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How Changes in Financial Incentives Affect the Duration of Unemployment

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This paper studies how changes in the two key parameters of unemployment insurance—the benefit replacement rate (RR) and the potential benefit duration (PBD)—affect the duration of unemployment. To identify such an effect we exploit a policy change introduced in 1989 by the Austrian government, which affected various unemployed workers differently: a first group experienced an increase in RR; a second group experienced an extension of PBD; a third group experienced both a higher RR and a longer PBD; and a fourth group experienced no change in the policy parameters. We find that unemployed workers react to the disincentives by an increase in unemployment duration, and our empirical results are consistent with the predictions of job search theory. We use our parameter estimates to split up the total costs to unemployment insurance funds into costs due to changes in the unemployment insurance system with unchanged behaviour and costs due to behavioural responses of unemployed workers. Our results indicate that costs due to behavioural responses are substantial.

1. INTRODUCTION

This paper studies how changes that make unemployment systems more generous affect the duration of unemployment. Most unemployment insurance systems are characterized by two major parameters: the earnings replacement rate (replacement rate RR), that is the level of unemployment benefits in relation to expected earnings; and the maximum duration that an unemployed worker can draw such benefits (potential benefit duration, PBD). A considerable theoretical literature has shown that a more generous unemployment insurance system (through a higher RR and/or a longer PBD) will reduce the optimal job search effort of an unemployed worker and hence result in longer unemployment duration. Moreover, these theories also offer sharp predictions on how changes in the key parameters in unemployment insurance will affect the unemployment exit rate. In particular, an increase in RR should have its strongest effect early in the unemployment spell, whereas an increase in PBD will have its strongest effect around the time when benefits expire, creating “spikes” in the unemployment exit rate.

The present paper identifies the causal effect of benefit duration on the willingness of individuals to accept jobs using a policy change that took place in Austria in 1989. The policy affected various unemployed workers differently: a first group experienced an increase in RR; a second group experienced an extension of PBD; a third group experienced both a higher RR and a longer PBD; and a fourth group experienced no change in the policy parameters. More precisely, unemployment benefits (and hence RR) were increased by about 15% for workers earning below a certain threshold whereas for workers above this threshold the RR remained unchanged. The size of the increase in PBD depended on age and experience: For workers below age 40 and/or for workers with little previous work experience PBD remained unchanged. For workers with high previous work experience PBD increased, respectively, from 30 to 39 weeks for the age group 40–49; and from 30 to 52 weeks for workers aged 50 and older. Hence, this policy change provides a nice empirical design, because it has elements of a "natural" experiment.

Furthermore, the policy change took place in 1989 which, in Austria, was quite a stable macroeconomic environment. This implies that our study is less subject to endogenous policy bias than other studies. Endogenous policy bias arises when more generous unemployment insurance rules are implemented in anticipation of a deteriorating labour market. Such a policy bias has been found important in several recent studies (Card and Levine, 2000; Lalive and Zweimüller, 2004*a*). To assess the effect of these changes in unemployment insurance rules, we use a large and informative data-set that allows us to trace workers' unemployment histories over an extended period of time. We compare the entire unemployment inflow that took place two years before the policy change to the entire inflow two years after the change. This leaves us with a rather large data-set and allows us to estimate the interesting policy parameters quite precisely.

We find that both the increase in RR as well as the extension of PBD significantly increases the duration of unemployment. In line with theoretical predictions, we find that most of the effect takes place early in the unemployment spell in the case of the RR increase and around the dates when benefits expired in the case of the PBD extension. Furthermore, our sensitivity analysis shows that, despite the clear heterogeneity in treated populations, results are rather robust. In particular, we find that the effect of the RR increase, the effect of PBD extensions, and the additive effect of simultaneous RR and PBD changes are independent of variations in the control groups used to identify the treatment effect. While the size of the effects are quite robust, there are two important exceptions. First, older workers react more strongly to PBD extensions than prime-age workers, and second, the additive effect of a simultaneous change in RR and PBD is larger for older workers than for prime-age workers. In a theoretical context, this stronger reaction of older workers to changes in unemployment insurance rules can be rationalized by their weaker labour-market position and/or by incentives created by the institutional environment (early retirement).

Reliable empirical evidence on how the two key unemployment insurance parameters affect the search behaviour of unemployed workers is crucial for designing appropriate policies. Most previous empirical studies identify such incentive effects from "exogenous" variation, separately, either in a change in benefit levels or a change in the PBD. While the *ceteris paribus* effects of changes in such parameters are clearly interesting *per se*, it does not allow to address potentially important questions of the appropriate design of unemployment insurance. To assess the relative importance of the two policy parameters we use our parameter estimates to split up the total costs to unemployment insurance funds into costs due to changes in the unemployment insurance system with unchanged behaviour and costs due to behavioural responses of unemployed workers. Our results indicate that costs due to behavioural responses are modest for increases in RR, but substantial for increases in PBD. From these results, a simple policy conclusion can be derived: If policy makers want to influence incentives, the potential duration of unemployment is a more effective tool than the level of unemployment benefits.

The paper is organized as follows. The following section discusses the relevant theoretical arguments and also provides a discussion of previous empirical evidence regarding the effects of the two key parameters of unemployment insurance. Section 3 gives the relevant institutional background on the changes to unemployment insurance in Austria that are used to identify their effects on unemployment durations. Section 4 provides a first descriptive analysis of the effects of unemployment insurance on unemployment duration. Section 5 presents the econometric analysis regarding the effects on the unemployment exit rate, and Section 6 concludes with a summary of the results and draws some policy conclusions.

2. THEORY AND PREVIOUS STUDIES

2.1. *Theory*

We assume that an unemployed worker is entitled to unemployment benefits for a fixed duration. After benefits expire he or she is entitled to unemployment assistance of infinite duration that is lower than unemployment benefits. We are interested to see how job search behaviour is affected by the presence of a fixed benefit duration.[1] The arguments in the optimization problem are expected costs and benefits, that is, the value of being unemployed compared to the value of having a job. The value of unemployment is determined by the level of the unemployment benefits, the search costs, the situation in the labour market (*i.e.* the way search intensity translates into job offers), the expected gain from accepting a job and the risk of not finding a job before unemployment benefits expire. While an increase in search intensity increases search costs it also increases the probability of finding a job and reduces the probability of running out of unemployment benefits. Optimal search intensity balances marginal costs and benefits of search.

At the start of the unemployment spell search intensity is low because the probability of finding a job before unemployment benefits expire is large anyway. With increasing duration of unemployment the risk of running out of benefits increases, which induces an unemployed worker to increase search intensity. Once an unemployed worker runs out of unemployment benefits, search intensity remains constant because the worker faces a stationary environment. The exit rate from unemployment is then determined by search costs, the level of unemployment assistance, and the situation in the labour market.[2]

Let us now consider an increase in the RR. A higher RR implies that, at the start of the unemployment spell, a worker will search less intensively than before, as the costs of being unemployed are lower. As benefit expiration comes closer, the risk of running out of the now higher benefits induces the unemployed worker to increase search intensity more strongly. Search intensity eventually becomes even larger than before due to an "entitlement effect". The value of a new job increases because losing this job is associated with less damage due to the higher RR. These predictions are summarized in the upper panel in Figure 1, which draws the unemployment exit hazard against the elapsed duration of unemployment for different RR (dashed for the higher RR).[3] The ratio between the exit rate in the new and the old system is initially less than unity, will

1. See, for example, Mortensen (1977, 1990), Burdett (1979), and Van den Berg (1990). The unemployed could also choose both an optimal search intensity and an optimal reservation wage, but this would not essentially change behaviour and is therefore ignored here. Note that, with a fixed reservation wage, there is a one to one relationship between search intensity and job-finding rate.

2. Mortensen (1977) argues that search intensity shifts down after benefit expiration if non-market time and market goods used in household production are substitutes, or it goes up if time and market goods are complements in household production.

3. This figure is based on a discrete time model in which job seekers receive unemployment benefits equal to 40% of their previous wage for six successive (five-week) periods and no benefits thereafter. Job seekers choose how intensively to search for jobs. The unemployment exit hazard is the product of the search intensity and the arrival rate of job offers, taken to be equal to 0·25 per period. The RR change increases the RR from 40% to 45%, whereas the benefit duration change increases PBD from 6 to 8 periods. Note that the increase in PBD with 30% and the increase in RR with about 10% mimic the Austrian changes in the UI system.

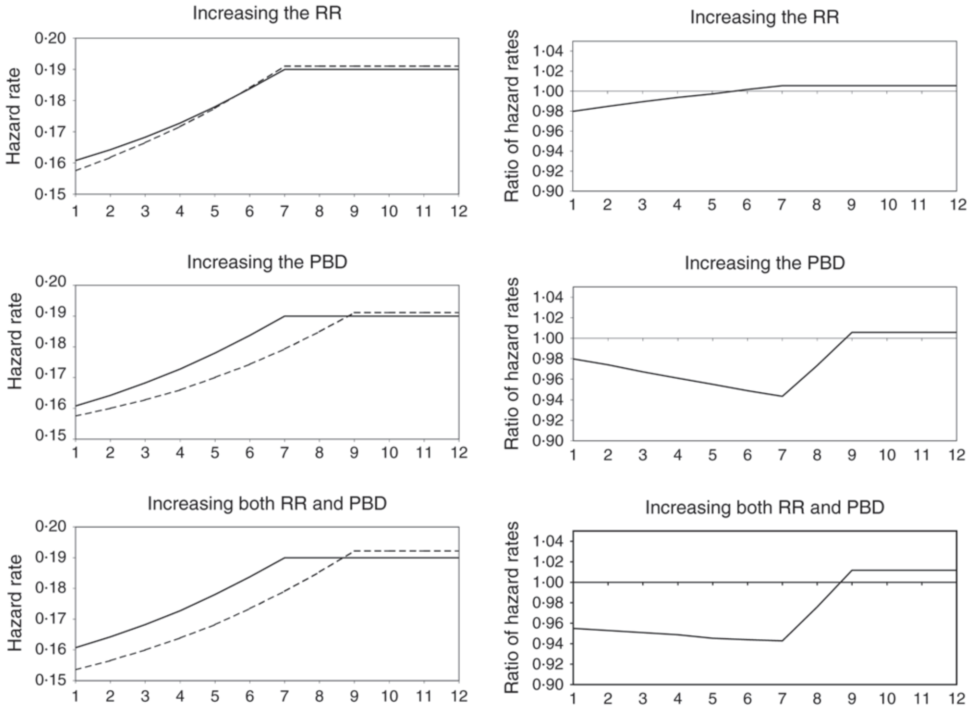


FIGURE 1

Change to replacement rate (RR) and potential benefits duration (PBD) and the unemployment exit hazard

increase with increasing unemployment duration, and will eventually exceed unity when benefits expire.[4]

In contrast, an extension of PBD entails only small immediate disincentive effects. The reason is that, at the beginning of the spell, the extension of PBD does not strongly affect the risk of running out of benefits. However, with increasing unemployment duration a longer PBD makes a difference. When unemployment duration is equal to PBD under the old system, the difference in search intensities of the two systems is largest. At that date, search intensity is at its maximum under the old system, while it is still comparably low under the new system. As unemployment duration comes closer to PBD under the new system, search intensity increases strongly. Eventually, it becomes even larger than under the old system, again due to the entitlement effect. As illustrated in the middle panel of Figure 1, most of the action will take place just before benefit exhaustion in the old system until benefit exhaustion in the new system. The ratio between the exit rate in the new and the old system will not only be less than unity, but also will be decreasing until benefits expire in the old system. Afterwards, the ratio increases strongly and eventually exceeds unity when benefits expire under the new system.

