The Spike at Benefit Exhaustion:
Leaving the Unemployment System or Starting a New Job?

By David Card, Raj Chetty, and Andrea Weber[1]

One of the best-known empirical results in public finance and labor economics is the "spike" in the exit rate from unemployment around the expiration of jobless benefits (see e.g., Robert Moffitt, 1985; Lawrence Katz and Bruce Meyer, 1990a; Katz and Meyer, 1990b). This sharp surge in the hazard rate is widely interpreted as evidence that recipients are waiting until their benefits run out to return to work. The spike in exit rates has become a leading example of the distortionary effects of unemployment insurance (UI) and social insurance programs more generally (see e.g., Martin Feldstein, 2005).

In this paper, we present the results of a meta-analysis of the literature on unemployment exit rates around benefit exhaustion, as well as new evidence using administrative data for a large sample of Austrian job losers. Our main finding is that the way in which unemployment spells are measured has a large effect on the magnitude of the spike at exhaustion, both in existing studies and in our Austrian data. Spikes are generally smaller when the spells are measured by the time to next job than when they are defined by the time spent on the unemployment system. In the Austrian data, we find a large spike in the exit rate from *registered unemployment* at the point of benefit exhaustion, consistent with earlier studies (e.g., Rafael Lalive et. al., 2007). However, the hazard of *re-employment* rises only slightly at the same point. Even recalls to the previous employer – which account for one-fifth of spell terminations in our data – increase by no more than 20% at benefit exhaustion.

We conclude that most job seekers in Austria are *not* waiting to return to work until their UI benefits are exhausted. Rather, a large fraction simply leave the unemployment registry once their benefits end and they are no longer required to register to maintain their eligibility for benefits. This finding underscores the importance of distinguishing between the effects of government programs such as UI on the decision to work vs. their auxiliary effects on the classification of non-working time.[2] The effect of UI on whether individuals choose to be

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[2]A similar point is made by David Card and W. Craig Riddell (1993) in explaining the much larger divergence between unemployment rates in Canada and the U.S. than between their employment rates.

classified as “unemployed” or “out of the labor force” can be relatively large because it requires little change in actual behavior. The distortionary costs of the program, however, ultimately depend on how UI affects the time spent working. Hence, the spike in unemployment-exit hazards may substantially overstate the degree of moral hazard induced by UI. The “reporting” effects induced by UI benefits can also lead to biases in comparisons of unemployment rates across countries that have different unemployment systems.

I Summary of the Existing Literature

We provide a detailed review of the existing literature on spikes at exhaustion in Card, Chetty, and Weber (2007). We summarize the main lessons that emerge from our meta-analysis here.

Existing studies have used three alternative measures of duration: the length of benefit receipt, the duration of registered unemployment, and the duration of non-employment (time to next job). Although these measures are equivalent in simple theoretical models, in practice there are important differences between them. The duration of benefit receipt is important for measuring program costs. The time to re-employment is more relevant for calculating efficiency costs and optimal benefits (e.g. using the elasticity approach of Martin Baily (1978) and Raj Chetty (2006)).

Studies that focus on durations of compensated unemployment find some evidence of a rise in exits from the unemployment system prior to benefit exhaustion. A concern in interpreting the pattern of exits from the unemployment system is that exit hazards may rise artificially prior to exhaustion because of measurement error in the computation of potential duration. In addition, Katz and Meyer (1990a) note that many UI recipients are eligible for only a small “partial payment” for their last week of benefits, and may fail to pick up their last check (and thereby exit the system early) for this reason.

Studies that focus on time to next job find little evidence of an increase in re-employment rates prior to benefit exhaustion. Some (but not all) of the studies using survey data to measure job starts find evidence of a spike in the re-employment hazard *at* benefit exhaustion. A concern in survey data is that recall errors can lead to irregular patterns; for example, Katz (1985) notes that survey-based spell data often exhibit heaping at exactly 26, 39, and 52 weeks. More recent studies using administrative data on time to next job find a relatively smooth re-employment hazard at benefit exhaustion.[3]

Overall, our reading of the existing literature is that spikes in hazards around benefit exhaustion are generally smaller when duration is measured as time to next job rather than time unemployed.

