The Effect of Unemployment Benefit Generosity on Unemployment Duration: Quasi-experimental Evidence from Slovenia

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PRELIMINARY DRAFT

Abstract

In this paper, we analyse the effects of an increase in the generosity of unemployment benefits on the duration of unemployment in Slovenia. Using registry data on the universe of Slovenian unemployment benefit recipients, we exploit legislative changes that selectively increased the replacement rates for certain groups of workers while leaving them unchanged for others. Applying this quasi-experimental approach, we find that increasing the replacement rate significantly decreased the hazard rate for exiting from unemployment to employment. Interestingly, the implementation of these legislative changes increasing the generosity of benefits coincided with a period of increased aggregate outflows from unemployment. (JEL J64, J65)

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1 Introduction

Although the generosity of unemployment insurance (UI) benefits is widely recognized as an important determinant of unemployment spell durations, the empirical literature on UI has largely focused only on one of its parameters – the potential benefit duration (PBD). The empirical evidence on the disincentive effects of increased PBD on duration of unemployment is well established, even if the precise magnitude of its effects may vary across countries, depend on macroeconomic circumstances (e.g., Schmieder, von Wachter and Bender, 2011) or definitions of precisely what is being measured (e.g., time until employment vs. time in registered unemployment, as in Card, Chetty and Weber, 2007). On the other hand, there is a paucity of evidence on the effects of another crucial component of UI on labor market outcomes – that of the replacement rate.

In this paper, we analyze the effects of an increase in the unemployment benefit replacement rate on the probability of employment in Slovenia. We exploit legislative changes which selectively increased the replacement rate for a certain group of recipients but not others. This created conditions for a quasi-natural experiment with treatment and control groups in which a difference-in-differences approach allows for the direct effect of the change to be isolated. Another distinguishing feature of the analysis is that is based on a rich administrative database tracking the universe of unemployment benefit recipients in Slovenia.

In standard job search models, a higher replacement rate level is presumed to decrease job search intensity and increase the reservation wage, both of which have the effect of increasing the duration of unemployment (e.g. Mortensen and Pissarides, 1999). In addition, the exit rate into employment increases when benefit exhaustion approaches, leading to a spike in the hazard rate concurrent with benefit exhaustion.

While the spike at benefit exhaustion and effect of increasing the PBD have been quite extensively documented, relatively few empirical studies have focused on the effects of the replacement rate on the hazard rate for exiting unemployment.[2] Carling et al (2001) document that find that following a decrease in UI benefits in Sweden, outflow from unemployment increased considerably (with an implied elasticity of the hazard rate relative to benefits of 1.6). Rosolia and Sestito (2012) find tentative evidence that an increase in the replacement rate in Italy increased the duration of unemployment, although the effects are not statistically significant. Lalive et al (2006) find a similar result based on an analysis of Austrian claimants: increasing the replacement rate has much weaker disincentive effects than increasing the maximum duration of benefits. Finally, Uusitalo and Verho (2010) find that a 15% increase in the UI benefit level in Finland extended time until re-employment 12%.

By focusing on the effects of the increase in the replacement rate on the hazard of exiting unemployment, this paper adds to the findings of the effects of unemployment insurance in Slovenia. Early related research on Slovenia examined determinants of exit from covered unemployment (Vodopivec, 1995). In one of the rare papers to pinpoint causality, van Ours and Vodopivec (2006) find that decreasing the PBD in Slovenia in the late 1990s strongly increased the hazard rate for exiting unemployment to employment. At the same time, van Ours and Vodopivec (2008) show that the resulting employment outcomes (duration of employment, wages and type of employment) were not affected by the shortened duration of unemployment.

The main finding of this paper is that the hazard rate of exiting from unemployment to employment decreased after legislative changes were enacted that increased the generosity of the UI system in Slovenia. Interestingly, the implementation of these legislative changes that coincided with a period of *increased* aggregate outflows from

[2] See Tatsiramos and van Ours (2013) for a recent survey of the labor market effects of unemployment insurance.

unemployment. We explain this ostensible discrepancy by noting changes in the macroeconomic environment that positively affected the job-finding rate (the baseline hazard rate) and changing compositional effects in the stock of unemployed individuals.

