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## Extended Unemployment Insurance and Unemployment Duration in The Great Recession: The U.S. Experience<sup>1</sup>

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### Abstract

The extent and duration of unemployment spells in the United States reached historically high levels in the Great Recession. In response, the availability of unemployment insurance (UI) benefits in the United States has been extended several times since mid-2008 beyond the normal 26 weeks, reaching a maximum of 99 weeks as of late 2009. We exploit variation in the timing and size of UI benefit extensions across states to estimate the overall impact of these extensions on exit from unemployment and unemployment duration. Using monthly matched individual data from the U.S. household survey (CPS) from 1994-2010, we estimate the effects of UI extensions on transitions out of unemployment conditional on state economic conditions and individual characteristics. The results reveal small but precisely estimated reductions in unemployment exits and small increases in unemployment durations arising from the UI extensions. The small size of our estimated effects are not surprising given that, even in the Great Recession, a large majority of spells of unemployment end within the first six months (the period of receipt of normal unemployment insurance benefits) while extended benefits apply only to those relatively few spells that last more than six months.

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# 1 Introduction

Compared with other advanced industrial countries that are part of the OECD, 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 the recent economic downturn (“The Great Recession”) however, UI benefit availability has been extended several times in the United States, reaching a maximum of 99 weeks as of late 2009. This unprecedented expansion of UI availability has been the subject of intense policy debate, which has largely revolved around the disincentive effects of UI payments on job search and reporting effects on labor force attachment (relevant research includes Chetty 2008, Card et al. 2007). 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, basing our analyses on longitudinal microdata for the period over which the extensions became effective.

Past empirical research has produced a range of estimates regarding the disincentive effects of UI benefits on 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 job rationing that occurs in recessions (e.g., Landais, Michaillat, and Saez 2010); it is likely that such considerations loomed especially large during the Great Recession.

Because extended unemployment insurance benefits were much more widely available during the Great Recession than during earlier periods, past empirical results cannot be reliably extrapolated to assess UI disincentive effects in the recent severe recession. We estimate these effects by developing a framework that relies on current labor market data and detailed information on the recent UI expansions. We 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 at the state level. The extension of UI

availability proceeded gradually and varied across states, with qualification for multiple UI extensions occurring based on the level and change in state unemployment 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 in single-risk (exit from unemployment) and competing-risk (exit to employment vs. exit from the labor force (NILF)) frameworks, conditional on state labor market conditions (the rates of unemployment and employment growth) and on individual characteristics.

We find that the availability of extended benefits reduced the rate of exit from unemployment after the first six months by a small amount. However, there is almost no effect of extended benefits on the unconditional expected duration of unemployment because the large majority of unemployment spells end within the first six months, before extended benefits are an issue. Our preferred estimates, which are based on “UI-eligible spells,” suggest that the expected duration of spells conditional on spells that last at least six months would be reduced by about two weeks if extended benefits were eliminated.<sup>1</sup> Among these spells, the estimates of the competing risk model imply that the elimination of extended benefits in these longer spells would reduce the time to employment by about one week and the time to exit from the labor force (NILF) by about 1.5 weeks.

## 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. The federal

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<sup>1</sup> We classify spells of unemployment as “UI-eligible” if the stated reason for unemployment is the loss of a job through no fault of their own. We discuss below the fact that this criterion does not ensure that an individual we classify as UI-eligible is, in fact, eligible for benefits.

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 are state specific but most commonly are based on an overall unemployment rate of 6.5 percent (for 13 weeks of extensions) 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 between 1958 and the most recent episode beginning in 2008.<sup>2</sup>

The severity of job loss and persistent labor market weakness during and after the recession of 2008-2009 has resulted in an unprecedented expansion of UI benefit availability and takeup. In addition to the automatic extensions under the EB program (up to 20 weeks), four additional extensions took effect between mid-2008 and late 2009, as specified by the Emergency Unemployment Compensation (EUC) program (currently authorized through January 3, 2012). 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 provided 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, measured relative to the total labor force) of at least 4 percent.

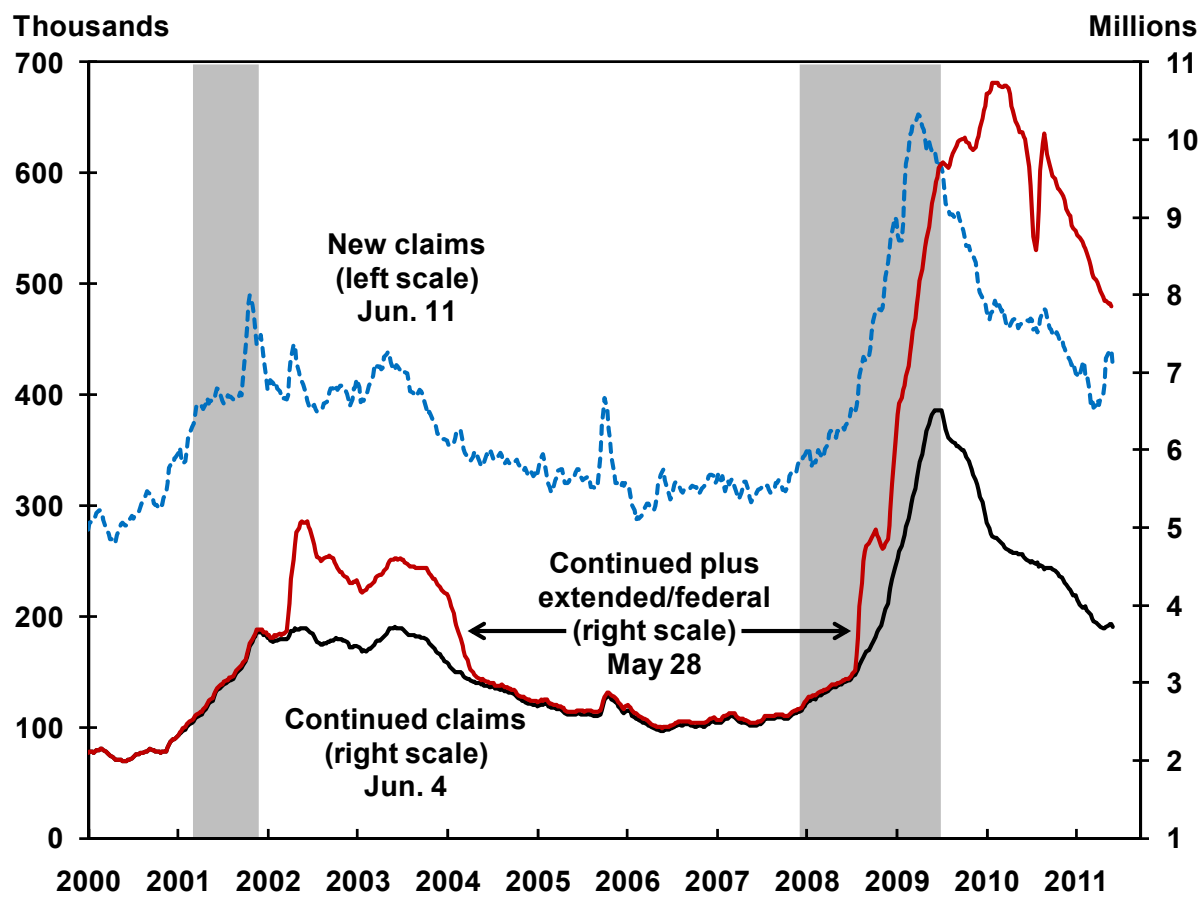
EUC Tiers III and IV were added in early November of 2009. At that point, Tier II eligibility was expanded by 1 week to 14, and the unemployment rate threshold was eliminated for Tier II eligibility and applied to Tier III eligibility (13 weeks). Tier IV eligibility (6 weeks) was based on a TUR of at least 8.5 percent or an IUR of at least 6 percent.<sup>3</sup> As of November of 2009 and continuing forward through early 2012, individuals in states that meet eligibility requirements for EB and all EUC tiers can receive 99 weeks of UI payments (26 weeks of regular benefits, 20 weeks of EB, and 53 weeks of EUC).

