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Benefit Duration, Unemployment Duration and Job Match Quality: A Regression-Discontinuity Approach

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Abstract

We use a sharp discontinuity in the maximum duration of benefit entitlement to identify the effect of extended benefit duration on unemployment duration and post-unemployment outcomes (employment stability and re-employment wages). We address dynamic selection, which may arise even under an initially random assignment to treatment, estimating a bivariate discrete-time hazard model jointly with a wage equation and correlated unobservables. Due to the non-stationarity of job search behavior, we find heterogenous effects of extended benefit duration on the re-employment hazard and on job match quality. Our results suggest that the unemployed who find a job close to and after benefit exhaustion experience less stable employment patterns and receive lower re-employment wages compared to their counterparts who receive extended benefits and exit unemployment at the same period. These results are found to be significant for men but not for women.

Keywords: Unemployment insurance, job search, post-unemployment outcomes, dynamic selection, effect heterogeneity.
JEL codes: J64, J65, C41.

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1 Introduction

One common feature in the design of most Unemployment Insurance (UI) systems in the OECD countries is the existence of a limited benefit duration. The non-stationarity that arises because of limited benefit duration implies that individuals with different lengths of benefit entitlement should have different optimal paths of reservation wage and search effort over time (Mortensen, 1977; Van den Berg, 1990). A number of studies have investigated the effect of potential benefit duration on the exit rate from unemployment finding that longer benefit duration leads to prolonged unemployment spells (e.g. Meyer, 1990; Katz and Meyer, 1990; Card and Levine, 2000 for the US, Roed and Zhang, 2003; Lalive, Van Ours, Zweimüller, 2006; Van Ours and Vodopivec, 2006; Card, Chetty and Weber, 2007a and 2007b for Europe). Another common finding of this literature is that limiting unemployment benefit duration tends to introduce a spike in the exit rate around benefit exhaustion.

The difference in the optimal job search behavior due to differences in the level or the potential duration of benefits may not only affect the unemployment exit rate but may also lead to different realized distributions of job match quality (e.g. Ehrenberg and Oaxaca, 1976; Belzil 1992, 1995, 2001; Marimon and Zilibotti, 1999). For example, unemployed workers who are approaching benefit expiration may accept jobs of lower quality because they become less selective. Alternatively, individuals with a given length of unemployment, the same level of benefits, but a longer period of remaining benefit entitlement may wait for job offers which are better either in terms of re-employment wages and/or employment stability. Extending benefit duration may, therefore, have an heterogenous effect on post-unemployment outcomes over the unemployment duration.

In this paper we investigate the heterogenous effects of extending benefit duration on the unemployment exit rate and on job match quality. As measures of job match quality we consider post-unemployment outcomes such as employment stability and re-employment wages. We use an inflow sample into unemployment for the years 2001 to 2003 based on detailed administrative records from Germany. The information includes a seven year labor market history and allows us to observe labor market states of individuals for three years after entering unemployment.

Our identification strategy relies on a sharp discontinuity in the maximum duration of unemployment benefits in Germany, which increases from 12 to 18 months at the age of 45. This creates a variation of benefit duration entitlement, which is based on the age at which a worker enters into unemployment. Comparing individuals who enter into unemployment just below the age threshold of 45 with those who become unemployed just above this threshold can be used to identify the effect of benefit extensions in a regression discontinuity (RD) framework. The main identification assumption is that the assignment into treatment (being entitled to extended benefits) is only determined by the age at entering into unemployment and is orthogonal to remaining unobserved heterogeneity. Moreover, it should be noted that our RD design identifies a local treatment effect which can be extended to population effects only with additional assumptions.

Our analysis based on a RD design may suffer from two potential selection issues that might invalidate our identification strategy. First, manipulation of the treatment status may occur which is related to skills and factors that are not measured such as motivation or honesty. We test the frequency of the inflow into unemployment and an extensive set of observable characteristics of the unemployed around the discontinuity and we find no evidence of selection on observables. However, selection on unobservables cannot be tested. Second, although the assignment into treatment at the beginning of the unemployment spell is based on this sharp discontinuity, there might be selection in the resulting sample of the re-employed based on observed and unobserved characteristics (Ham and LaLonde, 1996). We address the dynamic selection by estimating a bivariate discrete-time hazard rate model jointly with wages and allowing for potentially correlated unobserved heterogeneity.

We find that limited benefit duration is associated with non-stationary search behavior, which leads to a spike around benefit expiration. Most importantly, due to the non-stationarity of job search, our results suggest that there exist heterogenous effects of extended benefit entitlement duration on post-unemployment outcomes. In particular, the unemployed who receive benefits for 12 months and find a job close to and after benefit exhaustion tend to experience less stable employment patterns and to receive lower re-employment wages compared to their counterparts who receive extended benefits and exit unemployment at the same period. In other words, those with shorter benefit duration accept jobs they would otherwise reject, while those who receive extended benefits tend to accept jobs that last longer and pay higher wages. These results are found to be significant for men but not for women.

Previous research on the impact of UI on post-unemployment outcomes has focused on an homogenous or average effect offering mixed results; some studies find a positive effect (e.g. Ehrenberg and Oaxaca, 1976; Belzil, 1992, 1995, 2001; Tatsiramos, 2009) while others find no effect (e.g. Card, Chetty and Weber, 2007a and Van Ours and Vodopivec, 2008). Similar to our empirical approach, Card, Chetty and Weber (2007a) apply a RD design for

Austria based on the fact that previous employment duration determines the maximum length of unemployment benefit receipt of laid-off workers. Van Ours and Vodopivec (2008) apply a difference in differences approach based on the reduction of potential duration of UI benefits over time for some groups of unemployed workers in Slovenia. Both studies do not investigate potential effect heterogeneity with respect to elapsed unemployment duration and do not model selection into employment based on unobserved heterogeneity. Their findings suggest that longer potential benefit duration lowers job-finding rates but has no effect on subsequent job match quality. The contribution of our paper is to consider heterogeneous treatment effects over the duration of unemployment. This approach extends upon the previous studies that focus on average effects and enhances our understanding of the impact of unemployment benefits.

The paper is organized as follows: Section 2 outlines the theoretical framework and the existing empirical evidence. Section 3 describes the institutional background and the data. Section 4 presents the econometric model and discusses selectivity issues. The results of the empirical analysis are presented and discussed in Section 5 and Section 6 concludes.

2 Theoretical Framework and Empirical Evidence

Standard search theory predicts that an increase in UI benefit generosity, either in terms of benefit duration or replacement rates, has a negative impact on the job search activities of the unemployed increasing their unemployment duration. The value of unemployment depends on the benefit duration and the benefit level in a non-separable way, which implies that an increase of the benefit level increases the value of unemployment more if the potential benefit duration is longer. Unemployed workers exert lower search effort as the opportunity cost of search is lower and they choose higher reservation wages. Moreover, closer to the time of benefit exhaustion, the value of unemployment drops since the marginal benefit of search increases and the reservation wage falls, leading to a higher exit rate out of unemployment (Mortensen, 1977).

This non-stationarity implies that individuals with different lengths of benefit entitlement should have different optimal paths of reservation wage and search effort over time (Van den Berg, 1990). Consider the following stylized example of an extended benefit duration on the unemployment hazard rate. There are two groups of unemployed: a) those with benefit duration of 12 months (control group) and b) those who receive extended benefits of 18 months (treated group). These values correspond to the potential benefit durations for the groups we consider in the empirical analysis. At the beginning of the unemployment spell both groups exhibit a constant exit rate from unemployment, which is lower for the treated group because of the higher value of unemployment as a result of the longer potential benefit duration. As the control group approaches benefit expiration (month 12), the exit rate is increasing compared to the treated group and stays constant at a higher level after expiration. Similar to the control group, the exit rate for the treated group also starts increasing closer to benefit expiration (month 18) and remains constant and equal to the control group after expiration.[1] During the period between months 12 to 18, the exit rate of the control group is higher because benefits have been exhausted, while the treated receive benefits until month 18.

Considering the difference in the exit rates between the two groups of unemployed, during the initial period of unemployment this difference is positive but constant, which reflects the higher exit rate of the control group. Thereafter, the exit rate difference is increasing and reaches its maximum at the point of benefit exhaustion for the control group. After month 12 the difference starts decreasing as the exit rate for the treated group increases approaching benefit expiration, while the one for the control group remains constant. It is important to note that the difference in the exit rates between the two groups is at its maximum level when the benefits for the control group expire, while it goes to zero at the time of benefit expiration for the treated group.[2]

Numerous studies have investigated the effect of benefit duration on the exit rate from unemployment both in the U.S. and in Europe. Meyer (1990), Katz and Meyer (1990), Card and Levine (2000) and Addison and Portugal (2008) find for the US a sharp increase in the exit rate from unemployment before benefits are exhausted. Hunt (1995) for Germany, Winter-Ebmer (1998), Lalive and Zweimüller (2004), Lalive, Van Ours and Zweimüller (2006) and Lalive (2008) for Austria find that benefit extensions reduce job finding rates and create a spike around benefit exhaustion. Roed and Zhang (2003) find for Norway that a marginal increase in unemployment compensation reduces the unemployment exit rate and also find a sharp increase prior to benefit exhaustion, which is larger for women than for men. Van Ours and Vodopivec (2006) for Slovenia find that a reduction of benefit duration increases the transition out of unemployment and produces a shift of the hazard spike from the old to the new benefit exhaustion point.[3]

[1]In principle, those eligible to longer benefit duration should exhibit a higher exit rate after expiration because of the entitlement effect. However, the control group will also be eligible for longer benefits in the next unemployment spell because they will be older than 45 years. Therefore, we do not expect any differences in the hazard rates after month 18. We discuss the institutional details and the way we define the treated and control groups in Section 3.1.

[2]A graphical illustration of this example can be found in Figure B.1 in the Supplementary Appendix.