The structure of the paper is as follows. We first discuss how legislative changes in Slovenia that were enacted at the beginning of 2011 facilitate the identification strategy, briefly discussing the characteristics of the Slovenian UI program. Next, we discuss the registry data used in the study. In the results section, we present both Kaplan-Meier survival functions and estimates from Cox proportional hazard regression analysis. The final section concludes.

2 Identification Strategy

This paper analyses exit from unemployment for UI recipients in Slovenia, which has a UI system that is broadly similar to those in other European countries. Unemployed workers qualify for benefits following involuntary termination from indefinite employment contracts or after the conclusion of fixed-term employment contracts. In the case of fixed-term contracts, workers must have been employed for 9 months in the preceding 24 months in order to qualify for benefits (the condition was somewhat more strict prior to 2011 - 12 months in the preceding 24 months). The PBD is determined by the cumulative duration of employment engagements preceding the onset of unemployment, ranging from a minimum of three months up to a maximum of 25 months. The level of monthly benefits is subject to an absolute minimum and maximum in EUR terms, and conditional on being within these absolute amounts, the exact level is determined as a fraction of previous earnings (the replacement rate).

In January 2011, a new law went into effect which increased the replacement rate in the first three months of benefit entitlement. Previously, the replacement rate was 70 percent in the first three months of employment, and the new law increased this rate to 80 percent. After the first three months, the replacement rate decreases to 60 percent, which was the case both before and after the new law went into effect.

What is the effect of the increased replacement rate on the probability of becoming employed? In order to isolate the effect, we first calculate baseline hazard rates for exit from unemployment to employement before and after the legislative changes using a control group comprised of individuals whose replacement rate remained unchanged. These are used to establish a baseline for comparison and are assumed to account for period-specific effects, such as changes in the macroeconomic environment that could affect the job-finding rate. We then calculate “before and after" hazard rates for the group for which replacement rate increased. Comparing the differences in the changes of the hazard rates between treatment and control groups before and after the law change thus yields the direct effect of the leglislative change.

Formally, we estimate a Cox proportional hazard model with the following specification:

$$\lambda(t\,|T, P, I, \mathbf{X}) \;= \lambda_0(t) \cdot e^{\alpha T + \beta P + \gamma I + \delta' \mathbf{X}}$$

where $\lambda_0(t)$ denotes the baseline hazard, T is a binary variable equal to 1 for those who became unemployed in 2011, P is a binary variable equal to 1 for those who where affected by the policy change, and I is an binary interaction variable of P and T that captures the specific effects of the policy change on the treatment group. $\mathbf{X}$ is a vector of demographic characteristics; α, β, γ and δ pertain to coefficients that are to be estimated.

A necessary condition for a valid interpretation of the difference-in-differences coefficient as the treatment effect -- the increase in the replacement rate -- is the comparability of treatment and control groups. That is, both the treatment and control groups must have identical baseline characteristics (Lee & Lemieux, 2010).

In an attempt to avoid non-random sample selection stemming from strategic behavior of unemployment benefit recipients, we exclude unemployment spells that begin immediately preceding and immediately following the change in the legislation. Although eligibility for unemployment benefits is conditional on non-voluntary termination of employment (i.e., quitters are not eligible for unemployment benefits), the patterns of inflows to unemployment in December 2010 and January 2011 are indicative of strategic behavior on the part of claimants.[3] For this reason, we exclude

[3] In addition to the change in legislation concerning the generosity of unemployment benefits, significant reforms to the pension system were due to take effect in Slovenia beginning in January 2011. While the former gave most individuals an incentive to delay their unemployment claims, individuals close to fulfilling retirement criteria had an incentive to claim benefits under the preceding

unemployment spells beginning in December 2010 or January 2011. This approach is also used by van Ours and Vodopivec (2006) in their analysis of Slovenian unemployment registry data. However, owing to a shorter sample in our case, we cannot exclude a full three months of unemployment spells preceding and following the legislative change. Moreover, our sample data may be biased due to seasonal variation in the composition of inflows to unemployment and the job-finding rate; unfortunately, we do not have a full year of data for both groups which would ameliorate such concerns.

