THE IMPACT OF THE POTENTIAL DURATION OF UNEMPLOYMENT BENEFITS ON THE DURATION OF UNEMPLOYMENT

Lawrence F. KATZ*

*Harvard University and NBER, Cambridge, MA 02138, USA

Bruce D. MEYER*

*Northwestern University and NBER, Cambridge, MA 02138, USA

This paper examines the impact of the potential duration of unemployment insurance (UI) benefits on unemployment in the United States. First, we use a large sample of household heads to examine differences in the unemployment spell distributions of UI recipients and nonrecipients. Sharp increases in the escape rate from unemployment both through recalls and new job acceptances are apparent for UI recipients around the time of benefits exhaustion. Such increases are not apparent at similar points of spell duration for nonrecipients. Second, our analysis of accurate administrative data from 12 states indicates that a one week increase in potential benefit duration increases the average duration of the unemployment spells of UI recipients by 0.16 to 0.20 weeks.

1. Introduction

European countries with relatively generous unemployment insurance (UI) systems (such as Belgium, France, Germany, the Netherlands, and the United Kingdom) suffered much larger and more persistent increases in unemployment in the 1980s than the United States. These differences in European and U.S. unemployment experience are largely explained by the substantially longer duration of unemployment spells in Europe. Furthermore, much

*We thank Robert Moffitt for providing data and answering numerous questions. We are grateful to Gary Burtless, Dale Mortensen, Nick Stern, Lawrence Summers, and an anonymous referee for helpful comments. Financial support from the following sources is also gratefully acknowledged: National Science Foundation Grants SFS 88-09200 (Katz) and SFS 88-21721 (Meyer); an NBER Olin Fellowship in Economics (Katz); and the Industrial Relations Section at Princeton University (Meyer). The data used in this paper will be made available upon request.
microeconomic evidence indicates that there is a positive relation between the level of UI benefits received and the duration of the unemployment spells of UI recipients.¹ These observations have generated much interest among both academics [e.g. Minford (1985)] and the press (e.g. The Economist, 14–20 May 1988, p. 69) in the hypothesis that work disincentives arising from generous unemployment insurance (UI) systems played an important role in high and persistent European unemployment in the 1980s.

Burda (1988) provides some suggestive evidence that differences in the generosity of UI systems may help explain cross-country differences in unemployment performance. Burda finds a strong positive correlation (of 0.61) between a measure of the generosity of UI benefits available to a fully insured worker and the ratio of long-term unemployment to the labor force for fourteen OECD countries in 1985.² On the other hand, Burtless (1987) argues persuasively that European UI systems were generous well before the rise in European unemployment. Many European economies with very generous benefits had much lower unemployment than the less generous United States in the 1960s. Still, typical unemployment spell durations have tended to be quite long in economies with liberal UI systems, even in periods of relatively low unemployment.

In comparing the UI system in the United States with those in Europe, the major differences appear to be in the potential duration of benefits and eligibility requirements for benefits rather than in the weekly benefits level for qualified workers. In particular, insurance and other assistance lasts for more than twice as long in most European countries than in the United States. The potential duration of benefits varies dramatically across countries. The typical qualified worker is eligible for 6 months of benefits in the United States versus over a year of benefits in France, Germany, and Sweden [Burtless (1987)]. For some individuals benefits can last indefinitely (although typically at a reduced rate in the form of means-tested assistance after the first year) in Belgium, Ireland, and the United Kingdom. UI benefits that last for a long duration combined with limited monitoring of search

¹See, for example, Classen (1977) and Solon (1985) for estimates of the impact of benefit levels on spell duration in the United States, and Narendranathan et al. (1985) for estimates for the United Kingdom. Atkinson et al. (1984) present evidence that casts some doubt on the robustness of findings concerning the impact of benefit levels on the re-employment probability in the United Kingdom.

²Burda classifies as long-term unemployed those unemployed workers with current spells of 12 months or longer. The UI measure constructed by Burda combines information on the level of benefits, average manufacturing earnings, and the maximum duration of benefits. Burda’s measure is the present discounted value (using a 20 percent discount rate) of the maximum number of weeks of benefits available to an insured worker relative to the average weekly earnings for a manufacturing worker. This measure ignores taxes and substantial differences across countries in eligibility requirements for UI.
effort may make an economy more susceptible to increases in long-term unemployment in the face of adverse shocks. In fact, Jackman et al. (1989, ch. 4) find that the strong positive relationship between unemployment system generosity and the level of unemployment among OECD countries in the 1980s is driven more by differences in the duration of benefits than by differences in the replacement ratio.

While much microeconomic research has shown that higher levels of benefits are associated with longer durations of unemployment, there is much less empirical research on the impacts of the potential duration of benefits on the duration of unemployment. Since the length of available benefits varies substantially among UI systems, an understanding of how potential benefit duration affects the distribution of unemployment spells is important for determining whether UI differences help explain cross-country differences in unemployment.

In this paper we present new empirical evidence on the impact of the level and potential duration of benefits on the duration of unemployment in the United States. We look at two types of empirical evidence. The first part of our empirical work analyzes the unemployment spells of a sample of household heads from the Panel Study of Income Dynamics. These data allow us to compare spell distributions for UI recipients and nonrecipients and to look at differences in the time pattern of exits from unemployment by recalls and by new job acceptances. Sharp increases in both the recall and new job finding rates are apparent at durations when benefits are likely to lapse for UI recipients. The absence of such increases in the escape rate from unemployment for nonrecipients provides strong evidence of an impact of the potential duration of UI benefits on firm recall policies and workers’ willingness to start new jobs.

The second part of our empirical work examines the impact of the level and length of UI benefits on the escape rate from unemployment for a large sample of UI recipients. This Continuous Wage and Benefit History (CWBH) data set, extracted by Moffitt (1985a), has the advantage of providing detailed administrative records on the UI system parameters facing individuals. Since the data set covers spells in 12 states during the 1978–83 period, a fair amount of both cross-section and time-series variation in UI parameters is available. This variation allows us to directly estimate impacts on the escape rate from unemployment of differences in the level and length of

---

3Moffitt and Nicholson (1982), Moffitt (1985b), and Ham and Rea (1987) are among the few sophisticated econometric studies of the impact of potential benefit duration on the duration of unemployment.

4The prospect of rehire by one’s previous employer is important for a substantial fraction of UI recipients in the United States. For example, Meyer (1989) finds that over 65 percent of UI recipients in the states of Pennsylvania and Idaho from 1979 to 1984 expected to be recalled when they filed for UI benefits.
benefits and test the predictions of alternative models. We find substantial
effects of both the level and length of benefits on spell duration.

The remainder of the paper is organized as follows. Section 2 reviews the
predictions of search models of the effects of UI system parameters on the
escape rate from unemployment. Section 3 presents our comparison of the
unemployment spells of UI recipients and nonrecipients. Section 4 applies
econometric duration models to administrative data on the spells of UI
recipients. Section 5 uses our estimates to simulate the impact of changes in
the level and maximum duration of benefits on the mean duration of
unemployment, the fraction of workers exhausting benefits, and expected
expenditures on UI benefits. Section 6 provides some concluding remarks.