Finally, let us consider a simultaneous increase in RR and PBD. Now, an unemployed worker is facing not only a higher benefit over his initial PBD and a longer PBD with his initial benefits but he or she is also entitled to higher benefits over a longer PBD. Suppose PBD is

4. Note that the magnitude of the entitlement effect depends on the expected duration of employment compared to the expected duration of unemployment. If employment durations are long compared to unemployment durations, the entitlement effect is bound to be small.

increased from E_0 to E_1 and the RR increase raises benefits from b_0 to b_1. To get the intuition of how these changes affect incentives consider a worker who exhausts all his benefit claims. If this worker is eligible to the RR increase only, total benefit payments increase by $\Delta b E_0$. For a worker eligible to the PBD increase only, total benefit payments increase by $b_0 \Delta E$. A worker who is eligible to both an increase in RR and an increase in PBD, total benefits increase by $\Delta b E_0 + b_0 \Delta E + \Delta b \Delta E$. The latter interaction effect causes an additional disincentive effect and cannot be ignored when changes in b and E are substantial. Therefore, we expect that the disincentive effects of a simultaneous change in RR and PBD will cause an additional decline of the job-finding rate that is larger than the sum of two separate changes. Relative exit rates from a joint change are lower until benefits have expired in the old system increase more strongly thereafter and reach a higher level when benefits have expired in the new system.

The above discussion has obvious empirical implications for transitions from unemployment to employment. First, provided that entitlement effects are negligible and/or most unemployment exits take place before benefits expire, increases in RR and extensions in PBD should lead to a reduction in job search effort and hence to longer unemployment durations. Second, increases of RR should trigger strong behavioural responses early in the unemployment spell whereas extensions of PBD should lead to strong responses around the dates when benefits expire. Third, for individuals entitled to a simultaneous increase in RR and PBD, we expect an increase in unemployment durations larger than the sum of the increases from two isolated changes in the policy parameters.

Clearly, the exact size and nature of the relationship between the unemployment insurance parameters and the unemployment duration depends on the parameters of the model. In particular, differences in labour-market position and the institutional setting may explain why one group of workers reacts more strongly than another. For instance, older workers receive fewer job offers because employers often prefer to hire young or prime-age workers to fill their vacancies. Moreover, older workers are closer to (regular) retirement and/or have a greater chance to take up early retirement benefits. This implies that older workers benefit less from a given search effort and may react stronger to disincentives. This may be particularly relevant for large extensions of PBD. In that case, the consequences of not finding a job are less severe because the period between benefit expiration and the start of (early) retirement may be reduced substantially.

2.2. *Previous empirical studies*

Several U.S. studies estimate the effects on the exit rate from unemployment of variations in PBD that take place during recessions.[5] Early studies, including Moffitt and Nicholson (1982), Moffitt (1985), and Grossman (1989) find significantly negative incentive effects. Katz and Meyer (1990) and Meyer (1990) show that the exit rate from unemployment rises sharply just before benefits are exhausted. Such spikes are absent for non-recipients. More recent work by Addison and Portugal (2004) confirms these findings.[6]

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

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

6. 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, that is, 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 benefits expire.

the Austrian labour market. 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 labour-market programmes whereas the increase in the exit rate to employment is substantially smaller. Winter-Ebmer (1998) uses Austrian data and finds significant benefit duration effects for males, but not for females. Rød 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. 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. 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 exhaustion.

To estimate the effects of RR on unemployment durations, the early literature focuses on differences in benefit replacement ratios between individuals. Estimated elasticities of unemployment duration with respect to benefit levels range from 0·1 to 1·0 (Atkinson and Micklewright, 1991). This literature is problematic due to the possibility of unobserved heterogeneity distorting identification in cross-sectional data. More recent European studies have focused on the impact of policy changes affecting benefit levels. Carling, Holmlund and Vejsiu (2001), studying the effects of a reduction in the RR from 80% to 75% in Sweden in 1995, find that this policy change increased the job-finding rate by roughly 10%, implying an elasticity of around 1·7. Bennmarker, Carling and Holmlund (2004), studying changes in the Swedish system in the early 1990’s, find a smaller elasticity of around 0·6. Rød and Zhang (2003) estimate elasticities for Norway of around 0·95 for males and around 0·35 for females.

3. UNEMPLOYMENT INSURANCE AND THE AUSTRIAN LABOUR MARKET

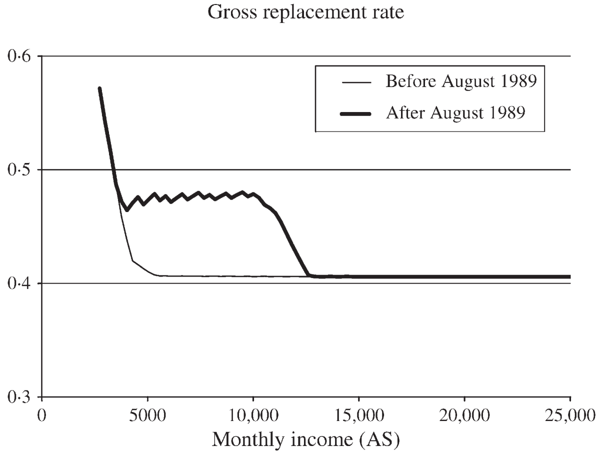
3.1. *The system before the policy change*

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). UB depend on previous earnings and, compared to other European countries, the replacement ratio (UB relative to *gross* monthly earnings) is rather low. Figure 2 shows that, before August 1989, the replacement ratio declined strongly from a maximum of about 63% (monthly income is below 2210 ATS) to 41% in the income range of between 3000 and 5000 ATS previous monthly earnings.[7] The benefit replacement ratio then stays just below 41% for incomes up to the cap of 27,430 ATS previous monthly earnings. Individuals earning more than 27,430 ATS get UB of 11,233 ATS per month irrespective of their income. Thus, the benefit RR decreases monotonically in previous monthly income for the high-income group.

On top of UB, family allowances are paid. UB payments are not taxed and not means tested. Voluntary quitters and workers discharged for misconduct cannot claim benefits during the first four weeks of a new unemployment spell. 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 four weeks.

Before August 1989, an unemployed person could draw regular UB for a maximum period of 20 weeks provided that he or she had paid unemployment insurance contributions for at least 52 weeks within the last two years. Individuals who had contributed to unemployment insurance (UI)

7. The median monthly income was about 16,400 ATS in the unemployment inflow from regular jobs in 1988.



Source: Austrian federal laws (Bundesgessetzblätter) no. 594/1983, 364/1989.

FIGURE 2

The change to the RR in August 1989

for at least 156 weeks in the last five years were eligible for 30 weeks of regular unemployment benefits.[8]

Once the period of regular unemployment benefits has expired, individuals can apply for "transfer payments for those in need" ("Notstandshilfe").[9] As the name indicates, these transfers are means tested, and the job seeker is considered eligible only if she or he faces economic difficulties. 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. These post-UB transfers are lower than UB and can be at most 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 system described above applies to all of Austria except for regions with a dominant steel industry. In these steel regions, a regional extended benefit programme was introduced in June 1988, that entails a dramatically different UI system for workers aged 50 or older. In order to focus attention on the policy change in August 1989, we focus on the non-steel regions that were never entitled to the regional extended benefit programme. See Winter-Ebmer (1998) and Lalive and Zweimüller (2004*a*,*b*) for analyses of this programme.

3.2. *The 1989 changes in policy parameters*

As of 1 August, 1989 the Austrian government enacted a series of important changes to the unemployment insurance rules (Arbeitslosenversicherungsgesetz). The first change was that the potential duration of UB payments became dependent not only on previous contributions, but

8. UB duration was 12 weeks for job seekers who did not meet either previous contribution requirement. In order to guarantee a minimum degree of homogeneity, this paper focuses on workers who have contributed for at least 52 weeks to UI in the two years prior to unemployment. Thus, all workers are eligible for at least 20 weeks of UB before the policy change.

9. This implies that job seekers who do not meet UB eligibility criteria can apply at the beginning of their spell.

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 has 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 has been employed for at least 468 weeks within the last 15 years.

The second important change was that the RR was increased from about 41% to roughly 47% in the previous income bracket between 5000 ATS and 10,000 ATS (Figure 2). After the income threshold of 10,000 ATS has been crossed, the RR is faded out quickly to reach the former level of 41% at 12,610 ATS previous monthly income. These changes affect all job seekers as of 1 August, 1989.

In the period that we study, there was a further slight increase to the benefit RR. As of 1 June, 1990, the Austrian government enacted an increase to the RR for those with previous income exceeding 12,610 but below 27,000 ATS. This change essentially ensures that the RR is faded out from a level of 47% to a level of 41% over the income range 10,000 ATS to 27,000 ATS instead of the range from 10,000 ATS to 12,610 ATS.

In this paper, we focus on the effect of the increase in PBD and the increase in RR in August 1989. Clearly, the change in policy parameters is different for different groups related to the three eligibility determinants: age, experience, and earnings. Earnings determine whether the RR goes up—only workers with low earnings get an increase in RR. Age and experience determine whether the PBD goes up—only workers from age 40 onwards with a high working experience get an increase in PBD. Furthermore, the size of the increase depends on the age of the individual. These policy changes created a nice empirical design that can be exploited for empirical research as it has elements of a "natural experiment".

We do not estimate the effect of the increase in RR in July 1990, for two reasons. First, this policy change entails a much weaker change to the benefit RR than the policy change in August 1989.[10] Second, the July 1990 change to the RR occurs rather shortly after the August 1989 change. This implies that the unemployment exit rate in the period before the July 1990 change, but after the August 1989 change is not identified in the period 48 weeks after entry and beyond. This time period is, however, of substantial interest in the present analysis. The fact that we do not account for the July 1990 change implies that the estimated treatment effects refer to treatments relative to a control treatment that does not leave unaffected all parameters of the unemployment insurance system. Thus, not accounting for the July 1990 policy changes the interpretation of the effects we report but it does not affect the internal validity of these effects. Moreover, since we evaluate the effects of these policy changes relative to a slight increase in benefits, our results give a *lower bound* on the effects relative to no change. Moreover, in Section 5.3 we will undertake sensitivity tests that account for possible effects of the 1990 policy change.

3.3. *The situation on the Austrian labour market 1987–1991*

Before we go into the details of data and statistical analysis, it is instructive to look at the situation on the Austrian labour market during the period 1987–1991. This is the period on which the empirical analysis below will be concentrated.

Table 1 shows that 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 favourable situation of the business cycle led to strong employment growth throughout the period under consideration. However, it did not show up in decreasing unemployment rates.

10. The average increase in the benefit RR is almost 6 percentage points for the August 1989 change. In contrast, the RR increases by a mere 1·3 percentage points due to the second policy change (Table 3).

TABLE 1

The Austrian labour market 1987–1991

	Real GDP growth	Employment growth	Unemployment rate
1987	1·7	0·6	5·6
1988	3·2	0·7	5·3
1989	4·2	1·4	5·0
1990	4·7	2·8	5·4
1991	3·3	2·8	5·8

Source: Statistics Austria.

