The Effect of Extended Unemployment Insurance Benefits: Evidence from the 2012-2013 Phase-Out

By Henry S. Farber, Jesse Rothstein, and Robert G. Valletta[*]

The duration of U.S. Unemployment Insurance (UI) benefits was expanded to an unprecedented degree in the Great Recession, reaching a maximum of 99 weeks in many states by 2010. These expansions were then rolled back in 2012 and 2013. Since January 2014, no state has had UI benefits available beyond the normal duration (26 weeks in most states).

Unemployment insurance extensions may raise measured unemployment, both by reducing the incentive for recipients to find jobs quickly and by bolstering the incentive to engage in and report active job search. But the relative magnitude of these effects is uncertain and may vary with economic conditions. In earlier work, we examined the effects of the 2008-2011 expansions, relying on cross-state and temporal variation in the duration of available UI benefits (Rothstein 2011; Farber and Valletta 2015). We found that benefit extensions slightly reduced the exit rate from unemployment, largely through increased labor force attachment rather than reduced job finding.[1] These estimates may be affected by the historically weak labor market conditions around the Great Recession, however, and they may not generalize to changes in UI durations under more favorable labor market conditions.

In this study, we update our earlier analyses to incorporate the phase-out of benefit extensions in 2012 and 2013. Figure 1 contains plots of median weeks of available UI, by quarter since 2007, along with a measure of labor market slack, the ratio of unemployment to job openings.[2] This figure shows the run-up of UI durations in the Great Recession, from the basic 26 weeks of UI early in 2008 to the maximum of 99 weeks late 2009, followed by a decline beginning in early 2012. The benefit extensions came during a period of sharply increasing slack, but by the time of the roll-backs the labor market was substantially tighter. It is plausible that UI extensions could have larger effects on job finding in a tighter labor market (Kroft and Notowidigdo 2011).

We estimate models for the likelihood that an unemployed individual exits employment, using two different empirical specifications designed to isolate variation in UI durations produced by policy changes (as distinct from changes in labor demand conditions). We estimate separate effects for 2008-2011, when UI durations were expanding or stable, and for 2012-June 2014, when they were contracting. We use a competing risks framework to model separate effects of UI extensions on exit to employment (job finding) and on exit from the labor force. Large effects on job finding would suggest important economic efficiency costs of UI extensions. By contrast, effects on labor force attachment have little or no implication for economic efficiency (Card, Chetty, and Weber 2007).

I. The Expansion and Reduction of UI Benefit Durations, 2008-14

The rapid decline in UI durations in Figure 1 reflects the largely automatic roll-

* Farber: Princeton University, Industrial Relations Section, Firestone Library, Princeton, NJ 08544, farber@princeton.edu. Rothstein: University of California, Berkeley, 2607 Hearst Avenue #7320, Berkeley, CA 94720, rothstein@berkeley.edu. Valletta: Federal Reserve Bank of San Francisco, 101 Market St., San Francisco, CA 94105, rob.valletta@sf.frb.org. The views expressed in this paper are solely those of the authors and should not be attributed to the Federal Reserve Bank of San Francisco or the Federal Reserve System.

[1] Our past and current analyses focus on worker search behavior. Other recent work examines the labor market effects of potential employer responses to UI extensions (Hagedorn et al. 2013).

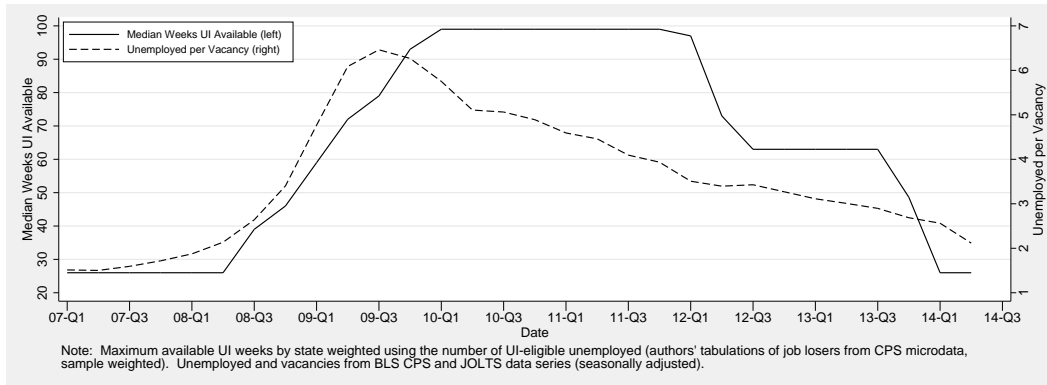[2] Weeks of available UI are measured monthly by state. We average this across months within quarters.

