The Effects of Unemployment Benefits on Unemployment and Labor Force Participation: Evidence from 35 Years of Benefits Extensions

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Abstract

This paper presents estimates of the effect of emergency and extended unemployment benefits (EEB) on the unemployment rate and the labor force participation rate using a data set containing information on individuals likely eligible and ineligible for EEB back to the late 1970s. To identify these estimates, we examine how exit rates from unemployment change across different points of the distribution of unemployment duration when EEB is and is not available, controlling for changes in labor demand and demographic characteristics. We find that EEB increased the unemployment rate by about one-third percentage point in the most recent recession but did not affect the participation rate. In previous recessions, the effect of EEB on the unemployment rate was even smaller.

JEL classifications: J6, E24.

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1. Introduction

During the 2008-2009 recession, U.S. unemployment insurance (UI) benefits were extended to unprecedented levels, with UI duration increasing from 26 weeks—the regular duration—to as much as 99 weeks in some states, prompting a lively debate in policy and academic circles about the adverse effects of such extensions on the search behavior of job seekers and thus on the unemployment and labor force participation rates.[1]

While recent studies have focused on the extension of UI benefits during the recent crisis, the extension of benefits is not specific to the 2008-2009 recession. In fact, the duration of UI benefits was increased in all recessions since the mid-70s.

In this paper, we exploit data on such UI benefits extensions going back to the mid-70s to estimate the effect of UI duration extension, also called Emergency and Extended unemployment Benefits (EEB), on the labor market, and more specifically the unemployment and participation rates.

We identify the effects of EEB from variations in the exit rates out of unemployment across similar individuals as EEB availability changes, controlling for the state of the business cycle. Importantly, we allow individuals with different unemployment duration to react differently to the availability of EEB. We then develop a mapping of our estimates of the effect of EEB on unemployment exit rates into estimates of the effect of EEB on unemployment and labor force participation, and we find that EEB increased the unemployment rate by one-third percentage point in the most recent episode and had hardly any effect on the participation rate.

To identify the effects of EEB, we consider individuals with similar observable characteristics with different lengths of unemployment spells across periods and states where EEB is and is not available. We find that EEB reduces substantially the Unemployment to Employment (U-E) transition rate of individuals receiving or about to receive EEB by about 16-17 percent. The effects of EEB on the Unemployment to Nonparticipation (U-N) transition rate are smaller and more uncertain, and only individuals receiving EEB benefits appear to be affected, with their U-N rate falling by about 7 percent. Importantly however, we find that EEB has little adverse effect on exit rates for individuals with less than 20 weeks of unemployment *or* for individuals with unemployment duration above the maximum duration of EEB.

To map our estimates of the effect of EEB on unemployment exit rates into estimates of the effect of EEB on unemployment and labor force participation, we use a first order approximation of steady state unemployment and labor force participation rates (Shimer, 2012) to decompose the unemployment and the labor force participation rates into linear functions of hazard rates of transitions across labor force states.

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[1] See e.g., Farber and Valletta (2013), Rothstein (2011), Bradbury (2014), Mazumder (2010), Fujita (2011), and Nakajima (2012).

Despite sizable effects of EEB on the exit rate of job seekers receiving or about to receive EEB, we find that EEB increased the unemployment rate by only one-third percentage point in the most recent episode and had hardly any effect on the participation rate. The effects in previous recessions were even smaller.

Our transparent mapping of the effects of EEB on unemployment, together with our estimates of how EEB affects different individuals with different unemployment durations, enable us to understand why the effects of EEB are relatively modest. The main reason is that EEB only affects eligible individuals with relatively long unemployment durations (greater than 18 weeks), and such individuals represent a relatively small fraction of the unemployment population (even in the recent recession where unemployment duration reached record highs, their share stood only at about 25 percent). In contrast, most of the unemployed are either simply ineligible for unemployment insurance (around 40 percent of the unemployed in recessions are non job losers), or, if eligible (job losers), respond little to the availability of EEB because they have either not been unemployed for very long (about one-third of the unemployed) or have durations above the maximum duration of EEB (a little over 5 percent of the unemployed). Regarding the participation rate, the effect is very small because (i) the average U-N rate is only modestly affected by EEB, and (ii) the average UE rate (which is somewhat affected by EEB) has only an indirect (and thus modest) effect on the participation rate.[2]

Our paper relates to a large literature on the effects of Unemployment Insurance on job seekers' unemployment duration. We contribute to the literature in several ways.

First, our identification strategy allows us to clearly understand why the effect of EEB on the aggregate unemployment rate has been estimated to be small (Fujita, 2011, Farber and Valletta 2013, Rothstein 2011), despite evidence that extensions of unemployment insurance do have a substantial effect on individuals' exit rate (Meyer 1990). Our identification strategy is similar to strategies used recently in the literature—Farber and Valletta (2013) and Rothstein (2011)—and our estimates of the effect of EEB on the unemployment rate are consistent. However, we rely on a dimension of the data (differences in effect by duration) not fully exploited previously, which allows us to estimate the effect of EEB for individuals across the whole spectrum of unemployment durations. As explained previously, we estimate that benefits extensions do have a substantial effect on an individual's exit rate, but *only* on individuals receiving or about to be receiving extended benefits. That latter estimate is the key to understand the small aggregate effects of EEB.

Second, we construct a data set that uses information on eligibility for extended and federal benefits across states and time and match this with individual CPS data on labor force transitions

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[2] Movements in the job finding rate have an effect on the participation rate because of a composition effect: Unemployed workers are much more likely to leave the labor force than employed workers. By raising the number of unemployed workers relative to the number of employed workers, a lower job finding rate increases the number of individuals who leave the labor force, and the participation rate goes down.

back to 1976. Our data set enables us to both compare the effects of EEB across different episodes and show how EEB has affected cyclical movements in the unemployment rate over time. Because our dataset spans several recessions, we can contrast our findings with those of an earlier, influential, literature which estimated the effects of EEB from recessions in the 1970s and 1980s, e.g. Meyer (1990) and Moffitt and Nicholson (1982). Two conclusions stand out. First, the effect of EEB on unemployment exit rates has been generally stable over time, allowing us to compare current estimates to studies of EEB in the 1970s and 1980s. Second, while previous studies suggest much larger effects than we estimate—Meyer (1990) and Moffitt and Nicholson (1982), and Nakajima, (2011)—we argue that extrapolating estimates from studies of previous episodes to calibrate EEB effects in the most recent episode can be misleading, because, unlike what is typically assumed, individuals nearing the exhaustion of regular UI benefits behave very differently from individuals who have exhausted their benefits. As a result, calibrations of EEB effects based on estimates such as those in Meyer (1990) that assume that individuals about to exhaust benefits are similar to those who have exhausted their benefits can substantially overestimate the effect of EEB. Third, the paper proposes a method to easily and transparently map the effect of EEB on unemployment hazard rates into the effect of EEB on unemployment. The ease and transparency of our approach allows us to compare our results to other estimates and account for differences. Finally, we estimate the effect of EEB on the labor force participation rate, a topic not addressed in previous papers.

The next section describes our data and method of identification. The third section presents estimation results. The fourth section maps these results to estimates of the effect of EEB on the unemployment and labor force participation rates. The fifth section compares results to others in the literature, and the sixth section concludes.

2. Data and Identifying the Effect of EEB

Data

To estimate the effect of EEB on unemployment and labor force participation, we combine two data sets; (i) a newly created dataset containing information on the maximum duration of regular UI benefits and EEB benefits by state and month over 1976-2012, and (ii) the CPS containing information on individuals' characteristics and labor force status across consecutive months as well as the unemployment duration of job seekers.

To construct data on the maximum duration of regular UI benefits and EEB by state and month, we compiled records published by the Employment and Training Administration (ETA). Our data contain information on the maximum allowable duration of combined EEB and regular UI benefits in a state in each month from 1976 through 2013. The appendix contains details on UI laws that have provided for emergency and extended unemployment insurance benefits.

Specification

We posit that the probability of transiting from unemployment to employment (or to out of the labor force) depends on the characteristics of an individual i at time t, X_{it}, and on the person's eligibility for extended unemployment benefits, τ.

$$y_{i,t}^{j} = f\left(X_{i,t}, \tau_{i,t}; \theta\right), \; j = U - E, U - N \tag{1}$$

where $y_{i,t}^{UE}$, for example, equals 1 if an individual transitions from unemployment (U) to nonparticipation (E) and 0 otherwise, and θ is a vector of aggregate parameters, capturing for instance aggregate labor demand conditions. Equation (1) can be derived from a simple model that relates the decision of whether to search for employment or not to the benefits and costs of searching versus not searching.[3] EEB, by providing a payment to searchers, is generally predicted to increase search on the extensive margin (i.e. reduce the hazard from unemployment to nonparticipation), but also to raise reservation wages and, thus, reduce the hazard from unemployment to employment.

When EEB benefits are available (typically during and/or immediately after a recession), jobless individuals become eligible for EEB after they have exhausted their regular UI benefits. Thus, to distinguish between individuals who are eligible for EEB and individuals who are not, i.e, $\tau_{it} = 0$ or $\tau_{it} = 1$, we combine information on individuals' unemployment duration with information on the availability of EEB in a given state at time t. We set $\tau_{it} = 1$ (EEB available) when the total length of benefits duration (including EEB and regular benefits) exceeds the unemployment duration of an individual.

Identification of the effect of EEB

To estimate the effect of EEB on an individual's exit rate from unemployment, we would ideally observe the same person under the same circumstances twice: once when the person is eligible for EEB and once when the person is not, and estimate the EEB effect as the difference in exit behavior

$$\sum_{i}\left(f\left(X_{i,t}, \tau_{i,t} = 1; \theta\right) - f\left(X_{i,t}, \tau_{i,t} = 0; \theta\right)\right) \tag{2}$$

Because such data are not available, we instead compare different individuals, some who are eligible for EEB and others who are not, while also controlling as best we can for the effect of individuals' characteristics (as well as aggregate labor demand conditions) on their probability of exiting unemployment. Specifically, we estimate

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[3] See, for example, Mortensen (1977) or Rothstein (2011).

$$\sum_{i \in N_1} f\left(X_{i,t}, \tau_{i,t} = 1; \theta\right) - \sum_{i \in N_2} f\left(X_{i,t}, \tau_{i,t} = 0; \theta\right)$$

(3)

where N_1 denotes individuals for whom EEB are available and N_2 denotes individuals for whom EEB are not available.

We control for individual characteristics and aggregate labor demand conditions in several ways.

First, we only include permanent job losers in our sample, thus limiting it to a relatively homogenous group. We include only job losers because in most states, one must have had a prior job and been laid off from that job to be eligible for unemployment benefits.[4] This excludes job quitters and entrants to the labor force. We also exclude individuals who are 20 years old or younger because, given the low employment rate and high transitions rates out of employment for this group, it is unlikely that this group will have high eligibility rates for unemployment insurance benefits. We also exclude temporary job losers. Because six months is the typical maximum duration of regular UI benefits, individuals on temporary layoff, who expect to be recalled to their former job within 6 months, may be less likely to be affected by the availability of EEB than permanent job losers.[5] In the appendix, we present estimates of the effect of EEB on temporary job losers and show it is quite similar to the effect on permanent job losers.

Second, we include several demographic variables in X_{it}, including gender, marital status, age, education, occupation, former industry of employment, and state of residence. Another crucial characteristic of individuals that affects the propensity to exit is the duration of unemployment. Specifically, as a spell length increases, the propensity to exit to employment declines, and the propensity to exit the labor force generally increases.[6] It is important to control for the tendency for exit hazards to change with duration so as not to confuse this effect with that of EEB. We do this by allowing duration to affect exit hazards non-parametrically.

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[4] Farber and Valletta (2013) match individuals identified as job losers in the monthly CPS with the same individuals in the subsequent year's March CPS, which asks about receipt of unemployment benefits. They find a strong correspondence between job loser status and subsequent reporting of receipt of UI benefits. If job loser status and eligibility for UI benefits are not perfectly correlated, then measurement error may bias our estimates of the effect of EEB downward if the extent of the discrepancy between UI eligibility and job loser status is related to the cycle, or more specifically to the implementation of EEB programs.

