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EFFECTS OF THE UI BENEFIT EXTENSIONS: EVIDENCE FROM THE MONTHLY CPS

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Effects of the UI Benefit Extensions: Evidence from the Monthly CPS*

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Abstract

Using the monthly CPS, I estimate unemployment-to-employment (UE) transition rates and unemployment-to-inactivity (UN) transition rates by unemployment duration for male workers. When estimated for the period of 2004-2007, during which no extended benefits are available, both of the transition-rate profiles show clear patterns consistent with the expiration of regular benefits at 26 weeks. These patterns largely disappear in the profiles for the period of 2009-2010, during which large-scale extensions have become available. I conduct counterfactual experiments in which the estimated profiles for 2009-2010 are replaced by the hypothetical profiles inferred from the ones for 2004-2007. The results indicate that the benefit extensions in recent years have raised male workers' unemployment rate by 0.9-1.7 percentage points. Roughly 50-60% of the total increase is attributed to the effects on UE transition rates and the remaining part is accounted for by the effects on UN transition rates.

JEL codes: J08, J64, J65.

Keywords: Unemployment duration, unemployment insurance, CPS.

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1 Introduction

Whether or not the large-scale extensions of UI benefits in recent years have contributed to raising the observed unemployment rate is an important policy question. As far as I know, there is no statistical examination of this issue directly using recent data, even though the effects of UI benefits on workers' search behavior in general have been studied many times in the past. In this paper, I use the monthly CPS data to estimate unemployment-to-employment (UE) transition rates and unemployment-to-inactivity (UN) transition rates by unemployment duration. I do this for all male workers.

Existing studies are often based on the UI administrative records and typically look at whether there is any spike in the *exit* rate around the expiration dates, without distinguishing between job finding and dropping out of the labor force.[1] It is important to distinguish between the two outcomes since a spike in the UN transition rate can occur without any change in workers' search behavior.[2]

I use the monthly CPS to estimate the two exit rates. The estimates based on survey data such as the CPS may not be as precise as those based on administrative records. However, there are several advantages. First, the CPS allows me to calculate the UE and UN transition rates separately, since tracking workers' labor market status is one of the main purposes of the CPS. Second, it is pretty much the only publicly available data source that covers the most recent recession. Last, it is used for the BLS's official labor market statistics such as the unemployment rate.[3] This allows me to easily translate the effects on transition rates into the unemployment rate.

I use the information regarding unemployment duration and worker transitions between labor market statuses obtained from matching individuals who are in the CPS for two consecutive months. In combining the two pieces of information, I first show that the classification error in the worker transition data creates a striking inconsistency between the duration information and the labor market transition data. I propose a simple yet novel procedure to correct the classification error. Using the classification-error adjusted data, I estimate UE and UN transition rates by unemployment duration, by estimating multinomial logit regressions. I estimate the transition-rate profiles for two different samples: 2004−2007 and 2009−2010 (up to July). I show that in the first sample period where no extended benefits are available, the UE transition rates exhibit a clear hump and UN transition rates exhibit a large spike around the time of the benefits expiration date. In the latter sample when benefits expiration dates are greatly extended, the pattern observed for the earlier sample

[1] The seminal work based on the administrative records includes Moffitt (1985), Katz and Meyer (1990), and Meyer (1990).

[2] Using high-quality administrative data in Austria, which can trace workers' labor market status after the exit, Card et al. (2007) find that the spike in the exit rate is largely driven by dropping out of the labor force and they call this effect the "reporting effect."

[3] The studies that use the survey data include Fallick (1991).

largely disappears.

I translate the effects on the transition-rate profiles into the unemployment rate by proposing counterfactual transition-rate profiles in the absence of the benefit extensions. The counterfactuals are inferred from the shape of the transition-rate profiles estimated for 2004-07. Using the counterfactual profiles together with the labor market flow data for 2009-2010, I estimate that the effect on the unemployment rate amounts to 0.9-1.7 percentage points, of which 50-60% are attributed to the effects on UE transition rates and the rest on UN transition rates. In Section 2, I briefly describe the methodology and the data issues, including the inconsistency between the duration information and the labor market transition data in the CPS. Section 3 estimates and compares the transition-rate profiles for the two time periods. Section 4 then translates the effects on the transition rates into the effects on the unemployment rate. Section 4 concludes the paper.

2 Data and Methodology

The CPS surveys a large sample of individual U.S. workers each month, ascertaining whether they are employed and, if not employed, whether they engaged in active job search activities over the preceding month. After entering the sample, a household is surveyed for four consecutive months. Following an eight-month hiatus, it is surveyed again for four consecutive months. Month-over-month transitions between employed, unemployed, and NILF (not-in-the-labor-force) status can be measured by matching individual workers that are in the CPS sample in two consecutive months. Owing to the sample rotation and the eight-month gap, at most 75% of individual workers in the sample can be matched. This study is based on this monthly matched CPS data.[4] Each two-period panel contains the information regarding workers' labor market transitions from the first month to the next month. It also includes unemployment duration (if the worker is unemployed) as well as other standard observable characteristics. I can thus calculate the UE transition rates and UN transition rates by duration. The analysis in this paper focuses on the sample of male workers, given that women are often secondary earners in a household, which may complicate the interpretation of the results.

