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## Do Extended Unemployment Benefits Lengthen Unemployment Spells? Evidence from Recent Cycles in the U.S. Labor Market<sup>1</sup>

Henry S. Farber  
Princeton University, NBER, IZA

Robert G. Valletta  
Federal Reserve Bank of San Francisco, IZA

### Abstract

In response to the Great Recession, the availability of unemployment insurance (UI) benefits was extended to an unprecedented 99 weeks in many U.S. states in the 2009-2012 period. We use matched monthly data from the CPS and exploit variation in the timing and size of the UI benefit extensions across states to estimate the overall impact of these extensions on individual exit from unemployment. We compare the estimated effect for the recent recession period with that for the prior extension of benefits during the much milder downturn in the early 2000s. In both periods, we find a small but statistically significant reduction in the unemployment exit rate and a small increase in the expected duration of unemployment. The effects on exits and duration are primarily due to a reduction in exits from the labor force rather than to a decrease in exits to employment (the job finding rate). Although the overall effect of UI extensions on exit from unemployment is small, it implies a substantial effect of extended benefits on the steady-state share of unemployment in the cross-section that is long-term.

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<sup>1</sup>Farber: Industrial Relations Section, Firestone Library, Princeton University, Princeton, NJ 08544. Phone: (609)258-4044. email: farber@princeton.edu. Valletta: Federal Reserve Bank of San Francisco, 101 Market St. San Francisco, CA 94105. Phone: (415)974-3345. email: rob.valletta@sfrb.org. We thank participants at numerous workshops and conferences since 2011 for their comments on various versions of this paper. We especially thank Jesse Rothstein for detailed discussions regarding data construction and Scott Gibbons of the U.S. Department of Labor and Julie Whittaker of the Congressional Research Service for their assistance with obtaining and interpreting data on extended UI benefits. We also thank Katherine Kuang and Leila Bengali for outstanding research assistance. The views expressed in this paper are solely those of the authors and should not be attributed to the Federal Reserve Bank of San Francisco or the Federal Reserve System.

# 1 Introduction

Compared with other advanced industrial countries, the United States is among the least generous with respect to the duration and level of unemployment insurance (UI) benefits (OECD 2007). Under normal economic circumstances, UI benefits in the United States are available for up to six months following job loss, compared with availability of a year or longer in many European countries. In response to the severe labor market downturn associated with “The Great Recession” of 2007-09, however, UI benefit availability was successively extended in the United States, reaching a maximum duration of 99 weeks as of late 2009 and continuing into 2012. This unprecedented expansion of UI availability has been the subject of intense policy debate, which has largely revolved around the incentive effects of UI payments on job search and prolonged labor force attachment.<sup>1</sup> In this paper, we provide an empirical assessment of the impact of extended UI on exit rates from unemployment and duration of unemployment in the United States. We focus in particular on a comparison between the effects of the recent UI extensions and those triggered by the earlier, less severe labor market downturn in the early 2000s.

Past empirical research has produced a range of estimates regarding the disincentive effects of UI benefits on job search in the United States (e.g., Moffitt 1985, Katz and Meyer 1990, Card and Levine 2000). However, as noted by others (e.g., Katz 2010), the impact of UI benefits on job search likely was higher in the 1970s and 1980s than it is now, due to the earlier period’s greater reliance on temporary layoffs and the corresponding sensitivity of recall dates to unemployment insurance benefits. Moreover, recent research suggests that the disincentive effects of UI are limited by the reduced returns to job search under weak labor market conditions (e.g., Landais, Michaillat, and Saez 2010; Kroft and Notowidigdo 2011); it may be that such considerations loomed especially large during the Great Recession. Rothstein (2011), who presents an analysis of the effects of the recent UI extensions that is related to ours, reports small effects of the recent UI extensions on unemployment exits, duration, and the overall unemployment rate.

Because extended UI benefits were much more widely available during the Great Recession than during earlier periods and because of the severity of the recent labor market downturn, earlier empirical results cannot be reliably extrapolated to assess UI disincentive effects in the recent episode. We estimate these effects by developing a framework that relies on current labor market data and detailed information on the recent UI expansions. We

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<sup>1</sup> Relevant prior research includes Chetty (2008) and Card, Chetty, and Weber (2007).

use microdata at the individual level from the monthly survey of households and individuals that is used to construct official unemployment and labor force statistics in the United States (the Current Population Survey, or CPS). We match observations on individuals across consecutive months of the data, which enables us to analyze transitions out of unemployment (exits), distinguishing between new job finding and labor force withdrawal. To assess the impact of UI extensions, we have compiled a detailed database of trigger dates and maximum available UI weeks at the state level. The extension of UI availability proceeded gradually, and its extent and timing varied across states. Qualification for multiple UI extensions at the state level occurred based on the level and change in state unemployment rates and UI reciprocity rates.

Exploiting the different timing and degree of extension activation across states, we estimate the effects of the extensions on unemployment exit rates and duration. We use both a single-risk framework based on overall exits from unemployment and a competing-risks framework that distinguishes between exits to employment and exits out of the labor force (the cessation of search activity). Identification of the UI effects is based on individual variation in benefit availability, conditional on state economic conditions and individual characteristics.

A key feature of our analysis is that we distinguish between exiting unemployment through finding a job and exiting unemployment by exiting the labor force. This is important because the efficiency implications of extended benefits differ depending on whether these benefits delay job finding or simply label those not employed as unemployed rather than not participating in the labor force.

Our analysis is related to that of Rothstein (2011), who found small effects of the recent UI extensions on unemployment exits (mainly transitions out of the labor force), duration, and the overall unemployment rate. However, our analysis differs from and extends his in a number of ways.

- We identify the effects of UI through variation in individuals' time to exhaustion, which varies across individuals and over time as a function of total weeks of UI available in a state and individuals' duration of unemployment.<sup>2</sup>
- We expand the period of analysis by comparing the impact of this recent episode of UI extensions with the earlier episode that followed the relatively mild 2001 recession.

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<sup>2</sup> While Rothstein presents some estimates that rely on a similar identification scheme, his primary estimates rely on variation across states and over time in total weeks of UI available.

This comparison is important for policy formulation. If the effects of UI extensions on job search differ substantially depending on aggregate labor market conditions, as suggested by recent work (Landais et al. 2010; Kroft and Notowidigdo 2011), the impact in the most recent recession may not provide a good guide for future policy.

- We extend the period of analysis through 2012 rather than stopping in early 2011. Our analysis therefore incorporates a period during which the labor market was slowly recovering from the Great Recession and extended benefits were being phased out on a state-by-state basis, through legislative changes and state-specific improvements in labor market conditions. Given the possibility of delayed behavioral responses to legislation that may have been regarded as temporary by job searchers, this expanded sample period enables us to estimate a more complete behavioral response to benefit expansions, and the rollback of extended UI has the potential to provide additional identifying information.
- We reconcile the duration of unemployment derived from our flow-based sample of ongoing unemployment spells (from the matched monthly CPS data) with the steady-state distribution of unemployment durations of spells in progress obtained from the monthly CPS cross-sections. This enables us to provide precise quantitative estimates of the contribution of extended UI to observed long-term unemployment, for which we find relatively large effects

To preview our results, we estimate small but statistically significant reductions in the unemployment exit rate arising from both sets of unemployment extensions, and we find that the magnitudes of these effects are similar across the two episodes of UI extensions. Our estimates further imply a small increase in the expected duration of completed unemployment spells. While the implied increase in expected duration is larger in the later (Great Recession) episode of UI extensions, this difference in magnitudes is due entirely to the fact that extended unemployment benefits were more widespread and more generous in the Great Recession period. Our competing risks analysis reveals that the effects of extended benefits on exit from unemployment occur primarily through a reduction in labor force exits rather than a reduction in job finding, with a particularly pronounced effect on labor force attachment in the recent episode. Interestingly, despite the relatively small effect of extended benefits on the expected duration of completed spells, our estimates imply a substantial effect of extended benefits on the steady-state share of unemployment in the cross-section that is long-term.

## 2 UI Program Characteristics and Research

### 2.1 Normal and extended benefits

UI benefits are normally available for 26 weeks in the United States under the joint federal-state Unemployment Compensation (UC) program established under the Social Security Act of 1935. Unemployed individuals are eligible to receive benefits if they lost a job through no fault of their own (typically a permanent or temporary layoff) and they meet state-specific minimum requirements regarding work history and wages during the 12 to 15 month period preceding job loss. Availability for work and active job search typically are required for ongoing receipt of UI benefits, although the exact rules vary across states.

Normal UI benefits periodically are supplemented and extended during episodes of economic distress, through a combination of permanent and temporary legislation.<sup>3</sup> The federal Extended Benefits (EB) program, permanently authorized beginning in 1970, provides up to 20 weeks of additional unemployment compensation for unemployed individuals who lost jobs in states where the level and change in the state unemployment rate is above a specified threshold. The thresholds or triggers are state specific but most commonly are based on an overall unemployment rate of 6.5 percent (for a 13-week extension) or 8.0 percent (for 20 weeks), combined with a 10-percent increase in the unemployment rate over the previous two years. The EB program has been supplemented by temporary programs that have been used eight times since 1958, with the most recent episode beginning in 2008. We focus on the two episodes of UI extensions since 2002.<sup>4</sup>

The severity of job loss and persistent labor market weakness during and after the recession of 2007-2009 resulted in an unprecedented expansion of UI benefit availability and takeup. Between mid-2008 and late 2009 a set of expansions resulted in availability of UI benefits up to a maximum of 99 weeks in many states. A similar but much more limited extension of UI benefits occurred through the Temporary Extension of Unemployment Compensation (TEUC) legislation that was effective from March 2002 through early 2004. A maximum of 72 weeks of total benefits were phased in during this period. We describe the timing of these expansions in detail in Appendix I.

As suggested by this discussion and the detailed description in Appendix I of the TEUC

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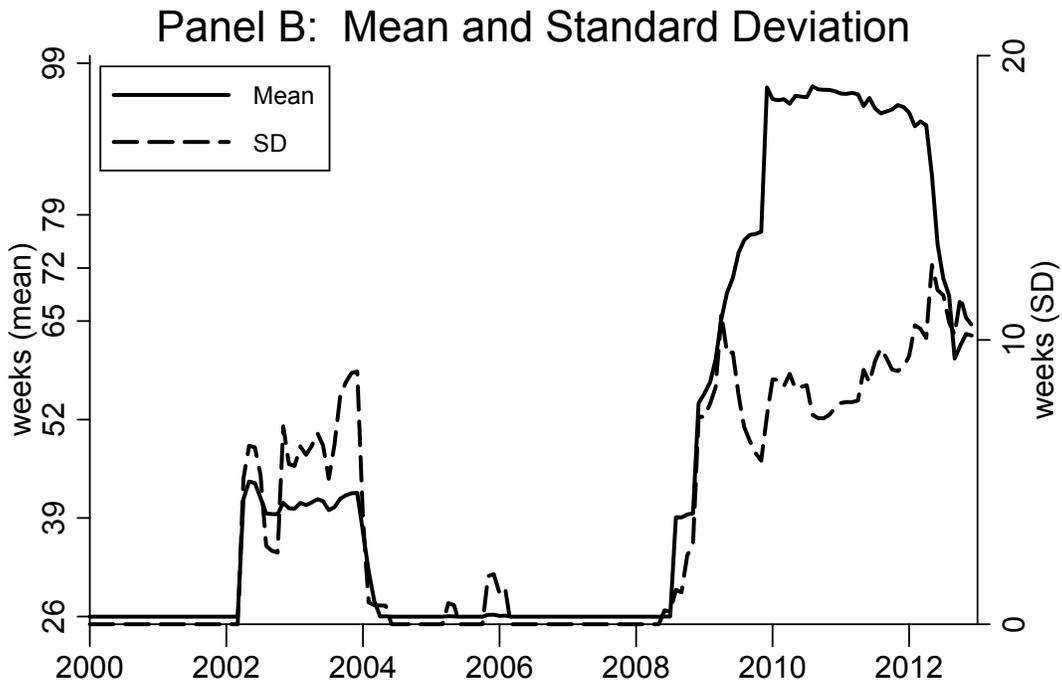
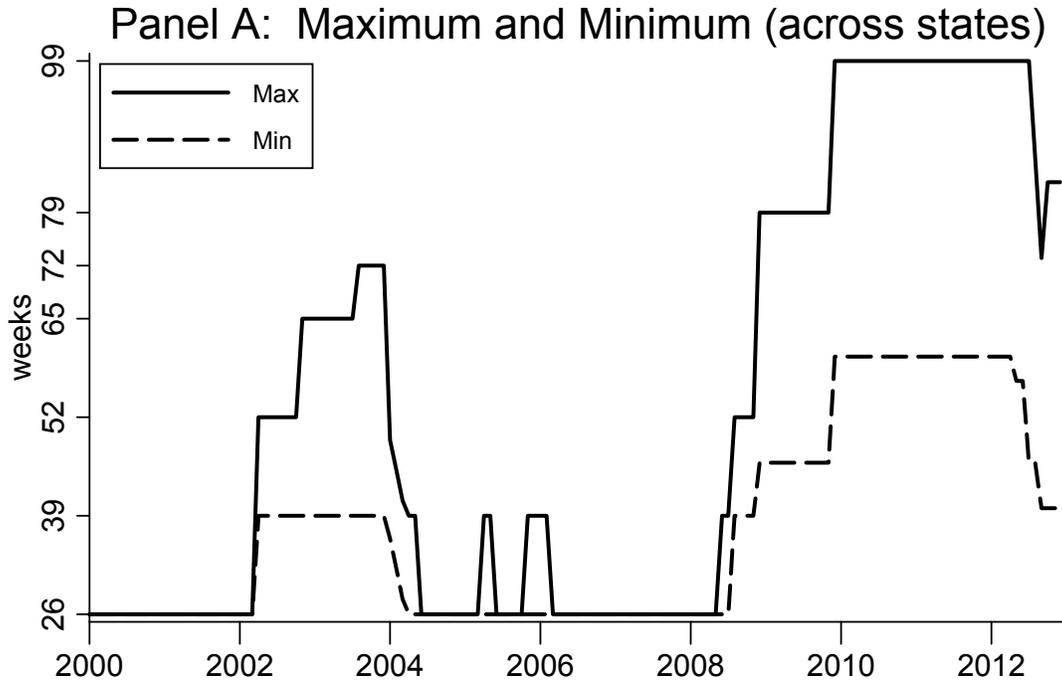
<sup>3</sup> See Whittaker (2008) and Whittaker and Isaacs (2012) for details regarding the various historical and current programs that provide extended UI benefits.