[5] In addition, Katz and Meyer (1990) illustrate that the evolution of exit hazards out of unemployment can be quite different for permanent and temporary job losers.

[6] The effect of duration on exit rates may be due to heterogeneity in exit rates and individuals' gradually learning about their true reemployment prospects. Under this hypothesis, individuals with low reemployment prospects become a higher fraction of the unemployment pool at higher durations, pushing down average exit rates to employment. Also, if these individuals learn about their relatively poor prospects, exit rates to out of the labor force should eventually increase. Alternatively, it could be that time spent unemployed, itself, lowers reemployment prospects so that the longer an individual is unemployed, the less likely he will find a job. Realizing this, as an individual's spell of unemployment increases in length, he eventually becomes more likely to drop out. Finally, it could be that many individuals who lose a job have very good job market contacts and find a new job fairly easily. As duration increases, this type of individual becomes less prevalent, leading exit rates to employment to decrease.

Finally, because EEB typically becomes available during recessions and becomes unavailable once a recovery is firmly established, it is also important to control for the influence of changes in aggregate economic conditions on the propensity to exit unemployment. We do this using the ratio of aggregate vacancies to aggregate unemployment.[7] This measure has been shown by Barnichon and Figura (2013) to be a very good predictor of average exit rates to employment. In addition, we control for local labor market conditions by including the unemployment rate in an individual's labor market segment, defined as four broad occupation groups (production workers, service workers, sales workers or clerks, and professionals) within the 50 states and the District of Columbia.[8]

To summarize our identification strategy, our estimator of the effect of EEB simply compares the behavior of individuals when EEB is available to the behavior of similar individuals when EEB is unavailable. In order for our estimate to correctly identify the effect of EEB on unemployment exit rates, it must be that our controls for aggregate demand and demographic influences are sufficient, so that the only reason exit rates for likely affected job losers change between periods when EEB is available and when it is not available, conditioning on an individual's duration, is the influence of EEB.

3. Results

This section presents our results on the effects of EEB. We find that EEB reduces the U-E transition rate of job losers about to exhaust regular UI benefits and of job losers receiving EEB benefits by a substantial 16-17 percent. The effects of EEB on the U-N transition rate are smaller and more uncertain. Only individuals receiving EEB benefits (group 4) appear to be affected, with their U-N rate falling by about 7 percent.

Preview of results

Before presenting the results of the complete specification with controls for individual characteristics, we first illustrate the idea behind our identification of the effect of EEB by comparing the unemployment exit rate by unemployment duration when EEB are available with the unemployment exit rate when EEB are not available.

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[7] We use the adjusted help wanted index computed by Barnichon (2010) as our measure of aggregate vacancies.
[8] In the appendix, Table A1 shows the number of observations and the average characteristics of individuals in the 5 groups when EEB is and is not available. It suggests that controlling for demographic characteristics is necessary in order to avoid biasing estimates of the effect of EEB, as some individual characteristics within groups vary across periods when EEB is and is not available. For instance, highly educated individuals are a much larger share of the unemployed during periods when EEB is available. In addition, males are more likely to have high duration unemployment and this is particularly true in periods when EEB is not available.

Table 1. Categorizing Permanent Job Losers when EEB is available

Group	Characteristics	
	Unemployment duration	Affected by EEB?
1	less than or equal to 8 weeks	No, or small effect
2	greater than 8 weeks, less than 18 weeks	Maybe
3	greater than or equal to 18 weeks, less than the maximum allowable duration of regular UI benefits	Yes
4	Figures 1a and 1b: greater than the maximum allowable duration of regular UI benefits, less than 60 weeks. Estimation: greater than the maximum allowable duration of regular UI benefits, less than the maximum allowable duration of EEB plus regular benefits	Yes
5	Figures 1a and 1b: greater than 60 weeks. Estimation: greater than the maximum allowable duration of regular plus EEB benefits	No

Figures 1a and 1b illustrate graphically our estimator of the effect of EEB on job losers'
unemployment to employment transition rate, U-E, and on job losers' unemployment to
nonparticipation transition rate, U-N. As described in Table 1, we split unemployed individuals
into 5 groups, determined by the length of their current unemployment spell. Groups 3 and 4
contain individuals who likely receive or may be about to receive EEB benefits.[9] In contrast,
groups 1, 2 and 5 contain individuals who are less likely to be affected by EEB, either because
they are early on in their receipt of regular UI benefits and thus unlikely to be significantly

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[9] Group 4 includes individuals with unemployment durations greater than the maximum allowable number of weeks
of regular UI benefits and less than the maximum allowable duration of regular plus EEB benefits. Individuals in
this group are likely receiving EEB benefits. Group 3 includes individuals with unemployment durations greater
than 18 weeks, i.e., individuals who are less than 8 weeks away from the expiration of regular UI benefits.
Individuals in this group are likely not yet eligible for EEB, but because they are nearing the termination of their
regular benefits, their behavior may be affected by the knowledge that EEB will be available if their regular benefits
are exhausted. The 18 weeks threshold is roughly consistent with Meyer (1990), who finds that a change in
behavior among UI recipients starts being noticeable six weeks prior to benefits expiration.

affected by the prospect of EEB (groups 1 and 2),[10] or because their unemployment duration exceeds the maximum duration of EEB (group 5). As noted in Table 1, to aid in the graphical exposition, in Figures 1a and 1b we simplify the categorization of individuals by using fixed unemployment duration cutoffs to place individuals in the five groups. But later when we estimate the effect of EEB on exit rates, we will use cutoffs that vary with the maximum allowable duration of both regular benefits and EEB in a state in a given month.

To construct Figures 1a and 1b, we first compute average U-E and U-N exit rates by week of unemployment duration for individuals when EEB is available in their state and for individuals when EEB is unavailable. These average exit rates are shown by the two lines in the top panel of Figures 1a and 1b. The basis for our estimator is the difference between the solid EEB line and the dashed non EEB line. Without controlling for demand, we can't properly gauge the magnitude of the effect, but if EEB is affecting exits, then the downward shift in exit hazards for groups 3 and 4 should be, and is, larger than for the groups that should be less affected or unaffected by EEB. The bottom panels of Figures 1a and 1b show the difference between the two lines in the top panel, and the horizontal bars show the average difference in exit rates for each group. In the bottom panel, it is obvious that the exit rate differences for groups 3 and 4 are larger than for the other groups. In sum, preliminary results without demand or demographic controls suggest that the change in exit hazards across periods when EEB is and is not available is significantly more negative for individuals likely receiving EEB benefits or about to exhaust regular UI benefits.

Specification and sample period

We now turn to our full specification which controls for a wide range of individual characteristics and changes in labor demand. In addition, we estimate EEB availability more rigorously by using state and time specific maximum allowable durations of EEB (rather than the fixed duration cutoffs used in Figures 1a and 1b) to determine EEB eligibility.

We estimate equation (1) using a logit specification and cluster standard errors for individuals in the same state at the same time. We use data from 1976 through 2013. As noted at the bottom of Table 2, our specification includes a measure of aggregate labor market tightness, unemployment rates for 204 state/occupation labor market segments, dummy variables for 10 age, 3 education, 13 industry, 4 occupation, and 100 unemployment duration classifications, as well as dummy variables for gender, marital status and state of residence.[11]

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[10]Both Meyer (1990) and Rothstein (2011) illustrate that job losers far removed from the expiration of benefits are likely to respond little to increases in the maximum duration of benefits because the probability of reaching the maximum duration of regular UI benefits is relatively small at the beginning of an unemployment spell.

[11] We also considered specifications that allow for interactions between demographic controls. For example, we included dummy variables constructed by interacting 2 gender groups, 2 marital status groups, 3 education groups, and 10 age groups. When we add these 120 dummies to our main specification, our estimates of the effect of EEB are little changed.

Exit Hazard from Unemployment to Employment

Starting with the effect of EEB on the Unemployment to Employment (U-E) transition rate, Figure 2a shows panels illustrating the effect of observable characteristics (other than eligibility for EEB) on the U-E rate. Exit rates decline as duration increases, but at a decreasing rate. Exit rates decrease in age, are higher for men and married individuals, increase in education and are highest for production occupations. Estimation results (not shown) reveal that, as expected, greater labor market tightness, and lower labor-market-segment-level unemployment, lead to higher transition rates to employment.

Estimates of the effect of EEB on U-E exit rates are shown in the first column of Table 2. Estimates for both group 3 and group 4 are negative and significant and suggest that EEB reduces exit rates to employment by about 17 percent for individuals about to exhaust their regular UI benefits and 16 percent for individuals receiving EEB benefits.[12] It also appears that EEB has a small negative effect on the U-E exit hazard of individuals at the beginning of their unemployment spells.

Exit Hazard from Unemployment to Nonparticipation

We now turn to the effect of EEB on the Unemployment to Nonparticipation (U-N) transition rate. First, the estimated effects of observable characteristics (other than eligibility for EEB) on the U-N exit rate are shown in Figure 2b. As shown in the top panel, the U-N exit rate generally rises as duration increases. As shown in the remaining panels, exit rates initially decline in age before increasing around age 45. Exit rates are higher for females, unmarried individuals, and individuals with less education and are lowest in professional occupations, as one might expect based on priors regarding the relative labor force attachment of these groups.

The second column of Table 2 shows estimates of the percent change in exit hazards when EEB is in effect for our 5 groups. Group 4, containing individuals likely receiving EEB, is the only group likely receiving UI or EEB benefits estimated to be significantly negatively affected by EEB. The estimate for group 4 suggests that EEB decreases exit rates for individuals receiving EEB by close to 7 percent.

Interestingly, we also estimate that EEB increases exit rates for groups 1 and 2, while reducing them for group 5. Overall, this pattern is difficult to interpret. The positive estimated coefficients for groups 1 and 2 suggest that some other factor may be increasing U-N rates in periods when EEB is available. If this factor affects U-N rates for all unemployed workers, not just those who have experienced short duration unemployment spells, then the estimated effect of

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[12] Theory would suggest that the effect of EEB on individuals nearing the maximum duration of their regular UI benefits is smaller than the effect on individuals already having exhausted their regular benefits. Thus, these results suggest that individuals nearing the termination of their regular benefits may be different than other individuals, perhaps because they having been waiting until the end of their benefit period to search for a job, see Katz and Meyer (1990).

EEB on group 4 may be larger than the -7 percent we estimate. On the other hand, the negative estimated coefficient for group 5 suggests that some other factor may be pushing down the U-N rates of long-duration unemployed when EEB is available, implying that the estimated coefficient on group 4 may overstate the negative effect of EEB. Thus, there is considerable uncertainty about the effect of EEB on U-N rates. Nevertheless, because the estimated effect is relatively small under either of the above scenarios, and because, as we show later, the effect of U-N rates on the unemployment rate is also relatively small, the amount of uncertainty about the effect of EEB on U-N rates does not translate into great uncertainty about the effect of EEB on the unemployment rate.

Separate estimates by EEB episode

Recent papers on the effect of EEB have used data only for the most recent recession.[13] However, as described in detail in the appendix, EEB benefits have been available during and after each major recession since the early 1970s. This raises the question of whether recent studies' estimates of the effect of EEB are influenced by the unusual circumstances surrounding the Great Recession of 2008-2009, or whether, instead, they are representative of earlier effects. As described below, we find that the 2008-2009 estimates are in line with estimates drawn from previous episodes.

Table 3 shows estimated effects for four different periods. Except for the early 80s recession, the estimated effects of EEB on U-E transitions are similar (and not statistically different) across the past three recessions: individuals about to exhaust regular UI benefits or receiving EEB see their exit rates fall by about 16 percent.

The estimated effect of EEB on U-N transitions is also fairly similar across episodes. In general, EEB only has negative effects on individuals about to exhaust regular UI benefits or individuals likely receiving EEB benefits, with effects estimated to be larger for the latter group. Estimated effects range from 0 to 12 percent and tend to be larger after the mid-1980s.