I run multinomial logit regressions where the dependent variable is the labor market status in the second month. There are three outcomes: "stay unemployed," "employed (E)," and "out of the labor force (N)." In addition to unemployment duration, time dummies, age, age squared, education, gender, and race are also included in the regressions. Using the estimated coefficients, I can calculate average predicted transition probabilities at each duration. The duration information is included as a categorical variable rather than a continuous variable.

[4] I use Robert Shimer's matching algorithm. I would like to thank him for posting the Stata code on his website.

Reported unemployment duration is classified into the following 12 categories: less than or equal to 4 weeks, 5-8 weeks, 9-12 weeks, 13-16 weeks, 17-20 weeks, 21-24 weeks, 25-28 weeks, 29-32 weeks, 33-40 weeks, 41-68 weeks, 69-96 weeks, 97 weeks or more.

I am interested in the comparison between 2004-2007 and 2009-2010 (January through August). During the former period, no UI benefit extensions were enacted. Therefore, the regular entitlement period of 26 weeks applies to this sample. In the latter sample, the entitlement period is extended a number of times in response to the severe recessions started in late 2007. The maximum entitlement period has reached 99 weeks in November 2009 and has not been changed since. An important assumption is that those who are in the latter sample expect to get some kind of extension of benefits beyond the regular 26 weeks (if not 99 weeks).[5]

2.1 Classification-Error Adjustment

In the CPS, individuals may misreport their employment status and the nature of their job search activities, causing transitions to be mis-measured. This is referred to as the classification error. Poterba and Summers (1984, 1986) made the earlier attempts to quantify the effects of the classification error for the aggregate labor flow data. Their corrections are made by comparing the labor market status reported in the original survey with that reported in the reinterview, which is considered to be truthful. This comparison makes it possible to estimate the response error matrix, which can be used to obtain the classification-error adjusted gross flows. Unfortunately, the reinterview data are no longer available.[6]

While other ways to correct the error without relying on the reinterview data are proposed in the literature, this paper proposes a very simple way to correct the classification error without using the reinterview data. I use the duration information to correct the misclassification of labor market status. But let me first highlight the striking inconsistency between the duration information and the labor-market transition data. Note first that, since the redesign of the CPS in 1994, if workers are unemployed in the two consecutive surveys, unemployment duration is simply extended by either 4 or 5 weeks in the second survey. If workers are in the incoming rotation groups (i.e., those who are in the first and fifth surveys after the eight-month hiatus) and are unemployed at that point, they are asked their duration. Last, when workers are employed or out of the labor force in the previous survey and are unemployed in the current month, the CPS asks their unemployment dura-

[5] See, for example, Whittaker (2008) and Fujita (2010) for details of the extensions. One may think that using the data after November 2009, I can examine if exit rates increase around 99 weeks. However, the CPS data include fewer observations at such long duration. Moreover, most of the observations are clustered at or around 104 weeks (2 years), which suggests that reported durations there are subject to serious measurement problems. This makes it difficult to distinguish workers based on the 99-week cutoff date.

[6] In Fujita and Ramey (2006), we use the response error matrix estimated by Poterba and Summers (1986) to obtain classification-error adjusted gross flow series for 1976-2005.

Table 1: Misclassification of Labor Market Status

Status in month 1	Duration in month 2 (weeks) 5 or less	6 or more	Total	Error rates (%)
Employed	9,155	2,506	11,661	8.0
	(78.5)	(21.5)	(100.0)	
Unemployed	2,065	21,360	23,425	–
	(8.8)	(91.18)	(100.0)	
NILF	4,339	5,527	9,886	17.6
	(44.0)	(56.0)	(100.0)	
Total	15,559	29,393	44,952	25.6
	(34.6)	(65.39)	(100.0)	

Notes: Distribution of unemployment duration by labor market status in the previous month's survey. Fractions within the previous month's labor market status are in parentheses. The last column reports the classification error rates calculated as the probabilities of misreporting given that the true state is unemployed. Sample period: 2004-07.

tion. For this last case, I can check whether the duration information reported is consistent with labor market status. That is, the duration of those who were employed or out of the labor force in the first survey cannot logically exceed 4 or 5 weeks depending on the month's calendar. However, this logical condition is greatly violated in the actual data.