[3]For Germany, the evidence of follow-up studies on the effects of institutional changes in the 1980s is rather

Besides the trade-off between insurance and the incentives to leave unemployment for a job, the relationship between the length of benefit entitlement and the optimal path of job search behavior might lead to a positive relationship between insurance and the quality of jobs obtained. The reason is that closer to benefit expiration and after benefits have expired workers may become less selective obtaining lower quality jobs. Therefore, the difference in the optimal job search behavior of individuals with different lengths of benefit entitlement over time could lead to different realized distributions of job quality. Individuals with a given length of unemployment, the same level of benefits, but a longer period of remaining benefit entitlement may wait for job offers which are better in terms of re-employment wages and employment stability.

Ehrenberg and Oaxaca (1976) were the first to consider the effect of UI on post-unemployment outcomes finding a positive effect of benefits on re-employment wages. Addison and Blackburn (2000) review the early literature and provide results which suggest a weak effect of UI on re-employment wages. More recently, Centeno and Novo (2009) exploit a reform of the Portuguese UI system that increased the entitlement period for some age-groups. They also find that the extension had a small but positive effect on re-employment wages, which is stronger at the bottom of the pre-unemployment wage distribution.

Another strand of the literature measures the effect of UI generosity on post-unemployment outcomes with the incidence of unemployment, or the time elapsed between re-employment and acceptance of a subsequent job, using job matching arguments based on Jovanovic (1979). In a series of papers, Belzil (1992, 1995, 2001) analyzes unemployment experience and employment duration in the context of the Canadian UI reform finding a weak positive relationship between re-employment duration and unemployment benefit generosity. Centeno (2004) studies the effect of the generosity of U.S. benefit levels and finds that larger UI benefits lead to longer subsequent employment spells. Tatsiramos (2009) analyzing data from eight European countries, comparing UI recipients with non-recipients, finds that jobs which are accepted while the unemployed worker is still insured last longer. In addition, this beneficial effect of unemployment insurance on employment stability is pronounced in countries with relatively more generous benefit systems.

mixed. Fitzenberger and Wilke (2010) based on administrative data find that firms and older workers make use of the extended benefit entitlement periods as part of early retirement schemes, but they do not find any impact on the time spent in unemployment before finding a new job. Schneider and Hujer (1997) based on the SOEP data also find no significant effect of the reforms on the duration of unemployment whereas Steiner (2001) finds a negative correlation of receiving unemployment benefits and the probability of leaving unemployment. For a detailed discussion of the literature on Germany see Fitzenberger and Wilke (2010).

Card, Chetty and Weber (2007a) apply a RD design and find for Austria that an increase in benefit entitlement length reduces job-finding rates but does not have any effect on subsequent job match quality, measured in wage growth and job duration. Similarly, Van Ours and Vodopivec (2008) find no effect of benefit duration cuts on the quality of post-unemployment jobs for Slovenia. Fitzenberger and Wilke (2010) also do not find evidence for an improved job match quality for older workers after extensions of the maximum entitlement periods for this group in Germany in the 1980s. In related work, Schmieder, von Wachter and Bender (2012) also make use of age discontinuities in the German UI system. Using a larger time span from 1987 to 2004 they focus on differences of the impact of UI generosity over the business cycle and find no evidence for heterogenous effects.

These empirical studies usually refer to different economies and different states of the business cycle. This is one reason why it is difficult to compare the results with each other. Moreover, they do not rely on the same identification conditions. For example, with respect to the studies analyzing post-unemployment outcomes, Belzil (1992, 1995, 2001), Centeno (2004) and Tatsiramos (2009) estimate parameters which refer to a more general population, while Card et al. (2007a), Van Ours and Vodopivec (2008), Fitzenberger and Wilke (2010) and Schmieder et al. (2012) use local identification. In this paper, we also rely on local identification and estimate the impact of benefit duration for a specific subgroup of unemployed individuals.

The job match quality aspect of UI has been also considered in a number of theoretical papers. Acemoglu and Shimer (1999) show that the increased utility of unemployment when receiving UI induces workers to search for higher wages and firms respond by creating high-wage, high-quality jobs. In their general equilibrium model, an economy with risk-averse workers requires a positive level of UI to maximize output, while an economy with risk-neutral workers achieves the highest output without any unemployment insurance. Marimon and Zilibotti (1999) show that UI can increase job match quality by helping workers to get jobs which are compatible with their skills and therefore less likely to dissolve. The empirical equilibrium search model presented by Van den Berg and Ridder (1998), which is based on the assumption of risk-neutral workers, implies that an increase of the benefit level does not affect unemployment as long as the benefit level does not exceed the productivity level. Their estimation results suggest that for their data a moderate increase of the benefit level has no welfare impact. Van Vuuren, Van den Berg and Ridder (2000) and Lise, Meghir and Robin (2011) consider the effect of labor market policies on welfare and find that the presence of UI can lead to increased welfare.

3 Institutional Background and Data

3.1 Institutional Background

Germany has undergone some major labor market reforms in the recent years including the *Hartz reforms* which consisted of, among other things, a change of the unemployment benefit and social assistance schemes (see Konle-Seidl, Eichorst and Grienberger-Zingerle, 2010 for a detailed description of the unemployment insurance system in Germany and its changes over time). In our empirical analysis we are focussing on an inflow sample of unemployed workers between 2001 and 2003, a period prior to the *Hartz reforms*.

Prior to the reforms, Germany had a system of income protection which was based on three pillars: 1) unemployment benefits, 2) unemployment assistance and 3) social assistance. Unemployment benefits (UB, *Arbeitslosengeld*) provide earnings-related income replacement and are based on an employment record in a reference period (see §127, Social Code III, *Sozialgesetzbuch III*). The replacement rate of UB depends on family status, while the duration depends on age and previous employment duration. Unemployed persons with at least one child are entitled to 67% of previous net remuneration and 60% otherwise; individual means or needs are not taken into account. The exact amount is calculated based on the average gross daily income within the assessment frame of twelve months from which social security contributions, income tax and the solidarity surcharge were subtracted to get the average net daily income which is the basis for the UB claim. The benefits are funded through employer and employee contributions and administered by the Public Employment Services (PES).

To generate a claim for UB workers had to be employed for at least 12 months in the last three years (*Rahmenfrist*) before entering unemployment; workers who have been employed less than 12 months within the last three years were not entitled for UB, but could receive means-tested social assistance. The maximum duration of unemployment benefits varied between 6 and 32 months. Depending on age and months worked in the last seven years, there exist several discontinuities in the maximum duration of unemployment benefits (see Table 1). For the purpose of our analysis we are focusing on the discontinuity at the age of 45 for which the maximum benefit duration increases by six months – from 12 to 18 months – given the workers have been employed for at least 36 months in the last seven years. Other discontinuities also appear at age 47 and 52 which lead to an increase of the maximum benefit duration by four months, conditional on previous employment duration of 44 and 52 months, respectively. We concentrate on the discontinuity at the age of 45 because the additional jumps from 18 to 22 and 22 to 26 occur at a very late stage in the unemployment spell and it seems reasonable to expect that the transition rate from unemployment to employment is quite low at this stage independent of receiving unemployment benefits or unemployment assistance. Finally, there is another discontinuity at age 57 of six months increase of benefit duration from 26 to 32 months. We do not consider this discontinuity either because it is very much related to early retirement (see Tatsiramos, 2010, for an analysis of unemployment and early retirement for older workers).

After the entitlement period of UB had expired unemployed individuals were eligible for principally unlimited and means-tested unemployment assistance (UA, *Arbeitslosenhilfe*). These benefits were still earnings-related (57%/53% replacement rate with/without children). In contrast to UB, the UA was granted for an unlimited period (as long as individuals were available for the labor market) and funded through the Federal budget, i.e., by general taxation. Finally, the social assistance (SA, *Sozialhilfe*), provided basic income protection on a means-tested and flat-rate basis for all German inhabitants. This assistance was independent of employment experience but conditional on not having other resources from earned income, other social benefits or family transfers. This makes clear that the benefits for unemployed individuals do not drop down to zero once the maximum duration for unemployment benefits is reached.

A worker who enters unemployment and is eligible for unemployment benefits keeps the entitlement for up to four years. The entitlement expires either after this time period or if a new entitlement emerges. To generate a new entitlement it is necessary to be employed for at least 12 months. In case that the worker still had months left from an old UB entitlement, the new entitlement is added to the old one up to the maximum possible entitlement according to age. This entitlement rule is important as it reduces the incentive for unemployed workers who are close to the age threshold to get a short job in order to be entitled for longer benefits.

3.2 Data and Sample

Our data are drawn from the *IZA Evaluation Data Set* which is an ongoing data collection process in order to provide a new data source for labor market research. Part of this data set is a random inflow sample into unemployment in Germany for the years 2001-2007 containing over 855,000 individuals corresponding to 4.7% of the total population of unemployment entrants (see Caliendo et al., 2010, for details). In this paper we use an inflow sample into unemployment from the years 2001 to 2003. The data is based on the 'Integrated Labor Market Biographies' (ILMB, *Integrierte Erwerbs-Biographien*) of the Institute of Employment Research (IAB), containing relevant register data from four sources: employment history, unemployment support recipience, participation in active labor market programs, and job seeker history. This gives us access to detailed daily information on employment subject to social security contribution, including occupational and sectoral information, and the receipt and level of transfer payments during periods of unemployment, job search, and participation in different programs of active labor market policy.

Furthermore, a large variety of socio-demographic and qualificational variables is available. We can use variables such as age, marital status, number of children and nationality (German or foreigner). A second class of variables refers to the human capital of the individual. The attributes available are school degree and job qualification. In addition, we can also draw on an extensive labor market history and career variables including nearly complete seven-year labor market history, daily earnings from employment, amount of daily unemployment benefits and previous profession. The employment outcomes of these individuals are observed for three years after entering unemployment.

