Taking the most out of unemployment insurance under nonstationary job search: unemployment duration, reemployment wages and job tenure*

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Abstract

This paper puts together the non-distortionary liquidity effect of unemployment insurance and job match quality. It takes advantage of sizeable (6 months) differences on benefits entitlements based only on the age of beneficiaries to set up a sharp regression discontinuity design. We identify a big impact on subsidized unemployment duration and a small impact on wages and tenure on the job that follows the unemployment spell. Wage gains are heterogeneous and concentrated on individuals at the bottom of the pre-unemployment income distribution. The impact on tenure is always rather small. The non-distortionary nature of the liquidity effect reduces the pressure on low income workers to accept lower productivity jobs.

Keywords: Unemployment insurance; Subsidized unemployment duration; Reemployment wages; Job tenure; Liquidity effect; Regression discontinuity.

JEL Codes: J38, J64, J65, J68.

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1 Introduction

This paper analyzes the impact of unemployment insurance (UI) generosity on subsidized unemployment duration and post-unemployment initial wages and tenure. We do this in the context of a nonstationary job search model (Mortensen (1986) and van den Berg (1990)), exploring the existence of constrained and unconstrained individuals, which generates a non-distortionary liquidity effect of UI (Chetty 2008). We put together the income effect and the potential gains to match quality from UI. If the liquidity effect is important, the disincentive of UI created through the substitution effect is mitigated, and UI becomes less distortionary than previously thought. The non-distortionary nature of the liquidity effect, reducing the pressure on low income workers to accept bad quality matches and allowing them to wait for a better match, should be associated with a greater (and positive) impact on post-unemployment outcomes, namely reemployment wages and tenure, for these workers.

The main novelty of this paper is the study of the importance of the liquidity effect in shaping job search duration and, more importantly, job search outcomes. This approach to the subject contributes directly to the policy debate on the distortionary impact of the UI system. As in Chetty (2008), we argue that the economic importance of the liquidity effect makes the UI system less distortionary than previously thought and that its reform should target low income individuals, where liquidity constraints are more alleviated.

Our exercise takes advantage of the rules of the Portuguese UI system to identify the causal effect of UI entitlement generosity on job search duration and reemployment match quality. The UI entitlement period is determined by the worker's age at the moment of unemployment, and this generates a sharp discontinuity at two age levels: 30 and 40 years. This makes it extremely appealing to use the regression discontinuity approach to identify and estimate the impact of UI on labor market outcomes. Regression discontinuity has been applied recently by Card, Chetty and Weber (2007), Lalive (2007) and Lalive (2008) to study the impact of UI and severance payments on unemployment duration and match quality in Austria and by Lemieux and Milligan (2008) in the context of other social assistance programs for Canada. Furthermore, the fact that these are prime-age individuals is quite important when evaluating labor market outcomes, specially under nonstationary environments. Also, the availability of multiple thresholds allows for further testing the assumptions of regression discontinuity, which may not hold at every age, specifically within a nonstationary search environment. Finally, there are sizeable differences, 6 months, in entitlement periods across the age thresholds. This is important, as it allows the estimation of meaningful economic impacts of UI generosity.

We extend this literature in one important way by studying whether the provision of more generous UI allows the unemployed to find better matches, and how the impact differs depending on the degree of the workers' liquidity constraints. The latter captures the impact of generosity via an liquidity effect. The existence of an liquidity effect of UI can be seen as the result of the nonstationarity in the job search environment (Mortensen (1986) and van den Berg (1990)). Nonstationarity stems from the fact that constrained and unconstrained individuals react differently to the degree of UI generosity, as they face diverse search environments when setting their optimal job search strategies. For example, in Mortensen (1986) a decreasing reservation wage, the most typical outcome of nonstationary environments, is the result of the limited capacity of individuals to self-finance their out-of-pocket job search expenses while unemployed. In van den Berg (1990), other factors leading to nonstationarity, for example, the time-dependence of the wage offers distribution and job offers arrival rate, reinforce the impact of nonstationarity on unemployed individuals' decisions.

We use Portuguese Social Security administrative data covering *all* subsidized unemployment spells that started during the period from July 1999 through December 2002. Several characteristics of this dataset make it particularly suitable for our study, specifically, the availability of information on (i) the salary and starting date of the first job following unemployment; (ii) the duration of the unemployment spell; (iii) the wage earned prior to entering unemployment; and (iv) the duration of the first job following unemployment.

Our results show a big impact of the benefit entitlement period on subsidized unemployment duration. The estimates point to an increase on subsidized unemployment of approximately 40 days at both discontinuity points as the result of the extra 180 days of entitlement. This extra search period translates into a rather small impact on reemployment wages.

Individuals with pre-unemployment income below the median extend their unemployment spells the most. This is in accordance with what we might expect from the liquidity effect of UI. This result is even clearer if we split the sample into the four quartiles of pre-unemployment income. The second quartile is the one that reacts more and the first and fourth the least. The weaker reaction of individuals in the first quartile can be explained by the strong effect of the nonstationarity of the search environment arising from a declining arrival rate of job offers and/or a deteriorating wage offers distribution faced by the very low income individuals.

The impact on reemployment wages is also rather heterogeneous. We estimate a significant wage loss for those above the median income and a gain, although smaller in size, for those below the median income. This gain is concentrated in the first quartile of pre-unemployment income, specially for those aged 40. In the same vein, the losses are larger for those previously at the fourth quartile.

For the most part, the impact of more generous UI entitlements on job tenure is negligible. Most point estimates are non-significant and, we only find a positive and statistically significant impact for older individuals, whose previous wages where in the top quartile. We speculate that the negative impact on wages for this group may have been compensated with longer job tenures.

From an international perspective, the Portuguese case is interesting because its rigid labor market institutions translate, as in other European countries, into high unemployment duration, and a substantial polarization of the labor market, with the unprotected fraction of the market experiencing significant flows between jobs (see Blanchard and Landier (2002) for a discussion on the impact of polarization in the labor market and Centeno, Machado and Novo (2008) for evidence on the Portuguese economy). Our results show the important role of individuals' liquidity constraints in shaping their reaction to UI generosity. The existence of an liquidity effect makes the global impact of UI less distortionary than previously thought. The impact on duration, together with the fact that the more constrained gain the most in terms of reemployment wages, gives better prospects to UI policies targeting these individuals. The general equilibrium effects obtained in Marimon and Zilibotti (1999) and Acemoglu and Shimer (2000) can be particularly important to this specific group. The reduction in the pressure to accept bad quality jobs is reduced through the liquidity effect, generating better matches.

2 Literature: Theory and evidence

The main theoretical results that motivate the empirical exercise in this paper are derived from nonstationary job search models, such as Mortensen (1986) and van den Berg (1990). These models assume that the unemployed is entitled to UI benefits for a limited period of time. We aim at studying how transitions out of subsidized unemployment are affected (possibly in a heterogeneous way) by this fixed benefit duration.

