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# Does Raising the Early Retirement Age Increase Reemployment of Older Unemployed Workers?

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## Abstract

This study examines an increase in the early retirement age from 60 to 63 for the group of older unemployed workers in Germany. The reform extends the distance to retirement for affected individuals by up to three years. We use this source of exogenous variation to estimate differences-in-differences of the reemployment hazard, thereby accounting for time-invariant unobserved heterogeneity and the possibility of competing risks. Previous to the estimation, we account for selection on observables by implementing matching based on entropy balancing. Our results indicate that - all other things equal - the sub-hazard of reemployment increases by 3.6% for each month by which the early retirement age is raised. The effect is large in absolute and relative terms and implies that unemployment as a pathway into early retirement is less prevalent. Thus, raising the distance to retirement substantially reduces public costs from non-employment.

**Keywords:** Retirement Age; Unemployment Duration; Reemployment; Entropy Balancing; Differences-in-Differences; Competing Risks

**JEL-Classification:** J14, J26, J64

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# 1 Introduction

Unemployment of older workers is often permanent and thus very costly. Raising the early retirement age (ERA) for unemployed individuals may not reduce the burden from social security if these individuals remain unemployed for a longer period instead of retiring. The purpose of this paper is to analyse the consequences of raising the ERA for older unemployed individuals who then face a larger distance to retirement. Our specific research question is whether raising the ERA affects the likelihood of reemployment for older unemployed workers. A higher prevalence of reemployment among treated individuals is likely to reduce public costs because reemployed individuals usually neither receive pension benefits nor unemployment benefits.

The ERA for unemployed individuals was raised in Germany from age 60 to 63. This reform was implemented in monthly steps for the birth cohorts 1946-1948 and phased in between January 2006 and December 2011. No old age pension due to unemployment is available at ages below 63 after full implementation (as of January 2012). Clearly, this policy change delays the availability of pension benefits by three years of age.

The empirical strategy of this paper is based on this natural experiment where exogenous variation from the policy change is exploited to estimate the effect of raising the ERA on reemployment prospects following a late-career job loss. We use survey data from the German Socio-Economic Panel (SOEP) for the years 1991 - 2012. Not only do these data contain detailed labour force states but, even more important, they provide rich information on the individual- and household-specific context. This enables us to produce estimates conditional on variables that control for selective unemployment inflow behaviour. We extract unemployment spells to examine potential transitions into employment, departing from a job loss where individuals face a situation of unemployment. Once entering unemployment, the trade-off is between three alternatives: *(i)* returning to a job (reemployment), *(ii)* remaining unemployed until retirement benefits are available or *(iii)* to exit the labour force.

The results suggest that raising the ERA does have a positive causal impact on reemployment prospects of older unemployed workers. Raising the ERA by one month increases the hazard to exit from unemployment into employment by roughly 3.6% for each month by which the ERA is raised, holding everything else constant. Assuming a binary treatment, the effect of the reform amounts to an increase in the reemployment hazard of about 180%. The effect is large in absolute and relative terms and implies that unemployment as a pathway into early retirement is less prevalent. Thus, raising the distance to retirement substantially reduces public costs from non-employment.

Our study is related to several strands of the literature. First of all, we are concerned with the distance to retirement and how it affects employment decisions. In an early contribution, Seater (1977) constructs a life-cycle model of labour supply and job search that reveals a declining job search intensity during older ages (eventually reaching zero in the moment of death). Recent work by Hairault et al. (2010) is more explicit about the distance effect. They show how the likelihood of employment for older workers is significantly affected by the distance to retirement to conclude that “the time to retirement is [then] key to understanding older workers’ employment”. Our paper is built on the grounds of this literature which yields the hypothesis, theoretically and empirically, that raising the distance to retirement for unemployed individuals may set incentives for reemployment. We add to this literature in that we provide quasi-experimental evidence on the relationship between the distance to retirement and employment prospects. The striking feature of our main finding of a positive impact on reemployment not only implies a reduction in payments of pension benefits but also reduces pressure from the unemployment insurance.

Second, the study that is probably the closest to ours is the one by Staubli and Zweimüller (2013), who examine the consequences of raising the ERA in Austria for relevant labour market outcomes. They find a considerable increase of both employment and unemployment among affected individuals and employment effects are the largest for high-wage and healthy workers. Our study carries on this research but, in contrast to Staubli and Zweimüller (2013), we explicitly take the perspective of older *unemployed* individuals to shed light on the impact of a raised ERA on reemployment. The explicit interplay be-

tween pension insurance and unemployment insurance is highly policy-relevant and is, to the best of our knowledge, a novel feature. The importance of this research stems from the fact that many countries attempt to increase labor force participation of older workers by pensions reforms, and the unemployment insurance typically plays a crucial role in this effort.

Third, a job loss has manifold implications for subsequent career patterns. Size and direction of a potential impact on reemployment outcomes is not at all clear a priori. The positive distance effect of a raised ERA on reemployment may be compromised by numerous aspects. One of them is a negative relationship between age and reemployment (see e.g. Chan and Stevens, 2001). Returning to a job may be an increasingly burdensome event as individuals grow older (Hernaes et al., 2013). Moreover, a job loss itself may alleviate the attractiveness of subsequent work due to considerable earnings reductions (Jacobson et al., 1993). If the expected future payoff from work decreases then the alternative of remaining unemployed is relatively more attractive (Chan and Stevens, 1999). Additionally, reemployment of displaced older workers is less likely in countries where unemployment insurance benefits are more generous (Tatsiramos, 2010). All these findings indicate that the question of whether raising the ERA affects the probability of reemployment depends on numerous aspects. We complement this literature as we employ an empirical strategy that is suitable to isolate the effect of a raised ERA on reemployment from mechanisms that may offset each other.

Finally, the methodology of this paper is indebted to a range of identification issues with respect to observable and unobservable heterogeneity. The estimation strategy combines selection on observables in the first step with time-invariant selection on unobservables in the second step. First, we implement matching based on entropy balancing to achieve exact covariate balance (Hainmueller, 2012; Hainmueller and Xu, 2013). This is of particular importance to reduce model dependence in the subsequent estimation of the treatment effect and improves on potentially poor properties of non-experimental estimates (Heckman et al., 1997). Second, we estimate differences-in-differences for a sample of recently unemployed individuals (age 55-63) following e.g. Hunt (1995), to identify the causal impact of

raising the ERA on the probability of reemployment.

The remainder of this paper is structured as follows. Section 2 provides an overview on recent employment patterns and the institutional setting. Section 3 describes data, identification, and empirical strategy. Section 4 provides empirical results and section 5 concludes.