FIGURE 1. WEEKS OF UI AVAILABLE AND UNEMPLOYED/VACANCIES, QUARTERLY, 2007-2014Q2

back of two programs that provided for extra benefits during the recession. These are the temporary Emergency Unemployment Compensation (EUC) program and the permanently authorized Extended Benefits (EB) program.

EUC was a federally-funded program, introduced and expanded in several steps beginning in July 2008. It provided for as many as 53 weeks of benefits through four separate "tiers," some of which were automatically activated for all states and some of which were conditioned on high state unemployment rates. The program expired at the end of 2013, producing the second step down in Figure 1, from 63 to 26 weeks.

EB provides for 13 or 20 additional weeks of benefits in states with high unemployment rates, following exhaustion of regular and emergency (EUC) benefits.[3] Most states became eligible for EB in early 2009. A provision in the EB rules restricts benefits to states with unemployment rates higher than they were three years prior. This became binding in early 2012, and by August 2012 no state was paying EB.[4] This produced the first step down in Figure 1, as median weeks of benefits fell from 99 weeks in 2011q4 to 63 weeks in 2012q3.

EUC and EB complemented regular state benefits. These are typically 26 weeks, so

the maximum duration of UI benefits available during the period we study was 99 weeks (26 weeks regular benefits, 53 weeks EUC, 20 weeks EB).[5] A few states cut regular benefit durations to less than 26 weeks in 2011 or later.

Our initial analysis exploits changes in benefit durations coming from changes in EUC rules, from the phase-out of EUC and EB, and from decisions by some states to cut regular benefits. We also present estimates below that focus on variation arising directly from the phase-out of EB in 2012 and the termination of EUC benefits at the end of 2013.

II. Sample Definition and Data Issues

We use Current Population Survey (CPS) microdata from January 2008 to August 2014 for individuals ages 18-69. We restrict our analyses to respondents who are unemployed and report job loss as the reason, and hence are potentially eligible for UI.[6] Given our focus on the effect of extended benefits, and because most new spells of unemployment end before extended benefits could be an important factor, we restrict attention to individuals who have been unemployed for at least 3 full months.

[3]State participation is optional. However, the federal government bore 100 percent of the cost of EB from 2009 through mid 2013, so state participation was high during the period of interest here.

[4]New York and Alaska briefly resumed EB payments later in 2012. No state paid EB benefits after April 2013.

[5]Some states had 63 weeks of EUC in early 2012, but this was conditional on not offering EB benefits, so the maximum duration never exceeded 99 weeks.

[6]Due to potential reporting errors and incomplete take-up rates for UI benefits, this is an imperfect means of identifying UI eligibility. See Rothstein (2011) and Farber and Valletta (2015) for further discussion.

We use the reported duration of unemployment to model the effects of extended benefits. We assume that job losers are eligible for the full duration of benefits and that each draws benefits continuously from the date of job loss until benefit expiration or exit from unemployment.

We use the panel structure of the CPS to identify exits from unemployment. Each sampled address (housing unit) is interviewed for 4 consecutive months, left alone for 8 months, then reinterviewed for another 4 months. This sample structure allows us to match unemployed individuals within households to month-ahead labor market outcomes for three consecutive months during each 4-month rotation.[7]

One key concern with regard to use of the matched data is the likelihood of spurious transitions due to mismeasurement of labor force status (see for example Abowd and Zellner 1985). To address this concern, we follow the recoding approach developed in our earlier work. For unemployed individuals in month one who report a transition out of unemployment in month two (to employment or non-participation) and a return to unemployment in month three, we recode the month two status as unemployed. We retain and use the resulting observations created by the recode (with associated variables such as reason for unemployment and duration imputed based on their month one values). Imposing this adjustment to observed transitions requires three consecutive matched months of data and hence restriction to respondents in the first two of each set of four consecutive CPS interviews. As such, although we use CPS data through August 2014, our last measured exit hazards are for June 2014.

