): 595-612.

Ridder, Geert and Wim Verbakel. "On the Estimation of the Proportional Hazard Model in the Presence of Unobserved Heterogeneity." Mimeo, University of Amsterdam, 1983.

Robertson, Matthew. "Temporary Layoffs and the Measurement of Unemployment: An Empirical Analysis for Canada." Mimeo, Employment and Immigration Canada, Ottawa, 1988.

Solon, Gary. "Unemployment Insurance, Filing Delay, and Unemployment Duration." Princeton University Industrial Relations Section Working Paper #144, October 1981.

Topel, Robert. "Unemployment and Unemployment Insurance." Research in Labor Economics 7 (1985): pp. 91-135.

Table 1: Descriptive Statistics for UI Recipients' Data Set

Missouri and Pennsylvania
Unemployment Spell Start Dates in 1979-80

Variable	Description	Mean (S.D.) MO	PA	Total
IUSR	weeks from UI claim date until reemployment or until interview date if spell is censored	16.64 (15.62)	12.91 (14.47)	14.92 (15.15)
FSPELL	weeks from end of pre-UI job until reemployment or until interview date if spell is censored	19.35 (16.66)	16.21 (16.54)	17.90 (16.67)
PAYSPELL	weeks from UI first payment date until reemployment or until interview date if spell is censored (Missouri only)	15.27 (14.81)	-	-
PD1	potential benefits duration in weeks at claim date	22.92 (4.52)	34.88 (4.49)	28.44 (7.47)
UI benefit	augmented weekly benefit amount	88.80 (17.64)	124.76 (42.38)	105.38 (36.28)
Pre-UI wage	usual weekly earnings on pre-UI job	258.35 (133.12)	256.13 (122.97)	257.33 (128.50)
EXPREC	= 1 if expect recall at time of claim	.74	.76	.75
DEFREC	= 1 if have definite recall date	.12	.25	.18
Recall	= 1 if spell ended in recall	.51	.64	.57
New Job	= 1 if spell ended in taking a new job	.40	.28	.34
Censored	= 1 if spell is censored at interview date	.09	.07	.08
Age	age in years	36.43 (13.19)	36.80 (13.59)	36.60 (13.37)
Female	= 1 if female	.33	.25	.30
Married	= 1 if married	.69	.63	.66
Education	years of schooling	11.37 (2.11)	11.56 (1.76)	11.46 (1.95)
Spwk	= 1 if spouse works	.45	.37	.41
PA	= 1 if Pennsylvania	.00	1.00	.46

Table 1: continued

Variable	Description	Mean (S.D.) MO	PA	Total
Industry Dummies				
Mining	= 1 if mining	.01	.03	.02
Construct	= 1 if construction	.30	.28	.29
Durables	= 1 if durable goods manufacturing	.21	.24	.22
Nondurables	= 1 if nondurable goods manufacturing	.16	.17	.16
Transport	= 1 if transportation, communications or utilities	.06	.04	.05
Trade	= 1 if wholesale or retail trade	.12	.13	.12
Admin	= 1 if public administration	.03	.03	.03
Service	= 1 if services	.11	.08	.10
Occupation Dummies				
Prof	= 1 if professional, technical, or managerial	.06	.05	.05
Clerical	= 1 if clerical or sales	.10	.09	.10
Supervisor	= 1 if supervisor	.06	.04	.05
Craft	= 1 if craft and related occupations	.34	.38	.36
Operator	= 1 if operator	.23	.29	.26
Laborer	= 1 if laborer	.21	.15	.18
Sample size		808	691	1499

Table 2: Characteristics of First Spells of Unemployment
Entire Sample: Missouri and Pennsylvania

Unemployment Measure = IUSR
n=1499

	Percentage of Spells	Percentage of Total Weeks of Unemployment	Mean Duration in Weeks
Spell Outcome			
Recall	57.2	32.4	8.4
New Job	34.4	39.1	17.0
Censored	8.4	28.5	50.6
Recall Expectations			
Expect Recall	75.2	63.8	12.7
Don't Expect Recall	24.8	36.2	21.8
Definite Recall			
Definite Recall Date	18.1	9.7	8.0
No Definite Recall Date	81.9	90.3	16.5

The length of the unemployment spell up to the interview date is utilized as the unemployment spell duration for censored spells in the percentage of unemployment and mean duration in weeks calculations.