4. Mapping Coefficient Estimates to EEB Effects on the Unemployment Rate and Labor Force Participation Rate

We now turn to mapping our estimates of the effect of EEB on exit rates into estimates of the effect of EEB on unemployment and labor force participation rates. Using a stock-flow accounting model of the labor market, we show that the effects of EEB on unemployment and participation depend on three factors: (1) the effect of EEB on workers' transition rates, (2) the distribution of job seekers across unemployment duration bins (recall that the effect of EEB on workers' transition rates is strongest for the unemployed receiving or about to receive EEB), and (3) how changes in a given transition rate (say U-E or U-N) affect the unemployment and participation rates.

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[13] Farber and Valletta (2013) are an exception as they also use data from the early 2000s.

We find that EEB have a small overall effect on unemployment. Importantly, this is not because the transition rates are not affected (on the contrary, recall that EEB lowered the UE transition rate by about 16 percent for job seekers receiving or about to receive EEB), but, instead because the groups who are affected represent a small fraction of all unemployed. Indeed, the unemployment pool includes mostly individuals with short unemployment duration, and EEB has little effect on these individuals.

The following system of differential equations expresses changes in the stocks of the three labor force states as functions of the instantaneous hazard rates of moving between states, the λ^{IJ} s.

$$\dot{U}_t = -\left(\lambda_t^{UE} + \lambda_t^{UN}\right)U_t + \lambda_t^{EU}E_t + \lambda_t^{NU}N_t$$

$$\dot{N}_t = -\left(\lambda_t^{NE} + \lambda_t^{NU}\right)N_t + \lambda_t^{UN}U_t + \lambda_t^{EN}E_t$$

$$\dot{E}_t = -\left(\lambda_t^{EU} + \lambda_t^{EN}\right)E_t + \lambda_t^{UE}U_t + \lambda_t^{NE}N_t \tag{4}$$

One can solve for the steady state value of each of the stocks as functions of the six hazard rates. The expressions for the unemployment rate and the labor force participation rate are

$$u_t^{ss} = \frac{s_t}{s_t + f_t} \tag{5}$$

$$lf_t^{ss} = \frac{s_t + f_t}{s_t + f_t + \lambda_t^{EU}\lambda_t^{UN} + \lambda_t^{UE}\lambda_t^{EN} + \lambda_t^{UN}\lambda_t^{EN}}$$

$$s_t = \lambda_t^{EU} + \lambda_t^{EN}\left(\frac{\lambda_t^{NU}}{\lambda_t^{NU} + \lambda_t^{NE}}\right) \tag{6}$$

$$f_t = \lambda_t^{UE} + \lambda_t^{UN}\left(\frac{\lambda_t^{NE}}{\lambda_t^{NE} + \lambda_t^{NU}}\right)$$

We construct estimates for the six hazard rates for each point in time by first using data on stocks of individuals in unemployment, employment, and out of the labor force and measures of transitions between these stocks to produce estimated labor force transition rates over time. We then adjust these transition rates for time aggregation bias using the method outlined by Shimer (2012). As noted by Shimer (2012), because of the relatively high transition rates between labor force states, the implied steady state measures track the actual levels of the unemployment rate and labor force participation rate quite closely.

Next, we take a first order approximation of expressions (5) and (6) to relate the deviation of the steady state unemployment rate and participation rate from the values they would have taken in

the absence of EEB to percent deviations of labor force flows from the values they would have taken absent EEB. We obtain

$$u_t^{EEB} - u_t = \alpha_{UN}^U \log\left(\frac{\lambda_t^{UN,EEB}}{\lambda_t^{UN}}\right) + \alpha_{UE}^U \log\left(\frac{\lambda_t^{UE,EEB}}{\lambda_t^{UE}}\right) \tag{7}$$

$$lf_t^{EEB} - lf_t = \alpha_{UN}^{lf} \log\left(\frac{\lambda_t^{UN,EEB}}{\lambda_t^{UN}}\right) + \alpha_{UE}^{lf} \log\left(\frac{\lambda_t^{UE,EEB}}{\lambda_t^{UE}}\right) \tag{8}$$

where $\lambda_t^{IJ,EEB}$ is the hazard rate for transition *I-J* at time *t* after taking into account the effect of EEB, λ_t^{IJ} is what the hazard rate for transition *I-J* at time *t* would have been if EEB were unavailable, and the α_{IJ} s are the first order coefficients describing the effect of a percent change in λ^{IJ} on the unemployment rate or participation rate.[14]

To construct an estimate of $\frac{\lambda_t^{IJ,EEB}}{\lambda_t^{IJ}}$, we decompose jobseekers into 5 groups, as described

previously, and we estimate the ratio $\frac{\lambda_t^{IJ,EEB,i}}{\lambda_t^{IJ,i}}$ for each of these five groups *i* and then sum these

estimates weighting each group by their unemployment share. Using *IJ* to denote the monthly transition probability we observe in the micro data and for which we estimate an effect of EEB, the hazard rate λ_t^{IJ} is given by $\lambda_t^{IJ} = -\ln(1-IJ)$, and we have

$$\frac{\lambda_t^{IJ,EEB,i}}{\lambda_t^{IJ,i}} = \frac{-\ln\left(1 - IJ_t * \left(1 + \beta^{IJ,EEB,i}\right)\right)}{-\ln\left(1 - IJ_t\right)}$$

where $\beta^{IJ,EEB,i}$ is our estimate from section 3 of the percent change in the transition rate *I-J* induced by EEB for individuals in group *i*. We then use CPS data on the number of permanent and temporary job losers by duration to estimate κ_t^i, *i=1, ...,4*, the fractions of the unemployed accounted for by group *i*.[15] (In the appendix we show that the effect of EEB on job losers on temporary layoff is similar to the effect of EEB on permanent job losers). Putting these pieces together, we construct $\frac{\lambda_t^{IJ,EEB}}{\lambda_t^{IJ}}$ as

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[14] Theoretically, EEB may affect all six labor force flows, but practically we did not find any significant effect on flows into the labor force or out of employment.

[15] To check that this estimate is reasonable, we compare the number of job losers with durations less than the maximum allowable duration of regular and EEB benefits from the CPS to published data on the number of UI and EEB beneficiaries published by the Department of Labor. The numbers match quite well.

$$\frac{\lambda_t^{IJ,EEB}}{\lambda_t^{IJ}} = \sum_i \frac{\lambda_t^{IJ,EEB,i}}{\lambda_t^{IJ.i}} \kappa_t^i = \sum_i \frac{-\ln\left(1 - IJ_t * \left(1 + \beta^{IJ,EEB,i}\right)\right) * \kappa_t^i}{-\ln\left(1 - IJ_t\right)}$$

(9)

Equations (7), (8), and (9) provide a transparent view of the factors that influence the effect of EEB on the unemployment and participation rates. These factors include: (1) the size of each group, captured by κ_t^i (2) the percent change in transition *I-J* caused by EEB, $\beta^{IJ,EEB,i}$, and (3) the first order coefficient describing the effect on the unemployment (participation) rate of a percent change in hazard *I-J*, α_{IJ}^u (α_{IJ}^{lf}).[16]

Table 4 shows the effect of EEB on the unemployment rate decomposed into the three factors (1) to (3). Panel 1 of the table shows the shares of unemployment accounted for by the five groups, the κ^i s.[17]

Panel 2 of the table shows estimates of the effect of EEB on exit rates by group. To get an upper bound of the adverse effects of EEB on exit rates out of unemployment, we impose the restriction that the effect of EEB on hazard rates is always less than or equal to 0.[18]

Panel 3 of Table 4 shows the average value of the first order coefficients mapping the percent change in hazard *I-J* to the change in the unemployment rate, the α_{IJ} s, over periods when EEB is available. It is worth noting that α_{UE} is nearly twice the level of α_{UN}. The effect of U-N rates on unemployment is relatively small because the direct effect, in which an increase in U-N lowers the unemployment rate, is compensated by the fact that a nonparticipant has a higher probability to flow into unemployment than into employment, which raises unemployment.

Panel 4 combines the information in the first 3 panels and reports the effect of EEB on the unemployment and labor force participation rates through the U-N channel, through the U-E channel, and through both channels combined. In the recent recession, the maximum effect of EEB on the unemployment rate is one-third percentage point, while the effect on the participation rate is essentially 0, as a small boost from a reduction in U-N transitions is more than offset by a small decline due to a reduction in U-E transitions.[19]

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[16] Our estimates of the effect of EEB on the unemployment and labor force participation rates assume that there is no secondary effect of EEB on exit hazards through the influence of EEB on firms' labor demand. Other estimates of the effect of EEB on unemployment, such as Rothstein (2011), are based on a similar assumption.

[17] We assume coefficient values of 0 for the fifth group because our prior is that EEB does not affect this group. Because this group is very small, using estimated coefficients for this group would not significantly change our estimates.

[18] Specifically, we set the positive estimates for the effect of EEB on U-N hazards for the first two groups to 0.

[19] The negative effect of a decrease in the U-E hazard on the participation rate depends on the implicit assumption that flows to out of the labor force are more likely, conditional on an individual being unemployed than conditional on an individual being employed. If, counterfactually, these probabilities were the same, then one could show using equation (6) that a change in the U-E hazard would have no effect on the participation rate.

The top panel of Figure 3 shows the total effect of EEB on the unemployment rate over time. In 2010 the estimated effect of EEB on the unemployment rate is one-third percentage point. The effect of EEB on the unemployment rate in previous episodes, between 0.1 and 0.2 percentage point, is considerably smaller than in the current episode. The lower panel of Figure 3 shows the estimated effect of EEB on the participation rate. It is quite small.

Overall, we find that EEB has a small effect on the unemployment rate. The most important reason is that most of job losers are relatively unaffected by EEB: Most of the unemployed eligible for EEB are concentrated in groups 1 and 2, for which EEB have little effect (panel 1 of Table 4). In other words, EEB could only have large effects on the unemployment rate if it had large effects on the U-E and U-N flows of individuals early in their unemployment spells. But as shown in Table 4, this is not the case.

Our results appears to contrast with earlier research (Moffitt and Nicholson (1982) and Meyer (1990), whose estimates seem to imply quite large effects of EEB on the unemployment rate, a point which we address in section 5 below.

5. Comparison with the Literature

We now put our findings in the context of other research on the effects of EEB on unemployment. In doing so, we seek to answer two questions: (1) is our finding of a relatively moderate effect of EEB on U-E transition rates consistent with earlier studies, which appeared to find substantial effects, and if so, why or why not? and (2) why are our estimates of the effect of EEB on the unemployment rate different from recent papers on the subject?

Meyer (1990) finds that exit hazards from unemployment increase dramatically as individuals approach the maximum duration of benefits. Specifically, he finds that exit hazards increase fourfold as an individual moves from 6 weeks prior to benefit expiration to 1 week prior to expiration, while changing relatively little prior to 6 weeks before benefit expiration. If exit hazards after the expiration of benefits are similar to those just prior to expiration, then Meyer's estimates imply that exit hazards after expiration are about four times greater than exit hazards for individuals receiving benefits but not yet near the expiration point. Thus, one way to extrapolate Meyer's results would be to assume that EEB reduces exit hazards by 75 percent, the difference between average hazards early on in unemployment spells, when the effect is small, and the estimated hazard rate right before benefit exhaustion. However, such an extrapolation may be misleading. Figure 4 illustrates why.

The thin solid line labeled "EEB Not Available" shows how one would expect the exit hazard to change as an individual nears the period when benefits expire in the absence of EEB, see Meyer (1990). As the time spent unemployed approaches 26 weeks, the exit rate increases until the point of benefit exhaustion (26 weeks) is reached, after which the hazard rate remains

unchanged. Meyer (1990) finds that exit rates are relatively flat in the early portion of a benefit period, region A in the figure, but then rise steeply with 1 to 2 months of benefits remaining, region B. As a result, an increase in the maximum duration of benefits would be predicted to reduce exit rates for individuals near the expiration of regular benefits. This is illustrated by the difference between the solid line in region B—representing exit behavior when EEB benefits are not available and the dashed line, representing exit behavior when benefits are available. An extrapolation of Meyer's (1990) results assumes that the effect on individuals that have already exhausted regular benefits, region C, is the same as the effect at 26 weeks or that the hazard in the absence of EEB, is the horizontal line labeled "Extrapolation of Meyer (1990)" in region C. However, this is likely not a good assumption because we find that the effect of EEB on exit hazards for individuals in region C is considerably smaller that this extrapolation would imply and similar to the average exit rate of individuals in region B. Our estimate is represented by the thick solid line labeled "Our Estimated Effect". The difference between this line and the dashed line represents our estimated effect of EEB on individuals in region C. Our estimated effect in region C is well below the extrapolated effect, the difference between the dashed line and the solid line labeled "Extrapolation of Meyer (1990)".