2. Theoretical background

In this section we analyze the likely impacts of the level and potential
duration of unemployment benefits on duration of unemployment and the
time pattern of the escape rate from unemployment. We discuss the
predictions of a standard job search model and of job search models that
incorporate the layoff–rehire process.5

2.1. Standard job search model with no recalls

Mortensen (1977) utilizes a dynamic search model with no recall possi-
bility, variable search intensity, a stationary known wage offer distribution,
and a constant arrival rate of job offers (for a given search intensity) to
analyze the effects of UI on the escape rate from unemployment. Mortensen
incorporates two key features of the UI system in the United States into the
model: benefits are assumed to be paid only for a specified duration rather
than in every period of an unemployment spell, and new entrants or workers
who quit jobs do not qualify for benefits.

As the remaining number of weeks of benefits available to a qualified
unemployed worker decreases, the value of remaining unemployed decreases.
This drop causes the reservation wage to fall and search intensity to increase
as an individual gets closer to when benefits lapse. These changes in behavior
imply that the escape rate from unemployment rises until the date of benefit
exhaustion. After exhaustion, the escape rate is constant given the assump-
tion that the environment is stationary. The time pattern of the exit rate for
an unemployed worker initially qualified for UI benefits with potential

5See Moffitt and Nicholson (1982) for a discussion of the predicted impacts of changes in UI
system parameters on unemployment duration (the fraction of time spent unemployed) in a
static labor supply model.
duration of $P_0$ periods is illustrated by the solid line in fig. 1(a). If individuals can locate jobs and arrange not to begin work until their benefits run out, one might observe a discrete increase in the escape rate near the point of benefits exhaustion followed by a discrete drop after exhaustion.

Mortensen’s model suggests that changes in the level and length of benefits have two opposing influences on the escape rate from unemployment. Increases in either of the benefit parameters have the standard disincentive effect of raising the value of being unemployed, but these increases also raise the value of being employed by increasing the utility associated with being laid off in the future. The second effect, known as the ‘entitlement’ effect, raises the escape rate from unemployment for workers who currently do not qualify for benefits and for qualified workers close to exhaustion.

The effect of an increase in the potential duration of benefits from $P_0$ to $P_1$ is illustrated in fig. 1(a). The standard disincentive effect reduces the escape rate from unemployment for a newly laid-off worker, but the entitlement effect leads to a higher escape rate as one approaches and passes the exhaustion point. The impact of an increase in the benefit level from $b_0$ to $b_1$ is illustrated in fig. 1(b).

---

6The figure is drawn assuming the marginal utility of leisure is independent of income.
The model suggests the following stylized, reduced-form specification for the escape rate from unemployment, \( \lambda \):

\[
\lambda = \lambda(P, P-t, b, b^*(P-t), X), \quad \text{for } P - t \geq 0,
\]

where \( t \) is the duration of the spell, \( P \) is potential duration of benefits, \( P-t \) is time until exhaustion, \( b \) is the level of benefits, and \( X \) is a vector of individual and labor market variables affecting the arrival rate of job offers, search intensity, and choice of reservation wage. The escape rate from unemployment increases as time until exhaustion declines. Higher benefits reduce the escape rate when time until exhaustion is high and increase the escape rate at around exhaustion. Since the entitlement effect is likely to be small relative to the standard search subsidy effect, the average duration unemployment is likely to rise with increases in both the level and potential duration of benefits.

2.2. Recall prospects, UI benefits, and unemployment spell duration

The standard job search model is not entirely appropriate for analyzing the unemployment spells of workers on layoff with some possibility of recall. The interpretation of empirical evidence on the duration of insured unemployment spells in the United States\(^7\) requires a consideration of the role played by recalls since the majority of insured unemployment spells appear to end in recall [Katz (1986), Katz and Meyer (1988)]. 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 a standard model of job search to include an exogenous probability of recall.\(^8\) He shows that under reasonable conditions better recall prospects reduce the new job finding rate by raising the reservation wage and reducing the likelihood of search.

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. The predictions of standard job search models for how variables 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).

Mortensen (1987) analyzes the effects of limited duration UI benefits in a

\(^7\)A majority of insured unemployment spells in Canada also end in recall by the previous employer. See Robertson (1988).

\(^8\)Burdett and Mortensen (1978) and Pissarides (1982) also analyze job search models that incorporate the possibility of recalls.
joint wealth maximizing model of job separations that allows for temporary layoffs. Layoffs occur in response to reductions in match-specific productivity. The reservation wage decreases over the course of an unemployment spell as a worker approaches benefit exhaustion. This induces an increasing new job finding rate and an increasing recall rate as well. Mortensen shows that for realistic parameter values most of the decline in the reservation wage should occur in the last week or two before exhaustion. 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 and that the new job finding rate should increase around exhaustion. The duration and incidence of unemployment spells are shown to rise with increases in the level and length of benefits.

3. UI, recalls, and unemployment spells: Evidence from the PSID

In this section we compare the distributions of unemployment spell durations of UI recipients and nonrecipients in the United States. We analyze employer-initiated unemployment spells in the 1980–81 period for a national sample of household heads. The data are derived from Waves 14 and 15 of the Panel Study of Income Dynamics (PSID). The interviews from these two waves of the PSID provide detailed information on each household head’s last unemployment spell at least partially contained in the calendar year preceding the interview date. For the last unemployment spell in the calendar year prior to the interview, respondents provide retrospective information on the spell duration and the start month of the spell.

The basic sample contains 1,115 layoff and plant-closing unemployment spells for household heads between ages 20 and 65. This data set has two major advantages. First, it contains a large sample of spells for both UI recipients and nonrecipients, and contains information on complete unemployment spells rather than just compensated unemployment. UI benefits...
were received during some part of the spell for 63 percent of the observations (703 spells). Second, our data set allows us to separate the escape rate from unemployment into its component parts: the new job finding rate and the recall rate.

The PSID data set also has some disadvantages. First, information is only available on whether UI is received at sometime during a spell. One cannot accurately identify the level or potential duration of the benefits available. Second, response biases for retrospective information on individual unemployment spells can be severe [Mathiowetz and Duncan (1988)].

Basic descriptive statistics for the entire sample, UI recipients, and UI nonrecipients are presented in table 1. The importance of recall for job losers in the United States is highlighted by the finding that 52 percent of the spells end in recall. The recall rate is 64 percent for manufacturing workers, 59 percent for construction workers, 43 percent for transportation workers, 35 percent for service workers, and 29 percent for trade workers.

There are sharp differences in the characteristics of UI recipients and nonrecipients. UI recipients have much higher wages than nonrecipients. Substantially larger fractions of the UI recipients are white, married, male, and manufacturing workers. The recall rate is also substantially higher for UI recipients. These differences in the characteristics of the two groups help explain the longer mean spell duration for nonrecipients.

Sample hazard functions for UI recipients and nonrecipients

The pattern of unemployment spell durations for UI recipients and nonrecipients from the PSID sample is illustrated in figs. 2(a) and 2(b). The figures plot the Kaplan-Meier empirical hazards for the two samples with the weekly duration data grouped into two-week intervals for ease of presentation. The overall empirical hazard for a given two-week period is the fraction of spells ongoing at the start of the period which end during the two-week interval. The recall and new job empirical hazards, plotted for the two samples in figs. 3(a) and 3(b), are analogously defined as the fraction of spells ongoing at the start of a period which end during the period through recall and through the finding of a new job, respectively.

