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Lalive, R.; van Ours, J.C.; Zweimüller, J.

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By Rafael Lalive, Jan C. van Ours, Josef Zweimüller

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How Changes in Potential Benefit Duration Affect Equilibrium Unemployment

Rafael Lalive* Jan C. van Ours† Josef Zweimüller‡

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Abstract

This paper uses microdata to evaluate the impact of an increase in maximum benefit duration on the steady-state unemployment rate. We draw on policy changes in Austria that extended maximum benefit duration from 30 to 52 weeks for individuals above age 50 and from 30 to 39 weeks for individuals between ages 40 and 49. We use these changes to estimate the causal impact of benefit duration on labor market flows and find that (i) the policy changes lead to an increase in the steady-state unemployment rate between 20 % and 50 %; (ii) surprisingly, most of the increase is due to an increase in the inflow into unemployment, whereas the decrease in the outflow from unemployment is modest; (iii) the effects are stronger for women than for men, but are otherwise rather robust across population subgroups.

JEL Classification: C41, J64, J65

Keywords: maximum benefit duration, unemployment duration, unemployment inflow, equilibrium unemployment

*IEW, University of Zurich, IZA, CEPR, and CESifo; e-mail: rlalive@iew.unizh.ch.
†Department of Economics, Tilburg University, CentER, CEPR, and IZA; e-mail: vanours@uvt.nl.
‡IEW, University of Zurich, IZA, CEPR, and CESifo; e-mail: zweim@iew.unizh.ch.

1 Introduction

The aim of this paper is to assess how the potential duration of unemployment benefit affects the equilibrium unemployment rate. Our main contribution is the analysis of the joint effects of benefit duraton on the outflow from *and* the inflow into unemployment. This is different from the literature which has studied one of these two effects in isolation. The majority of previous studies has concentrated on the effects of the generosity of the benefit system on the probability that unemployed workers find regular jobs while a smaller literature has looked on the role of benefit rules on the probability to enter unemployment.

According to standard job search theory, more generous unemployment benefits increase the unemployment rate by reducing the search effort of workers thereby reducing the unemployment outflow rate. This prediction has been studied extensively in theoretical and empirical work and has proved to be empirically relevant and quantitatively important. The general finding from the empirical literature which will be discussed in more detail below is that it takes about 14 weeks of benefit duration to increase unemployment duration by one week.

The benefit system may affect unemployment not only via a reduced outflow from unemployment but also via a higher inflow into unemployment. One prominent argument, due to Mortensen and Pissarides (1994), holds that idiosyncratic shocks to workers' productivities let firms' optimal layoff rule depend on the wage rate – which in turn is affected by the prevailing unemployment benefit system. If the benefit system becomes more generous newly established jobs become unprofitable more quickly. As a result, a more generous benefit system will lead to an increase in the steady state flow from employment to unemployment. Alternatively, when workers' preferences (rather than their productivity) change randomly over time, a sufficiently negative shock may induce an employed worker to "quit" and collect benefits. More generous benefits will induce individuals to quit more easily rasing the inflow into unemployment. As we discuss below, and in contrast to outflow effects, empirical evidence on the effects of benefit effects on the unemployment inflow is much more scarce and far from conclusive.

Our analysis is based on a change in the Austrian unemployment insurance system that lead to a quasi-experimental situation allowing us to estimate benefit-duration effects on flows in and out of unemployment. In August 1989, the Austrian government made unemployment insurance more generous by increasing the maximum duration of unemployment benefits for certain groups of workers. Depending on age and previous work experience, the potential duration of regular benefits was raised from 30 to 52 weeks for one group, from 30 to 39 for a second group, and remained unchanged for a further group. We exploit this policy change and its differential treatment of these various groups of workers to assess the impact of benefit duration on unemployment inflows and outflows.

A particular advantage of our analysis is a very large and informative data drawn from two sources: the Austrian unemployment register and the Austrian Social Security Data (ASSD). These data sources contain the universe of all employed and unemployed Austrian workers. We observe these worker over a period of four years, two years before the policy change, i.e. from August 1987 to July 1989; and two years after this policy change, from August 1989 to July 1991. A further advantage of our study concerns the fact that the period during which the policy change took place was quite stable from a macroeconomic environment. This implies that our study is not subject to the endogenous policy which arises when more generous unemployment insurance rules are implemented in anticipation of a deteriorating labor market. Such a policy bias has been found important in several recent studies (Card and Levine, 2000, Lalive and Zweimüller, 2004a). The absence of an endogenous policy bias, the large size and the low measurement error in our data set allow us to estimate the relevant policy parameters quite precisely.

Understanding the inflow and outflow effects of the unemployment benefit system is crucial for labor market policy. First, the overall effect of a policy change remains unclear without a comprehensive understanding of both the inflow and the outflow channel. The risk is that policy makers may underestimate the implications of extended benefits for steady state labor market outcomes. Second, it is also crucial to understand the relative importance of the inflow and the outflow channel from a welfare point of view. Generous benefits that prolong unemployment spells can be problematic because long-term unemployment can cause skill depreciation. Skill depreciation is less of a concern when generous benefits mainly reduce job duration. As previous studies were typically concerned either with the inflow effect or with the outflow effect, the relative size of these two effects remains unclear. The current study aims to shed light on their relative importance. As far as we know, this is the first paper that investigates the implications of the unemployment benefit system from a comprehensive perspective.[1]

The set-up of the paper is as follows. In section 2 we review the relevant theoretical and empirical literature. Section 3 discusses the characteristics of the Austrian unemployment insurance system and briefly describes the Austrian labor market during the period when the change in maximum unemployment benefits was implemented. Section 4 presents the data we use in our analysis and discusses our empirical strategy. Section 5 presents parameter estimates and section 6 uses our estimates to simulate the implied effects for the steady-state unemployment rate. Section 7 concludes.