In the model, the change in UI level that occurs when benefits expire has an impact on the search behavior of the unemployed that can be interpreted as a consequence of liquidity constraints faced by the unemployed. The ability of constrained individuals to finance the out-of-pocket cost of search improves with UI recipiency, allowing them to smooth consumption during this period. UI generates a liquidity effect, similar to the one described in Chetty (2008).

The unemployed search behavior is determined by the comparison of the value of being unemployed and the value of having a job. The key determinants of the value of being unemployed are the level of UI benefits net of search costs, the labor market prospects (e.g., the probability of receiving a given wage offer) and the risk of being jobless when UI expires. The worker sets a reservation wage that equates optimally the job search marginal benefits and costs. For instance, while a higher reservation wage increases the expected re-employment wage, it also increases the expected duration of unemployment and the probability of exhausting the UI benefits before getting a job. Thus, the model predicts that at the start of the unemployment spell the reservation wage is high because the probability of finding a job before UI expires is large. However, the limited duration of UI benefits generates a nonstationary search environment, and, consequently, a declining reservation wage.[1]

The model predicts that longer entitlement periods will lead to longer spells of subsidized unemployment. However, an extension of the entitlement period has only a small impact on the reservation wage at the beginning of the unemployment spell. As discussed above, the reason is that early on the risk of not getting a job prior to UI exhaustion is not significantly affected by the extension. As the unemployment spell progresses a longer entitlement period becomes more important and has an increasing impact on the optimal reservation wage. The difference in the optimal response before and after the extension will be greatest when unemployment duration reaches the entitlement period prevailing prior to the reform.

The model also predicts an heterogeneous impact of UI generosity associated with the individuals' differences in the probability of becoming long-term unemployed. Clearly, the relative magnitude of the relationship between UI and unemployment duration depends on the parameters of the nonstationary model and the labor market position of specific groups of workers. This is particularly important in UI regimes in which the extension of benefits accrues only after a long period of subsidized unemployment. In such cases, it is expected that the nonstationarity of the job search environment will affect one group of workers more than another.

[1] In van den Berg (1990) other exogenous variables are considered as sources of nonstationary, namely the arrival rate of job offers and the wage offer distribution. The model results discussed herein are only reinforced if these variables are added.

The extra time to search for a new job can also have a positive impact on post-unemployment outcomes, improving match quality. The impact of the UI system on productivity and job mismatch has recently been examined in several theoretical papers. Marimon and Zilibotti (1999) present a model of the role of UI on mismatch and unemployment and show the positive impact of the UI system on the reduction of job mismatch. In a related paper, Acemoglu and Shimer (2000) analyze the productivity gains from more generous UI systems. Considering risk-averse workers, they show that UI increases labor productivity by encouraging both workers to seek higher productivity jobs and firms to create such jobs. In their setting, the UI is more than a search subsidy, and affects the type of jobs that workers look for and accept.

The impact of UI on match quality remains, nonetheless, an empirical issue. There are only a limited number of studies addressing the impact of UI on post-unemployment outcomes, and they have concentrated almost exclusively on the wage dimension of match quality. Belzil (2001) looks at job duration by exploring a reduction in the initial entitlement period rule in Canada to study the impact of UI duration on subsequent job duration for young individuals, and reports a weak but positive impact. Centeno (2004) and Centeno and Novo (2006) look at the US system, using UI variation across states with data from the NLSY, and find evidence that more generous UI increases the tenure of reemployment and that this impact is stronger at longer durations. They also show a positive impact on post-unemployment wage distribution. More recently, McCall and Chi (2008) using also data from the NLSY, find a positive impact of UI on reemployment wages.

Recently, a number of papers considered the impact of UI on post-unemployment outcomes using data for European countries. Lalive (2008) and Lalive (2007) use Austrian data from an extension of UI benefits and report a significant impact on unemployment duration but no impact on wages. Similar results are obtained in the studies by Fitzenberger and Wilke (2007) for Germany (in the context of a nonstationary model) and van Ours and Vodopivec (2006) for Slovenia. Card et al. (2007) also uses data from Austria and finds some impact of severance payments on reemployment job tenure, but no impact on wages.

3 The unemployment insurance system and identification

One peculiar feature of the Portuguese UI system is the definition of the entitlement period. For individuals younger than 45 years at the moment of entering unemployment, the entitlement period is fully determined by the individuals' age at the beginning of the unemployment spell. The legislation foresees four age-groups with different entitlement periods, ranging from 12 months to a maximum of 38 months (see Table 1).

[TABLE 1 HERE (see page 27)]

The characteristics of the system result in two sharp discontinuities at ages 30 and 40, which we will explore to identify the impact of UI with a regression discontinuity design. The third discontinuity at age 45 faces one major practical identification problem within this approach. For the younger cohorts, age is the forcing variable, i.e., age fully determines the entitlement period. This is not the case for individuals aged 45 or more; these individuals have an entitlement period of 30 months plus 2 months per each set of 5 years with social security contributions. However, in the data, we do not have the complete record of social contributions, nor the information on the actual entitlement period awarded to each individual. Therefore, we chose to analyze only the discontinuities at ages 30 and 40, where age is the only forcing variable.

Additionally, the Portuguese UI system has two distinct characteristics that are useful for the economic analysis of its impact on labor market outcomes. First, the variation in the entitlement periods is rather sizeable, with 6 additional months at each discontinuity point. This enhances the conditions for identification of an impact. Furthermore, because there are two discontinuity points, it is possible to simultaneously identify different impacts at different locations of the age distribution within the same setting. This is important given the time and duration dependence introduced by the nonstationary search environment, which impacts differently at these age locations.

In terms of the financial generosity of the Portuguese UI system, benefits are set as a percentage of the 12-month average of the previous wages. Figure 1 illustrates the financial generosity of the system expressed in terms of the gross replacement rate (GRR). In international terms, the financial generosity is similar to other countries, particularly European.

[FIGURE 1 HERE (see page 24)]

4 Data

Our study uses administrative data collected by the Portuguese government's agency *Instituto de Informática* (II) of the Portuguese Social Security bureau. The dataset registered *all* subsidized unemployment spells initiated between July 1999 and the end of 2002, some of which ended with claimers reemployed as salaried workers. Our study concentrates on reemployed workers, aged 25 to 44 years, resulting in a sample size of 18,457 observations. The dataset contains very detailed and reliable information on the type, amount and duration of benefits, the previous income. The socio-demographic variables available are limited to gender, age, and place of residence. The first reemployment wage and starting date of the job are also observed; eventually, the last day in the new job is also observed, otherwise the job tenure is considered right-censored.

We are, however, able to follow individuals until their benefits expired or they exited the system voluntarily. Thus, we consider complete spells of subsidized unemployment, which correspond to a single-cycle/flow sampling scheme as defined in Lancaster (1992). Table 2 contains summary statistics of the key variables by age group.