The reason was primarily a strong increase in labour supply (a strong increase in immigration and rising female labour force participation). That is why unemployment rose slightly despite a strong employment growth.

It is worth noting that this situation is favourable for our empirical strategy. Employment growth during the treatment period was stronger than before. Hence it is unlikely that our results from a comparison of the labour-market experiences of older workers between the period prior to the policy change to the post-policy period is strongly driven by a deteriorating labour market.

4. DESCRIPTIVE ANALYSIS

4.1. *Data*

To assess the impact of changes to financial incentives on transition rates out of unemployment, we use longitudinal individual data from two different sources: (i) the *Austrian social security database*, 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 extract a sample that contains all unemployment entrants in the period between 1 August, 1987 (two years before the policy change) and 31 July, 1991 (two years after the policy change). We concentrate on job seekers in the age bracket 35–54, who have at least 52 weeks within the last two years and with residence in regions that were never eligible for a special regional extended benefit programme. Furthermore, in order to isolate the effects of changes in PBD, we concentrate on workers who either fulfil both previous contribution requirements[11] or neither. We end up with 225,821 unemployment spells. The median duration of unemployment is 12 weeks. More than 85% of spells end in a job, 14% of spells in a non-job exit destination (long-term sickness, pension, unknown). Since spells are observed until May 1999, only 1% of spells are right censored.

4.2. *Construction of groups*

Table 2 summarizes the changes to unemployment insurance that were enacted in August 1989. Eligibility depends on three criteria: previous gross monthly income, age, and previous work experience. Thus, in the data it is possible to distinguish four groups of job seekers. The first group comprises job seekers with monthly income exceeding 12,610 ATS who are aged 40 or older with much previous work experience (6 out of previous 10 years and 9 out of previous 15

11. That is, at least 312 weeks contributions in the previous 10 years and at least 512 weeks contributions in the previous 15 years.

TABLE 2

Eligibility (e) for change to the relative rate (RR) and for change to potential benefit duration (PBD)

		Age			
		Younger than 40 years		40 years and older	
		Work experience		Work experience	
		Low	High	Low	High
Monthly income	≤12,610 AS	eRR	eRR	eRR	ePBD–RR
	>12,610 AS	Control	Control	Control	ePBD

Notes: Work experience "low" refers to less than 6 out of previous 10 years and less than 9 out of previous 15 years work experience. Work experience "high" refers to worked more than 6 out of previous 10 and worked more than 9 out of previous 15 years. RR, replacement rate; PBD, potential benefits duration; ePBD, eligible for increase in potential benefit duration; eRR, eligible for increase in benefit RR; ePBD–RR, eligible for both.

years). This group is eligible for the change to PBD (the ePBD group). The second group consists of job seekers with income lower than 12,610 ATS and age less than 40 years or age exceeding 40 years but with little work experience. This group benefits from the increase in the RR (the eRR group). The third group contains job seekers with income lower than 12,610 ATS and aged 40 or older with much previous work experience. This group is affected by both, the increase in the benefit RR and the increase in PBD (the ePBD–RR group). The fourth group contains job seekers with income exceeding 12,610 ATS who are either young or have little work experience. For those individuals there was no change in the two central parameters of the unemployment insurance system (control group). The heterogeneity in treatment is obvious from the existence of four groups. There is not only heterogeneity in terms of the nature of the treatment but also in the size of the treatment. The RR increases with about 15% whereas PBD increases either with 30% or 75%. This additional variation expands the range over which parameters are estimated. Table 3 reports the means of the RR and of PBD together with the number of spells in the respective groups both before August 1989 and after August 1989. The first row shows that almost all job seekers entering unemployment before August 1989 in the ePBD group were eligible for the maximum duration of regular benefits of 30 weeks.[12] In contrast, the PBD was almost 43 weeks for spells in the ePBD group after the policy change. Thus, this group experienced an increase by 13·5 weeks in PBD. The second row in Table 3 shows that there is a slight increase in the RR. This increase is due to the fact that there is a second policy change in June 1990 affecting the high-income workers in the ePBD group. The third row shows that this group is the largest in size and that the number of spells in the ePBD group increases. However, note that this probably reflects the fact that eligibility to the change in the RR depends on previous income in nominal terms rather than the fact that more individuals register to collect unemployment insurance because benefits are more generous. In line with economic growth over the period 1987–1991 (Table 1), the total number of spells is almost identical before and after the policy change.

The second set of rows in Table 3 refers to the eRR group. PBD is lower on average in the RR group than in the PBD group. This reflects the fact that many spells in the RR group do not fulfil the 30-week requirement that at least 60% out of the previous five years have to be spent working. While there is virtually no change to PBD, the RR increases strongly, from 41% to 47%. This group is smaller than the first group, reflecting the fact that median income in the

12. The empirical analysis below accounts for the fact that unemployment insurance parameters vary due to the fact that the changes enacted in August 1989 pertain to *all* spells, not just to those started after August 1989.

TABLE 3

The changes to unemployment insurance in August 1989

	Before August 1989	After August 1989	Change (after–before)	Diff-in-diff (change compared to "control")
ePBD group				
PBD (weeks)	29·5	42·5	13·0	13·5
RR (%)	40·0	40·9	0·9	−0·3
N	48,294	51,110		
eRR group				
PBD (weeks)	25·1	24·6	−0·5	0·0
RR (%)	41·4	47·3	5·9	4·6
N	17,160	15,310		
ePBD–RR group				
PBD (weeks)	29·0	42·6	13·6	14·1
RR (%)	41·3	47·0	5·7	4·4
N	11,992	9182		
Control group				
PBD (weeks)	27·4	26·9	−0·5	—
RR (%)	40·2	41·5	1·3	—
N	33,815	38,958		
Total				
PBD (weeks)	28·1	34·8	6·7	—
RR (%)	40·4	42·5	2·0	—
N	111,261	114,560		

Notes: Diff-in-diff, difference-in-difference; PBD, potential benefits duration; RR, replacement rate; ePBD, eligible for increase in potential benefit duration; eRR, eligible for increase in benefit RR; ePBD–RR, eligible for both.
Source: Own calculations, based on Austrian Social Security Data.

unemployment inflow is higher than the nominal income threshold of 12,610 AS. The decrease in the number of spells in the group is due to the fact that previous income must be below a nominally fixed income threshold in order to be a member of the RR group.

The third set of rows in Table 3 refers to the ePBD–RR group. This group very much resembles a combination of the ePBD and of the eRR group exhibiting long PBD as in the ePBD group, and a rather high RR of 41% as in the eRR group. Interestingly, this group experiences an increase in PBD and in the RR in exactly the same magnitude as both previously discussed groups. This is the smallest group. Again, the number of spells allocated to this group declines since all individuals must have earned less than the nominal income threshold of 12,610 AS in order to be allocated to the ePBD–RR group.

The fourth set of rows in Table 3 refers to the *control group*. This second largest group has rather long PBD and relatively low RR before the policy change. There is a slight decrease in PBD over time and a slight increase in the RR (reflecting the policy change in June 1990) over time.

The last column in Table 3 reports difference-in-difference (diff-in-diff) estimates of the effect of the policy change on both parameters of the unemployment compensation system. This column shows that a corresponding diff-in-diff calculation for some outcome, identifies the effect of extending the PBD by 14 weeks (starting from 30 weeks), the effect of increasing the RR by 4·6 percentage points (starting from 41%), and the effect of increasing both PBD by 14 weeks and increasing the RR by 4·4 percentage points. Thus, this design allows for an exhaustive analysis of how financial incentives affect the duration of unemployment.

TABLE 4

Average unemployment duration in first 104 weeks (measured in weeks)

	Before August 1989	After August 1989	Change (after–before)	Diff-in-diff (change compared to "control")
ePBD group	16·25	18·67	2·42	1·13
	(0·08)	(0·09)	(0·12)	(0·18)
	48,294	51,110		
eRR group	17·79	20·03	2·24	0·96
	(0·12)	(0·16)	(0·20)	(0·24)
	17,160	15,310		
ePBD–RR group	19·01	23·55	4·53	3·25
	(0·17)	(0·24)	(0·20)	(0·24)
	11,992	9182		
Control group	15·24	16·52	1·29	—
	(0·08)	(0·09)	(0·13)	
	33,815	38,958		

Notes: Standard errors in parentheses. Diff-in-diff, difference-in-difference; RR, replacement rate; PBD, potential benefits duration; ePBD, eligible for increase in potential benefit duration; eRR, eligible for increase in benefit RR; ePBD–RR, eligible for both.
Source: Own calculations, based on Austrian Social Security Data.

4.3. *Unemployment duration*

Table 4 reports average unemployment duration in the first 104 weeks by programme implementation status and by group.[13] Let t_u denote the realized duration of unemployment measured in weeks. Unemployment duration in the first two years is $t_u^{104} \equiv \min(t_u, 104)$.

The first column in Table 4 shows that average unemployment duration is longest in the ePBD–RR group and shortest in the control group in spells that started before August 1989. The second column in Table 4 shows average unemployment duration after August 1989. The third column in Table 4 shows that unemployment duration increases in all groups. However, as column 4 in Table 4 shows, the change in unemployment is stronger in groups, which are eligible for either the change to PBD or RR or both. For instance, unemployment duration increases by 1·13 weeks more strongly in the ePBD group that is eligible for the extension of PBD but not for the increase in RR. There is a slightly weaker increase by 0·96 weeks in unemployment duration in the eRR group, which is eligible for the increase in RR but not for the extension of PBD. The strongest effect is in the group that is eligible for both RR and PBD. In the ePBD–RR group, unemployment duration increases by 3·25 weeks more strongly than in the control group. Furthermore, there appears to be an excess increase of 1·16 weeks (= 3·25 − 1·13 − 0·96) in unemployment duration in the sense that this effect exceeds the sum of the effects reported in the ePBD and in the eRR groups.

Obviously, Table 4 provides only a first crude check on how the policy changes may have affected durations. Interestingly, the results are in line with theoretical predictions discussed in Section 2. It is worth noting, however, that the diff-in-diff estimates of Table 4 are based on rather different groups. Moreover, unbiased estimates are obtained only if there are no group-specific trends in unemployment durations. To check for potential biases resulting from such group-specific trends, Section 5.3 below will provide a variety of robustness tests, including a focus on more narrowly defined groups and on groups that are just below or just above the

13. We report average unemployment in the first 104 weeks for two reasons. First, we will calculate the contribution to the total change as a function of elapsed unemployment duration. Second, right censoring is not an issue in the first 104 weeks. Moreover, note that results referring to total unemployment duration are very similar.

eligibility threshold. Within such more homogenous groups differential trends should be of minor importance.[14]

4.4. *Survivor functions*

Job search theory predicts that financial incentives in UI affect the shape of the unemployment exit hazard depending on the time to benefit exhaustion. In order to study this prediction, it is useful to decompose the effect on average unemployment duration reported in Table 4 as follows. It is well known that expected unemployment duration in the first two years is given by $E(T_u^{104}) = \int_0^{104} S(z)dz$ where $S(z) = \exp\left(-\int_0^z \theta(x)dx\right)$ is the survivor function, that is, the probability that unemployment spells are longer than z weeks and where $\theta(x)$ is the unemployment exit hazard, that is, $\theta(x) = f(x)/S(x)$ where $f(x)$ is the density of unemployment spells (Lancaster, 1990). This says, for instance, that the increase in average unemployment duration in the ePBD group by 2·42 weeks is due to the fact that the survivor function in the ePBD group after August 1989 was higher than the corresponding survivor function in the ePBD group based on spells that started before August 1989. Moreover, the difference in these survivor functions integrates to 2·42 weeks. Thus, analysing the effect of the policy change on survivor functions allows decomposing the total change in unemployment duration into contributions to this change as a function of duration.[15]

Figure 3 shows the Kaplan–Meier survivor functions for the four groups. The top left subfigure contrasts the survivor function after the policy change with the survivor before the policy change in the ePBD group. Clearly, after 15 weeks of elapsed unemployment duration, the survivor function after the policy change starts to diverge from the corresponding function before the policy change. The difference widens until about 40 weeks have elapsed. After this point, the difference becomes smaller again and stays constant at a low level after 65 weeks of elapsed duration. Thus, extending the PBD appears to create a "lens" that starts 15 weeks before benefit exhaustion in the old system, and that ends about 15 weeks after benefit exhaustion takes place in the new system.