Table 3: Characteristics of First Spells of Unemployment
Missouri Only

Unemployment Measure - PAYSPELL
n-808

	Percentage of Spells	Percentage of Total Weeks of Unemployment	Mean Duration in Weeks
Spell Outcome			
Recall	51.1	30.2	9.0
New Job	39.6	39.7	15.3
Censored	9.3	30.1	49.6
Recall Expectations			
Expect Recall	74.4	65.7	13.5
Don't Expect Recall	25.6	34.3	20.4
Definite Recall			
Definite Recall Date	12.0	7.1	9.1
No Definite Recall Date	88.0	92.9	16.1

The length of the unemployment spell up to the interview date is utilized as the unemployment spell duration for censored spells in the percentage of unemployment and mean duration in weeks calculations.

Table 4: Recall Expectations and Unemployment Outcomes
Entire Sample: Missouri and Pennsylvania

Unemployment Measure = IUSR
n=1499

Expect Recall

Spell Outcome	Percentage of Spells	Percentage of Total Weeks of Unemployment	Mean Duration in Weeks
Recall	71.7	46.4	8.2
New Job	22.2	29.0	16.5
Censored	6.1	24.6	50.8

n=1127

Don't Expect Recall

Spell Outcome	Percentage of Spells	Percentage of Total Weeks of Unemployment	Mean Duration in Weeks
Recall	13.4	7.6	12.3
New Job	71.2	57.0	17.4
Censored	15.3	35.4	50.4

n=372

Table 5: Search Behavior of UI Recipients
Entire Sample: Missouri and Pennsylvania
n=1499

	Percent Who Searched	Mean Search Hours Per Week of Those Who Searched	Unconditional Mean Search Hours Per Week
Entire Sample	59	12.1	7.1
Outcome			
Recall	41	9.8	4.0
New Job	85	14.3	12.1
Censored	78	11.3	8.8
Recall Expectations			
Expect Recall	52	10.9	5.7
Don't Expect Recall	83	14.5	12.0
Definite Recall			
Definite Recall Date	33	11.7	3.8
No Definite Recall Date	65	12.2	7.9

The percent who searched calculations are based on the yes-no answers of workers to the following question: "I'd like to ask you about the period of time after that job [pre-UI job] ended. Did you look for work at that time?" Workers who answered yes to this question were later asked "And about how many hours per week on the average would you say you spent looking for work?"

Table 6: Distribution of Total Compensated Unemployment in Benefit Year

Missouri Only
n=808

	Percentage of Individuals	Mean Weeks of Compensated Unemployment	% of Total Compensated Unemployment
Outcome of First Spell			
Recall	51.1	11.3	40.7
New Job	39.6	15.3	42.9
Censored	9.3	25.1	16.4
First Spell Recall Expectations			
Expect Recall	74.4	13.2	69.5
Don't Expect Recall	25.6	16.9	30.5

Entire Sample

Mean Weeks of Compensated Unemployment = 14.2

S.D. of Weeks of Compensated Unemployment = 9.9

Table 7: Post-UI Job Earnings Relative to Pre-UI Job Earnings for those Reemployed by the Interview Date

Earnings Change Measure = Log(Post-UI Earnings / Pre-UI Earnings)
Missouri and Pennsylvania

Entire Sample

	Sample Size	Change in Log Weekly Earnings: Mean	Change in Log Weekly Earnings: Median	Change in Log Hourly Earnings: Mean	Change in Log Hourly Earnings: Median
Spell Outcome					
Recall	838	-.059 (.011)	-.046	-.014 (.009)	-.023
New Job	493	-.156 (.023)	-.103	-.128 (.019)	-.089

New Job Finders

	Sample Size	Change in Log Weekly Earnings: Mean	Change in Log Weekly Earnings: Median	Change in Log Hourly Earnings: Mean	Change in Log Hourly Earnings: Median
Recall Expectations					
Expect Recall	240	-.201 (.034)	-.141	-.151 (.028)	-.104
Don't Expect Recall	253	-.113 (.031)	-.081	-.106 (.027)	-.081
Whether Exhausted					
Exhausted Benefits	67	-.520 (.078)	-.425	-.301 (.058)	-.246
Didn't Exhaust Benefits	426	-.098 (.023)	-.086	-.101 (.020)	-.085

The numbers in parentheses are the standard errors of the means. Earnings are deflated by average hourly earnings of U.S. private nonagricultural workers (series AHEEAP from DRI). The base period for the deflator is the second quarter of 1979. Pre-UI job earnings are deflated from the end date of the pre-UI job. Post-UI job earnings are deflated from the interview date.