<sup>4</sup> Analysis of UI extensions prior to 2002 is precluded by the difficulties of obtaining precise data on the timing of state-level extensions.

(2002-04) and EUC (2008-forward) programs, the timing of the extended UI triggers and consequent maximum duration of UI eligibility has varied substantially across states and over time. Different states surpassed the trigger levels for EB and TEUC/EUC availability at different times; some states never achieved the unemployment rates necessary for the complete extensions; and states saw available weeks rolled back as labor market conditions improved in 2011-12 and also as a result of the legislated rollback of maximum weeks available through the EUC program in late 2012.

Figure 1 illustrates the variation in eligibility for extended UI over time (years 2000-2012) based on the various programs in effect. Panel A displays the maximum and minimum number of total UI weeks available across states, and Panel B displays the average and standard deviation of the distribution of total weeks of UI available across unemployed individuals (measured using a sample of all individuals identified as unemployed and eligible to receive UI in the CPS microdata; see our definition of eligibility below). The spread between the maximum and minimum number of weeks was similar between the most recent episode and the preceding episode in the early 2000s, at about 26-27 weeks. However, the number of states at or near the minimum in the recent episode, and their size, was much smaller than in the preceding episode. This is reflected in Panel B, which shows that the average weeks of total UI eligibility reached about 96 in late 2009, implying that the typical unemployed individual was located in a state in which maximum UI eligibility was 99 weeks. In the early 2000s, maximum weeks of eligibility reached 72, but few states triggered on to the maximum extensions, and only about 13 additional weeks of UI beyond the normal 26 were available to the typical unemployed individual. The standard deviation displayed in Panel B (on a separate scale, on the right side of the graph) indicates that the dispersion in total weeks available was only slightly higher in the recent episode than in the preceding one, implying that there is a similar degree of cross-state variation used for our estimates in both episodes. Panel B shows a sharp drop in 2012 in the average number of weeks of UI for which unemployed individuals qualify, as implied by the discussion of the legislation in Appendix I.

Figure 2 illustrates the expansion of UI receipt during the recent recession and subsequent reduction as the labor market recovery has proceeded. The weekly flow of new UI claims peaked at about 660 thousand in early 2009 (slightly below the peak of nearly 700 thousand reached in late 1982; not shown). As of early 2013, new UI claims had declined to nearly their pre-recession level. In addition to the weekly flow of new UI claims, two series for the level of ongoing UI claims are displayed in figure 2: 1) regular UI claims (26 or fewer



Note: Panel B series calculated using monthly CPS observations (weighted) for UI-eligible unemployed individuals (see Section 3.1).

Figure 1: Variation in Total Weeks of UI Available

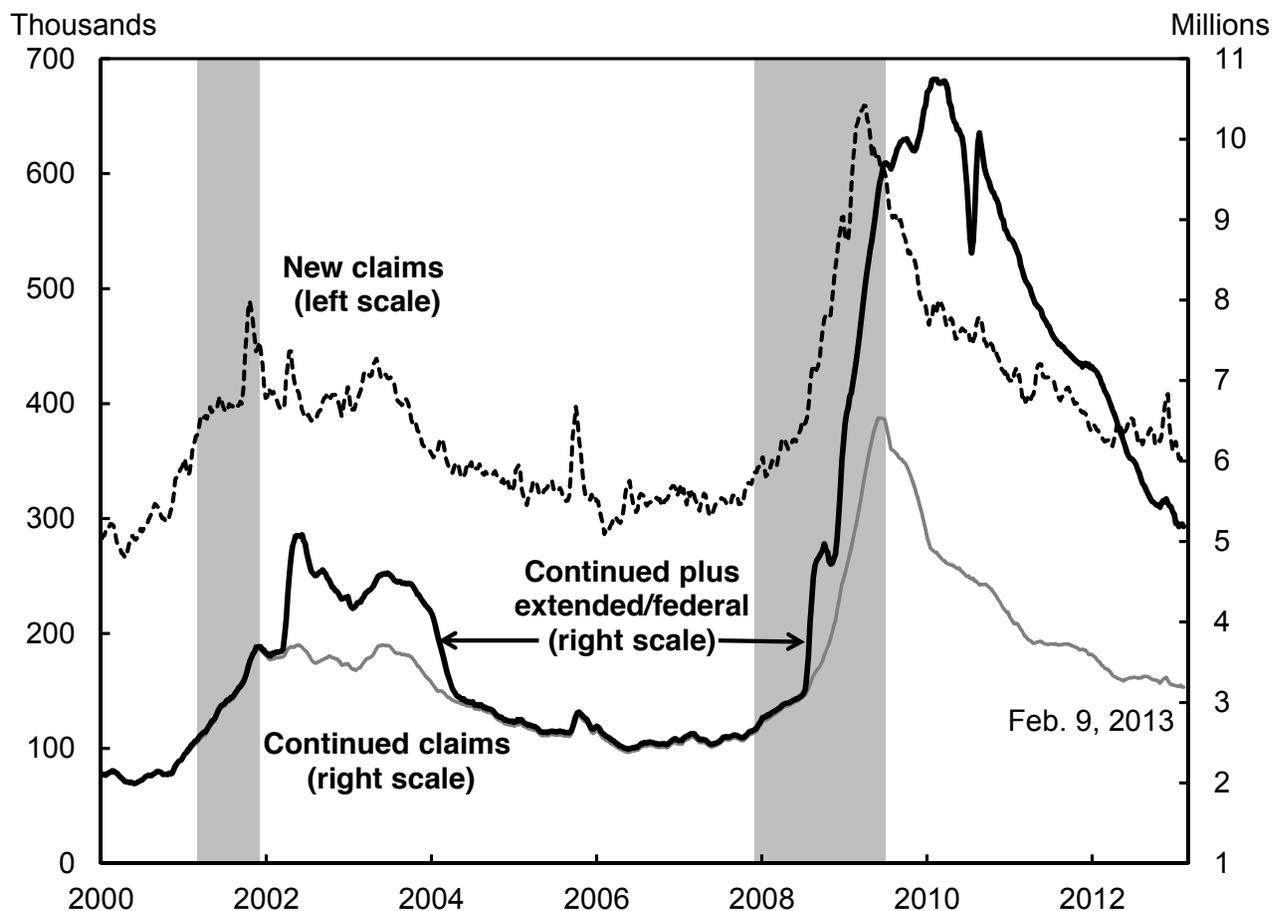


Figure 2: Unemployment Insurance Claims (Regular and Extended)

weeks) and 2) regular UI claims plus UI claims available through extensions. The level of both series has declined by about half since peaking in late 2009 and early 2010. The sharp, temporary drop in mid-2010 corresponds to the suspension period of June-July 2010. A much smaller but still substantial number of extended claimants also were present during the earlier episode of UI extensions during 2002-2004.

## 2.2 Past research on UI disincentive effects

Standard search models of unemployment imply that UI payments are likely to reduce job finding and prolong unemployment spells for eligible individuals, and a voluminous empirical literature exists that attempts to quantify the size of such effects. Much of this research has focused on unemployment exit rates around the time of UI benefit exhaustion, often using administrative data on UI recipients (e.g., Moffitt 1985, Katz and Meyer 1990, Meyer

1990, Card and Levine 2000, Jurajda and Tannery 2003). The findings from this research generally reveal the expected disincentive effects of UI availability on unemployment exits. Some analysts have relied on such estimates to simulate the likely effects of extended UI on unemployment durations and the unemployment rate during the recent downturn (e.g. Fujita 2010a, Mazumder 2011). However, the estimated magnitude of such effects has varied widely across different studies. Recent research suggests that the disincentive effects of extended UI are likely to be muted relative to the effect implied by earlier research based on the exhaustion of regular UI benefits. This may be due to the decreased reliance on temporary layoffs over the past few decades and the unusually weak labor market conditions prevailing during the recent extension episode (see e.g., Katz 2010, Landais, Michailat, and Saez 2010, and Kroft and Notowidigdo 2011).

Given the difficulty in relying on past research findings to assess disincentive effects of extended UI in the recent episode, some researchers have relied on concurrent labor market data to directly assess the impact of extended UI. These analyses have focused on comparing unemployment durations for unemployed individuals distinguished by their likely eligibility for UI (Valletta and Kuang 2010) and examining changes in transition rates out of unemployment measured across duration spans that are distinguished by their exposure to extended UI (e.g. Fujita 2010b, Howell and Azizoglu 2011). Such studies provide broad empirical guidance about the likely impacts of extended UI. However, they are not based on formal statistical analyses that attempt to isolate the direct effects of UI extensions while controlling for the effects of labor market conditions and individual characteristics that are related to extended UI eligibility and receipt. As such, it is difficult to conclude that the relationships between extended UI and exit from unemployment found in these papers are causal.

We build on the existing literature by exploiting directly variation in the duration of extended UI available in different months in different states. In addition, our matched CPS data enable us to incorporate detailed controls for state economic conditions and individual characteristics into the analyses. Rothstein (2011) conducted an investigation that is closely related to ours for the recent episode of extended benefits only.<sup>5</sup>

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<sup>5</sup> Our work goes beyond Rothstein's by including the earlier (2002-04) episode in our analyses and by covering the period of both the growth and rollback of extended UI availability in 2008 forward. We also rely on a different set of identification conditions and perform different simulations to assess the overall impact of the programs, as described in more detail below.

### 3 Econometric Framework and Identification

The rotation group structure of the CPS allows us to follow individuals for four consecutive months if they have not changed residence. Respondents are asked their labor force state (employed, unemployed, not in the labor force (NILF)) in each month, and the unemployed are asked how long they have been unemployed.<sup>6</sup> In this way, spells are “joined in progress,” leading to the classic problem of length-biased sampling. This produces a sample of longer spells than the overall distribution of unemployment spells, and our econometric model needs to account for this.<sup>7</sup>

We use a simple discrete-time hazard specification to model the probability that an unemployment spell ends at duration  $S$  given that it has lasted at least until  $S$ . This hazard function is  $h(S)$ , where  $h(\cdot)$  is a probability function (e.g., probit or logit) that will also depend on a set of individual and labor market characteristics as well as the duration of unemployment. Assuming independence across months, the unconditional probability that a spell of unemployment ends at duration  $S$  (to either employment or NILF) is

$$P(D = S) = h(S) \prod_{t=1}^{S-1} (1 - h(t)). \quad (1)$$

In the case where a spell remains in progress with duration  $S$  at the last survey in which it is observed, what is known is that the duration of the spell is at least  $S$ . The unconditional probability that a spell of unemployment has duration at least  $S$  (the survivor function) is

$$G(S) = P(D \geq S) = \prod_{t=1}^S (1 - h(t)). \quad (2)$$

The short-panel structure of the CPS implies that only unemployment spells lasting long enough to make it to the survey date are measured. Let  $S_0$  represent the duration of an unemployment spell when it is first observed in the CPS. Now suppose there are

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<sup>6</sup> The CPS documentation implies that the unemployed are asked their duration only in the first month they report being unemployed and that this duration is automatically incremented in subsequent months for which they report being unemployed. In fact, the sequences of spell durations are not this clean. In addition, as reported by Elsby, Hobijn, Sahin, and Valletta (2011), in the matched CPS data many individuals identified as newly unemployed in a month report durations of unemployment that substantially exceed one month.