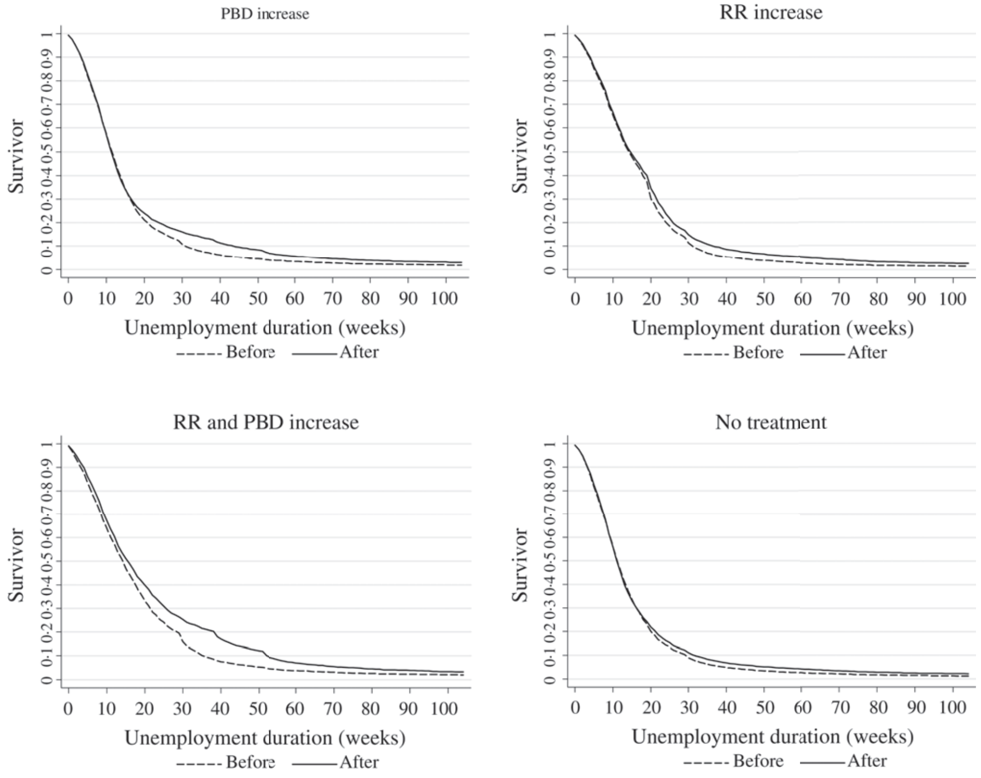
The top right subfigure in Figure 3 reports the difference in the survivor functions in the eRR group. In contrast to the previous findings, we note a slight increase in the survivor function that takes place almost from the start of the unemployment spell. The difference between the survivor functions becomes larger after 20 weeks of unemployment duration and again after week 30. The survivor function remains at a higher level than before the policy change even in the period when benefits have been exhausted, after week 30.

The bottom left subfigure in Figure 3 reports the survivor function analysis for the ePBD–RR group. This figure very clearly combines aspects of both policy changes. On one hand, we note a relatively strong increase in the survivor function right from the start of the unemployment spell. Again, there a "lens" starts to appear after 15 weeks of unemployment duration, which disappears only after 65 weeks of unemployment duration have elapsed.

The bottom right panel reports the survivor functions in the "control" group. There is no difference at all in the survivor functions up to 20 weeks of unemployment duration. Thereafter, we note a slight upward shift in the survivor function. Thus, in order to isolate the effects of the changes to the unemployment compensation system in August 1989, it is necessary to net out the change occurring over time from the raw effects on the survivor functions in the previous three subfigures.

14. A further reason why results in Table 4 could be biased are substitution effects. As individuals from the treatment group are less likely to accept jobs, individuals from the control group may be able to leave unemployment more quickly.

15. Note that elapsed unemployment duration is not time to benefit exhaustion. However, recall that in the old system, only two levels of the PBD prevailed, that is, 20 and 30 weeks. Thus, elapsed duration is very closely related to time to benefit exhaustion.



Notes: Before, spell starts before August 1989; after, spell starts August 1989.
Source: Own calculations, based on Austrian Social Security Data.

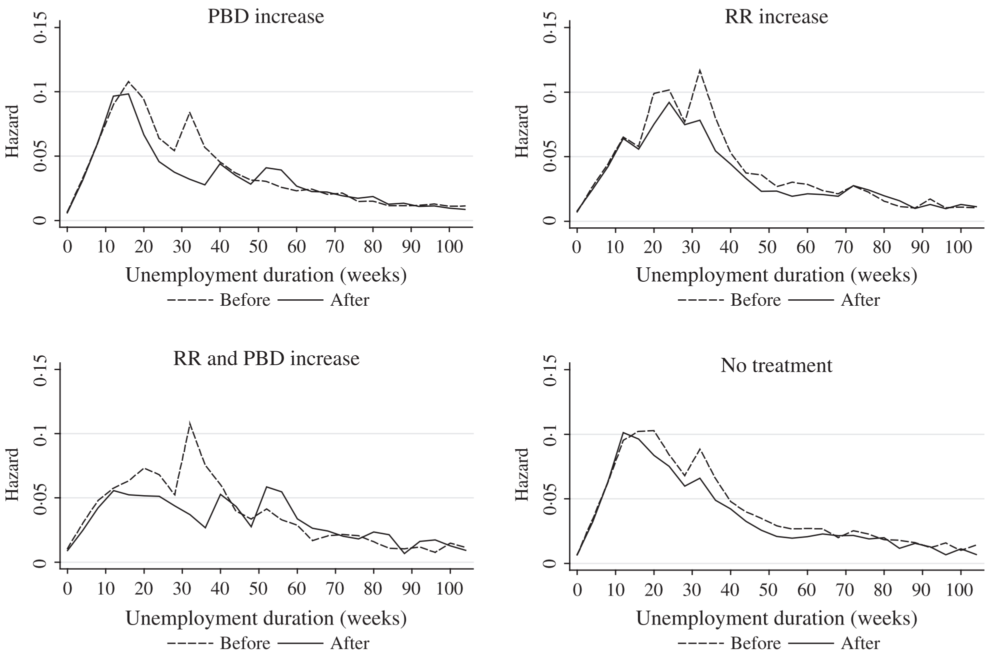
FIGURE 3

Kaplan–Meier survivor functions

4.5. *Exit hazards*

Figure 4 reports the Kaplan–Meier estimates of the unemployment exit hazard by period and group. The top left subfigure refers to the ePBD group. The unemployment exit rate before the policy change (dashed line) is very low at the start of the unemployment spell, reaches a maximum of 0·1 per week after 20 weeks of unemployment have elapsed, and declines gradually to a very low level. Interestingly, there is an important spike in the unemployment exit rate in week 30—the week when regular unemployment benefits are exhausted for almost all individuals in this group. This replicates the important findings in Meyer (1990). There are two important differences between the unemployment exit rate before August 1989 and the corresponding rate after August 1989. First, the spike that was observed in week 30 "moves" to weeks 39 and 52. Second, the unemployment exit rate is strongly depressed in the period from week 20 and ending in week 40. This is the period just before exhaustion in the old system and in between old and new exhaustion weeks.

The exit rate in the eRR group is characterized by two spikes in the old system in weeks 20 and 30 (top right subfigure). In the new system, the exit rate is slightly depressed already from the start of the unemployment spell. Thus, an important difference between changes to PBD and to RR emerges. In line with theoretical predictions, the exit rate is depressed from the start



Notes: Before, spell starts before August 1989; after, spell starts August 1989.
Source: Own calculations, based on Austrian Social Security Data.

FIGURE 4

Kaplan–Meier unemployment exit rates

with changes to RR. In contrast, with changes to PBD this rate is initially unaffected, but varies strongly around dates of benefit expiration.

The combined effects of RR and PBD can potentially be studied in the bottom left subfigure (ePBD–RR group). In the old system, the unemployment exit rate is rather low until 30 weeks of regular benefits have elapsed. In the new system, we observe a depressed hazard from the start of the spell. Moreover, whereas the unemployment exit rate shoots upward between week 30 and week 40 in the old system, the exit rate is strongly depressed in the corresponding period in the new system. Furthermore, there are two notable spikes centred after 39 and 52 weeks in the new system.

The remaining subfigure shows the unemployment exit rate for the group that was not affected in August 1989. We note that even though there was no change to unemployment insurance parameters in this group, the exit rate after August 1989 appears to be lower from 10 weeks until 65 weeks of elapsed unemployment duration. There are at least two reasons for this reduction in the exit rate. First, real GDP growth was lower in 1991 than in 1990 (Table 1). Second, in June 1990 this group was affected by a slight increase in the benefit level. Both factors may have contributed to a lower unemployment exit hazard after August 1989 compared to the period before August 1989.

This descriptive analysis already provides important insights into the mechanism by which financial incentives in UI affect unemployment duration. However, a number of important aspects were not accommodated so far. First, job search theory models the unemployment exit hazard for homogeneous workers. So far, however, we have discussed unemployment exit rates that refer to very heterogeneous groups of job seekers. It has been shown that failure to account for heterogeneity biases the duration dependence of the unemployment exit hazard and, potentially, the effects of financial incentives on unemployment. Second, the change to PBD is heterogeneous: PBD increases by 9 weeks for workers aged 40–49 years and by 22 weeks for workers aged 50 years and older. These problems will be addressed in the context of an econometric model of the unemployment exit rate in the following section.

5. RESULTS

5.1. *Statistical model*

To estimate how financial incentives affect the unemployment exit hazard, we apply a proportional hazard model.[16] The proportional hazard model posits the following specification for the exit rate $\theta(t_u \mid x) = \lambda(t_u)\exp(x\beta)$, where $\lambda(t_u)$ captures the baseline duration dependence of the hazard (in weeks), and x are the observed characteristics of the individuals.[17] The baseline duration dependence is of central interest in this paper because it refers to the exit rate for a homogeneous group of workers. We specify the duration dependence of the hazard as a piecewise constant function of elapsed duration as follows:

$$\lambda(t_u) = \exp\left(\sum_{l=0}^{14} \lambda_l I(4l < t_u \leq 4(l+1)) + \lambda_{15} I(t_u > 60)\right). \tag{1}$$

Thus, the hazard rate shifts in every four-week interval. Because there are very few transitions beyond week 60, the last time interval covers the entire remaining duration of the spell as of week 60.