[TABLE 2 HERE (see page 27)]

The average spell of subsidized unemployment increases with age: 188, 255, 327 days, respectively for the 25-29, 30-39 and 40-45 age groups. Similarly, average pre-unemployment wages are also an increasing function of age, although the differences do not exceed 46 euros (at 1999 prices), a value that corresponds to approximately 8 percent of the average wage in the total sample. The average GRR is similar across the age groups at 67 percent. Although not reported in the table, the median GRR is 65 percent. Reemployment wages are on average smaller than pre-unemployment wages for all age groups. The percentage of female UI claimers decreases with age, which also conforms to the age-profile of women's participation rate.[2]

5 Graphical analysis

The nature of the RD design and the sizeable change in the entitlement period suggest that the impact of the treatment should be visually observable in a simple plot of the average outcome of the variable of interest against the forcing variable (in our case, age) at each discontinuity point. We start with such an informal assessment of the impact of longer UI benefits at the identified discontinuity points, 30 and 40 years old on both outcomes of interest. Moreover, to

[2]The data contains also information about the region, UI claiming date and reemployment date, which we do not summarize here, but use in the regression setting.

assess the empirical validity of the identification strategy, we will also plot covariates against the forcing variable and the density of the forcing variable itself.

5.1 Subsidized unemployment duration, reemployment wages, and job tenure by age

Figures 2 summarizes the essence of the identification strategy followed in this paper. On the left-hand side panel, which plots average subsidized unemployment duration (in days) by age, there are two clear discontinuity points at ages 30 and 40. These coincide with the 6 months discontinuities introduced by the legislator in the age-based UI entitlement periods: from 12 to 18 months at age 30 and 18 to 24 months for older individuals. To these increases in UI generosity, individuals respond unequivocally by extending their subsidized unemployment spells; there are significant discontinuities at ages 30 and 40, which suggests that adding a sizeable 6 months to potential benefits results in longer average subsidized unemployment duration.

[FIGURE 2 HERE (see page 24)]

The right-hand side of Figure 2 presents evidence of the impact of longer UI entitlement periods on post-unemployment job match quality, indexed by the log reemployment wages. Overall, this figure suggests a much smaller impact than for the case of duration. Indeed, superimposed regression lines at the discontinuity points suggest a tiny negative impact at the first discontinuity (29/30 years), and a small impact for older individuals (39/40 years). In the following section, these results will be further exploited in an appropriate econometric setting, along with alternative regression specifications.

[FIGURE 3 HERE (see page 25)]

The impact on post-unemployment job quality indexed by the duration of the reemployment jobs is depicted in Figure 3. Contrarily to the subsidized unemployment spells, where all observations are complete, the measure of job tenure includes some right-censored observations. Here, we choose to plot in the left panel only the complete job tenures, while in the right panel there are only incomplete job tenures. Both plots suggest that the impact of UI on the duration of reemployment is negligible. Interestingly, the picture shows that complete job tenure increases with age, but it decreases for incomplete spells. This find is consistent with the higher hiring and separation rates found for younger workers in Centeno et al. (2008) for the Portuguese economy in the 2001-2006 period.

5.2 Covariates by age

A crucial identification assumption in the regression discontinuity design is that there are no other discontinuities in the conditioning variables that may confound the identification of the impact with the discontinuity of treatment based on the forcing variable. As already mentioned, the set of available conditioning variables is limited. Nonetheless, it contains an important labor market variable, namely, (average) pre-unemployment wages, which ought to capture differences in the characteristics of the individuals. Thus, to assess the possibility that there are other discontinuities in the data, we plot the average pre-unemployment wages against age, splitting the data by gender (Figure 4).

[FIGURE 4 HERE (see page 25)]

The wage profiles depicted in Figure 4 follow the general pattern observed in the Portuguese economy. As expected, males have larger average wages than female. In terms of possible discontinuities, there are no evident unexpected increases or drops in the age-wage profile, and in particular at the discontinuity points. In fact, at the discontinuity points under analysis, the differences are statistically (and economically) insignificant. Nonetheless, the women's profile is somewhat less 'smooth' than the men's. However, in the regression setup any such local differences may be addressed by appropriately conditioning the estimates on the impact of such characteristics (Lee 2008).

5.3 Density of the forcing variable

An alternative possibility to assess the manipulation of the pool of treatment participants is to identify discontinuities in the density of the forcing variable. In the present case, due to more attractive entitlement periods at older ages, it is possible that individuals aged 29 and 39 selected themselves into the older groups by postponing unemployment entry. Figure 5 shows, however, that there are no such visible discontinuities; there are no sings of claims concentration at ages 30 and 40, where UI entitlement periods are more generous.

[FIGURE 5 HERE (see page 26)]

The set of graphical evidence presented hitherto suggests that longer entitlement periods increase significantly the duration of subsidized unemployment spells, and have only minor impacts on reemployment wages. Furthermore, the battery of additional plots did not expose any evident violation of the standard identification hypotheses underlying the regression discontinuity design.

6 Econometric results

Our identification strategy relies on the discontinuity points of UI entitlement periods observed at ages 30 and 40. To explore this characteristic in a technically appropriate fashion, we rely on the regression discontinuity design, which considers a model of the form:

$$y_i = \alpha + \tau\ D_i + \xi(a) + \epsilon_i, \tag{1}$$

where y_i is the outcome variable for individual i, and the effect of the forcing variable a (age in our case) on the dependent variable is captured by the function $\xi(\cdot)$. D_i is a dichotomous treatment variable, assuming value 1 if the individual i belongs to the group with longer entitlement period. In particular,

$$D_i = \begin{cases} 1, \text{ if } 30 \leq a \leq 39 \\ 0, \text{ if } 25 \leq a \leq 29 \end{cases} \qquad \text{or} \qquad D_i = \begin{cases} 1, \text{ if } 40 \leq a \leq 44 \\ 0, \text{ if } 30 \leq a \leq 39, \end{cases}$$

depending on the discontinuity threshold that is under analysis. In both cases, we have a denominated 'sharp' regression discontinuity design (Hahn, Todd and van der Klaauw 2001). The impact of participation in the treatment is given directly by the estimate of the parameter τ at the point where the treatment variable switches from 0 to 1.

The key underlying assumption of the regression discontinuity design is the smoothness (continuity) of the ξ function. Following a Mincerian formulation of wages formation (Mincer 1974), there are reasons to believe that reemployment wages are a smooth function of age; this was shown to be the case of pre-unemployment wages (see Figure 4). The relatively smooth age-profile of unemployment duration (increasing with age) is also established in the literature (van den Berg, van Lomwel and van Ours 2003). In practice, we will consider alternative polynomial functions of age to check for the sensitivity of the estimates to alternative specifications.