<sup>7</sup> The durations-to-date of the spells in progress in a cross-section is a misleading guide to the duration of a randomly selected set of completed spells. Later, we use our estimates of the model of exit from unemployment to simulate the effect of extended benefits on the steady-state distribution of duration of spells in progress in a cross-section.

$n$  observations subsequent to the first observation of the unemployment spell where the individual remains unemployed or is first observed to have exited unemployment (either to employment or NILF). Given that spells are not observed unless they reach duration  $S_0$ , the appropriate conditional probability of a spell ending with duration  $S$  is

$$P(D = S | D \geq S_0) = \frac{h(S) \prod_{t=1}^{S-1} (1 - h(t))}{\prod_{t=1}^{S_0-1} (1 - h(t))} = h(S) \prod_{t=S_0}^{S-1} (1 - h(t)). \quad (3)$$

Analogously, the appropriate conditional probability for a spell that remains in progress with duration  $S$  when it is last observed is

$$P(D \geq S | D \geq S_0) = \frac{\prod_{t=1}^S (1 - h(t))}{\prod_{t=1}^{S_0-1} (1 - h(t))} = \prod_{t=S_0}^S (1 - h(t)). \quad (4)$$

These conditional probabilities appropriately account for the length-biased sampling problem and allow inference about the overall distribution of unemployment durations.

The likelihood function appropriate to this model is derived from equations 3 and 4. Assuming a standard normal CDF for the hazard probability, the result is a probit model where each monthly observation on an unemployment spell (matched to the succeeding month) contributes to the likelihood function. Monthly observations where the spell has not ended by the next month have a “zero” outcome with the probability specified in equation 4. Monthly observations where the spell has ended by the next month have a “one” outcome with the probability specified in equation 3. Each spell in the sample has at most one observation with a “one” outcome (the end of the spell).

A specification choice must be made regarding how to characterize the availability of extended benefits in the model. Based on past research findings regarding exhaustion spikes and our preliminary specification search, the approach we selected includes two indicator variables for the availability of unemployment insurance benefits to the worker.

1.  $EB_{it}$  – An indicator for availability of extended benefits at time  $t$  which equals one if 1) individual  $i$  has been unemployed for fewer months than the number of months of UI available (including extended benefits) in the relevant state and months and 2) some extended benefits are available in the relevant state and month.
2.  $Last_{it}$  – An indicator which equals one if individual  $i$  is in the last month of his/her UI availability at time  $t$ . This is meant to allow for a spike in the exit hazard at exhaustion.

The probit model is specified by assuming a spell ends in a given month if an unobserved latent variable for spell  $i$  in month  $t$  ( $y_{it}$ ) is positive. This latent variable is

$$y_{it} = X_{it}\beta + \delta_1 EB_{it} + \delta_2 Last_{it} + \epsilon_{it}, \quad (5)$$

where  $X_{it}$  is a vector of individual and economic variables,  $\beta$  is a vector of parameters,  $\delta_1$  is a parameter measuring the marginal effect of extended benefits on the hazard of an unemployment spell ending,  $\delta_2$  is a parameter measuring the marginal effect on the hazard of being in the last month of UI eligibility, and  $\epsilon_{it}$  is an error term with a standard normal distribution. The hazard of a spell ending is then

$$h(t) = P(Y_{it} > 0) = P(-\epsilon_{it} < X_{it}\beta + \delta_1 EB_{it} + \delta_2 Last_{it}) = \Phi(X_{it}\beta + \delta_1 EB_{it} + \delta_2 Last_{it}), \quad (6)$$

where  $\Phi(\cdot)$  is the standard normal cumulative distribution function.

The estimated model includes in the  $X$  vector a set of standard personal characteristics that are systematically related to labor market outcomes: 4 education categories, 5 age categories (covering the included ages 20-64), and indicators for female, married, female\*married, and nonwhite individuals. In order to account for local labor market conditions, the model includes a cubic in the monthly seasonally adjusted state unemployment rate and a cubic in the 3-month annualized growth in seasonally adjusted log non-farm payroll employment. To allow for a flexible baseline hazard and to account for the effects of normal UI benefits, the model also includes a set of indicators for the first 6 months of unemployment (0-6) and single indicators for months 7-9, months 10-12, and months 13-28 (9 categories in total). We also include a complete set of date (year-month) and state indicators, which provide additional controls for relative economic conditions that are shared across all states but vary over time and relative conditions that are state-specific and fixed over time.

### 3.1 A Competing Risk Model

We apply the probit model developed here to the probability that an individual exits unemployment, regardless of the subsequent labor force state. It is also interesting to explore how extended benefits affect both the probability of exiting unemployment to a job (employment) and probability of exiting unemployment to leave the labor force (cease searching). A competing risk framework is natural for this purpose and can be implemented as a straightforward generalization of the discrete choice hazard model outlined above. The key is to define cause-specific hazard functions and use them in the same likelihood context.

In the competing risk framework, there are two durations associated with each spell of unemployment:

1. Time until exit to employment. Call this time the UE duration.
2. Time until exit to NILF. Call this time the UN duration.

There are three types of spells in the data.

1. Spells that end in employment. For these spells, we observe the end of a UE spell. The UN spell is censored at the duration of the UE spell.
2. Spells that end in exit to NILF. For these spells, we observe the end of a UN spell. The UE spell is censored at the duration of the UN spell.
3. Spells that do not end during the observation period. For these spells, both the UE and UN durations are censored at the last observed unemployment duration.

We estimate three versions of the unemployment exit model: 1) exit to either employment or NILF, 2) exit to employment, and 3) exit to NILF in order to determine the effect of extended benefits on each of these exit rates.

### **3.2 Identifying the Effects of Extended Benefits**

Our hazard model of the exit from unemployment includes fixed effects for each state and each month, along with individual duration, demographic characteristics, and measures of local economic conditions. The effects of extended UI are identified by exploiting differences in UI availability at the individual level, which are determined by the comparison of individual unemployment duration with the number of weeks of UI available in that individual's state at the same point in time (measured at a monthly frequency). These effects are estimated conditional on individual characteristics and differences in local labor market conditions. Given that UI extensions and rollbacks are triggered by deterioration or improvement in state labor market conditions, including the measures of local economic conditions in the model is critical for this identification scheme.

Consider this specific example of identifying information. Between April and May of 2012, maximum available weeks of UI benefits fell from 99 weeks to 73 weeks in Washington State and from 99 weeks to 89 weeks in Oregon. This was due to legislative changes and idiosyncratic features of prior and contemporaneous labor market conditions in those states. During this time, the actual unemployment rates were stable at 8.4 percent in Washington State and 8.8 percent in Oregon. Consider an individual in each state who had been unemployed for 78 weeks (18 months) as of April 2012 and who were observationally equivalent

to each other more generally. In our model, both individuals have UI available to them in April 2012, because their unemployment duration is less than their maximum eligibility in that month. If they remain unemployed in the subsequent month, however, their unemployment duration will reach 82-83 weeks. At that time, the individual in Oregon will still have UI available, while the individual in Washington State will not. Our model will uncover an increase in unemployment duration due to extended UI if the individual in Washington State is more likely to exit unemployment between April and May of 2012 than is the individual in Oregon (conditional on their observational equivalence and unrestricted state and time effects). In this specific example, the difference in unemployment rates between the two states is small, and its impact on job prospects and search behavior for the two individuals is captured in our model by the local labor market controls (polynomials in the state unemployment rate and rate of payroll job growth).

It is important to note that, because our empirical models include indicator variables for each of the first six months of unemployment, the EB availability indicator captures the effects of extended benefits only for individuals unemployed for at least six months. In other words, the specification excludes any potential effects of EB availability on search behavior for the short-term unemployed. Our specification was chosen to capture the rapid decline in unemployment exit rates during the first six months of unemployment and to avoid the estimation biases that might result from a parametric specification of duration dependence. The resulting restriction is innocuous: Rothstein (2011) found that the recent episode of extended benefits had meaningful effects on exit rates only for individuals unemployed for at least 6 months. More generally, our specification enables separate estimation of the effects of UI remaining and unemployment duration at the individual level.

## 4 Sample Definition and Data Issues

We use CPS data from January 2000 to December 2005 (2000m1-2005m12) and from January 2007 to December 2012 (2007m1-2012m12) for individuals ages 20-64 to examine the effects of the two episodes of extended benefits in the 2000s.

### 4.1 Defining the Relevant Sample: UI-Eligibility

An appropriate sample for analysis of the effect of UI benefits is a sample of unemployed individuals who are eligible to receive UI. However, there is no direct information in the CPS on UI eligibility, and we rely on a proxy measure based on the reported reason for

unemployment.

Unemployed individuals who report having lost a job as the reason for unemployment are, in principle, eligible to receive UI. In contrast, individuals who report voluntarily leaving a job or new- or re-entry into the labor force as the reason for unemployment are, in principle, not eligible to receive UI. While losing a job is a necessary condition for being eligible to receive UI benefits, it is not sufficient. For example, a worker who reports a job loss may not have sufficient previous employment experience to qualify for unemployment insurance or may have been fired for cause.<sup>8</sup>

Because we do not have the detailed work-history information needed to impute eligibility, we proceed by classifying unemployed job losers as the UI-eligible group that we use in our analysis. Job leavers and new labor force entrants are classified as UI-ineligible and are not included in our analysis. Later, we present a “placebo” analysis of the effect of extended benefits on exit from unemployment for the UI-ineligible sample (job leavers and new entrants to the labor force). We expect the effects of extended UI for this group to be smaller than for UI eligible individuals. However, if there is variation in economic conditions across states over time that is correlated with the availability of extended benefits but that is not accounted for by the variables in the model, we could find that exits from unemployment for the ineligible group is related to extended UI simply due to this omitted variable problem. Such a finding would imply that our estimates of the effect of extended UI on the UI-eligible group may be too large and would represent an upper bound on the true effect. Working in the opposite direction, there may be spillover effects on job search and job finding from eligible to ineligible individuals that would show up as a positive relationship between exit from unemployment and extended UI for the UI-ineligible group (Levine, 1993).<sup>9</sup>

Table 1 provides supporting evidence for our working definition of UI eligibility. Each March, the regular monthly CPS is accompanied by an extensive set of supplemental questions regarding income receipt in the prior calendar year; UI is separately identified as an individual income source. The rotating sample structure of the CPS enables matching of observations on unemployed individuals for selected months in year  $t$  with the information

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<sup>8</sup> Another consideration is that an eligible worker may choose not to take up benefits. However, the decision to take up benefits may be affected by the structure of the UI program, including the availability of extended benefits, so that it would not be appropriate to restrict the analysis only to those eligible workers who choose to take up benefits. And since we have no data in the monthly CPS that would allow us to determine actual UI receipt, we cannot make this sample distinction in any case.

<sup>9</sup> Levine’s empirical results suggest that the UI-induced reduction in job finding within the group of UI-eligibles increases job availability and hence the job-finding rate for ineligibles.