2 How potential benefit duration affects unemployment

2.1 Theory

Denote by $\theta_{u,t}(x|T)$ the probability that an unemployed worker with personal characteristics x finds a job in calendar time interval t when T is the maximum benefit duration (or potential benefit duration – PBD); and by $\theta_{e,t}(x|T)$ the probability that an employed worker with these characteristics looses his/her job in calendar time interval t. The steady state unemployment rate of the group of workers with characteristics x is then

$$u^*(x|T) = \frac{\theta_e(x|T)}{\theta_e(x|T) + \theta_u(x|T)}. \tag{1}$$

Consider the effects of a change in the maximum benefit duration T from the perspective of search theory. According to Mortensen (1977) expanding the duration of benefits has two opposite effects on the exit rate out of unemployment. First, the value of being unemployed increases so there is a disincentive effect that leads an unemployment worker to search less intensively. Second, the value of being employed also increases (because the value of being unemployed in the future has increased) which has a positive effect on the exit rate. For short-term unemployed the disincentive effect dominates, for unemployed near the point of benefit exhaustion (and beyond) the incentive effect dominates.

[1] There are cross-country studies that relate aggregate parameters of the unemployment insurance system - i.e. average replacement rate and average benefit duration - and other labor market institutions in various countries to the aggregate unemployment rates in these countries. See for an overview Layard and Nickell (1999).

Therefore, if there is an extension of benefit duration this will have a negative effect on the exit rate out of unemployment for short-term unemployed but it will have a positive effect on the exit rate for long-term unemployed. While first effect has been found often in empirical research, evidence for the second effect is scarce (Fredriksson and Holmlund, 2003).

The increase in the value of being unemployed through the extension of the potential benefit duration may also induce an increase in the inflow into unemployment. There are various reasons why this could be the case. For instance, the standard search and matching model with endogenous job destruction (Mortensen and Pissarides, 1994, and Pissarides, 2000, chapter 2) assumes that a worker's productivity on the job is subject to idiosyncratic shocks and firms require a minimum productivity level that prevents them from firing the worker and destroying the job. The firms' reservation productivity increases with more generous unemployment benefits, because more generous benefits push up wages requiring a higher average productivity on the job. Alternatively, assume a worker's disutility of labor (rather than his/her productivity) is subject to idiosyncratic shocks. In that case, more generous benefits will induce a worker to quit his/her job more easily. A more generous unemployment benefits may also increase the take-up of unemployment benefits. Conditional on loosing his/her job, a worker may be more inclined to apply for a benefit. Finally, it may be that the separation rate increases because a worker reduces his/her effort on the job and is more likely to be dismissed because he/she is less likely to take actions to prevent job loss.[2]

In conclusion, from a theoretical point of view, it is likely that $\partial\theta_u(x|T)/\partial T < 0$ and $\partial\theta_e(x|T)/\partial T > 0$. Therefore, an extension of the maximum benefit duration will increase the equilibrium unemployment rate:

$$\frac{\partial u^*(x|T)}{\partial T} > 0. \tag{2}$$

[2] Note, however, that according to Fredriksson and Holmlund (2003) there is not much empirical evidence in support of such an effect.

2.2 Empirical literature

Several US studies estimate the effects on the unemployment exit rate of variations in PBD that take place during recessions.[3] Early studies, including Moffitt and Nicholson (1982), Moffitt (1985), and Grossman (1989) find significantly negative incentive effects. Meyer (1990) and Katz and Meyer (1990) show that the exit rate from unemployment rises sharply just before benefits are exhausted. Such spikes are absent for nonrecipients. More recent work by Addison and Portugal (2004) confirms these findings.[4]

A common objection against these studies is policy endogeneity. Benefits are typically extended in anticipation of a worse labor market for the eligible workers. Card and Levine (2000) exploit a variation in benefit duration that occurred independently of labor market condition and show that policy bias is substantial. Lalive and Zweimüller (2004b) find similar evidence for Austria.

Evidence on the effect of PBD in European studies is mixed. Hunt (1995) finds substantial disincentive effects of extended benefit entitlement periods for Germany. Carling, Edin, Harkman and Holmlund (1996) find a big increase in the outflow from unemployment to labor market programs whereas the increase in the exit rate to employment is substantially smaller. Puhani (2000) finds that reductions in PBD in Poland did not have a significant effect on the duration of unemployment whereas Adamchik (1999) finds a strong increase in re-employment probabilities around benefit expiration. Roed and Zhang (2003) find for Norwegian unemployed that the exit rate out of unemployment increases sharply in the months just prior to benefit exhaustion where the effect is larger for females than for males. Winter-Ebmer (1998) and Lalive and Zweimüller (2004b) show that extending the potential duration of benefits had significant disincentive effects in Austria. Van Ours and Vodopivec (2006) studying PBD reductions in Slovenia find both strong effects on the exit rate out of unemployment and substantial spikes around benefit

[3] Fredriksson and Holmlund (2003) give a recent overview of empirical research related to incentives in unemployment insurance. See Green and Riddell (1993, 1997), and Ham and Rea (1987) for studies that focus on Canada.

[4] Note that there is no theoretical explanation for the existence of end-of-benefit spikes. It could be that the spikes have to do with strategic timing of the job starting date, i.e. workers have already found a job but they postpone starting to work until their benefits are close to expiration. Card and Levine (2000) point at the possibility that there is an implicit contract between the unemployed worker and his previous employer to be rehired just before benefit expire.

exhaustion.

Empirical studies on the unemployment inflow effect of a lengthening of the maximum benefit duration are more rare. Most of these studies focus on requirements concerning entrance into unemployment insurance. Christofides and McKenna (1995, 1996) for example find a clear relationship between entrance requirements of Canadian UI and employment durations. The exit rate from employment to unemployment increases substantially as soon as the workers satisfy the number of weeks worked in order to qualify for UI benefits. Anderson and Meyer (1997) investigate the take up rate of unemployment benefit insurance of workers separating from their employer. They find that both the level and the maximum duration of benefits have a significant positive effect. Green and Riddell (1997) study the effect of changes on entrance requirements on the inflow into Canadian unemployment finding that changes in entrance requirements have a significant impact on employment durations. They also find that many employment spells that just qualify under the old system are extended to just qualify under the new system. And they find that all of the response is in layoffs, not quits, which suggest that employers play an important role in the adjustment of employment durations. Green and Sargent (1998) analyze Canadian data and also find evidence of concentrations of job spell durations at the entrance requirement point and at the point at which individuals have qualified for the maximum possible weeks of UI receipts. Winter-Ebmer (2003) finds strong inflow effects of the Austrian regional extended benefit program which granted very long benefits for older wokers in certain regions.[5] These results are in line with those of Lalive and Zweimüller (2004a) who also find signficant inflow effects which were particularly strong immediately before this program was abolished.