Another earlier paper with results seemingly in contrast to ours is Moffitt and Nicholson (1982). They find that an extra week of benefits leads, on average across all individuals, to an increase of 0.1 week of unemployment. Card and Levine (2000) estimate similar results for a one-time increase in benefits in New Jersey in the 1990s. Applying this estimate to the most recent episode of EEB would imply that the average unemployment duration of individuals eligible for UI increased by nearly 6 weeks. While it is difficult to map this precisely into the effect of EEB on exit hazards, given that the mean duration of unemployment prior to the advent of EEB was around 17 weeks, it would seem to imply a reduction in the unemployment exit hazard of at least 35 percent (assuming that unemployment duration equals 1 divided by the exit hazard), considerably larger than the effect we estimate, especially once one takes into consideration that this estimate applies to all eligible unemployed workers.[20] However, Moffitt and Nicholson (1982) caution that their sample, composed of approximately 1000 individuals observed over 1975-1977, who received federal supplemental benefits, is unlikely representative of the universe of individuals eligible for UI, as it includes mainly individuals who have suffered relatively long periods of unemployment.[21] For different reasons, it is also likely misleading to extrapolate the Card and Levine (2000) estimate to the most recent episode of EEB. This is because it is not

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[20]The rule of thumb mapping from average unemployment duration to exit hazard that we use requires data on the average expected duration of all unemployment spells. However, we only observe the average duration of ongoing spells of unemployment, which will be greater than the duration of all spells because individuals with long duration spells are more likely to be sampled. Thus, the reduction in exit hazards implied by a 6 week increase in the average duration of a spell of unemployment is likely greater than 40 percent.

[21]Moreover, it is not clear how robust the Moffitt and Nicholson (1982) result is to small changes in specification. For example, table 4 of their paper suggests that changes in the assumption needed to empirically identify individuals at the kink point in the budget sets can lead to dramatic reductions in the estimated effect of an extra week of benefits on mean unemployment durations.

clear that the relationship between benefits and average durations of unemployment is linear. In particular, the estimates in Meyer (1990) and the estimates in this paper, suggest that individuals nearing the end of their regular benefits respond strongly to an extra 6-8 weeks of benefits, but would not respond very much to any further increase in benefit durations. Thus, applying estimates of the effect of an extra week of benefits in a context in which the maximum duration of benefits was changed relatively little to a highly unusual situation in which the length of benefits increased by over 60 weeks in many states may be misleading.[22]

Turning to more recent studies, our paper relates to the work of Fujita (2011), Rothstein (2011), Nakajima (2012), Farber and Valletta (2013) and Bradbury (2014). Similar to Rothstein (2011) and Farber and Valletta (2013), we find that EEB does reduce exit hazards to nonparticipation. Unlike these studies, however, we find that EEB also has a sizable effect on exit hazards to employment.

Fujita (2011) estimates a ¾ percentage point effect of EEB on the unemployment rate in the most recent episode. He uses a regression discontinuity approach to identify the EEB effect as the sharp change in unemployment exit behavior of prime-age males at 26 weeks during the period when EEB was available in the most recent episode compared to prior to the recession. This is similar to our approach except that Fujita (2011) uses the abrupt change in behavior at an unemployment duration of 26 weeks, whereas we compute the change in average exit rates for all individuals likely affected by EEB to exit rates of similar individuals when EEB is not available. Also, while Fujita (2011) focuses on prime age men, we estimate effect for adult permanent (and temporary) job losers.

Nakajima (2012) calibrates a model of search behavior and finds much larger effects (1.4 percentage points on the unemployment rate) than we do. The discrepancy with our findings likely owes to the paper's choice of calibration. The paper uses the estimate of the effect of an extra week of UI benefits on the average duration of unemployment from Moffitt and Nicholson (1982) and Card and Levine (2000) to calibrate the cost of search, the key parameter determining the effect of EEB.

Finally, comparing unemployment in counties on borders between states having different maximum durations of EEB, Hagedorn, Karahan, Manovskii, and Mitman (2013), HKMM, find that EEB increased unemployment by 3.6 percentage points. They justify this outsized effect by showing that it is theoretically possible for higher wage demands from workers eligible for EEB to cause firms to drastically lower vacancies (or increase layoffs). This general equilibrium effect of EEB (operating through job creation rather than worker's search effort or reservation wages) is not captured by our estimates (which only reflect the influence of EEB on search and

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[22] Another reason to be wary of extrapolating the Card and Levine (2000) results is that New Jersey, where both the UI benefit replacement rate and the average benefit exhaustion rate were near well above average levels in the United States, may not be representative of the United States.

reservation wages), or other empirical work discussed above. As a result, their results could be consistent with ours if the effect of EEB on vacancies and job creation is extremely large.

6. Conclusion

We use differences in the behavior of unemployed individuals at different unemployment durations to identify and understand the magnitude of the effect of EEB on unemployment. Because EEB has significant effects on U-E hazard rates, we estimate that it had non-negligible effects on the unemployment rate in recent years. At the same time, because EEB has very little effect on the behavior of the relatively large share of individuals who are at the beginning of their unemployment spells, the overall effect of EEB on the unemployment rate is fairly modest; at its peak (in terms of the average number of benefit weeks provided) EEB boosted the unemployment rate by one-third percentage point. Because the changes imparted to U-E and U-N exit rates by EEB have offsetting implications for the participation rate, the effect of EEB on the participation rate is estimated to have been quite small.

Using a data set on EEB that goes back to the 1970s, we are able to compare the effects of EEB on unemployment in the most recent episode with effects in previous episodes. We find that for the most part, estimated effects of EEB on U-E and U-N hazards are quite similar over history to estimated effects in the most recent episode. Because the number of individuals likely receiving EEB was much smaller, as a share of unemployment, in these episodes, the effect of EEB on the unemployment rate is estimated to have been quite modest.

**Figure 1a. Permanent Job Loser U-E Exit Rates
(with and without EEB)**
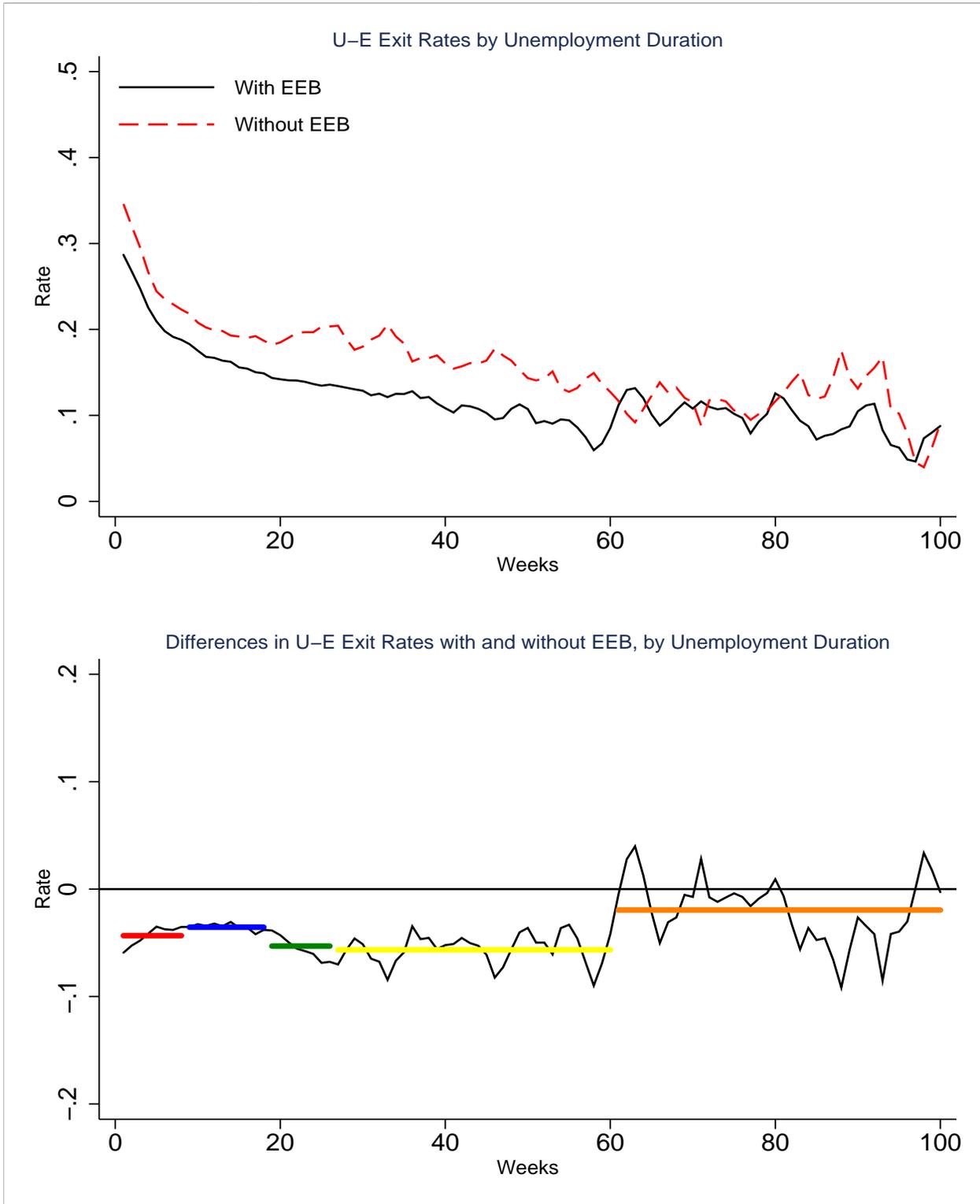
Note. Rates are smoothed using a five-week centered moving average with weights of .1, .2, .4, .2, .1. Data adjusted for potential break in exit rates following the 1994 CPS redesign.