Our matched CPS sample covers January 2008 through June 2014 and contains 56,491 monthly observations on 37,059 spells of unemployment for eligible workers

out of work three months or more and aged 18-69. To compare our estimates across periods of extended benefit expansion and contraction, we split this sample in two: 2008-2011 and 2012-2014m6. In the earlier period, we use 24,735 spells of unemployment, of which 15.4 percent are observed to end in exit to employment and 13.4 percent are observed to end in exit from the labor force. In the later period, we use 12,324 spells, of which 16.4 percent are observed to end in exit to employment and 16.4 percent are observed to end in exit from the labor force.

III. The Effects of Extended Benefits on Exits from Unemployment

Figure 2 contains plots of seasonally adjusted monthly exit rates from unemployment (averaged by quarter) for the spells in our analysis sample. We show series for all exits, for exits to employment, and for exits out of the labor force. Total exits and exits to employment fell sharply in 2008 as the recession deepened. Both reemployment and labor force exit rates have risen gradually since the recession ended in mid-2009. There is no visible change in the rates of increase as extended benefits were phased out in 2012 and 2013.

Our primary estimates of the effects of UI durations come from a simple logistic discrete-choice model of exit from unemployment. The model is specified by assuming a spell ends in a given month t if an unobserved latent variable for spell i in state s and month t (y_{ist}^*) is positive. This latent variable is

$$
\begin{aligned}
y_{ist}^* = \ & X_{ist}\beta + Z_{st}\lambda + UI_{ist}\delta \\
& + \omega_s + \psi_t + \epsilon_{ist},
\end{aligned}
$$

where X_{ist} is a vector of individual variables, Z_{st} is a vector of state economic variables, ω_s and ψ_t are vectors of state and date (month-year) effects respectively, β and λ are vectors of coefficients, and ϵ_{ist} is an error term with a logistic distribution. UI_{ist} (with coefficient δ) is an indicator that equals one if individual i, identified as unemployed in state s and month

---

[7]Failures to match occur primarily when a household moves to a new housing unit between interviews (less than five percent of cases). To ensure valid matches of individuals across months, we dropped a small number of observations for which reported age, gender, race, or educational attainment is not consistent across months (e.g., age changes by more than 1 year).
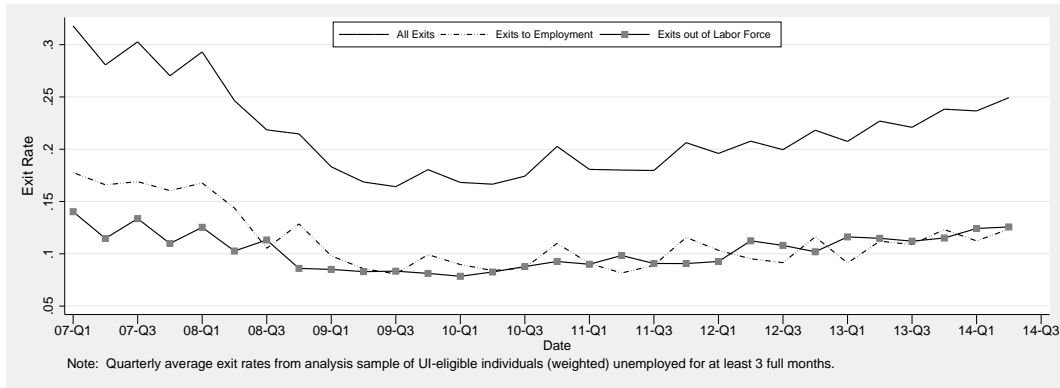
FIGURE 2. UNEMPLOYMENT EXIT RATES, QUARTERLY, 2007-2014Q2 (SEASONALLY ADJUSTED)

$t-1$, will have been unemployed in month t (the potential transition month) for fewer weeks than the number of weeks of UI benefits available.[8] We estimate the parameters specified in this relationship using a logit model of the observed monthly spell outcomes (exit or continuation). To examine the separate effects of UI_{ist} on exit to employment and exit out of the labor force, we use a competing risks version of this model. We assume that the two types of exit are independent events and treat each realized event as censoring the time until the other type of exit.