It is worth knowing 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

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<sup>2</sup> See Whittaker (2008) and Isaacs, Whittaker, and Shelton (2010) for details regarding the prior and most recent extended benefits programs.

<sup>3</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.



Source: U.S. Department of Labor

Figure 1: Unemployment Insurance Claims (Regular and Extended)

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.

Figure 1 illustrates the expansion of UI receipt through the regular and extended programs. 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). The level of ongoing UI claims is displayed for regular claims (26 or fewer weeks) separately and with claims through available extensions added on. The overall level of continued claims has dropped sharply on net since early 2010. The decline in recipiency was largely accounted for by falling

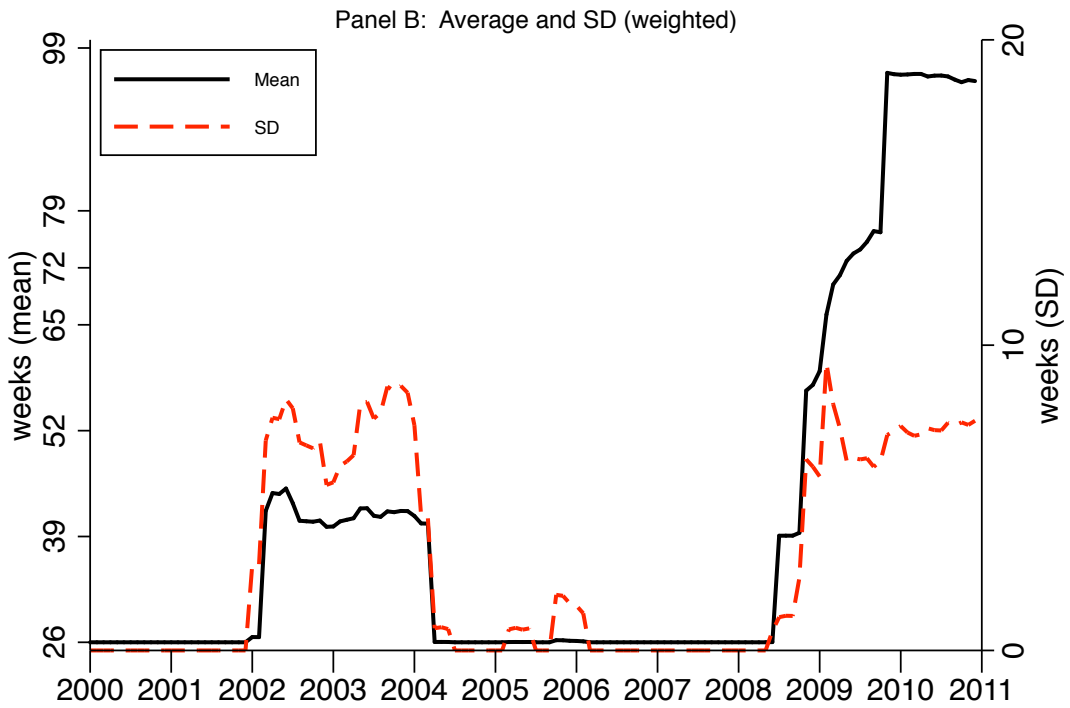
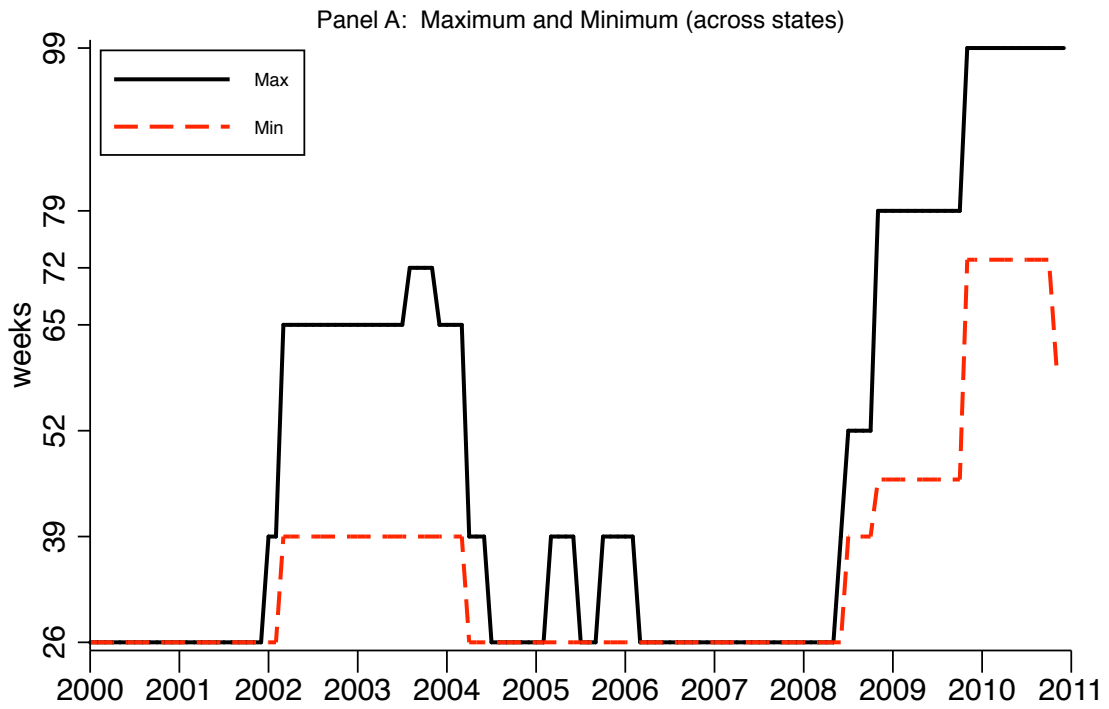
regular claims, but the number of claims under extended programs has fallen as well. The sharp drop followed by a similar increase in mid-2010 corresponds to the suspension period of June-July 2010. A much smaller but still substantial number of extended claimants also were present during the immediately preceding episode of extended claims, from 2002-2004. Between the EB program and temporary legislation enacted in 2002, UI eligibility reached a maximum duration of 72 weeks in 2003. We will also use data for this earlier period in our subsequent empirical analyses.<sup>4</sup>

As suggested by the preceding discussion of EB and EUC program elements, the timing of the extended UI triggers and consequent maximum duration of UI eligibility has varied substantially across states. Different states surpassed the trigger levels for EB and EUC availability at different times, some states never achieved the unemployment rates necessary for the complete extensions up to 99 weeks, and some states triggered off EUC tiers as labor market conditions improved in 2010.