Table 1: UI Income Receipt, by Eligibility Status  
(unemployed in base month, matched to subsequent March survey)

<u>Income year</u>	UI status:	<u>March to March</u>		<u>December to March</u>	
		<u>Eligible</u>	<u>Ineligible</u>	<u>Eligible</u>	<u>Ineligible</u>
2005		0.298	0.052	0.264	0.049
2006		0.281	0.038	0.335	0.054
2007		0.299	0.052	0.297	0.074
2008		0.319	0.088	0.438	0.068
2009		0.490	0.069	0.505	0.144
2010		0.474	0.117	0.456	0.122
2011		0.401	0.055	0.353	0.087

**Note:** Fraction of unemployed individuals from monthly CPS (as indicated) who report receiving annual UI income in subsequent matched March CPS (weighted). Eligibles are job losers/layoff, ineligible are quits/entrants.

on their income receipt in year  $t$  recorded in the year  $t + 1$  March supplement. Based on this match, table 1 provides a breakdown of UI income receipt (percent reporting positive UI income) by measured eligibility status for individuals who are unemployed in March or December of the calendar year corresponding to reported income receipt, for income years 2005-2011. The percentage of the UI-eligible who report receiving UI income reached about 50 percent in 2009 and has declined somewhat since then. The percentage of the UI-ineligible who report receiving UI income is much lower, usually at about 5-10 percent, with a few higher values recorded. Because individuals may be subject to multiple unemployment spells over a calendar year, an unknown proportion of individuals identified as ineligible in a particular month may have received UI income based on a separate spell of unemployment that year, for which they were eligible for UI receipt based on their stated reason for unemployment. On balance, we interpret these figures as suggesting that our eligibility indicator is strongly correlated with actual eligibility status, although take-up of unemployment insurance by those eligible is clearly not universal.<sup>10</sup>

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<sup>10</sup> See Blank and Card (1991) and Anderson and Meyer (1997) for earlier studies of incomplete take-up rates for unemployment insurance.

## 4.2 Sample Construction: Matching the Monthly CPS

The rotation group structure of the CPS visits a given address (housing unit) for four months, does not interview for eight months, and revisits the address for four more months. In other words, over a 16 month period, the household is surveyed for four months, left alone for eight months, and surveyed for four more months. This sample structure allows us to match households in consecutive months up to three times. Failures to match primarily occur when a household moves to a new housing unit between interviews. This generally occurs less than five percent of the time.

One key concern with regard to use of the matched data is the likelihood of spurious transitions in labor force status, particularly spurious exits from unemployment, that can lead to an overestimate of the probability of exit from unemployment (e.g., Abowd and Zellner, 1985; Poterba and Summers, 1986, 1995). As demonstrated by Rothstein (2011), the unemployment durations implied by the frequency and duration structure of exits from unemployment are much lower than the reported durations of spells in progress in the cross-section.<sup>11</sup> This difference in distributions is due in part to the length-biased sampling inherent in an analysis of durations of spells in progress in a cross-section. But the difference is also the result of the likely presence of spurious transitions. These spurious transitions are potentially problematic for the estimation of our models, since errors in the identification of exit from unemployment could seriously bias the estimates of the effects of extended benefits.

Following Rothstein (2011), we address this problem through direct adjustments to reported transitions following specific patterns that are indicative of reporting errors. In particular, for individuals who report a transition out of unemployment in month one followed immediately by an entry back to unemployment in month three, we recode the intervening month as a continuation of the initial unemployment spell. We do this whether the unemployment exit is due to job finding or labor force withdrawal. In other words, letting U represent unemployment, E employment, and N out of the labor force, we recode 3-month transition patterns of UEU and UNU to UUU, and we retain both of the resulting UU "transitions" in the matched data. We recoded 3,016 UEU transitions and 7,119 UNU transitions (a total of 10,135 transitions) in our estimation samples in this way.<sup>12</sup>

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<sup>11</sup> We confirmed that this is not due to the construction of the matched CPS sample. The distribution of reported durations from the sample, treated as a series of cross-sections, is essentially identical to the distribution from the complete CPS cross sections.

<sup>12</sup> Of these 10,135 recoded transitions, 5,008 were for UI-eligible spells and 5,127 were for UI-ineligible spells.

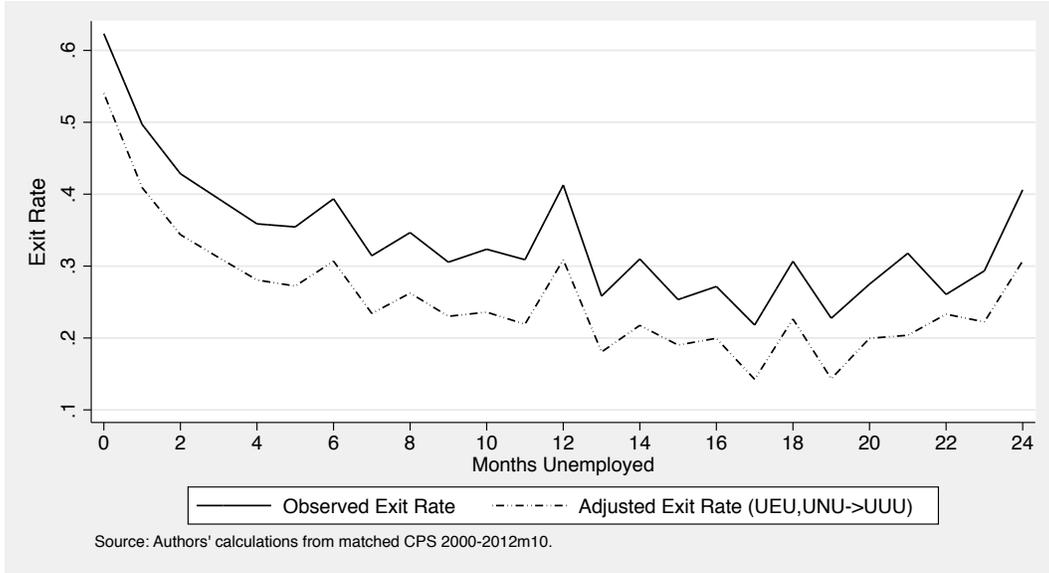


Figure 3: Exit Rate from Unemployment, Observed and Adjusted (UEU, UNU → UUU)

Some support for the idea that UEU and UNU transitions are likely to be spurious is that reported unemployment durations in the third month of observed UEU and UNU transitions is, on average, far greater than one month (average of 5.7 months, median of 3 months). As noted by Elsbey et al. (2011), it is likely that individuals' reported unemployment duration reflects the time elapsed since the loss of a salient job, which is likely the one that enabled them to qualify for UI benefits. The reported durations are therefore much more likely to capture the duration of a spell of ongoing UI receipt than are the durations implied by reported unemployment exits, which may reflect stopgap jobs and temporary labor force withdrawals in addition to direct reporting errors (see also Poterba and Summers, 1995).

Figure 3 contains a plot of average (over the 2000-2012 period) exit rates from unemployment by duration of unemployment. Two exit rates are presented: 1) the observed exit rate and 2) the exit rate adjusted by recoding UNU and UEU sequences to UUU (no exit). Clearly, the adjustment substantially reduces the exit rates from unemployment, implying an increase in the survivor function and associated unemployment durations.<sup>13</sup>

Imposing this adjustment to observed transitions requires restricting the set of observations we use to those from the first two of each set of four consecutive CPS rotation groups, so that we have at least two subsequent matched observations. We impose this restriction

<sup>13</sup> Rothstein (2011) presents a graph (Figure 7, page 182) of the Kaplan-Meier survivor functions with and without recoding of the UEU and UNU transitions to UUU.

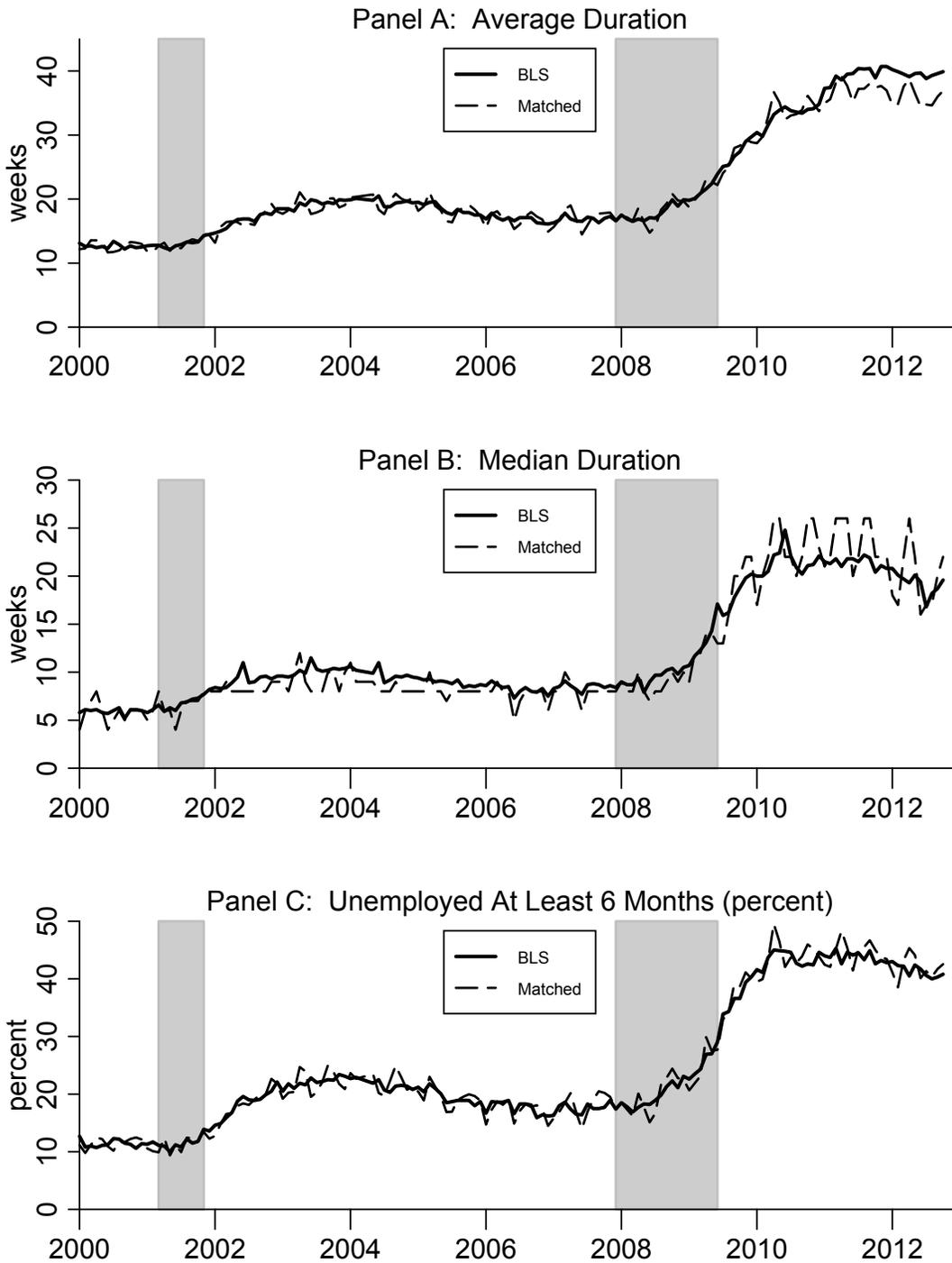
Table 2: Sample Breakdown by Eligibility for Unemployment Insurance

2000-2005				
Sample	Spells	End in UE	End in UN	Censored
UI-Eligible	39,155	13,499	5,775	19,881
UI-Ineligible	41,810	11,666	13,233	16,911
All Spells	80,965	25,165	19,008	36,792
2007-2012m10				
Sample	Spells	End in UE	End in UN	Censored
UI-Eligible	59,157	15,387	7,978	35,792
UI-Ineligible	46,747	9,554	14,580	22,613
All Spells	105,904	24,941	22,558	58,405

and the transition adjustment for all of our analysis samples. As such, although we have CPS data through December 2012, our final observation is for October 2012. In addition, to ensure valid matches of individuals across months, we dropped a small number of observations for which reported age, gender, race, and educational attainment are not consistent across months (i.e., age changes by more than 1 year, etc.).

Table 2 contains, for each period and by eligibility status, counts of the number of spells that end with employment, end with NILF, and are censored. When considering the competing risk model, 1) the number of censored UE spells is the sum of the number of spells that end in NILF and the number that do not end within the observation period, and 2) the number of censored UN spells is the sum of the number of spells that end in employment and the number that do not end within the observation period.

Our matched CPS sample covering the 2000m1-2005m12 period contains 109,190 monthly observations on 80,965 spells of unemployment for workers ages 20-64. The analogous sample for the 2007m1-2012m10 period contains 151,189 monthly observations on 105,904 spells of unemployment. In order to examine the representativeness of our matched sample, we compared the reported unemployment duration in our sample, based on the survey responses regarding in-progress spells, with published statistics on unemployment duration from the U.S. Bureau of Labor Statistics (BLS). Figure 4 contains plots of three statistics from the two distributions of unemployment durations. These are mean duration, median duration, and the fraction long-term unemployed (at least six months). While the BLS series is seasonally adjusted and our series is not, it is clear that there is little difference between our sample



Note: From U.S. BLS (seasonally adjusted) and authors' tabulations of matched CPS data (weighted). Based on reported duration of in-progress spells. Gray areas denote NBER recession dates.