The estimation of the model uses local linear regression methods for their superior statistical efficiency in these settings (Hahn et al. 2001, Imbens and Lemieux 2008). If we assume a linear specification of ξ, then the local linear regression parameter estimates are obtained by solving the following optimization problem:

$$\min_{\alpha,\beta,\tau,\gamma} \sum_{i=1}^{N} (y_i - \alpha - \tau\ D_i - \beta(a_i - c) - \gamma(a_i - c)D_i)^2\, \mathbf{1}\{c - h \le a_i \le c + h\}, \tag{2}$$

where c is the discontinuity threshold (30 or 40 years). The specification above considers a rectangular kernel, represented by the indicator function $\mathbf{1}\{\cdot\}$ and the sensitivity of the estimates is assessed by considering alternative values to the kernel bandwidth parameter, h. Imbens and Lemieux (2008) argue that more sophisticated kernels rarely make a significant difference, and that if estimates vary considerably with the choice of the kernel, then they are sensitive to the bandwidth choice in the first place. We will test the results for different values of the bandwidth.

The estimation of the impact of UI on the duration of post-unemployment jobs poses additional difficulties due to the fact that some jobs are on going. To handle job tenure right-censored observations, we use a proportional hazard model. The hazard function represents the probability of job termination conditional on the duration lasting up to t, formally:

$$\lambda(t|x) = \lambda_0(t) \exp(x\zeta),$$

where the x vector includes individual characteristics. In this equation $\lambda_0(t)$ is the baseline hazard rate at time t for the covariate vector $x = 0$. The estimation of ζ is based on the standard Cox (1972) semi-parametric estimator, accounting appropriately for the censored observations. The associated baseline hazard estimates, λ_0, are obtained nonparametrically with the product limit estimator (Kalbfleisch and Prentice 2002). The model results reported throughout the text are based on the specific formulation:

$$\lambda(t|X) = \lambda_0(t) \exp\{\alpha + \tau\ D_i + \beta(a_i - c) + \gamma(a_i - c)D_i + x\zeta\}\, \mathbf{1}\{c - h \le a_i \le c + h\}, \tag{3}$$

where t is the duration of the reemployment job (in months); the remaining variables and coefficients are as defined above.

6.1 Subsidized unemployment duration, reemployment wages, and job tenure

Table 3 summarizes the regression discontinuity design estimates for the two thresholds (30- and 40-year-old) and for the three outcome variables (unemployment duration, reemployment wages, and job tenure).

Confirming the visual inspections of the previous section, the results in the left part of Table 3 show a sizeable impact on duration for the younger cohort. Indeed, the estimates point to an average increase on subsidized unemployment duration varying between 37 and 45 days, estimated quite precisely; individuals aged 30 take up one-quarter of the additional entitlement period. This range of estimates corresponds to different functional form specifications (linear (column 2) and quadratic (column 3) in age; and with control variables (column 5)), but also to different kernel bandwidth choices (column 4).[3] The stability of the point estimates, in columns (2) to (5), together with the previous graphical evidence, suggests that the current setting does not violate the standard assumption of the RD design. The same conclusion holds for the older individuals.[4]

[TABLE 3 HERE (see page 28)]

For the universe of subsidized unemployed, the increase in the entitlement period and the additional duration of subsidized unemployment did not bring about, on average, higher reemployment wages. The impact measured in the log difference between pre-unemployment and reemployment wages varies between -1.0 and 2.4 percent at the 30 years discontinuity point; none of the estimates is statistically significant. Similarly, the impact on the tenure of the reemployment job is always non-significant.[5] Notice, however, that the pool of unemployed is rather heterogeneous and there is, for example, a wide range of pre-unemployment income, leaving open the possibility of differentiated impacts along the distribution of income.

For the discontinuity at the older age (40 years), the overall assessment is similar to the younger one. The range of estimates is almost the same: between 36 and 44 days for the impact on subsidized unemployment duration and -1.5 and 1.9 percent for the impact on reemployment wages. Notice, however, that the somewhat erratic values for the outcomes at age 44 (Figure 2) seem to influence the simple linear adjustment (column (7)), but once those values are excluded (through a tighter bandwidth) or the functional form adjusted (columns (8) to (10)), the estimates range only between 36 and 41 days. The impact on job tenure is slightly negative, although not always statistically significant.

[3]The set of control variables included are listed in the notes to Table 3. We experimented with other bandwidths, but the results resemble those reported in the Table.

[4]Addison and Portugal (2007) study the impact of the Portuguese UI rules on transitions out of unemployment into employment, using survey data. They find higher rates of job finding for individuals with shorter entitlement periods. They do not study the impact on reemployment wages.

[5]In this version of the paper, we consider the tenure results preliminary; please do not quote.

The addition of control variables does not have any significant impact on the estimates (compare columns (2) and (5), and columns (7) and (10)). Therefore, for reasons of statistical efficiency (Lee 2008), we will use the specification with control variables in the following analysis.

It seems adequate to conclude in favor of a meaningful economic impact of longer entitlement periods: unequivocal longer subsidized unemployment spells (approximately $^1/_4$ of the additional 6 months is taken up) and small impact on job match quality, as measured by reemployment wages and job tenure. The evidence on the impact on reemployment wages and job tenure is somewhat less compelling, which also conforms to the apparently less unanimous evidence found in the literature, a point addressed below in the context of the UI liquidity effect.

6.2 Falsification test

The graphical analysis has already provided compelling evidence that our setting does not violate standard regression discontinuity identification hypotheses. Nonetheless, we can be reassured that this is the case by means of a falsification test.

For this purpose, we take individuals aged 35 to 39 and presume that they were entitled to longer UI benefits than those aged 30 to 34. In the regression discontinuity design, this amounts to reestimate equation (2) with D_i defined as

$$D_i = \begin{cases} 1, \text{ if } 35 \leq a \leq 39 \\ 0, \text{ if } 30 \leq a \leq 34 \end{cases}$$

The results are presented in Table 4. As expected, they reveal non-significant impacts on both duration and reemployment wages.

[TABLE 4 HERE (see page 29)]

Our choice of the pseudo-discontinuity point was intentional, because before July, 1999, there was indeed a discontinuity point at ages 34 and 35, of the type described in the previous sections; individuals aged 35-39 were entitled to 90 more days of unemployment benefits than those aged 30-34. Centeno and Novo (2007), using the quantile treatment effects and the Kaplan-Meyer estimators, show that there were indeed differences in the survival rates of these two groups. This reinforces our causal interpretation of the previous results.

6.3 Gender and replacement rate effects

The above analysis collapses the impact into a simple average, which may be hiding heterogeneous behaviors within the pool of unemployed. In this section, we address two sources of heterogeneity. First, we split the sample by gender; Figure 4 had already depicted wage differences between genders that may reflect themselves in differences in the responses to the same UI incentive. Second, given the wide range of pre-unemployment income, there are also rather different GRRs (see Figure 1) that could lead to different sensitivities to UI generosity. Indeed, as pointed out by Fitzenberger and Wilke (2007), large GRRs are associated with larger labor supply disincentives. To control for the impact of such types of disincentives, we take advantage of the way in which the financial generosity is defined in the Portuguese UI system and reestimate the impact of longer entitlement periods for a subsample of unemployed with GRR of 65 percent (the flat portion of the curve in Figure 1).[6] In practical terms, this means that we keep in our sample individuals earning between 1.5 minimum wages and 4.5 minimum wages per month. Given the universal coverage of our sample, this still leaves us with a large sample. The results are presented in Table 5.