Figure 2 illustrates the variation in eligibility for extended UI over time (years 2000- 2010), which includes the impact of EB and a temporary program in the early 2000s in addition to the recent episode. 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 in the CPS microdata). 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 95 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 the variation needed for estimation is available for both episodes.

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<sup>4</sup> Incorporating information on UI extension episodes prior to 2001 will have limited payoff in our setting because the matched CPS data that we use for our analyses is not generally available prior to 1994.



Note: Panel B series calculated using CPS weighted unemployment observations.

Figure 2: Variation in Total Weeks of UI Available

## 2.2 Past research on UI disincentive effects

Economists have long recognized 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, leaving the results of such an exercise contingent on the study chosen. Moreover, existing research suggests that the disincentive effects of extended UI are likely to be muted relative to the effect implied by past research based on the exhaustion of regular UI benefits, particularly given 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 2010).

Given the drawbacks to reliance 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 on examination of 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 but useful empirical guidance about the likely impacts of extended UI. However, because such studies are not based on formal statistical tests that attempt to isolate the direct effects of UI extensions while conditioning out the effects of state labor market conditions and individual characteristics that are related to extended UI eligibility and receipt, the reliability of such results is difficult to assess.

We propose to build on the existing literature by directly exploiting the variation in the duration of extended UI available in different months in different states, discussed above. In addition, our matched CPS data enable us to incorporate detailed controls for state economic conditions and individual characteristics into the analyses. We describe our empirical framework in more detail in the next section.



### 3 Econometric Framework and Specification

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>5</sup> In this way, spells are “joined in progress” leading to the classic problem of length-biased sampling, which results in a sample of longer spells than the overall distribution of unemployment spells, and our econometric model needs to account for this.

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 that last 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  $n$  observations subsequent to the first observation of the unemployment spell ( $n \leq 3$ ) 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

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

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<sup>5</sup> In theory, the unemployed are asked their duration only in the first month they report being unemployed, and this is automatically incremented in future months in which they report being unemployed. In fact, the spell durations are not this clean.

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 likelihood function is a standard probit 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. Monthly observations where the spell has ended by the next month have a “one” outcome. Each spell in the sample has at most one observation with a “one” outcome (the end of the spell).

A probit model is specified by assuming that 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

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

where  $X_{it}$  is a vector of individual and economic variables,  $\beta$  is a vector of parameters,  $EB_{it}$  is an indicator for the availability of extended benefits in person  $i$ 's state in month  $t$ ,  $\delta$  is a parameter measuring the marginal effect of the availability of extended benefits on the hazard of an unemployment spell ending, and  $\epsilon_{it}$  is an error term with a standard normal distribution. The hazard of a spell ending is then

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

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

The estimated model includes controls for age and its square, education, female, married, female\*married, nonwhite, indicators for year, calendar month (to control for seasonality), and state. In order to account for local labor market conditions, the model also includes controls for the monthly seasonally adjusted state unemployment rate and the 3-month annualized growth in seasonally adjusted log non-farm payroll employment. To allow for a flexible baseline hazard, the model also includes a set of indicators for the first 6 months of unemployment (0-6) and single indicators for months 7-11, month 12, months 13-17, month 18, months 19-23, and month 24 and later (13 categories in total). The key variable with respect to extended benefits is the indicator for the availability of extended benefits in a

given state and month for a worker with the observed duration level ( $EB_{it}$ ). Note that this variable equals zero for all observations with six or fewer months of unemployment. Workers at these durations are eligible for normal unemployment insurance, and the effect of normal unemployment insurance is captured by the unemployment duration indicators for these months.

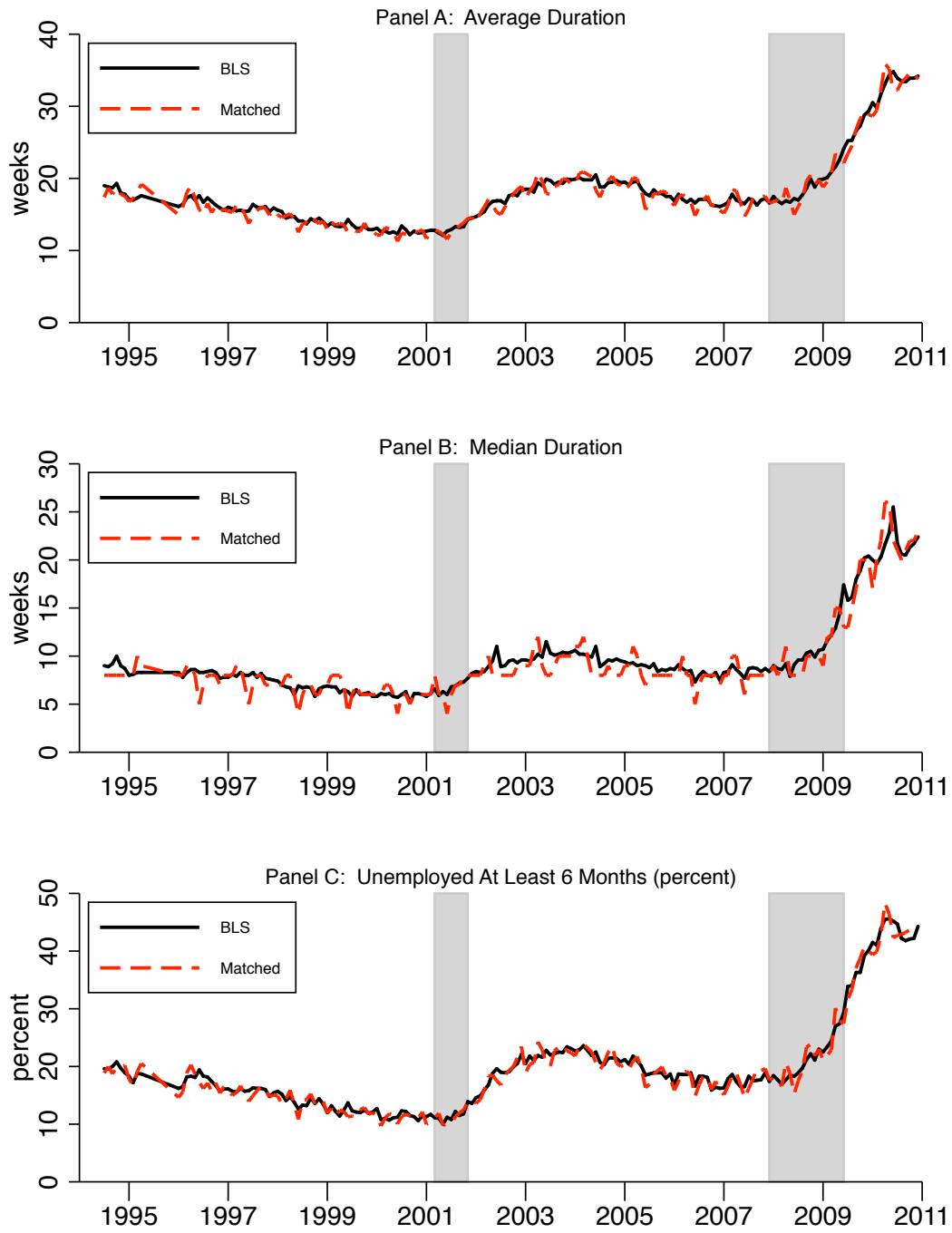
### 3.1 Sample and Identification

The rotation group structure of the CPS revisits surveys a given address 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 month-to-month up to three times. Failures to match occur when a household moves to a new housing unit between interviews. This generally occurs less than five percent of the time.