II Austrian Data and Institutional Background

The Austrian labor market is comparable to the U.S. in terms of its job turnover and unemployment rates (Alfred Stiglbauer et. al., 2003). In Austria, job losers who have worked for 12 months or more in the preceding two years are eligible for benefits that replace approximately 55% of their previous (after-tax) wage. The maximum duration of benefits depends discontinuously on the number of months that the individual worked in the past five years. Individuals with under 36 months of employment can receive up to 20 weeks of benefits, while those who have worked for 36 months or more can receive 30 weeks of UI.

We analyze the effect of benefit expiration on labor market transitions using data from the Austrian social security registry. Our sample includes all spells of UI associated with job losses between 1981 and 2001, with three restrictions. First, we limit the sample to individuals between age 20 and 50. Second, we focus on individuals who take up UI benefits within 28 days of job loss, thereby excluding voluntary quitters (who face a 28-day waiting period for eligibility). Finally, we focus on individuals who (a) have worked at their prior employer for at least one year and (b) have worked for between 33 and 38 months (3 years +/-3 months) in the past five years. These restrictions allow us to focus on a relatively homogeneous sample of job losers who are eligible either for 20 or 30 weeks of benefits, depending on whether they worked a little more or a little less than 36 months in the past 5 years, as in a regression-discontinuity analysis. Our final sample contains 92,969 unemployment spells, of which 47% are eligible for 20 weeks of UI benefits and 53% for 30 weeks of benefits. Additional details on the database and summary statistics are given in Card, Chetty, and Weber (2007).

There are two measures of spell length in the data. The first is "unemployment duration,"

[3] An exception is a study by van Ours and Vodopivec (2006), which finds a large and clearly discernible spike in the re-employment hazard at the point of UI exhaustion in Slovenia. One explanation for the large spike in the Slovenian case – emphasized by Vodopivec (1995) – is that many benefit recipients are working in the informal sector and waiting until their benefits expire to return to the formal sector. Such behavior is presumably less likely in more developed countries, where the informal sector is small.

the measure used in Lalive et. al.'s (2007) analysis of the Austrian data. This is defined as the total number of days that an individual is registered with the unemployment agency. Individuals are required to register while they are receiving benefits, and can remain registered afterward to take advantage of job search assistance services. Spells of registered unemployment are relatively short, as in the U.S.: the median spell length is less than 3 months, 67% of spells end within 20 weeks, and 95% end within a year.

The second duration measure, which we term "time to next job," is the amount of time that elapses from the end of the previous job to the start of the next job. The median time to next job is 3.9 months. Fifty-five percent of individuals find a new job within 20 weeks, and 80% find a new job within one year.

III Empirical Results

Let $h^U_{t,T}$ denote the unemployment exit hazard in week t for an individual who is eligible for T weeks of UI benefits. Similarly, let $h^J_{t,T}$ denote the job finding hazard. Figure 1a plots $h^U_{t,20}$ and $h^J_{t,20}$. Figure 1b shows the corresponding series for individuals eligible for 30 weeks of benefits. In both figures, there is a sharp spike in the unemployment exit hazard in the week of benefit exhaustion ($t = T$), and a relatively high unemployment exit rate in the weeks immediately after exhaustion (consistent with Lalive et al., 2007). The corresponding changes in the job finding hazards, however, are very small.