[5] The regional extended benefit program was implemented in 1987 and ended in 1993 and was directed to a subset of Austrian regions. (See Winter-Ebmer, 1998, 2002 and Lalive and Zweimüller, 2004a, 2004b). The policy change analyzed here applies to workers in all other regions and excludes regions that were subject to the regional extended benefit program.

3 The labor market in Austria

3.1 Unemployment benefits

Like in a number of other countries the Austrian unemployment insurance system is characterized by a limited period over which unemployed individuals can draw 'regular' unemployment benefits (UB). Unemployment benefits depend on previous earnings and, compared to other European countries, the replacement ratio (UB relative to *gross* monthly earnings) is rather low. In 1990, the replacement ratio was 40.4 % for the median income earner; 48.2 % for a low-wage worker who earned half the median; and 29.6 % for a high-wage worker earning twice the median. On top, family allowances are paid. UB payments are not taxed and not means-tested. Voluntary quitters and workers discharged for misconduct cannot claim benefits until a waiting period of 4 weeks has passed. UB recipients are expected to search actively for a new job that should be within the scope of the claimant's qualifications, at least during the first months of the unemployment spell. Non-compliance with the eligibility rules is subject to benefit sanctions that can lead to the withdrawal of benefits for up to 4 weeks.

Once the period of regular unemployment benefits has expired, individuals can apply for "transfer payments for those in need".[6] As the name indicates, these transfers are means-tested and the job seeker is considered eligible only if she or he is in trouble. These payments depend on the income and wealth situation of other family members and close relatives and may, in principle, last for an indefinite time period. These transfers are granted for successive periods of 39 weeks after which eligibility requirements are recurrently checked. The post-UB transfers are lower than UB and can at most be 92 % of UB. In 1990, the median post-UB transfer payment was about 70 % of the median UB. Note however, that individuals who are eligible for such transfers may not be comparable to individuals who collect UB because not all individuals who exhaust UB pass the means test. The majority of the unemployed (59 %) received UB whereas 26 % received post-UB transfers. In sum, the Austrian unemployment insurance system is less generous than many other continental European systems and closer to the U.S. system.[7]

[6] This so-called "Notstandshilfe" implies that job seekers who do not meet UB eligibility criteria can apply at the beginning of their spell.

[7] See Nickell and Layard (1999). It is interesting to note that the incidence of long-term unemployment in Austria is closer to U.S. figures than to those of other European countries. In 1995, when our sample

Before August 1989, an unemployed person could draw regular unemployment benefits (UB) for a maximum period of 30 weeks provided that he or she had paid unemployment insurance contributions for at least 156 weeks within the last 5 years.[8] In August 1989 the potential duration of UB payments became dependent not only on previous experience but also on age at the beginning of the unemployment spell. Benefit duration for the age group 40-49 was increased to 39 weeks if the unemployed had been employed 312 weeks of employment within the last 10 years prior to the current spell. For the age group 50 and older, UB-duration was increased to 52 weeks if the unemployed had been employed for at least 468 weeks within the last 15 years.

3.2 The labor market 1987-1992

Before we go into the details of data and statistical analysis, it is instructive to briefly review the situation on the Austrian labor market during the period 1987 to 1992. This is the period on which the empirical analysis below will be concentrated. In 1987 the economy was at the end of a recession and started to improve. Real GDP growth was 1.7 % in 1987 and then started to grow to as much as 4.7 % in 1990. The favorable situation of the business cycle lead to strong employment growth throughout the period under consideration. The unemployment rate was rather stable over this period due to an increase in labor supply (immigration and rising female labor force participation).

Aggregate flows into and out of unemployment did not dramatically change during the period under consideration. The aggregate quarterly unemployment inflow rate (new unemployment spells that started in given quarter relative to the total stock of employment and out-of-labor-force) was fluctuating around 2.75 percent and the average duration of unemployment (spells completed during respective year) was roughly stable at somewhat less than 4 months. The average unemployment rate during the post-treatment period 1989-1991 was as high as during the pre-treatment period 1987-1988. Furthermore, employment growth during the treatment period was even somewhat stronger than before.

It is worth noting that this situation is favorable in terms for our empirical strategy.

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period ends, 17.4 % of the unemployment stock were spells with an elapsed duration of 12 months or more. This compares to 9.7 % for the U.S. and to 45.6 % for France, 48.3 % for Germany, and 62.7 % for Italy (OECD, 1995).

[8] UB duration was 20 weeks for job-seekers who did not meet this requirement. This paper focuses on individuals who were entitled to at least 30 weeks of benefits.

It is unlikely that our results from a comparison of the labor market experiences of older workers before and after the policy change are driven by a deteriorating labor market. As the state of the business cycle does not differ much between the two periods, an increased unemployment inflow and/or longer durations of unemployment for the subgroup of workers eligible to longer PBD can be attributed to more generous benefits.

4 Data and Empirical Strategy

To assess the impact of the above increase in benefit duration on unemployment outflow and inflow rates, we use longitudinal individual data from two different sources: (i) the *Austrian Social Security Database* (ASSD) which contains detailed information on the individuals' employment, unemployment and earnings history since the year 1972, and some information on the employer like region and industry affiliation; and (ii) the *Austrian unemployment register* from which we get information on the relevant socio-economic characteristics.

From these data we drew two samples, a "before-policy" sample and an "after-policy" sample, as follows. For both samples we selected individuals who were at least 35 years and at most 54 years old. Furthermore, we included only individuals with a continuous work history. To be included in the sample, an individual had to have a job for at least 6 out of the last 10 years *and* for at least 9 out of the last 15 years. Hence all individuals in our sample satisfy the work experience criteria for eligibility to extended benefit duration (see above). Furthermore, we excluded all individuals living in regions subject to the regional extended benefit program.[9] Furthermore, we considered only workers with previous income above ATS 12,610 (Euros 916). Workers above this income threshold experienced an increase in maximum benefit duration but no further change in the benefit rules. For workers below this threshold, the 1989 policy change also raised benefits levels. The analysis here is confined to an evaluation of maximum benefit duration on the level of unemployment rate. Restricting the analysis to higher-earnings workers avoids confounding benefit level effects with maximum benefit duration effects.