Eligibility for unemployment benefits is based on age and previous employment experience. We restrict our sample to men and women from West Germany who have been employed for at least 36 months in the last seven years when entering unemployment and have been working for 12 months in regular employment in the last year prior to entering unemployment. This ensures that extended benefit duration i) only depends on age and ii) that all individuals are eligible for 12 and 18 months of benefit entitlement, respectively, since they have generated a new benefit entitlement (see Section 3.1 for details).

We further restrict our sample for men to be aged between 44 and 46 years which leaves us with 2,241 male unemployed (see Table 2, first line). For women we choose a slightly larger age range from 43.5 to 46.5 years which results in 2,776 female unemployed. This wider age range for women ensures that we have enough observations for both males and females in order to allow for heterogeneous effects.

We only consider two labor market states in our analysis: unemployment and employment. The unemployment state includes registered unemployment with or without receiving benefits, participation in active labor market programs and job-seeking (if not in regular employment at the same time). Unemployment also includes "out of labor force" because we are interested in the effect on the time until the next job and not on the time being officially registered as unemployed. Since we do not have any information about self-employment in the administrative data, the latter might also include people who became self-employed, which is a relatively small group in Germany. The employment state includes only individuals who exit unemployment and who are in regular employment, i.e. those who do not fall in one of the mentioned unemployment categories. Participants in public work programs or individuals receiving wage subsidies are not treated as regular employed.

Based on these definitions, Table 2 contains the number of transitions between the two states. For unemployment we report these transitions in monthly intervals up to month 18. Across gender we observe approximately 25%-27% of the observations as right-censored in unemployment. That is, these individuals do not leave unemployment within our observation period of 36 months. Around 35%-42% leave unemployment for a job within the first six months, while the majority is unemployed for more than six months. Conditional on having made a transition from unemployment to employment we also see that around 46% of the men and 58% of the women remain in this state until the end of our observation window.

4 Empirical Approach

The goal of our empirical analysis is to examine the effects of extended benefit duration on unemployment duration and post-unemployment outcomes, such as the stability or duration of the subsequent employment spell and re-employment wages. We have outlined in Section 3.1 that the German legislation for unemployment benefits contains sharp discontinuities with respect to age which we will exploit as a source for identification. With a RD approach we will be able to measure the effects of the treatment at some threshold, see e.g. Hahn, Todd and Van der Klaauw (2001) for a discussion of the RD approach.

Our estimations are based on a specific group of unemployed individuals. By conditioning on individuals with more than 36 months of labor market experience in the last 7 years we are focusing on a sub-population with a higher labor market attachment, which is potentially eligible – depending on age – for the maximum benefit duration. Conditional on a large set of controls, which includes among others detailed past employment history and earnings, our maintained assumption is that whether an individual receives longer benefit duration is only a function of age at entering into unemployment and is orthogonal to remaining unobserved heterogeneity. In this sense, the results from our RD approach are only relevant for this sub-population – unemployed workers aged around 45 who have worked at least 36 months in the last 7 years – and cannot be extended to other segments of the population without imposing the strong assumption of common treatment effects for all unemployed individuals.

4.1 Econometric Model

The analysis is based on a bivariate discrete-time hazard model for the transition from unemployment to employment and for the transition from employment back to unemployment (for those who obtain a job). This model is jointly estimated with the wage equation.[4] Estimating a discrete-time hazard model allows us to take into account a number of important aspects. First, we can easily estimate effect heterogeneity of the treatment depending on the time spent in unemployment. Second, we can handle the many right-censored observations in our sample for which we do not observe the end of the unemployment spell and the end of the subsequent employment spell, respectively, before the end of our observation window. Third, we can allow for time-varying factors like the local unemployment rate.[5]

Although the assignment into treatment at the beginning of the unemployment spell is based on a sharp discontinuity and is therewith assumed to be orthogonal to the error term, there might be dynamic selection in the resulting sample of the re-employed based on observed and unobserved characteristics. For a similar argument in the context of experimental data on training see Ham and LaLonde (1996). In our specification, we control for a broad range of observable characteristics like education, type of occupation and wage in the last job, which might be important for the transitions processes. In order to take dynamic selection based on unobservable characteristics into account, we estimate the bivariate hazard rate model with potentially correlated unobservables influencing both the duration of unemployment and the duration of subsequent employment. Additionally, we allow the unobservables to be correlated with the realized wages.

Following the standard practice in the literature (see, for instance Ham and LaLonde, 1996; Cameron and Heckman, 1998; Bover, Arellano and Bentolila, 2002, among others), the discrete hazard function is specified as the logistic. The transition rate from unemployment to

[4]An alternative modeling strategy is to model explicitly benefit receipt as a time-varying monthly indicator, which is a function of previous employment experience, age and other individual characteristics without any further sample restriction on previous labor market experience. This equation could be then jointly estimated with the unemployment and employment transitions and the wage equation allowing for correlated unobserved heterogeneity. Studies following this approach include Bover, Arellano and Bentolila (2002) for the effect of benefit receipt on unemployment duration, and Tatsiramos (2009) which extends this approach to subsequent employment stability. In these models one can allow for heterogenous treatment effects, for example with respect to age or elapsed unemployment duration, by interacting corresponding observable characteristics with the treatment indicator. The difference between this alternative modeling strategy and the papers which rely on instrumental variables or regression discontinuities is that the latter assume separability between benefit duration and unobserved heterogeneity.

[5]Since we observe the duration in the two states on a daily basis, this would principally allow us estimating a continuous time duration model. However, as in Germany most of the employment spells start at the beginning of a month (and unemployment spells last until the end of a month), we construct discrete time spell data in which one month corresponds to one time unit.

employment λ_u in period t can be written as:

$$\lambda_u(t) = (1 + exp(-y_{ui}(t)))^{-1},$$

where

$$\begin{aligned} y_{ui}(t) &= \beta_{0u} + \sum_{d=2}^{k} \beta_{1ud} I_{uid}(t_u) + \sum_{d=1}^{k} \delta_{ud} I_{uid}(t_u) D_i + \beta_{2u} X_{it} \\ &\quad + \beta_{3u} D_i (Age_i - Age_0) + \beta_{4u}(1 - D_i)(Age_i - Age_0) + \eta_{ui}. \end{aligned} \tag{1}$$

The effect of duration dependence is modeled in a flexible way by using time dummy variables denoted as $I_{uid}(t_u)$, which are equal to one when duration in unemployment t_u is within the duration intervals denoted by the subscript $d = (2, ..., k)$. The treatment indicator D_i takes the value of one if the unemployed is above the age of 45 at the time of entering unemployment and zero otherwise. As we have outlined in Section 2, we expect a higher exit rate out of unemployment closer to the time of benefit exhaustion. Corresponding to that, we expect the treatment effect to vary over time spent in unemployment. Therefore, we interact the treatment indicator D_i with time dummy variables denoted as $I_{uid}(t_u)$. The coefficient δ_{ud} captures the causal effect of the increase in the maximum benefit duration on the hazard rate from unemployment to employment in the interval d of the unemployment spell with $d = (1, ..., k)$. We also estimate the more restrictive specification in which we consider only the average effect of being treated, which is captured by the single coefficient of the treatment indicator D_i.

The parameters β_{3u} and β_{4u} capture the effects of the assignment variable age below and above the threshold on the probability of leaving unemployment for a job. This ensures that δ_{ud} does not capture a general age effect but the causal impact of the discontinuity in the benefit duration. In addition, we control for observable characteristics X_{it} including time-invariant individual characteristics like education and previous employment history and the time-varying local unemployment rate. Finally, η_{ui} describes unobserved heterogeneity influencing the transition process from unemployment to employment.

Besides the transition process from unemployment to employment a main focus of our study is on the effect of extended benefit duration on the stability of new employment. The corresponding transition rate from employment to unemployment λ_e in period t is given by:

$$\lambda_e(t) = (1 + exp(-y_{ei}(t)))^{-1},$$

where

$$
\begin{aligned}
y_{ei}(t) &= \beta_{0e} + \sum_{d=2}^{k} \beta_{11ed} I_{eid}(t_e) + \sum_{d=2}^{k} \beta_{12ed} I_{uid}(T_{ui}) + \sum_{d=1}^{k} \delta_{ed} I_{uid}(T_{ui}) D_i \\
&\quad + \beta_{2e} X_{it} + \beta_{3e} D_i (Age_i - Age_0) + \beta_{4e}(1 - D_i)(Age_i - Age_0) + \eta_{ei}. \qquad (2)
\end{aligned}
$$

Similar to the transition process from unemployment to employment the effect of duration dependence is modeled flexibly by using time dummy variables denoted as $I_{eit}(t_e)$, with t_e corresponding to the duration in employment. In addition to time in the current spell we control for duration of the previous unemployment spell T_{ui} by using time dummy variables denoted as $I_{uid}(T_{ui})$. Since we expect that workers might become less selective the closer they are to benefit expiration and after their benefit entitlement has expired, we additionally interact the impact of being eligible for 18 months of benefits ($D_i = 1$) with the time interval in which individuals left the unemployment spell for a job, described by the time dummy variables $I_{uid}(T_{ui})$. The causal effect of extended benefit duration on subsequent job stability for individuals who left unemployment in interval d is given by δ_{ed}. The specification for the employment hazard includes controls for the observable characteristics X_{it}, for different effects of age below (β_{3e}) and above (β_{4e}) the threshold, and for unobserved heterogeneity influencing the transition process from employment to unemployment denoted by η_{ei}.

We implicitly control for the benefit level by including the previous wage as a regressor. An alternative approach would be to additionally allow for effect heterogeneity with respect to the benefit level. This would be in line with the sequential job search model, which implies non-separability between benefit level and benefit duration. However, given our limited sample size and the focus of the paper on effect heterogeneity with respect to elapsed unemployment duration, we estimate common effects independent of the benefit level.