Next, we study the effect of potential duration on the hazard rate by examining the difference in the hazard rates between individuals eligible for 20 and 30 weeks of UI. Define $d^U_t = h^U_{t,20} - h^U_{t,30}$ and $d^J_t = h^J_{t,20} - h^J_{t,30}$. Figure 3 plots d^U_t and d^J_t. Observe that d^J_t is positive for all $t < 20$, indicating that individuals eligible for 20 weeks of benefits search harder to find a job throughout the unemployment spell (anticipating their shorter duration of benefits), and not just at the point when benefits are exhausted. There is a small increase in d^J_t from week 19 to 23: 20-week eligibles are somewhat more likely to find jobs in those weeks than the 30-week eligibles. Conversely, individuals eligible for 30 weeks of benefits are slightly more likely to find jobs from week 29 to 32. These patterns contrast sharply with the corresponding series of differences in unemployment exit rates (d^U_t), which exhibits sharp spikes at and after $t = 20$ and $t = 30$.

To quantify the size of the spikes in job-finding and unemployment-exit rates around benefit exhaustion, we estimate Cox proportional hazard models of the following form:

$$h_t = \alpha_t \exp(f(T-t) + \delta(T=30) + \gamma X) \tag{1}$$

where h_t denotes the hazard rate in week t, α_t denotes the "baseline" hazard rate in week t, X denotes a set of covariates, $f(T-t)$ is a function of the time-to-exhaustion $(T-t)$, and δ captures a proportional shift in the hazard for people with 30 weeks of eligibility. We censor all spells at 50 weeks to focus on hazards in the year after job loss. We use a spline function for f in order to allow different effects at different weeks, as in Meyer (1990). The coefficients of the exhaustion spline are identified despite the nonparametric baseline hazard by the difference in the hazard rates between the 20-week and 30-week eligibility groups at each t, as in Figure 3. The identification assumption is that the two groups would have similar hazard rates at each duration in the absence of their differential UI eligibility. In Card, Chetty, and Weber (2006), we argue that this is a plausible assumption because sample members with 20 or 30 weeks of eligibility differ only in whether they worked a little more or a little less than 36 months in the past 5 years, and thus have very similar observable characteristics.

The specification in (1) assumes that the difference in potential duration induces a permanent difference in the hazard rate (δ) and a deviation that depends on time to exhaustion. More flexible specifications – e.g. introducing separate time-to-exhaustion splines at week 20 and week 30 or implementing a regression-discontinuity estimator using a cubic control function for months worked – yield similar estimates of the spikes around benefit exhaustion.

Table 1 reports estimates of the coefficients of the time-to-exhaustion spline for variants of (1). These represent the ratio of the hazard rate in time-to-exhaustion interval j relative to the rate in the first eight weeks of the spell (the omitted time-to-exhaustion interval).[4] In specifications 1 and 2, we estimate the model using unemployment duration as the measure of spell length, defining the failure event as exiting the unemployment registry. In specifications 3 and 4, we measure duration by time to next job, defining the failure event as starting a new job. Specifications 1 and 3 include no controls, while specifications 2 and 4 include the following

[4]More precisely, the weeks that are more than 12 weeks before benefit exhaustion are omitted from $f(T-t)$. Hence, the baseline hazards correspond to the hazard rates in these weeks, and the hazard ratios that are reported are relative to this baseline.

covariates: age and its square, log wage and its square, gender, "blue collar" status, Austrian nationality, region dummies, industry dummies, and prior firm size.

The estimates from specifications 1 and 2 show that the unemployment exit hazard is approximately 2.4 times higher in the week of benefit exhaustion than in the reference period, and remains elevated for the next 8 weeks. In contrast, as shown in columns 3 and 4, the job-finding hazard is only 1.15 times larger in the week of exhaustion than the reference period, and remains elevated for only 2 weeks post-exhaustion. Despite the differences post-exhaustion, the unemployment-exit and job-finding hazards track each other closely prior to exhaustion. Neither hazard shows an increase just prior to exhaustion, and both imply that eligibility for 30 weeks of UI benefits reduces the average hazard by approximately 6%.

A substantial fraction of Austrian unemployment is attributable to workers in seasonal industries who return each year to the same employer (Emilia Del Bono and Andrea Weber 2006). Indeed, 20% of the unemployment spells in our sample end in recall. We have estimated a set of independent competing risk models for the time to the next job that distinguish between recalls and new job starts. The estimates (available in the working paper version) show that the hazard of recall rises by a slightly larger amount and for a longer period after benefit exhaustion than the hazard of new job starts, consistent with Katz (1985) and Katz and Meyer (1990b).