Table 1 presents the distribution of unemployed workers by duration. The first row looks at employed (E) workers in the previous month's survey, the second row looks at those who previously report "unemployed" (U), and the third row looks at those who report "not-in-the-labor force" (N). If the labor market status in the previous month's survey is correctly reported, unemployment duration for those who report E or N in the previous month's survey needs to be less than 4 or 5 weeks by construction. However, a large fraction of workers give unemployment durations inconsistent with the previous month's labor market status. The first row shows that more than 20% of individuals who report being employed in month 1 say "unemployed" in month 2 *and* give an unemployment duration of more than 5 weeks in month 2. The problem is a lot more severe for those who report being out of the labor force in month 1. Roughly 60% of these workers who say "unemployed" in month 2 report durations inconsistent with the previous month's labor market status. There are two ways to go from here. The first is to take the labor market status as truth, overwriting the duration information, in which case, the reported duration information is overwritten as "0-4" or "0-5 weeks."[7] The second approach is to take the duration information as given and renew the

[7] I cannot pin down the exact length because the CPS does not specify which week workers become

labor market status in the previous month. That is, if the duration is greater than or equal to 4 or 5 weeks, I replace the previous month's labor market status with "unemployed" and calculate the associated duration by subtracting 4 or 5 from the unemployment duration reported in month 2. I follow this second approach. The reason is that it is well known that the classification error is highest for unemployed individuals. For example, Poterba and Summers (1986) estimate that the error reporting rate that unemployed workers are misclassified as being out of the labor force is greater than 11%. Similarly, the error rate that unemployed workers are misclassified as having a job is around 4%. I can also calculate error reporting rates using the data presented in Table 1 under the assumption that the duration information is correctly reported.[8] They can be calculated by

$$\Pr(S_R = E | S_T = U) = \frac{2,506}{2,506 + 23,425 + 5,527} = 0.080,$$

$$\Pr(S_R = N | S_T = U) = \frac{5,527}{2,506 + 23,425 + 5,527} = 0.176,$$

where S_T is the true labor market status and S_R is the reported labor market status. These error rates are somewhat higher than those reported by Poterba and Summers (1986). However, other, more recent studies that use different methodologies and data from different sample periods estimate the error rates that are surprisingly close to my estimates.[9]

In particular, the classification error that creates the spurious flow from NILF to unemployment is an easily understandable phenomenon given the difficulty in drawing a clear line between unemployment and NILF. For example, it is easy to imagine a situation in which an individual has been looking for a job for a long time yet reports being out of the labor force in one particular survey. This situation is consistent with the data shown in Table 1.

An important consequence of this adjustment is the reduction in exit rates from the unemployment pool. Recall first that UE and UN transition rates are calculated by taking gross worker flows of workers who report being unemployed in month 1 and either employed or NILF in month 2, and divide each flow by all workers who report being unemployed in month 1. The classification error adjustment proposed above adds unemployed workers to month 1 based on the duration information observed in month 2. In the corrected data, these workers are recorded as unemployed in both months. Therefore, this adjustment makes both transition rates lower. Another way to appreciate this adjustment is that, without the adjustment, the exit rate from the unemployment pool based on gross flows (i.e., the job finding rate plus the dropout rate) is at a level much higher than those implied by the

unemployed.

[8]More precisely, the only necessary assumption is that a worker knows whether he or she has been looking for a job (i.e., unemployed) for more than one month. I do not need to assume that workers are reporting exactly how many weeks they are unemployed.

[9]See Table 3 in a recent paper by Feng and Hu (2010), which lists the error reporting rates estimated by seven different studies, including Poterba and Summers (1986).