The estimated model includes in the X vector a set of standard personal characteristics that are systematically related to labor market outcomes: 4 education categories, 6 age categories (decade indicators covering the included ages 18-69), indicators for female, married, female*married, race/ethnicity, and indicators for 13 broad industry categories. In order to account for state labor market conditions over time (Z_{st}), the model includes a cubic in the monthly seasonally adjusted state unemployment rate and a cubic in the 3-month annualized growth in seasonally adjusted

log non-farm payroll employment. To allow for a flexible baseline hazard and to account for the effects of normal UI benefits, the model also includes a set of indicators for months 4, 5, and 6 of unemployment and single indicators for months 7-9, months 10-12, and months 13 and beyond (6 categories in total).[9] We also include a complete set of state and date (year-month) indicators (ω_s and ψ_t) in the model.

The first two columns of Table 1 contain estimated marginal effects of UI availability on the probability of exit from unemployment. We present three estimates for each of the two time periods (2008-2011 and 2012-2014m6) we study: One for all exits from unemployment, one for exit to employment, and one for labor force exit. The estimates for the single risk model (row 1) imply that the availability of UI benefits to an unemployed worker has a significant negative effect on the probability of exit from unemployment. The estimated effect is about 3.5 percentage points in 2008-2011 and 2.7 percentage points in 2012-2014. With an average exit hazard around 20% (figure 2), these estimates imply that the availability of extended benefits reduced the monthly exit rate from unemployment by about 15 percent.

---

[8]This specification follows Farber and Valletta (2015). Their models include an additional indicator for being in the final month of UI benefits. We have explored specifications that include this indicator as well. It was never significant, however, and its inclusion did not affect the estimated coefficient δ, so for parsimony we exclude it here. Rothstein (2011) takes a somewhat different approach to modeling extended UI effects but obtains results consistent with those presented here.

[9]Given the detailed controls for unemployment duration included in the model, our estimated effects of UI availability primarily reflect variation in extended benefits (EUC and EB). However, the estimated effects also reflect in part reductions of regular UI durations below 26 weeks in 8 states since 2011.

TABLE 1—ESTIMATED AVERAGE MARGINAL EFFECTS ON PROBABILITY OF EXIT FROM UNEMPLOYMENT

Model	Specification 1		Specification 2
	2008-2011	2012-2014m6	2012-2014m6
(1) Single Risk	-0.035	-0.026	0.008
	(0.008)	(0.011)	(0.013)
(2) Exit to Emp	-0.001	-0.002	0.004
	(0.007)	(0.009)	(0.009)
(3) Exit from LF	-0.031	-0.021	0.004
	(0.005)	(0.008)	(0.008)

Note: Columns 1 and 2 present the average marginal effect on the exit probability of an indicator for availability of UI benefits (a transformation of $\hat{\delta}$). Column 3 presents the average marginal effect of an indicator for the *loss* of benefits (a transformation of $\hat{\theta}$), controlling for an indicator for simulated benefit eligibility in the pre-expiration period. See text for a list of other controls included in the models and for spell counts.

Consistent with our earlier work, the estimates for the competing risks of exit to employment and exit from the labor force (rows 2 and 3 of the table respectively) imply that the UI effect on overall exits is primarily driven by the participation margin rather than the employment margin. We find very small, statistically insignificant effects on exit to employment in both periods. However, available UI benefits do have statistically and economically significant negative effects on labor force (LF) exit. The effect is 3.1 percentage points in 2008-2011 and 2.1 percentage points in 2012-2014. With an average rate of labor force exit of about 10 percentage points (figure 2), these estimates imply that the availability of extended benefits reduced the rate of labor force exit by those eligible for extended UI benefits by 20 to 30 percent. We cannot reject equality of coefficients across the two periods in any of our three models.

Column 3 contains results from an alternative specification that identifies the effect of UI availability only from variation due to the phase-out of EUC and EB. Specifically, we identify duration ranges that were covered by EUC and EB in each state prior to each program's disappearance, and we examine how unemployment exit rates for individuals in those ranges change following the relevant program's disappearance.