## **2 Recent Employment Trends and Institutional Setting**

### **2.1 Recent Trends in Employment and Retirement Behaviour**

Labour force participation of older workers is gaining importance in Germany. While the number of individuals in gainful employment has increased by remarkable 9% between 2005 and 2011 (Federal Statistical Office, 2012), more than one half of this increase (about 1.8 Mio.) can be attributed to the group of older individuals between age 55 to 65. Furthermore, unemployment among older individuals (age 50 - 64) declined by 25% from 1.2 Mio. (2005) to 0.9 Mio. (2011) (Federal Employment Agency, 2012). Meanwhile, the average duration of completed unemployment spells decreased from 38.4 weeks in 2005 to 36.9 weeks in 2011 (Federal Employment Agency, 2012). And finally, the take-up rate of old age pensions due to unemployment has decreased by more than 40% from 135,991 (2005) to 59,027 (2011) (German Federal Pension Insurance, 2013), indicating that direct transitions from unemployment into retirement are less frequent. In summary, increasing labour force participation among older workers with more gainful employment/less unemployment, shorter unemployment spells and fewer old age pensions due to unemployment highlight the importance of this age group. All these developments are potentially linked to a raised ERA for old age pensions due to unemployment.

### **2.2 Institutional Setting**

Recently unemployed older individuals face alternative exit routes out of unemployment. Whether individuals leave unemployment and when such transitions take place are likely

to be incentivised by the unemployment insurance and - specifically for older workers - the pension insurance. In what follows, we briefly outline the administrative rules in Germany that are relevant for this study.

### **2.2.1 Unemployment Insurance**

The duration and level of unemployment benefit entitlements influences the incentives for reemployment (Hunt, 1995; Lalive and Zweimüller, 2004; Kyryä and Ollikainen, 2008; Lalive, 2008; Grogger and Wunsch, 2012). The current replacement rate of the unemployment insurance in Germany is 67% of the previous net income if dependent children live in the household and 60% without dependent children.<sup>1</sup> As of January 2008, the duration of benefit entitlements is limited to 18 months (age 55 - 57) and 24 months for unemployed individuals of age 58 and older.<sup>2</sup> Both level and duration of unemployment benefits have been subject to reforms in Germany. However, such modifications for the German case as examined by Hunt (1995) and Grogger and Wunsch (2012) do not challenge the identification of the effect of raising the ERA on reemployment if both control- and treatment group are affected equally. We need this requirement to hold because our estimation strategy builds on differences-in-differences and otherwise the common trend assumption would be compromised.<sup>3</sup>

### **2.2.2 Public Pension System and Reform**

This study is concerned about a change in the German public pension system, which has been converted into a pure pay-as-you-go pension system after World War II. In the course of a major reform in 1972, the generosity of the public pension scheme was increased dramatically. Inevitably, this system ran into severe financing problems due to demographic change and major disincentives (see e.g. Börsch-Supan and Schnabel, 1998; Börsch-Supan, 2000).

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<sup>1</sup>See § 149 *SGB III Grundsatz* of the German social security legislation for details about the level of unemployment benefits, i.e. the net replacement rate.

<sup>2</sup>See § 147 *SGB III Grundsatz* of the German social security legislation for details about benefit duration and age.

<sup>3</sup>For a discussion of the common trend assumption, see section 3.4.

A series of reforms, beginning in the early 1990's, was motivated by the demographic change and to make the system suitable for future generations. It seems natural in such a setting, that the burden of a changing old age dependency ratio of pension benefit recipients to contributors must somehow be allocated among those groups who participate in the public pension system.

Therefore a policy reform was introduced into the German social security legislation<sup>4</sup> that reorganises the specific retirement rules for unemployed individuals.<sup>5</sup> The reform was originally enacted in March 2001.

The relevant change is that the former minimum age to receive an old age pension due to unemployment is raised from 60 to 63 years in monthly steps. This increase is realised for the birth cohorts 1946 to 1948 (see table 4 in appendix for the relevant part of *supplement 19 SGB VI* in the German social security legislation). Raising the ERA by one additional month for each month of birth implies that the reform phased in between January 2006 and December 2011. For instance, a person born in April 1947 faces a postponement of 16 months until eligibility for old age pension due to unemployment is achieved. In other words, this person cannot receive pension benefits due to unemployment as formerly starting in April 2007 (age 60) but instead in August 2008 (age 61 years and 4 months).

While the increase in the ERA for old age pensions due to unemployment affects both sexes equally, women still have the option to retire by the age of 60 due to specific legislative rules, i.e. “womens’ pension” (§ 237a SGB VI). However, women are only eligible for this alternative, if they meet certain requirements. Women need a minimum of 15 contribution years and at least ten years of social security contributions after age 40. We identify the impact of the increased ERA using differences-in-differences estimation. This strategy crucially relies on the possibility to distinguish a control group from a treatment group. The existence of a “womens’ pension” classifies eligible women to be a sufficient control group (see section 3.2 for a detailed discussion).

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<sup>4</sup>The legislative change was part of the so-called *Altersvermögensergänzungsgesetz* (AVmEG) from March 2001, as published in *BGBI. I 2001, Nr. 13, p. 403*. The new version of social security code including the corresponding changes was announced on February 19, 2002, and published in *BGBI. I 2002, Nr. 12, p. 754*.

<sup>5</sup>The relevant clause in the German social security legislation (*Sozialgesetzbuch*) is § 237 Absatz 3 SGB VI in combination to *supplement 19 SGB VI*.



## 3 Data and Identification Strategy

### 3.1 Data

For this study we use data provided by the German Socio-Economic Panel (SOEP). It includes about 11,000 German households with some 20,000 individuals to be interviewed in subsequent years. The SOEP is a representative sample of the German population. It allows access to a rich set of socio-demographic characteristics on the individual level as well as on the household level (see Haisken-DeNew and Frick, 2005). We make extensive use of this kind of information as it is highly relevant for retirement behaviour.

The key information used in this paper is based on calendar records of employment status, unemployment status and retirement status. We use precise monthly information which is based on retrospective questions for the panel waves from 1991 to 2012. Individuals who participate in the survey are asked to answer these retrospective questions corresponding to the year before being interviewed. Due to the retrospective information of calendar records, the employment status is only available until 2011 and hence the observation period is 1991 to 2011.

After we restrict our sample to individuals (males and females) of age 55 to 63 (4,014 individuals) we identify only those individuals that enter unemployment in the relevant observation period between 1991 and 2011 (1,304 individuals).<sup>6</sup> Furthermore, civil servants are excluded from the analysis due to different legislative rules. After conditioning on a set of control variables (see table 1), the final estimation sample consists of 971 individuals (19,795 person-months at risk).

### 3.2 The Separation of Control- and Treatment Group

To identify the causal effect of raising the ERA on the reemployment probability of older unemployed individuals, we combine matching based on entropy balancing to the estimation of differences-in-differences.