The treatment effect can be identified in a (log) diff-in-diff setting. Denote eligibility for the extension of PBD from 30 to 39 weeks by $\text{eP39} = I$ $(\text{ePBD} = 1, \text{age} < 50)$, eligibility for an extension to 52 weeks is denoted by $\text{eP52} = I$ $(\text{ePBD} = 1, \text{age} \geq 50)$. Second, introduce the calendar time varying function $A89(t_c) = I(t_c \geq \text{mdy}(8, 1, 1989))$ where t_c measures calendar time in days since 1 January, 1960, and $\text{mdy}(x, y, z)$ gives the number of days since 1 January, 1960 of day y in month x and year z. The function $A89(t_c)$ records the moment a spell enters the period after the policy change has taken place. Thus, the interaction term $\text{eP39}{*}A89(t_c)$ indicates that an individual satisfying all eligibility criteria for the extension to 39 weeks has entered the period when this policy change has been enacted. The duration dependence of the hazard rate is specified as follows:[18]

$$\begin{aligned}
\lambda_l = {} & \beta_{0l} + \beta_{1l}\text{eP39} + \beta_{2l}\text{eP52} + \beta_{3l}\text{eRR} \\
& + \beta_{4l}(\text{eP39} + \text{eP52}) * \text{eRR} + \beta_{5l} A89 \\
& + \delta_{1l}\text{eP39} * A89 + \delta_{2l}\text{eP52} * A89 + \delta_{3l}\text{eRR} * A89 \\
& + \delta_{4l}\text{eP39} * \text{eRR} * A89 + \delta_{5l}\text{eP52} * \text{eRR} * A89 \\
l = {} & 0, \ldots, 15.
\end{aligned} \tag{2}$$

16. See Van den Berg (2001) for a recent survey of the properties of the mixed proportional hazard model and for applications of this model to duration data.

17. These are age, marital status, female, education, log(previous monthly income), recall status, blue collar, seasonal industry, manufacturing industry, time spent non-employed, tenure, and quarter of inflow. Use of the $\exp(\cdot)$ function guarantees that the hazard rate is non-negative in the entire domain of $x\beta$.

18. For ease of exposition, we suppress dependence of $A89(t_c)$ on calendar time t_c and write $A89$.

The set of β parameters captures *ex ante* differences between groups ($\beta_{1l},\ldots,\beta_{4l}$) and changes to duration dependence occurring over time (β_{5l}). Note that we assume that there are no *ex ante* differences between those individuals who are eligible for the change to RR and an extension of PBD from 30 to 39 weeks compared to individuals eligible for the change to RR and the change to PBD from 30 to 52 weeks. The set of δ parameters measure the change in the duration dependence of the hazard rate due to changes in financial incentives. There are five sets of δ parameters because the policy change entails five interventions (P39, P52, RR, and combinations). δ_{1l} and δ_{2l} capture the effect of extending PBD, δ_{3l} captures the effect of increasing the gross RR. The parameters δ_{4l} and δ_{5l} test whether changes to both dimensions of unemployment insurance affect the unemployment exit rate in a way that would not be expected from two separate changes to one dimension only. Thus, these parameters address the issue whether increasing the generosity of the unemployment insurance system due to simultaneous changes in RR and PBD generates disincentive effects beyond the effects expected from two uni-dimensional changes.

With this proportional specification of the exit hazard, the individual likelihood contribution is

$$L_i = \theta(t_{u_i} \mid x_i)^{1-c_i} S(t_{u_i} \mid x_i), \tag{3}$$

where $c_i = 1$ if the spell is right censored, and $c_i = 0$ otherwise; and $S(t_{u_i} \mid x_i) = \exp\big(-\int_0^{t_i} \theta(z \mid x_i)dz\big)$ is the survivor function. The likelihood function is obtained as the product of the individual likelihood contributions. The parameter estimates are obtained by maximizing the log likelihood.[19]

The conditional hazard estimates address two important issues. First, the model allows for the differences across treatments. Second, the hazard rate is identified conditional on important differences across individuals captured by x.

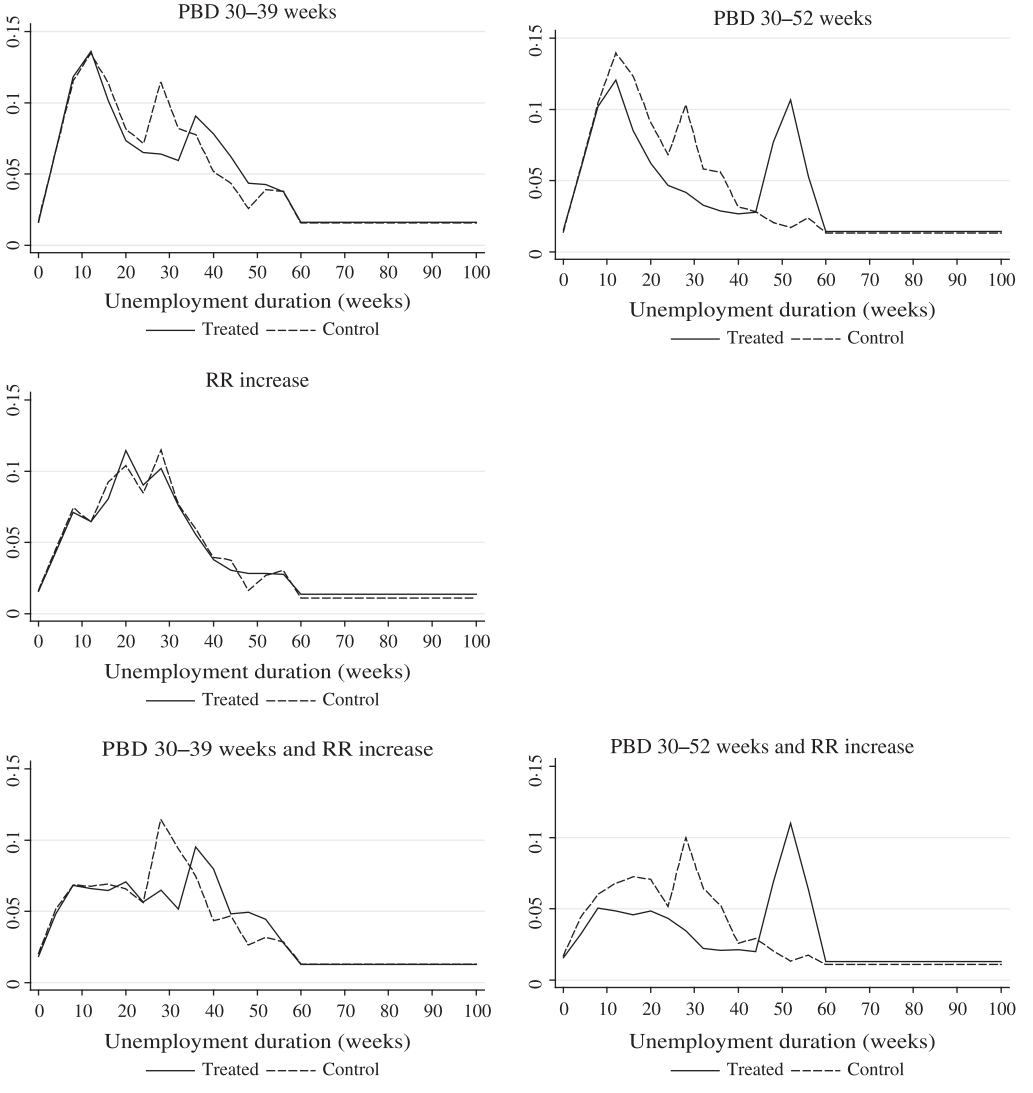
5.2. *Main results*

Figure 5 reports the difference in the factual hazard rate with treatment to the counterfactual hazard rate without treatment.[20] To obtain the factual hazard rate (solid line of Figure 5), we calculate the prediction for the individual hazard rate with treatment $\hat{\theta}_1(t_u \mid x_i)$, $t_u = 0, 4, \ldots, 56, 60, 100$, and take the average within the group of treated individuals, that is, we average with respect to the distribution of individual characteristics x_i in the population receiving the treatment. To obtain the counterfactual hazard rate (dashed line in Figure 5) we impose all treatment effects δ to be 0 to get the individual hazard rate $\hat{\theta}_0(t_u \mid x_i)$ and again average across treated individuals. The difference between the two hazards is equal to the “average treatment effect on the treated”.[21] The top left subfigure reports the effect of extending the PBD from 30 to 39 weeks. Increasing PBD does not affect the unemployment exit rate until an elapsed duration of about 16 weeks. In week 30 (the old exhaustion date), the differences in hazard rates between the new and the old system become very large. Thereafter, the hazard rate in the new system increases strongly compared to the corresponding rate in the old system. In weeks 42–50—when benefits have expired also in the new system—the unemployment exit hazard is higher in the new system. The higher exit rates after benefit expiration under the new system could be taken as evidence for a (transitory)

19. This specification does not allow for unobserved heterogeneity. It turns out that allowing for unobserved heterogeneity does not affect results. Results are available upon request from the authors.

20. See table A1 on `http://www.restud.org/supplements.htm` for the coefficient estimates.

21. Note that the hazard rate is constant over successive four-week periods. Figure 5 plots the hazard rate with respect to the start of the four-week period.



Notes: *x*-axis gives the beginning of the duration interval.
Source: Own calculations, based on Austrian Social Security Data.

FIGURE 5

Estimated average treated and control hazard rates (based on table A1)

entitlement effect (discussed in Figure 1). Alternatively, it could simply mean that success from increased search effort takes some time to materialize. From week 54 onwards, there is no effect of extending PBD on the unemployment exit rate.

In contrast, extending benefits from 30 to 52 weeks reduces the exit hazard in an earlier stage of the unemployment spell (top right subfigure). While the unemployment exit rates under the two systems are initially very similar, they start to diverge at week 14. The largest difference occurs, again, in week 30—the period when benefits expire in the old system. Between weeks

30 and 42 the hazard under the old system continues to be above the one of the new system. Thereafter, the benefit exhaustion effect of the new system dominates, reaching its peak at week 54 (which covers the period of benefits under the new system) and stays higher until week 58. In week 62 (which covers the entire period from week 60 to the end of the spell) there is no effect of extending the PBD.

Increasing the RR tends to depress the unemployment exit rate much less strongly than changing PBD (middle subfigure). Individuals with access to more generous unemployment benefits tend to leave unemployment less rapidly in the covered period (weeks 2–30, exceptions are weeks 22 and 26). After benefits have expired for all individuals there is no effect on the unemployment exit rate (with the exception of weeks 50 and 62).