[9] The so-called "Krisenregionsregelung" accounted for about 15 % of all observations. In these crises-ridden regions even more generous unemployed insurance policies were implemented between 1988 and 1993. For empirical analyses of these programmes, see Winter-Ebmer (1998 and 2003) and Lalive and Zweimüller (2002, 2004).

For the "before-policy" sample we selected all individuals who were either (i) employed as white or blue collar workers or who were (ii) unemployed at one of the four mid-quarter baseline dates (10th February, 10th May, 10th August, 10th October) in the year 1986. We then follow these individuals up until quarter IV.1988. Similarly, the "after-policy" sample contains all individuals who were either unemployed or employed as blue- or white-collar workers at each mid-quarter baseline date in the year 1990. We then follow these individuals up until quarter IV.1992. Note that this design allows individuals in the before-policy (after-policy) sample to be *out of labor force* only in the years 1987 and 1988 (1991 and 1992). Hence this restriction reinforces our focus on attached workers. Note further, that we do not consider observations for the year 1989. This procedure minimizes potential biases resulting from anticipation effects that may arise due to behavioral changes of individuals that were unemployed under initial policy rules but were anticipating that rules will become more generous.

Table 1: *Descriptive statistics*

Table 1 compares the characteristics of the two groups. There are basically two major differences between the two groups. First, we see that after the policy change, somewhat more than a quarter (half) of the sample is eligible to additional 22 (9) weeks of potential benefits duration. While average age in the before-policy sample is only slightly younger (by 0.4 years) than the after-policy sample, the distribution across relevant age groups is more strongly affected. Second we see that the after-policy sample has a higher fraction of females.[10] Otherwise, the differences between sample are minor. Real earnings are slightly higher in the after-policy sample. Also the years of work experience within the last 15 years ("Experience") and the duration of the current job ("Tenure"; for the non-employed: tenure in the last job) is slightly higher in the after policy sample. Moreover, the number of white collar workers and the industry distributions of the two samples are very similar.

[10] The higher fraction of ages 50+ is because the big birth cohorts of 1940 - 1942 are in the age group 40-49 in the before-policy sample whereas they are in the age group 50+ in the after-policy sample. The higher fraction of females in the after-policy sample is most likely due to the fact that the cohorts that are in the after-policy but not in the before-policy sample have a high labor force participation and are relatively large (vintages in the mid 1950s). In contrast, the cohorts that are in the before-policy sample but not in the after-policy sample (vintages of the early 1930s) do have a low labor force participation and are comparably small.

Our analysis of the impact of the maximum duration of benefits on the steady-state unemployment rate is based on an analysis of individual transition probabilities to and from unemployment. To assess the effect of the maximum benefit duration on these transition probabilities we use a simple difference-in-difference estimator in the context of a probit model for quarterly transition probabilities for observation i in quarter t

$$\begin{aligned} \theta^*_{yit} &= \delta_{y1} ELIG52_{it} + \delta_{y2} ELIG39_{it} + \gamma_y A_{it} + x_{it}\beta_y + \varepsilon_{yit} \\ \theta^*_{yit} &\geq 0 \text{ if } \theta_{yit} = 1 \quad \text{and} \quad \theta^*_{yit} < 0 \text{ if } \theta_{yit} = 0 \end{aligned} \tag{3}$$

where y is a subscript indicating whether the transition concerns outflow from unemployment ($y = u$) or inflow into unemployment ($y = e$). The variables $ELIG52_{it}$ and $ELIG39_{it}$ are indicator variables that take value 1 when observation i is eligible for at most 52 or at most 39 benefit weeks, respectively.[11] Furthermore, δ_{y1} and δ_{y2} are the corresponding differences-in-differences estimators, the dummy variable A_{it} indicates the after-policy period and γ_y measures the calender time effect on transition y that is irrespective of observation i's eligibility status. Finally, x_{it} is a vector in individual characteristics, β_y is a vector of parameters that estimate the impact of these characteristics on transition y,[12] and the error term ε_{yit} capturing unobservable heterogeneity are assumed to be standard normally distributed.[13]

Obviously, whether the difference-in-difference estimator identifies the causal effect of the increase in benefit duration on the unemployment risk hinges upon whether or not the policy change was exogenous.[14] There are two reasons why policy endogeneity is most likely of minor importance in the present context. The first reason is that the economy was doing badly before the policy change (in the years 1987 and 1988). After the policy change (in the years 1989, 1990, 1991) the economy was in a boom. To the extent that all age groups were benefitting from this situation, policy endogeneity is not

[11] All observations in our samples for which both $T39_i = 0$ and $T52_i = 0$ are eligible for at most 30 weeks of benefits.

[12] The vector of individual characteristics include the individual's age, an age spline, dummies for the inflow quarter, log daily wage, experience, tenure, broad occupation (blue/white collar), sex, and industry (manufacturing, construction/tourism, other industries).

[13] The analysis below will be undertaken also for more flexible specifications of age and calendar time, and will be estimated for various subgroups to assess the robustness of the results.

[14] If policy was implemented because policy makers became concerned with worse labor market prospects for older individuals there would be policy endogeneity.

an issue. Second, one reason for the implementation of the policy may have been equity concerns. In 1988, the Austrian government implemented a very generous program that was targeted towards older steel in crises ridden steel regions. This 'Austrian regional extended benefit program' granted 4 years of unemployment benefits to eligible older workers in crisis-ridden steel regions. Hence political pressure to treat older unemployed workers in non-eligible regions more generously was one reason for changing the benefit rules. To the extent that such equity concerns were the reason for the policy change, the increase in benefit duration can be regarded as exogenous with respect to labor market outcomes of the eligible individuals in our sample.

5 Empirical estimates

We proceed in two steps. We first show the regression results of our basic statistical model for the unemployment flows, separately for the unemployment outflow and the unemployment inflow. We next check (i) whether our estimated effects of benefit duration extension are robust to at a more flexible specification of the age and calendar time variables; and (ii) how the estimated effects differ across various population subgroups. Using our outflow- and inflow-estimates, we can then discuss the question of our main interest: How do the changes in maximum benefit duration affect the steady-state unemployment rate? This is done in the next Section below.