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<sup>6</sup>The data used in this paper was extracted using the Add-On Package PanelWhiz for Stata. PanelWhiz (<http://www.PanelWhiz.eu>) was written by Dr. John P. Haisken-DeNew ([john@PanelWhiz.eu](mailto:john@PanelWhiz.eu)). See Haisken-DeNew and Hahn (2010) for details. The PanelWhiz generated DO file to retrieve the data used here is available upon request. Any data or computational errors in this paper are the authors'.

Identification is based on the exogenous change in the distance to retirement, i.e. the raised ERA. Whether an individual is subject to the raised ERA is a question of being born before January 1946 (not treated) or after December 1945 (treated). Despite the fact that men and women are affected equally, women who meet certain requirements (section 2) do have an outside option in the shape of a “womens’ pension” which is available at age 60. We exploit this institutional rule to separate the control- and treatment group in the differences-in-differences framework. Women who meet all requirements have a relevant alternative (i.e. an outside option) and old age pensions due to unemployment are effectively irrelevant. Thus, the control group consists of women who are eligible for a “womens’ pension”. Contrarily, the treatment group consists of all men and those women who do not qualify for a “womens’ pension”.

Following this proceeding, the vast majority of women in the sample is part of the control group, leaving only few women in the treatment group.<sup>7</sup> The explanation is that entering unemployment is a prerequisite for individuals to enter the estimation sample. All women that we observe are unemployed by definition and therefore must have a relatively strong labour force attachment due to their employment biographies. At the same time, these women are likely to fulfill the requirements for a “womens’ pension” (e.g. at least 15 contribution years).

### 3.3 Matching Based on Entropy Balancing

Combining matching to the estimation of differences-in-differences improves the potentially poor performance of a non-experimental estimator (Heckman et al., 1997). The principle idea of entropy balancing is to take selection on observables into account via exact covariate balance to reduce model dependence in the subsequent estimation of the treatment effect.

As apparent from table 1, means and standard deviations for most of the conditioning variables differ across treatment and control group (column 3-6). In order to produce a balanced sample with respect to covariates across treatment and control group, we implement a matching method based on entropy balancing as proposed by Hainmueller

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<sup>7</sup>For this reason we exclude the dummy that indicates sex from the estimation and let the treatment indicator capture variation between male and female individuals.

Table 1: Descriptive Statistics.

	Full Sample		Treatment		Control: Unbalanced		Control: Balanced	
	Mean	Std.Dev.	Mean	Std.Dev.	Mean	Std.Dev.	Mean	Std.Dev.
Reduction-Free Ret.-Age	63.65	(1.85)	64.09	(1.50)	62.95	(2.12)	64.09	(1.50)
Age (Years)	57.51	(1.60)	57.54	(1.63)	57.47	(1.54)	57.54	(1.63)
West Germany	0.49	(0.50)	0.48	(0.50)	0.49	(0.50)	0.48	(0.50)
Migrant	0.13	(0.34)	0.15	(0.36)	0.11	(0.32)	0.15	(0.36)
Handicap Level	7.03	(17.95)	7.72	(18.73)	5.94	(16.62)	7.72	(18.73)
Married	0.78	(0.42)	0.81	(0.40)	0.73	(0.44)	0.81	(0.40)
Home Owner	0.50	(0.50)	0.47	(0.50)	0.53	(0.50)	0.47	(0.50)
Dep. Children in HH	0.07	(0.25)	0.09	(0.29)	0.03	(0.17)	0.09	(0.29)
Years of Educ.	11.43	(2.31)	11.63	(2.49)	11.11	(1.95)	11.63	(2.49)
Regional UE Rate	13.05	(4.87)	13.09	(4.81)	12.98	(4.97)	13.09	(4.81)
Particip. in Pre-Ret.	0.07	(0.25)	0.07	(0.26)	0.05	(0.22)	0.07	(0.26)
Unemployed Partner	0.13	(0.34)	0.12	(0.33)	0.14	(0.35)	0.12	(0.33)
Retired Partner	0.19	(0.39)	0.11	(0.31)	0.31	(0.46)	0.11	(0.31)
Past UE Experience	1.56	(2.13)	1.49	(2.08)	1.66	(2.19)	1.49	(2.08)
Income Available	0.86	(0.35)	0.86	(0.35)	0.85	(0.35)	0.86	(0.35)
N	971		594		377		377	

Note: Own calculations based on the SOEP (1991-2012). In subsequent regressions, age-in-year dummies are included to allow the most flexible form. The separation of control- and treatment group almost perfectly determines sex and is thus not part of the estimation.

(2012); Hainmueller and Xu (2013). This method allows to achieve exact balancing of pre-specified sample moments. We balance our sample by the first two moments, i.e. mean and variance, of each conditioning variable to be used in the subsequent estimation of treatment effects. The target is to calibrate individual weights that allow to balance the control group in such a way that the distribution of its covariates is similar to the treatment group. The last two columns in table 1 show that mean and standard deviation are perfectly adjusted for the balanced control group.

The difference in the reduction-free retirement age is likely due to sample composition which underlines the importance of selection on observables and the need for an adequate matching procedure. Previous to the balancing step, the mean for the reduction-free retirement age differs by roughly one year between treatment and control group (column 3 and 5, table 1). After the balancing step, the means are equalised (column 3 and 7). The variable ‘‘Reduction-Free Retirement Age’’ contains information on a recent reform in Germany, which has raised the retirement age without reductions from actuarial adjustments for different types of old age pensions (implementation 1997 - 2004). The reform is important, because it sets financial incentives to retire later, that must be incorporated in our analysis. Relevant for this paper is the increase of the reduction-free retirement

age for old age pensions due to unemployment from 60 to 65 for the birth cohorts 1937 to 1941 in monthly steps.

Pre-retirement agreements are a tool to downsize the workforce. Contracts are canceled by mutual agreements and employees receive a compensation. On this pathway, individuals enter unemployment using the compensation and unemployment benefits to bridge the gap until they reach the ERA. The take-up of pre-retirement programmes may introduce selection and crucially affect reemployment due to the absorbing nature of pre-retirement. However, we observe the participation in such programmes and thus control for the take-up of pre-retirement using a respective dummy indicator.