3 Data

The data used in this study consists of registry data covering all unemployment spells from January 2010 to December 2012 in Slovenia.[4] Analysis time is measured in days, as the administrative nature of the data means every change in the unemployment status of the individual is precisely determined and monitored for compliance with eligibility criteria. For each individual unemployment spell, the register contains the following information:

– starting date of unemployment,

– reason for onset of unemployment,

– date of employment or censoring due to sickness, if applicable,

– PBD and level of unemployment benefits, if applicable,

– wage at previous job (basis for calculating unemployment benefits), if applicable,

legislation, which offered more generous conditions for retirement. In the end, the proposed pension reform was retroactively struck down in a referendum.

[4] The source of the data are two databases from the government agency tasked with administering the unemployment insurance system, the Employment Service of Slovenia (ESS). The first database, the registry of unemployed individuals, covers individuals who are registered as unemployed and who may not necessarily be receiving benefits. The second database covers unemployment benefit recipients. Individuals who are not recipients of unemployment benefits may register at the ESS because it provides job placement and training services; in addition, being listed in registry is a precondition for receiving certain categories of social assistance payments.

– personal demographic characteristics (age, education, gender, region).

The final dataset contains approximately 100 thousand unemployment spells and 30 thousand "events” (i.e., becoming employed).

Defining when an individual is at risk and when they are censored is based on an elaborate set of eligibility criteria. The onset of risk is defined as the beginning of unemployment; this does not necessarily coincide with the date of registration at the unemployment office because unemployment benefit claimants are entitled to full benefits if they apply up to one month after the onset of unemployment (the median difference between the application date and date of unemployment is 4 days). Individuals are presumed to be at risk so long as they are registered as unemployed. Censoring occurs for the duration of sick or maternity leave or during periods of declared vacation days (unemployment benefit recipients are entitled to a basic number of vacation days). For the following events, individuals are no longer under observation and are presumed to be censored (in addition, such individuals are no longer eligible for unemployment benefits): unavailability of unemployed individual to ESS staff monitoring unemployment status, not actively seeking formal employment, retirement, rejection of job offers, full-time enrollment in an educational institution, voluntary suspension of unemployment status, and emigration to a foreign country.

Furthermore, exiting to employment was defined as exiting to either full-time employment or entering public works programs. Unemployed individuals who gain part-time employment are entitled to benefits that are proportional to the number of hours that they are lacking to full-time employment, and are still considered to be at risk (i.e., unemployed).[5] Individuals' eligibility for unemployment benefits is suspended for the duration of their engagement in public works, although individuals

[5] In estimates of the hazard functions, a time-varying covariate denoting the degree of part-time employment is included in the model to allow for a potentially different hazard rate for exiting to employment.

may automatically receive unemployment benefits at the conclusion of such engagements (such instances are construed as the beginning of a new unemployment spell). Exiting to full-time employment includes becoming self-employed.

4 **Results**

In order to describe aggregate differences in job-finding rates for unemployment benefits recipients, we first turn to Kaplan-Meier survival functions for unemployment benefit recipients. Despite the increase in generosity under the new law, the survival rate for exiting unemployment to employment was slightly lower under the new law (Figure 1). After three months -- at which point the replacement rate under both laws decreases to 60 percent -- the survival rate for recipients under the less generous regime stands at around 0.8. For UB recipients under the new law, there is no discernible difference in the survival rate in the first two months of unemployment, but after the third month, the survival rate is several percentage points *lower* than the respective survival rate for the other group of UB recipients. Although the difference in the survival rates is not large, the direction of the difference points to the importance of other factors in determining exit to employment - and underscores the relevance of the difference-in-difference approach. The divergence between the survival rates becomes progressively larger in subsequent months.

Consistent with other empirical findings (e.g. Card et al 2007, van Ours and Vodopivec 2006), the survival functions exhibit relatively pronounced decreases after benefit exhaustion (Figure 2). At almost every unemployment duration, a longer PBD is monotonically associated with a higher survival rate in unemployment: longer potential benefit durations are associated with longer unemployment spells. The sole exception is for individuals with 6 month PBD, who exhibit lower survival rates than their counterparts with 3 month PBD after approximately 7 months. Paralleling the above finding of a slightly lower survival rate under the new, more generous law, the survival rates in unemployment under the new law are generally lower. The survival functions under the new law are not as smooth at longer unemployment durations due to the decreased sample size at such periods (all unemployment spells are censored in December 2012).