5.1 Unemployment outflow and inflow

Table 2 show the results of equation (3) both for the unemployment outflow (column 1) and the unemployment inflow (column 2). (Notice that the coefficients displayed in the Table are probit-coefficients rather than marginal effects). The probit estimation of column 1 includes all 44,909 unemployment cases in our sample that are observed in one of the eight quarterly baseline dates. Similarly, the estimation of column 2 is based on the 1,433,259 employment cases observed in our sample.

The diff-in-diff estimators are in line with the theoretical predictions. Eligibility to longer benefits reduces the outflow rate from unemployment (column 1) and increases the inflow probability into unemployment (column 2). All coefficients have the expected sign. The effect of increasing PBD by 22 weeks ($ELIG52$) is particularly strong, both in the outflow and in the inflow equation. The effect of increasing PBD by 9 weeks ($ELIG39$) is weaker and statistically significant only in the inflow equation.

To translate the probit-coefficents of Table 2 into changes in in- and outflow probabilities, we evaluate the coefficients at the sample mean of the outflow rate (= .586) and the inflow rate (= .0175). The $ELIG52$-coefficient of $-.183$ in the outflow equation implies that the extending the maximum duration of benefits from 30 to 52 weeks decreases the quarterly outflow rate by about 7 percentage points.

Table 2: *Effect of PBD-increase on unemployment-outflow and -inflow*

We find a strong impact of extending the maximum benefit duration on the unemployment inflow. The effects are not only highly significant, they are also quantitatively substantial. Evaluated at the sample mean of the outflow rate (= .0175) the $ELIG52$ probit-coefficient of $+.143$ implies that, as a result of an increase in maximum benefit duration from 30 to 52 weeks, the monthly inflow rate increases by .50 percentage points. Similarly, the $ELIG39$-coefficient of $+.054$ implies a corresponding increase of .18 percentage points in the quarterly inflow rate.[15]

Table 2 also displays the coefficient for the control variables included in the regressions. The results indicate that age has a negative (but barely significant) impact on the probability to exit unemployment, and a significantly positive impact on the probability to enter unemployment. Apart from the continuous effect of *age* (in years), there is no significant additional impact of age for the eligibility-relevant age bracket 40-49.[16] The coefficient for "Age Spline 40" implies that the age-effect for individuals below age 40 and in the age group 40-49 is not significantly different, both in the outflow- and the inflow-regressions. However, we find that an additional year of age beyond age 50 ("Age

[15] With respect to the effect of PBD on the unemployment outflow, our results are in line with the estimates in Lalive et al. (2006) who find that the increase in PBD from 30 to 52 weeks lead to an increase in the expected duration of unemployment of 12.3 percent and who find a very small effect of the increase in PBD from 30 to 39 weeks. Our results are also similar to previous estimate to Winter-Ebmer (2003) who finds substantial effects of PBD on the unemployment inflow for a different policy change in Austria, which extended PBD for older worker in certain regions.

[16] The regression of Table 2 assumes a piecewise linear effect of age on the probit index $\beta_1 * age + \beta_2 * (age - 40) * I(age \geq 40) + \beta_2 * (age - 50) * I(age \geq 50)$, where $I(age \geq x)$ indicates whether is of age x or older. The variable Age Spline 40 is given by $(age - 40) * I(age \geq 40)$ and the variable Age Spline 50 is given by $(age - 50) * I(age \geq 50)$.

Spline 50") strongly affects the outflow from (but not the inflow into) unemployment. This may be a result of social security legislation that makes access to early retirement benefits more generous for individuals beyond age 50.

Similarly, we do not find significant of *calendar time* effects (as indicated by coefficient of the after-policy dummy), both in the outflow- and in the inflow-regression. This confirms our assumption that labor market conditions have not been worse after the policy-change than before. Furthermore, our estimates indicate substantial *seasonal fluctuations* both in inflow and outflow rates. We also find that *high-wage* workers have substantially lower turnover: not only is their risk of unemployment is lower, but high-wage workers have also lower chances to exit unemployment. Similar effects are found with respect to pervious *work experience and tenure* in the previous job. In contrast, *white-collar* workers have both a higher exit rate and a lower unemployment entry rate. Results also show significant differences between *industries* with, unsurprisingly, turnover being higher in seasonal industries (construction, tourism) and somewhat higher in manufacturing than in other industries (the reference category). Finally, we find that *women* have significantly worse chances than men to exit unemployment. Furthermore, there are significant gender differences in the risk of unemployment.

5.2 Sensitivity analysis

We now check the sensitivity of our parameter estimates. We look first for a more flexible specification of age and/or calendar time (Table 3). This could be important, as the diff-in-diff coefficient that identifies the effect of the increase in maximum benefit duration, is based on interactions with age brackets 50+ and 40-49, respectively, as well as on calendar time. Hence appropriate control for age and calendar time effects seems crucial.

Table 3: *More flexible effect of age and calendar time*

The results reported in Table 3 indicate that the estimated effects are quite robust with respect to a more flexible specifications of age and calendar time. The results reported in the first column control for the particular calendar time quarter (rather than a single after-policy dummy and seasonal quarter-dummies as in Table 2). Results in the second column are based on a regression that accounts for single-year age dummies (rather than the piecewise linear age effects captured by the age and age-spline variables in Table

2). Column 3 contains the most flexible specification with both quarterly calendar time dummies and single-year age-dummies. In all cases, the point estimates are quite similar to the one obtained in Table 2 above. Furthermore, $ELIG52$-coefficients are all highly significant and the point estimates are very close to the ones obtained in Table 2. Similarly, just like in Table 2, the $ELIG39$-coefficients are statistically significant only in the inflow- but not in the outflow-regression. Again, the point estimates are very similar. Hence we conclude that our results are rather robust with respect to the particular specification of age and calendar time.

As a further test for the robustness of our results, we look at the effects of the change in maximum benefit duration once we split our sample into various subgroups (Table 4). As a general picture, Table 4 reveals that increasing the maximum benefit duration from 30 to 52 weeks significantly affects both the outflow from (column 1) and the flow into unemployment (column 3) in most subgroups. In contrast, increasing the maximum benefit duration from 30 to 39 weeks, yields insignificant effects on the unemployment outflow in all subgroups (except for men), but still highly significant effects on the unemployment inflow for all groups.