[TABLE 5 HERE (see page 29)]

The results by gender reveal that men respond more to the additional entitlement period than women, with the larger gender difference observed at the older discontinuity point. At the 30 years discontinuity threshold, older men spend an additional 52 days unemployed, while women increase unemployment duration by only 36 days. The difference between men aged 39 and 40 years old is 46 days of additional unemployment, while 40-year-old women spend 24 additional days in subsidized unemployment. The gender differences in terms of reemployment wages are significant only for younger women, who may lose up to 3.5 percent relative to 29-year-old women. For the remaining cases, the impacts are statistically non-different from zero.

The more homogeneous sample obtained by restricting the GRR at 65 percent yields smaller impacts on unemployment duration at both the 30-year-old cutoff (52 vs 44 days), and at the 40-year-old cutoff (44 vs 36 days).[7] The difference between the two samples, of about one-week,

[6] In the data, GRRs are not computed exactly to 65 percent, therefore, we keep the unemployed with GRR in the 63 to 67 percent interval.

[7] We take as benchmark the full sample results reported in columns (5) and (10) of Table 3, which have the same functional form and bandwidth choice.

can be explained by the behavior of individuals at the extremes of the GRRs distribution. For individuals with previous wages above 4.5 minimum wages (with GRRs below 63 percent), the relative cost of unemployment is higher, resulting in smaller increases in unemployment duration. For individuals at the bottom of the income distribution (GRRs between 67 and 100 percent), their large GRRs may be associated with larger labor supply disincentives (Fitzenberger and Wilke 2007). However, these individuals may also face larger liquidity constraints, which limit their capacity to extend unemployment duration, resulting in smaller impacts (an effect explored in the following section).

In terms of reemployment wages, the difference between the two samples is much smaller. The estimates are -2.0 and -1.0 percent for the GRRs restricted subsample, while in the full sample the estimates were -1.0 and -1.2 percent, respectively, at the younger and older discontinuity points.

6.4 Liquidity effect

We look now into the possibility that UI generates a non-distortionary liquidity effect. The underlying mechanism has been previously laid out in the work of Mortensen (1986) and van den Berg (1990) on nonstationary job search models. Both models explore the variation in the degree of liquidity constraints faced by the unemployed, predicting that more constrained individuals, those less able to self-finance their job search process, will react more to increased UI generosity.

The identification of the liquidity effect rests on individual differences in the degrees of liquidity constraints. However, the 'constrained' status is a latent variable. As such, it is not feasible to classify individuals, directly from the data, into distinct groups of liquidity constraints. The approach followed to identify these distinct groups was to split the sample according to the 12-month average of pre-unemployment wages, which serves as an index for the distribution of liquidity constraints. We resort to wages because our data lacks the information on asset holdings for the unemployed, a more direct measure of their degree of liquidity constraints.

Notwithstanding these drawbacks, the work by Ziliak (2003) shows that wages are the leading factor driving differences between poor and rich households in terms of net worth to permanent income ratio. Furthermore, he shows also that many low-lifetime-income households accumulate little wealth relative to their incomes. A similar behavior for the Portuguese economy is reported in Centeno and Novo (2007). They assess the quality of pre-unemployment wages as an index for the distribution of savings in the Portuguese economy using data from the Household Income

Survey for 2000. For instance, for a sample of individuals aged $[30, 39]$, those in the bottom wages quintile hold financial assets worth only 2.9 times of their median wage, while the top two quintiles hold assets worth 7.5 times their median wage.

In the analysis of the liquidity effect, we limit our attention to the restricted sample of individuals with GRR $\in [63, 67]$. For the purpose of our identification, the more homogeneous sample eliminates a source of distinct labor supply disincentives coming from differences in the UI replacement rates as described above.[8]

Table 6 reports evidence on the impact of UI on duration, reemployment wages and job tenure obtained by splitting the sample into individuals with pre-unemployment wages below and above the median wages computed for the full sample of individuals. We consider individuals with wages below the median as constrained, and those above the median as unconstrained; we will assess the sensitivity of the results to this threshold. The results support the existence of a UI liquidity effect and confirm those obtained by Chetty (2008) for the US, who concluded that constrained individuals react more – i.e. they increase their unemployment spells more – than unconstrained individuals. In our case, that is true at the older discontinuity point (48.4 vs 39.5 days for, respectively, constrained and unconstrained). However, no difference is observed with the younger group (49.2 vs 52.1 days), but the reason for this result will become clearer below. In terms of reemployment wages, there is a sharp distinction for individuals below and above the median threshold at both discontinuities. The variation in log wages is positive for constrained individuals, a 2.5 and 2.0 percent increase in wages at the younger and older discontinuity points, and negative for unconstrained individuals, -5.3 and -4.4 percent at the 30 and 40 year thresholds, respectively. These results suggest that low income individuals benefit more from increased UI generosity than higher income individuals. In terms of job tenure, the impact is non-significant both below and above the median, and for both age thresholds.

[TABLE 6 HERE (see page 30)]

To better understand the behavior of workers at each extreme of the income distribution, we further split the sample into 4 subsamples, each corresponding to an income quartile of the pre-unemployment income. In other words, each of the previous below/above median subsamples

[8]Notice that the GRR restriction applies only to pre-unemployment income, since in the post-unemployment period, we require only individuals to have a full-time job, meaning that the lower bound of reemployment wages is equal to the minimum wage.

is further divided into two. This creates a finer grid of the degree of constraints faced by unemployed individuals.

[TABLE 7 HERE (see page 31)]

The impact of UI generosity on the duration of subsidized unemployment follows a hump-shape pattern by degree of liquidity constraint; the most and the least constrained react the least, those at the second quartile react the most. This pattern is common to both discontinuities. This result is compatible with the liquidity effect reported above, but highlights the existence of other mechanisms that condition the response of individuals at each extreme of the income distribution. Except for the unemployed in the first quartile, the response of the remaining individuals by degree of liquidity constraint is as expected: smaller increases in unemployment spells for individuals with lower liquidity constraints. The behavior of the more constrained individuals deviates from this pattern. The explanation we put forward for this behavior relies on the nonstationarity characteristics of the labor market faced by these individuals, a fact that may hinder their capability to respond to the increased generosity of the UI system.

In the model of van den Berg (1990), the time and duration dependence of the arrival rate of job offers and wage offers distribution can cause nonstationarity in labor supply decision variables, for example, the reservation wage. The literature on the nonstationary job search model, recently reviewed in Eckstein and van den Berg (2007), points to the importance of these exogeneous variables in shaping the search environment at the individual level. As shown in Addison, Centeno and Portugal (2004), for a sample of European households, this environment has a great deal of heterogeneity among the unemployed. In particular, their results show that low-wage, older and less educated workers have a lower job offers arrival rate.[9] In turn, as reported in Ziliak (2003) for the US, these individual characteristics are highly correlated with the existence of liquidity constraints. If more constrained individuals face a worse labor market environment, nonstationary job search models predict that they will react less to increased generosity, which translates into a hump-shape relationship between the UI impact on unemployment duration and the degree of liquidity constraints.