(a) UE Transition Rates (b) UN Transition Rates

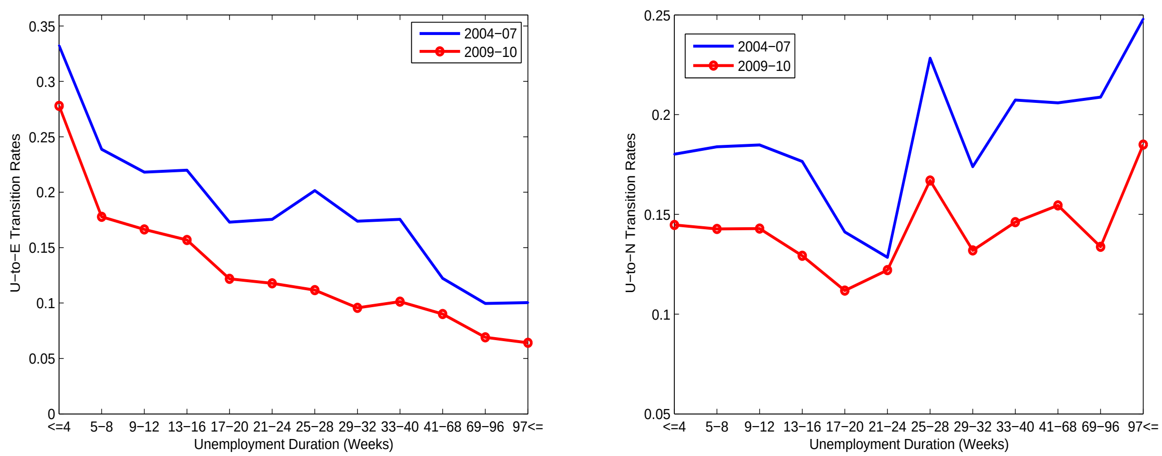


Figure 1: Average Predicted Exit Rates from Unemployment

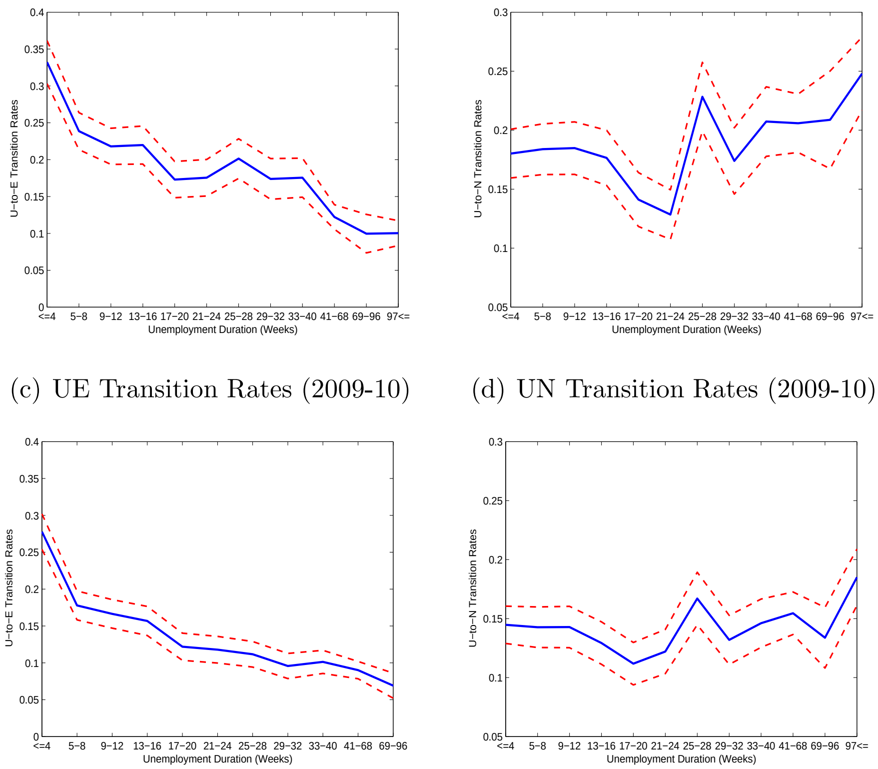
Notes: The figures plot average predicted transition rates for each duration category for all males. Sample sizes: 73,775 (2004-2007) and 55,251 (2009-2010). R^2: 0.0397 (2004-2007) and 0.0458 (2009-2010). The 2010 data include observations up to July.

duration information. This is especially important in recent years, when aggregate duration statistics such as median unemployment duration are at record-high levels. The classification error adjustment I propose makes the transition rates based on gross flow data more in line with the duration information.

3 Transition Rates by Duration

Figure 1 presents the average predicted transition rates by duration for the two time periods, using the classification-error adjusted data. Panel (a) looks at UE transition rates. Observe first that there is duration dependence for both periods. While neither of them shows a clear "spike," one can clearly see a hump around the 23-29 week bin for 2004-2007. For the recent one-and-half-year period, the declines in UE transition rates continue all the way to the longest duration category.

Panel (b) presents the transition rates into the not-in-the-labor force status. First, consider the one for the recent years (red line). While it shows some spiking at the 25-28 week bin, it is roughly constant all the way to the longest duration category. On the other hand, the UN transition rates for 2004-2007 show a much larger spike at the 25-28 week bin, which amounts to more than 10 percentage points and then stays at a level higher than before. It seems plausible to interpret this spike as being induced by the expiration of the UI benefits. Part of the observed jump may have to do with other factors such as measurement errors,



(a) UE Transition Rates (2004-07)
(b) UN Transition Rates (2004-07)
(c) UE Transition Rates (2009-10)
(d) UN Transition Rates (2009-10)

Figure 2: Exit Rates from Unemployment with Confidence Intervals

Notes: See notes to Figure 1. Dashed lines represent 95% confidence intervals that are calculated by the delta method.

given that UN transition rates for the 2009-10 sample years also show a spike at the same bin. However, the magnitude of the spike in the earlier period is much larger, and the fact that the UN rates increase to a permanently higher level is consistent with the idea that the jump is largely driven by the expiration of UI benefits. Figure 2 presents the same transition rates by duration, including the 95% confidence intervals. One can see that all of the predicted probabilities are fairly tightly estimated.