13The extremely high proportion of nonwhites in the sample results from the oversampling of low-income households in the PSID. The empirical findings are qualitatively quite similar to those presented in this section when only observations from the 'random' portion of the PSID are used.

14It is likely that this is an underestimate of the recall rate for this sample since some of the spells censored at the interview date probably ended in recall.

15Formal duration model estimates reported in Katz (1986) that include demographic variables, industry, occupation, the county unemployment rate, and a plant-closing dummy indicate that the difference in the mean spell duration for the two groups is largely explained by differences in these observed variables.
Table 1
Descriptive statistics for PSID unemployment spell sample; UI recipients and nonrecipients; unemployment spells initiated by plant closings, layoffs, and firings.

<table>
<thead>
<tr>
<th>Variable</th>
<th>Description</th>
<th>Mean (S.D.)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>UI=0</td>
</tr>
<tr>
<td>Duration</td>
<td>Unemployment spell duration in weeks</td>
<td>18.64</td>
</tr>
<tr>
<td>Recall</td>
<td>= 1 if spell ended in recall</td>
<td>0.48</td>
</tr>
<tr>
<td>New job</td>
<td>= 1 if spell ended in taking a new job</td>
<td>0.28</td>
</tr>
<tr>
<td>Censored</td>
<td>= 1 if spell is censored at interview date</td>
<td>0.24</td>
</tr>
<tr>
<td>UI</td>
<td>= 1 if received UI during some part of spell</td>
<td>0.00</td>
</tr>
<tr>
<td>Unemp. rate</td>
<td>= county unemployment rate</td>
<td>7.09</td>
</tr>
<tr>
<td>PC</td>
<td>= 1 if spell initiated by plant closing</td>
<td>0.12</td>
</tr>
<tr>
<td>Wage</td>
<td>Average hourly earnings in calendar year prior to interview</td>
<td>5.69</td>
</tr>
<tr>
<td>Age</td>
<td>Age in years</td>
<td>32.44</td>
</tr>
<tr>
<td>Nonwhite</td>
<td>= 1 if nonwhite</td>
<td>0.61</td>
</tr>
<tr>
<td>Female</td>
<td>= 1 if female</td>
<td>0.19</td>
</tr>
<tr>
<td>Married</td>
<td>= 1 if married</td>
<td>0.52</td>
</tr>
<tr>
<td>Education</td>
<td>Years of schooling</td>
<td>11.31</td>
</tr>
<tr>
<td>Mining</td>
<td>= 1 if in mining or agriculture</td>
<td>0.04</td>
</tr>
<tr>
<td>Construct</td>
<td>= 1 if construction</td>
<td>0.21</td>
</tr>
<tr>
<td>Durables</td>
<td>= 1 if durable goods manufacturing</td>
<td>0.16</td>
</tr>
<tr>
<td>Nondurables</td>
<td>= 1 if nondurable goods manufacturing</td>
<td>0.09</td>
</tr>
<tr>
<td>Transport</td>
<td>= 1 if transportation or utilities</td>
<td>0.10</td>
</tr>
<tr>
<td>Trade</td>
<td>= 1 if wholesale or retail trade</td>
<td>0.16</td>
</tr>
<tr>
<td>Service</td>
<td>= 1 if services</td>
<td>0.24</td>
</tr>
<tr>
<td>White collar</td>
<td>= 1 if managerial, professional, clerical or sales worker</td>
<td>0.38</td>
</tr>
<tr>
<td>Sample size</td>
<td></td>
<td>412</td>
</tr>
</tbody>
</table>

The figures reveal substantial differences in the pattern of the escape rate from unemployment for UI recipients and nonrecipients. The total hazard rates are initially downward sloping for both groups. The total hazard increases substantially in the 25–40 week interval for UI recipients. There are large spikes in the escape rate from unemployment at 26 weeks and at 39 weeks for UI recipients. Spikes of similar magnitude at 26 and 39 weeks are not apparent for UI nonrecipients. While the exact placing of the spikes may be an artifact of the tendency for individuals to report long unemployment rates as lasting exactly half a year, three-quarters, or one year, the much greater importance of these spikes for UI recipients strongly suggests that
Fig. 2. Total hazard.
Fig. 3. Recall and new job hazards.
they may be related to the limited duration of UI benefits. Most UI recipients during this period were eligible for either 26 or 39 weeks of benefits in a benefit year. The escape rate from unemployment appears to increase substantially around when many UI recipients would be exhausting benefits and a much smaller increase in the escape rate is apparent for nonrecipients.

The total hazard rates presented in figs. 2(a) and 2(b) mask sharp differences in the new job finding and recall rates illustrated in figs. 3(a) and 3(b). The recall hazard drops sharply with spell duration for both UI recipients and nonrecipients. Most recalls occur within 8 weeks of the start date of a spell. The new job finding rate differs substantially for UI recipients and nonrecipients. The new job finding rate starts out quite a bit lower and is much more upward sloping for UI recipients. The lower initial new job finding rate for UI recipients may be an artifact of the one-week waiting period before UI eligibility in most states. Individuals expecting quite short spells may also not bother to apply for benefits. On the other hand, these factors could not plausibly account for the differences in the new job finding rate patterns for UI recipients and nonrecipients after the first few weeks. The low initial new job finding rate and apparent positive duration dependence in the new job finding rate for UI recipients provides support for the prediction that UI depresses new job finding when time until exhaustion is large and that the escape rate rises with time until exhaustion. The jumps in the recall and new job finding rates for UI recipients at likely exhaustion points (26 and 39 weeks) are strong evidence for the prediction that firms take into account the duration of UI benefits in designing recall policies [as suggested by the Mortensen's (1987) model of the impact of limited duration UI on recall policies] and that workers become much more likely to take new jobs as their benefits run out. The absence of such patterns for nonrecipients in the PSID sample strongly suggests that these patterns represent behavioral responses by firms and workers to the incentives created by a UI system with limited benefit duration.

The sample hazard functions plotted in the figures do not take into account heterogeneity among individuals in the sample. Although uncontrolled heterogeneity biases estimates of duration dependence in the total 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 model. If uncontrolled factors that raise the recall rate also reduce the new job finding rate, then one can (at least in theory) generate spurious positive duration dependence in the new job hazard. In fact, Katz (1986) finds that positive duration dependence in the new job finding rate for UI recipients is more prevalent when controls for observables are included in formal duration model estimates using this PSID data set. The differences between the escape rates for UI recipients and nonrecipients and the spikes
near exhaustion points for UI recipients also remain when we estimate competing risks models with nonparametric baseline hazards and controls for observables are estimated. Furthermore, Katz and Mayer (1988) find substantial increases in both the recall and the new job finding rates near the week of benefits exhaustion and find strong positive duration dependence in the new job finding rate for a sample of UI recipients in Missouri.

4. Hazard model estimates using the Moffitt data set

This section reports hazard model estimates of the effect of the level and length of UI benefits on unemployment durations. We use Continuous Wage and Benefit History (CWBH) UI administrative records on the compensated unemployment spells of a sample of 3,365 males from 12 states during the period 1978–83. The sample is drawn from a data set previously analyzed by Moffitt (1985a). CWBH data provide accurate information for each individual on the level of UI benefits and their potential duration. Potential duration varies over time since benefits are frequently extended. The number of weeks of benefit receipt is also known exactly. This avoids many of the measurement error problems common in other data sources.