Our matched CPS sample, covering the 1994-2010 period, has data for 590,698 monthly observations on 362,889 spells of unemployment. In order to examine the representativeness of our matched sample, we compared the reported unemployment duration in our sample with published statistics on unemployment duration from the U.S. Bureau of Labor Statistics (BLS). Figure 3 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 of unemployment durations from the matched CPS and the BLS reported statistics from all sampled unemployment spells in the CPS.

Figure 3 also provides a sense of changes in unemployment duration over the 1994-2010 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 3 comes 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 1, 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 full sample and for



Note: From U.S. Bureau of Labor Statistics (BLS; seasonally adjusted) and authors' tabulations of matched CPS data (weighted). Based on duration of in-progress spells. Recession periods in gray.

Figure 3: Unemployment Duration, BLS and Matched CPS Data

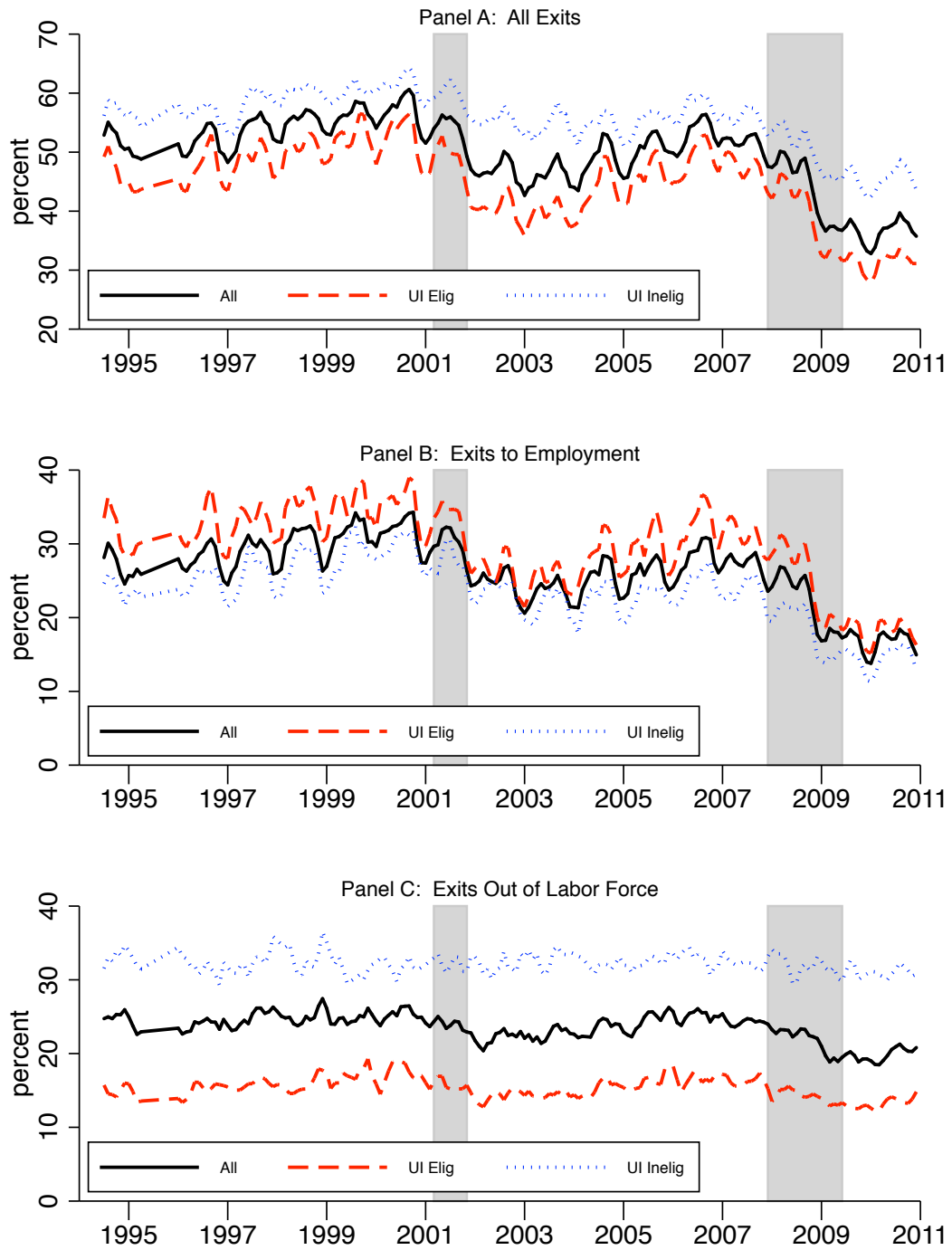
Table 1: Unemployment Survivor Rates, by Duration and UI Eligibility

<u>Duration</u> <u>(months)</u>	1994-2010			2008-2010		
	<u>Full</u> <u>sample</u>	<u>UI</u> <u>eligibles</u>	<u>UI</u> <u>ineligibles</u>	<u>Full</u> <u>sample</u>	<u>UI</u> <u>eligibles</u>	<u>UI</u> <u>ineligibles</u>
0	0.354	0.384	0.324	0.414	0.439	0.375
1	0.164	0.204	0.129	0.220	0.260	0.168
2	0.086	0.121	0.058	0.130	0.169	0.085
3	0.049	0.077	0.029	0.082	0.117	0.047
4	0.030	0.051	0.016	0.055	0.083	0.028
5	0.018	0.034	0.008	0.036	0.059	0.016
6	0.011	0.022	0.004	0.024	0.042	0.009
7	0.007	0.015	0.002	0.017	0.032	0.005
8	0.004	0.010	0.001	0.011	0.023	0.003
9	0.003	0.007	0.001	0.008	0.017	0.002
10	0.002	0.005	0.000	0.006	0.013	0.001
11	0.001	0.004	0.000	0.004	0.010	0.001
12	0.001	0.002	0.000	0.003	0.007	0.000

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

the recent recession and its aftermath (years 2008-2010). The exit rates from unemployment are sufficiently high that very few individuals remain unemployed after the first six months. Even during the recent recession episode, only about 2.5 spells last past 6 months. The survivor rates are higher in general for UI-eligible than for ineligible individuals, indicating that they have characteristics associated with longer unemployment spells (lower exit rates). This latter point emphasizes the importance of controlling for observable characteristics such as age in our analyses of the effects of UI.

Figure 4 complements table 1 by showing monthly exit rates from unemployment over our complete sample frame, 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 in 2010. Higher exit rates are evident for UI-ineligible individuals than for eligible individuals, particularly for exits out of the labor force



Note: Authors' tabulations from matched CPS data (weighted), expressed as 3-month moving averages. Recession periods in gray.

Figure 4: Unemployment Exit Rates, Matched CPS

Table 2: Sample Breakdown by Eligibility for Unemployment Insurance

Sample	Spells	Uncensored	Censored	Total Obs
UI-Eligible	176,580	98,378	78,202	308,877
UI-Ineligible	186,309	120,211	66,098	281,821
All Spells	362,889	218,589	144,300	590,698

Based on matched CPS sample 1994-2010. A spell is censored if it does not end within the observation period.

in Panel C. The rates of labor force exit exhibit very little cyclicity, with only a slight net decline evident during the recent recession and essentially no cyclicity evident for UI ineligible.