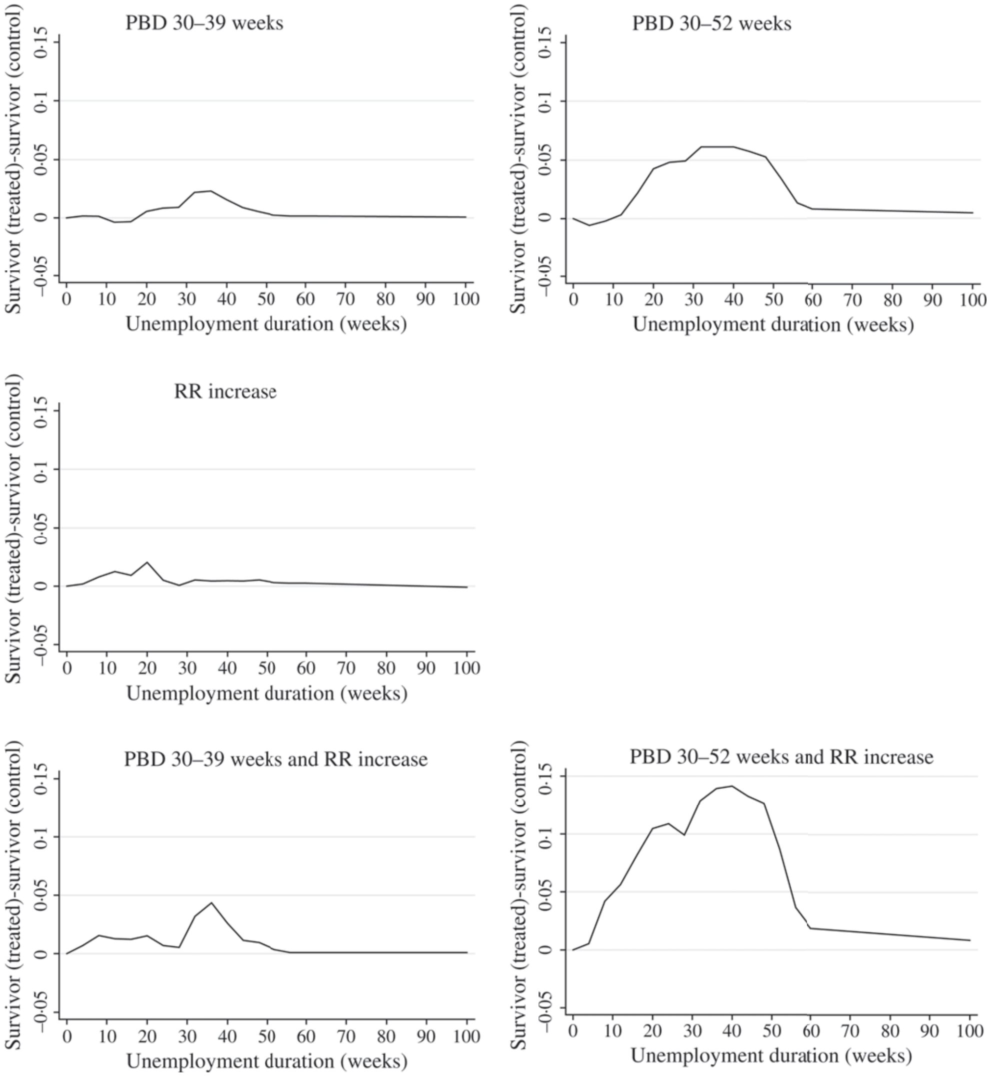
Does it matter whether individuals are affected by a combined change instead of two separate changes to financial incentives? The bottom left subfigure answers this question by comparing exit rates for the group that experienced both an increase in RR and a nine-week increase in PBD. The hazard rate is slightly (yet insignificantly) lower in the first four-week interval. Thereafter, the hazard rates are very similar—with the exception of the hazard at week 18. The main difference between the two graphs shows up at weeks 26, 30, and 42 (around old and new benefit expiration dates). The hazard under the new system is higher until week 60. Thereafter, no differences remain.

The bottom right subfigure shows the hazard rates for the group, which experienced both an increase in RR and a 22-week increase in PBD. The two graphs diverge already from the beginning of the unemployment spell. The hazard rate under the old system lies above the one under the new system until week 48. Only around the date of benefit expiration (weeks 46, 50, and 58) hazard rates are higher under the new system. No difference remains after week 60.

So far we have discussed hazard rate estimates. These are important because job search theory offers sharp predictions regarding the impact of unemployment insurance parameters on the exit hazard. However, policy makers are interested in the implied effects on unemployment durations or, equivalently, effects on survivor functions. Figure 6 reports the difference in the factual survivor function with treatment to the counterfactual survivor function without treatment. Specifically, in the first step we estimate the survivor function with treatment for each individual as implied by the hazard rate estimates $\widehat{S}_1(t_u \mid x_i) = \exp\big(-\int_0^t \hat{\theta}_1(z \mid x_i)dz\big)$ with $t_u = 0, 4, \ldots, 56, 60, 100$. The corresponding survivor function without treatment is $\widehat{S}_0(t_u \mid x_i) = \exp\big(-\int_0^t \hat{\theta}_0(z \mid x_i)dz\big)$. The counterfactual hazard rate without treatment is obtained by imposing all treatment effects δ to be 0. This gives the change to the survivor function in the treated group at the individual level. In the second step, we average this change with respect to the distribution of individual characteristics x_i in the population receiving the treatment. This gives the change in the population survivor function reported in Figure 6.

Extending PBD by 9 weeks entails a positive contribution to the change in expected unemployment duration in the time period between 20 and 50 weeks (top left subfigure). Both, the periods before 20 weeks of elapsed duration and the period after 50 weeks of duration have elapsed do not contribute to increasing expected duration. The maximum contribution arises in week 35, exactly in between the old benefit exhaustion week (30) and the new benefit exhaustion week (39).

Results are qualitatively similar but quantitatively much stronger for an increase of PBD by 22 weeks (top right subfigure). Again, the unemployment spell can be divided in three periods. From week 0 to week 12, the contribution to expected unemployment duration is slightly negative, from week 12 to week 60, the contribution is strongly positive, and from week 60 onwards, the contribution to expected unemployment duration is positive but small. Again, the maximum contribution occurs at week 40, which is roughly in between the old exhaustion week (30) and



Notes: *x*-axis gives the end of the duration interval. Difference refers to the end of the duration interval.

Source: Own calculations, based on Austrian Social Security Data.

FIGURE 6

Simulations: difference between survior function with treatment and survivor function without treatment (control)

the new exhaustion week (52). However, the strongest difference lies in the magnitude of the contribution. Whereas extending duration by 9 weeks generates a maximum contribution on the order of 2·5 percentage points, the corresponding maximum contribution due to a 22-week increase exceeds 5 percentage points.

In contrast to benefit duration extensions, an increase in the benefit RR generates a positive contribution to expected unemployment duration right from the start of the unemployment spell (middle subfigure). Most of the prolonging contribution occurs in the covered period of the unemployment spell (weeks 0–30). There is also a positive, but much weaker contribution to

TABLE 5

Simulated effects on expected duration in first 104 weeks

	Treated	Control	Effect
No treatment	16·91	16·91	0·00
Change to one parameter			
PBD 30–39 weeks	17·53	17·08	0·45
PBD 30–52 weeks	20·62	18·35	2·27
RR increase	20·97	20·60	0·38
Change to two parameters			
PBD 30–39 and RR increase	21·95	21·09	0·86
PBD 30–52 and RR increase	29·43	23·70	5·72

Notes: Based on population receiving the treatment in the period after the policy change. RR, replacement rate; PBD, potential benefits duration.
Source: Own calculations, based on Austrian Social Security Data.

expected unemployment duration in the period that is no longer covered by regular unemployment benefits (week 30 onwards).

The bottom two subfigures report results for interventions that increase PBD as well as RR. Two interesting results emerge in comparison with isolated changes to PBD (top two subfigures). First, the contribution to expected unemployment duration is positive from the start of the unemployment spell. This is clearly the impact of the RR on top of the PBD effect. Second, the maximum contribution to expected unemployment duration increases strongly from 2·5 percentage points to more than 4 percentage points (PBD 30–39 weeks) and from 6 percentage points to almost 14 percentage points (PBD 30–52 weeks).

To indicate the effects of the changes in financial incentives Table 5 shows the average unemployment duration in the first 104 weeks of the unemployment spell.[22] The first column in Table 5 gives the factual expected unemployment duration with treatment for the five treated groups and the group that is not affected by an intervention in August 1989.[23] The second column in Table 5 gives the counterfactual expected unemployment duration without treatment for the five treated groups. The third column gives the effect of the interventions on expected unemployment duration.[24]

Extending the PBD by nine weeks tends to increase expected unemployment duration by 0·45 weeks or by 0·05 weeks per additional week of PBD (second row). Increasing PBD by 22 weeks generates about 2·3 additional weeks of unemployment (third row). Thus, the second PBD extension produces twice as many weeks of unemployment per additional week of PBD (0·10). This result lies within the range of previous findings regarding the effect of PBD on unemployment duration (see Section 2) and is also similar to the estimates of Lalive and Zweimüller (2004*a*) who find a disincentive effect of 0·05 weeks per additional week of PBD. In contrast,

22. We report expected unemployment duration in the first 104 weeks because, in order to estimate total expected unemployment duration we need to know the survivor function until infinity. Since inference on the survivor function tends to become ever more unreliable as we extend the duration of the unemployment spell, we arbitrarily limit our discussion to the first 104 weeks, which are quite well identified in our large data-set. Expected unemployment duration is obtained by integrating the population survivor function with respect to time up to 104 weeks.

23. Recall that average unemployment duration in the control group is 16·5 weeks in the period after 1989 (Table 4). The corresponding number implied by the econometric model is 16·9 weeks (top, left cell). This is strong evidence that the econometric model fits the data well. The resulting difference is due to the fact that average unemployment duration treats spells, which are right censored in the first 104 weeks as completed, whereas the econometric model accounts for right censoring.

24. Note that the simulation results in Table 5 give the "effect of treatment on the treated". A concern with these simulations is that the treated groups differ from the control group. We deal with this concern below in the sensitivity analysis.

increasing RR by 6 percentage points tends to prolong unemployment duration by 0·38 weeks (fourth row). This implies an elasticity of unemployment duration with respect to the RR of about 0·15, which is small compared to the results of other studies discussed in Section 2.

Individuals eligible to a combined nine-week increase to PBD and a 6 percentage point increase in RR are unemployed for 0·86 weeks longer than in the counterfactual situation without this combined intervention (fifth row). Interestingly, this change to unemployment duration equals almost exactly to the prediction obtained from two separate changes, that is, $0{\cdot}38 + 0{\cdot}45 = 0{\cdot}83$. Individuals who get both, a 22-week increase in PBD and a 6 percentage point increase in RR are unemployed much longer than in the counterfactual situation of no intervention (5·72 weeks, sixth row). Note that the "adding up result" no longer obtains for this group of individuals, that is, $2{\cdot}27 + 0{\cdot}38 = 2{\cdot}65$ weeks instead of 5·72 weeks.[25] This result is obvious already from Figure 6. Whereas the survivor curve difference in the case "PBD 30–39 weeks and RR Increase" is approximately the sum of "PBD 30–39 weeks" and "RR Increase", this is not true for the case "PBD 30–52 weeks and RR Increase".