The data set provides enough variation in UI system parameters within and across states and over time to get accurate estimates of the impact of the level and length of UI benefits and the time until benefit exhaustion. The sources of variation in benefit levels in our data are nonlinearities in the benefit schedules (different minima and maxima across states), legislative changes during the sample period, and the erosion of real benefit levels due to inflation between legislative changes. Benefit maxima differ substantially across states. For example, the maximum benefit in Missouri is below the mean benefit in Pennsylvania in our data set. It is hard to make a plausible case for endogeneity of these sources of variation given that we are controlling for the previous wage, as well as state characteristics through state fixed effects.

There are several sources of variation in the potential duration of benefits

16Han and Hausman (1986) have developed an estimator to handle unobserved heterogeneity that is correlated among the risks in a competing risks model. They have implemented their estimator on a sub-sample of our PSID data set that excludes spells initiated by plant closings. Their results indicate that allowing for correlated, unobserved heterogeneity does not qualitatively affect one's inferences for this data set.
17The 12 states are Georgia, Idaho, Louisiana, Missouri, New Mexico, New York, North Carolina, Pennsylvania, South Carolina, Vermont, Washington, and Wisconsin.
18The original Moffitt (1985a) data set contains 4,628 observations. 1,227 observations are excluded because of missing data on age, schooling, dependents or marital status. 36 observations are excluded because the recorded spell is longer than the reported potential duration of benefits.
19The spells in the Moffitt data are periods of benefit receipt. Spells that are interrupted by short periods when benefits are not received are concatenated. This modified spell of benefit receipt may do a better job of grouping together periods of similar behavior. See Moffitt (1985b) for more discussion.
in the data. First, there is variability across states in the length of regular benefits provided. During the sample period, Louisiana typically provided 28 weeks, Pennsylvania provided 30 weeks, while most other states provided 26 weeks of benefits. Second, benefits were extended during periods of high unemployment under several federal programs. The Extended Benefits program extended benefits 50 percent beyond state durations, up to a maximum of 39 weeks, whenever the insured unemployment rate was above a trigger level. Two other programs provided supplemental benefits. At the beginning of the sample period, the Federal Supplemental Benefits program provided up to a total of 65 weeks of benefits. Beginning in the Fall of 1982 the Federal Supplementary Compensation program provided up to 62 weeks of benefits. Finally, within a state at a point in time the length of benefits may depend on an individual’s work history.

A disadvantage of the CWBH data is that it only covers compensated unemployment so that one cannot use it to make inferences about what happens to individuals after benefits are exhausted. The data set also does not permit one to identify whether spells end through recall or the finding of a new job. Thus, one can only analyze the overall unemployment escape rate. Spells that end with the exhaustion of benefits are treated as right-censored since we only know that the spell is at least as long as its length at exhaustion. Spells that are ongoing when the data collection ends are also treated as right-censored. In both cases the appropriate term for right-censoring is included in the likelihood function.

The duration of unemployment spells is analyzed using hazard model techniques. 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 little theoretical justification.

Formally, we parameterize the overall hazard rate of exit 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 the hazard at spell length $t$ is:

$$\lambda_i(t) = \lim_{h \to 0^+} \frac{\text{prob}[T_i \geq t + h \mid T_i \geq t]}{h} = \lambda_0(t) \exp \{z_i(t)'\beta\},$$

where
\( \lambda_0(t) \) 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 
\( \beta \) is a vector of parameters which is unknown.

Our approach estimates \( \beta \) and the baseline hazard parameters \( \gamma(t) \) using maximum likelihood techniques, where

\[
\gamma(t) = \ln \left( \int_t^{t+1} \lambda_0(u) \, du \right).
\]

The effects of unemployment insurance are measured using functions of the benefit level and the length of benefits. The level of benefits and pre-UI earnings after state and federal taxes are used in the specifications below. We measure the effect of an individual's remaining potential duration of unemployment benefits on the hazard rate using the variables UI 1 to UI 41-54 which form a spline in the time until benefit exhaustion. The coefficient on UI 2-5 is the additional effect on the hazard of having moved 1 week closer to exhaustion when one is 2-5 weeks away. The coefficient on UI 1 is the additional effect on the hazard when one moves from 2 to 1 week from exhaustion. Thus, the effect of moving from 6 weeks away to 1 week is four times the UI 2-5 coefficient plus the UI 1 coefficient. The other UI spline coefficients have analogous interpretations.

Formally, let \( \tau \) be the number of weeks until benefits lapse. \( \tau \) decreases by one each week, unless benefits are extended thus increasing \( \tau \). Then

\[
\begin{align*}
\text{UI 1} &= 1, & \text{if } \tau &= 1, \\
&= 0, & \text{otherwise};
\end{align*}
\]

\[
\begin{align*}
\text{UI 2-5} &= \min(6 - \tau, 4), & \text{if } \tau &\leq 5, \\
&= 0, & \text{otherwise}
\end{align*}
\]

\[
\begin{align*}
\text{UI 6-10} &= \min(11 - \tau, 5), & \text{if } \tau &\leq 10, \\
&= 0, & \text{otherwise},
\end{align*}
\]

and similarly for the remaining spline variables.

In some of the later specifications the potential duration of benefits is directly included as an explanatory variable along with the time until exhaustion spline.\(^{20}\) In addition, interaction terms suggested by the theor-

\(^{20}\)It is possible to allow the hazard rate to depend on time \( t \), time until exhaustion \( (P_r - t) \), and potential benefit duration \( P_r \), because \( P_r \) is a time-varying covariate in this sample. \( P_r \) changes over the course of many of the spells as extended or supplemental benefits begin or expire.
ethical models reviewed in section 2 and previous empirical work are included. These variables interact the level of benefits with age, the unemployment rate, and the time until exhaustion. Since benefits are extended in the course of many spells, a variable is included which equals 1 in week $t$ if anytime during a spell it was expected that benefits would lapse in week $t$. The unemployment rate variable used in the specifications is the monthly state unemployment rate, which is interpolated to give a weekly series. State dummy variables (fixed effects) are included in the specifications to account for unobserved differences across states that may be correlated with both the generosity of the state UI system and the character of unemployment in the state. The other included variables are age, race, education, marital status, and the number of dependents.

Discussion of the estimates

The coefficient estimates from the specifications are reported in table 2. Meyer (1988) provides a detailed discussion of the estimates from similar specifications using this data set. In all of the specifications tried the UI benefit level has a strong negative effect on the hazard rate. Specification (1) indicates that a 10 percent increase in the benefit level is associated with a 5.4 percent decrease in the hazard. A strong effect of weeks until benefits lapse is seen from the exhaustion spline coefficient estimates in Specification (1). The hazard increases 94 percent when one moves from 6 weeks to 2 weeks before benefits expire. In the last week the hazard increases an additional 78 percent. Cumulatively, the hazard more than triples as one moves from 6 weeks to 1 week before exhaustion. When it is more than 6 weeks until exhaustion, the time pattern of the hazard is not precisely estimated.