So far, I have focused on all male unemployed workers. However, some of the past research focused on a subset of unemployed workers, namely, "job losers" mainly because "job leavers (or quitters)" in principle do not qualify for UI benefits.[10] There are a couple of reasons why I would like to take the results for all male unemployed workers as my benchmark. First, in Section 4, I map the results on the transition rates into the effects on the unemployment rate,

[10]Note that in the CPS, there are six possible reasons for unemployment: (i) temporary layoffs, (ii) permanent job losers, (iii) persons who completed temporary jobs, (iv) job leavers, (v) reentrants, and (vi) new entrants.

and therefore, I would like to cover a lager population. Second, the distinction between job leavers and job losers in a household survey such as the CPS is inherently vague. Third, the qualification status of “labor market entrants” is not clear, either. It seems safe to assume that new labor market entrants do not qualify for UI benefits. However, the qualification status of “reentrants” is simply difficult to discern. In particular, as I have shown in the previous section, flows between U and N involve serious classification errors.[11]

Despite these reasons, it is still useful to see if I obtain results similar to those for all male unemployed workers even when I restrict the sample to (male) job losers. This category includes the second and third categories mentioned in footnote 10. In addition to excluding job leavers and labor market entrants, I exclude those who are on temporary layoffs because recalls to previous jobs that coincide with benefit expirations can create a spurious hump or spike around the 25-28 week bin. Figure 3 presents the results for this group. In general, the results here are stronger than those for all male unemployed workers (which makes sense). The hump in the UE transition rates for the 2004-07 period become even clearer, whereas there is no discernible hump in the UE transition rates for the sample of 2009-2010. The upward jump in the UN transition rates for the earlier sample period is again clearer: Although the size of the jump at the 25-28 week bin is now smaller (5.7 percentage points instead of 10 percentage points), that the UN transition rates go up to a permanently higher level after the 25-28 week bin is much easier to observe. On the other hand, the UN transition rates for the recent sample period are roughly flat, as in the case of including all reasons.

4 Mapping into the Unemployment Rate

The purpose of this section is to infer the differences in the steady-state unemployment rate with and without extensions of UI benefits, using the estimated transition rates. The idea is to replace the estimated profiles of UE and UN transition rates for 2009-2010 with the ones whose shape mimic the ones for the 2004-2007 period, taking into account the difference in the business cycle condition.

[11]A recent article by Valletta and Kuang (2010) estimates the effects of the UI benefit extensions under the assumption that job leavers (whose share in the unemployment pool is roughly 8% in the last two years) and labor market reentrants (whose share is roughly 25% in the last two years) do not qualify for UI benefits. The second and third reasons I mentioned above make me somewhat uncomfortable with the premise of their exercise. In fact, when I estimate the regressions using the sample of job leavers and labor market entrants, I observe a similar hump in the UE transition rates and a large spike in the UN transition rates around the 25-28 week bin for the 2004-2007 sample. Moreover, as mentioned in Barnichon and Figura (2010), a sizable fraction of workers who report being a job leaver in one month report being a job loser in the following month's survey, which is another indication that the distinction between the two groups of workers is quite vague in the CPS.

(a) UE Transition Rates (b) UN Transition Rates

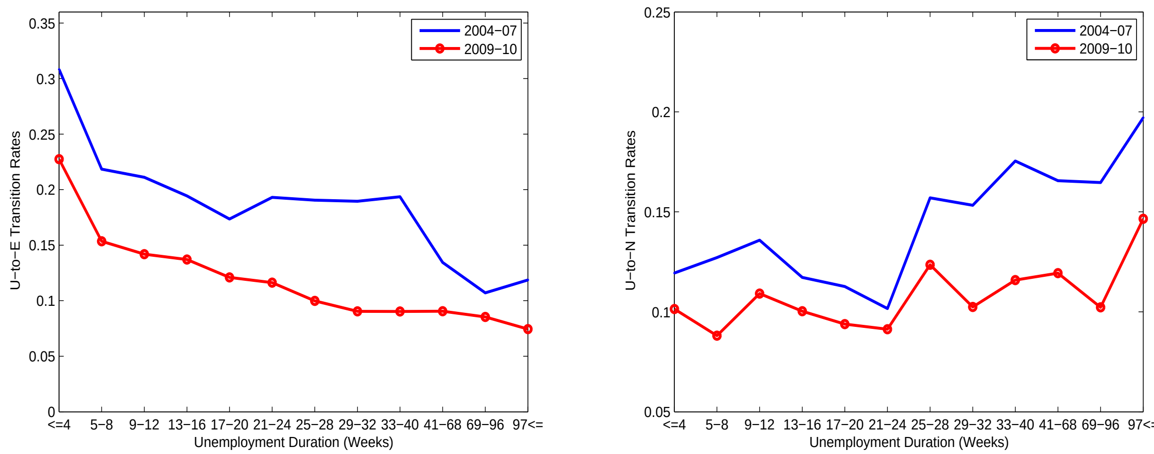


Figure 3: Exit Rates from Unemployment: Job Losers

Notes: The figures plot average predicted transition rates for each duration category for all males who are job losers excluding those who are on temporary layoffs. Sample sizes: 25,800 (2004-2007) and 28,226 (2009-2010). R^2: 0.0348 (2004-2007) and 0.0330 (2009-2010). The 2010 data include observations up to July.