Table 4: *Probit-results for various subgroups*

Furthermore, the results in Table 4 show that the effects are heterogenous across subgroups. For instance, labor market flows of women are more strongly affected by the increase in maximum benefit duration from 30 to 52 weeks. Interestingly, the opposite is the case for the increase from 30 to 39 weeks. The large increase affects the outflow rates of blue-collars more strongly than those of white-collars. However, the opposite is true with respect to the unemployment inflow. A similar picture emerges when we split the sample, respectively, into low- and high-wage workers and into low- and high-tenure workers. Finally, and in line with expectations, non-seasonal workers are more strongly affected by the policy changes, while only weak effects are detected for high-turnover seasonal workers.

6 Benefit duration and equilibrium unemployment

Using the parameter estimates of the inflow and outflow probabilities we consider how the maximum benefit duration affects equilibrium unemployment. Our thought experiment is the following. Let us take our estimates of the increase in PBD at face value and consider a steady-state situation in which the inflow into and the outflow from are identical. Which unemployment rate is implied by the system before the policy change as compared to the system after the change. Ignoring effects of personal characteristics x we have

$$u^*(T) = \frac{\hat{\theta}_e(T)}{\hat{\theta}_e(T) + \hat{\theta}_u(T)} \tag{4}$$

The policy changes we are analyzing are discrete, and amount to a substantial increase in maximum benefit duration for the concerned groups. In order to assess the effect of the change in benefit duration on equilibrium unemployment, we perform a comparative static analysis. If T_1 and T_2 are the maximum benefit durations before and after the policy change, the change in equilibrium unemployment equals

$$\Delta u^* = u^*(T_2) - u^*(T_1) \tag{5}$$

Furthermore, it is straightforward to decompose this change into (i) a change due to a lower outflow from unemployment, (ii) a change due to a higher inflow into unemployment, and (iii) to an interaction effect involving higher-order terms

$$\Delta u^* = \Delta u^*(out) + \Delta u^*(in) + \text{interaction effect}$$

where the inflow- and outflow-effects are given by[17]

$$\Delta u^*(out) = \frac{\hat{\theta}_e(T_1)}{\hat{\theta}_e(T_1) + \hat{\theta}_u(T_2)} - \frac{\hat{\theta}_e(T_1)}{\hat{\theta}_e(T_1) + \hat{\theta}_u(T_1)}$$

$$\Delta u^*(in) = \frac{\hat{\theta}_e(T_2)}{\hat{\theta}_e(T_2) + \hat{\theta}_u(T_1)} - \frac{\hat{\theta}_e(T_1)}{\hat{\theta}_e(T_1) + \hat{\theta}_u(T_1)}$$

We are now ready to present our simulation results that show how the more generous potential benefit duration affects the steady-state unemployment rate (Table 5). We proceed as follows. To get the effect of the benefit duration increase from 30 to 52 weeks, we utilize the entire sample (all age groups, both before- and after-policy sample). Using our regression results of Table 2 we estimate, for each observation, the inflow-

[17] Note that the interaction effect captures both changes in flows between employment and non-employment (which are likely to be negligible) and multiplicative terms of changes in unemployment-inflow and -outflow rates that arise due to the discrete nature of our decomposition analysis.

and outflow-probability with and without benefit duration extended to 52 weeks. With these estimates, we can calculate the implied steady-state unemployment rate with and without extended benefit duration. Moreover, using these estimates we can decompose the estimated increase in the steady-state unemployment rate into an inflow- and an outflow-component applying the procedure described. For the PBD extension from 30 to 39 weeks, we proceed in an analogous way.

Table 5 reports the result from this simulation exercise. In the upper panel of Table 5 we show the effects of the PBD extension from 30 to 52 weeks. The exit rate from unemployment (first column) falls from .5369 before the policy change and to .4768 after the change, implying a 11.2% reduction in the unemployment outflow probability. Similarly, the entry rate into unemployment (second column) increases from .0157 to .0209, which amounts to a 33.4% increase in the unemployment inflow probability. Taken together, these estimates imply an increase in the steady-state unemployment-population ratio from 2.84 % before the policy change to 4.20 % after the policy change. In other words, the unemployment rate increases by 1.07 percentage points or by almost 50 %.

Table 5: *Effects of PBD increase on inflow to and outflow from unemployment and on the equilibrium unemployment rate*

The lower panel of Table 5 applies the same procedure the estimate the effects of the PBD extension from 30 to 39 weeks. While qualitatively all effects go in the same direction they are quantitatively much smaller. The outflow probability decreases from .5809 to .5695, and the inflow probability increases from .0161 to .0180. These effects imply an increase in the equilibrium unemployment ratio of 0.4 percentage points (or 13.8%), from 2.7 percent before the change to 3.1 percent after the change.

Table 5 shows a further interesting result. Decomposing the increase in the unemployment ratio into an inflow- and an outflow-component reveals that the bulk of the increase is due to the larger unemployment *inflow* rate. The effect of extended PBD on the unemployment outflow is much smaller. For the PBD increase to 52 weeks, 66.7% of the increase in the unemployment ratio can be attributed to an increase in the entry rate, whereas only 25.5% is to due a lower exit rate from unemployment. For the PBD increase from 30 to 39 weeks, an even larger fraction of the increase in the unemployment ratio (83.8%) is due to the increase in the inflow-rate, whereas only 13.5% can be assigned to the lower unemployment exit probability.

A further interesting indicator shows that the increase in PBD raises the unemployment ratio by .062 percentage points *per additional PBD week* for the extension from 30 to 52 weeks; and by .041 percentage points for the extension from 30 to 39 weeks. Interestingly, for both policy changes, the estimated effect per additional PBD week attributable to the unemployment *inflow*, is similar for the short (30 to 39) and the long (30 to 52) PBD increase. The isolated effect of one additional PBD week on the unemployment inflow indicates and increase in the unemployment ratio by .041 percentage points (increase from 30 to 52 weeks) and by .034 percentage points (increase from 30 to 39 weeks). The effects on the unemployment outflow are much smaller. We find that one additional PBD-week increases the equilibrium unemployment ratio by .016 percentage points for the policy change from 30 to 52 weeks, whereas the corresponding estimate for the policy change from 30 to 52 weeks amounts to only .005 percentage points.