How quantitatively important is the spike in the re-employment rate at benefit exhaustion? To answer this question, we use our estimated hazard models to calculate the excess fraction of spells of joblessness that end at or within the month after benefit exhaustion, relative to a counterfactual in which there is no spike in re-employment rates during this interval. Note that about 80% of jobless spells end prior to benefit exhaustion, and that the average re-employment rate at and just after exhaustion is roughly 4%. Hence, a 20% higher re-employment rate at exhaustion and in the following 4 weeks implies that an extra 0.8% of spells of joblessness end at or just after the expiration of UI benefits than would do so in the absence of the spike.[5] This calculation suggests that the re-timing of job starts to coincide with benefit exhaustion is quantitatively a less important behavioral response to the provision of UI benefits than the smooth reduction in job-finding hazards that occurs throughout the spell.

[5] A similar calculation based on the estimates reported by Katz and Meyer (1990b) shows that an extra 3.5% of jobless spells end precisely at the date of UI exhaustion in their sample.

IV Conclusions

Our analysis of job losers in Austria and meta-analysis of the existing literature both indicate that re-employment hazards rise much less than unemployment exit hazards when benefits expire. This is because many individuals leave the unemployment system when their benefits expire without returning to work.

While our findings for Austria are broadly consistent with existing studies of other countries, we caution that the magnitude of any spike in the re-employment rate depends on institutional factors and labor market conditions that may differ across countries or over time. Some important factors include the availability of post-exhaustion benefits (Michele Pellizzari, forth.), the participation of UI recipients in the uncovered sector (Vodopivec, 1995), and the incentives for firms to cycle workers through temporary unemployment (Feldstein, 1976; Katz, 1985). We conclude that the size of the spike in re-employment rates at exhaustion in the current U.S. labor market remains an open question. Further work on estimating re-employment hazards using administrative measures of time to next job would be valuable.

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Table 1

Cox Hazard Model Estimates of Spike at Exhaustion

	1	2	3	4
	unemp dur		time to next job	
	no cntrls	w/cntrls	no cntrls	w/cntrls
9-12 weeks before exhaustion	0.978	0.970	0.999	1.002
	(0.019)	(0.020)	(0.021)	(0.022)
5-8 wks before	1.018	1.018	1.010	1.014
	(0.022)	(0.023)	(0.023)	(0.024)
3-4 wks before	1.054	1.070	1.051	1.042
	(0.033)	(0.035)	(0.035)	(0.036)
1-2 wks before	1.045	1.042	1.007	1.027
	(0.038)	(0.040)	(0.039)	(0.042)
0: at exhaustion	**2.404**	**2.373**	**1.150**	**1.148**
	(0.093)	(0.097)	(0.059)	(0.062)
1-2 weeks after	1.478	1.455	1.199	1.187
	(0.056)	(0.058)	(0.048)	(0.050)
3-4 wks after	1.369	1.375	1.035	1.062
	(0.062)	(0.066)	(0.049)	(0.053)
5-8 wks after	1.226	1.224	1.009	1.012
	(0.050)	(0.053)	(0.044)	(0.046)
9-12 wks after	1.165	1.147	0.929	0.927
	(0.053)	(0.055)	(0.046)	(0.048)
>12 wks after	0.958	0.967	0.772	0.782
	(0.052)	(0.055)	(0.044)	(0.047)
eligible for 30 weeks of UI	0.932	0.954	0.923	0.950
	(0.010)	(0.011)	(0.011)	(0.012)
Num. of spells	92,969	83,209	92,969	83,209

Note: Estimates shown are hazard ratios. For exhaustion spline, estimates can be interpreted as ratio of hazard rate in a given interval relative to the baseline hazard >12 weeks before exhaustion. Standard errors in parentheses.