**Figure 1b. Shift in Permanent Job Loser U-N Exit Rates
(Across Periods with and without EEB)**
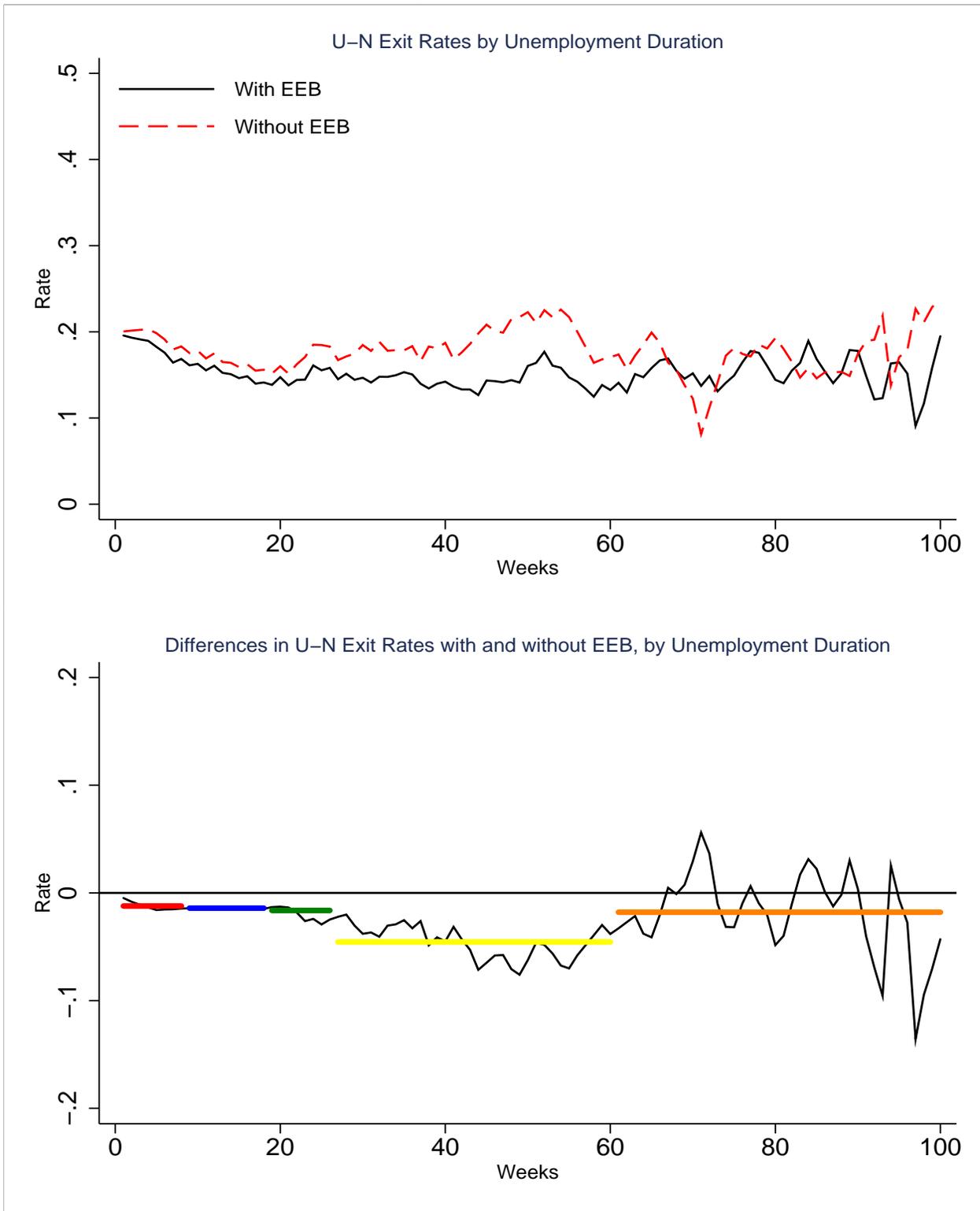
Note. Rates are smoothed using a five-week centered moving average with the following weights: .1, .2, .4, .2, .1.
Data adjusted for potential break in exit rates following the 1994 CPS redesign.

Figure 2a. Effect of Observable Characteristics on U-E Exit Rates (Permanent Job Losers)
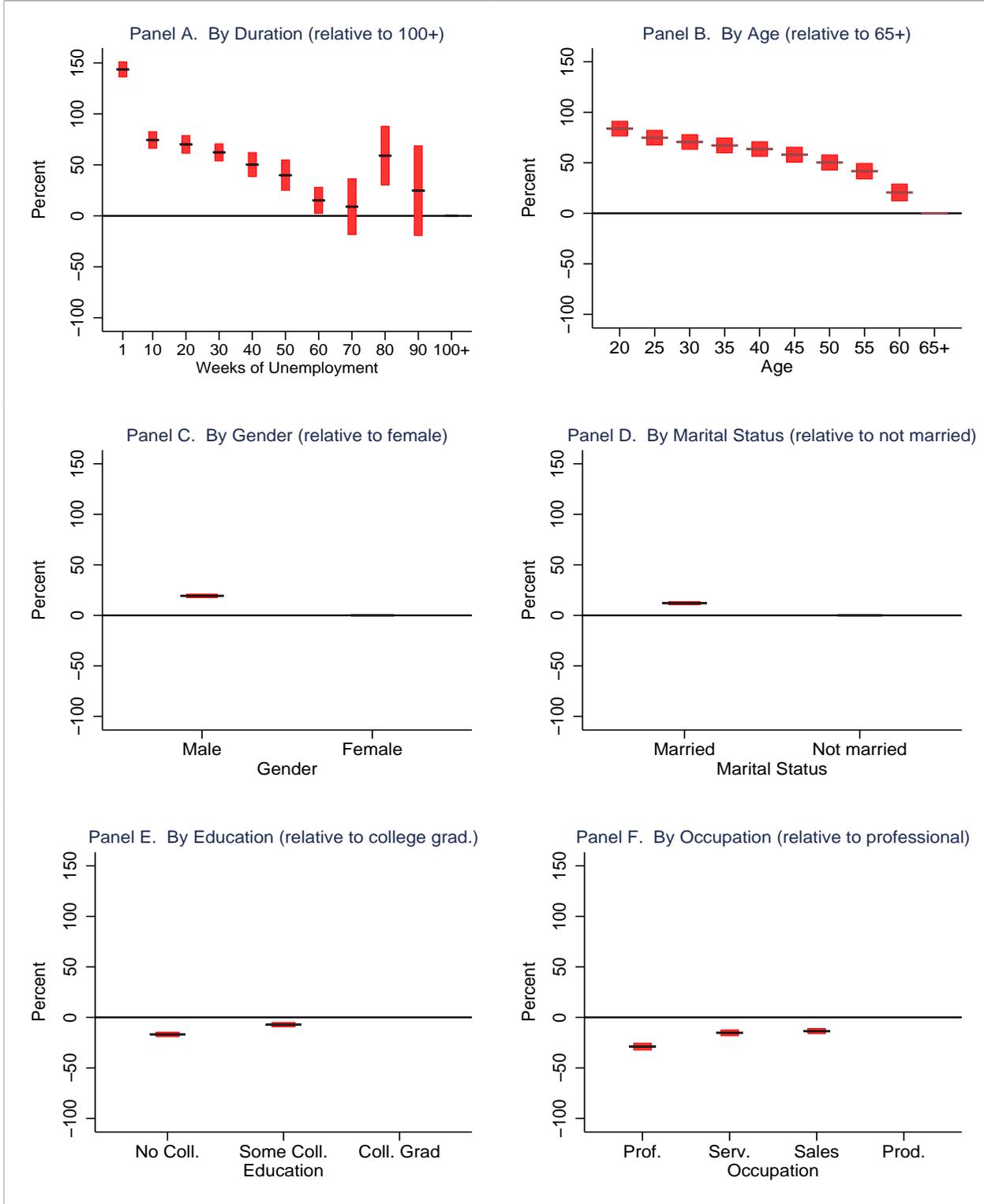
Note. Thick black lines represent point estimates. Red bars span point estimate plus or minus two standard deviations.

Figure 2b. Effect of Observable Characteristics on U-N Exit Rates (Permanent Job Losers)
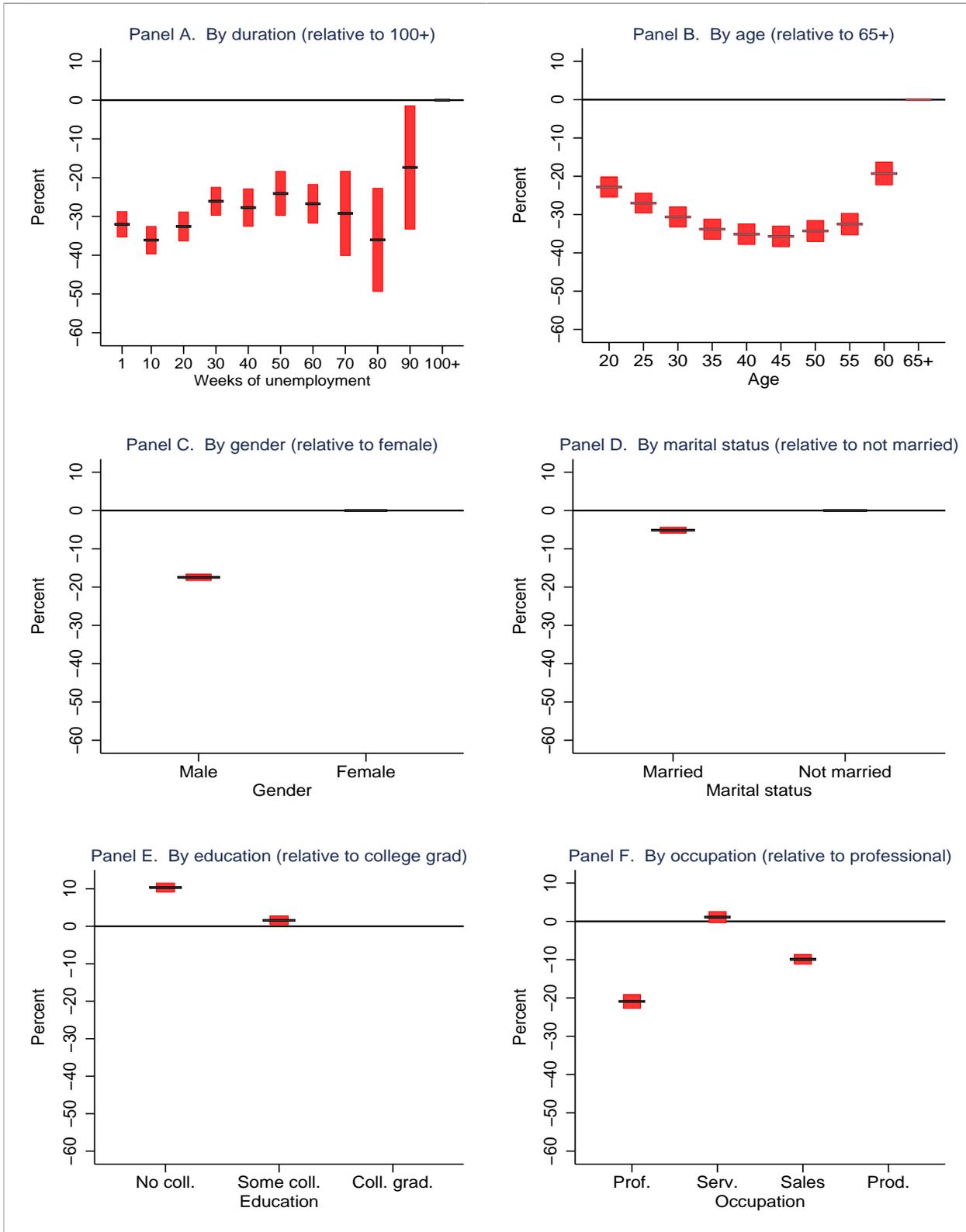
Note. Thick black lines represent point estimates. Red bars span point estimate plus or minus two standard deviations.