Our matched sample includes 176,580 UI-eligible spells and 186,309 UI-ineligible spells. Table 2 contains sample counts and information on censoring by type of spell. We classify UI eligibility based on the reported reason for unemployment, and it will not always reflect actual UI eligibility or receipt.<sup>6</sup> For example, a worker who reports a job loss may not have sufficient previous employment to qualify for unemployment insurance, may have been fired for cause, or may be an eligible worker who has not chosen to take up benefits.

Table 3 provides supporting evidence for our working definition of UI eligibility. Each March, the regular monthly CPS survey 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 on their income receipt in year  $t$  recorded in the year  $t + 1$  March supplement. Based on this match, the table 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 2005-2009. The percent of UI eligibles who report UI income reached about 50 percent in 2009. The percentage of ineligibles reporting UI income is much lower, generally on the order of 5-10 percent, although it reached nearly 15 percent for individuals recorded as unemployed in December of 2009. Because individuals may be subject to multiple unemployment spells over a calendar year, an unknown but possibly substantial proportion of individuals identi-

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<sup>6</sup> We classify a spell as UI-eligible where the individual reports being unemployed due to loss of a job. We classify a spell as UI-ineligible where the individual reports being unemployed after leaving a job voluntarily or after entry to the labor force.

Table 3: UI Income Receipt, by Eligibility Status  
(unemployed in base month, matched to subsequent March survey)

Income year	UI status:	March to March		December to March	
		<u>Eligible</u>	<u>Ineligible</u>	<u>Eligible</u>	<u>Ineligible</u>
2005		29.8%	5.2%	26.4%	4.9%
2006		28.1%	3.8%	33.5%	5.4%
2007		29.9%	5.2%	29.7%	7.4%
2008		31.9%	8.8%	43.8%	6.8%
2009		49.0%	6.9%	50.5%	14.4%

Note: Percent of unemployed individuals who report receiving UI income, matched CPS data (as indicated). Eligibles are job losers/layoff, ineligible are quits/entrants.

fied 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 largely reflects actual eligibility status, although take-up of unemployment insurance by those eligible is likely not universal.<sup>7</sup>

Clearly, the distinction between the UI-eligible and UI-ineligible is useful. We expect that the availability of extended benefit in a state at a point in time will have a first order effect on the UI-eligible spells but not on the UI-ineligible spells. Despite the fact that extended benefits should not have a first order effect on the duration of ineligible spells, we present estimates for the overall sample (including both UI-eligible and UI-ineligible spells) in order to provide a baseline estimate of the effect of the availability of extended benefits on the overall distribution on unemployment durations.

Next, we restrict our model to the 176,580 UI-eligible spells of unemployment to derive our preferred estimate of the effect of extended benefits on the unemployment durations of those directly affected. Identification of the effect of extended benefits in this model comes from variation over time and across states in the availability of extended benefits controlling for local economic conditions. In others words, there have been times and places

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<sup>7</sup> See Blank and Card (1991) and Anderson and Meyer (1997) for earlier studies of incomplete take-up rates for unemployment insurance.



with particular weak labor markets where extended benefits have been available and other times and places with weak labor markets where extended benefits have not been available. It is this contrast, together with our linear specification of economic conditions, that provides identification.<sup>8</sup>

An alternative/additional source of identification could come from a control group of unemployed workers who are not eligible for extended benefits. This group includes job leavers and new labor market entrants. The availability of extended benefits may have some effect on the unemployment experience of those not eligible through general equilibrium effects on the operation of the labor market. However, this is likely to be second order compared to the effect of the availability of extended benefits on the unemployment experience of those eligible. Importantly, the unemployment experience of the UI-ineligible group may be related to the availability of extended benefits through a correlation between activation of extended UI and economic conditions not captured directly by variables in the model.

We present difference-in-difference (DID) estimates of the effect of extended benefits on the hazard. These are calculated by estimating the hazard model specified above separately for the UI-eligible and the UI-ineligible spells and differencing the estimates of  $\delta$  from these models. Specifically, the DID estimate of the effect of extended benefits is

$$\Delta\Delta = \hat{\delta}_{elig} - \hat{\delta}_{inelig}, \quad (7)$$

where  $\hat{\delta}_{elig}$  and  $\hat{\delta}_{inelig}$  are the estimates of  $\delta$  for the UI-eligible and UI-ineligible samples respectively.

The UI-ineligibles serve as a control group that can account for a correlation of the availability of extended benefits with labor market conditions. However, it may well be that the underlying dynamics of unemployment are different for job leavers and new entrants relative to job losers. To the extent that this is the case, the UI-ineligible are not a useful control group, and we do not place much weight on the DID results.

To summarize, we present several sets of estimates of the effect of extended benefits based on the hazard in equation 6:

- Estimates based on the overall sample as a baseline.
- Estimates based on the UI-eligible sample, without an explicit control group, which rely on variation over time and across states for identification.

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<sup>8</sup> The linear specification implies that the marginal effect of economic conditions on the latent variable determining the hazard of leaving unemployment is the same at lower and higher levels of unemployment.

- DID estimates using UI-ineligible spells as a control group.

### 3.2 A Competing Risk Model

It is of interest to understand how extended benefits affect both the time it takes an unemployed worker to find employment and the time before an unemployed worker leaves the labor force. 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.

Table 4 contains counts of number of spells that end with employment, end with NILF, and are censored. 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. 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.

## 4 Results

We present estimates of various probit models of the probability that an unemployment spell ends in a given month. The key parameter we estimate is the normalized effect of the availability of extended benefits on the probability of a spell ending. As specified in

Table 4: Sample Breakdown by Eligibility for Unemployment Insurance  
Competing Risk: Employment vs. NILF

Sample	Spells	End in UE	End in UN	Censored
UI-Eligible	176,580	64,464	33,914	78,202
UI-Ineligible	186,309	51,024	69,187	66,098
All Spells	362,889	115,488	103,101	144,300

Based on matched CPS sample 1994-2010. A spell is censored if it does not end within the observation period.

equation 5, the underlying estimated probit parameter is  $\hat{\delta}$ , and the average marginal effect on the hazard is computed from this as

$$\hat{\delta}^* = \hat{\delta}\phi(\bar{X}\hat{\beta} + \hat{\delta}\bar{E}B), \quad (8)$$

where  $\phi(\cdot)$  is the standard normal probability density function,  $\bar{X}$  is the average value of the  $X$  vector, and  $\bar{E}B$  is the average value of the extended benefit indicator.