Most states limit the total dollar amount of benefits paid to an individual during a benefit year. The limit typically depends on an individual's base period earnings. If this limit on total payments is binding, an individual may receive a smaller payment in his or her last week of eligibility. Some individuals may not bother to pick up this smaller final check. The possibility that these smaller final payments could spuriously generate the rise in the hazard just before exhaustion was examined using an additional CWBH sample. The sample includes 38,472 individuals from eight states during 1979–84. The eight states include seven of the 12 in our subset of the CWBH data set previously analyzed by Moffitt. We compared the hazard rate in the week before regular benefits lapse when the benefit payment was its full amount and when it was reduced because of the limit on total dollar benefits.

When UI benefits and previous earnings are entered in logarithms, the effect of benefits is somewhat higher. See Meyer (1988).
Table 2

Semiparametric hazard model estimates: Moffitt data set (n = 3,365).

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean (S.D.)</th>
<th>Specification</th>
</tr>
</thead>
<tbody>
<tr>
<td>UI benefit level (1977 dollars)</td>
<td>104.23 (27.91)</td>
<td>-0.0053 (0.0014) - 0.0040 (0.0015)</td>
</tr>
<tr>
<td>Pre-Ul weekly earnings after taxes (1977 dollars)</td>
<td>169.51 (66.52)</td>
<td>0.0026 (0.0005) 0.0025 (0.0005)</td>
</tr>
<tr>
<td>State unemployment rate</td>
<td>8.70 (2.08)</td>
<td>0.0006 (0.0002) 0.0006 (0.0002)</td>
</tr>
<tr>
<td>Age 17–24</td>
<td>0.16 (0.086)</td>
<td>0.234 (0.201) 0.688</td>
</tr>
<tr>
<td>Age 25–34</td>
<td>0.34 (0.076)</td>
<td>0.117 (0.076) 0.118</td>
</tr>
<tr>
<td>Age 35–44</td>
<td>0.24 (0.079)</td>
<td>0.112 (0.079) 0.109</td>
</tr>
<tr>
<td>Age 45–54</td>
<td>0.14 (0.083)</td>
<td>0.034 (0.083) 0.032</td>
</tr>
<tr>
<td>Exhaustion spline:</td>
<td></td>
<td></td>
</tr>
<tr>
<td>UI 1</td>
<td>0.577 (0.249)</td>
<td>0.551 (0.250)</td>
</tr>
<tr>
<td>UI 2–5</td>
<td>0.166 (0.062)</td>
<td>0.036 (0.099)</td>
</tr>
<tr>
<td>UI 6–10</td>
<td>0.005 (0.032)</td>
<td>-0.011 (0.037)</td>
</tr>
<tr>
<td>UI 11–25</td>
<td>-0.006 (0.007)</td>
<td>-0.031 (0.015)</td>
</tr>
<tr>
<td>UI 26–40</td>
<td>0.006 (0.007)</td>
<td>-0.019 (0.019)</td>
</tr>
<tr>
<td>UI 41–54</td>
<td>0.021 (0.138)</td>
<td>-</td>
</tr>
<tr>
<td>Benefits previously expected to lapse</td>
<td>1.537 (0.188)</td>
<td>1.578 (0.189)</td>
</tr>
<tr>
<td>Potential duration of benefits</td>
<td>-</td>
<td>-0.0247 (0.0153)</td>
</tr>
<tr>
<td>Interaction of benefit level and age 17–24</td>
<td>-</td>
<td>-0.0048 (0.0019)</td>
</tr>
<tr>
<td>Interaction of benefit level and ≤ 3 weeks until exhaustion</td>
<td>-</td>
<td>0.0039 (0.0027)</td>
</tr>
<tr>
<td>Log-likelihood value</td>
<td>-8,905.1</td>
<td>-8,900.6</td>
</tr>
</tbody>
</table>

*Variables for education, race, marital status, number of dependents, and 11 state dummy variables are also included. The UI benefit level and pre-Ul weekly earnings variables are in 1977 dollars. The numbers in parentheses are asymptotic standard errors.

benefits. Those who received Extended Benefits or Federal Supplemental Compensation benefits were excluded. The hazard was 21 percent higher (0.257 compared to 0.212) when the last payment was less than the full weekly amount. While these comparisons are somewhat crude since we did not control for individual attributes, they support the hypothesis that a lower benefit amount may cause people not to claim their last week of benefits.
Because only a slim majority of the final payments are less than the full amount, this effect could only cause a 12 percent increase in the overall hazard the week before benefits lapse. Thus, only a small part of the rise in the hazard just before exhaustion could be explained by this fact.

The estimates reported in table 2 also indicate that the probability of a spell ending is very high in a week in which benefits were scheduled to lapse at some point earlier in a spell. One interpretation of this result is that some firms plan the timing of recalls and some workers arrange the starting of a new job to coincide with the end of eligibility for benefits, but do not alter these plans in the face of an extension of benefits. Alternatively, this result could reflect that some people eligible for extended benefits do not claim them. The estimated effects of UI on spell length are nearly identical if this variable is omitted since only a small fraction of spells have unexpected benefit extensions, and the coefficient affects the hazard in only one week.

The estimated effect of the benefit level on the length of unemployment spells is at the high end of the distribution of recent estimates. A consensus of the previous estimates of the effect of a ten percentage point increase in the replacement ratio might be a 0.5–1-week increase in the length of spells. Here the estimate is around 1.5 weeks. Larger estimated effects are a plausible result of better data on spell length and the level and length of benefits. Many other studies have had to impute the level of benefits for individuals and often it is not known who is even eligible for benefits.

Specification (2) in table 2 includes several additional variables. The potential duration of benefits is included as in the original Moffitt (1985a) paper. This variable is time-varying and often increases from 26 to 39 in the course of a spell as benefits are extended. The entitlement effect captured by Mortensen's (1977) job search model leads to the prediction that the coefficient on potential benefit duration should be positive when time until exhaustion is also included as a covariate. In fact, the coefficient estimate is negative and substantial in magnitude, although it is not quite significant at conventional levels. A negative coefficient is consistent with the income effects from more generous benefits postulated by Moffitt and Nicholson (1982). The coefficient estimate implies that a 13-week extension of benefits is associated with a 27 percent decline in the hazard.

Two benefit level interaction variables are also included in Specification (2). Benefits are interacted with the dummy variable Age 17–24. This variable has a large and significant negative coefficient, indicating that the response of younger people to the benefit level is much more elastic. A larger elasticity for younger people was previously found in England by Narendranathan et al. (1985).

An interaction between the level of benefits and time until exhaustion less

\[^{22}\text{See Hamermesh (1977) and Burtless (1986) for surveys of estimates based on U.S. data.}\]
than or equal to 3 weeks was also added. This coefficient tends to support the hypothesis of Mortensen (1977) that higher benefits will have less of an effect near exhaustion (and may even raise the hazard), but the positive coefficient is not quite significant.

Several other specifications were tried, but are not reported. If Specification (2) is estimated without state fixed effects, the benefit level coefficient almost doubles in absolute value to 0.0073 (standard error 0.0011), while the estimated time until exhaustion and potential benefit duration effects are not greatly altered. The exclusion of fixed effects also causes the state unemployment rate to change sign while retaining statistical significance. It appears that higher unemployment states also tend to have longer spells, but when the unemployment rate rises within a given state the mean spell length drops. This finding may be the result of a composition effect arising from the greater frequency of short temporary layoffs in downturns.