Figure 3. Effect of EEB on the Unemployment Rate over Time
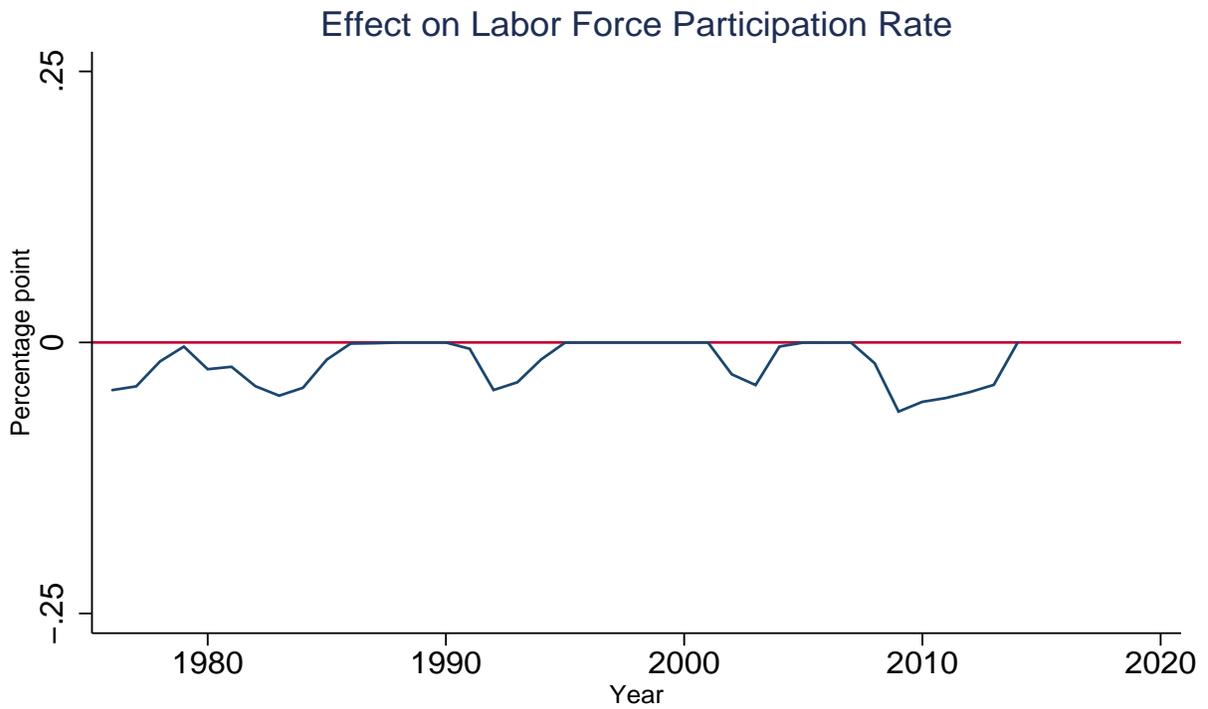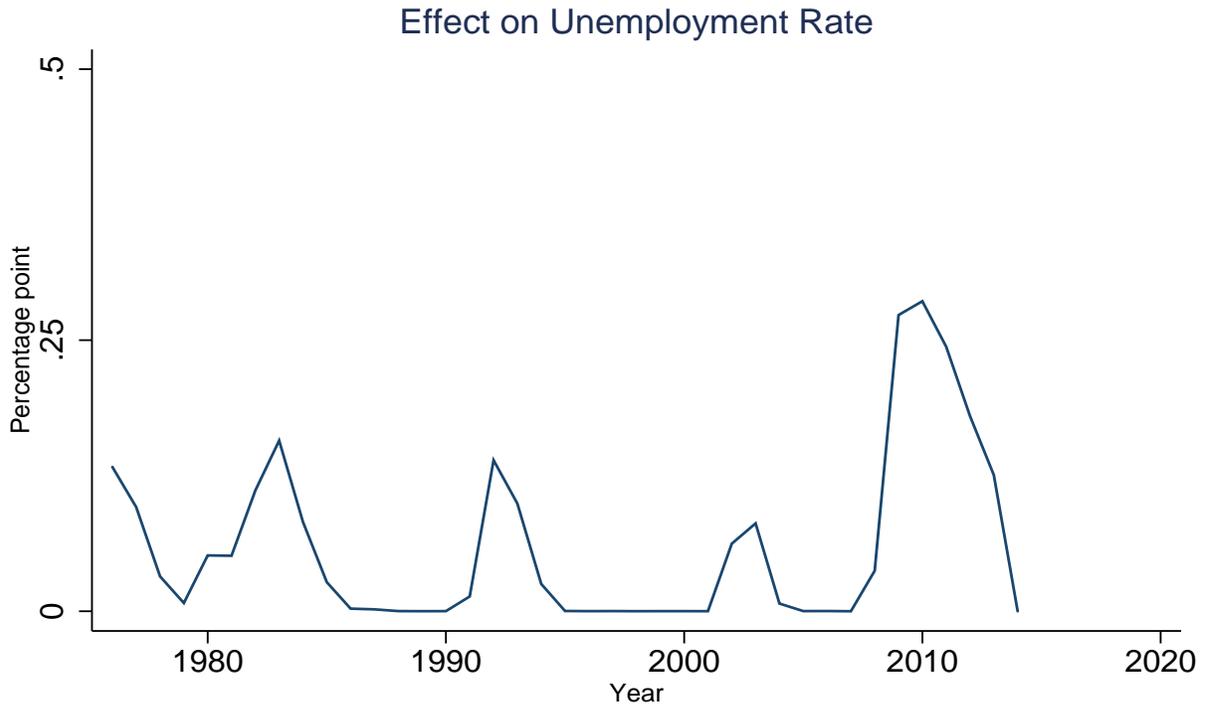
Figure 4. Hypothetical Effects of EEB on U-E Hazard
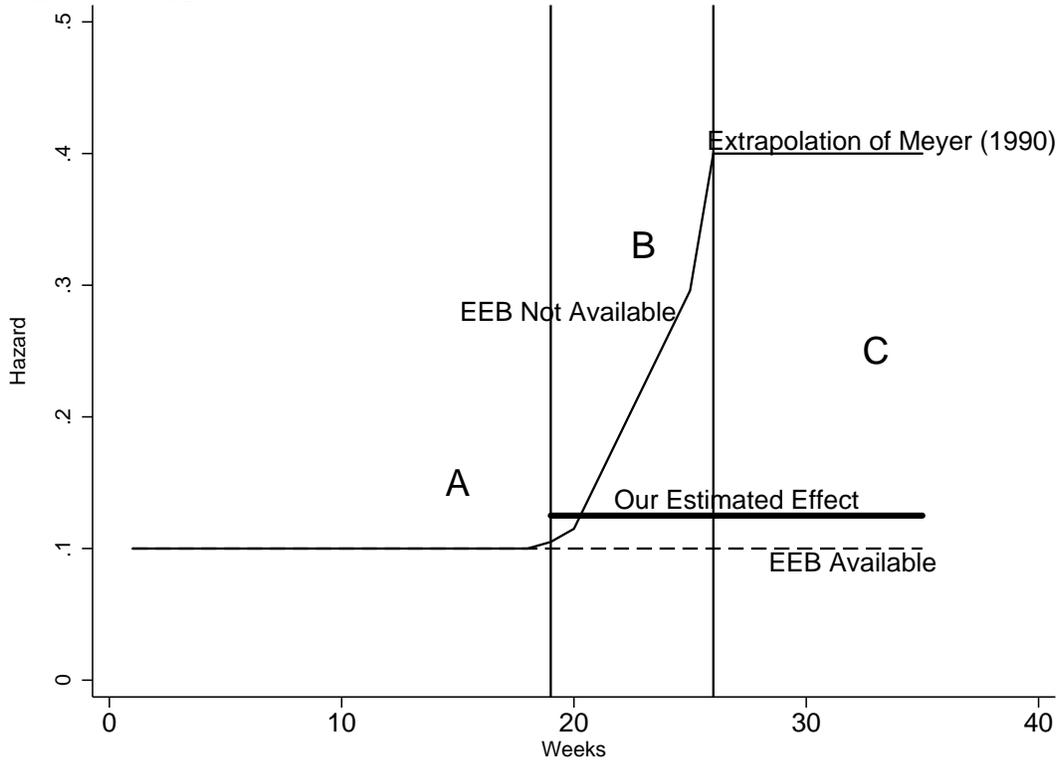
Note. The flat thin solid and dashed line at the bottom of the figure represents a hypothetical hazard rate for an individual when EEB benefits are available. Because the expiration of benefits for the individual is far into the future, the hazard rate does not change as weeks of benefit receipt increase. The thin solid line represents the hazard rate when EEB benefits are unavailable. Because benefits expire at 26 weeks, hazard rates rise through section B of the graph as weeks of benefit receipt near the termination of benefits. The amount of the increase of the hazard in section B is roughly calibrated using results from Meyer (1990). The thin solid line in section C of the graph shows the hazard rate after benefits are exhausted and EEB benefits are not available. This line extrapolates the hazard right before the termination of benefit recept, as roughly calibrated using results from Meyer (1990), to the period when the individual is no longer receiving benefits. The difference between the thick solid lines and the bottom thin solid and dashed line corresponds roughly to our estimate of the effect of EEB on the hazard rate. In section C of the graph, it is well below what an exptrapolation of results from Meyer would suggest.

Table 2. EEB Effect on Permanent Job Loser U-E and U-N Exit Rates (percent change, sample period: 1976-2013)

	U-E	U-N
Group 1	-3.55	4.32
	(1.24)	(1.12)
Group 2	-2.84	3.99
	(1.66)	(1.47)
Group 3	-17.43	-0.86
	(2.28)	(1.93)
Group 4	-16.23	-6.73
	(2.10)	(1.63)
Group 5	0.03	-7.93
	(2.19)	(1.62)

Note. Standard errors clustered by state and time are in parenthesis. Estimates are from a logit specification that also includes 100 unemployment duration dummy variables, a measure of aggregate labor market tightness, and the unemployment rates for 204 state/occupation labor market segments. Demographic controls include dummy variables for 10 age, 3 education, 13 industry, and 4 occupation classifications, as well as dummy variables for gender, marital status and state of residence. Estimated effects are computed assuming observable characteristics are equal to their mean values when EB is available. Standard errors for the percent change in exit rates are computed using the delta method.

Table 3. Estimates of EEB Effect on Permanent Job Loser U-N and U-E Exit Rates by Episode (percent change)

	1976-1985		1986-1995		1996-2005		2005-2013	
	U-N	U-E	U-N	U-E	U-N	U-E	U-N	U-E
Group 1	0.58	2.42	8.34	-9.51	3.03	-9.39	7.22	-3.93
	(1.53)	(1.96)	(2.51)	(3.30)	(3.47)	(2.96)	(3.94)	(4.05)
Group 2	1.43	-2.07	7.26	-8.45	8.81	-5.15	3.24	-1.67
	(2.22)	(2.75)	(3.16)	(4.09)	(4.01)	(4.04)	(4.63)	(4.86)
Group 3	-3.19	-14.00	3.49	-25.85	-3.85	-18.31	6.92	-16.15
	(2.98)	(3.82)	(3.84)	(5.21)	(5.39)	(5.45)	(5.65)	(6.48)
Group 4	-7.13	-2.06	-7.08	-22.25	-11.59	-13.01	-7.66	-17.84
	(2.90)	(3.98)	(3.82)	(5.46)	(5.81)	(5.75)	(4.36)	(5.19)
Group 5	-7.73	1.79	4.55	-4.19	-0.29	0.10	-4.43	-8.60
	(2.54)	(3.70)	(3.29)	(4.94)	(4.64)	(5.62)	(5.04)	(6.18)

Note. Standard errors clustered by state and time are in parenthesis. Estimates are from a logit specification that also includes a measure of aggregate labor market tightness, unemployment rates for 204 state/occupation labor market segments, dummy variables for 10 age, 3 education, 13 industry, 4 occupation, and 100 unemployment duration classifications, as well as dummy variables for gender, marital status and state of residence. Estimated effects are computed assuming observable characteristics are equal to their mean values when EB is available. Standard errors for the percent change in exit rates are computed using the delta method.

Table 4. Factors Influencing the Effect of EEB on Unemployment

Panel 1. Percent of unemployment				
	1982	1992	2003	2010
Groups				
Group 1. Duration<8 weeks, EEB available	18.03	24.80	24.84	16.54
Group 2. Duration<=18 weeks, duration>8 weeks, EEB available	8.84	11.26.0	11.49	10.13
Group 3. Duration<=Reg. UI max, duration>18 weeks, EEB available	5.06	6.29	5.75	5.95
Group 4. Duration>Reg. UI max , duration < EEB max	3.04	6.58	4.46	19.67
Group 5. Duration>EEB max, EEB available	5.11	6.64	7.37	7.37

Panel 2. Effect of EEB on hazard rates (percentage points)				
Hazard		Groups		
	1	2	3	4
U-N	0	0	-0.86	-6.73
U-E	-3.55	-2.84	-17.43	-16.23

Panel 3. Effect of hazard rates on unemployment and participation (percentage point)

	Unemployment rate	Participation rate
First Order Effect of 10 percent change in Hazard Rate on the Unemployment Rate		
U-N	-0.25	-0.14
U-E	-0.47	0.10

Panel 4. Effect of EEB on unemployment and participation

	1982	1992	2003	2010
Unemployment rate				
U-N	.01	.01	.01	.04
U-E	.14	.16	.11	.28
Total	.15	.17	.11	.32
Participation rate				
U-N	.00	.01	.00	.02
U-E	-.03	-.03	-.03	-.06
Total	-.02	-.03	-.02	-.04

Note. Group 1 includes permanent job losers and temporary job losers with durations greater than 26 weeks.
Group 2 includes only permanent job losers.