The results of the multivariate analysis, presented in Table 1, show that the Kaplan-Meier survival functions fail to capture a more complex underlying story. The table presents the results from a Cox proportional hazards model for the first three months of benefit receipt. The coefficient referring to the hazard ratio for UB recipients under the newer, more generous regime is greater than unity, indicating a slightly higher hazard rate of exit to employment under the new law. This result is consistent with the Kaplan-Meier survival functions indicating a lower survival rate for individuals whose onset of unemployment was in 2011 (after the legislative change). The coefficient for the treatment group refers to the relative hazard of the group whose replacement rate changed under the new law. The fact that it is not statistically significantly different from unity supports the assumption that the treatment and control groups are comparable: under the previous legislation, their hazard rate for exiting unemployment was indistinguishable from the hazard rate of the control group.

The coefficient on the interaction term, which we interpret as the pure effect of the increase in the replacement rate, is statistically significantly smaller than unity. The magnitude of the coefficient is quite large; a strict interpretation of the coefficient would suggest an approximately 18% lower hazard rate for exiting unemployment associated with the 14% increase in the replacement rate for this group (from 70% to

80%), suggesting an implied elasticity of the hazard rate with respect to benefits of 1.3.

The other coefficients in Table 1 are also of interest. Men are not observed to have a statistically significantly different hazard rate relative to women. This result stands in contrast to the findings of van Ours and Vodopivec (2006), who document higher hazard rates for exiting from unemployment for men. One possible explanation for this finding is that Slovenia's weak economic environment has had a disproportionally adverse impact on traditionally male-dominated fields such as construction. Older workers have lower hazard rates for exiting unemployment, a finding that can be explained with their generally longer PBD. Interestingly, higher levels of education are not definitively associated with increased hazard rates of exit from unemployment: the largest hazard rate is for individuals with completed technical secondary education.

5 Conclusion

Exploiting legislative changes enabling a quasi-experimental approach and using registry data on the universe of Slovenia UI recipients, we find that the 2011 increase in unemployment benefit generosity decreased the hazard rate for exiting unemployment. The estimated magnitude of the decrease is relatively large, suggesting an implied elasticity of the hazard rate with respect to benefits of 1.3.

How can we reconcile this decrease with the aggregate increase in outflows from unemployment? There are several possible explanations to explain this discrepancy. First, the increase may be attributable to a change in the composition of newly unemployed. During the summer of 2011, a large wave of bankruptcies in Slovenia, particularly amongst construction firms, flooded the labor market with many job seekers. In contrast to individual layoffs, during which firms preferably lay off underperforming or less desirable workers (Galuscak et al, 2012), such bankruptcies indiscriminately affect all workers. Second, after experiencing an extremely sharp recession in 2009, the macroeconomic outlook was relatively favorable in Slovenia during the first half of 2011 (although the situation deteriorated thereafter), which had a positive impact on labor market. Thirdly, spurned by a government eager to reduce the public deficit, the Employment Services of Slovenia increased inspections of eligibility compliance for UB receipt in 2011, conceivably accelerating the job matching process. To the extent that these factors affected all groups of unemployed workers equally, however, this would affect only the baseline hazard rate from unemployment and not our estimates of the pure effect of the legislative change.

Our results suggest that, at least within a local vicinity of the replacement rate observed in Slovenia (70-80%), decreasing benefit generosity can considerably increase the rate at which job-seekers become employed.[6] In June 2012, the Slovenian government in fact adopted another change to UI benefit legislation, considerably decreasing the benefit generosity. Our results suggest that, in addition to the direct effects, this move will reduce government expenditures also due to higher exit rates from unemployment.

[6] As Kroft and Notowidigdo (2012) point out, to the extent that optimal UI benefits should respond to shifts in labor demand, these findings may only be directly applicable under similar macroeconomic circumstances.