Table 8: Empirical Hazards for Missouri Sample
Using PAYSPELL Unemployment Spell Concept

Weeks Unemployed	Risk Set	Number of Spells That End: Total	Recall	New Job	Empirical Hazard: Total	Recall	New Job
1	756	75	59	16	0.0992	0.0780	0.0212
2	681	48	36	12	0.0705	0.0529	0.0176
3	633	46	27	19	0.0727	0.0427	0.0300
4	587	36	23	13	0.0613	0.0392	0.0221
5	551	24	15	9	0.0436	0.0272	0.0163
6	527	35	23	12	0.0664	0.0436	0.0228
7	492	29	14	15	0.0589	0.0285	0.0305
8	463	40	24	16	0.0864	0.0518	0.0346
9	423	16	12	4	0.0378	0.0284	0.0095
10	407	32	14	18	0.0786	0.0344	0.0442
11	375	19	9	10	0.0507	0.0240	0.0267
12	356	45	28	17	0.1264	0.0787	0.0478
13	311	20	12	8	0.0643	0.0386	0.0257
14	291	16	11	5	0.0550	0.0378	0.0172
15	275	22	14	8	0.0800	0.0509	0.0291
16	253	27	16	11	0.1067	0.0632	0.0435
17	226	12	6	6	0.0531	0.0265	0.0265
18	214	15	8	7	0.0701	0.0374	0.0327
19	199	9	4	5	0.0452	0.0201	0.0251
20	190	14	6	8	0.0737	0.0316	0.0421
21	176	3	1	2	0.0170	0.0057	0.0114
22	173	12	2	10	0.0694	0.0116	0.0578
23	161	7	3	4	0.0435	0.0186	0.0248
24	154	11	6	5	0.0714	0.0390	0.0325
25	143	5	1	4	0.0350	0.0070	0.0280
26	138	8	3	5	0.0580	0.0217	0.0362
27	130	4	0	4	0.0308	0.0000	0.0308
28	126	5	3	2	0.0397	0.0238	0.0159
29	121	4	1	3	0.0331	0.0083	0.0248
30	117	2	0	2	0.0171	0.0000	0.0171
31	114	1	0	1	0.0088	0.0000	0.0088
32	113	1	0	1	0.0088	0.0000	0.0088
33	112	2	0	2	0.0179	0.0000	0.0179
34	110	3	0	3	0.0273	0.0000	0.0273
35	107	5	2	3	0.0467	0.0187	0.0280
36	102	4	1	3	0.0392	0.0098	0.0294
37	98	6	1	5	0.0612	0.0102	0.0510
38	92	3	1	2	0.0326	0.0109	0.0217
39	89	4	0	4	0.0449	0.0000	0.0449
40	85	2	1	1	0.0235	0.0118	0.0118

Table 9: Empirical Time Until Exhaustion Hazards for Missouri Sample Using PAYSPELL Unemployment Spell Concept