4.1 Steady-State Conditions

First, I describe the derivation of the steady-state labor market quantities, when UE transition rates and UN transition rates vary by unemployment duration. I define the frequency distribution of unemployment over duration by $u(m)$. This function gives the number of unemployed at duration m, which is measured in months. The duration in the CPS is measured in weeks, and thus I convert the reported information into months because the labor flow series I need to use are available only at monthly frequency. Let w be unemployment duration measured in weeks. I convert w into m by the floor function:

$$m = \lfloor w/4 \rfloor.$$

Let $f(m)$ and $d(m)$ be the duration-dependent UE transition rates and UN transition rates. Note that the following relationship holds for $m > 1$:

$$u(m) = [1 - f(m-1) - d(m-1)]u(m-1). \tag{1}$$

For $m = 0$, the following relationship holds:

$$u(0) = p^{EU}e + p^{NU}n, \tag{2}$$

where p^{jk} represents the transition probability from state j to state k, and e and n are the stock of employed workers and those who are out of the labor force, respectively. Equation

(2) simply equates the number of the unemployed whose duration is less than a month and the inflows into the unemployment pool from the two possible sources. The recursion in equation (1) together with equation (2) can be expressed as:

$$u(m) = (p^{EU}e + p^{NU}n) \prod_{i=1}^{m+1} [1 - f(m+1-i) - d(m+1-i)], \tag{3}$$

for all $m \geq 0$ with $f(0) \equiv 0$ and $d(0) \equiv 0$. The total number of unemployment (u) can be calculated by

$$u = \sum_{m=0}^{\infty} u(m) = (p^{EU}E + p^{NU}N) \sum_{m=0}^{\infty} \prod_{i=1}^{m+1} [1 - f(m+1-i) - d(m+1-i)]. \tag{4}$$

Note also that hires from unemployment (h) and the flow from unemployment to out-of-the-labor force (g) are calculated by

$$h = \sum_{m=0}^{\infty} f(m)u(m), \tag{5}$$

and

$$g = \sum_{m=0}^{\infty} d(m)u(m). \tag{6}$$

The stock-flow balance equations for employment (e), unemployment (u), and not-in-the-labor force (n) are written as:

$$(p^{EU} + p^{EN})e = h + p^{NE}n, \tag{7}$$

$$h + g = p^{NU}n + p^{EU}e, \tag{8}$$

$$(p^{NU} + p^{NE})n = p^{EN}e + g, \tag{9}$$

$$\bar{p} = e + u + n, \tag{10}$$

where $\bar{p}$ is the working-age population. Note that Equation (8) is equivalent to Equation (4). Also note that one of the first three equations above is not independent. By dropping one of the three and adding the last equation for normalization, I can solve this system for e, u, n, h, g and $u(m)$ taking all transition rates as given.

4.2 Calibration

The basic idea of the experiment is to infer the counterfactual unemployment rate using the information on the transition-rate profiles observed in 2004-2007, when workers had no expectations of receiving UI benefits beyond 26 weeks. First, I have to set the four transition rates p^{eu}, p^{en}, p^{nu}, and p^{ne}. I use the BLS research series on labor force status flows from the CPS. The BLS releases all labor market flows for males and females separately. I calculate the four monthly transition rates based on the seasonally adjusted data for all males. They are set at the following values: $p^{eu} = 0.0200, p^{en} = 0.02194, p^{nu} = 0.0444,$ and $p^{ne} = 0.0472$, which correspond to the average rates over the period between January 2009 and August 2010.

I specify $f(m)$ and $d(m)$ as follows. I have to translate the duration-dependent transition rates estimated above into these two variables. Up to 32 weeks, duration originally reported in the CPS is classified into 8 four-week bins. After that, the widths of the bins are set longer than 4 weeks to accommodate smaller sample sizes at longer duration ranges in the CPS. Thus, for $m \leq 8$, I directly apply the estimated UE and UN transition rates to $f(m)$ and $d(m)$. For $9 \leq m \leq 10$, $11 \leq m \leq 17$, $18 \leq m \leq 24$, and $25 \geq m$, the same transition rates are applied within each bin.[12]

Having assigned numerical values to all parameters, I solve the system of equations for all quantities as follows. I first set an initial guess for $u(0)$ denoted by $u^{(1)}(0)$, then I can use (3) through (10) to solve for n, e, n and all $u(m)$ for $m \geq 1$. I can then calculate $u^{(2)}(0)$ from (2). Let $u^{(i)}(0)$ be the guess of $u(0)$ in the ith iteration. When $|u^{(i)}(0) - u^{(i-1)}(0)| > \kappa$ where κ is some small number, I set a new guess by $u^{(i+1)}(0) = \frac{1}{2}[u^{(i)}(0) + u^{(i-1)}(0)]$. If $|u^{(i)}(0) - u^{(i-1)}(0)| \leq \kappa$, then the algorithm stops.