5.3. *Sensitivity analysis*

A question that arises from the previous subsection is to what extent the results are driven by the heterogeneity of treatment groups and control group. For example, if the exit rates of these groups are subject to different time trends, estimated treatment effects will be biased. To investigate how robust our results are, we perform two types of sensitivity analyses focusing on the main issue of heterogeneity in treatment and control groups. First, we redo the analysis for several specific subgroups that are more similar than treated and controls in the baseline model in terms of age, monthly income, and previous wage. Second, we redo the analysis for a subsample that is closer to the "policy intervention threshold" in terms of calendar time and age.

Columns 2–6 of Table 6 present the treatment effects when identification is based on smaller, but more similar groups. Column 1 of Table 6 reproduces our baseline estimates for ease of comparison.

In column 2 we identify the RR-effect from restricting the sample to young, low-wage individuals. Restricting the sample to individuals below age 40 means that included workers are not eligible to extended PBD. Restricting the sample to individuals with income below ATS 17,610 lets us compare low-wage individuals below the RR-eligibility threshold (ATS 12,610) to low-wage individuals not widely above the threshold. This is in the spirit of regression discontinuity analysis (see Hahn, Todd and Van der Klaauw, 2001), which exploits situations where assignment to treatment is a discontinuous function of some given variable (in the present context, an individual's pre-unemployment income). Column 2 shows that expected duration is 0·31 weeks longer for treated individuals, an effect, which is very close to our baseline estimate of 0·38 weeks.

Columns 3 and 4 of Table 6 identify the PBD effect from comparing unemployment durations of individuals with a sufficiently high pre-unemployment income (to ensure that RR remains unchanged) who are either older than 40 (column 3) or have high work experience (column 4). In other words, in column 3 treated and controls are more similar along the dimension of age and income, whereas in column 4 the groups are more similar along the dimension of experience and

25. We have also investigated the adding up result by comparing expected duration when we ignore the parameter estimates that measure the additional effect of the joint change to expected duration that allow for the effect due to the joint change. This analysis indicates that the effect of changing RR and PBD for the prime-age workers is the same as the effect predicted from two separate changes. In contrast, two separate changes in the PBD and in the RR for the older workers are predicted to increase unemployment duration by 2·98 weeks whereas the total change is 5·72 weeks.

TABLE 6

Sensitivity analyses: making the control group more similar. Expected duration (weeks)

	Baseline result (1)	Identify RR (2)	Identify PBD (3)	Identify PBD (4)	Identify PBD+RR (5)	Identify PBD+RR (6)	1990 reform
Sample restrictions							
Age	—	<40	≥ 40	—	≥ 40	≥ 40 or <40	—
Monthly income	—	$<17,610$	$>12,610$	$>12,610$ high	$\leq 12,610$ and high or $>12,610$ and low	$\leq 12,610$ high or $>12,610$ high	—
Work experience	—	—	—			—	
Change to one parameter							
PBD 30–39 weeks	0·45	—	0·36	0·63	—	—	0·51
PBD 30–52 weeks	2·27	—	1·99	2·29	—	—	2·17
RR increase	0·38	0·31	—	—	—	—	0·82
Change to two parameters							
PBD 30–39 and RR increase	0·86	—	—	—	0·76	0·65	1·50
PBD 30–52 and RR increase	5·72	—	—	—	5·97	6·25	6·79
N	225,821	39,685	117,208	142,523	38,978	64,293	225,821
$\ln L$	−792,903	−134,303	−415,646	−489,566	−145,098	−219,244	−792,445

Notes: "High" refers to individuals with 6 years out of previous 10 and 9 years out of previous 15 work experience; "low" refers to individuals who satisfy neither criterion. RR, replacement rate; PBD, potential benefits duration.
Source: Own calculations, based on Austrian Social Security Data.

income. The resulting PBD-treatment effects indicate that the increase in PBD from 30 to 39 weeks increases average unemployment duration by 0·36 weeks (column 3) and 0·63 weeks (column 4), respectively. Similarly, the increase in PBD from 30 to 52 weeks increases average unemployment duration by 1·99 and 2·29 weeks. Notice that our baseline estimates of 0·45 weeks (PBD 30–39 weeks) and 2·27 weeks (PBD 30–52 weeks) are within the range of these estimates.

Finally, columns 5 and 6 identify the joint effect of RR and PBD increases. In column 5 we confine the sample to workers older than 40 and in column 6 we concentrate only on high-experience workers. Variation in RR plus PBD eligibility relies, respectively, on differences across groups in experience and income (column 5), and on differences in age and income (column 6). Depending on the included groups, an increase in RR together with an increase in PBD from 30 to 39 weeks raises the average duration of unemployment by 0·76 and by 0·65 weeks, respectively. Again, this is very close to our baseline estimate of 0·86 weeks. The same holds for the joint RR and PBD increase from 30 to 52 weeks. Confining samples to older workers yields an estimated increase in unemployment duration of 5·97 weeks. Alternatively, when we confine samples to high-experience workers this estimate is 6·25 weeks. Both estimated effects are very close the 5·72 weeks baseline estimate. In sum, the sensitivity analysis of Table 6 makes us confident that our baseline results are rather robust and not particularly sensitive to permutations in the samples used to identify the treatment effects.

A final sensitivity analysis conducted in Table 6 concerns the 1990 policy change in RR that has so far been disregarded. Recall that in June 1990 policy change led to an increase in RR, which was roughly linear in income (RR = 0·47 with income 12,610 AS and RR = 0·41 with income 27,000 AS). On average, this led to an increase in the gross RR of about 2 percentage points (Table 3). Column 7 shows that controlling explicitly for the 1990 policy change leads to slightly higher effects of the estimated parameters, but does not affect the general picture.[26]

The results of the second type of sensitivity analysis are shown in Table 7. Here, in the spirit of regression discontinuity analysis, we adjust groups in such a way that they are close to the eligibility threshold. More precisely, we only include workers aged 38–41 and 48–51, respectively (as opposed to workers aged 35–54 in the baseline model); and we confine our analysis to the inflow to one year before and one year after the policy change (as opposed to two years before and after in the baseline model). This reduces the size of the sample strongly, from more than 225,000 workers in the baseline case to less than 37,000 workers in the threshold sample. The former sample restriction should substantially reduce heterogeneity between groups as age is a very important determinant of unemployment exit rates. The latter restriction is a simple check whether the results in the baseline model are driven by a trend that adversely affects treated workers. If this is indeed the case we would expect smaller treatment effects in the "threshold" analysis, which is based on a much more narrow time window. For ease of comparison column 1 in Table 7 reproduces our baseline results, column 2 shows the results from the threshold sample.

The results from the threshold sample indicate that increasing PBD from 30 to 39 weeks increases the expected duration of unemployment by 1·38 weeks, whereas increasing PBD from 30 to 52 weeks increases expected duration by 2·56 weeks. While the latter effect is of the same order of magnitude as in the baseline model (2·27 weeks) the former effect is now larger than in the baseline model. Similarly, we find that the increase in RR has a stronger effect on expected duration in the threshold sample where it raises expected duration by 1·50 weeks as opposed to 0·45 weeks in the baseline model. The joint PBD and RR changes lead to respective increases in unemployment duration by 1·40 weeks (PBD 30–39) and by 6·23 weeks (PBD 30–52). The

26. We control for the 1990 policy change by allowing for a change in the baseline hazard rate affecting all spells as of 1 July, 1990. A second change is modelled only for the high-income group that is affected by the 1990 change. This implies that the 1989 policy parameters are identified using information from the two pre-programme years and the post-programme period lasting from August 1989 to July 1990.

TABLE 7

Sensitivity analysis: getting close to age and calendar time threshold. Expected duration (weeks)

	Baseline result	Threshold
Sample restrictions		
Age	35–54	38–41 or 48–51
Calendar time	1 August, 1987 to 31 July, 1991	1 August, 1988 to 31 July, 1990
Change to one parameter		
PBD 30–39 weeks	0·45	1·38
PBD 30–52 weeks	2·27	2·56
RR increase	0·38	1·50
Change to two parameters		
PBD 30–39 and RR increase	0·86	1·40
PBD 30–52 and RR increase	5·72	6·23
N	225,821	36,587
$\ln L$	−792,903	−130,269

RR, replacement rate; PBD, potential benefits duration.
Source: Own calculations, based on Austrian Social Security Data.

former is somewhat larger than, the latter similar to the baseline results. However, the general picture is similar than in the baseline model. Changing RR and PBD simultaneously leads to an increase in expected unemployment duration larger than the effects of two isolated increases for the older workers (PBD 30–52) whereas this is not the case for younger workers (PBD 30–39). The results in Table 7 also indicate that group heterogeneity in the baseline model does not bias our results upwards. In contrast, we find somewhat stronger treatment effects in the threshold analysis.

In addition to these two types of sensitivity analysis we performed a number of additional analyses. We investigated to what extent unobserved heterogeneity matters, what happens if we exclude seasonal workers and whether the results change if we focus on exits to regular jobs instead of exits out of unemployment. Basically, our main results do not change.[27]

5.4. *Empirical estimates and theoretical predictions*

The empirical analysis in Sections 5.2 and 5.3 has revealed several interesting pieces of evidence. *First*, we have seen that an isolated increase in RR by about 15% (6 percentage points) leads to an increase in unemployment duration of roughly equal size as the 30% increase in PBD (9 weeks). *Second*, the increase in unemployment durations is relatively small for the PBD extension from 30 to 39 weeks, but is much larger for the increase from 30 to 52 weeks. Whereas for the PBD extension from 30 to 39 weeks the extended unemployment duration is 0·35 days for every week of extra PBD, this is 0·70 days for every week of extra PBD for the increase from 30 to 52 weeks. *Third*, the effect of a joint increase in RR and PBD from 30 to 39 weeks is only slightly larger than the sum of the effects from two separate changes whereas the effect of joint increase in RR and PBD from 30 to 52 weeks is much larger than the effects of two separate changes.

In Section 2 we have reviewed the main predictions from job search theory. Let us now discuss to which extent our empirical results are consistent with these predictions and which pieces of evidence are not. A first prediction from job search theory was that if entitlement effects are

27. We dealt with unobserved heterogeneity by allowing for a discrete distribution of unobserved heterogeneity with two mass points (Heckman and Singer, 1984). Although we find evidence of the presence of these two mass points the estimated effects of a change in RR and PBD are not affected by taking unobserved heterogeneity into account. We did not find evidence of the presence of more than two mass points. All results of the additional analyses are available on request.

negligible and/or most unemployment exits take place before benefits expire, increases in RR and extensions in PBD should lead to a reduction in job search effort and hence to longer unemployment durations. This prediction is clearly consistent with the evidence. In all our estimations we found that increasing RR and/or extending PBD resulted in an increase in unemployment duration. A second prediction was that increases of RR should trigger strong behavioural responses early in the unemployment spell, whereas extensions of PBD should lead to strong responses around the dates when benefits expire. Also, this prediction is clearly consistent with data. We find that increases in PBD are associated with strong changes in exit rates from unemployment around (old and new) dates of benefit expiration, whereas behavioural responses due to an increase in RR are distributed more uniformly over the unemployment spell. This pattern also helps to rationalize the result that a 15% increase in RR can lead to roughly the same effect on unemployment duration than a 30% extension of PBD. This is because the change in RR affects behaviour strongest from the start of the unemployment spell (hence affecting *all* unemployed) while the extension of PBD has the largest effects around dates of expiration (when many have already left unemployment). A third prediction was that a simultaneous increase in RR and PBD should lead to an increase in unemployment durations larger than the sum of the increases from two isolated changes in these policy parameters. Our estimates clearly confirm this prediction for older workers whereas for prime-age workers only small (and statistically insignificant) additive effects were found.