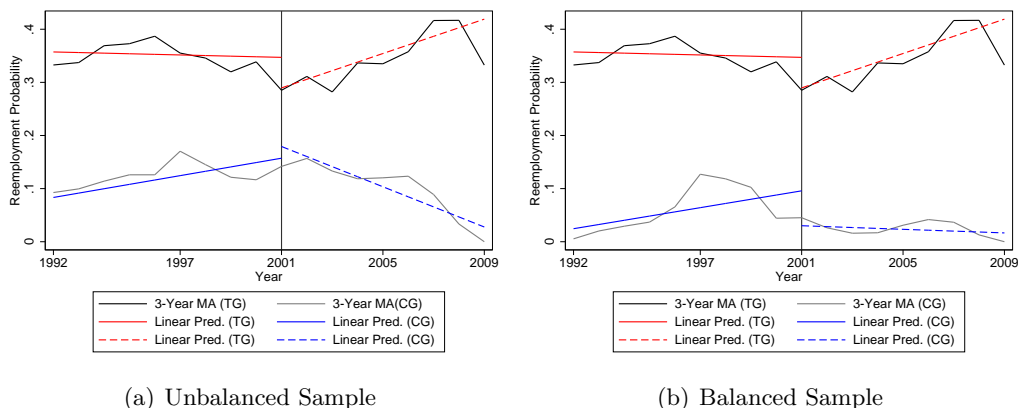
Time fixed-effects are important in the context of this analysis, but we do not include them in our model for a number of reasons. First, we make use of annual regional unemployment rates, which account for most of the variation regarding unemployment and calendar time. Second, since treatment status is a function of the birth cohort, the combination of the treatment dummy and age-in-years perfectly determine calendar time for effectively treated individuals after 2005.

### **3.4 The Common Trend Assumption**

Identification is based on variation over time (before and during/after implementation of the reform) and variation across treated and non-treated individuals. The identifying assumption is that, in absence of the reform, treated individuals had behaved in the same way as otherwise similar non-treated individuals (common trend). Figure 1 illustrates the common trend with respect to the reemployment rate across control- and treatment group. It is important to distinguish between two principal periods because the reform was announced in the year 2001. Consequently, we distinguish the pre-treatment period (1991 - 2000) from the treatment period (2001 - 2011). Here, we implicitly assume that as of 2001 all treated individuals were informed about the raised ERA.

We combine 3-year moving averages of the sample reemployment rate with predictions from a linear regression of the average annual reemployment rate on calendar years. The figures resemble the same principal pattern for both the unbalanced sample (panel a) and

Figure 1: Reemployment across Control- and Treatment Group.



*Note:* Own calculations using the SOEP (1991-2012). 3-year moving averages (3-year-MA) for reemployment of sample individuals are plotted in combination to regression lines (linear predictions), regressing calendar time on the annual reemployment rate. Linear predictions are computed for the periods 1992 - 2001 and 2001 - 2009. CG and TG abbreviate control- and treatment group respectively. We draw on all 1,304 individuals aged 55 - 63 who enter unemployment in the SOEP.

the balanced sample (panel b). While the reemployment rate is much lower within the control group in absolute terms, control- and treatment group follow a common trend with respect to reemployment rates until the announcement period (2001).<sup>8</sup> Thereafter, the reemployment rate clearly exhibits an upward trend in the treatment group which is in sharp contrast to the downward trend within the control group. Our goal is to isolate the difference of this difference as the causal effect of raising the ERA on the reemployment rate in the subsequent estimation step. In contrast to the subsequent estimation, the patterns displayed in figure 1 are not conditional on observed variables such as the regional unemployment rate.

### 3.5 Estimation of Unemployment Duration: Competing Risks

We estimate the probability for reemployment in the context of a model of unemployment duration. By restricting the sample to individuals who enter unemployment during the sample period, we can rule out problems that result from left-censoring, as we precisely know when unemployment spells are to begin. Subsequently, we follow these unemployment spells over time. We either observe spells to end in a transition to reemployment,

<sup>8</sup>This argument is supported by a t-test on the equality of the slope parameters for both groups in the pre-treatment period.

to other undefined states<sup>9</sup>, or to be right-censored. In the case of right-censoring we do not know when a spell is to end since no transition out of unemployment has occurred until the end of the observation period. In the estimation sample, 238 individuals exhibit a transition into employment, 186 exit into other states and 547 spells are right-censored (table 2).

Table 2: Transitions out of Unemployment and Number of Spells.

Transition into...	Treatment Group		Control Group		Total Spells
	BC < 1946	BC > 1945	BC < 1946	BC > 1945	
Employment	112	80	30	16	238
Other State	67	56	43	20	186
Censored	208	71	174	94	547
Group Total	387	207	247	130	
Total Spells	594		377		971

Note: Own calculations based on the SOEP (1991-2012). BC = birth cohort.

The treatment group accounts for a total number of 594 observations (about 61%), while the control group consists of 377 observations (roughly 39%). Table 2 separates these two groups into pre-reform birth cohorts (BC < 1946) and post-reform birth cohorts (BC > 1945). Our interest is in the treatment effect as measured for the post-treatment cohorts (i.e. birth cohorts after 1945) within the treatment group, which accounts for 80 transitions into full-time employment and thus corresponds to about 34% of all transitions into employment (238) and about 8% of the total number of spells (971).

Although we face a binary decision on either staying unemployed or leaving unemployment into employment, linear probability models or probit models are inappropriate. First, the probit framework very much relies on normally distributed duration times. Second, we specifically investigate transitions out of unemployment and therefore need to take duration dependence into account. Third, even if unemployment spells are not terminated by the event of reemployment, right censoring may prevent these spells to last until the end of the observation period for other potentially unknown reasons.<sup>10</sup> To overcome these

<sup>9</sup>In the subsequent estimation, we account for alternative exit routes such as early retirement due to medical indication for disability. While explicitly allowing for this competing risk, a precise definition of these states is not possible.

<sup>10</sup>Although the probit framework may condition on the length of unemployment spells which could be

problems, we apply duration models which are widely used in the relevant literature (see e.g. Lancaster, 1979; Nickell, 1979; Hunt, 1995; Chan and Stevens, 2001; Steiner, 2001; Tatsiramos, 2010). The failure variable is defined as a binary indicator which is zero as long as an individual remains unemployed and takes the value one as soon as an individual enters full-time employment.

Our primary interest is on reemployment but unemployed individuals may exit into alternative states other than employment. Such alternatives are difficult to define and the data does not allow to distinguish such pathways because they are manifold and often subject to exceptional rules. We subsume all alternative pathways out of late-career unemployment under “other exits” and thus it is straightforward to implement a competing risk model between alternatives that are mutually exclusive. We implement a model as proposed by Fine and Gray (1999), where competing failure types, i.e. reemployment and “other exits”, are not assumed to be independent from each other. Thus, the risk set of the hazard function accounts for two competing event types which can be written as

$$\lambda(t) = \lim_{h \rightarrow 0} \frac{\text{Prob}[t \leq T < t + h, \varepsilon = 1 | T > t \text{ or } (T \leq t, \varepsilon \neq 1)]}{h} \quad (1)$$

where  $\varepsilon = 1, 2$  is the event type. Hence, the probability that an individual experiences event type 1, conditional on not having failed before or at least not having experienced event type 1 before, is estimated as the corresponding hazard function

$$\lambda(t|x, \varepsilon) = \lambda_\varepsilon(t) \exp[X\beta_1 + \beta_2 \text{Treat} + \beta_3 \text{AddMonths} + \beta_4 \text{Treat} * \text{AddMonths}] \quad (2)$$

such that we model an event-type-specific baseline hazard  $\lambda_\varepsilon(t)$  and  $X$  is a set of time-invariant covariates. We exploit exogenous variation as exposed by the reform steps to estimate differences-in-differences, where “Treat” is a dummy variable that indicates whether an individual belongs to the treatment group and “AddMonths” is a variable that indicates the number of potential additional months after age 60 until the ERA for an old age pension due to unemployment is achieved. The variable “AddMonths” is a linear combination of calendar time after the policy change including the implementation period, precisely measuring the treatment intensity. Following equation (2), the treatment effect of the included as a regressor, it does not provide appropriate mechanisms to account for right censoring.