For the estimation of the effect of extended benefits on wages we estimate the following linear regression:

$$
\begin{aligned}
\log(w_i) &= \beta_{0w} + \sum_{d=2}^{k} \beta_{12wd} I_{uid}(T_{ui}) + \sum_{d=1}^{k} \delta_{wd} I_{uid}(t_u) D_i + \beta_{2w} X_{it} \\
&\quad + \beta_{3w} D_i (Age_i - Age_0) + \beta_{4w}(1 - D_i)(Age_i - Age_0) + \eta_{wi} + \epsilon_{it} \qquad (3)
\end{aligned}
$$

Similar to the transition process from employment to unemployment, we control for duration of the previous unemployment spell T_{ui} by using time dummy variables. The causal effect of extended benefit duration on re-employment wages, for individuals who left unemployment in interval d, is given by δ_{wd}. The specification for wages also includes controls for the observable characteristics X_{it}, for the effects of age below (β_{3w}) and above (β_{4w}) the threshold and for the impact of unobserved heterogeneity on wages, which is denoted by η_{wi}. For the main specification we assume that the error term follows a normal normal distribution with $\epsilon_{it} \sim N(0, \sigma^2)$. We additionally estimate a more flexible specification assuming a two-dimensional mixture of normals for the stochastic wage component. In this specification we assume that $\epsilon_{it} = p_1\epsilon_{1it} + (1 - p_1)\epsilon_{2it}$, with $\epsilon_{1it} \sim N(0, \sigma_1^2)$ and $\epsilon_{2it} \sim N(0, \sigma_2^2)$.[6]

The indicators τ_u and τ_e take on the value one if a transition to employment or to unemployment, respectively, is observed and zero otherwise. The likelihood contribution of an individual i with an unemployment spell of j_u intervals, a subsequent employment spell of j_e intervals and an observed wage w_i for given unobserved characteristics η_{ui}, η_{ei} and η_{wi} for the basic specification is given by

$$\begin{aligned} l_i(\eta_{ui}, \eta_{ei}, \eta_{wi}) \quad = \quad & \prod_{t=1}^{j_u-1}(1 - \lambda_u(t|\eta_{ui}))(1 - \lambda_u(j_u|\eta_{ui}))^{(1-\tau_u)}\lambda_u(j_u|\eta_{ui})^{\tau_u} \\ & \prod_{t=T_{ui}+1}^{T_{ui}+j_e-1}(1 - \lambda_e(t|\eta_{ei}))(1 - \lambda_e(T_{ui} + j_e|\eta_{ei}))^{(1-\tau_e)}\lambda_e(T_{ui} + j_e|\eta_{ei})^{\tau_e} \\ & \frac{1}{\sqrt{2\pi\sigma^2}}e^{(-\frac{(\log(w_i) - \widehat{\log w_i})^2}{2\sigma^2})} \end{aligned} \tag{4}$$

Following Heckman and Singer (1984), the unobserved heterogeneity distribution is defined as a discrete distribution with the support points denoted by (η_{up},η_{ep},η_{wp}) and the corresponding probability mass given by $P(\eta_{ui} = \eta_{up}, \eta_{ei} = \eta_{ep}, \eta_{wi} = \eta_{wp})$=$\pi_p$. Each unobserved factor is assumed to be time invariant and individual-specific for each state. This flexible specification allows for free correlations between the different terms of unobserved heterogeneity. In the estimation we increase the number of support points until the model cannot be improved further by an additional support point, evaluated on the basis of the Akaike Information Criterion (AIC). The unobserved factors are allowed to be different and correlated across unemployment and employment spells, and they are assumed to be uncorrelated with observable characteristics X_{it} and the treatment indicator D_i, whereby the latter assumption is in line with our regression discontinuity design. The sample likelihood is given by

[6] Alternatively, one could model the impact of benefit entitlement on reservation wages instead of wages. We decided to follow the empirical literature and to analyze the impact on realized wages, since this is the more relevant measure of job quality for both the job-seeker and the policy maker. For example, tax payment and future unemployment benefits depend on the realized wages.

$$L = \prod_{i=1}^{n} \sum_{p=1}^{P} \pi_p l_{ip}, \tag{5}$$

where the individual likelihood contribution given the unobserved characteristics η_{up}, η_{ep} and η_{wp} is denoted by l_{ip}.

One major concern with respect to the bivariate hazard rate model in the context of single spell data might be that identification of the correlation based on unobserved characteristics is mainly driven by the functional form of the model. However, as discussed e.g. by Eberwein, Ham and LaLonde (1997) and Gaure, Roed and Zhang (2007), this concern is more serious if the model only includes time-invariant variables. Time-varying variables like the local unemployment rate, which is included in our specification, provide a more robust source of identification. In addition, we model both the duration dependence and the unobserved heterogeneity in a quite flexible way, which ensures that the results are not driven by the pre-specified functional forms or distributional assumptions with respect to these dimensions.

4.2 Selectivity Issues

One important identification assumption is that the assignment to treatment around the threshold is random. However, rational agents will take the distance remaining before the age of 45 into account when deciding about accumulation of work experience. The incentives to work longer are higher for individuals who are eligible for 18 instead of 12 months of unemployment benefits. In our estimations, we compare individuals below and above the age threshold, conditional on having worked for at least 36 months in the last seven years. Due to the different incentives depending on the distance from the age threshold, this might lead to differences in types below and above the threshold, i.e. to differences in observed and unobserved characteristics between treated and non-treated individuals.[7] Moreover, firms and workers may alter the timing of layoffs leading to non-random selection around the threshold in order to qualify for a more beneficial treatment. It is not possible to test for differences in unobserved characteristics. However, if there is selection around the threshold, we would expect differences in the inflow probability with higher inflow probabilities above the threshold, and we would expect differences in observable characteristics between treated and non-treated individuals. In order to test for this we i) report the smoothness of individuals characteristics around the threshold

[7] We thank an anonymous referee for pointing this out.

as suggested by DiNardo and Lee (2004) and ii) apply the density test proposed by McCrary (2008).

Table 3 contains descriptive characteristics of job losers below and above the threshold. In order to examine whether there are differences between both groups we perform *t-tests* on the quality of means and *Kolmogornov-Smirnov-tests* of the equality of the distributions of the relevant variables; p-values for both tests are reported in the Table. We find that individuals are very similar with respect to (nearly) all of the variables. For men we do not observe any significant difference in the distribution of the different variables. There is only one slightly significant difference in the share of individuals with a high school degree. For women we find some significant mean differences in marital status, number of children below the age of 10, which can be explained by the fact that individuals below the threshold are on average one year younger, and school degree. Additionally, the time spent in employment in the years 4-7 before entering unemployment is approximately one month larger for women above the threshold which makes not only the mean but also the distribution significantly different. These differences become largely insignificant once we use only women in the age range from 44 to 46 years as we do for men. Comparisons at other 'artificial' thresholds which are unrelated to ours (e.g. at age 41) show similar patterns. This indicates that these significant differences may be driven by age and not by selection around the threshold.

To explore the selectivity issue in more detail we apply the density test suggested by McCrary (2008). It is based on an estimator for the discontinuity at the cutoff in the density function of the running variable. The test is implemented as a Wald test of the null hypothesis that the discontinuity is zero. It can be seen as an extension of the local linear density estimator proceeding in two steps. In the first step, one obtains a finely gridded histogram and in the second step, one smoothes the histogram using local linear regression, separately on either side of the cutoff. The density test is informative if the manipulation is monotonic, i.e. if agents adjust the running variable in one direction only which would clearly be the case in our application. Figure 1 presents the first-step histogram along with estimation results for the height difference in the distributions. For men the estimated height difference between the group below and above the age threshold is 0.178 (standard error: 0.120) and not statistically different from zero. The same holds true for women (difference: -0.030, standard error: 0.105) which makes us comfortable in arguing that there is no manipulation in the running variable. Both, the comparison of the observable characteristics and the density test indicate that workers do not delay the timing of job separation to benefit from a longer entitlement period.

Before we provide some intuition why this might be the case we make a last validation exercise and compare the time between the end of the last employment spell and the start of the unemployment spell below and above the threshold. If there would be any strategic behavior we would expect people below the threshold to have a longer duration in between the two spells. We find that on average the time between the end of the employment and the beginning of the unemployment spell is below three days and there are no systematic differences in this variable below and above the threshold.[8]

Looking once again at the institutional setting shows that this finding is not surprising. In order to avoid benefit sanctions individuals are required to register as a job-seeker with the local labor office already three months before an employment contract ends (§38, Social Code III). If the time between the moment they are informed about the termination and the actual end of the contract is less than three months they have to register within three days. If individuals do not comply with this rule they can be sanctioned with benefit withdrawal of up to 12 weeks (§144, Social Code III). Apart from this institutional feature, the probability of benefiting from the additional six months of benefit duration is quite low at the beginning of the unemployment spell, because many individuals find a new job within the first 12 months of unemployment. Therefore, we find convincing evidence that there is no manipulation in the running variable, which allows us to identify causal effects of the extended benefit duration.

5 Results

5.1 Descriptive Evidence

We first provide some descriptive evidence on the effect of extended benefit duration based on the observed unemployment duration for different groups by age of entering unemployment. For both men and women the average unemployment duration is higher for the treated who enter unemployment after the age of 45 (with 18 months of potential benefit duration) compared to their younger counterparts below the age threshold (with 12 months of potential benefit duration). In addition, there is a slight increase in unemployment duration for the treated as they get older (see Figure B.3 in the Supplementary Appendix). We also estimate a logit model for the probability of finding a job within a given interval after entering unemployment. Table 4 reports the results showing that being eligible for six additional months of benefits significantly

[8]For additional information see Figure B.2 in the Supplementary Appendix which shows the difference along the age distribution and the relevant test statistics.

reduces the probability of finding a job within 6, 12 and 18 months, respectively. For example, the estimates for the job finding probability within 12 months imply that extended benefits are associated with a reduced probability of around 23% for men and around 34% for women.