Wks Until Exhaustion	Risk Set	Number of Spells That End: Total	Recall	New Job	Empirical Hazard: Total	Recall	New Job
25	415	41	33	8	0.0987	0.0795	0.0192
24	404	30	25	5	0.0743	0.0618	0.0123
23	403	30	18	12	0.0744	0.0447	0.0298
22	404	26	20	6	0.0644	0.0495	0.0149
21	410	26	19	7	0.0634	0.0463	0.0170
20	422	32	23	9	0.0758	0.0545	0.0213
19	425	25	15	10	0.0588	0.0352	0.0235
18	417	34	18	16	0.0815	0.0432	0.0384
17	412	23	14	9	0.0558	0.0340	0.0218
16	417	21	12	9	0.0504	0.0288	0.0216
15	422	19	10	9	0.0450	0.0237	0.0213
14	431	40	24	16	0.0928	0.0557	0.0371
13	414	31	14	17	0.0749	0.0338	0.0411
12	409	21	14	7	0.0513	0.0342	0.0171
11	408	25	13	12	0.0613	0.0319	0.0294
10	398	24	9	15	0.0603	0.0226	0.0377
9	374	21	13	8	0.0561	0.0348	0.0213
8	350	27	14	13	0.0771	0.0400	0.0371
7	320	8	4	4	0.0250	0.0125	0.0125
6	304	12	7	5	0.0395	0.0230	0.0164
5	291	13	3	10	0.0446	0.0103	0.0344
4	268	16	7	9	0.0597	0.0261	0.0336
3	254	11	6	5	0.0433	0.0236	0.0197
2	244	17	9	8	0.0697	0.0369	0.0328
1	220	17	6	11	0.0773	0.0273	0.0500
0	201	26	11	15	0.1294	0.0547	0.0746
-1	164	7	2	5	0.0427	0.0122	0.0305
-2	151	4	2	2	0.0265	0.0132	0.0132
-3	143	4	2	2	0.0280	0.0140	0.0140
-4	137	2	1	1	0.0146	0.0073	0.0073
-5	129	4	0	4	0.0310	0.0000	0.0310
-6	115	3	0	3	0.0261	0.0000	0.0261
-7	107	4	1	3	0.0374	0.0093	0.0280
-8	98	4	1	3	0.0408	0.0102	0.0306
-9	88	0	0	0	0.0000	0.0000	0.0000
-10	85	1	0	1	0.0118	0.0000	0.0118
-11	72	0	0	0	0.0000	0.0000	0.0000
-12	59	1	1	0	0.0169	0.0169	0.0000
-13	54	1	0	1	0.0185	0.0000	0.0185
-14	51	2	0	2	0.0392	0.0000	0.0392
-15	47	2	0	2	0.0426	0.0000	0.0426
-16	43	1	0	1	0.0233	0.0000	0.0233
-17	40	0	0	0	0.0000	0.0000	0.0000
-18	36	1	0	1	0.0278	0.0000	0.0278
-19	34	1	1	0	0.0294	0.0294	0.0000
-20	32	2	0	2	0.0625	0.0000	0.0625

Table 10: Semiparametric Hazard Model Estimates[a]
Missouri UI Recipients (n=756)

Variable	Total Hazard	Recall Hazard	New Job Hazard
Expect Recall	.423	2.236	-.500
	(.099)	(.272)	(.135)
Definite Recall	.445	.509	.218
	(.138)	(.148)	(.282)
UI Benefit ($100's), Pre-Exhaust[b]	.381	1.640	-1.115
	(.322)	(.438)	(.447)
UI Benefit ($100's), Post-Exhaust	.496		-.150
	(.838)		(1.136)
Pre-UI Net Weekly Wage ($100's)	-.026	-.075	.048
	(.045)	(.059)	(.061)
Age	-.043	-.039	-.054
	(.024)	(.031)	(.040)
Age Squared / 100	.046	.041	.053
	(.029)	(.039)	(.050)
Pre-UI Job Tenure (years)	.0139	.0260	-.0304
	(.0073)	(.0088)	(.0191)
Education	.032	-.049	.128
	(.018)	(.030)	(.029)
Black	-.404	-.392	-.459
	(.193)	(.247)	(.288)
Female	-.161	.027	-.416
	(.118)	(.145)	(.182)
Time Until Exhaustion Dummies[c]:			
UI 0	.928	.835	.789
	(.235)	(.371)	(.329)
UI 1	.393	.385	.410
	(.300)	(.479)	(.405)
UI 2-5	-.090	-.045	-.164
	(.194)	(.273)	(.291)
UI 6-10	-.167	-.166	-.182
	(.146)	(.208)	(.220)
UI ≤-1	-.636	-.470	-1.423
	(.732)	(.416)	(.976)
Log Likelihood Value	-2416.2	-1388.4	-1275.6

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Table 10 continued

[a]The unemployment spell duration measure utilized is PAYSPELL. Other controls included in each of the specifications are number of dependents, spouse works and married dummies, a dummy indicating whether the spell started before February 1, 1980, 6 industry dummies, weeks from end of pre-UI job until claim date, weeks from claim date until first payment date. In the total and new job hazard models individual baseline hazard parameters are estimated for weeks 1 to 52; spells longer than 52 weeks are censored at 52. In the recall hazard parameters are estimated for the first 30 weeks, after which spells are censored. The numbers in parentheses are asymptotic standard errors.