policy change is estimated by the coefficient  $\beta_4$ , which measures the difference between treatment group (Treat = 1) and control group (Treat = 0) in the difference before the policy change (AddMonths = 0) and after (AddMonths > 0). From a decision-theoretical point of view, we model individual behaviour taking the perspective of a forward-looking individual from the point in time that corresponds to a job loss. Thus we make use of time-invariant regressors, where the individual socio-demographic situation is captured once an individual enters unemployment but not thereafter.

### 3.5.1 Sensitivity Analysis

To examine the robustness of the main result, we alternate distributional assumptions including a semi-parametric Cox model (Cox, 1972; Cox and Oakes, 1984), a fully parametric Weibull distribution and a frailty distribution (inverse Gaussian) for evolving unobserved heterogeneity (Gutierrez, 2002). Unobserved characteristics such as the attitude towards work may affect reemployment and the sample of unemployed individuals may become more homogeneous in this respect when spell lengths become long. In other words, specific types of individuals will remain in unemployment for a longer period, making the sample more homogeneous after a while.

Moreover, we contrast our results from the continuous treatment to the assumption of a binary treatment which assumes a sharp cut-off from the pre-treatment period to the treatment period. So far, we have implemented the raised ERA exactly as it has been introduced into the German social security legislation. However, we account for selection on observables previous to the estimation via matching based on entropy balancing which assumes a binary treatment. Precisely, calibrating weights for balancing the control group ignores the fact that we actually observe 36 different treatment groups that differ by their treatment intensity. To be consistent with the matching procedure, we drop some information and estimate all models using a binary treatment indicator. This treatment indicator takes the value one for all individuals who are born after December 1945, and is zero for all individuals who are born before January 1946.

For a further robustness check, we construct an artificial reform to test against placebo effects in all previous models. For this purpose, we simulate a situation where the reform



is introduced for the birth cohorts 1940 - 1942 and is fully implemented for the cohorts 1943 - 1945. Furthermore, we exclude all individuals who were treated by the true reform as of 1946. None of the estimated coefficients indicating the treatment effect is statistically significant and thus we can rule out placebo effects.<sup>11</sup>

## 4 Results

### 4.1 The Effect of Raising the ERA on Reemployment

Our interest is on the identification of the causal effect of raising the ERA on the exit rate from unemployment into employment. Table 3 provides estimated coefficients for the sub-hazard of reemployment. All models are estimated in two versions, each being based on the unbalanced and balanced control group respectively. The discussion is restricted to the results on the balanced sample because matching based on entropy balancing reduces model dependency of the treatment effect. The interaction term “Treat \* AddMonths” identifies the treatment effect of the raised ERA on the reemployment probability.

Table 3 provides results of the main specification. While the event of primary interest is the sub-hazard of reemployment, we explicitly allow for competing events that are defined to be all exits other than reemployment. The decision between three alternatives (i.e. remaining unemployed, reemployment, and transition to other states) that are mutually exclusive implies that choosing one alternative is equivalent to not choosing the remaining ones.

The estimated treatment effect of raising the ERA by one month implies that the sub-hazard of reemployment increases by 3.6%, holding everything else constant. As discussed in the previous section, we estimate all models assuming a binary treatment to establish consistency with the matching step that assumes a binary treatment over the observed treatment period from 2006 to 2011 (i.e. sharp cut-off). The corresponding “as-if” binary treatment effect implies that raising the ERA increases the sub-hazard of reemployment by 180%.<sup>12</sup> The probability, that these results occur by chance 1.1% (Continuous Treat-

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<sup>11</sup>The results from this exercise are available from the authors upon request.

<sup>12</sup>Note that the interpretation is restricted to estimations as based on the balanced control group. The

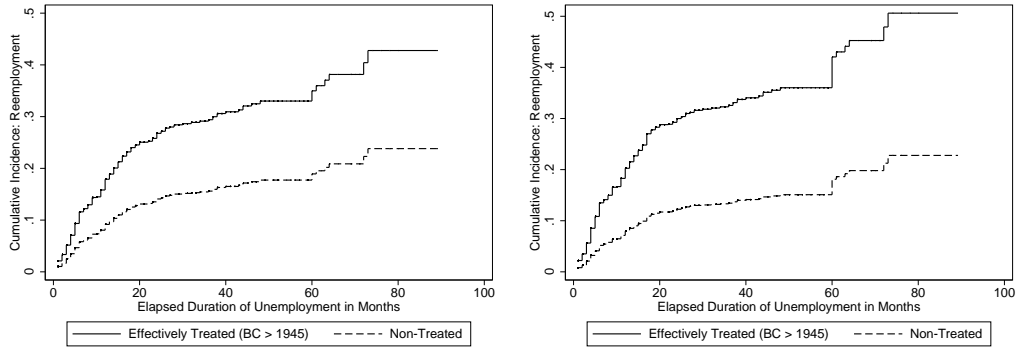
Table 3: Reemployment Probability: DiD considering “Other Exits” as Competing Risk.