Further specifications were estimated which explicitly accounted for the potential impact of omitted variables (unobserved heterogeneity) on the hazard model estimates. Specifications which allowed unobserved heterogeneity with a gamma distribution were tried. The signs and statistical significance of the main coefficients did not change appreciably, but the benefit coefficients (benefit level, benefit and age interaction, potential duration of benefits) rose by about 25 percent.

5. Simulating the impact of changes in the length and level of UI benefits

In this section we simulate the effect of changes in the level and length of UI benefits on the duration of unemployment spells, the exhaustion rate, and the amount of benefits paid. The simulations are somewhat speculative since they only use information on the unemployment spell durations of UI recipients and thereby do not illuminate how changes in UI parameters may affect wages, the incidence of layoff unemployment, or the unemployment experience of non-UI recipients through possible displacement effects.23

Our simulations use the actual sample distribution of the time-invariant covariates from our subsample of the Moffitt data. The parameter estimates used in the simulations are those in table 2. The simulations assume that state unemployment rate is constant during each spell and equal to the start of spell value for each individual in the sample. It is assumed that everyone is eligible for 39 weeks of benefits (the standard potential duration in a period in which extended benefits are triggered).

The most speculative part of the procedure is making an assumption

23See Burtless (1987) for a discussion of the difficulties in going from micro estimates of the effect of UI parameters on unemployment duration to conclusions concerning aggregate unemployment.
concerning behavior after benefits are exhausted. Since the data set covers only compensated unemployment, one cannot use it for inferences concerning post-exhaustion escape rates. We assume that after exhaustion the baseline hazard is equal to the average baseline hazard in our sample and the benefit level is zero. We set the exhaustion spline in the post-exhaustion period equal to 15 weeks before exhaustion. The rationale for this treatment of the exhaustion spline is to avoid the high escape rate from temporary layoffs at the early part of spells and the exhaustion spike found close to the exhaustion point. Katz and Meyer (1988) find for a sample of UI recipients in Missouri that the overall hazard does decline substantially after exhaustion. This decline is largely accounted for by an extremely low recall rate after exhaustion.

5.1. Simulation methodology

The next few paragraphs formally describe the simulation methodology. The key quantity used in the simulations is the predicted survivor function for each individual in week \( t \), conditional on the individual's covariates, \( z(t) \), up until \( t \). The predicted survivor function in week \( t \) is the predicted probability of a spell lasting at least until \( t \) and it is defined by:

\[
\hat{S}(t) = S(t \mid \hat{\gamma}(t), z(t) \beta; t-0, \ldots, t-1) - \exp \left\{ - \int_0^t \hat{\lambda}(u) \, du \right\}
\]

\[= \exp \left\{ - \sum_{\tau=0}^{t-1} \exp [\hat{\gamma}(\tau) + z(\tau) \beta] \right\}, \tag{3}\]

where a hat above an expression denotes an estimated quantity. The aggregate survivor function for the sample is then defined by:

\[
\bar{S}(t) = \frac{1}{N} \sum_{i=1}^{N} \hat{S}_i(t), \tag{4}
\]

where \( N \) is the sample size. Given the aggregate survivor function, the predicted mean weeks of unemployment is calculated using the rolling sum \( M(t) \), which is the predicted weeks of unemployment accumulated by week \( t \):

\[24\text{The simulations are not much changed if the exhaustion spline is treated as if one who exhausts benefits has the exhaustion spline values of an individual at 25 weeks before exhaustion.}\]
Table 3
Simulations using Specification (1) in table 2 (the numbers in parentheses are percentage changes from the base case).

<table>
<thead>
<tr>
<th>Scenario</th>
<th>Predicted mean weeks of unemployment</th>
<th>Predicted mean weeks compensated</th>
<th>Predicted benefits paid per spell</th>
<th>Predicted percentage exhausting</th>
</tr>
</thead>
<tbody>
<tr>
<td>Base case (39 weeks)</td>
<td>18.4</td>
<td>16.6</td>
<td>$1,796</td>
<td>12.9</td>
</tr>
<tr>
<td>Benefit level reduced 10%</td>
<td>16.9</td>
<td>15.5</td>
<td>1,503</td>
<td>10.4</td>
</tr>
<tr>
<td></td>
<td>(-8.2)</td>
<td>(-6.6)</td>
<td>(-16.3)</td>
<td>(-19.4)</td>
</tr>
<tr>
<td>Benefit level reduced 20%</td>
<td>15.4</td>
<td>14.4</td>
<td>1,236</td>
<td>8.2</td>
</tr>
<tr>
<td></td>
<td>(-16.3)</td>
<td>(-13.3)</td>
<td>(-31.2)</td>
<td>(-36.4)</td>
</tr>
<tr>
<td>Benefit level reduced 30%</td>
<td>14.1</td>
<td>13.3</td>
<td>996</td>
<td>6.3</td>
</tr>
<tr>
<td></td>
<td>(-23.4)</td>
<td>(-19.9)</td>
<td>(-44.5)</td>
<td>(-51.2)</td>
</tr>
<tr>
<td>Potential benefit duration</td>
<td>17.6</td>
<td>15.7</td>
<td>1,690</td>
<td>14.6</td>
</tr>
<tr>
<td>reduced to 35 weeks</td>
<td>(-4.3)</td>
<td>(-5.4)</td>
<td>(-5.9)</td>
<td>(13.2)</td>
</tr>
<tr>
<td>Potential benefit duration</td>
<td>16.2</td>
<td>13.6</td>
<td>1,461</td>
<td>20.7</td>
</tr>
<tr>
<td>reduced to 26 weeks</td>
<td>(-12.0)</td>
<td>(-18.0)</td>
<td>(-18.7)</td>
<td>(60.5)</td>
</tr>
</tbody>
</table>

weeks accumulated by $t \equiv M(t) \equiv \sum_{\tau=1}^{t} \bar{S}(\tau)$.

(5)

In all of the simulations $M(104)$, the number of weeks accumulated by the end of two years, was calculated. Since the sum converged rapidly the simulation results would not be very different if we had truncated the sum at 1 or 3 years instead. Thus, the predicted mean weeks of unemployment is defined by $M(104)$.

Predicted mean weeks compensated is defined by $M(d)$, where $d$ is the potential duration of benefits. Predicted benefits paid per spell, $B(d)$, is defined by:

$$B(d) \equiv \frac{1}{N} \cdot \sum_{i=1}^{N} \sum_{\tau=1}^{d} \bar{S}(\tau) \cdot b_i,$$

(6)

where $b_i$ is the UI benefit for individual $i$. Finally, the predicted percentage exhausting UI benefits equals $\bar{S}(d)$.

5.2. Simulation results

Simulations which use Specification (1) are reported in table 3, and simulations using Specification (2) are reported in table 4. The base case scenario predicted values differ appreciably for the two simulations, but the effects of policy changes are very similar. The base case difference occurs
L. F. Katz and B. D. Meyer, Duration of unemployment benefits

Table 4
Simulations using Specification (2) in table 2 (the numbers in parentheses are percentage changes from the base case).