Hence, while there are several theoretical predictions that are clearly consistent with the data, we are left with two empirical results that are *prima facie* difficult to reconcile with theoretical predictions. The first concerns the heterogeneity of PBD effects across age groups. The reason why older workers react much more strongly than prime-age workers has to do with conditions that are specific to this group. Older workers have a weaker labour-market position. They get fewer job offers because employers preferably fill their vacancies by hiring young or prime-age workers. Moreover, the particular institutional environment may have an important effect on older workers' search efforts. Older workers are close to retirement (or face a higher probability of becoming eligible to early retirement benefits than prime-age workers). Hence, they have a lower incentive to search because the value of finding a job is small. For this reason, a more generous unemployment insurance system may induce older workers to reduce their search efforts more strongly than prime-age workers.

The second result that is difficult to reconcile with theoretical prediction concerns the age differences in reactions to a joint variation in policy parameters. For the negligible interaction effect for prime-age workers we lack a theoretical explanation. One reason could be that the size of the two policy changes is simply too small meaning that such an interaction effect is difficult to measure anyway. In contrast, the large interaction effect for older workers can be rationalized in much the same way as the large isolated PBD effects discussed above: labour-market conditions together with the institutional environment most likely explain the strong behavioural responses of the older workers to a joint increase in RR and PBD.[28]

Finally, note that our sensitivity analysis of Table 6 above shows that parameter estimates are otherwise quite robust. The effect of the RR increase, the effect of PBD extensions, and the additive effect of simultaneous RR and PBD changes are quite robust to variations in the control groups used to identify the treatment effect. In particular, the estimated treatment effect of an RR remains unchanged when the control group is confined to workers younger than 40 who are slightly above the earnings threshold. The effects of a given PBD extension and of a simultaneous RR and PBD change remain unchanged when the control group either consists

28. Note that the early retirement explanation is supported by the facts. Given an elapsed duration of unemployment of at least 30 weeks, the probability to exit to an early retirement scheme (some form of disability insurance) is about 20% among individuals aged 40–49 and is about 60% for individuals aged 50–54.

only of too unexperienced workers (of the same age) or of too young workers (with the same work experience). This suggests that, apart from the much larger reactions to PBD extensions and joint RR and PBD changes for older workers mentioned above, the estimated effects are quite homogenous across the various groups and largely in line with theoretical predictions.

5.5. *Disincentive effects*

This final subsection compares the disincentive effects of changes in PBD and changes in RR. Since such a comparison needs a common denominator, we study the effects in terms of the total increase in benefit payments. Suppose, for instance, that policy makers increase the RR. This raises the total amount of benefit payments for two reasons. First, the RR increase will raise benefit payments even if individuals do not change their behaviour, simply because higher benefits have to be paid for the same number of days individuals spend in unemployment. Second, the RR increase will induce individuals to stay longer in unemployment, thus raising benefit payments further. An intuitively appealing measure splits up the total costs into the fraction of direct costs (without behavioural changes) and the fraction of indirect costs resulting from individuals' change in behaviour.

Specifically, we estimate direct cost due to the change in the UI system from an increase in the benefits level from b_{0i} to b_{1i} and/or an increase in PBD from E_{0i} to E_{1i} as follows. The expected duration of unemployment without treatment is $\int_0^{E_{0i}} S_{0i}(t)dt$, where $S_{0i}(t) \equiv S_0(t \mid x_i)$ is the counterfactual survivor function without the policy change.[29] The direct costs due to the change in the UI system are given by $b_{1i} \int_0^{E_{1i}} S_{0i}(t)dt - b_{0i} \int_0^{E_{0i}} S_{0i}(t)dt$. The indirect costs due to the change in the behaviour of job seekers are given by $b_{1i} \int_0^{E_{1i}} S_{1i}(t)dt - b_{1i} \int_0^{E_{1i}} S_{0i}(t)dt$ where $S_{1i}(t) \equiv S_1(t \mid x_i)$ is the factual survivor function for treated individuals and $\int_0^{E_{1i}} S_{1i}(t)dt$ is the expected duration of unemployment in the new system that takes behaviour changes into account. Denote by D_S the fraction in total additional costs due to the direct effect of the change in the system and by D_B the fraction due to the indirect effect from changes in individuals' behaviour, with $D_\mathrm{S} + D_\mathrm{B} = 1$. Formally, D_S and D_B are given by

$$D_\mathrm{S} = \frac{E_{x_i}\left[b_{1i} \int_0^{E_{1i}} S_{0i}(t)dt - b_{0i} \int_0^{E_{0i}} S_{0i}(t)dt\right]}{E_{x_i}\left[b_{1i} \int_0^{E_{1i}} S_{1i}(t)dt - b_{0i} \int_0^{E_{0i}} S_{0i}(t)dt\right]}$$

and

$$D_\mathrm{B} = \frac{E_{x_i}\left[b_{1i} \int_0^{E_{1i}} (S_{1i} - S_{0i})(t)dt\right]}{E_{x_i}\left[b_{1i} \int_0^{E_{1i}} S_{1i}(t)dt - b_{0i} \int_0^{E_{0i}} S_{0i}(t)dt\right]}. \tag{4}$$

Table 8 shows the estimated split of the total increase in benefit payments into D_S (first column) and D_B (second column) of the various changes for the respective treated groups. For the PBD change from 30 to 39 weeks, more than 80% of additional total costs are direct. Obviously, these costs result only from spells that lasted longer than 30 weeks already under the old system and the behavioural effects, less than 20% of total additional costs, reflect an increase in the length of all spells. This comparably small number reflects the relatively small behavioural changes we have estimated above. For the RR change, the direct cost component is even larger and amounts to about 90%. This is not surprising, because with an increase in RR all eligible spells are affected.

29. Note that $b = Bw$, where B is the gross RR and w is gross weekly income. Moreover, this *ex ante* expected cost measure does not account for unemployment assistance payments, which are available after unemployment benefits have run out. The data do not contain information on unemployment assistance. Also, results are qualitatively not sensitive to using imputed unemployment assistance payments.

TABLE 8

Simulated effects on benefit payments

	Percentage of total change in cost due to change		
	UI system	Behaviour of job seekers	Total
Change to one parameter			
PBD 30–39 weeks	82·0	18·0	100·0
PBD 30–52 weeks	54·8	45·2	100·0
RR increase	89·9	10·1	100·0
Change to two parameters			
PBD 30–39 and RR increase	84·3	15·7	100·0
PBD 30–52 and RR increase	51·1	48·9	100·0

Notes: Based on population receiving the treatment in the period after the policy change. RR, related rate; PBD, potential benefits duration.
Source: Own calculations, based on Austrian Social Security Data.

Results are different for the groups eligible to the strong increase in PBD from 30 to 52 weeks. In this group behavioural effects are much stronger and of roughly equal size as the direct effects. In other words, individuals react strongly to the increase in benefit duration, and these behavioural changes are the main factor driving the total additional costs of the policy change. Differences in replacement ratios are here of lesser importance. While a somewhat higher fraction of additional total costs has to be attributed to behavioural changes, the split is of a similar order of magnitude for changes in two parameters as for the corresponding change in only one parameter.

6. CONCLUSIONS

This paper addresses the issue of how financial incentives embedded in the unemployment insurance system affect the duration of unemployment. This issue is important for a number of reasons. On the one hand, the years since the turn of the century have witnessed important reforms to unemployment insurance in many (particularly European) countries. On the other hand, studies of how simultaneous changes of UI parameters affect the unemployment exit rate are lacking. Hence, it is difficult to compare the relative effects of changes in the policy instruments.

This paper relies on a change to unemployment insurance in the late 1980's in Austria. This reform leads both to extensions of the potential duration (PBD) of regular unemployment benefits for a first group of individuals; to an increase in the earnings replacement rate (RR) for a second group of individuals; to both extended PBD and higher RR for a third group of individuals; and to no changes for a final group of individuals. This means that it is not only possible to study the relative magnitudes of two key parameters determining the generosity of unemployment compensation, but also to analyse whether there are excess effects from combined changes in PBD and RR.

Age, experience, and earnings determine the treatment group. Earnings determine whether the RR goes up—only workers with low earnings get an increase in RR. Age and experience determine whether the PBD goes up. Only workers from age 40 onwards with a high working experience get an increase in PBD, where the size of the increase depends on the age of the individual. Individuals between 40 and 50 get an increase in PBD from 30 to 39 weeks; individuals from age 50 onwards get an increase from 30 to 52 weeks. Furthermore, the changes in the UI system are introduced in a period of employment growth so that any increases in unemployment duration were not related to labour-market conditions. The Austrian case has clear elements of a "natural experiment".

Although it is clear that there is heterogeneity in the treated populations we find that the effects are quite robust. In particular, we find that both the estimated effects are invariant to variations in the control groups used to identify the treatment effect. This suggests that the various groups do not react systematically different to financial incentives. However, there are two exceptions. First, the relative effect of a PBD extension is smaller for prime-age workers. For prime-age workers unemployment duration increases by 0·35 days per week of extra PBD, while for older workers it increases by 0·70 days per week of extra PBD. Second, the additive effect of a simultaneous change in PBD and RR is larger for older workers. The reason why older workers react differently is because they face different labour-market conditions and a different institutional setting. They get fewer job offers because employers preferably fill their vacancies by hiring young or prime-age workers. Furthermore, they are close to retirement, and the closer they get to retirement the less incentive they have to search for a job because the value of finding a job is reduced. Both labour-market conditions and institutional setting may cause older workers to react stronger to disincentives. In short, despite the apparent heterogeneity in response workers react to incentives in line with several predictions from theory.

The changes in incentives are interesting from a scientific point of view to investigate to what extent results are in line with theoretical predictions. Theory provides predictions about the pattern of changes in search intensity over the duration of unemployment and about the way separate changes in incentives interact. Nevertheless, theory only provides a prediction about the sign of the effects and does not give an indication about how large these effects should be. With respect to the stronger reaction of older workers there is not much that can be derived from theory to explain the exact difference in response. However, it does give some idea about potential causes of the differences: labour-market conditions and institutional settings.

From a policy point of view our study is interesting as well. An innovative part of our analysis concerns the way we use our parameter estimates to split up the total costs to unemployment insurance funds into costs due to changes in the unemployment insurance system with unchanged behaviour and costs due to behavioural responses of unemployed workers. For the increase in RR we find that additional costs due to behavioural responses are modest, of the order of 10% of total costs. This is different for increases in PBD where the behavioural component amounts to about 20% in case of the PBD increase from 30 to 39 weeks and to close to 50% in case of the PBD increase from 30 to 52 weeks. Taking these results at face value, we conclude that PBD is a more effective policy parameter than RR to affect individual's job search behaviour and unemployment durations. From this we derive two simple policy recommendations. First, if governments change various incentives for unemployed workers they should be aware of behavioural effects, and above all they should take behavioural effects related to interactions between the changes in incentives and interaction between changes in incentives and institutional settings into account. Second, if they want to influence incentives the potential duration of unemployment is a more effective policy tool than the level of the unemployment benefits.

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