	Continuous Treatment		Binary Treatment	
	CR	CR Balanced	CR	CR Balanced
Treat X AddMonths	0.031 (0.012)	0.035 (0.014)	0.719 (0.361)	1.029 (0.440)
Treat	0.538 (0.220)	0.641 (0.282)	0.576 (0.229)	0.629 (0.292)
AddMonths	-0.024 (0.012)	-0.029 (0.014)	-0.673 (0.355)	-0.991 (0.437)
Reduction-Free Ret.-Age	0.227 (0.057)	0.138 (0.070)	0.237 (0.059)	0.157 (0.073)
Age 55	-0.502 (0.763)	-0.474 (0.860)	-0.545 (0.794)	-0.516 (0.884)
Age 56	-0.801 (0.762)	-0.732 (0.856)	-0.845 (0.793)	-0.768 (0.881)
Age 57	-1.201 (0.760)	-1.234 (0.866)	-1.260 (0.791)	-1.297 (0.890)
Age 58	-1.563 (0.763)	-1.521 (0.867)	-1.601 (0.794)	-1.538 (0.890)
Age 59	-1.546 (0.771)	-1.380 (0.881)	-1.603 (0.802)	-1.432 (0.907)
Age 60	-2.024 (0.884)	-2.257 (0.972)	-2.051 (0.911)	-2.288 (0.995)
Age 61	-1.576 (0.976)	-1.584 (1.056)	-1.545 (0.996)	-1.554 (1.070)
West Germany	-1.115 (0.306)	-0.992 (0.403)	-1.136 (0.308)	-1.006 (0.403)
Migrant	-0.239 (0.263)	-0.205 (0.302)	-0.268 (0.264)	-0.241 (0.302)
Handicap Level	-0.024 (0.006)	-0.023 (0.007)	-0.023 (0.006)	-0.022 (0.006)
Married	0.140 (0.173)	0.098 (0.192)	0.126 (0.175)	0.087 (0.194)
Home Owner	0.015 (0.133)	-0.015 (0.169)	0.017 (0.133)	-0.020 (0.170)
Dep. Children in HH	-0.434 (0.287)	-0.438 (0.285)	-0.444 (0.286)	-0.448 (0.285)
Years of Educ.	0.001 (0.028)	0.016 (0.041)	0.004 (0.027)	0.016 (0.041)
Regional UE Rate	-0.025 (0.031)	-0.032 (0.036)	-0.030 (0.032)	-0.036 (0.036)
Particip. in Pre-Ret.	-1.583 (0.608)	-1.738 (0.607)	-1.584 (0.605)	-1.758 (0.608)
Unemployed Partner	-0.311 (0.221)	-0.163 (0.253)	-0.317 (0.220)	-0.170 (0.252)
Retired Partner	-0.170 (0.203)	-0.157 (0.242)	-0.149 (0.203)	-0.147 (0.245)
Past UE Experience	-0.013 (0.039)	0.036 (0.042)	-0.007 (0.038)	0.039 (0.042)
Income Available	0.056 (0.173)	0.227 (0.198)	0.056 (0.171)	0.223 (0.196)
N	971	971	971	971

Note: Own calculations based on the SOEP (1991-2012). Robust standard errors in parentheses. All exits to other states than employment are considered as the competing risk (e.g. disability pension). Age in years is modeled using binary indicators. Abbreviations: CR = Competing Risk; UE = Unemployment. Both models estimated in two versions with and without the balanced control group respectively.

ment, CR Balanced) and 1.9% (Binary Treatment, CR Balanced) and thus the estimated treatment effects are statistically significant assuming conventional error probabilities.

Figure 2: Cumulative Incidence Functions: Failure Event Reemployment.



(a) Unbalanced Sample

(b) Balanced Sample

*Note:* Own calculations using the SOEP (1991-2012). Cumulative incidence functions as predicted from competing risks regression. Remaining covariates are held constant at their mean values. Effectively treated individuals belong to the treatment group and are born after 1945 while non-treated individuals are the remaining observations from both control- and treatment group.

Figure 2 plots cumulative incidence functions as predicted from the competing risks model for both the unbalanced and balanced sample (table 3, column 3 and 4, binary treatment). We separate the estimation sample in two groups to compare effectively treated individuals to non-treated individuals, using the interaction term from the differences-in-differences estimation.<sup>13</sup> In this sense, the difference between the cumulative incidence functions between groups in figure 2 can be understood as a graphical illustration of the treatment effect. Fixing all remaining covariates at their mean values, the model predicts that no more than 12% of the non-treated unemployed individuals exhibit a reemployment after 24 months while this is the case for about 30% among the effectively treated individuals. After the maximum observed unemployment duration of 96 months (i.e. 8 years), the model predicts that only about 20% of non-treated individuals are reemployed while about 50% of effectively treated older workers exhibit the event of reemployment. This difference is large in absolute (30 percentage points) and relative terms (150%).

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hazard ratios are  $\exp(0.035) = 1.0356$  (Continuous Treatment, CR Balanced),  $\exp(1.029) = 2.798$  (Binary Treatment, CR Balanced).

<sup>13</sup>Effectively treated individuals thus belong to the treatment group and are born after 1945, while non-treated individuals are all remaining observation of the estimation sample, i.e. the full control group and those individuals of the treatment group who were born before 1946.

The estimated coefficient for “Treat” is positive and large in magnitude, indicating that reemployment in the treatment group is larger compared to the control group. The large difference between the two groups is explained by the fact that the treatment indicator (variable “Treat”) almost perfectly determines the subgroups of the two sexes; this result is an artefact of the separation between control- and treatment group. Finally, the estimated coefficients for the variable “AddMonths” are significantly negative, which indicates a negative time trend in reemployment.

To examine the robustness of the main result, we alternate distributional assumptions including a semi-parametric Cox model (SP), a fully parametric Weibull distribution (FP) and a frailty distribution (FR) for evolving unobserved heterogeneity (see table 5 and 6 in the appendix). The interpretation of results differs from the previous ones due to the absence of a competing event. The principal result of an increasing reemployment hazard is confirmed, however differing by the distributional assumptions of the respective model. Raising the ERA by one month increases the reemployment hazard by 3.3% (SP Balanced), 2.9% (FP Balanced) and 4.7% (FR Balanced), holding everything else constant.<sup>14</sup> The probability, that these results occur by chance vary between 1.8% (SP Balanced), 4.5% (FP Balanced) and 2.7% (FR Balanced).

Taking into account that the perspective here is an “as-if” binary treatment, the estimated average treatment effect is large (table 6). The binary treatment effect of raising the ERA by three years increases the reemployment probability by 163% (SP Balanced), 136% (FP Balanced) and 288% (FR Balanced), holding everything else constant.<sup>15</sup> The probability, that these results occur only by chance are 3.0% (SP Balanced), 6.8% (FP Balanced) and 4.4% (FR Balanced).

The existing literature supports the finding of a positive effect of raising the ERA on reemployment. Staubli and Zweimüller (2013) find that raising the ERA does increase employment in the relevant age group by restricting retirement to higher ages. Moreover, magnitude and direction of our result are supported by the finding that the distance to retirement has a significant positive impact on employment (Hairault et al., 2010).

<sup>14</sup>The hazard ratios are  $\exp(0.033) = 1.0336$  (SP Balanced),  $\exp(0.029) = 1.0294$  (FP Balanced) and  $\exp(0.046) = 1.0471$  (FR Balanced).

<sup>15</sup>The hazard ratios are  $\exp(0.966) = 2.6274$  (SP Balanced),  $\exp(0.859) = 2.3608$  (FP Balanced) and  $\exp(1.357) = 3.8845$  (FR Balanced).

## 5 Conclusion

This paper analyses the impact of an increased early retirement age on the probability of reemployment and early retirement behaviour for older unemployed individuals in Germany. In the course of this reform, the ERA for an old age pension due to unemployment is increased by a maximum of three years from age 60 to 63 in monthly steps.