Appendix

Characteristics of Sampled Individuals

Table A1. Characteristics of Individuals at Different Unemployment Durations when EEB is and is not Available

	Group 1		Group 2		Group 3		Group 4		Group 5	
	No EEB	EEB	No EEB	EEB	No EEB	EEB	No EEB	EEB	No EEB	EEB
Age	34	34	37	37	38	38	40	41	43	42
Married	42	44	47	48	50	49	49	50	52	51
Male	61	62	62	63	62	63	66	61	75	68
Low Education	53	40	50	39	49	39	52	44	49	41
Medium Education	35	36	35	36	36	35	33	32	33	34
High Education	12	24	15	25	15	26	15	24	17	26
Professional	15	15	17	18	18	19	18	22	19	19
Service	17	17	14	15	13	14	14	14	14	13
Sales	22	21	23	22	24	22	21	24	17	21
Production	45	45	45	44	44	45	46	40	50	46
No. of Obs.	106,972	101,652	48,724	59,376	22,531	32,967	29,828	54,144	10,945	46,011

Note. The figures in the "Age" column are the average age of individuals in the relevant groups. The other figures in the table, except for the final row, are the percent of the relevant group with the relevant characteristic. For example, the number "61" in the third row of the first column is the percent of individuals in group 1 that are male when EEB is not available.

<u>Temporary Layoff Unemployment</u>

Estimates presented in section 3 of the paper are for permanent job losers, who make up the large majority of individuals most likely affected by EEB. However, it is also possible that temporary job losers, particularly those nearing the exhaustion of regular UI benefits, are also affected by EEB. To test this possibility, we use an equation similar to the one used for permanent job losers. Because there is a break in the definition of temporary layoff unemployment in 1994, we only use data after 1993. Figures A1 and A2 show estimates of coefficients other than those capturing the effect of EEB. Because there are so few observations on temporary layoff unemployment with durations above 35 weeks, we only show coefficient estimates up to this duration. In general, duration profiles have the same qualitative features as those for permanent job losers—upward sloping for U-N transitions and downward sloping for U-E transitions—but the downward and upward movements are more abrupt and occur earlier in the duration profile. Otherwise, coefficients on demographic characteristics are similar to those estimated for permanent job losers.

Table A2 shows estimates of the effect of EEB on the U-E and U-N transitions of temporary job losers. For the most part, these are quite similar to those for permanent job losers. For U-E transitions, the largest negative effects are for groups 3 and 4, and the estimated effect on these groups is similar to those estimated for permanent job losers. For U-N transitions, the largest negative response is for group 4, and the estimate for this group is similar in magnitude to that for permanent job losers (though it is not statistically significant). The group with significantly different estimated behavior is group 5. But there are very few observations to estimate the effect for this group, and it is likely, given the very long duration of unemployment in this group that these individuals are more similar to permanent job losers than temporary job losers. Overall, these results suggest that we can treat temporary job losers similarly to permanent job losers when considering the effects of EEB.

**Table A2. EEB Effect on Temporary Layoff U-N and U-E Exit Rates
(percent change, sample period: 1994-2013)**

	U-E	U-N
Group 1	-3.93	3.81
	(2.44)	(4.00)
Group 2	-2.59	0.19
	(3.47)	(6.08)
Group 3	-18.21	-3.93
	(5.67)	(8.48)
Group 4	-12.46	-7.09
	(6.25)	(9.26)
Group 5	-11.78	28.20
	(10.34)	(12.87)

Note. Standard errors clustered by state and time are in parenthesis. Estimates are from a logit specification that also includes 100 unemployment duration dummy variables, a measure of aggregate labor market tightness, and the unemployment rates for 204 state/occupation labor market segments. Demographic controls include dummy variables for 10 age, 3 education, 13 industry, and 4 occupation classifications, as well as dummy variables for gender, marital status and state of residence. Estimated effects are computed assuming observable characteristics are equal to their mean values when EB is available. Standard errors for the percent change in exit rates are computed using the delta method.

Figure A1. Effect of Observable Characteristics on U-E Exit Rates (Temporary Job Losers)
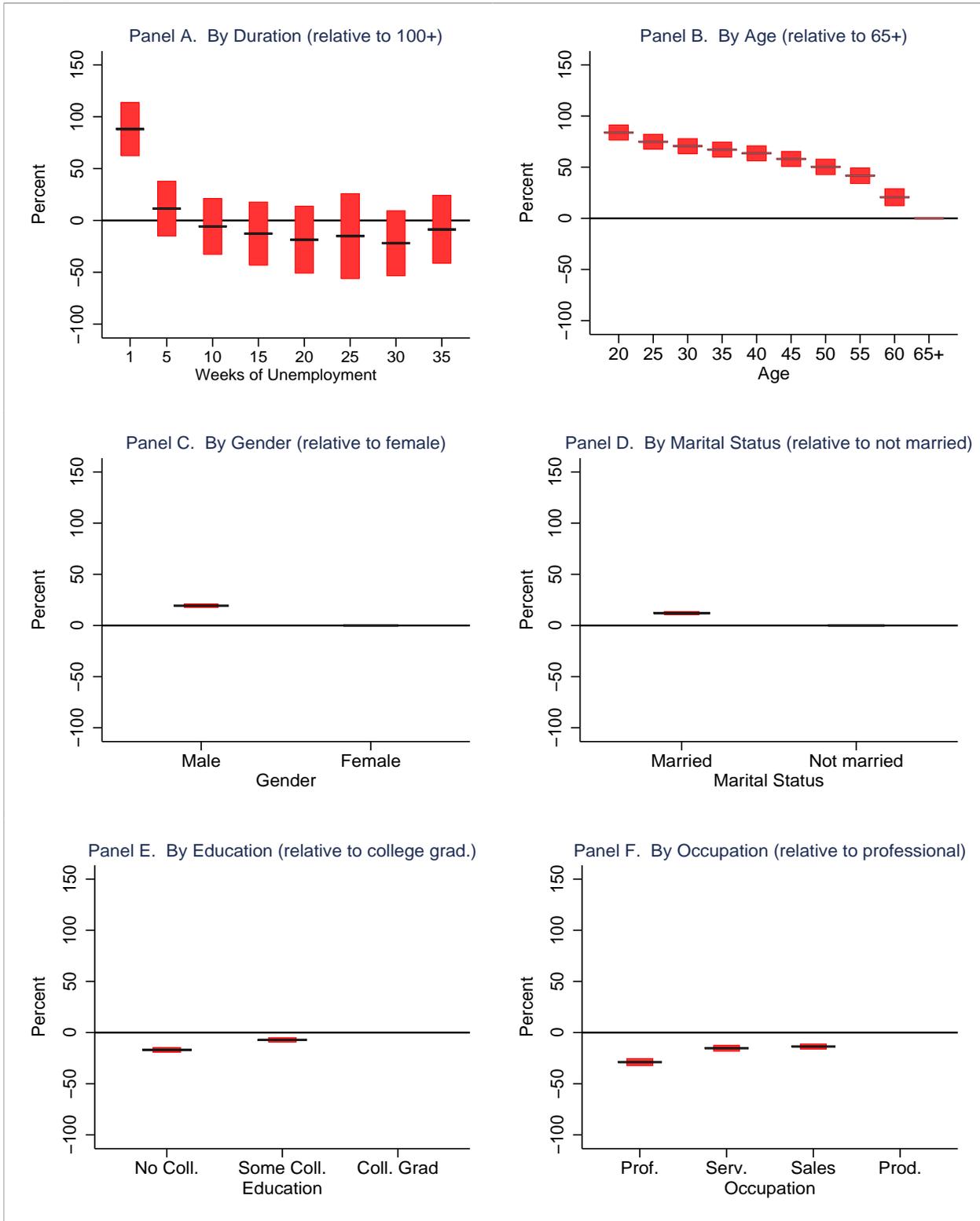
Note. Thick black lines represent point estimates. Red bars span point estimate plus or minus two standard deviations.

Figure A2. Effect of Observable Characteristics on U-N Exit Rates (Temporary Job Losers)
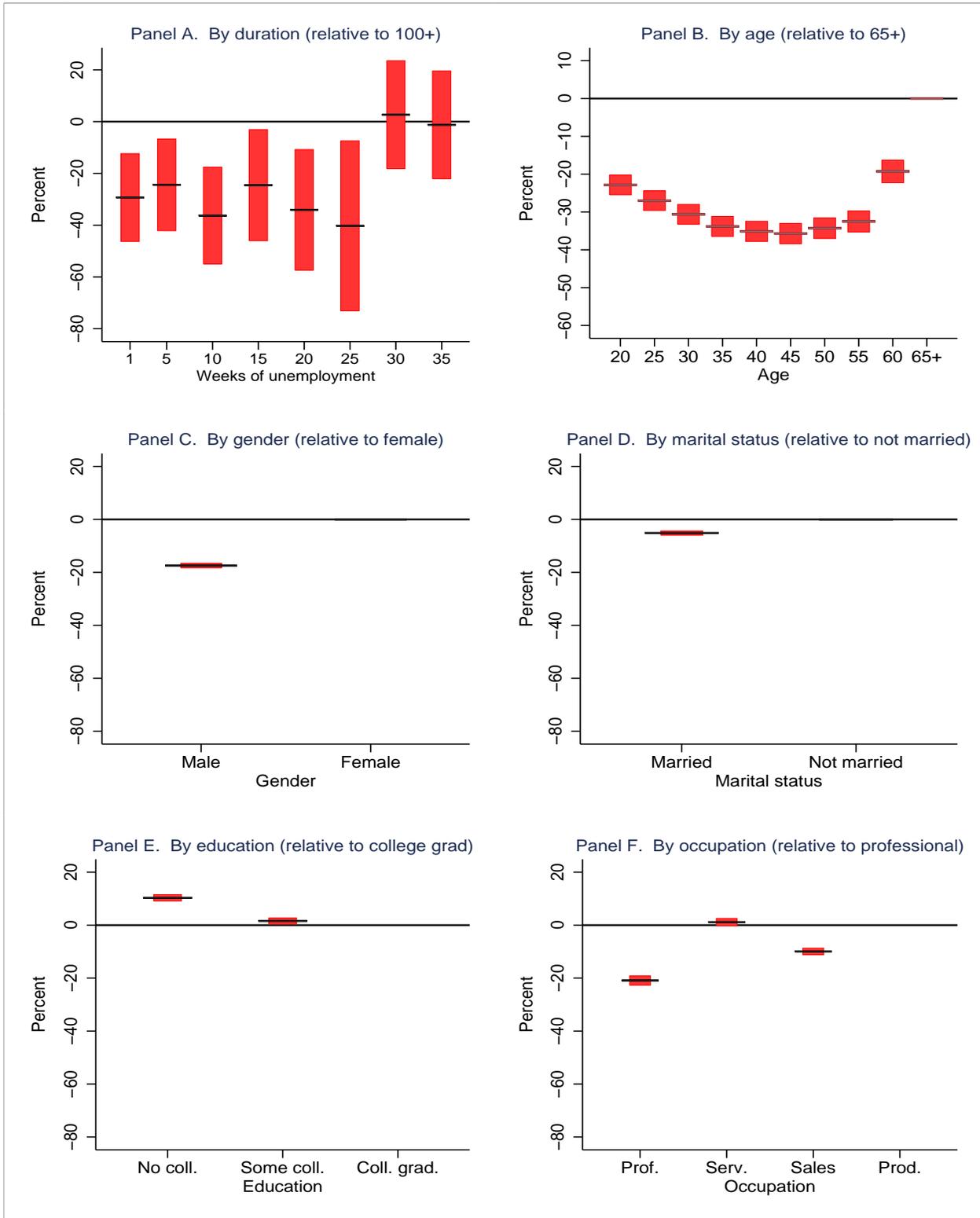
Note. Thick black lines represent point estimates. Red bars span point estimate plus or minus two standard deviations.