4.3 Counterfactual Experiments

A couple of important assumptions underlie the counterfactual experiments below. First, it is assumed that in the absence of the benefit extensions, the shapes of the UE and UN transition rate functions are the same as those estimated for 2004-2007. In other words, I assume that the business cycle conditions would have only shifted the profiles down. This assumption excludes the possibility that the exit rates around the 25-28 week bin are affected more strongly by the recession. I am not aware of any theories that generate such nonlinear effects. Second, I also assume (in the benchmark experiment) that the transition rates in the absence of the benefit extensions cannot be lower than those with the benefit extensions (i.e., the ones actually estimated from the 2009-2010 data) at *any* unemployment duration. This second assumption is plausible given that extensions of UI benefits cannot lower the value of job search at any duration relative to the other two labor market states. Also note that this assumption assumes away the magnification of the direct effect that could result from the interaction between firms' vacancy posting and workers' search activity. That is, if workers' search efforts are adversely affected by the benefit extensions, firms would also have weaker incentives to post vacancies, which would, in turn, lower workers' job finding rates. This magnification effect exists in a standard search/matching model of Mortensen and

[12]The largest number of m I consider is 100. There are virtually no unemployed workers at $m = 100$.

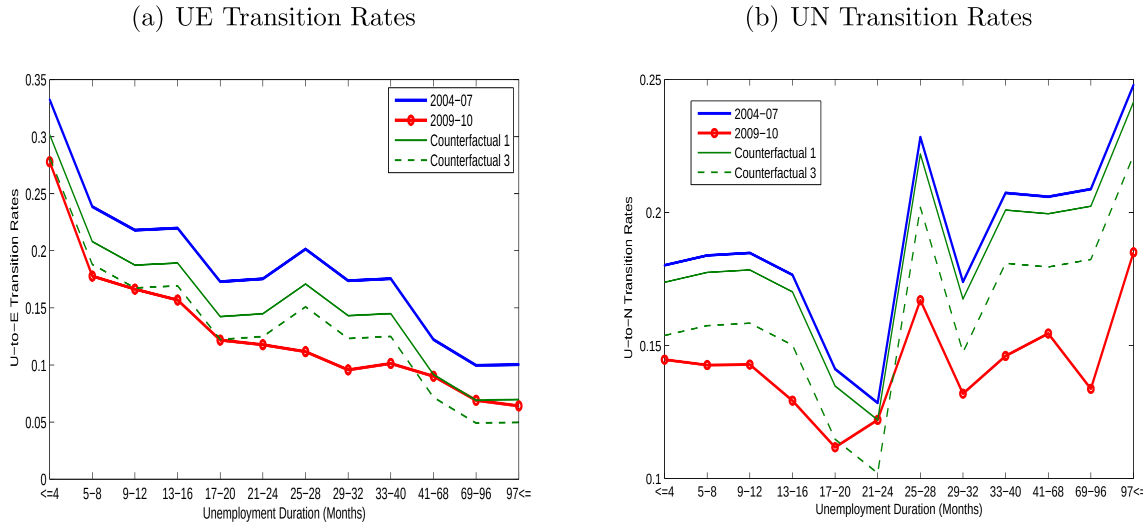


Figure 4: Counterfactual Transition Rate Profiles

Notes: The green lines give counterfactual UE and UN transition-rate profiles under the absence of the benefit extensions, obtained by shifting downward the estimated profiles for 2004-2007. The solid green lines ("Counterfactual 1") are obtained as the largest shifts under the condition that transition rates under the presence of the benefit extensions cannot be lower than those in the absence of the extensions at any duration. The dashed green lines ("Counterfactual 3") assume additional shifts by 2 percentage points for both profiles.

Pissarides (1994), where vacancy posting is endogenously influenced by the workers' search activity. I ignore this magnification effect in my experiments because I cannot quantify this effect precisely. But note that ignoring this effect only makes my experiments more "conservative," (i.e., it makes the effects of the extensions smaller).

Figure 4 presents what the counterfactual profiles look like. The solid green lines (labeled as "Counterfactual 1") correspond to the profiles that come out under the two assumptions I have just made. The blue and red lines are estimated profiles previously presented in Figure 1. Note that the second condition determines the size of the shift by the smallest difference between the 2004-07 profile and 2009-2010 profile across all duration bins. For example, it occurs at the 69-96 week bin for the UE profile and at the 21-24 week bin for the UN profile. The sizes of the shifts are 3.06% and 0.64%, respectively. First, consider panel (a). One can see that the green solid line never goes below the red dotted line. Under this counterfactual scenario, the unemployment rate would be lower, because workers find jobs faster at all durations. Of course, that is particularly true around the 25-28 week bin. Calculating the steady-state unemployment rates under the observed profiles for 2009-2010 and the counterfactual profiles, I find that the difference in the steady-state unemployment rates amounts to 1.7 percentage points. I can also calculate the contribution of each transition rate by shifting only one profile, keeping the other at the estimated level for 2009-10. The contribution of the UE profile amounts to 1.09 percentage points, whereas that of the UN

Table 2: Effects on the Unemployment Rate

		ΔUN			2009
		-0.64	-1.64	-2.04	-10
	-3.06	1.74	1.56	1.37	1.09
ΔUE	-4.06	1.43	1.23	1.02	0.71
	-5.06	1.09	0.88	0.65	0.31
$2009-10$		0.84	0.60	0.36	–