Since the impact of a longer benefit entitlement may vary over time spent in unemployment, it is informative to look at the empirical hazard rates. Figure 2 shows a spike in the probability of leaving unemployment for a job around month 12, the last month of benefit receipt for the younger group, which is larger for women. However, the increase in the exit probability is rather moderate. This might be due to the fact that transfers do not drop to zero once the receipt of UB expires (see Section 3.1). Especially if the household does not have any other sources of income, the drop is relatively small compared to other countries like the U.S.

5.2 Specification and Estimation of the Model

We estimate the econometric model described in Section 4.1 for men and women separately under two different specifications regarding the way we capture the treatment effect. In the first specification, we include a dummy for being treated which identifies an average effect of extending benefits on the three outcomes of interest (unemployment transition, employment transition and re-employment wages) imposing a common effect of the treatment over the unemployment duration. In the second specification, we relax this restriction by interacting the treatment dummy with elapsed unemployment duration for the unemployment transition and with previous unemployment duration for the employment transition and the wage equation. We specify unemployment duration flexibly by monthly intervals until month 18. For longer durations we consider group intervals for months 19 to 24, 25 to 30 and 31 to 36 because the cell size by treatment and control group becomes small. With this specification we can identify the effect of treatment on the exit rate from unemployment at different lengths of unemployment. Most importantly, we can also identify the effect of treatment at different lengths of completed unemployment duration on subsequent job match quality.

The Log-Likelihood of the estimation with unobserved heterogeneity clearly indicates an improved model-fit compared to the model without unobserved heterogeneity, so we focus the discussion on these results.[9] The preferred specification for the flexible model allowing for heterogeneous treatment effects is based on five mass points for males and on three mass points

[9]For example, the Log-Likelihood significantly increases from $-10,497.43$ to $-10,144.66$ for men in the restricted model and from $-10,461.52$ to $-10,110.07$ in the model including interaction effects of the treatment dummy with elapsed unemployment duration. The results for the independent transitions without controlling for unobserved heterogeneity are not reported but they are available from the authors.

for females. For both men and women, our results indicate that unobserved characteristics influencing unemployment duration and employment stability are negatively correlated, whereby this correlation is close to zero for men. Moreover, individuals with unobserved characteristics associated with higher wages tend to have longer employment spells, while the correlation between unobserved characteristics of the wage equation and the probability of leaving unemployment for a job is positive (see Table A.1 in the Supplementary Appendix). Finally, in the case of male workers, we find evidence that the error term of the wage equation follows a two dimensional mixture of normals, while this is not the case for females.

5.3 Homogeneous Treatment Effect

Starting with the homogenous treatment effect the results in Table 5 show that eligibility to extended unemployment benefits (Treated) has a negative effect on the unemployment transition for both males and females. We report estimates under different specifications of the discrete distribution of unobserved heterogeneity. The average treatment effect of extended benefits is significant at the 5% level in all specifications for females. The point estimates are quite similar and they suggest that the probability of finding a job in a given month t is about 26% lower for those females who are eligible to extended benefits compared to the non-eligible ones. For males the coefficient of being treated on the unemployment transition remains negative but loses its significance once we increase the mass points from 4 to 5, which goes along with a better model fit, evaluated on the basis of the AIC.[10]

The average effect of extended benefit duration on the employment transition is not significantly different from zero for both genders. The finding of no significant average effect of changes in benefit duration on employment stability is in line with the studies by Card et al. (2007a) and Van Ours and Vodopivec (2008). We also consider the wages received in the sample of re-employed individuals. If eligibility of additional six months of unemployment benefits induces unemployed individuals to set higher reservation wages, we may observe higher wages for the treatment group. Our results show that benefit extension has a positive average effect on wages of around 3.7% for men, but this effect is not statistically significant. For females, the point estimate is negative but it is also not statistically different from zero. Centeno and Novo (2009) find also small or no effect of benefit duration on wages in Portugal.

[10]The coefficients of the complete model are reported in the Supplementary Appendix in Tables A.2, A.3 and A.4 for the unemployment duration, the employment transitions and the wage equation, respectively.

5.4 Time-Varying Treatment Effects

5.4.1 Unemployment Hazard

In Table 6 we present the heterogenous treatment results from the interaction of the treatment dummy with unemployment duration grouped in intervals. For both males and females we observe very similar patterns in the exit rate from unemployment. There are three different effects worth emphasizing which are in line with the discussion in Section 2. First, we find evidence for differences in the exit rate at the beginning of the unemployment spell between the two groups. The treated, who receive extended benefits for 18 months, exhibit a lower exit rate compared to the controls with 12 months of benefit duration. This difference is significant in months 2, 3 and 6 for men, and in months 1, 2 and 6 for women. Second, starting from month 8 the control group exhibits a significantly higher exit rate from unemployment as it approaches benefit expiration. This results in an increasing difference in the exit rate between the two groups with a peak in months 11 and 12.[11] Third, the difference in the exit rates between the two groups remains negative also after month 12 and it starts decreasing towards zero as they approach the benefit expiration of the treated group in month 18. A joint test for the coefficients for the months 8 to 15 suggests that they are jointly significant at the 10% level for men and at a 1% level for women.

These patterns suggest that both groups react to the incentives induced by the design of the UI system with limited benefit duration. The control group exhibits higher unemployment exit rates from the beginning of the spell with a spike close to benefit expiration around month 12, which leads to the largest difference with the treated group. The treated group exhibits lower exit rates all the way up to month 15 with an increasing exit rate closer to benefit expiration in month 18, which reduces the difference in the exit rate from unemployment with the control group to zero. We also find that the difference in the exit rate around benefit expiration in month 12 is larger for females than males, which is consistent with the evidence provided by Roed and Zhang (2003) for Norway. For men we find an additional significant difference at month 18, which we do not find for women. This effect is in contrast to our expectations. However, investigating further this difference it turns out that both groups exhibit a persistent drop in the exit rate from unemployment, which occurs at month 18 for the treated and at month 19 for the controls. Given these similar patterns over time, we consider the significant

[11] Because we define unemployment by the status in the middle of the month there could be a discrepancy of about one month for those spells which start right after the middle of the month. For this reason the spike may not coincide exactly with month 12.

difference in month 18 the result of a small difference in the timing of the response between the two groups, which does not seem to be driven by any behavioral or institutional reasons.

After both groups have exhausted their benefits we should observe no difference in their exit rate from unemployment. For females we indeed do not find any significant differences in the probability of finding a job after 18 months. For men we observe a lower probability of finding a job for the months 19 to 24, which is significantly different from zero at the 10% level. However, if we conduct a joint test for the coefficients after 18 months of unemployment, this test suggests that the differences between individuals with 12 and 18 months of eligibility for unemployment benefits do not differ from each other significantly.

5.4.2 Post-Unemployment Outcomes

We turn now to the effect of benefit extension on post-unemployment outcomes allowing for heterogenous effects by the length of the previous unemployment spell. As we discussed in Section 2 there are two periods of interest. The first is the period until month 12 where both the treated and controls receive benefits, while the control group gets closer to the time benefits expire. The second period is the one between months 12 and 18 in which only the treated group above age 45 receives benefits.

For men, we observe in Table 6 a negative effect of being treated on the exit rate from employment and a positive effect on re-employment wages for jobs obtained close to the time benefits expire for the control group (month 12) and close to the time benefits expire for the treated group (month 18). The negative effect on the employment hazard during the first period (months 9 to 11) suggests that even though both the treated and the control groups find a job while still being benefit recipients, the control group obtains jobs that last for a shorter period as they accept them closer to benefits expiration. The impact at 10 months of elapsed unemployment duration is significant at a 10% level. The effect in the second period (months 16 to 18) is also negative suggesting a lower hazard from employment for the treated group with a significant effect at 16 months of elapsed unemployment duration. Moreover, a joint test indicates that the effects in months 16 to 18 are significantly different from zero at a 10% level. For wages, there is an increasing positive difference for the treated from 3.1% in month 7 to 30% in month 9 and 17% in month 10, which are significantly different from zero. For jobs obtained close to benefit expiration of the treated group (months 16 and 17), we also find that wages are higher for the treated compared to the control group although these effects are not statistically significant.[12]

These findings suggest that jobs that are found while still receiving benefits last longer and are better paid jobs compared to those found after remaining in unemployment for the same period but with benefits expiring or without benefits. We observe, therefore, that although both groups react to benefit expiration by exiting faster from unemployment, the treated group with longer benefit duration obtains jobs with higher employment stability and higher wages, while the control group faces higher employment instability and lower wages.[13]

For women, although we find that their unemployment exit rate reacts very strongly to benefit expiration we do not find significant effects on post-unemployment outcomes. From Table 6 we observe a negative treatment effect on the employment hazard for jobs found between 12 to 16 months in unemployment. However, none of these coefficients is statistically significant. In the model without controlling for unobserved heterogeneity we get a significant difference in the employment stability at the 10% level for females who found a job at month 16. Once we control for dynamic selection, this coefficient decreases slightly and looses its significance. For wages we do not find any effect of benefit extension either before or after benefit exhaustion.

5.5 Simulation of Treatment Effects

In order to evaluate the importance of being eligible for additional 6 months of unemployment benefits on the expected unemployment and employment duration, we simulate the corresponding treatment effects. The simulations are conducted for the average unemployed individual in our sample and are based on the estimated coefficients discussed above. In every simulation we calculate the difference in expected unemployment and employment durations conditional on 12 versus 18 months of benefit eligibility. The simulations for the model with a homogenous treatment effect suggest that for men receiving 18 instead of 12 months of unemployment benefits increases the expected unemployment duration from 11.2 to 13.5 months. However, these estimates are based on insignificant coefficient for the treatment indicator. For females, we get an increase in the expected unemployment duration from 8.8 to 11.3 months. If we instead

[12]Note that we do not measure hourly wages but daily wages in our data. Therefore, the measured impact might be a mixture of working hours and hourly wages. Moreover, we measure the wage at the beginning of the employment spell, which implies that the observed effect does not have to hold for the entire duration of the post-unemployment job.