In order to quantify the effect of extended benefits on unemployment duration, we also calculate two other quantities based on our estimates. The first is the effect of extended benefits on the expected duration of unemployment spells ( $E(D)$ ), and the second is the effect of extended benefits on the expected duration of unemployment spells conditional on spells lasting at least six months ( $E(D|D \geq 6)$ ). This conditional expectation is of particular interest because extended benefits can have a first-order effect only on spells long enough to exhaust the 26 weeks of standard UI benefits. We calculate these expected durations based on our estimates from the 1994-2010 period by creating a data set of hypothetical spells of unemployment, all of which begin in October 2008. We create a spell of unemployment (starting in October 2008) for each individual in our estimation sample who we observed with a spell of unemployment at any time in the 2008-2010 period. We calculate the expected duration of each spell as

$$E(D_i) = \left[ \sum_{s=1}^{27} s \cdot h_i(s) \cdot G_i(s-1) \right] + 30 \cdot G_i(27), \quad (9)$$

where  $G_i(\cdot)$  is the survivor function for spell  $i$  and  $h_i(\cdot)$  is the hazard function for spell  $i$ . We calculate the conditional expectation of duration for spells lasting at least six months as

$$E(D_i|D_i \geq 6) = \frac{[\sum_{s=6}^{27} s \cdot h_i(s) \cdot G_i(s-1)] + 30 \cdot G_i(27)}{G_i(6)} \quad (10)$$

The survivor and hazard functions for each spell are calculated using the estimates of the probit model and the characteristics of each spell. The last term in equations 9 and 10 reflects an assumption that all spells that last past the 27th month end in the 30th month.<sup>9</sup>

We calculate each of these expectations two ways. First, we calculate the expected durations based on our estimates of the hazard model incorporating the actual timing and availability of extended benefits. Second, we calculate the expected durations based on the same estimates of the hazard model but assuming that there have been no extended benefits in any state at any time. We then average each set of estimated durations over the set of hypothetical unemployment spells, and we focus on these average expected durations. Differencing these two estimates of the average expectation yields an estimate of the effect of eliminating extended benefits on expected unemployment duration.

## 4.1 The Baseline: All Unemployment Spells

We start by presenting baseline results for the sample of all unemployment spells (both those eligible for UI and ineligible for UI). The first column of table 5 contains estimates of the marginal effect of the availability of extended UI benefits in a given state and month on the probability of exit from unemployment. These estimates show that there is a significant negative effect of the availability of extended benefits on the hazard of an unemployment spell ending. We also estimate that the availability of EB reduces the hazard of exit to both employment and NILF.

Consider first the single-risk estimate. We predict that the average hazard of a spell of unemployment ending in the seventh or later month would increase 5 percent from 0.306 to 0.321 if extended benefits were eliminated. While this seems substantial, it does not follow that there is a substantial effect of the availability of EB on the expected duration of unemployment spells ( $E(D)$ ). This is because the large majority of unemployment spells end before 6 months duration. The predicted survivor probability at the sixth month in our baseline model is 0.049, implying that more than 95 percent of unemployment spells end before extended benefits can be an issue.<sup>10</sup> One metric that illustrates this is the expected

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<sup>9</sup> In order to compute an expectation, we need to assume something about the end of long spells. Our assumption of 30 months is benign given the extremely small survival probability at 27 months ( $\approx 0.0002$ ).

<sup>10</sup> The finding that 95 percent of unemployment spells end in the first six months contrasts sharply with the published CPS-based statistics on the distribution of unemployment durations. In those data spells of unemployment are many times longer than our estimates suggest. The explanation is that the CPS, being fundamentally a cross section of spells in progress at a point in time, yields a length-biased sample where long unemployment spells, because they are more likely to survive until the survey date, are substantially

Table 5: Estimated Effect of Extended Benefits on Exit from Unemployment and on Expected Duration, Baseline

	$\hat{\delta}^*$	$E(D)$			$E(D D \geq 6 \text{ Months})$		
		EB	No EB	$\Delta$	EB	No EB	$\Delta$
Single Risk	-0.0269 (0.0043)	2.16	2.15	-0.01	12.61	12.38	-0.23
Exit to Emp	-0.0165 (0.0040)	5.78	5.77	-0.01	16.16	16.07	-0.09
Exit to NILF	-0.0223 (0.0035)	5.80	5.69	-0.11	13.38	13.02	-0.36

The reported value of  $\hat{\delta}^*$  is the estimated marginal effect of the availability of extended unemployment insurance benefits on the hazard of an unemployment spell ending in a given month. This is derived from a probit model that includes controls for age and its square, education, sex, marital status, the interaction of sex and marital status, calendar year, calendar month, the state monthly unemployment rate, and the rate of growth of log employment. The model also includes controls for the length to-date of the unemployment spell and the availability of extended UI benefits. The estimates are weighted by the CPS sampling weights, and asymptotic standard errors are in parentheses. The sample includes 444,885 monthly observations on 299,645 spells of unemployment in the 1994-2010 period.

duration of unemployment spells. The estimates in the first row table 5 imply that the elimination of extended benefits would reduce the expected duration of an unemployment spell trivially from 2.16 to 2.15 months. These expected durations are so short because the hazard of a spell ending is very high in the early months.<sup>11</sup> This is why the 6 month survivor rate is only 0.05. The elimination of extended benefits can have no effect on the hazard in the first six months in our model, and the estimated five percent increase in the exit hazard after the first six months, noted above, applies only to the five percent of unemployment spells that have lasted that long.

The right hand panel of table 5 contains estimates of expected duration of unemployment spells conditional on the spells lasting at least 6 months ( $E(D|D \geq 6)$ ). Since these are the spells for which extended benefits are relevant, we expect larger effects of the elimination of extended benefits. Indeed, the elimination of extended benefits reduces  $E(D|D \geq 6)$  by two percent, from 12.61 months to 12.38 months. While the effect is relatively small, there does appear to be some measurable relationship of extended benefits with the duration of long unemployment spells.

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over-represented.

<sup>11</sup> The hazard is estimated to be 0.58 in the first month, 0.45 in the second month, and so on.

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.

The second and third rows of table 5 contain the estimates of the competing risk model. Extended benefits significantly reduces the rate of exit both to employment and to NILF. As is reasonable, the cause-specific expected durations are longer than the expected durations of the single risk model. Intuitively, this is because the expected duration in the single risk model is essentially the expected minimum of time to exit to employment and the time to exit to NILF. Elimination of extended benefits reduces unconditional expected time to employment only by a small amount. The effect of elimination of extended benefits on the expected time to employment conditional on a spell lasting at least six months is also relatively small at -0.09 months (0.6 percent). The effect of elimination of extended benefits on time until exiting the labor force (NILF) is larger. Elimination of extended benefits reduces unconditional expected time to NILF by 0.11 months (1.9 percent). The effect of elimination of extended benefits on the expected time to NILF conditional on a spell lasting at least six months is at -0.36 months (2.6 percent). We conclude from this baseline analysis that the reduction in unemployment durations for long spells from the elimination of extended benefits is associated largely with a reduction in the time to NILF rather than any substantial effect on the time to employment. In other words, it may be that some marginally attached workers who receive extended benefits remain in the labor force in order to receive these benefits.

## 4.2 Unemployment Among Job Losers: The UI Eligibles

While the baseline results are suggestive and summarize the overall relationship between unemployment spells and the availability of extended benefits, the primary effect of extended benefits must be on those workers who became unemployed for reasons that would make them eligible for unemployment insurance including extended benefits, when and where available.