Figure 4: Unemployment Duration, BLS and Matched CPS Data

Table 3: Unemployment Survivor Rates, by Duration in Months (UI Eligible Sample)

Months Duration	(1) 2000-2005	(2) 2002m3-2004m2	(3) 2007-2012m10	(4) 2009-2011
1	0.474	0.499	0.501	0.523
2	0.296	0.328	0.337	0.366
3	0.202	0.233	0.246	0.276
4	0.143	0.166	0.185	0.215
5	0.105	0.127	0.143	0.168
6	0.077	0.099	0.113	0.137
7	0.053	0.074	0.088	0.111
8	0.041	0.059	0.072	0.092
9	0.030	0.044	0.059	0.076
10	0.023	0.035	0.048	0.062
11	0.018	0.027	0.040	0.053
12	0.013	0.021	0.033	0.044

Note: Authors' calculations from matched CPS data (weighted).

of unemployment durations from the matched CPS and the BLS reported statistics from all sampled unemployment spells in the CPS.

Figure 4 also provides a sense of changes in unemployment duration over the 2000-2012 period. The sharp increase in duration during and after the most recent recession is evident in all panels, with each of the duration measures increasing by about double or more relative to their pre-recession values.

The unemployment durations in figure 4 come from the cross-sectional distribution of spells in progress. As noted earlier, due to the length-biased sampling problem, this distribution is likely to overstate the duration distribution for all spells of unemployment. This is implied by the tabulations of survivor rates of unemployment spells in table 3, for which we used our matched data to calculate survivor rates across months during the first year of unemployment. This table displays the tabulations separately for our two estimation samples (2000-2005 in column 1 and 2007-2012m10 in column 3). In order to highlight the higher unemployment survivor rates in the weakest labor market periods, the table also displays survivor rates for sub-periods with availability of extended benefits and relatively high unemployment rates (2002m3-2004m2 in column 2 for the earlier period and 2009-2011 in column 4 for the later period). The exit rates from unemployment are sufficiently high that only a small fraction of individuals remain unemployed after the first six months. In the weak labor market from 2002-2004, only about 10 percent of unemployment spells for job

losers (our UI-eligible sample) lasted past 6 months. In the very weak labor market from 2009-2011, 13.7 percent of job losers remained unemployed for at least 6 months.

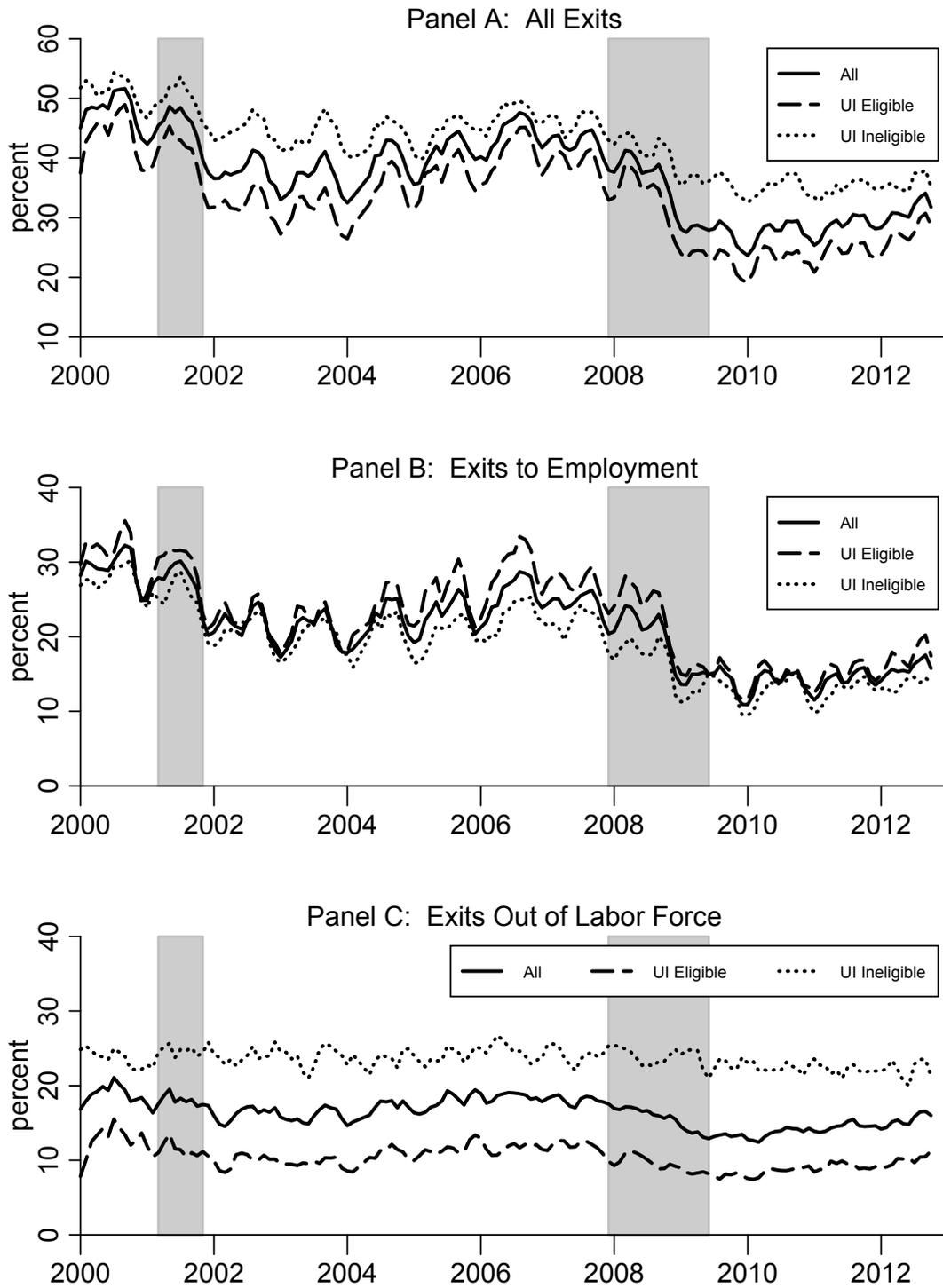
On inspection, it appears that the survivor rates presented in table 3 are not consistent with the cross-section distribution of durations of incomplete spells calculated directly from the CPS. For example, in the 2009-2011 period, the survival rate at 6 months 13.7 percent while 39.8 percent of spells in progress in the cross-section in this period were at least 6 months. This is largely a result of the length biased sampling built into the cross-section that results in over-sampling of long spells. Later we present estimates of the cross-section distribution of spells in progress implied by our model of exits that reconciles much of this difference. Our analysis highlights the inappropriateness of making inferences about the distribution of completed spells from the cross-section distribution of spells in progress.

Figure 5 complements table 3 by showing monthly exit rates from unemployment over our complete sample tabulated for all exits in Panel A and exits by type (to employment or out of the labor force) in Panels B and C. In each case, we display the exit rates for all unemployed individuals and also for UI eligible and ineligible individuals separately. A sharp decline in exit rates, particularly exits to employment, is evident during the recent recession, with a minor rebound evident beginning in 2010. Overall exit rates are higher for UI-ineligible individuals than for eligible individuals, which reflects the large gap between the two groups for exits out of the labor force (Panel C). The rates of labor force exit from unemployment exhibit very little cyclicalicity, with only a slight net decline evident during the recent recession and essentially no cyclicalicity evident for UI ineligible.

## 5 Estimation of the Probit Model of Exit

We present estimates of the probit model, specified in equations 5 and 6, of the probability that an unemployment spell ends in a given month. The key parameters we estimate representing the effect of extended benefits on the unemployment exit probability are  $\hat{\delta}_1$  (the coefficient on the *EB* indicator) and  $\hat{\delta}_2$  (the coefficient on the *Last* indicator for the final month of UI eligibility). As specified in equation 5, the underlying estimated probit parameter on *EB* is  $\hat{\delta}_1$ , and the average marginal effect of *EB* on the probability of exit from unemployment is computed from this as

$$\hat{\delta}_1^* = \hat{\delta}_1 \frac{1}{N} \sum_{i,t} \phi(X_{it}\hat{\beta} + \hat{\delta}_1 EB_{it} + \hat{\delta}_2 Last_{it}), \quad (7)$$



Note: Authors' tabulations from matched CPS data (weighted), expressed as 3-month moving averages. Gray areas denote NBER recession dates.

Figure 5: Unemployment Exit Rates, Matched CPS

Table 4: Estimated Average Marginal Effects on Probability of Exit from Unemployment  
 UI Eligible Sample

Model	2000-2005m2		2007-2012m10	
	$\hat{\delta}_1^*$	$\hat{\delta}_2^*$	$\hat{\delta}_1^*$	$\hat{\delta}_2^*$
Single Risk	-0.0583 (0.0138)	0.0538 (0.0156)	-0.0500 (0.0064)	0.0220 (0.0199)
Exit to Emp	-0.0212 (0.0121)	0.0263 (0.0150)	-0.0099 (0.0065)	0.0208 (0.0129)
Exit to NILF	-0.0372 (0.0106)	0.0287 (0.0098)	-0.0340 (0.0033)	0.0040 (0.0109)

Note:  $\hat{\delta}_1^*$  is the average marginal effect on the exit probability of the indicator for availability of extended benefits.  $\hat{\delta}_2^*$  is the average marginal effect of the indicator for the last month of availability of benefits. These are calculated using equations 7 and 8. The probit model also includes controls for 4 education categories, 5 age categories (covering the included ages 20-64), female, married, female\*married, nonwhite, year-month indicators, state indicators, a cubic in the monthly seasonally adjusted state unemployment rate, a cubic in the 3-month annualized growth in seasonally adjusted log non-farm payroll employment, a set of indicators for the first 6 months of unemployment (0-6) and single indicators for months 7-9, months 10-12, and months 13-28 (9 categories for duration in total). The estimates are weighted by the CPS sampling weights, and robust asymptotic standard errors clustered by state are in parentheses. The sample for 2000-2005m12 includes 44,367 matched monthly observations on 31,925 spells of unemployment for job losers. The sample for 2007-2012m10 includes 81,472 matched monthly observations on 54,928 spells of unemployment.

where  $\phi(\cdot)$  is the standard normal probability density function and  $N$  is the sample size. Analogously, the average marginal effect of exhaustion of UI (indicated by  $Last$ ) on the probability of exit is

$$\hat{\delta}_2^* = \hat{\delta}_2 \frac{1}{N} \sum_{i,t} \phi(X_{it}\hat{\beta} + \hat{\delta}_1 EB_{it} + \hat{\delta}_2 Last_{it}). \quad (8)$$

Note that both of these marginal effects could be important in measuring the effect of extended benefits because extended benefits increase the number of periods for which an individual can receive UI and any spike in the exit probability at exhaustion is pushed further into the unemployment spell when extended benefits are available.

Estimates of the key parameters from six versions of the probit model of exit from unemployment are presented in table 4. There are three models for each of the two time periods (2000-2005 and 2007-2012m3) with a period of extended benefits. Within each time period, there is a model of exit from unemployment and two models representing the competing

risks of exit to employment and exit to NILF. There is a clear pattern to the estimates. In both time periods and in the single risk and in the competing risk model for exit from the labor force, the availability of extended benefits has a substantial and precisely estimated negative effect on the probability of exit from unemployment (indicated by  $\hat{\delta}_1^*$ ). These effects are comparable across the two time periods, and they are primarily driven by the effects of extended benefits on labor force exits. The effect of extended benefits on the probability of exit to employment (the job-finding rate) is much smaller and less precisely estimated than is the effect on labor force exits, although the job-finding effect is marginally significant for the earlier sample period (p-value of about .08). This suggests that the negative effect of extended benefits on the exit rate from unemployment is driven largely by individuals staying in the labor force longer, perhaps to collect benefits, rather than by individuals reducing search effort and taking longer to find jobs. This pattern is largely consistent across the two time periods.

The evidence in table 4 on the effect of exhaustion of benefits on exit from unemployment is mixed. This is represented by  $\hat{\delta}_2^*$ , the average marginal effect of being in the last month of UI eligibility on the probability of exit. In the earlier time period, the exit rate from unemployment and the exit to NILF are significantly higher in the last month of eligibility for UI. This is another mechanism through which extended benefits can increase the duration of unemployment spells. The availability of extended benefits pushes the exhaustion spike deeper into the spell of unemployment. There is no significant effect of exhaustion of benefits on exit to employment (the job-finding rate) in the earlier period. The exhaustion of benefits does not have a statistically significant effect on any of the exit measures in the later period.