[13]The difference in the job match quality between the treated and the controls may be underestimated at the time of benefit expiration because some workers in the control group whose benefits expire at month 12 may have already obtained a job before benefit expiration (e.g. in months 9 and 10) but they could have postponed the starting date to gain from the remaining months of UI entitlement until month 12. Boone and van Ours (2009) suggest that the spike in the unemployment hazard rate can be partly explained by this behavior.

simulate the same effects based on the model with heterogenous treatment effects, the results indicate that the expected unemployment duration increases for men from 9.7 to 13.8 months. For females, the expected unemployment duration increases from 9.6 to 12.5 months. For the employment duration, in the specification based on an homogenous treatment effect, we do not find evidence for an impact of receiving 6 additional months of unemployment benefits on the expected employment duration. However, for male workers, we find some evidence for positive treatment effects, depending on the time the job seekers find a job. Our simulations suggest that individuals who leave unemployment for a job in month 10 stay employed for around 8.5 months if they have 2 remaining months of unemployment benefits (controls) and 17.4 months in case of 8 remaining months of eligibility (treated). If we calculate the corresponding effects for exits from unemployment after 16 months of unemployment the estimates indicate that the expected employment duration increases from 17.2 to 27.5 months.

5.6 Sensitivity

We perform additional analyses to further investigate the sensitivity of our results. First, we estimate the model for men allowing for a two dimensional mixture of normals in the error term of the wage equation and we find very similar effects as in the model in which we assume normality. If anything, the effects with the mixture of normals lead to slightly larger effects (see Table A.5 in the Supplementary Appendix). For women, we could not find evidence that the error term in the wage equation follows a two dimensional mixture of normals.

Second, we perform the analysis on the inflow sample of the unemployed without the condition on having worked for 12 months in regular employment in the last year prior to entering unemployment. The main results remain qualitatively similar, although after conditioning the sample on past employment we obtain more precise estimates since the new benefit entitlement leads to a uniform increase in potential benefit duration for the unemployed worker above the age threshold from 12 to 18 months. Third, we estimate both transitions after considering exits to inactivity (out of labor force) as right-censored spells. In the analysis so far, those who exit the labor force were considered as continued unemployment spells since we are interested in the time until re-employment. Our results for the unemployment hazard are robust to this sampling strategy. By censoring spells of unemployed workers who exit the labor force we find even a larger spike at benefit exhaustion. Similarly, for the employment hazard we find similar effects with those of the model which we present. Fourth, we check the sensitivity of our estimates to the functional form assumption by estimating the model with a clog-log specification instead of the logistic. We find that the results are not sensitive to this choice. Finally, we estimate the model considering the exit rate out of the subsequent job instead of overall employment duration, which includes job-to-job transitions. Again our findings are very similar for both the unemployment and employment transitions.

6 Conclusion

Besides creating only disincentives unemployment insurance may also improve job match quality. The non-stationarity of job search behavior with limited benefit duration may not only lead to increasing unemployment exit rates, as the unemployed approach the time of benefit expiration, but may also affect the quality of the job match if they become less selective close to and after benefit expiration. Exploiting a sharp discontinuity in the maximum benefit duration in Germany, which increases from 12 to 18 months at the age of 45, we investigate the heterogeneous effects of extended benefit duration on unemployment duration and post-unemployment outcomes, such as employment stability and re-unemployment wages.

The analysis addresses the two important selection issues that might invalidate our design. The first is related to the manipulation of the running variable, for which we do not find any evidence, and the second is related to dynamic selection issues, which we take into account by estimating a bivariate discrete-time hazard model jointly with the wage equation allowing for correlated unobserved heterogeneity.

Our findings for the unemployment hazard confirm previous evidence that limited benefit duration is associated with non-stationary search behavior, which leads to a spike around benefit expiration. The results of the main focus of the paper suggest that, due to the non-stationarity of job search, there exist time-dependent heterogeneous effects of extended benefit duration on job match quality. In particular, jobs which are accepted close to and after benefits have expired are associated with a lower stability and lower wages, while those unemployed who exit unemployment when they are still insured and could therefore reject job offers tend to find jobs which last longer and pay higher wages. These results are found to be significant for men but not for women.

A role of policy might be to smooth the transition rate out of unemployment to prevent workers from being forced to obtain low quality jobs. These effects are likely to be mitigated by the possibility to receive unemployment assistance after unemployment insurance runs out, which is the case in Germany. Future research should shed more light on the interaction of unemployment insurance and unemployment assistance on the unemployment exit rate and the post-unemployment outcomes. Understanding for which subgroups of the population the unemployment insurance job matching effect matters more is also another important question. It is also important to emphasize that the finding of positive heterogeneous (over the duration of unemployment) post-unemployment effects of extended benefits might be short-run effects. This is the case because the observation period extends to only 3 years since the inflow into unemployment such that we can only focus on the first observed wage after exit from unemployment and on the duration of the first employment spell. A natural and relevant extension of this study is to investigate longer-term effects of benefit entitlement on employment outcomes with the use of data that allow for a longer observation period.

References

Acemoglu D, Shimer R. 1999. Efficient unemployment insurance. *Journal of Political Economy* **107**: 893–928.

Addison JT, Blackburn ML. 2000. The effects of unemployment insurance on postunemployment earnings. *Labour Economics* **7**: 21–53.

Addison JT, Portugal P. 2008. How do different entitlements to unemployment benefits affect the transitions from unemployment into employment? *Economics Letters* **101**: 206–209.

Belzil C. 1992. On the empirical relationship between unemployment duration, unemployment insurance and voluntary unemployment. *Economics Letters* **39**: 235–240.

Belzil C. 1995. Unemployment insurance and unemployment over time: An analysis with event history data. *Review of Economics and Statistics* **77**: 113–26.

Belzil C. 2001. Unemployment insurance and subsequent job duration: job matching vs. unobserved heterogeneity. *Journal of Applied Econometrics* **16**: 619–633.

Boone J, van Ours JC. 2009. Why is there a spike in the job finding rate at benefit exhaustion? *IZA Discussion Paper* **No.4523**.

Bover O, Arellano M, Bentolila S. 2002. Unemployment duration, benefit duration and the business cycle. *Economic Journal* **112**: 223–265.

Caliendo M, Falk A, Kaiser L, Schneider H, Uhlendorff A, Van den Berg G, KF Zimmermann. 2011. The IZA Evaluation Dataset. *International Journal of Manpower* **32**: 721–752.

Cameron SV, Heckman JJ. 1998. Life cycle schooling and dynamic selection bias: Models and evidence for five cohorts of american males. *Journal of Political Economy* **106**: 262–333.

Card D, Chetty R, Weber A. 2007a. Cash-on-hand and competing models of intertemporal behavior: New evidence from the labor market. *Quarterly Journal of Economics* **122**: 1511–1560.

Card D, Chetty R, Weber A. 2007b. The spike at benefit exhaustion: Leaving the unemployment system or starting a new job? *American Economic Review* **97**: 113–118.

Card D, Levine PB. 2000. Extended benefits and the duration of ui spells: Evidence from the new jersey extended benefit program. *Journal of Public Economics* **78**: 107–138.

Centeno M. 2004. The match quality gains from unemployment insurance. *Journal of Human Resources* **34**: 839–863.

Centeno M, Novo AA. 2009. Unemployment insurance generosity and post-unemployment wages: Quantile treatment effects. *mimeo* .

DiNardo J, Lee D. 2004. Economic impacts of new unionization on private sector employers: 1984-2001. *Quarterly Journal of Economics* **119**: 1383–1441.

Eberwein C, Ham JC, LaLonde R. 1997. The impact of being offered and receiving classroom training on the employment histories of disadvantaged women: Evidence from experimental data. *Review of Economic Studies* **64**: 655–682.

Ehrenberg RG, Oaxaca RL. 1976. Unemployment insurance, duration of unemployment, and subsequent wage gain. *American Economic Review* **66**: 754–766.

Fitzenberger B, Wilke RA. 2010. Unemployment Durations in West-Germany Before and After the Reform of the Unemployment Compensation System during the 1980s. *German Economic Review* **11**: 336–366.

Gaure S, Roed K, Zhang T. 2007. Time and causality: A monte carlo assessment of the timing-events approach. *Journal of Econometrics* **141**: 1159–1195.

Hahn J, Todd P, Van der Klaauw W. 2001. Identification and estimation of treatment effects with a regression-discontinuity design. *Econometrica* **69**: 201–209.

Ham JC, LaLonde R. 1996. The effect of sample selection and initial conditions in duration models: Evidence from experimental data on training. *Econometrica* **64**: 175–205.

Heckman JJ, Singer B. 1984. A method for minimizing the distributional assumptions in econometric models for duration data. *Econometrica* **52**: 271–320.

Hunt J. 1995. The effect of the unemployment compensation on unemployment duration in germany. *Journal of Labor Economics* **13**: 88–120.

Jovanovic B. 1979. Job matching and the theory of turnover. *Journal of Political Economy* **87**: 972–990.

Katz LF, Meyer BD. 1990. The impact of the potential duration of unemployment benefits on the duration of unemployment. *Journal of Public Economics* **41**: 45–72.

Konle-Seidl R, Eichhorst W, Grienberger-Zingerle M. 2010. Activation Policies in Germany: From Status Protection to Basic Income Support. *German Policy Studies* **6**: 59–100.