Table 6: Estimated Effect of Extended Benefits on Exit from Unemployment and on Expected Duration, UI-Eligible Spells

	$\hat{\delta}^*$	$E(D)$			$E(D D \geq 6 \text{ Months})$		
		EB	No EB	$\Delta$	EB	No EB	$\Delta$
Single Risk	-0.0368 (0.0056)	2.47	2.43	-0.04	12.66	12.26	-0.40
Exit to Emp	-0.0305 (0.0053)	4.85	4.81	-0.04	15.92	15.68	-0.24
Exit to NILF	-0.0172 (0.0037)	8.12	7.97	-0.15	14.39	14.03	-0.36

The reported value of  $\hat{\delta}^*$  is the estimated marginal effect of the availability of extended unemployment insurance benefits on the hazard of an unemployment spell ending in a given month. This is derived from a probit model that includes controls for age and its square, education, sex, marital status, the interaction of sex and marital status, calendar year, calendar month, the state monthly unemployment rate, and the rate of growth of log employment. The model also includes controls for the length to-date of the unemployment spell and the availability of extended UI benefits. The estimates are weighted by the CPS sampling weights, and asymptotic standard errors are in parentheses. The sample includes 229,994 monthly observations on 146,068 spells of unemployment that are eligible for UI (job losers) in the 1994-2010 period.

In this section, we repeat our analysis limiting the sample to UI eligible spells.<sup>12</sup> This is our preferred specification.

The single risk estimates, in the first row of table 6 show a significant negative 3.7 percentage point effect of the availability of extended benefits on the hazard of an eligible unemployment spell ending. The effect is to reduce expected duration by 0.04 months or about 1.6 percent of the mean of 2.47 months. As in the baseline, the effect of eliminating extended benefits on the expected duration conditional on a spell lasting at least six months ( $E(D|D \geq 6)$ ) is larger at -0.4 months or about 3.2 percent of the conditional mean of 12.7 months. The estimated effects for the UI-eligible spells are larger than we found in the baseline (table 5) for the good reason that the baseline includes UI-ineligible spells where extended benefits are not likely to have substantial effects.

The second and third rows of table 6 contain the estimates of the competing risk model for the UI-eligible spells. As in the baseline, extended benefits significantly reduce the rate of exit both to employment and to NILF. Elimination of extended benefits reduces

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<sup>12</sup> We make the distinction between UI-eligible and UI-ineligible spells based on the stated reason for entering unemployment. Some of the spells we classify as UI-eligible are not, in fact, UI eligible because of insufficient prior work history or firing for cause.

unconditional expected time to employment by 0.04 months (0.8 percent). The effect of elimination of extended benefits on the expected time to employment conditional on a spell lasting at least six months is somewhat larger at -0.24 months (1.5 percent). The effect of elimination of extended benefits on time until exiting the labor force (NILF) is larger still. Elimination of extended benefits reduces unconditional expected time to NILF by 0.15 months (1.8 percent). The effect of elimination of extended benefits on the expected time to NILF conditional on a spell lasting at least six months is at -0.36 months (2.5 percent). This modifies our conclusion from the baseline estimates that the reduction in unemployment durations for long spells from the elimination of extended benefits is associated largely with a reduction in the time to NILF rather than any substantial effect on the time to employment. The estimates based on the UI-eligible spells implies that extended benefits reduces the rate of exit from unemployment both to employment (perhaps by reducing search intensity and increasing reservation wages) and to NILF (perhaps by keeping workers in the labor force to receive benefits).

### 4.3 Unemployment Among Job Leavers and New Entrants: The UI Ineligibles as a Control Group

Table 7 contains estimates from the unemployment hazard model for UI-ineligible spells (job leavers and new entrants). We present these results not because we think that extended benefits have a direct first-order effect on the duration of unemployment for spells that are not eligible for extended benefits but because we wish to explore using these spells as a control group for the UI-eligible spells. To the extent that the variables included in our model do not capture all variation in labor market conditions that affect the availability of extended benefits and the duration of insured unemployment, the UI-ineligible spells can serve as a control group that will allow us to difference out the effect of unmeasured economic conditions on the exit rate from unemployment. Despite our earlier finding that ineligible rarely receive UI payments, they may not be a very good control group because the dynamics of unemployment are different for job leavers and new entrants (the UI-ineligible group) relative to job losers (the UI-eligible group).

The single risk estimates, in the first row of table 7 show no effect ( $\hat{\delta}^* = 0.00$ ) of the availability of extended benefits on the hazard of an ineligible unemployment spell ending. The second and third rows of table 6 contain the estimates of the competing risk model for the UI-eligible spells. While there is a small and marginally significant *positive* point estimate on the availability-of-EB variable in the exit-to-employment model, there is no measurable



Table 7: Estimated Effect of Extended Benefits on Exit from Unemployment and on Expected Duration, UI-Ineligible Spells

	$\hat{\delta}^*$	$E(D)$			$E(D D \geq 6 \text{ Months})$		
		EB	No EB	$\Delta$	EB	No EB	$\Delta$
Single Risk	0.0002 (0.0066)	1.82	1.82	0.00	12.96	12.96	0.00
Exit to Emp	0.0125 (0.0061)	6.95	6.95	0.00	15.86	15.91	0.04
Exit to NILF	-0.0228 (0.0061)	4.07	4.01	-0.06	12.72	12.43	-0.29

The reported value of  $\hat{\delta}^*$  is the estimated marginal effect of the availability of extended unemployment insurance benefits on the hazard of an unemployment spell ending in a given month. This is derived from a probit model that includes controls for age and its square, education, sex, marital status, the interaction of sex and marital status, calendar year, calendar month, the state monthly unemployment rate, and the rate of growth of log employment. The model also includes controls for the length to-date of the unemployment spell and the availability of extended UI benefits. The estimates are weighted by the CPS sampling weights, and asymptotic standard errors are in parentheses. The sample includes 214,437 monthly observations on 153,570 spells of unemployment that are not eligible for UI (job leavers and new entrants) in the 1994-2010 period.

effect of the elimination of extended benefits on unconditional expected time until exit to employment. This point estimate implies a small positive effect of eliminating extended benefits on  $E(D|D \geq 6)$  of 0.04 months (0.25 percent). A larger negative effect is found for the elimination of extended benefits on the time until exit to NILF. The unconditional expected duration of time to NILF is about 1.5 percent lower (0.06 months) when extended benefits are eliminated. The conditional expected duration ( $E(D|D \geq 6)$ ) of time to NILF is about 2.3 percent lower when extended benefits are eliminated.

We are not arguing for a causal relationship between extended benefits and the duration of unemployment for UI-ineligible spells. The small relationships we find may reflect omitted variables, and we now use these estimates to calculate difference-in-difference estimates of the effect of extended benefits on the exit from unemployment and unemployment duration.