We infer from the estimated coefficients in Table 4 that the main effects of extended UI are on labor force attachment rather than job finding. However, assessment of the complete magnitudes of these effects requires a simulation that takes the full set of coefficients into account. We discuss those results below, in Section 6. Our simulations will rely on the complete set of estimated coefficients for the effects of extended UI, without exclusion of any coefficients based on pre-set levels of statistical significance. We first discuss the results from a placebo test based on a sample of UI-ineligible individuals.

## 5.1 A Placebo Test: UI-Ineligible Spells

We defined the UI eligible group to be those who report a job loss as their reason for unemployment. The remaining unemployed report being a job leaver (quit) or a labor force entrant (new entry or re-entry) as their reason for unemployment, and we classify these

individuals as UI ineligible. While this classification scheme is not perfect, we presented evidence in table 1 based on the March CPS that only a small fraction of job leavers and new entrants report having received UI. This group should be largely unaffected by the availability of extended benefits. On this basis, we re-estimate our probit model of exit from unemployment on samples of UI ineligible unemployed individuals in order to provide a placebo test of our estimation strategy. If we find effects of extended benefits on exit from unemployment for the ineligible that are similar to those we present in table 4, it could be that we have not adequately controlled for state/month specific economic conditions and that our estimates of the effect of extended benefits are too large. A potential factor working in the opposite direction is that there may be a positive effect of extended benefits on exit from unemployment among the UI-ineligible resulting from spillovers from the eligible to the ineligible. As we noted earlier, the reasoning is that, if extended benefits reduce the job finding rate among the UI-eligible, there may be improved job opportunities for the UI-ineligible that increase their exit rate from unemployment (Levine, 1993).

Table 5 contains the estimates of the key UI parameters ( $\delta_1^*$  and  $\delta_2^*$ ) estimated for the sample of UI ineligible spells. In the earlier period, we estimate there to be no effect of either the availability of extended benefits or the exhaustion of benefits on the exit rate in either the single risk or the competing risks model. In the later period, we do find a statistically significant negative effect of the availability of extended benefits on the unemployment exit rate in the single risk model and in the exit-to-NILF model. These estimates suggest that our estimates for the UI eligible sample in table 4 may be biased upward.

Interestingly, our estimates show a marginally significant (though small) increase in the exit rate to employment due to extended benefits for the UI ineligible individuals in the 2007-2012 period. This may reflect a spillover from those eligible for UI to those not eligible for UI. If more generous extended benefits lead to longer durations of unemployment for those eligible, this may increase opportunities for those who are not eligible.<sup>14</sup>

Our conclusion from the placebo exercise is that in the earlier period we appear to have adequately controlled for labor market conditions while in the later period it may be the case that the availability of extended benefits is correlated with unmeasured (unfavorable) economic conditions. As such, our estimates of the effect of extended benefits on exit from unemployment in the later period could be considered an upper bound on the true effect.

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<sup>14</sup> Levine (1993) makes this argument and presents evidence on this point.

Table 5: Estimated Average Marginal Effects on Probability of Exit from Unemployment  
 UI Ineligible (Placebo) Sample

Model	2000-2005m2		2007-2012m10	
	$\hat{\delta}_1^*$	$\hat{\delta}_2^*$	$\hat{\delta}_1^*$	$\hat{\delta}_2^*$
Single Risk	-0.0053 (0.0195)	0.0138 (0.0252)	-0.0320 (0.0098)	-0.0065 (0.0330)
Exit to Emp	0.0034 (0.0184)	-0.0080 (0.0208)	0.0152 (0.0084)	-0.0181 (0.0184)
Exit to NILF	-0.0094 (0.0189)	0.0251 (0.0206)	-0.0466 (0.0101)	0.0123 (0.0243)

Note:  $\hat{\delta}_1^*$  is the average marginal effect on the exit probability of the indicator for availability of extended benefits.  $\hat{\delta}_2^*$  is the average marginal effect of the indicator for the last month of availability of benefits. See note to table 4 for details. The sample for 2000-2005m12 includes 30,722 matched monthly observations on 23,445 spells of unemployment. The sample for 2007-2012m10 includes 44,590 matched monthly observations on 32,290 spells of unemployment for job leavers and new entrants to the labor force.

## 6 How Large is the Effect of Extended Benefits?

In order to quantify the effect of extended benefits on unemployment duration, we use our estimates to calculate the distribution of duration of unemployment spells under a set of three alternative scenarios regarding the availability of extended benefits. We select these alternative scenarios to facilitate comparison across the two episodes of extended benefits we study of 1) the effects of the observed extended benefit programs and 2) the potential effects of comparably-sized extended benefit programs. The three scenarios are

1. Observed-EB: An baseline scenario that uses the actual availability of extended benefits in each state in a given month. This is meant to provide a reference prediction of the survivor function and the expected duration.
2. No-EB: A scenario that assumes no extended benefits were available in any state in any month, regardless of local labor market conditions. In comparison with the baseline, this provides a measure of the extent to which the actual program of extended benefits affected the distribution of duration of unemployment spells.
3. Full-EB: A scenario that assumes that 99 weeks of extended benefits were available in all states and months during the hypothetical spells. In comparison with the no-extended-benefits scenario, this provides a measure of the extent to which a universal

unemployment insurance program offering 99 weeks of benefits would affect the distribution of duration of unemployment spells. Without this comparison, it would be difficult to compare the 2002-2004 with the 2008 and later experiences with extended benefits because extended benefits in the earlier period were much less widespread.

For each scenario we estimate the expected duration of unemployment spells ( $E(D)$ ) and various quantiles of the distribution of unemployment durations (the complement of the survivor function). We also use our estimates to calculate estimates of the distribution of unemployment durations that would be observed in a cross-section in a steady state and how this steady state distribution is affected by the availability of extended benefits.

We begin by creating a set of individuals starting unemployment spells based on the characteristics of unemployed workers over the 2000m1-2012m10 period. We create a spell for each unemployed job loser (the “eligible” group) aged 20-64. This set of 96,575 spells reflects the wide distribution of observable characteristics among the unemployed including demographics, human capital, and state of residence. We use this set of individuals to create two hypothetical sets of unemployment spells that cover the two periods of extended benefits in our sample:

1. March 2002 – June 2004. This period covers the early period of extended UI benefits (running from March 2002 through February 2004).<sup>15</sup>
2. January 2009 – April 2011. This period covers the end of the Great Recession and the immediate post-recession period. Extended benefits were generally available at a very high level throughout this period. We consider this period because the labor market generally lags the NBER business cycle dates. For example, the peak unemployment rate since 2007 was in 2009q4, while the NBER dated the end of the Great Recession as 2009q2.<sup>16</sup>

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<sup>15</sup> There were also extended benefits available in Alaska in March and April 2005 due to a weak seasonal labor market and in Louisiana October 2005 through January 2006 due to Hurricane Katrina. These are included in the estimation of the model for the earlier period (2000-2005), but no effort is made to quantify the effects of these small episodes.

<sup>16</sup> We also investigated hypothetical spells for two other time periods related to the great recession. One ran from July 2008 – October 2010. Given the fact that the labor market lags recession timing, this turns out to be a bit early to see the very low unemployment exit rates characteristic of the weak labor market subsequent to the Great Recession. Another ran from July 2010 – October 2012. Given that the maximum extended benefits only began to phase out late in this period, the effects of extended benefits on the duration distribution are virtually identical to those for the January 2009 – April 2011 period.

For each of the three scenarios regarding extended benefits described above and based on the estimates from the relevant probit model of unemployment exit, we predict for each spell the monthly hazard of exit from unemployment for each month for the first 28 months of the spell.<sup>17</sup> We use these predicted hazards to estimate the expected duration and the survivor function of each spell, and we present the average across spells of these quantities.

The estimated survivor function of spell  $i$  at duration  $t$  is

$$\hat{G}_i(t) = \prod_{s=1}^t (1 - \hat{h}_i(s)), \quad (9)$$

where  $\hat{h}_i(s)$  is the estimated unemployment exit probability for individual  $i$  in month  $s$ . In order to compute the expected spell duration, we need to assume something specific about the distribution of long spells. We assume that the monthly hazard of a spell ending after month 28 is constant for each individual at the average value for that individual of the hazard from months 24-28. The constant hazard feature of the conditional distribution of spells longer than 28 months implies that the conditional distribution is exponential and that the expected duration of spells from this point is simply the inverse of the constant hazard. On this basis, the expected duration of each spell is

$$E(D_i) = \left[ \sum_{s=1}^{28} s \hat{h}_i(s) \hat{G}_i(s-1) \right] + \hat{G}_i(28) \frac{1}{\bar{h}_i}, \quad (10)$$

where  $\bar{h}_i$  is the average across months 24-28 of the estimated monthly hazard of the unemployment spell of individual  $i$  ending.

## 6.1 The Effect of Extended Benefits on the Duration Distribution

Figure 6 contains plots of the inverse of the CDF of unemployment duration for the sets of hypothetical spells of unemployment corresponding to the weakest labor market in the early 2000s (spells starting in 2002) and the weakest labor market later in the decade (spells starting in 2009).<sup>18</sup> In other words, these plots show the number of months required to reach a given quantile of the duration distribution (time-to-quantile).

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<sup>17</sup> We use the exit model estimated over the 2000-2005m12 period for the March 2002 – June 2004 hypothetical spells. We use the model estimated over the 2007-2012m10 period for the later spells. See table 4.

<sup>18</sup> The inverse CDF plots months of unemployment on the vertical axis against quantiles of the duration distribution on the horizontal axis.

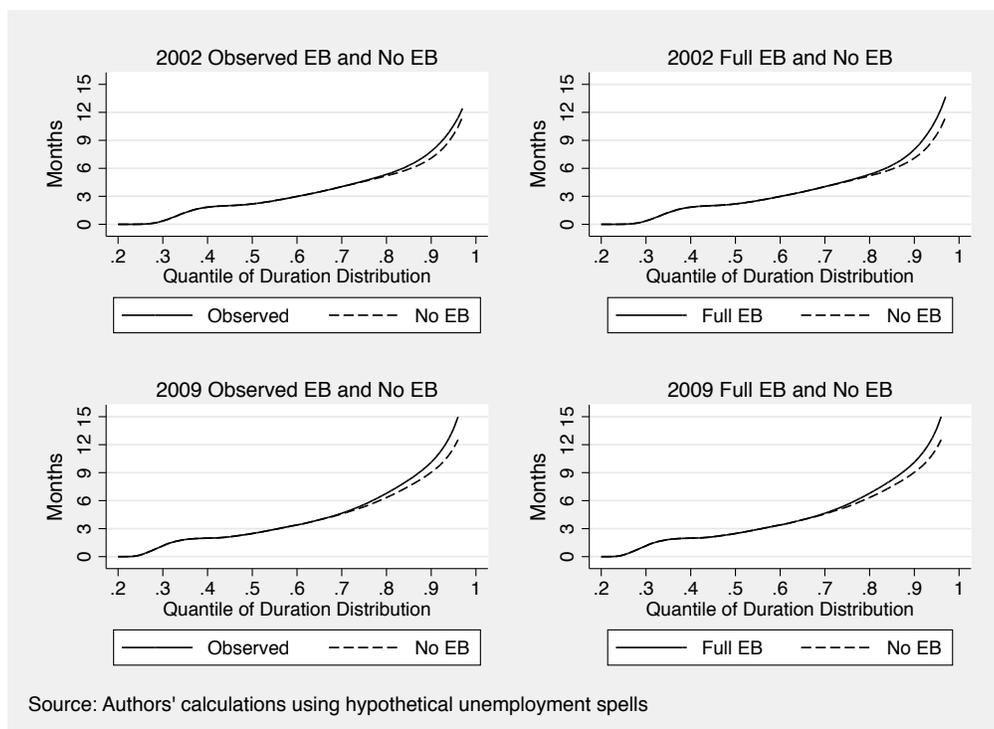


Figure 6: Comparisons of Time-to-Quantile: Observed-EB, No-EB, and Full-EB Scenarios.

Each panel of figure 6 presents a comparison of the inverse CDF of durations for two scenarios. The first row of the figure shows comparisons for the hypothetical spells beginning in 2002 while the second row shows comparisons for hypothetical spells beginning in 2009. The left panel in each row shows a comparison of the Observed-EB and No-EB scenarios, while the right panel in each row shows a comparison of the Full-EB and No-EB scenarios.