6.1 Simulations for subgroups

We find that the increase in maximum benefit duration increases equilibrium unemployment, to some extent because the outflow from unemployment goes down but mainly through an increase in the inflow into unemployment. To investigate whether this results also holds for subgroups we use the parameter estimates presented in Table 4 to perform similar simulations as before, but now separately for each subgroup. Table 6 presents the simulation results. The upper part presents the results for the PBD change from 30 to 52 weeks, the lower part gives the simulation results for the PBD change from 30 to 39 weeks. For reasons of comparison the first rows of each part of the table replicates the main results of Table 5.[18]

Table 6: *Effects of PBD increase on inflow to and outflow from unemployment and on the equilibrium unemployment rate; various subgroups*

As shown the PBD change from 30 to 52 weeks increases equilibrium unemployment for every subgroup with the increase for women, low wage workers and non-seasonal workers being larger than for their counterparts. There is hardly any difference between blue collar and white collar workers and between workers with low tenure and high tenure.

[18] Note that in the simulations we use all estimated parameters of Table 6 irrespective of whether or not they are significantly different from zero at conventional levels of significance.

For every subgroup the contribution to the change in equilibrium unemployment of the change in inflow is larger than that of the change in outflow.

Also for the PBD change from 30 to 39 weeks we find that the increase in equilibrium unemployment is mostly due to the increase in the inflow into unemployment and to a much smaller extent due to the decrease in the outflow from unemployment.

7 Conclusion

According to job search theory an increase in the maximum unemployment benefit duration affects the unemployment rate both through a decrease in the outflow from unemployment and through an increase in the inflow to unemployment. These theoretical predications are confirmed by empirical research. However, empirical research has been on either the outflow from unemployment or the inflow into unemployment. There are no studies that investigate both effects simultaneously. So, it is not clear to what extent effects on inflow and outflow affect the unemployment rate.

This paper uses microdata to evaluate the impact of an increase in maximum benefit duration on the steady-state unemployment rate distinguishing between these two effects. We draw on policy changes in Austria that extended maximum benefit duration from 30 to 52 (30 to 39) weeks for individuals above age 50 (between ages 40 and 49) with a continuous work history. We find that this policy change lead to 50 % increase in the steady-state unemployment rate for the older age group and a 20 % increase in the steady-state unemployment rate for the younger age group. Surprisingly, most of the increase in equilibrium unemployment is due to an increase in the inflow into unemployment, whereas the effect of the decrease in the outflow from unemployment is modest. We also find that the effects are stronger for women than for men. There may be institutional reasons this as conditional on age women are closer to (early) retirement, and it is in line with the general notion the women react more strongly to incentives – wage elasticities of labor supply are larger for women than for men. Otherwise our results are rather robust across population subgroups.

Our results show that the PBD extension had much bigger effects on the behavior of individual unemployed workers than it had on individual employed workers. The PBD extension made it a lot more attractive for unemployed workers to reduce the search activities and thus lower their job finding rate. The PBD extension made it only a little bit more attractive for employed workers to become unemployed. Nevertheless, since there are many more employed workers that are affected the aggregate unemployment inflow effect is larger than the aggregate unemployment outflow effect. It is the sheer mass of employed workers that cause the inflow effect to be larger.

From a policy point of view it is important to know that the inflow effect is larger than the outflow effect. Should this not be taken into account the effects of a change in PBD will be seriously underestimated. The fact that changes in PBD have quite a large – aggregate – inflow effect also means that PBD could be an instrument to increase the employment rate. If the PBD is shortened, firms will become more reluctant to destroy jobs and / or it will be less attractive for workers to “quit" into unemployment. We also note that our results are partly based on older (50+) workers which have low employment rates in many countries. Taking into account inflow effects for these groups seems highly relevant from a policy perspective.

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Table 1:
Descriptive statistics

	Before-policy sample		After-policy sample	
	mean	std.dev.	mean	std.dev.
ELIG52	0	0	0.310	0.463
ELIG39	0	0	0.537	0.499
Age 50+	0.249	0.432	0.310	0.463
Age 40-49	0.576	0.494	0.537	0.499
Age	45.473	5.477	45.952	5.436
After-policy	0	0	1	0
1st quarter	0.252	0.434	0.251	0.434
3rd quarter	0.249	0.432	0.249	0.433
4th quarter	0.248	0.432	0.248	0.432
log(wage)	6.559	0.279	6.585	0.300
Experience[a] (in years)	14.192	1.306	14.138	1.370
Tenure (in years)	9.566	5.824	10.341	7.137
White collar	0.541	0.498	0.550	0.497
Manufacturing	0.357	0.479	0.332	0.471
Construction, toursim	0.139	0.346	0.136	0.342
Women	0.310	0.462	0.358	0.480

a) Work experience during last 15 years

Table 2:
Probit results on unemployment flows

dependent variable:	unemployment outflow	unemployment inflow
mean of dep. variable	0.586	0.0175
ELIG52	-0.183	0.143
	(4.96)**	(8.77)**
ELIG39	-0.035	0.054
	(1.04)	(3.78)**
Age	-0.017	0.011
	(1.64)	(2.26)*
Age Spline 40	-0.003	-0.003
	(0.22)	(0.64)
Age Spline 50	0.047	0.000
	(7.13)**	(0.07)
After policy	-0.046	-0.014
	(1.51)	(1.03)
1st quarter	0.513	0.159
	(28.97)**	(16.32)**
3rd quarter	-0.109	0.154
	(5.34)**	(16.17)**
4th quarter	-0.199	0.644
	(10.32)**	(76.20)**
log(wage)	-0.191	-0.169
	(7.95)**	(13.45)**
Experience[a] (in years)	-0.015	-0.128
	(3.27)**	(67.96)**
Tenure (years)	-0.028	-0.056
	(20.68)**	(91.88)**
White collar	0.155	-0.163
	(7.67)**	(21.53)**
Manufacturing	-0.014	0.085
	(0.88)	(11.54)**
Construction, tourism	0.719	0.442
	(43.43)**	(58.18)**
Women	-0.127	0.016
	(8.61)**	(2.17)*
Constant	2.229	0.279
	(5.08)**	(1.39)
Observations	44,909	1,433,259