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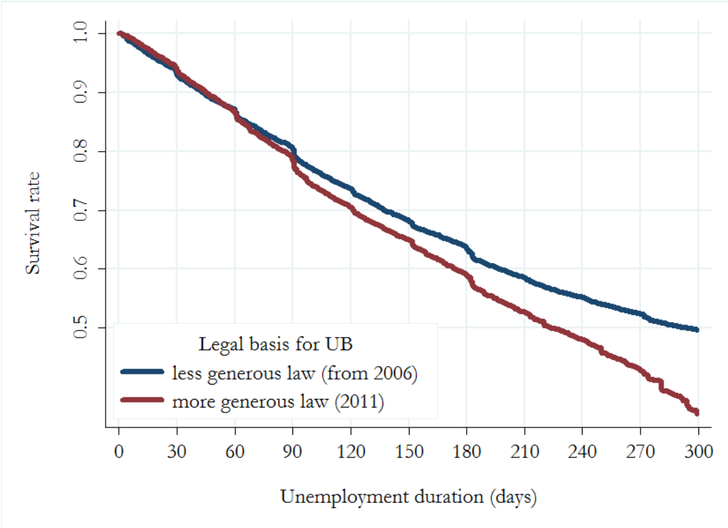
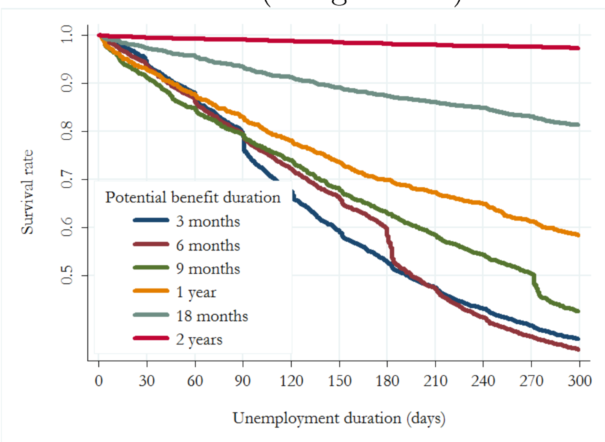
Figure 1: Kaplan-Meier Survivor Functions, Previous and New Law



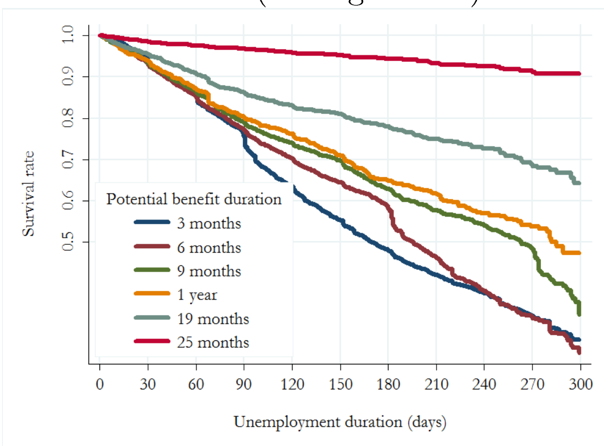
Note: Kaplan-Meier survival functions of individuals who were eligible for unemployment benefits at the onset of unemployment. Failure is defined as exiting to employment; other exits from unemployment registry database are construed as censoring. See data section for details.

Figure 2: Kaplan-Meier Survivor Functions, Previous and New Law

Old law (less generous)



New law (more generous)



Note: Kaplan-Meier survival functions of individuals who were eligible for unemployment benefits at the onset of unemployment. Failure is defined as exiting to employment; other exits from unemployment registry database are construed as censoring. See data section for details.

Table 1: Estimates from Cox proportional hazards model

	Hazard ratio	Standard Error
Policy and time-varying variables (omitted group: control group under old law)		
New (generous) law	1.081*	-0.0502
Treatment group	0.971	-0.0391
Interaction	0.824***	-0.0426
Gender (Omitted group: women)		
Men	0.972	-0.0206
Age (Omitted group: under 25 years old)		
25-29	1.022	-0.0463
30-39	0.906**	-0.0389
40-49	0.850***	-0.0372
50+	0.262***	-0.013
Education (Omitted group: Primary school or less)		
Secondary school (technical)	1.248***	-0.0387
Secondary school (general)	0.894***	-0.0297
2-year tertiary	0.936	-0.0404
4-year tertiary (or greater)	1.181***	-0.0487
Unemployment benefit (UB) recipient? (Omitted group: unemployed persons receiving UB)		
Not receiving unemployment benefits	2.921***	-0.0191
Partial UB recipient (Omitted group: No)		
Yes (part-time employed)	2.130***	-0.0508
Number of observations	87,395	

*** p<0.01, ** p<0.05, * p<0.1