In neither case is there a difference in time-to-quantile for quantiles below 0.65 or so. This is because this quantile is reached well before 6 months and extended benefits do not have a measurable effect on exit so early in spells. At higher quantiles for spells beginning in 2002, the Observed-EB - No-EB difference in time-to-quantile is quite small while the Full-EB - No-EB difference is somewhat larger. This reflects the fact, shown in figure 1, that extended benefits were relatively less generous in the 2002-2004 period, so that the No-EB scenario is much closer to the observed EB scenario than to the Full-EB scenario. The plots in the second row of figure 6 for spells beginning in 2009 show an analogous pattern. At higher quantiles, the Observed-EB - No-EB difference in time-to-quantile is similar in magnitude to the Full-EB - No-EB difference in time-to-quantile. This reflects the fact that 99 weeks of extended benefits were almost universally available in the 2009-2011 period, so that the

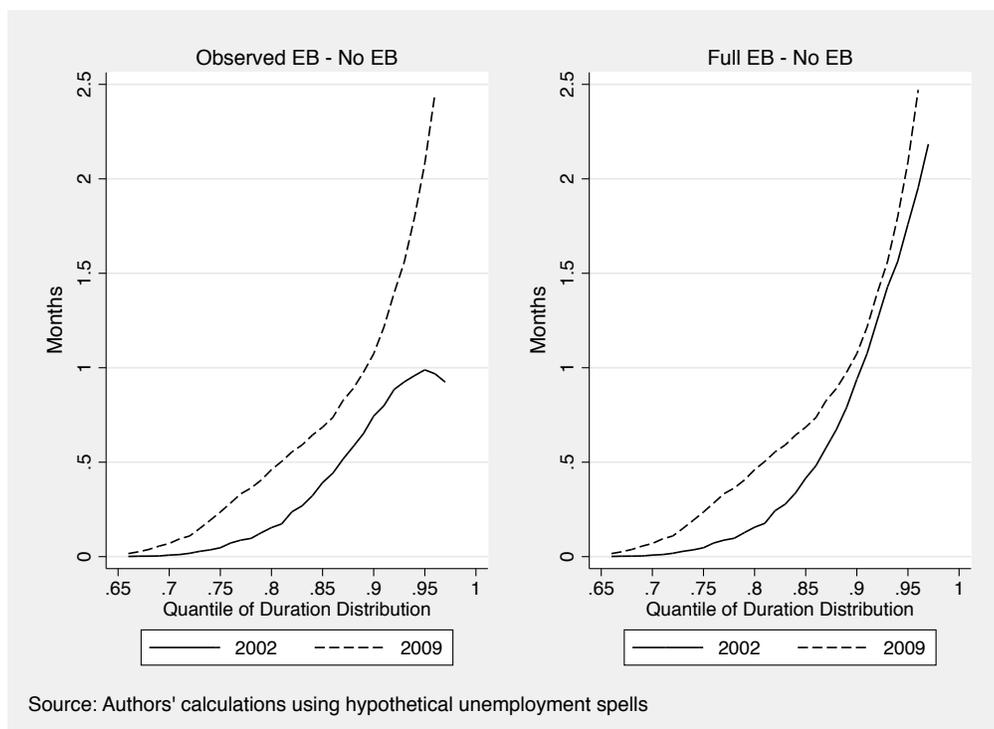


Figure 7: Difference in Time-to-Quantile: Observed-EB-No-EB and Full-EB-No-EB.

Observed-EB and Full-EB scenarios are very similar to each other and far different from the No-EB scenario.

Figure 7 contains plots, for the 2002 and 2009 hypothetical spells, of the difference in time-to-quantile between the Observed-EB and No-EB scenarios (left panel) and the difference between the Full-EB and No-EB scenarios (right panel). The comparison of the Observed-EB and No-EB scenarios suggests that extended benefits had a much larger effect on the duration of unemployment at higher quantiles in the later period than in the earlier period. However, the comparison of the Full-EB and No-EB scenarios makes it clear that the difference across periods is largely due to the higher level of availability and generosity of UI in the 2009-2011 period relative to the 2002-2004 period.

Taken together, figures 6 and 7 imply that there is a substantial effect of a generous extended benefits program on unemployment durations for a small fraction of unemployment spells. Focusing on the Full-EB – No-EB comparison in the right hand columns of the two figures, there is no measurable effect of extended benefits on time-to-quantile for any quantile lower than 0.65. In the 2009-2011 period, the time to the 0.8 quantile is about 2 weeks longer with extended benefits than without extended benefits. In the 2002-2004 period the

extended benefit effect at the 0.8 quantile was about 1 week. To put this in context, the average observed time to the 0.8 quantile was 6.8 months in the 2009-2011 period and 5.4 months in the 2002-2004 period. The extended benefit effect on time-to-quantile is larger at higher quantiles. The average observed time to the 0.9 quantile is about 1 month longer with full extended benefits in both periods. Again to put this in context, the average time to the 0.9 quantile was 10.1 months in the 2009-2011 period and 7.8 months in the 2002-2004 period. While a 10 percent or larger increase in the time-to-quantile seems substantial, this increase only applies to the small fraction of spells that survive to this point.

## 6.2 The Effect of Extended Benefits on Expected Duration

A useful summary measure of the overall effect of extended benefits on the distribution of duration of unemployment is the expected duration of an unemployment spell. Table 6 contains our estimates of the expected duration of unemployment spells for the 96,575 hypothetical spells for each of the three scenarios (Observed-EB, No-EB, Full-EB). Each panel of the table contains estimates for one of the two time periods we defined earlier. The first row of each panel contains predicted expected durations at the observed distribution of extended benefits for the single risk and competing risk models. The second and third rows of the table contain predicted expected durations for the two counterfactuals of 1) No-EB – no extended benefits available and 2) Full-EB – 99 weeks of extended benefits available in all states and months. The next section of the panel shows both the absolute and proportional differences between the average expected duration in the Observed-EB and No-EB scenarios and the final section shows the absolute and proportional differences between the average expected duration in the Full-EB and No-EB scenarios.

Consider first the estimates for Observed-EB in the single risk model in the first column. The expected duration of unemployment spells beginning in March 2002 (Panel 1) averaged 3.56 months. This expected duration is intermediate between those for the No-EB and Full-EB scenarios. Our estimates suggest that the expected duration of unemployment was 4 percent higher due to the existence of extended benefits during the 2002-2004 period. The average expected duration of unemployment spells was substantially higher (4.89 months) for spells beginning in January 2009 (Panel 2). Our estimates suggest that the expected duration of unemployment was 7 percent higher due to the existence of extended benefits during the 2009-2011 period. As we noted earlier, the larger effect of extended benefits in the 2009-2011 period is due to the wider availability of generous extended benefits during this period. This is demonstrated in the last section of each panel, which show the proportional

Table 6: Estimated Effect of Extended Benefits on Expected Duration (in Months)  
UI Eligible Spells

Panel 1: March 2002 – June 2004			
Scenario	Single Risk	Exit to Emp	Exit to NILF
Observed-EB	3.56	5.55	9.05
No-EB	3.42	5.41	8.61
Full-EB	3.65	5.65	9.59
Observed-EB - No-EB	0.14	0.14	0.43
(Obs EB - No-EB)/No-EB	0.04	0.03	0.05
Full-EB - No-EB	0.23	0.24	1.02
(Full-EB - No-EB)/No-EB	0.07	0.04	0.12

Panel 2: January 2009 – April 2011			
Scenario	Single Risk	Exit to Emp	Exit to NILF
Observed-EB	4.89	7.85	10.29
No-EB	4.55	7.62	9.32
Full-EB	4.89	7.85	10.24
Observed-EB - No-EB	0.34	0.23	0.97
(Obs EB - No-EB)/No-EB	0.07	0.03	0.10
Full-EB - No-EB	0.34	0.23	0.92
(Full-EB - No-EB)/No-EB	0.07	0.03	0.10

The estimates based on hypothetical samples of 96,575 spells of unemployment for each time period. The counterfactuals are based on the estimates of the probit model of exit from unemployment presented in table 4. The expected duration is calculated from equation 10. See text for details.

difference in average expected duration between the Full-EB and No-EB scenario (thereby holding constant the extent of availability). This comparison shows very similar difference in expected durations across the two periods (7 percent in both periods).

An interesting question is what the effect of extended benefits is on the measured unemployment rate. One very simple approach is based on two assumptions: 1) extended benefits have no effect on the rate of entry into unemployment or into the labor force and 2) extended benefits have no effect on exit from unemployment for job leavers or new entrants (our UI-ineligible sample). In this case, the proportional effect of extended benefits on the unemployment rate is equal to their proportional effect on the duration of unemployment spells multiplied by the fraction of spells that are UI-eligible.<sup>19</sup> On this basis, extended benefits accounted for 1) 0.12 percentage points (2.2 percent) of the 5.4 percent unemployment

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<sup>19</sup> We use the proportional difference between the Observed-EB expected duration and the No-EB expected duration in table 6 as our measure of the proportional effect of extended benefits on expected duration. In 2003, 55.6 percent of spells were UI eligible, while in 2010 63.0 percent of spells were UI eligible.

rate in 2003 and 2) 0.40 percentage points (4.4 percent) of the 9 percent unemployment rate in 2010.

### **6.2.1 Time to Exit to Employment and NILF: The Competing Risk Model**

There are at least two pathways through which extended benefits could reduce exit from unemployment:

1. The unemployed could reduce search effort or maintain a higher reservation wage. Either results in longer time until an unemployment spell ends in a new job.
2. The unemployed could remain attached to the labor force and searching (perhaps minimally) when, without extended benefits, they would exit the labor force.

The competing risk model is well suited to investigating the extent to which extended benefits works through these pathways, and the estimated effects of extended benefits on the times until exit to employment and exit to NILF are shown in the second and third columns of table 6 for each scenario and in each time period.

Recall that the estimated marginal effects from the probit model do not show a significant effect of extended benefits on exit to employment (table 4). The point estimates themselves imply only a small proportional effect of extended benefits on the expected time until exit to employment. The estimated proportional effect of extended benefits on expected time until exit to employment is only 3-4 percent in each of the time periods (column 2 of table 6), even for the stark comparison of the Full-EB and No-EB scenarios.

The estimated marginal effect of extended benefits on time to exit to NILF were statistically significant in the probit model (table 4) in both time periods. The effects of extended benefits on time to exit to NILF are substantial in each of our sets of hypothetical spells of unemployment (column 3 of table 6). In the set of hypothetical spells starting in March 2002, the difference between the observed and No-EB time to exit to NILF is 0.43 months (5 percent of average expected duration in the No-EB scenario). The effect is relatively small because extended benefits were not widespread and were less generous in this period. The Full-EB - No-EB difference during this period is 1.02 months (12 percent of average expected duration in the No-EB scenario), reflecting the large marginal effect of extended benefits in table 4. In the 2009-2011 time period, the difference between the observed and No-EB time to exit to NILF is almost a full month (10 percent of observed expected time to exit).

These results from the competing risk model are clear cut. We do not find a substantial effect of extended benefits on time to exit to employment. This implies that there is not a

significant reduction in search effort or increase in the reservation wage due to the availability of extended benefits. However, we do find a significant effect of extended benefits on time to exit to NILF. This implies that there may be individuals who remain attached to the labor force, perhaps searching at a low level, because extended benefits are available. In our view, this latter effect of extended benefits does not have first-order efficiency consequences on the level of employment. It reflects mainly a redistribution to long-term job losers who, without extended benefits, would have left the labor force.<sup>20</sup>

### 6.3 The Cross-Section Distribution of Unemployment Duration in a Steady State

The weak labor market in the Great Recession and its aftermath is notable not only for the high level of unemployment but also for the historically high incidence of long-term unemployment, as measured based on the reported duration of in-progress spells (figure 4).<sup>21</sup> An interesting question regards the extent to which these long durations observed in the cross-section are attributable to the availability of extended benefits. The effects of extended benefits on expected durations and on the distribution of spells more generally pertains to the distribution of completed spells. The cross-sectional distribution of durations of spells in progress results from the dynamic process of births and deaths of spells over time, and it is known to over-represent longer spells (length-biased sampling). The result is that the survivor function for a set of spells starting at a point in time appears on its face to be inconsistent with the observed cross-sectional distribution. For example, raw survivor rates calculated from the matched CPS data (table 3) show that for the 2009-2011 period, 13.7 percent of unemployment spells last at least six months while the cross-sectional distribution of unemployment spells for the same period shows 39 percent of unemployment spells in progress have lasted at least 6 months. The same comparison for 2002m3-2004m6 shows that 9.9 percent of unemployment spells from the matched CPS last at least 6 months while the cross-sectional distribution of unemployment spells shows that 22 percent of spells in progress have lasted at least 6 months.