Note: Probit coefficients (not marginal effects), absolute value of z statistics in parentheses
* significant at 5%; ** significant at 1%

a) Work experience during last 15 years

Table 3:
PBD-coefficients: Allowing for more flexible effects of age and calendar time

unemployment outflow			
ELIG52	-0.176	-0.193	-0.180
	(4.75)**	(4.80)**	(4.48)**
ELIG39	-0.035	-0.041	-0.035
	(1.04)	(1.12)	(0.95)
unemployment inflow			
ELIG52	0.137	0.143	0.135
	(8.36)**	(8.01)**	(7.53)**
ELIG39	0.050	0.067	0.061
	(3.49)**	(4.22)**	(3.80)**
Calendar time dummys (for each quarter)	yes	no	yes
Age dummys (for each year)	no	yes	yes

Note: Probit coefficients (not marginal effects), absolute value of z statistics in parentheses
* significant at 5%; ** significant at 1%

Table 4:
Diff-in-diff Probit coefficient of PBD-effects: various subgroups

	unemployment outflow		unemployment inflow	
	ELIG 52	ELIG 39	ELIG 52	ELIG 39
Whole sample	-0.183	-0.035	0.143	0.054
	(4.96)**	(1.04)	(8.77)**	(3.78)**
Women	-0.248	-0.019	0.164	0.003
	(3.84)**	(0.32)	(5.61)**	(0.10)
Men	0.000	0.079	0.120	0.089
	(0.00)	(1.96)	(5.87)**	(4.97)**
Blue collar	-0.146	-0.028	0.139	0.071
	(2.68)**	(0.58)	(5.99)**	(3.51)**
White collar	-0.005	0.083	0.212	0.078
	(0.07)	(1.46)	(7.65)**	(3.17)**
Low wage	-0.160	0.032	0.145	0.022
	(3.39)**	(0.74)	(6.60)**	(1.14)
High wage	-0.052	-0.002	0.138	0.063
	(0.97)	(0.05)	(5.95)**	(3.10)**
Low tenure	-0.123	0.003	0.100	0.042
	(3.02)**	(0.08)	(5.27)**	(2.62)**
High tenure	-0.160	0.009	0.157	0.052
	(1.94)	(0.11)	(4.92)**	(1.75)
Seasonal	-0.073	0.086	0.022	-0.038
	(1.06)	(1.40)	-0.7000	(1.38)
Non-seasonal	-0.091	0.041	0.192	0.087
	(2.11)*	(1.05)	(9.84)**	(5.01)**

Note: Probit coefficients (not marginal effects), absolute value of z statistics in parentheses
* significant at 5%; ** significant at 1%

Table 5:
Effects of PBD increase in inflow, outflow and unemployment population ratio

	Quarterly outflow	Quarterly inflow	Interaction	Implied steady-sate unemployment ratio (%)
PBD change 30 to 52 weeks				
Before policy-change	.5369	.0157		2.84
After policy-change	.4768	.0209		4.20
Implied increase in u* (p.p.)	.35	.91	.11	1.37
(percentage due to ...)	(25.5%)	(66.7%)	(7.9%)	(100.0%)
Implied increase in u* per additional PBD week (p.p.)	.016	.041	.005	.062
PBD change 30 to 39 weeks				
Before policy-change	.5809	.0161		2.70
After policy-change	.5695	.0180		3.07
Implied increase in u* (p.p.)	.05	.31	.01	.37
(percentage due to ...)	(13.5%)	(83.8%)	(2.7%)	(100.0%)
Implied increase in u* per additional PBD week	.005	.034	.001	.041

Table 6:
Decomposing the increase in the unemployment population ratio, various subsamples

Subsample	u	change in u	due to outflow		due to inflow		due to interaction	
			absolut	%	absolut	%	absolut	%
BPD 30 to 52								
full sample	.0284	.0137	.0035	25.4%	.0091	66.8%	.0011	7.8%
females	.0305	.0231	.0073	31.8%	.0128	55.6%	.0029	12.6%
males	.0268	.0066	.0000	0.0%	.0066	100.0%	.0000	0.0%
blue collar	.0286	.0127	.0027	20.9%	.0092	72.6%	.0008	6.5%
white collar	.0195	.0133	.0001	0.7%	.0131	98.9%	.0001	0.5%
low wage	.0418	.0178	.0041	23.0%	.0126	70.5%	.0012	6.5%
high wage	.0204	.0077	.0007	8.9%	.0068	88.2%	.0002	2.9%
low tenure	.0506	.0134	.0033	24.2%	.0096	71.5%	.0006	4.3%
high tenure	.0153	.0116	.0027	23.1%	.0076	65.8%	.0013	11.2%
seasonal industries	.0668	.0036	.0015	42.7%	.0020	56.1%	.0000	1.2%
other industries	.0220	.0147	.0018	12.0%	.0120	81.7%	.0009	6.3%
BPD 30 to 39								
full sample	.0270	.0037	.0005	14.2%	.0031	84.2%	.0001	1.6%
females	.0269	.0005	.0004	69.6%	.0002	30.0%	.0000	0.4%
males	.0263	.0037	-.0009	-25.1%	.0047	129.5%	-.0002	-4.4%
blue collar	.0287	.0049	.0004	8.2%	.0044	90.6%	.0001	1.2%
white collar	.0182	.0024	-.0013	-52.6%	.0039	163.6%	-.0003	-11.0%
low wage	.0408	.0011	-.0007	-61.2%	.0018	163.7%	.0000	-2.5%
high wage	.0193	.0028	.0000	0.8%	.0028	99.0%	.0000	0.1%
low tenure	.0452	.0036	-.0001	-1.8%	.0036	102.0%	.0000	-0.1%
high tenure	.0118	.0017	-.0001	-5.5%	.0018	106.3%	.0000	-0.8%
seasonal industries	.0689	-.0050	-.0016	31.8%	-.0035	69.7%	.0001	-1.5%
other industries	.0209	.0039	-.0006	-15.6%	.0046	119.0%	-.0001	-3.4%