Lalive R. 2008. How do extended benefits affect unemployment duration? a regression discontinuity approach. *Journal of Econometrics* **142**: 785–806.

Lalive R, Van Ours J, Zweimüller J. 2006. How changes in financial incentives affect the duration of unemployment. *Review of Economic Studies* **73**: 1009–1038.

Lalive R, Zweimüller J. 2004. Benefit entitlement and unemployment duration: The role of policy endogeneity. *Journal of Public Economics* **88**: 2587–2616.

Lise J, Meghir C, Robin JM. 2011. Matching, sorting and wages. *Working Paper* .

Marimon R, Zilibotti F. 1999. Unemployment vs. Mismatch of Talents: Reconsidering Unemployment Benefits. *Economic Journal* **109**: 266–291.

McCrary J. 2008. Manipulation of the running variable in the regression discontinuity design: A density test. *Journal of Econometrics* **142**: 698–714.

Meyer BD. 1990. Unemployment Insurance and Unemployment Spells. *Econometrica* **58**: 757–782.

Mortensen D. 1977. Unemployment insurance and job search decisions. *Industrial and Labor Relations Review* **30**: 505–517.

Roed K, Zhang T. 2003. Does unemployment compensation affect unemployment duration? *Economic Journal* **113**: 190–206.

Schmieder JF, von Wachter T, Bender S. 2012. The effects of unemployment insurance on labour supply and search outcomes. *Quarterly Journal of Economics* **forthcoming**.

Schneider H, Hujer R. 1997. Wirkungen der unterstützungsleistungen auf die arbeitslosigkeitsdauer in der bundesrepublik deutschland: Eine analysye der querschnitts- und längschnittdemension. In R Hujer GW U Rendtel (ed.) *Wirtschafts- und Sozialwissenschaftliche Panel-Studien, Datenstrukturen und Analyseverfahren. Sonderhefte zum Allgemeinen Statistischen Archiv, Bd. 30*. Goettingen: Vandenhoeck and Ruprecht, 71–88.

Steiner V. 2001. Unemployment persistence in the west german labour market: negative duration dependence or sorting? *Oxford Bulletin of Economics and Statistics* **63**: 91–113.

Tatsiramos K. 2009. Unemployment insurance in europe: Unemployment duration and subsequent employment stability. *Journal of the European Economic Association* **7**: 1225–1260.

Tatsiramos K. 2010. Job displacement and the transitions to re-employment and early retirement for non-employed older workers. *European Economic Review* **54**: 517–535.

Van den Berg GJ. 1990. Nonstationarity in job search theory. *Review of Economic Studies* **57**: 255–277.

Van den Berg GJ, Ridder G. 1998. An empirical equilibrium search model of the labor market. *Econometrica* **66**: 1183–1221.

Van Ours J, Vodopivec M. 2006. How shortening the potential duration of unemployment benefits entitlement affects the duration of unemployment: evidence from a natural experiment. *Journal of Labor Economics* **24**: 351–378.

Van Ours J, Vodopivec M. 2008. Does reducing unemployment insurance generosity reduce job match quality? *Journal of Public Economics* **92**: 684–695.

Van Vuuren A, Van den Berg GJ, Ridder G. 2000. Measuring the equilibrium effects of unemployment benefits dispersion. *Journal of Applied Econometrics* **15**: 547–574.

Winter-Ebmer R. 1998. Potential unemployment benefit duration and spell length: Lessons from a quasi-experiment in austria. *Oxford Bulletin of Economics and Statistics* **60**: 33–45.

Tables

Table 1: Maximum Duration of Unemployment Benefit

Length of Benefit Entitlement (in months)	Age (in years)	Months worked in last 7 years
6	-	12
8	-	16
10	-	20
12	-	24
14	45	28
16	45	32
18	45	36
20	47	40
22	47	44
24	52	48
26	52	52
28	57	56
30	57	60
32	57	64

Source: Social Code III (§117 et seq.)

Table 2: Number of Observations and Transitions - Below and Above the Threshold

	Men			Women		
	Below	Above	Total	Below	Above	Total
Observations	1,147	1,094	2,241	1,442	1,334	2,776
Transitions						
From UE to E	857	780	1,637	1,110	977	2,087
in month 1	221	201	422	232	205	437
2	97	77	174	84	56	140
3	74	51	125	58	67	125
4	42	54	96	53	42	95
5	32	34	66	46	46	92
6	34	26	60	46	34	80
7	27	36	63	31	32	63
8	25	23	48	36	29	65
9	15	16	31	39	34	73
10	19	18	37	41	26	67
11	27	11	38	64	23	87
12	32	18	50	35	22	57
13	17	18	35	23	28	51
14	20	17	37	35	24	59
15	14	14	28	26	16	42
16	12	17	29	26	23	49
17	11	15	26	20	30	50
18	16	8	24	22	22	44
19-24	52	48	100	81	100	181
25-30	37	47	84	68	65	133
31-36	33	31	64	44	53	97
Censored	290	314	604	332	357	689
From E to UE	463	431	894	474	411	885
Censored	394	349	743	636	566	1,202

Note: The table shows the number of transitions from unemployment (UE) to employment (E) and the transitions back to unemployment for those who exit unemployment and are not censored for the two samples.
This is an inflow sample into unemployment aged 44-46 for men and aged 43.5-46.5 for women conditional on having been employed for 36 months in the last seven years and on been working for 12 months in regular employment in the year prior to entering unemployment. For men the age range below the threshold of 45 is 44-44.99 and above is 45-45.99. For women the age range is 43.5-44.99 (below) and 45-46.5 (above).

Table 3: Descriptive Statistics and Equality Tests by Gender

Age Group	Men				Women			
	Below	Above	p-values t-test[1]	p-values KS-test[2]	Below	Above	p-values t-test[1]	p-values KS-test[2]
N	1,147	1,094			1,442	1,334		
Age (in years)	44.49	45.50	0.00	0.00	44.24	45.74	0.00	0.00
Marital status								
Married	0.63	0.64	0.70	1.00	0.62	0.65	0.07	0.44
Nationality								
Non-German	0.09	0.09	0.71	1.00	0.06	0.05	0.33	1.00
Migration background	0.04	0.03	0.72	1.00	0.02	0.02	0.35	1.00
Children $\leq$ 10 years	0.18	0.15	0.12	0.90	0.07	0.05	0.01	0.81
School Degree								
No degree	0.08	0.09	0.47	1.00	0.05	0.06	0.45	1.00
Low	0.56	0.54	0.53	1.00	0.47	0.52	0.01	0.05
Medium	0.18	0.15	0.13	0.92	0.28	0.27	0.26	0.97
High	0.18	0.21	0.09	0.77	0.19	0.15	0.01	0.22
Apprenticeship (yes)	0.79	0.78	0.43	1.00	0.81	0.81	0.58	1.00
University Degree (yes)	0.14	0.16	0.16	0.96	0.12	0.09	0.06	0.89
Occupational Group								
Agriculture, Other	0.02	0.02	0.76	1.00	0.01	0.01	0.69	1.00
Manufacturing	0.46	0.46	0.91	1.00	0.15	0.15	0.50	1.00
Technical Occupations	0.08	0.08	0.80	1.00	0.02	0.02	0.41	1.00
Services	0.44	0.44	0.91	1.00	0.82	0.82	0.83	1.00
Labor Market History								
Last daily income (in Euro)	81.44	82.00	0.72	0.34	53.22	52.63	0.63	0.53
Employment last 3 years (in months)	33.82	33.80	0.91	0.98	33.84	33.91	0.67	1.00
Employment last 4-7 years (in months)	39.41	39.44	0.95	0.99	36.15	37.31	0.03	0.00
Unemployed last 7 years (in months)	3.21	3.16	0.80	1.00	2.87	2.64	0.20	0.81
Months in employment - Year t-1	12.00	12.00		1.00	12.00	12.00		1.00
t-2	11.11	11.19	0.45	1.00	11.22	11.24	0.76	0.99
t-3	10.71	10.62	0.45	0.99	10.63	10.68	0.66	1.00
Year cohort								
2001	0.26	0.26	0.84	1.00	0.26	0.28	0.14	0.79
2002	0.37	0.36	0.70	1.00	0.34	0.33	0.55	1.00
2003	0.37	0.38	0.57	1.00	0.40	0.39	0.45	1.00

Source: IZA Evaluation Data Set, own calculations.
[1] p-value for t-test of mean equality in the characteristics between groups below and above the age threshold.
[2] p-value for Kolmogorov-Smirnov test of distribution equality between groups below and above the age threshold.

Table 4: Logit Results for the Effect of Extended Benefits on the Unemployment Exit Rate

Male Sample

	6 months		12 months		18 months	
	Coeff	SE	Coeff	SE	Coeff	SE
Treated	−0.266	0.177	−0.319	0.175*	−0.335	0.181*

Female Sample

	6 months		12 months		18 months	
	Coeff	SE	Coeff	SE	Coeff	SE
Treated	−0.423	0.161**	−0.423	0.154***	−0.387	0.158**

Note: ***/**/* indicate significance at the 1%/5%/10% levels. Logistic regressions for the probability of exiting unemployment within 6, 12 or 18 months. Treated denotes those who are above age 45 at the time of entering into unemployment and are eligible for 18 months of unemployment benefits. The reference group includes those who enter into unemployment below age 45 and are eligible for 12 months of benefits. The model includes a set of additional controls which are listed in Table 3.