<table>
<thead>
<tr>
<th>Scenario</th>
<th>Predicted mean weeks of unemployment</th>
<th>Predicted mean weeks compensated</th>
<th>Predicted benefits paid per spell</th>
<th>Predicted percentage exhausting</th>
</tr>
</thead>
<tbody>
<tr>
<td>Base case (39 weeks)</td>
<td>14.7</td>
<td>13.5</td>
<td>$1,455</td>
<td>9.0</td>
</tr>
<tr>
<td>Benefit level reduced 10%</td>
<td>13.5</td>
<td>12.5</td>
<td>1,215</td>
<td>7.2</td>
</tr>
<tr>
<td></td>
<td>(-8.2)</td>
<td>(-7.4)</td>
<td>(-16.5)</td>
<td>(-20.0)</td>
</tr>
<tr>
<td>Benefit level reduced 20%</td>
<td>12.4</td>
<td>11.6</td>
<td>999</td>
<td>5.7</td>
</tr>
<tr>
<td></td>
<td>(-15.6)</td>
<td>(-14.1)</td>
<td>(-31.3)</td>
<td>(-36.7)</td>
</tr>
<tr>
<td>Benefit level reduced 30%</td>
<td>11.3</td>
<td>10.8</td>
<td>807</td>
<td>4.4</td>
</tr>
<tr>
<td></td>
<td>(-23.1)</td>
<td>(-20.0)</td>
<td>(-44.5)</td>
<td>(-51.1)</td>
</tr>
<tr>
<td>Potential benefit duration reduced to 35 weeks</td>
<td>13.9</td>
<td>12.8</td>
<td>1,376</td>
<td>9.8</td>
</tr>
<tr>
<td></td>
<td>(-5.4)</td>
<td>(-5.2)</td>
<td>(-5.4)</td>
<td>(8.9)</td>
</tr>
<tr>
<td>Potential benefit duration reduced to 26 weeks</td>
<td>12.6</td>
<td>11.4</td>
<td>1,222</td>
<td>13.4</td>
</tr>
<tr>
<td></td>
<td>(-14.3)</td>
<td>(-15.6)</td>
<td>(-16.0)</td>
<td>(51.1)</td>
</tr>
</tbody>
</table>

because the simulations assume that potential duration does not change during the unemployment spells, while in the actual Moffitt data the potential duration of benefits rises in the course of many spells as benefits are extended. This effect is captured through the baseline hazard estimates in Specification (1) rather than directly in the potential duration of benefits coefficient as in Specification (2). In both sets of simulations the potential duration of benefits is assumed to be constant over time, but in Specification (1) the baseline hazard estimates implicitly incorporate increases in the potential duration of benefits from extended benefits triggers turning on during the course of a spell.

We simulated the impact of changes in UI parameters on the predicted mean completed spell of unemployment and on the mean weeks of compensated unemployment. Tables 3 and 4 report the following policy experiments: 10, 20 and 30 percent reductions in the level of benefits, and changes in the potential duration of benefits from 39 to either 35 or 26 weeks. A change in maximum potential benefit duration from 26 to 39 weeks is exactly the natural policy experiment that occurs when extended benefits are triggered in the United States.

Changes in the level of benefits and changes in the potential length of benefits have substantial effects on the mean duration of unemployment of UI recipients. The two sets of simulations provide quite similar estimates of the impact of policy changes. An increase in the potential duration of benefits from 26 to 39 weeks is predicted to raise the mean unemployment spell duration by 2.1 weeks in both simulations. This is a surprisingly large
effect given that most spells are completed well before the 26 weeks of regular benefits run out. An increase in potential benefit duration will mechanically increase the mean compensated spell duration even if benefits have no incentive effect since previously uncompensated unemployment will be classified as compensated unemployment. The predicted effect of an extension of benefits from 26 weeks to 39 weeks if there are no incentive effects from extending benefits [using Specification (2)] is a 0.9 week increase in mean compensated spell duration. Thus, most of the impact of extended benefits on compensated unemployment arises through the negative effects of UI on the escape rate from unemployment.

We conclude from an examination of a variety of simulated changes in potential benefit durations that a 1-week extension of benefits increases the mean duration of an unemployment spell by approximately 0.16–0.20 week. One caveat in interpreting these estimates is that some of the variation in the potential length of benefits arises from the extension of benefits in times of poor macroeconomic conditions. If the time-varying state unemployment rate variable included in our specifications does not fully capture labor market conditions, then our estimates of the impact of increases in potential benefit duration may partially reflect that potential benefit duration is high when job availability is low.

Our estimates of the impact of potential benefit duration on the average unemployment spell duration of UI recipients are a bit larger than most of those in the literature. Our estimates are slightly larger than Moffitt's (1985a) estimate of 0.15 week from a model that does not include state dummy variables. Moffitt and Nicholson (1982) find, using a static labor supply estimation framework, that a 1-week extension raises the average unemployment duration by 0.10 week. Moffitt and Nicholson's sample consists of 'long-term' unemployed workers who had exhausted their benefits; such individuals may plausibly be less sensitive to benefits than a more representative group of UI recipients such as in our data set. In a study of Canadian UI recipients, Ham and Rea (1987) find that a 1-week increase in the duration of benefits increases the mean duration to the start of a new job by 0.26–0.33 week in a competing risks framework.25

The hypothetical UI parameter changes examined in the simulations have substantial effects on the amount of benefits paid per spell. A reduction in the level of benefits by 10 percent has an impact on the UI budget similar to a reduction in the potential duration of benefits from 39 to 26 weeks. The simulations indicate that increases in potential benefit duration have much larger adverse incentive effects on unemployment than do changes in the level of benefits that have the same effect on the UI budget. The simulations

---

25Since some of the spells in their sample end in recall, it is difficult to translate this finding into an estimate of the effect on the mean duration of unemployment.
in table 4 show that the budget cut from the base case accomplished through a 10 percent reduction in benefits reduces mean unemployment by 1.2 weeks, while a similar budget cut done through eliminating extended benefits generates almost twice as large a reduction in unemployment.

These findings suggest that a government with a fixed UI budget faces a sharp trade-off between incentives and insurance in the design of the level and time sequence of UI payments. A balanced-budget combination of a reduction in the level with an increase in the maximum duration of benefits has strong adverse incentive effects although it does provide greater protection for those who are unlucky in their attempts to gain re-employment.

5.3. The impact of extended benefits on the income of the unemployed

Broadly, our results suggest that the behavioral effects of UI are extremely important. The estimated incentive effects of extended benefits are large enough to allow the possibility that a benefit extension could actually reduce the total money income of UI recipients. If the benefit extension did not affect unemployment duration or re-employment earnings, then increasing the weeks of unemployment in which benefits are received would raise the income of the unemployed. On the other hand, if a higher duration of benefits increases unemployment duration and does not affect re-employment wages, the extension of benefits may reduce the income (although probably not the welfare) of UI recipients if re-employment wages are higher than UI benefits.