Figure 1a

Job Finding vs. Unemployment Exit Hazards: 20 Week UI

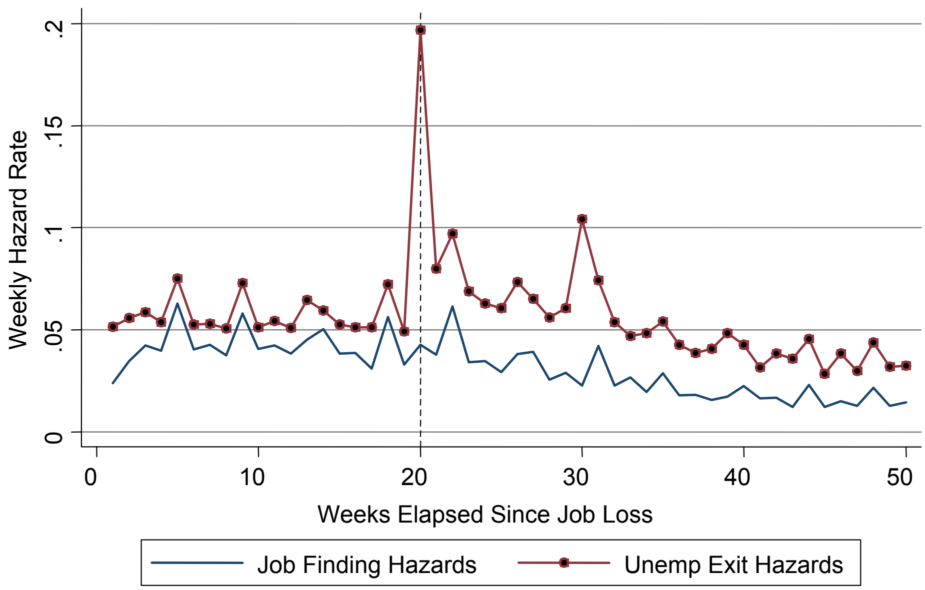

Figure 1b

Job Finding vs. Unemployment Exit Hazards: 30 Week UI

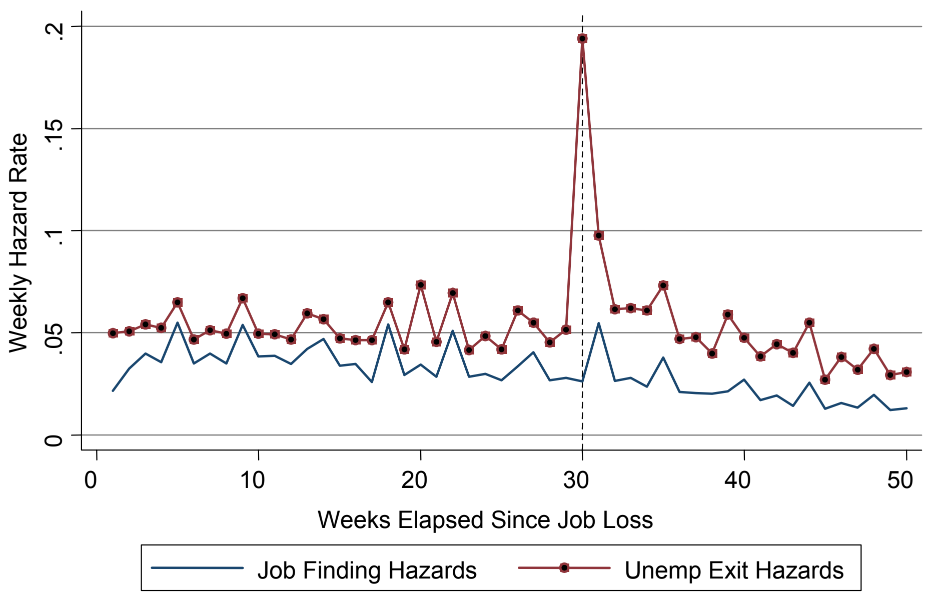

Figure 2

Benefit Expiry Effect on Hazard Rates