Table 5: Estimation Results of Eligibility to Extended Benefits for Homogeneous Treatment

			Male Sample			
	Unemployment 4 mass points		Employment 4 mass points		Wages 4 mass points	
	Coeff	SE	Coeff	SE	Coeff	SE
Treated	−0.380	0.185**	−0.050	0.147	0.037	0.029
Log-Likelihood			-10155.69			
	Unemployment 5 mass points		Employment 5 mass points		Wages 5 mass points	
	Coeff	SE	Coeff	SE	Coeff	SE
Treated	−0.217	0.148	0.012	0.152	0.037	0.027
Log-Likelihood			-10144.66			
			Female Sample			
	Unemployment 3 mass points		Employment 3 mass points		Wages 3 mass points	
	Coeff	SE	Coeff	SE	Coeff	SE
Treated	−0.311	0.132**	−0.098	0.150	−0.034	0.034
Log-Likelihood			-13267.61			
	Unemployment 4 mass points		Employment 4 mass points		Wages 4 mass points	
	Coeff	SE	Coeff	SE	Coeff	SE
Treated	−0.303	0.132**	−0.105	0.147	−0.027	0.033
Log-Likelihood			-13218.64			

Note: ***/**/* indicate significance at the 1%/5%/10% levels. Treated denotes those who are above age 45 at the time of entering into unemployment and are eligible for 18 months of unemployment benefits. The reference group includes those who enter into unemployment below age 45 and are eligible for 12 months of benefits. The model includes a set of controls, which are listed in Table 3, and duration dependence dummies for the unemployment and employment transitions and is estimated jointly allowing for correlated unobserved heterogeneity as described in Section 4.1. For both the male and the female sample we present the results when we allow for different number of mass points for the distribution of unobserved heterogeneity.

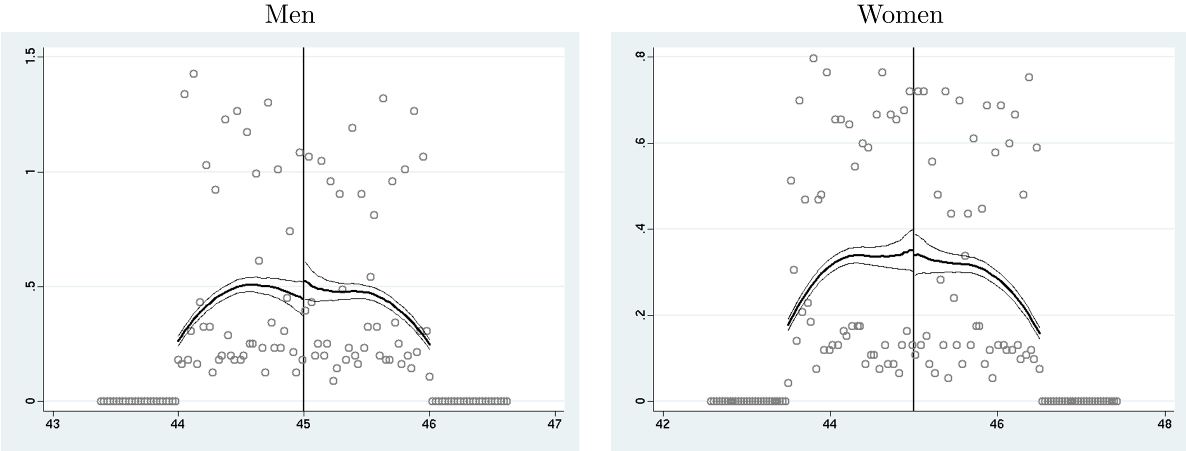
Table 6: Estimation Results of Eligibility to Extended Benefits for Heterogeneous Treatment

	Male Sample						Female Sample					
	Unemployment Transition		Employment Transition		Wages		Unemployment Transition		Employment Transition		Wages	
	Coeff	SE	Coeff	SE	Coeff	SE	Coeff	SE	Coeff	SE	Coeff	SE
Treated * Month 1	−0.281	0.206	0.197	0.196	−0.005	0.033	−0.130	0.170	−0.077	0.206	−0.037	0.045
Treated * Month 2	−0.463	0.253*	−0.106	0.268	−0.003	0.046	−0.488	0.232**	0.085	0.313	−0.030	0.067
Treated * Month 3	−0.698	0.271**	0.084	0.286	−0.024	0.048	0.083	0.239	0.272	0.317	−0.022	0.075
Treated * Month 4	−0.038	0.297	−0.063	0.328	0.074	0.058	−0.328	0.265	−0.308	0.344	0.071	0.085
Treated * Month 5	−0.234	0.334	0.250	0.330	0.009	0.056	−0.110	0.274	0.120	0.357	−0.038	0.083
Treated * Month 6	−0.596	0.348*	−0.020	0.435	−0.004	0.079	−0.479	0.290*	−0.097	0.378	0.006	0.083
Treated * Month 7	0.001	0.341	0.002	0.391	0.031	0.070	−0.162	0.321	−0.183	0.440	−0.149	0.100
Treated * Month 8	−0.376	0.370	−0.085	0.480	0.072	0.091	−0.437	0.317	0.362	0.412	−0.090	0.125
Treated * Month 9	−0.237	0.438	−0.623	0.518	0.305	0.104***	−0.376	0.303	0.171	0.420	−0.029	0.091
Treated * Month 10	−0.360	0.405	−0.856	0.500*	0.173	0.089*	−0.721	0.318**	−0.400	0.438	0.093	0.104
Treated * Month 11	−1.275	0.429***	−0.317	0.602	−0.139	0.101	−1.423	0.308***	0.524	0.397	0.017	0.099
Treated * Month 12	−0.992	0.374***	−0.071	0.510	0.078	0.080	−0.841	0.330**	−0.155	0.508	−0.135	0.140
Treated * Month 13	−0.370	0.414	0.713	0.575	0.130	0.100	−0.104	0.336	−0.407	0.488	−0.098	0.109
Treated * Month 14	−0.605	0.413	−0.700	0.589	0.047	0.097	−0.663	0.316**	−0.585	0.532	0.153	0.100
Treated * Month 15	−0.436	0.451	0.145	0.637	0.121	0.115	−0.771	0.366**	−0.529	0.555	−0.074	0.121
Treated * Month 16	−0.066	0.437	−1.095	0.521**	0.129	0.132	−0.385	0.327	−0.783	0.498	0.092	0.126
Treated * Month 17	−0.081	0.466	−0.864	0.796	0.053	0.151	0.190	0.323	−0.075	0.543	−0.127	0.152
Treated * Month 18	−1.138	0.497**	−1.063	0.749	−0.096	0.121	−0.209	0.336	0.560	0.541	−0.224	0.148
Treated * Month 19-24	−0.520	0.284*	−0.492	0.336	0.007	0.050	0.032	0.190	−0.035	0.285	−0.095	0.069
Treated * Month 25-30	−0.194	0.309	−0.459	0.384	0.070	0.061	−0.211	0.209	−0.481	0.358	−0.106	0.071
Treated * Month 31-36	−0.484	0.339	−0.305	0.668	0.122	0.068*	0.023	0.238	0.082	0.585	0.147	0.114
Log-Likelihood			-10110.07						-13191.72			

Note: ***/**/* indicate significance at the 1%/5%/10% levels. Treated denotes those who are above age 45 at the time of entering into unemployment and are eligible for 18 months of unemployment benefits. The reference group includes those who enter into unemployment below age 45 and are eligible for 12 months of benefits. The months interacted with the treatment indicator denote the month of unemployment in the unemployment transition and the month of having exited unemployment in the employment transition and the wage equation. The model includes a set of controls, which are listed in Table 3, and duration dependence dummies for the unemployment and employment transitions and is estimated jointly allowing for correlated unobserved heterogeneity as described in Section 4.1.

Figures

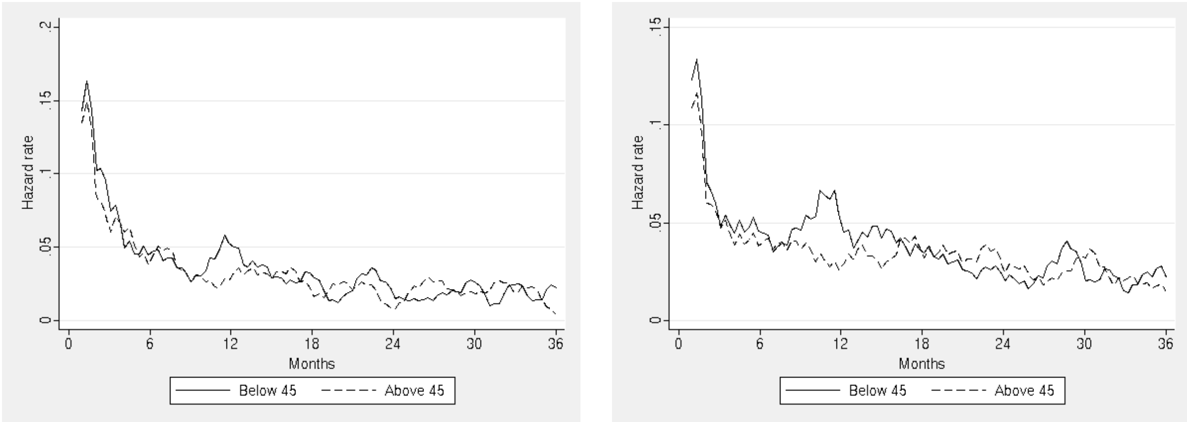
Figure 1: Density Test of Manipulation in the Running Variable



Source: IZA Evaluation Data Set, own calculations.
Note: The density test is based on McCrary(2008) and is implemented using the `DCDensity.ado` routine in `Stata`.
Discontinuity estimates (s.e. in parentheses):
Men: .178 (.120)
Women: -.030 (.105)

Figure 2: Empirical Unemployment Hazard Function by Treatment

Men Women



Source: IZA Evaluation Data Set, own calculations.
Note: Individuals are an inflow sample into unemployment aged 44-46 for men and aged 43.5-46.5 for women conditional on having been employed for 36 months in the last seven years and having worked for 12 months in regular employment in the year prior to entering unemployment. Those who enter unemployment below the age of 45 are eligible to 12 months of benefits, and those above 45 to 18 months of benefits.