A similar heterogeneous pattern is observed for reemployment wages; gains are concentrated at the low end of the income distribution, while losses appear to be the rule at the high end. At 30-year-old threshold, relatively to the counterfactual (29-year-old), the sign on log variation at

[9]Evidence for the nonstationarity of the arrival rate of job offers is presented in the Appendix.

the quartiles below the median is positive (3.4 and 2.7 percent), marginally significant, while the impact on the top two quartiles is negative (reaching -6.5 percent for the 4th quartile) with $p-$values sufficiently close or below 0.10. Older individuals in the first quartile experience a positive and sizeable impact (7.8 percent); at the other quartiles the impact on reemployment wages is negative, reaching -3.9 percent, although none is statistically different from zero.

In terms of job tenure, although the majority of (Cox proportional) estimates suggest a reduction in the duration of post-unemployment jobs, these estimates are not statistically significant. The only exception is the significant reduction in the hazard rate of job tenure for older individuals in the top income quartile; notice, however, that these individuals had a negative (non-significant) impact on reemployment wages.

The hump-shaped sensitivity of unemployment duration to UI entitlement generosity along the income distribution conforms with the elasticities of unemployment duration with respect to the level of benefits reported in van den Berg (1990). Similarly, the obtained responses of reemployment wages to the observed changes in unemployment duration matches the pattern of the elasticities of reservation wages to UI generosity also reported in van den Berg (1990). Furthermore, across age groups, the results conform to a larger labor supply elasticity of older individuals.

The identification of an liquidity effect together with the fact that more constrained individuals present positive variation in their reemployment wages and no impact on post-unemployment job tenure resets the discussion of the UI system beyond the typical moral hazard effects on labor supply.

7 Conclusions

The impact of unemployment insurance programs has long attracted the attention of empirical economics. However, this attention has been primarily directed towards estimating the distortionary disincentive effects of these programs. The purpose of this paper is to analyze the existence of a non-distortionary effect of UI, the liquidity effect, and its relationship with the quality of post-unemployment job matches (measured by the post-unemployment wages and tenure).

We find evidence that UI generosity increases job search duration, and has a small impact on wages and tenure in the subsequent employment spell. This results conform to the ones reported in Belzil (2001), Centeno (2004), and McCall and Chi (2008) and show slightly higher effects than those reported in other recent papers, such as van Ours and Vodopivec (2008), Chetty (2008), and Lalive (2007), who find a null impact on wages from changes in UI generosity.

When we consider the existence of an liquidity effect of UI, we detect a positive impact in match quality for the more constrained individuals, those at the bottom of the income distribution and who gain from the liquidity effect. The non-distortionary nature of the liquidity effect, which reduces the pressure on low income workers to accept low productivity jobs and allows them to wait for a better match, is associated with a greater (and positive) impact on reemployment wages and tenure for these workers. The results reinforce the scope for a reform of the UI system targeting low income individuals, those who benefit the most from it.

A Appendix

We computed the arrival rate of job offers by age group and unemployment duration for a sample of unemployed from the European Community Household Panel (ECHP) between 1994 and 2001. The ECHP survey interviews individuals once a year, and asks them a range of questions relating to their labor market experience during the preceding calendar year. We report results from 5 countries, namely, France, Germany, Portugal, Spain, and the United Kingdom.

To measure the job offer rates we exploit the question in the survey asking each unemployed individual whether or not (s)he received a job offer in the previous month. The actual question is: "Have you received any job offer during the past 4 weeks?"

To this reported number of offers we added those individuals who were employed at the survey date but unemployed one month earlier and who therefore obtained a job during the month preceding the interview. The general pattern does not change.

[TABLE A.1 HERE (see page 32)]

The results are presented in Table A.1. In the Portuguese labor market, the arrival rate of job offers for individuals with more than one year of unemployment is less than one third the one observed for those with less than 6 months of unemployment. The decreasing pattern of the arrival rate is clear and common to all other countries. Additionally, in Portugal individuals in the 35-39 age group face an arrival rate of job offer $^3/_4$ smaller than their 25-29 counterparts. Again similar results are obtained for the others considered. This is quite important since the observed patterns seem to prevail in markets with rather rigid labor market institutions, as the ones in Continental Europe, and also in those with more flexible institutional settings, as is the case of the United Kingdom.

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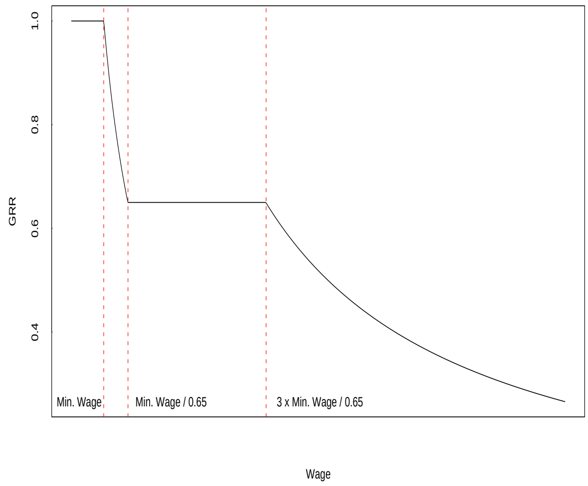

Figure 1: Financial generosity of the Portuguese UI system: Gross Replacement Rates (GRR)

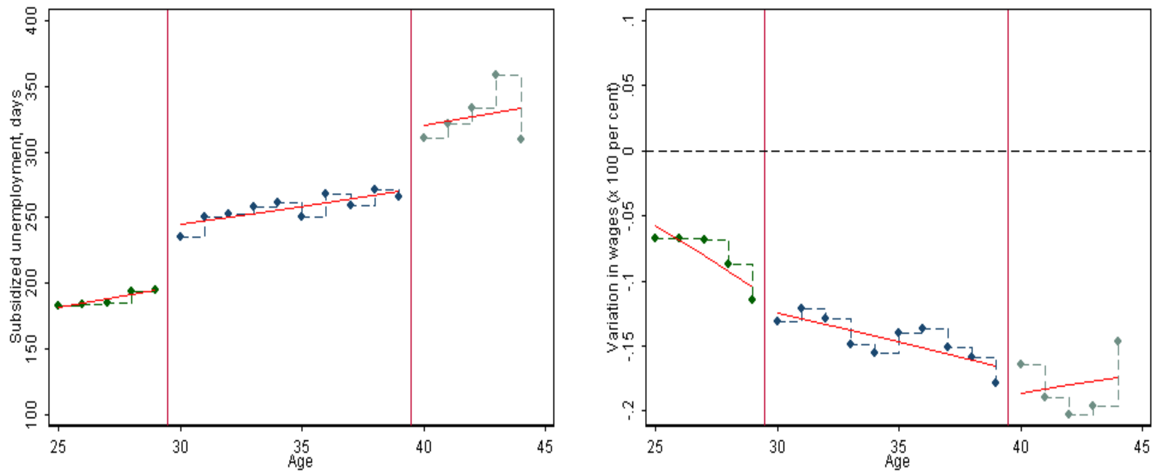

Figure 2: Average subsidized unemployment spells by age (left panel) and log difference between pre-unemployment and reemployment wages (right panel). Individuals aged 30-39 have an entitlement period 6 months longer than the youngest cohort and 6 months shorter than the oldest cohort, 40-44. See Table 1 for exact entitlement periods.