[b]The UI benefit level variable is constrained to have the same effect before and after exhaustion in the recall hazard model.

[c]The time until exhaustion dummy variables are defined in the text.

Table 11: Semiparametric Hazard Model Estimates Allowing Gamma Heterogeneity[a]
Missouri UI Recipients (n=756)

Variable	Total Hazard	Recall Hazard	New Job Hazard
Expect Recall	.634	2.744	-.874
	(.171)	(.330)	(.253)
Definite Recall	.523	.626	.193
	(.211)	(.229)	(.396)
UI Benefit ($100's), Pre-Exhaust[b]	.602	2.237	-.834
	(.441)	(.599)	(.683)
UI Benefit ($100's), Post-Exhaust	1.397		.395
	(1.058)		(1.522)
Pre-UI Net Weekly Wage ($100's)	-.045	-.107	.075
	(.064)	(.084)	(.100)
Age	-.072	-.066	-.003
	(.038)	(.045)	(.067)
Age Squared / 100	.080	.076	.013
	(.046)	(.055)	(.082)
Pre-UI Job Tenure (years)	.0187	.0339	-.0646
	(.0121)	(.0136)	(.0296)
Education	.029	-.057	.189
	(.032)	(.041)	(.058)
Black	-.748	-.554	-.934
	(.284)	(.354)	(.503)
Female	-.173	.075	-.686
	(.181)	(.218)	(.311)
Time Until Exhaustion Dummies[c]:			
UI 0	.925	.799	.851
	(.276)	(.412)	(.390)
UI 1	.314	.319	.439
	(.341)	(.533)	(.454)
UI 2-5	-.187	-.121	-.172
	(.225)	(.318)	(.330)
UI 6-10	-.232	-.215	-.190
	(.162)	(.228)	(.244)
UI ≤-1	-1.200	-.552	-1.538
	(.869)	(.485)	(1.255)
heterogeneity variance	.836	.880	1.611
	(.257)	(.368)	(.589)
Log Likelihood Value	-2405.2	-1384.3	-1265.3

Continued on the next page

Table 11 continued

[a]The unemployment spell duration measure utilized is PAYSPELL. Other controls included in each of the specifications are number of dependents, spouse works and married dummies, a dummy indicating whether the spell started before February 1, 1980, 6 industry dummies, weeks from end of pre-UI job until claim date, weeks from claim date until first payment date. In the total and new job hazard models individual baseline hazard parameters are estimated for weeks 1 to 52; spells longer than 52 weeks are censored at 52. In the recall hazard parameters are estimated for the first 30 weeks, after which spells are censored. The numbers in parentheses are asymptotic standard errors.

[b]The UI benefit level variable is constrained to have the same effect before and after exhaustion in the recall hazard model.

[c]The time until exhaustion dummy variables are defined in the text.