The following simple calculations are instructive concerning the incentive effects of increases in benefit duration. We make the strong assumption that re-employment weekly wages are unaffected by the availability of extended benefits. Under this assumption, the change in UI recipient's money income arising from the extension of benefits is given by the formula:

\[
\Delta \text{income} = \{\Delta(\text{weeks of compensated unemployment}) \times (\text{weekly UI benefit})\} - \{\Delta(\text{total weeks of unemployment}) \times (\text{re-employment weekly wage})\}.
\]

In table 5 we present the predicted impact on the income of a UI recipient (with pre-UI weekly earnings equal to our sample average of $170) of an increase in potential benefit duration from 26 to 39 weeks. We use the simulations discussed above based on Specifications (1) and (2) from table 2. We consider three cases for each specification. The first case assumes there are no behavioral effects of extended benefits. The second case assumes that extended benefits increase unemployment by the amounts shown in our simulation results (in tables 3 and 4) and that re-employment wages are 90 percent of pre-UI weekly wages. The final case assumes these same behavioral effects on the duration of unemployment, but assumes that re-employment wages are only 60 percent of pre-UI weekly wages. Katz and
Table 5
The effect of an increase in potential benefit duration from 26 to 39 weeks on the income of a typical UI recipient.

<table>
<thead>
<tr>
<th>Scenario</th>
<th>Change in unemployment income</th>
<th>Change in wage income</th>
<th>Net change in income</th>
</tr>
</thead>
<tbody>
<tr>
<td>Specification (1) from table 2</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>No behavioral effects of UI</td>
<td>$179</td>
<td>$0</td>
<td>$179</td>
</tr>
<tr>
<td>Behavioral effects:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Re-employment weekly wage equals 90% of pre-UI weekly earnings</td>
<td>336</td>
<td>-337</td>
<td>-1</td>
</tr>
<tr>
<td>Re-employment weekly wage equals 60% of pre-UI weekly earnings</td>
<td>336</td>
<td>-225</td>
<td>111</td>
</tr>
<tr>
<td>Specification (2) from table 2</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>No behavioral effects of UI</td>
<td>$98</td>
<td>$0</td>
<td>$98</td>
</tr>
<tr>
<td>Behavioral effects:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Re-employment weekly wage equals 90% of pre-UI weekly earnings</td>
<td>223</td>
<td>-321</td>
<td>-98</td>
</tr>
<tr>
<td>Re-employment weekly wage equals 60% of pre-UI weekly earnings</td>
<td>223</td>
<td>-214</td>
<td>9</td>
</tr>
</tbody>
</table>

*Note:* These calculations assume a pre-UI weekly wage of $170. All figures are in 1977 dollars. The change in income of a UI recipient arising from the extension of benefits is given by the formula:

\[
\Delta \text{income} = [\Delta(\text{weeks of compensated unemployment}) \times (\text{weekly UI benefit})] - [\Delta(\text{total weeks of unemployment}) \times (\text{re-employment weekly wage})].
\]

Meyer (1988) find that the typical UI recipient who gains re-employment within a year of layoff has (initial) re-employment weekly earnings that are approximately 10 percent less than pre-UI weekly earnings. On the other hand, those who were not recalled and exhausted benefits averaged 50 percent losses in weekly earnings.

The calculations presented in table 5 suggest that extending benefits may reduce the total money income of UI recipients. If one assumes that re-employment earnings are 90 percent of previous earnings, then both specifications yield the prediction that the income of the typical UI recipient actually falls in response to an increase in potential benefit duration from 26 to 39 weeks. If workers whose behavior is most strongly affected by extended
benefits have low re-employment wages, then it is likely that extended benefits raise the money income of UI recipients. Of course, these simple calculations ignore the increases in leisure accruing to UI recipients from greater unemployment and do not take into account the possibility that extended benefits may allow workers to make better job matches raising future earnings from employment.

6. Conclusions

The evidence presented in this paper indicates that the potential duration of UI benefits has a strong impact on the duration of the unemployment spells of UI recipients in the United States. Our examination of data from the PSID indicates that the distributions of unemployment spell durations of UI recipients and nonrecipients are quite different. Substantial increases in both the recall rate and new job finding rate are apparent for UI recipients around the time when benefits are likely to lapse. Large increases in the escape rate from unemployment in the several weeks before exhaustion are also apparent for a large sample of UI recipients for which administrative data allow us to accurately date the end of the spell and the point at which benefits are exhausted. It seems safe to conclude that potential benefit duration has significant behavioral effects on firm recall policies and worker new job finding strategies. Furthermore, our estimates indicate that policies that extend benefits have much greater adverse incentive effects on the duration of unemployment than policies with the same predicted impact on the government budget which raise the level of benefits.

Our results indicate that a 1-week increase in potential benefit duration increases the average duration of the unemployment spells of UI recipients by about 0.16 to 0.20 week. These estimates can be used guardedly to make a rough guess as to what the impact of longer potential benefit durations in Europe than in the United States is on the mean duration of unemployment. An increase in potential benefit duration from 6 months to 1 year is predicted to increase mean duration of unemployment by 4–5 weeks, and an increase from 6 months to 2 years is predicted to generate a 13–16-week increase in unemployment duration. The average uncompleted duration of ongoing spells was 63 weeks (14.8 months) in the United Kingdom in 1984 versus 18 weeks in the United States [Jackman et al. (1989)]. The fraction of the unemployed covered by benefits is also much lower in the United States than in the United Kingdom. Thus, longer duration of benefits may be able to explain about 10–30 percent of the difference in mean unemployment spell durations between the United States and the United Kingdom.

Two caveats about our results should be kept in mind. First, while lower

\[^{26}\text{Blank and Card (1989) and Kane (1988) document the recent decline in the fraction of the unemployed receiving UI in the United States.}\]
unemployment benefits might decrease the length of UI recipients' spells, the spells of non-U1 recipients might rise due to congestion/displacement effects. If aggregate employment is determined by the level of demand, and the matching of particular workers to jobs is not important, shorter unemployment spells for one group would imply longer spells on average for others. This effect would imply that our estimates of the microeconomic effects of UI on unemployment are an overestimate of the macroeconomic effects. Second, we have concentrated on transitions in one direction between only two of the possible labor market states. Clark and Summers (1982) and Topel (1985) have emphasized the effects of UI on other transitions. A more encompassing analysis of the effects of UI might yield different conclusions about the aggregate effects of changes in the level and length of UI benefits.

References

Han, Aaron and Jerry A. Hausman, 1986, Semi-parametric estimation of competing risks models, MIT, unpublished manuscript.
Jackman, R., R. Layard, S. Nickell and S. Wadwhani, 1989, Unemployment (London School of Economics) unpublished manuscript.
Kane, Thomas J., 1988, What happened to unemployment insurance? Administrative reforms and compositional changes (Harvard University, Kennedy School of Government) unpublished manuscript.
Meyer, Bruce D., 1988, Unemployment insurance and unemployment spells, NBER working paper no. 2546.
Meyer, Bruce D., 1989, An event study approach to the effects of unemployment insurance (Northwestern University) unpublished manuscript.
Mortensen, Dale T., 1987, A structural model of UI benefit effects on the incidence and duration of unemployment (Northwestern University) unpublished manuscript.
Robertson, Matthew, 1988, Temporary layoffs and the measurement of unemployment: An analysis for Canada (Employment and Immigration Canada, Ottawa) unpublished manuscript.