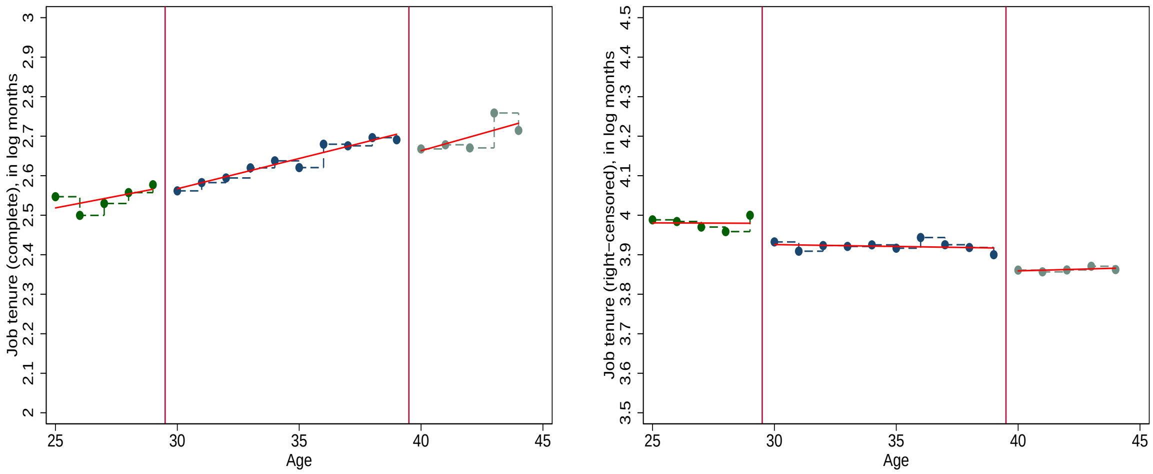

Figure 3: Average job tenure by age. Complete observations in the left panel, and (right) censored observations in the right panel.

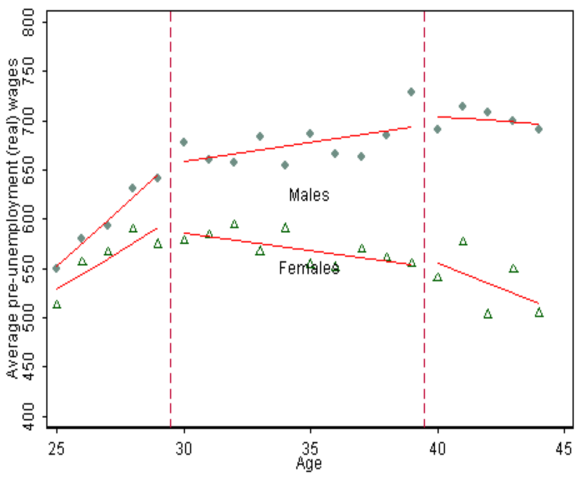

Figure 4: Pre-unemployment average wages and the forcing variable, age: There are no visible discontinuities.

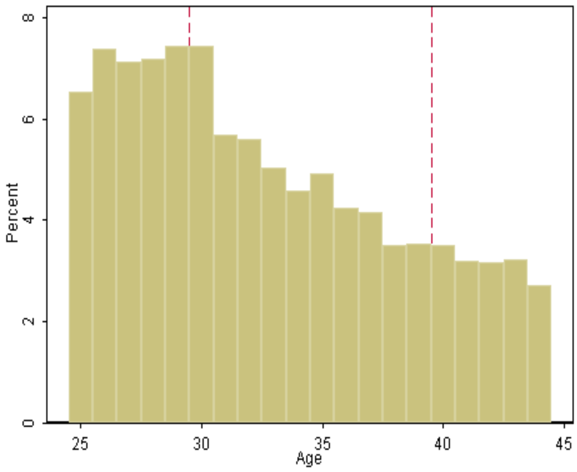

Figure 5: Density of UI claims by age. There are no noticeable discontinuities in the density of the forcing variable. In particular, at the discontinuity points 29/30 and 39/40, the percentage of claims is virtually identical. This suggests that there is no marked bias coming from self-selection into more generous UI benefits.

Table 1: Entitlement periods (in months)

Age (years)†	Entitlement period
[15, 29]	12
[30, 39]	18
[40, 44]	24
[45, 64]	30(+8)*

† Age at the beginning of the unemployment spell.
* For those aged 45 or older, 2 months can be added for each 5 years of social contributions during the previous 20 calendar years.

Table 2: Summary statistics: Mean values

	Age groups			
	[25, 29]	[30, 39]	[40, 44]	All
Duration of subsidized spell (in days)	188.4	255.4	326.6	243.0
	(129.5)	(195.8)	(266.0)	(195.3)
Pre-unemployment wage (1999 euros)	578.8	625.2	637.0	610.6
	(228.4)	(313.2)	(356.4)	(294.9)
Gross replacement ratio	67.8	67.1	67.1	67.4
	(6.4)	(7.0)	(7.6)	(6.9)
Reemployment wage (1999 euros)	550.1	548.3	526.7	545.5
	(268.0)	(280.8)	(263.3)	(273.6)
Female (Proportion)	0.54	0.49	0.39	0.49
	(0.50)	(0.50)	(0.49)	(0.50)
No. of observations	6,566	8,931	2,960	18,457

Notes: II's dataset with authors' computation. Data include all subsidized unemployment spells initiated during the period July, 1999 to December, 2002. The dataset includes also observations on region, date of UI claim, and date of reemployment. Standard deviations in parentheses.