Rules Governing Eligibility for Extended and Emergency Benefits (EEB)

After individuals exhaust their regular unemployment insurance benefits, which have a maximum duration of 26 weeks in almost all states, they can receive benefits from two other programs: extended unemployment insurance and federal supplemental or emergency benefits.

Extended Benefits

The Extended Unemployment Compensation program was available in all states beginning in 1972. It provided for a maximum of 13 extra weeks, or ½ the individual's entitlement to regular benefits, whichever was less, after an individual exhausts her regular benefits when both the national insured unemployment rate and the state insured unemployment rate satisfied certain conditions. Specifically, both the national and state-level insured unemployment rates (IUR) were required to be above a certain level; in addition, the state IUR was required to be a certain percentage greater than its level one and two years earlier, henceforth termed the comparison requirement. Amendments to this act have mostly focused on adjusting these triggers by, for example, raising and lowering the national and state IUR levels, allowing triggers tied to the state total unemployment rate (TUR), or suspending the comparison requirement. Below are listed some of the more notable changes.

- From October 1972 until the end of 1976, the comparison requirement was temporarily waived.
- Effective January 1, 1975 and ending on December 31, 1975, the national IUR trigger was lowered from 4.5 percent to 4 percent.
- In the first quarter of 1977, the triggers in the extended benefit (EB) program were modified to trigger on when either of the following two conditions were met: (1) the seasonally adjusted national IUR reached at least 4.5 percent over the most recent 13 week period, or (2) the state IUR reached at least 4.0 percent over the most recent 13 week period and was 20 percent higher than the state's average IUR in the corresponding 13 week period in each of the two preceding two years. States were allowed to waive the comparison requirement if the state's IUR reached 5 percent or more.
- Effective August 13, 1981, the national trigger for extended benefits was eliminated. In addition, individuals on extended benefits could not be counted as part of a state's IUR that was used to trigger extended benefits. After Sept. 25, 1982, the state trigger was raised to at least a 5 percent IUR that had to be 20 percent greater than the IUR in the corresponding period of each of the two preceding years. The comparison requirement was waived for an IUR of 6 percent of greater.
- On July 3, 1992, a new optional extended benefit trigger was created, under which EB would trigger on if the seasonally adjusted TUR was at least 6.5 percent for the most recent three month period and 110 percent of the rate in the comparable period in either or both of the two preceding years. Under this option, an additional 7 weeks of EB

become available if the state TUR reached at least 8 percent and was 110 percent of the comparable period in either or both of the two preceding years.

Emergency or Supplemental Benefits

In every recession since 1974, Congress has passed legislation allowing for additional weeks of federally funded unemployment insurance benefits to be paid after individuals have exhausted regular benefits or both regular and extended benefits. Usually, the duration of these benefits has been tied to conditions in a state, such as its IUR or TUR. Also, the weeks of benefits allowable, their relationship to state economic conditions, and the duration of the supplemental or emergency program have been amended frequently. Typically, the programs have lasted two to three years. Below we list the five program since the early 1970s and their most important characteristics.

1972-1973

- The Emergency Unemployment Compensation Act of 1971 provided 13 weeks of federally funded unemployment insurance when a state's IUR plus an adjustment for UI exhaustees reached 6.5 percent or more and the state had triggered onto EB or was not on EB solely because of the comparison requirement. In addition, the state must have triggered onto EEB at some point over the previous 52 weeks. The program lasted from January 30, 1972 until December 31, 1972 for initial claims and March 31, 1973 for continuing claims.

1975-1977
- The Emergency Unemployment Compensation Act of 1974 created Federal Supplementary Benefits (FSB) payable to individuals who had exhausted regular and extended benefits.
    - The program was extended multiple times and lasted from Jan. 1, 1975 until Nov. 1, 1977 for new claims and Feb. 1, 1978 for existing claims.
    - Originally, the maximum benefit duration was 13 weeks. The program trigger was the same as the EB trigger.
    - On March 29, 1975, the benefit period was extended to the lessor of 26 weeks or 100 percent of regular benefit duration for states with an IUR of 6 percent over the preceding 13 weeks. Individuals in states with an IUR of less than 6 percent could receive 13 weeks of federal benefits. In addition, the law was amended on June 30, 1975, so that FSB became only available in states in which extended benefits had triggered on and the IUR over the preceding 13 week period was 5 percent.
    - On April 12, 1977, effective April 30, 1977, the maximum duration of benefits was reduced back to 13 weeks.

<u>1982-1985</u>
- Beginning Sept. 12, 1982, Federal Supplemental Compensation (FSC) became available for individuals exhausting regular and extended benefits.
    o The program was extended multiple times and lasted from Sept. 12, 1982 until March 31, 1985 (after March 31, 1985, individuals could continue to draw FSC for consecutive weeks up until their FSC limit was reached as long as they were unemployed).
    o Initially, benefit durations were
        ▪ 10 weeks for individuals in states that were in EB status on or after June 1, 1982.
        ▪ 8 weeks in states with an IUR over the recent 13 week period of at least 3.5 percent.
        ▪ 6 weeks in all other states.
    o On Jan. 6, 1983, weeks of benefits were increased to:
        ▪ 16 weeks in states with an IUR of at least 6 percent.
        ▪ 14 weeks in states with an IUR less than 6 percent but in which EB had triggered on between June 1, 1982, and January 6, 1983.
        ▪ 12 weeks in states where the IUR was less than 6 percent and at least 4.5 percent, or if the IUR was greater than 6 percent but the state had triggered onto EEB after Jan. 6, 1983.
        ▪ 10 weeks in states with an IUR of at least 3.5 percent but less than 4.5 percent and that had not triggered onto EB during the life of FSC.
        ▪ 8 weeks elsewhere.
        ▪ States triggering into a higher FSC status remained there for at least 4 weeks.
    o After April 1, 1983, new recipients of FSC received benefits for:
        ▪ 14 weeks in states with an IUR of 6.0 percent or more.
        ▪ 12 weeks with an IUR between 5.0 and 5.9 percent.
        ▪ 10 weeks in states with an IUR between 4.0 and 4.9 percent.
        ▪ 8 weeks in all other states.
    o On October 24, 1983, benefit durations were changed to
        ▪ 14 weeks in states with an IUR of 6 percent or more or a long-term IUR of 5.5 percent or more.
        ▪ 12 weeks in states with an IUR of 5 to 5.9 percent or long-term IUR of 4.5 to 5.4 percent.
        ▪ 10 weeks for states with an IUR between 4 and 4.9 percent
        ▪ 8 weeks in states with an IUR of less than 4 percent.
        ▪ A long-term IUR was defined as the average from Jan 1, 1982 to the last week of the second calendar quarter ending before an unemployment spell begins.

- The Emergency Unemployment Compensation Act of 1991 was extended multiple times and lasted from Nov., 17, 1991 until Feb. 5, 1994 for new claims. No payments were allowed after April 30, 1994.
    - Initially, beneficiaries received benefits for:
        - 20 weeks if the average adjusted IUR, AIUR, was 5 percent or greater, or the 6 month total unemployment rate (TUR) was 9 percent or greater.
        - 13 weeks if the state AIUR was at least 4 percent, or if the state AIUR was at least 2.5 percent and the state exhaustion rate for the most recent month was at least 29 percent.
        - 6 weeks in all other states.
        - The AIUR was defined as the insured unemployment rate with exhaustees in the last 3 months included in the numerator.
    - The program also allowed for "reach-back", which allowed workers exhausting regular benefits after February 28, 1991 but before the passage of this act to be eligible for benefits if the AIUR was at least 3 percent.
    - On Dec. 4, 1991, the act was amended to leave only two tiers defined for high unemployment states and all others:
        - 20 weeks in states with an AIUR of 5 percent or more, or a six-month TUR of 9 percent or more.
        - 13 weeks in all other states.
        - Reach-back was made available in all states.
    - On February 7, 1992, the limits were changed to
        - From Feb., 7, 1992 until June 13, 1992
            - 33 weeks in high unemployment states (as defined previously).
            - 26 weeks in all other states.
        - After June 13, 1992
            - 20 weeks in high unemployment states
            - 13 weeks in all other states.
    - On July 3, 1992, benefit lengths were changed to depend on the national TUR:
        - Until the national TUR averaged 7 percent for two consecutive months, benefits were
            - 26 weeks in states where the AIUR equaled or exceeded 5 percent, or where the TUR exceeded 9 percent.
            - 20 weeks in all other states
        - After the average TUR falls below 7 percent and before it falls below 6.8 percent.
            - 15 weeks in high unemployment states.
            - 10 weeks in all other states.

- After the average TUR falls below 6.8 percent
  - 14 weeks in high unemployment states.
  - 7 weeks in all other states.
    - On Nov. 24, 1993, the duration of benefits was changed to
      - 13 weeks in high unemployment states (defined as above).
      - 7 weeks in all other states.

2002-2003

- On March 9, 2002, the Job Creation and Worker Assistance Act created temporary emergency unemployment compensation (TEUC).
  - The program lasted from March 9, 2002 until December 28, 2003 for new claims. Payments could extend through March 31, 2004.
  - The program provided:
    - The lessor of 13 weeks or ½ the duration of regular UI benefits.
    - 26 weeks if when an individual exhausts the first TEUC entitlement, the states is in an EB period or would have been if the IUR requirement had been 4 percent instead of 5 percent.

2008-2013

- On June 30, 2008, the Emergency Unemployment Compensation program was created.
  - The program was extended multiple times and lasted from June 30, 2008 until the Dec. 31, 2013.
  - Initially, the program provided 13 additional weeks of benefits.
  - On Nov. 21, 2008, benefit periods were increased to
    - 20 weeks for all states
    - 33 weeks in high unemployment states, defined as states where
      - EB has triggered on
      - The state IUR over the preceding 13 weeks was 4 percent.
      - A state's 3 month seasonally adjusted TUR was at least 6 percent
  - On Nov., 6, 2009, benefit durations were changed to
    - 34 weeks for all states
    - 47 weeks in states where
      - The state IUR was at least 4 percent
      - The 3 month seasonally adjusted TUR was at least 6 percent
    - 53 weeks if
      - A state was on EB and the state thirteen-week IUR was at least 6 percent, or

- The three-month seasonally adjusted TUR was at least 8.5 percent.
  - For states not on extended benefits,
    - From March 1, 2012 to May 31, 2012, Tier 4 benefits were changed to the lesser of 62 percent of regular UI or 16 weeks (effectively increases the max federal benefits available to 63 weeks)
    - From June 1, 2012 to August 31, 2012, Tier 4 benefits were reduced to the lesser of 24 percent of regular UI benefits or 6 weeks (the maximum of federal benefits was reduced back to 53 weeks.
  - From September 2, 2012 to December 31, 2013 EUC benefit durations were
    - 14 weeks in all states
    - 28 weeks in states with seasonally adjusted TUR of 6 percent.
    - 37 weeks where seasonally adjusted TUR at least 7 percent.
    - 43 weeks where seasonally adjusted TUR at least 9 percent.

The Employment and Training Administration publishes trigger notices listing the states where extended and federal benefits are available. These notices are available on line back to 1987. Prior to 1987, we use hard copies of trigger notices to identify the level of EEB benefits available in a state in each week.

Other issues

Lapses and extensions

At times, federal benefit programs expire for a short period before legislation is passed to extend them. Usually, the unemployed are granted the benefits they would have been entitled to had the program been extended without a lapse. In our estimation, we use the average maximum duration of allowable benefits over a month. If the benefit lapse is long enough, it can have a noticeable effect on the average maximum duration of benefits in a month. We assume that behavior during these lapses is similar to other periods, that is, it is governed only by the individual's unemployment duration relative to the average maximum allowable duration of benefits in the individual's state at that time.

Payments at the end of a program

Typically, at the end of EUC programs, individuals are no longer allowed to apply for federal benefits, but if they are already receiving benefits, they are allowed to continue receiving benefits (typically for another three months) if they are still unemployed and if they have not reached the maximum weeks of federal benefits. Because of the uncertainty about whether individuals during these three month periods are receiving EEB payments, we exclude these three-month periods from our sample.

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