After implementing a matching procedure based on entropy balancing, we estimate differences-in-differences and account for the intensity of the individual treatment, i.e. by how many months the distance to retirement is raised in correspondence to the reform. Using data from the German Socio-Economic Panel allows to control for a rich set of socio-demographic factors that are part of a complex relationship between unemployment and reemployment for older unemployed workers.

From the perspective of an older unemployed individual raising the ERA may change the expected lifetime utility for the alternatives unemployment versus employment. Raising the ERA extends the distance to retirement in total by three years. Holding everything else constant, we investigate whether this exogenous extension of the unemployment horizon stimulates reemployment. Our reasoning is that raising the ERA may incentivise older unemployed individuals to return to work instead of using unemployment as a bridge until an old age pension is available.

Estimating the average treatment effect indicates that raising the ERA for an old age pension due to unemployment has a significant positive effect on reemployment. Holding everything else constant, the reemployment probability increases by 3.6% for each month by which the ERA is raised. Assuming a binary treatment, the effect amounts to an increase in the reemployment hazard of 180%. The effect is large in absolute and relative terms because a critical share of unemployed older workers exhibits the event of reemployment because no old age pension is available during age 60 to 63. The result implies that unemployment as a pathway into early retirement is less prevalent and thus

raising the distance to retirement substantially reduces public costs from non-employment.

In line with previous findings (Lalive, 2008; Hairault et al., 2010) our results suggest that the distance to (early) retirement does matter. Recent population ageing induces an increasing demand for labour force participation at higher ages. This study shows that restricting old age pensions to higher ages for older unemployed workers does stimulate reemployment. We conclude that such a policy substantially reduces public costs from non-employment.

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## Appendix

Table 4: Reform Steps of Raising the ERA (Excerpt from Social Security Code).

Year of Birth	NRA		ERA	
	Years	Months	Years	Months
<b>Before 1946</b>	65	0	60	0
<b>1946</b>				
January	65	0	60	1
February	65	0	60	2
March	65	0	60	3
April	65	0	60	4
May	65	0	60	5
June	65	0	60	6
July	65	0	60	7
August	65	0	60	8
September	65	0	60	9
October	65	0	60	10
November	65	0	60	11
December	65	0	61	0
<b>1947</b>				
January	65	0	61	1
February	65	0	61	2
March	65	0	61	3
April	65	0	61	4
May	65	0	61	5
June	65	0	61	6
July	65	0	61	7
August	65	0	61	8
September	65	0	61	9
October	65	0	61	10
November	65	0	61	11
December	65	0	62	0
<b>1948</b>				
January	65	0	62	1
February	65	0	62	2
March	65	0	62	3
April	65	0	62	4
May	65	0	62	5
June	65	0	62	6
July	65	0	62	7
August	65	0	62	8
September	65	0	62	9
October	65	0	62	10
November	65	0	62	11
December	65	0	63	0
<b>1949 - 1951</b>	65	0	63	0

Source: German Social Security Code (Anlage 19, SGB VI).

Note: NRA abbreviates normal retirement age; ERA abbreviates early retirement age. The figures in column 4 and 5 display the reform steps of raising the ERA for an old age pension due to unemployment.

Table 5: Reemployment Probability: Continuous Treatment DiD

	SP	SP Balanced	FP	FP Balanced	FR	FR Balanced
Treat X AddMonths	0.028 (0.012)	0.033 (0.014)	0.026 (0.012)	0.029 (0.015)	0.041 (0.018)	0.046 (0.021)
Treat	0.563 (0.224)	0.673 (0.283)	0.636 (0.222)	0.770 (0.298)	0.979 (0.377)	1.197 (0.484)
AddMonths	-0.024 (0.011)	-0.029 (0.014)	-0.023 (0.011)	-0.027 (0.014)	-0.034 (0.017)	-0.040 (0.020)
Reduction-Free Ret.-Age	0.212 (0.059)	0.124 (0.071)	0.209 (0.058)	0.120 (0.073)	0.329 (0.094)	0.189 (0.114)
Age 55	-0.555 (1.043)	-0.534 (0.871)	-0.593 (1.043)	-0.512 (0.860)	-0.383 (1.874)	-0.383 (1.577)
Age 56	-0.849 (1.042)	-0.769 (0.867)	-0.897 (1.041)	-0.784 (0.857)	-0.978 (1.878)	-0.883 (1.581)
Age 57	-1.212 (1.042)	-1.250 (0.876)	-1.216 (1.041)	-1.213 (0.864)	-1.528 (1.888)	-1.595 (1.591)
Age 58	-1.582 (1.042)	-1.532 (0.878)	-1.524 (1.040)	-1.420 (0.868)	-2.016 (1.909)	-1.970 (1.612)
Age 59	-1.498 (1.048)	-1.330 (0.889)	-1.347 (1.045)	-1.118 (0.881)	-1.642 (1.897)	-1.399 (1.599)
Age 60	-1.889 (1.132)	-2.142 (0.980)	-1.742 (1.129)	-1.918 (0.968)	-2.241 (1.995)	-2.606 (1.692)
Age 61	-1.591 (1.172)	-1.592 (1.065)	-1.411 (1.169)	-1.341 (1.049)	-1.500 (2.025)	-1.619 (1.820)
West Germany	-1.114 (0.308)	-0.983 (0.408)	-1.130 (0.310)	-0.955 (0.425)	-1.508 (0.502)	-1.244 (0.638)
Migrant	-0.252 (0.273)	-0.220 (0.301)	-0.276 (0.271)	-0.256 (0.304)	-0.466 (0.415)	-0.539 (0.458)
Handicap Level	-0.023 (0.006)	-0.021 (0.007)	-0.023 (0.006)	-0.021 (0.007)	-0.036 (0.010)	-0.035 (0.010)
Married	0.135 (0.174)	0.090 (0.197)	0.130 (0.174)	0.071 (0.201)	0.091 (0.277)	0.006 (0.321)
Home Owner	0.025 (0.136)	-0.007 (0.172)	0.034 (0.136)	-0.003 (0.181)	0.047 (0.218)	-0.064 (0.286)
Dep. Children in HH	-0.426 (0.262)	-0.445 (0.282)	-0.466 (0.262)	-0.469 (0.304)	-0.666 (0.430)	-0.737 (0.458)
Years of Educ.	0.001 (0.028)	0.016 (0.042)	-0.002 (0.028)	0.012 (0.043)	0.008 (0.046)	0.047 (0.072)
Regional UE Rate	-0.028 (0.031)	-0.035 (0.037)	-0.026 (0.031)	-0.029 (0.038)	-0.022 (0.048)	-0.032 (0.057)
Particip. in Pre-Ret.	-1.535 (0.584)	-1.703 (0.606)	-1.589 (0.584)	-1.759 (0.614)	-2.147 (0.800)	-2.454 (0.811)
Unemployed Partner	-0.306 (0.210)	-0.159 (0.259)	-0.324 (0.210)	-0.165 (0.271)	-0.396 (0.339)	-0.093 (0.421)
Retired Partner	-0.134 (0.211)	-0.130 (0.247)	-0.110 (0.211)	-0.133 (0.273)	-0.180 (0.330)	-0.161 (0.425)
Past UE Experience	-0.014 (0.039)	0.033 (0.042)	-0.014 (0.039)	0.033 (0.044)	-0.006 (0.060)	0.064 (0.070)
Income Available	0.104 (0.188)	0.262 (0.203)	0.097 (0.188)	0.227 (0.215)	0.280 (0.307)	0.548 (0.338)
Constant			-15.597 (3.800)	-10.418 (4.508)	-23.831 (6.119)	-15.632 (7.076)
Weibull Parameter $\alpha$			0.831 (0.056)	0.829 (0.046)	1.282 (0.138)	1.323 (0.078)
Frailty Parameter $\theta$					2.272 (0.980)	2.588 (0.568)
N	971	971	971	971	971	971