Table 3: Impact on duration, reemployment wages and job tenure

	Age discontinuity: 30 years					Age discontinuity: 40 years				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Duration	67.0	44.8	37.3	36.9	43.5	71.3	44.4	40.7	35.7	36.1
	(2.78)	(5.82)	(10.83)	(8.00)	(5.59)	(4.57)	(8.46)	(12.19)	(15.79)	(8.10)
	0.000	0.000	0.001	0.000	0.000	0.000	0.000	0.001	0.024	0.000
Wages	-0.061	-0.009	0.024	0.012	-0.010	-0.038	-0.015	0.015	0.019	-0.012
	(0.006)	(0.013)	(0.024)	(0.018)	(0.012)	(0.008)	(0.015)	(0.022)	(0.027)	(0.015)
	0.000	0.480	0.317	0.493	0.416	0.000	0.343	0.487	0.487	0.429
Tenure		-0.029	0.077	0.016	-0.014		0.075	0.056	0.065	0.054
		(0.027)	(0.051)	(0.039)	(0.027)		(0.033)	(0.048)	(0.056)	(0.033)
		0.275	0.129	0.684	0.619		0.024	0.246	0.247	0.104
No of observations	15497	15497	15497	8249	15497	11891	11891	11891	4583	11891
Polynomial order	0	1	2	1	1	0	1	2	1	1
Bandwidth	∞	∞	∞	3 yrs	∞	∞	∞	∞	3 yrs	∞
Control variables	No	No	No	No	Yes	No	No	No	No	Yes

Notes: II's dataset for the period July, 1999 to December, 2002. The estimates are based on local linear regression with a rectangular kernel. The impact is expressed in days for duration, as a log difference for reemployment wages, and for tenure as the impact on the hazard rate of a Cox proportional model. Standard deviations are presented in parentheses, and below are the corresponding $p-$values. The bandwidth "∞" weights equally all available observations. For instance, at the 30 years discontinuity point, estimates to the left are based on '4' age points and to the right on '10'. The 3-years bandwidth is balanced around the discontinuity point. The control variables included in the indicated regressions are: female, region, month of unemployment, month of reemployment, year of unemployment, year of reemployment, and regional dummies. For the duration regression, pre-unemployment wages are also included in the set of control variables.

Table 4: Falsification test at age 35: Before and after July, 1999

	Duration (1)	Wages (2)
	-6.3	0.024
	(8.22)	(0.017)
	0.442	0.144
No of observations	8927	8927
Polynomial order	1	1
Bandwidth	∞	∞
Control variables	Yes	Yes

Notes: Before July, 1999, individuals aged 35 were entitled to 90 more days of UI benefits than those aged 34. But after July, 1999, the entitlement period is the same for both ages. Thus, we use age 35 as a pseudo-discontinuity point and perform a falsification test (column (1)). In column (2), we verify if indeed there were difference in labor market behavior between individuals aged 34 and 35, as predicted by theory. See notes in Table 3 for other details. Standard deviations are presented in parentheses, and below are the corresponding $p-$values.

Table 5: Heterogeneous impact on duration and reemployment wages: By gender and gross replacement ratio

	Age discontinuity: 30 years			Age discontinuity: 40 years		
	Male (1)	Female (2)	GRR (3)	Male (4)	Female (5)	GRR (6)
Duration	51.9	36.1	52.2	46.2	24.4	43.8
	(8.21)	(7.60)	(6.44)	(11.01)	(11.97)	(9.53)
	0.000	0.000	0.000	0.000	0.041	0.000
Wages	0.008	-0.035	-0.020	-0.005	-0.014	-0.009
	(0.018)	(0.015)	(0.014)	(0.020)	(0.019)	(0.017)
	0.657	0.019	0.159	0.787	0.444	0.578
No of observations	7603	7894	11823	6410	5481	9033
Polynomial order	1	1	1	1	1	1
Bandwidth	∞	∞	∞	∞	∞	∞
Control variables	Yes	Yes	Yes	Yes	Yes	Yes

Notes: See notes in Table 3 for details. Estimates in columns (3) and (6) are based only on individuals with GRR $\in [63, 67]$ percent. Standard deviations are presented in parentheses, and below are the corresponding $p-$values.

Table 6: UI liquidity effect: below and above median pre-unemployment wages

	Age discontinuity: 30 years		Age discontinuity: 40 years	
	Below median	Above median	Below median	Above median
	(1)	(2)	(3)	(4)
Duration	49.2	52.1	48.4	39.5
	(9.14)	(9.07)	(13.34)	(13.68)
	0.000	0.000	0.000	0.004
Wages	0.025	-0.053	0.020	-0.044
	(0.016)	(0.023)	(0.019)	(0.029)
	0.127	0.022	0.291	0.130
Tenure	-0.044	0.003	0.047	0.019
	(0.049)	(0.044)	(0.060)	(0.063)
	0.369	0.945	0.434	0.766
No of observations	5927	5896	4473	4560
Polynomial order	1	1	1	1
Bandwidth	∞	∞	∞	∞
Control variables	Yes	Yes	Yes	Yes

Notes: All estimates based on individuals with GRR $\in [63, 67]$. Estimates in columns (1) and (3) are for the subsample of individuals with pre-unemployment wages below the median value, and in columns (2) and (4) for the subsample above the median wage. Standard deviations are presented in parentheses, and below are the corresponding $p-$values. See notes to Table 3 for other details.

Table 7: UI liquidity effect: By quartiles of the pre-unemployment wages distribution

	Age discontinuity: 30 years Pre-unempl. wages quartiles				Age discontinuity: 40 years Pre-unempl. wages quartiles			
	1st	2nd	3rd	4th	1st	2nd	3rd	4th
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Duration	39.7	58.5	53.0	51.4	34.2	64.1	55.0	20.5
	(12.73)	(13.20)	(12.43)	(13.28)	(19.02)	(18.82)	(20.05)	(18.73)
	0.002	0.000	0.000	0.000	0.072	0.001	0.006	0.273
Wages	0.034	0.027	-0.043	-0.065	0.078	-0.036	-0.019	-0.039
	(0.021)	(0.025)	(0.028)	(0.037)	(0.025)	(0.028)	(0.036)	(0.043)
	0.110	0.273	0.122	0.081	0.002	0.193	0.588	0.374
Tenure	-0.128	0.061	0.160	0.113	0.006	0.092	0.092	-0.241
	(0.131)	(0.128)	(0.122)	(0.108)	(0.120)	(0.125)	(0.126)	(0.135)
	0.330	0.635	0.189	0.298	0.961	0.460	0.465	0.074
No of observations	2945	2982	2985	2911	2189	2284	2211	2349
Polynomial order	1	1	1	1	1	1	1	1
Bandwidth	∞	∞	∞	∞	∞	∞	∞	∞
Control variables	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: All estimates based on individuals with GRR $\in [63, 67]$. Standard deviations are presented in parentheses, and below are the corresponding $p-$values. See notes to Table 3 for other details.

Table A.1: Arrival rate of job offers by age and unemployment duration

	Age			Unemployment duration (in months)			
	25-29	30-34	35-39	< 6	$6 - 12$	$12 - 24$	> 24
France	0.103	0.091	0.088	0.086	0.067	0.076	0.040
Germany	0.200	0.143	0.138	0.142	0.148	0.034	0.026
Portugal	0.037	0.032	0.029	0.032	0.019	0.012	0.000
Spain	0.101	0.095	0.065	0.096	0.061	0.053	0.044
United Kingdom	0.135	0.069	0.068	0.085	0.056	0.027	0.022

Source: European Community Household Panel (ECHP), 1994-2001. Authors' computations. Job offer rates are computed from the answers to the question: "Have you received any job offer during the past 4 weeks?"