We use our estimates of the model of exit from unemployment together with the steady-state assumption that there is a constant monthly inflow into unemployment to estimate the effect of extended benefits on the cross-sectional distribution of duration of spells in

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<sup>20</sup> Card, Chetty, and Weber (2007) term this the reporting effect of unemployment insurance.

<sup>21</sup> Long-term unemployment is defined by the BLS as durations of at least 27 weeks.

progress. We calculate these cross-sectional distributions for each of the two time periods (2002m3-2004m6 and 2009m1-2011m4) and for the Observed-EB and No-EB scenarios. The idea is that, in a steady state, the estimate of the survivor probability ( $G(\cdot)$ ) at a given duration  $s$  is also an estimate of the probability in a cross-section that a spell that started  $s$  periods earlier is still in progress. On this basis, the fraction of spells in progress in the cross-section with duration greater than  $s$  is

$$P(D \geq s) = \frac{\sum_{j=s}^{\infty} \hat{G}(j)}{\sum_{j=1}^{\infty} \hat{G}(j)}, \quad (11)$$

where  $\hat{G}(j)$  is the average predicted probability that a spell of unemployment that started  $j$  periods earlier is still in progress (the survivor probability in the steady state).<sup>22</sup>

Figure 8 contains plots for the 2002 and 2009 samples of simulated spells of the fraction of spells in a cross-section that have durations at least as long as the indicated number of months. This is calculated using the estimates for the relevant time period contained in table 4 and the two sets of hypothetical spells of unemployment described earlier. In order to implement the steady state assumption, exit probabilities were predicted using the estimated parameters and assuming that the time varying measures (the month indicators and the measures of economic activity in the state) are held constant at their state-specific mean values for each time period. As such we are assuming not only that entry into unemployment is time invariant but also that economic conditions are stable as well. The figures contain a vertical line at 6 months, indicating where the BLS defines unemployment to be “long term.”

The left-hand panel contains the plot for the 2002m3-2004m6 simulated samples of unemployment spells of the fraction of spells in a cross-section that have durations at least as long as the indicated number of months. In the Observed-EB scenario, 20.3 percent of spells in the cross-section are long term ( $\geq 6$  months). In the No-EB scenario, 17.3 percent of spells in the cross-section are long term. This implies that 14.8 percent of long-term unemployment for UI-eligibles in this period was due to the availability of extended benefits.

The right-hand panel contains the analogous plot for the 2009m1-2011m4 simulated samples of unemployment spells. In the Observed-EB scenario, 31.4 percent of spells in the cross-section are long term ( $\geq 6$  months). In the No-EB scenario, 24.3 percent of spells in the cross-section are long term. This implies that 22.6 percent of long-term unemployment

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<sup>22</sup> In practice, we sum the survivor probabilities to 28 months (rather than an infinite sum). Contributions to the sums decline sharply with duration and are trivial after about 24 months.

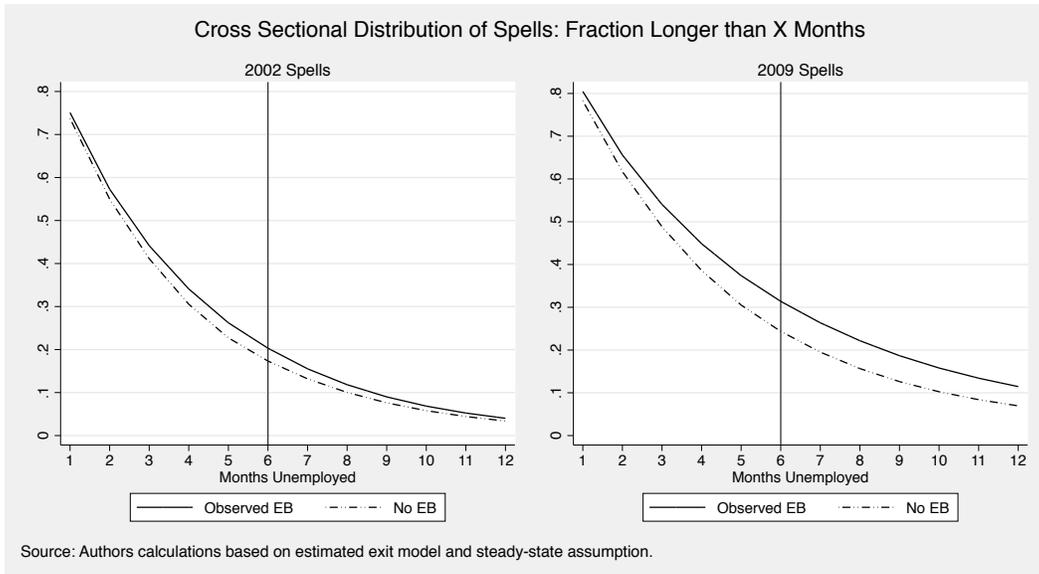


Figure 8: Cross-Sectional Distribution of Duration of Unemployment Spells in Progress in the Steady State.

for UI-eligibles in this later period was due to the availability of extended benefits.

Note that our steady-state calculation cannot mimic the observed cross-section distribution of unemployment durations because of the extreme assumptions of constant flows into unemployment and a stable economic environment. However, this calculation largely reconciles the low survivor rates for individual spells with the relatively high share of long-term unemployment in total unemployment in the cross-section. In the 2002-2004 period, our steady-state calculation for the Observed-EB scenario indicates that 17.3 percent of unemployment in the cross-section is long term, compared with a long-term share of 22 percent based on reported duration in the CPS. In the 2009-2011 period, our steady-state long-term share is 31.4 percent while the reported CPS long-term share is 39 percent.

The results of these calculations show that a substantial fraction (15-25 percent) of the long-term unemployment observed in the cross-section is due to the availability of extended UI benefits.

## 7 Concluding remarks

We examined the impact of the unprecedented extensions of UI benefits in the United States over the past few years on unemployment dynamics and duration and compared their effects

with the extension of UI benefits in the milder recession of the early 2000s. We found small but statistically significant reductions in unemployment exits and small increases in unemployment durations arising from both sets of UI extensions. The magnitude of these overall effects is similar across the two episodes, after we account for the large differences in the amount of UI extension that occurred in each case. From a policy perspective, the similar impact of extended UI across the two episodes suggests that optimal UI policy may be invariant to aggregate labor market conditions. This implication contrasts with the findings from other recent work (Landais et al. 2010; Kroft and Notowidigdo 2011) and indicates the need for additional research on behavioral responses to variation in UI availability and generosity under different labor market states.

Potential caveats to our findings include the possible influence of mis-classification of labor force states leading to mis-measurement of exits from unemployment and the possible mis-classification of individuals as UI-eligible. The former mis-classification is likely to lead to a downward bias in our estimated effects of extended benefits. On the other hand, the results of our placebo analysis of UI-ineligibles raises the possibility that our indicators of UI availability are correlated with unobserved features of state economic conditions, and this would lead to an upward bias in our estimated effects of extended benefits. Thus, these biases likely run in opposite directions and may offset to a greater or lesser extent.

We find that the effect on exit from unemployment occurs primarily through a reduction in labor force exits rather than through exit to employment (job finding). This is important because it implies that extended benefits do not delay the time to re-employment substantially and so do not have large first-order efficiency consequences. The major effect of extended benefits is redistributive, providing income to job losers who would have exited the labor force otherwise (consistent with Card et al. 2007).

While we find small effects of extended benefits both on exits from unemployment and on the expected duration of completed unemployment spells, our estimates imply a substantial effect of extended benefits on the share of unemployment in the cross-section that is long-term. For example, we find that, during the recent downturn, extended benefits increased the expected duration of unemployment by about 7 percent on average but accounted for about 22 percent of long-term unemployment in the cross-section. This apparent conflict arises because extended benefits have only small effects on the distribution of unemployment duration early in spells ( $\leq 6$  months), and the substantial majority of spells end in the first 6 months.

Overall, our estimates suggest that extending unemployment insurance benefits in weak

labor markets has virtually no effect on the rate of job finding but that, on average, unemployment spells are somewhat longer as a subset of UI recipients remain nominally unemployed rather than exit the labor force. In addition to these limited implications for economic efficiency, we find only small impacts on the aggregate labor market. We estimate that extended UI increased the overall unemployment rate by only about 0.4 percentage points in the recent episode, which is small in comparison with the peak unemployment rate of 10 percent.

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## Appendix I – The Timing and Extent of Unemployment Insurance Extensions

The temporary UI extensions initiated in 2008 were historically unprecedented. In addition to the automatic extensions under the EB program (up to 20 weeks) in 2008, four additional extensions took effect between mid-2008 and late 2009, as specified by the Emergency Unemployment Compensation (EUC) program. The initial extension in mid-2008 provided for 13 additional weeks of UI benefits in all states. This was followed by implementation of a second tier (tier II) in late 2008 that allowed receipt up to 33 weeks of extended benefits, with eligibility for weeks 21-33 predicated on a state total unemployment rate (TUR) of at least 6 percent or an insured unemployment rate (IUR, based on the number of regular UI claimants) of at least 4 percent.

EUC Tiers III and IV were added in early November of 2009. At that point, Tier II eligibility was expanded by 1 week to 14, and the unemployment rate threshold was eliminated for Tier II eligibility and applied to Tier III eligibility (13 weeks). Tier IV eligibility (6 weeks) was based on a TUR of at least 8.5 percent or an IUR of at least 6 percent.<sup>23</sup> As of November of 2009 and continuing through early 2012, individuals in states that met eligibility requirements for EB and all EUC tiers could receive up to 99 weeks of UI payments (26 weeks of regular benefits, 20 weeks of EB, and 53 weeks of EUC).

The maximum of 99 weeks through existing legislation continued into early September 2012. By that point, however, many states had "triggered off" (i.e., became ineligible for) available EUC tiers and EB benefits based on declines in their unemployment rates. In addition, as of early September 2012, the number of weeks available through the various EUC tiers was reduced slightly, lowering the maximum number of UI weeks available from 99 to 93. By the end of 2012, no state qualified for all EUC tiers and maximum EB benefits; the highest number of weeks available through all programs in any state at that point was 83. In addition, in 2011 and 2012, a small number of states (seven) reduced their normal UI availability below 26 weeks, in some cases based on formulas that depend on the existing state unemployment rate. These reductions in normal UI weeks also reduced the number of weeks of extended UI benefits available in those states, through the formulas specified in the

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<sup>23</sup> The EUC triggers also are based on a 10 percent increase in the unemployment rate over the preceding 2 years. This constraint has become binding on a widespread basis as labor market weakness persisted. As of late 2010, the legislation was changed so that the unemployment rate 3 years earlier was used; some states chose to opt out of the higher tiers subsequent to that change.

EUC legislation. We include these features of state law in our data on available UI weeks.

It is important to note that the EUC program was subject to several temporary suspensions arising from legislative disagreements that precluded seamless reauthorization of the program. Suspensions of 12 days or longer occurred during April, June-July, and December of 2010. Although otherwise eligible individuals were unable to apply for a new tier of EUC benefits during those months, these individuals were permitted to attain the new tier after program reinstatement and receive retroactive benefits. Because reinstatement was expected and individuals were allowed to receive benefits on their existing tiers during the suspensions, it is likely that the suspension periods did not significantly offset any behavioral responses to the overall extension programs. In all of our analyses, we record UI availability during the EUC suspension months as being identical to availability in the month immediately preceding the suspension, under the assumption that job search behavior was not affected by the suspensions.

A similar but much more limited extension of UI benefits occurred through the Temporary Extension of Unemployment Compensation (TEUC) legislation that was effective from March 2002 through early 2004. In all states, individuals who exhausted their regular state UI benefits were automatically eligible for an additional 13 weeks of benefits. Individuals in states classified as "high unemployment" (based on the level and change in the insured unemployment rate) were eligible for an additional 13 weeks of UI benefits. In conjunction with EB eligibility of up to 20 weeks, a maximum of 46 weeks of extended benefits and 72 weeks of total benefits were available during the period of TEUC authorization. The TEUC program was phased out in early 2004, with individuals who filed a valid TEUC claim prior to the end of 2003 able to continue receiving those benefits for as long as they lasted in 2004 (up to 13 weeks, so potentially through March).