FIGURE 1

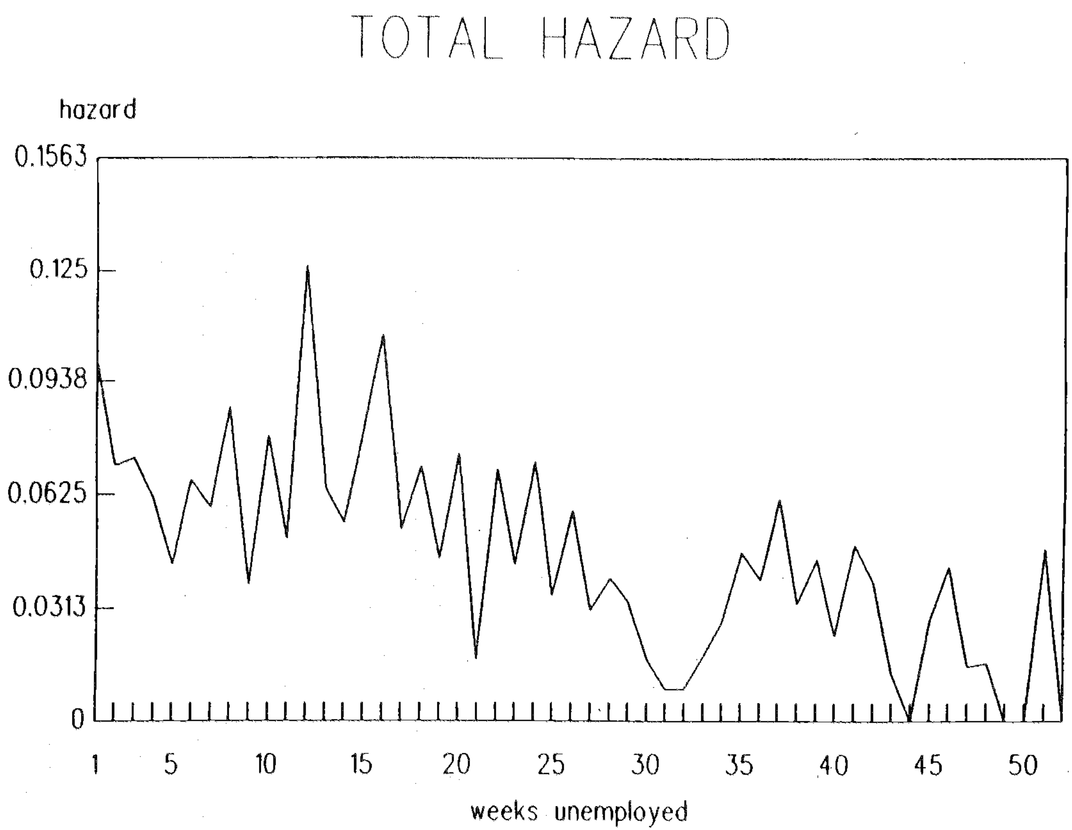
TOTAL HAZARD
hazard
0.1563
0.125
0.0938
0.0625
0.0313
0
1
5
10
15
20
25
30
35
40
45
50
weeks unemployed

FIGURE 2

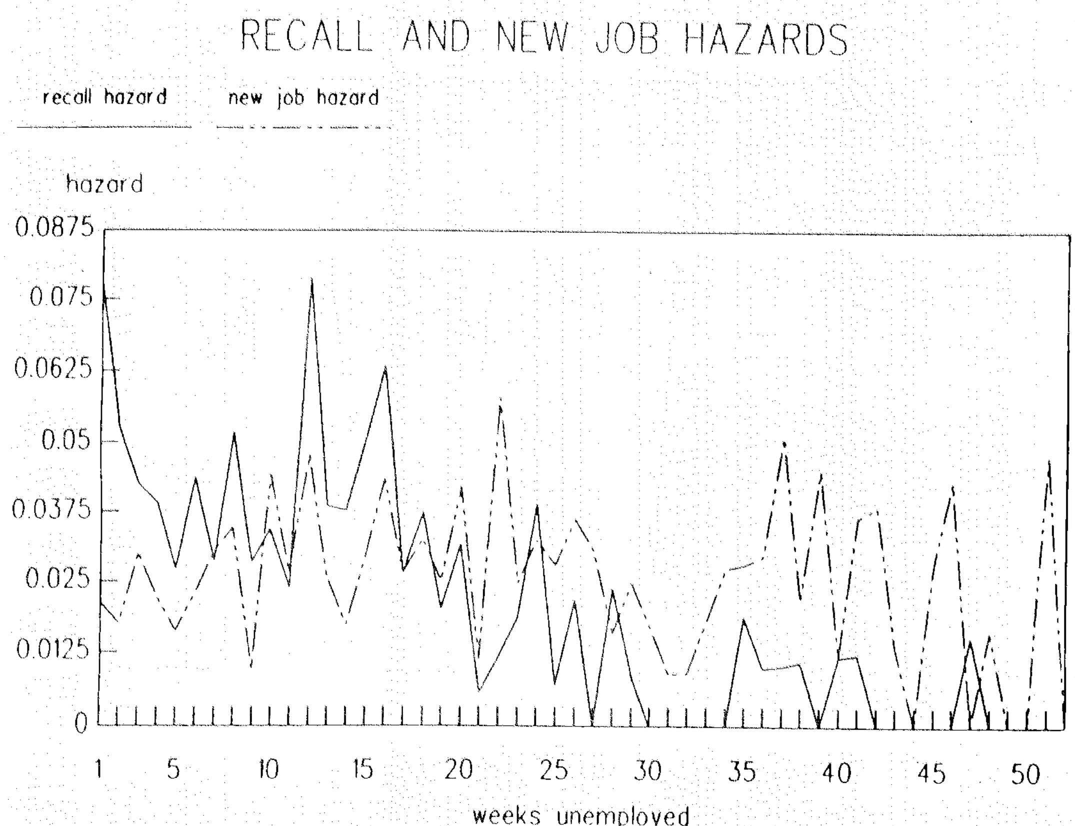
RECALL AND NEW JOB HAZARDS
recall hazard
new job hazard
hazard
0.0875
0.075
0.0625
0.05
0.0375
0.025
0.0125
0
1
5
10
15
20
25
30
35
40
45
50
weeks unemployed