#### 4.4 Difference-in-Difference Estimates of the Effect of Extended Benefits

Table 8 contains the components of the calculation of the difference-in-difference estimates of the effect of extended benefits on unemployment duration. These components are derived from the results presented in tables 6 and 7. Given the generally small relationships of

Table 8: Estimated Effect of Removing Extended Benefits on Expected Duration  
 DID estimates using UI-Ineligible Spells as Control

	$\hat{\delta}^*$			$E(D)$			$E(D D \geq 6 \text{ Months})$		
	$\hat{\delta}_{elig}^*$	$\hat{\delta}_{inelig}^*$	$\Delta\hat{\delta}^*$	$\Delta_{elig}$	$\Delta_{inelig}$	$\Delta\Delta$	$\Delta_{elig}$	$\Delta_{inelig}$	$\Delta\Delta$
Single Risk	-0.0368 (0.0056)	0.0002 (0.0066)	-0.0366 (0.0087)	-0.04	0.00	-0.04	-0.40	0.00	-0.40
Exit to Emp	-0.0305 (0.0053)	0.0125 (0.0061)	-0.0430 (0.0081)	-0.04	0.00	-0.04	-0.24	0.04	-0.28
Exit to NILF	-0.0172 (0.0037)	-0.0228 (0.0061)	0.0056 (0.0071)	-0.15	-0.06	-0.09	-0.36	-0.29	-0.07

The estimates for the UI-eligible models are derived from table 6. The estimates for the UI-ineligible models are derived from table 7. See the notes to those tables for details.

extended benefits with the outcomes of UI-ineligible spells shown in table 7, the DID results look very much like our preferred estimates for the UI-eligible sample in table 6. We begin with the difference between the estimated marginal effects of extended benefits ( $\hat{\delta}^*$ ) in the first three columns of tables 8. The estimated differenced marginal effect in the single-risk model is significantly negative and is virtually identical to the estimate for the UI-eligible model. The DID estimates of the conditional and unconditional expected durations are also identical to the difference estimate for the UI-eligible model. This is due to lack of any relationship between extended benefits and the hazard of a spell ending for UI-ineligible spells.

The results are slightly different in the competing risk model. The effect of extended benefits on exit to employment for the control group is positive which implies a slightly larger DID estimate. The DID estimates imply that the expected conditional duration ( $E(D|D \geq 6)$ ) until exit to employment is 0.28 months shorter when extended benefits are removed. There is a larger effect of using the control group on time to exit from the labor force. The DID estimates suggest at most a small effect of removing extended benefits on the time to NILF.

## 5 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. Our single-risk and competing-risk models of unemployment exits over the period 1994-2010 yielded precisely estimated reductions in exit rates from unemployment for individuals who are likely to be

eligible for UI and have been unemployed for at least 6 months (the point at which extended UI becomes available). However, because the overwhelming majority of unemployment spells never reach six months, the impact on overall unemployment duration is quite limited. We found that the observed empirical pattern of extended UI availability increased overall unemployment durations by about 1 to 2 percent. For individuals unemployed for at least 6 months, the increase is about 2 to 3 percent. Our competing risks analysis indicates that a reduction in exits out of the labor force accounts for a slightly larger share of the reduction in overall exit rates than does the reduction in exits to employment, suggesting that the reporting effect (Card, Chetty, and Weber 2007) is relatively large compared with the adverse efficiency effects of extended UI.

Our analysis focused narrowly on behavioral responses by individuals affected by extended UI, with the ultimate goal of understanding what portion of the elevated U.S. unemployment rate arises from its availability. We plan to extend our framework to conduct such an analysis in future versions of the manuscript. For now, the estimated limited impact on unemployment durations suggests that the impact on the overall unemployment rate will be small. This implication is important for assessing the likely future path of the U.S. unemployment rate after extended UI legislation expires (scheduled for January 2012). Moreover, our framework and findings ignore the aggregate demand stimulus provided by UI payments; recent research suggests that multiplier effects of normal and extended UI payments are quite large (e.g., Vroman 2010), and the reduction in unemployment from this channel is likely to exceed the increase from the behavioral channel that we examine. Further research regarding the job search and consumption behavior of UI recipients would be useful in order to address such questions.

## 6 References

- Anderson, Patricia M. and Bruce D. Meyer. 1997. *Quarterly Journal of Economics*, 112 (3, August): 913-937.
- Blank, Rebecca M. and David E. Card. 1991. "Recent Trends in Insured and Uninsured Unemployment: Is There an Explanation?" *Quarterly Journal of Economics*, 106 (4, November): 1157-1189.
- Card, David, Raj Chetty, and Andrea Weber. 2007. "The Spike at Benefit Exhaustion: Leaving the Unemployment System or Starting a New Job?" *American Economic Review*, Papers and Proceedings 97: 113-118.
- Card, David, and Phillip B. Levine. 2000. "Extended benefits and the duration of UI spells: evidence from the New Jersey extended benefit program." *Journal of Public Economics* 78: 107-138.
- Chetty, Raj. 2008. "Moral Hazard versus Liquidity and Optimal Unemployment Insurance." *Journal of Political Economy* 116(2, April): 173-234.
- Farber, Henry S. 2010. "Job Loss and Unemployment in the 21st Century: The Great Recession in Labor Market Perspective." Working Paper, Industrial Relations Section, Princeton University, December.
- Fujita, Shigeru. 2010a. "Economic Effects of the Unemployment Insurance Benefit." Federal Reserve Bank of Philadelphia Business Review, Fourth Quarter.
- Fujita, Shigeru. 2010b. "Effects of the UI Benefit Extensions: Evidence from the Monthly CPS." Working Paper No. 10-35, Federal Reserve Bank of Philadelphia, November.
- Howell, David R. and Bert M. Azizoglu. 2011. "Unemployment Benefits and Work Incentives: The U.S. Labor Market in the Great Recession." Manuscript, The New School, May (forthcoming in *Oxford Review of Economic Policy*).
- Isaacs, Katelin P., Julie M. Whittaker, and Alison M. Shelton. 2010. "Temporary Extension of Unemployment Benefits: Emergency Unemployment Compensation (EUC08)." Report RS22915, Congressional Research Service, December.

- Jurajda, S. and F.J. Tannery. 2003. "Unemployment Durations and Extended Unemployment Benefits in Local Labour Markets." *Industrial and Labor Relations Review* 56(2): 324- 348.
- Katz, Lawrence. 2010. "Long-Term Unemployment in the Great Recession." Testimony to the Joint Economic Committee, U.S. Congress. April 29, 2010.
- Katz, Lawrence, and Bruce Meyer. 1990. "The Impact of the Potential Duration of Unemployment Benefits on the Duration of Unemployment." *Journal of Public Economics* 41 (Feb.): 45-72.
- Kroft, Kory, and Matthew J. Notowidigdo. 2010. "Should Unemployment Insurance Vary with the Local Unemployment Rate? Theory and Evidence." Working paper, Yale University School of Management, November.
- Landais, Camille, Pascal Michailat, and Emmanuel Saez. 2010. "Optimal Unemployment Insurance over the Business Cycle." NBER Working Paper 16526, November. Cambridge, MA: National Bureau of Economic Research.
- Mazumder, Bhashkar. 2011. "How did unemployment insurance extensions affect the unemployment rate in 2008-10?" Chicago Fed Letter 285, April.
- Meyer, Bruce. 1990. "Unemployment Insurance and Unemployment Spells." *Econometrica* 58 (July): 757-782.
- Moffitt, Robert. 1985. "Unemployment Insurance and the Distribution of Unemployment Spells." *Journal of Econometrics* 28: 85-101.
- OECD. 2007. *Benefits and Wages 2007: OECD Indicators* (Paris: OECD).
- Valletta, Robert G., and Katherine Kuang (2010). "Extended Unemployment and UI Benefits," *FRB SF Economic Letter* 2010-12.
- Vroman, Wayne. 2010. "The Role of Unemployment Insurance As an Automatic Stabilizer During a Recession." Report, Urban Institute and IMPAQ International, July.
- Whittaker, Julie M. 2008. "Extending Unemployment Compensation Benefits During Recessions." Report RL34340, Congressional Research Service, December.