Notes: Each entry in the table gives the estimate of the effect of the benefit extensions on the unemployment rate. ΔUE and ΔUN, respectively, give the size of the shift of the UE and UN transition-rate profiles. The fourth column and fourth row, respectively, give the effect when either the UN or UE profile is held fixed at the estimated level for 2009-2010 and thus give the contribution from UE and UN transition rates. The combination $(\Delta UE, \Delta UN) = (-3.06, -0.64)$ corresponds to the largest shifts in the profiles under the condition that transition rates in the presence of the benefit extensions cannot be lower than those without extensions at any duration.

profile is 0.84 percentage point.[13]

Table 2 presents the effects on the unemployment rate under various scenarios. The upper-left entry corresponds to the benchmark counterfactual scenario just discussed. The other scenarios considered in the table are more "conservative" in the sense that the sizes of the shifts are larger than those examined in the benchmark counterfactual scenario. There are a couple of reasons to consider these conservative scenarios. The first is based on the statistical precision of the estimates of the profiles, even though the confidence intervals are quite narrow as shown before. The second reason applies to the UN profile: As the blue line in panel (b) in Figure 4 shows, UN transition rates for 2004-2007 exhibit significant declines before the large jump at the 23-28 week bin. A part of this behavior seems to be due to measurement problems because the same qualitative pattern is observed for the 2009-2010 profile as well. However, the initial decline and the subsequent jump are much more pronounced for the 2004-2007 profile. To understand the behavior, suppose that a worker has no willingness to find a job but receives UI benefits by manipulating his eligibility. Consider the two different economies facing this worker: one where UI benefits last 26 weeks and the other where UI benefits last longer. Further, assume that receiving UI benefits involves some time cost (filing the form, reporting to the UI office, etc.). The worker in the first economy has a strong incentive to substitute the time for receiving the benefits for his leisure time,

[13]The sum of the two contributions do not exactly add up to the total effect of 1.7 percentage points because of nonlinearity.

especially around the expiration date. In this case, the UN profile in the first economy may come above that in the second economy, especially around the expiration date. Therefore, I also consider two other counterfactuals that allow the transition rates to be lower at some points over duration. The two alternative counterfactuals add 1 percentage point and 2 percentage points to the shift of the two profiles.

The profiles corresponding to shifting an additional 2 percentage points are plotted by dashed green lines ("Counterfactual 3") in Figure 4. I consider this scenario to be the upper bound on the size of the downward shifts. As can be seen from panel (a), up to the 21-24 week bin, the counterfactual UE profile almost coincides with the actual profile, which is, theoretically speaking, quite unlikely. Furthermore, at the last three bins, the counterfactual UE transition rates are lower than the actual rates. Recall that, since November 2009, the actual maximum entitlement period has been 99 weeks.[14] Given that UI benefits would not be available beyond 26 weeks under the counterfactual scenario, the fact that UE transition rates at the last three bins become lower under the counterfactual scenario does not appear plausible.

Table 2 summarizes the effects on the unemployment rate under various combinations of the counterfactual UE and UN profiles. The most conservative case of $(\Delta UE, \Delta UN) = (-5.06, -2.64)$ estimates the effect on the unemployment rate being 0.65 percentage point. Excluding this case, the smallest effect is 0.88 percentage point. Thus, an implication of Table 2 is that the UI benefit extensions have had the effect of raising the unemployment rate by 0.9-1.7 percentage points.

The figures in the last column and last row allow me to decompose the total effect into the contribution of each transition-rate profile. For example, in the benchmark counterfactual scenario, roughly 60% is attributed to the effects through the changes in UE transition rates and the remaining part is explained by the effects through UN transition rates.

5 Conclusion

This paper has estimated the effects of the UI benefit extensions in recent years on the two exit rates from the unemployment pool. The results are based on the monthly CPS data that are adjusted for the classification error. I find large differences in the shape of UE and UN transition-rate profiles between the 2004-2007 period and the 2009-2010 period. The inference I draw from the empirical exercise is that the UI extensions have had measurable impacts on workers' exit rates from the unemployment pool. Using counterfactual transition-rate profiles inferred from the 2004-07 data, I find that the benefit extensions have raised

[14]This does not mean that all unemployed workers receive UI benefits for 99 weeks. Individuals can only complete the "tier" that they are in at the time of the expiration date, which is currently set at the end of November 2010. See, for example, Whittaker (2008) and Fujita (2010) for details.

the unemployment rate for male workers by 0.9-1.7 percentage points.

The purpose of this paper is *not* to provide any normative conclusions regarding the welfare implications of the extended UI benefits but to conduct a statistical analysis using the currently available information. It is possible that longer duration is welfare improving, for example, due to the liquidity effect (Chetty (2008)) and the job creation effect (Acemoglu and Shimer (2000)). I hope that the estimates given in this paper provide an important basis for those normative discussions.

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