FIGURE 3

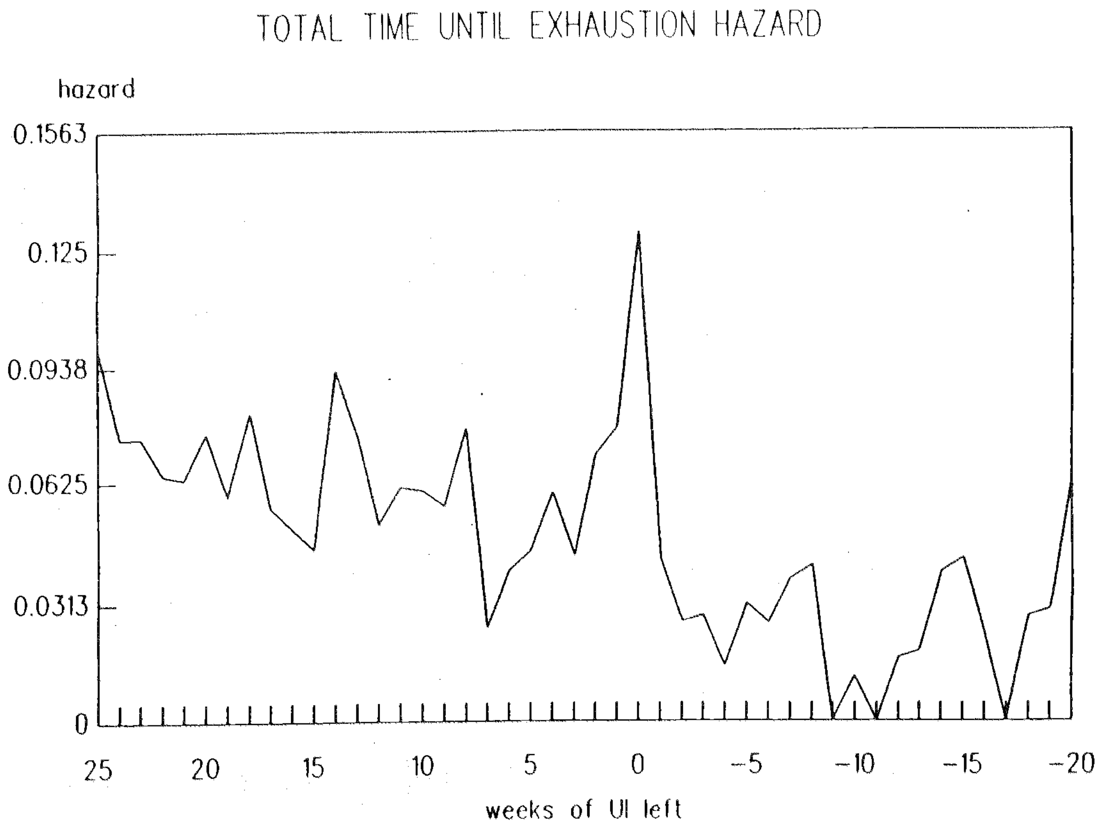
TOTAL TIME UNTIL EXHAUSTION HAZARD
hazard
0.1563
0.125
0.0938
0.0625
0.0313
0
25
20
15
10
5
0
−5
−10
−15
−20
weeks of UI left

FIGURE 4

TIME UNTIL EXHAUSTION
RECALL AND NEW JOB HAZARDS