Note: Own calculations based on the SOEP (1991-2012). Robust standard errors in parentheses.  $\alpha$  is the additional parameter of the Weibull distribution that allows for duration dependence, where  $\alpha = 1$  implies the exponential distribution.  $\theta$  is the heterogeneity parameter of the inverse Gaussian distribution, where  $\theta = 0$  implies that no unobserved heterogeneity is present. Age in years is modeled via respective dummies. Abbreviations: SP = Semi-Parametric; FP = Fully Parametric; FR = Frailty; UE = Unemployment. All models estimated in two versions with and without the balanced control group respectively.

Table 6: Reemployment Probability: Binary Treatment DiD

	SP	SP Balanced	FP	FP Balanced	FR	FR Balanced
Treat X AddMonths	0.653 (0.362)	0.966 (0.444)	0.570 (0.361)	0.859 (0.470)	0.922 (0.546)	1.357 (0.672)
Treat	0.605 (0.234)	0.664 (0.294)	0.680 (0.233)	0.764 (0.313)	1.063 (0.401)	1.188 (0.511)
AddMonths	-0.649 (0.346)	-0.978 (0.442)	-0.622 (0.345)	-0.914 (0.464)	-0.909 (0.516)	-1.322 (0.658)
Reduction-Free Ret.-Age	0.221 (0.061)	0.141 (0.074)	0.217 (0.060)	0.137 (0.077)	0.340 (0.099)	0.212 (0.119)
Age 55	-0.602 (1.044)	-0.580 (0.887)	-0.643 (1.044)	-0.563 (0.870)	-0.471 (1.888)	-0.446 (1.599)
Age 56	-0.893 (1.043)	-0.807 (0.883)	-0.944 (1.042)	-0.824 (0.867)	-1.067 (1.893)	-0.939 (1.602)
Age 57	-1.268 (1.043)	-1.311 (0.892)	-1.275 (1.042)	-1.277 (0.875)	-1.631 (1.905)	-1.680 (1.614)
Age 58	-1.621 (1.043)	-1.555 (0.892)	-1.567 (1.041)	-1.449 (0.877)	-2.111 (1.926)	-2.005 (1.635)
Age 59	-1.553 (1.049)	-1.383 (0.906)	-1.405 (1.046)	-1.174 (0.893)	-1.753 (1.913)	-1.478 (1.623)
Age 60	-1.920 (1.132)	-2.180 (0.996)	-1.779 (1.130)	-1.963 (0.979)	-2.315 (2.013)	-2.658 (1.718)
Age 61	-1.574 (1.172)	-1.579 (1.071)	-1.410 (1.169)	-1.343 (1.052)	-1.511 (2.039)	-1.608 (1.833)
West Germany	-1.126 (0.308)	-0.985 (0.406)	-1.137 (0.310)	-0.951 (0.422)	-1.536 (0.506)	-1.257 (0.631)
Migrant	-0.280 (0.273)	-0.255 (0.302)	-0.302 (0.272)	-0.290 (0.306)	-0.505 (0.420)	-0.587 (0.461)
Handicap Level	-0.022 (0.006)	-0.021 (0.006)	-0.023 (0.006)	-0.021 (0.007)	-0.035 (0.010)	-0.034 (0.010)
Married	0.117 (0.177)	0.075 (0.199)	0.107 (0.177)	0.053 (0.205)	0.069 (0.282)	-0.004 (0.324)
Home Owner	0.024 (0.136)	-0.012 (0.172)	0.031 (0.136)	-0.010 (0.182)	0.052 (0.219)	-0.064 (0.287)
Dep. Children in HH	-0.431 (0.261)	-0.448 (0.283)	-0.466 (0.262)	-0.468 (0.303)	-0.681 (0.433)	-0.751 (0.458)
Years of Educ.	0.003 (0.028)	0.016 (0.042)	0.001 (0.028)	0.012 (0.043)	0.010 (0.046)	0.046 (0.072)
Regional UE Rate	-0.032 (0.031)	-0.037 (0.037)	-0.030 (0.031)	-0.031 (0.038)	-0.027 (0.048)	-0.036 (0.056)
Particip. in Pre-Ret.	-1.543 (0.584)	-1.727 (0.608)	-1.602 (0.584)	-1.785 (0.617)	-2.177 (0.805)	-2.487 (0.815)
Unemployed Partner	-0.311 (0.209)	-0.166 (0.258)	-0.328 (0.209)	-0.171 (0.270)	-0.404 (0.341)	-0.105 (0.423)
Retired Partner	-0.111 (0.210)	-0.116 (0.250)	-0.087 (0.210)	-0.121 (0.277)	-0.131 (0.330)	-0.122 (0.431)
Past UE Experience	-0.010 (0.039)	0.034 (0.042)	-0.010 (0.038)	0.034 (0.043)	-0.002 (0.060)	0.065 (0.069)
Income Available	0.104 (0.187)	0.258 (0.201)	0.094 (0.187)	0.222 (0.214)	0.277 (0.308)	0.546 (0.338)
Constant			-16.064 (3.909)	-11.381 (4.701)	-24.347 (6.344)	-16.881 (7.334)
Weibull Parameter $\alpha$			0.834 (0.056)	0.831 (0.046)	1.294 (0.140)	1.325 (0.083)
Frailty Parameter $\theta$					2.331 (1.006)	2.596 (0.624)
N	971	971	971	971	971	971

Note: Own calculations based on the SOEP (1991-2012). Robust standard errors in parentheses.  $\alpha$  is the additional parameter of the Weibull distribution that allows for duration dependence, where  $\alpha = 1$  implies the exponential distribution.  $\theta$  is the heterogeneity parameter of the inverse Gaussian distribution, where  $\theta = 0$  implies that no unobserved heterogeneity is present. Age in years is modeled via respective dummies. Abbreviations: SP = Semi-Parametric; FP = Fully Parametric; FR = Frailty; UE = Unemployment. All models estimated in two versions with and without the balanced control group respectively.