To implement this, we define four new variables: $postEUC_t$ and $postEB_{st}$, indicators for observations after the end of EUC and EB benefits, respectively, in state s, and $EUCrange_{ist}$ and $EBrange_{ist}$, indica-

tors for unemployment durations that were covered by EUC and EB in the last month in which the relevant program was available in the state. These vary across states based on the specific benefit durations. For example, California offers 26 weeks of regular benefits. Before the state lost EB eligibility in April 2012, it had 53 weeks of EUC and 20 weeks of EB benefits. Thus, in California $EBrange_{ist}$ is an indicator for an unemployment duration between 80 and 99 weeks. Subsequent EUC changes reduced California's EUC benefits to 37 weeks before it expired in December 2013, so in California $EUCrange_{ist}$ is an indicator for a duration between 27 and 63 weeks. Across states, individuals with $EUCrange_{ist} = 1$ would have received EUC benefits in the months before December 2013 but not afterward, while those with $EBrange_{ist} = 1$ would have received EB benefits in the months before their state's EB program phased out but not afterward.

We model the latent propensity to exit unemployment as

$$
\begin{aligned}
y^*_{ist} = {} & X_{ist}\beta + Z_{st}\lambda + \gamma postEB_{st} \\
& + \pi_{EUC} EUCrange_{ist} + \pi_{EB} EBrange_{ist} \\
& + \theta \max\{EUCrange_{ist} * postEUC_t, \\
& \qquad\qquad EBrange_{ist} * postEB_{st}\} \\
& + \omega_s + \psi_t + \mu_{ist}.
\end{aligned}
$$

The coefficient of interest is θ, capturing the change in the exit hazard at the relevant durations for those who have lost benefits due to elimination of the EB or EUC pro-

grams in a state.[10] We expect θ to have the opposite sign from δ, as it reflects the effect of *not* having access to UI benefits.[11] Point estimates are much smaller in magnitude here than in Columns 1-2. We can never reject zero effects and, based on the implied confidence intervals, we can rule out quantitatively large effects on either exit margin.

IV. Final Comments

We draw two primary conclusions from our analysis. First, there is no evidence that the effects of UI were larger in the later period. This is consistent with the finding from Farber and Valletta (2015) of little difference in UI effects between the Great Recession period and the earlier episode of benefit extensions during the less severe recession of the early 2000s. Second, the primary effects of extended UI are on labor force attachment rather than job finding. From this perspective, the phasing out of extended and emergency benefits in 2013-2014 reduced the unemployment rate mainly by moving people out of the labor force rather than by increasing the job-finding rate.

One implication of our results is that the expiration of extended UI benefits may have put downward pressure on the labor force participation rate in 2012 and thereafter. However, our estimates imply that this effect is small. At any given time only about 15effect of UI eligibility on the probability of labor force exit of 2.1 percentage points (row 3, column 2 of Table 1), a back-of-the-envelope calculation implies that the rollback of extended UI reduced the labor force participation rate in early 2014 by only about 0.1 percentage point. We conclude that the phaseout of extended unemployment benefits is not an important factor in explaining the failure of the labor force to grow rapidly, despite steadily improving labor market conditions, since 2012.

An even clearer conclusion is that the UI extensions have not had large moral hazard effects on recipients' job-finding rates, either during the worst period of the Great Recession or during the subsequent recovery. Our point estimates suggest a near-zero effect in each period, and confidence intervals are small enough to rule out any quantitatively important effect. This suggests that the recent UI extensions have had very limited impacts on labor market efficiency.

REFERENCES

Abowd, John M., and Arnold Zellner. 1985. "Estimating Gross Labor-Force Flows." *Journal of Business and Economic Statistics* 3 (3): 254-83.

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.

Farber, Henry S., and Robert G. Valletta. 2015. "Do Extended Unemployment Benefits Lengthen Unemployment Spells? Evidence from Recent Cycles in the U.S. Labor Market." *Journal of Human Resources* (in press).

Hagedorn, Marcus, Fatih Karahan, Iourii Manovskii, and Kurt Mitman. 2013. "Unemployment Benefits and Unemployment in the Great Recession: The Role of Macro Effects." NBER Working Paper No. 19499, (2013).

Kroft, Kory, and Matthew J. Notowidigdo. 2011. "Should Unemployment Insurance Vary with the Local Unemployment Rate? Theory and Evidence." NBER Working Paper No. 17173.

Rothstein, Jesse. 2011. "Unemployment Insurance and Job Search in the Great Recession" *Brookings Papers on Economic Activity* (Fall).

---

[10]Note that the main effect of $postEUC_t$ is absorbed by the calendar month controls.

[11]In a linear probability model for UI_{ist} that includes the other controls, the coefficient on the benefit loss interaction is -0.58 (S.E. 0.01).