Retirement Decisions: The Role of Labour Supply, Financial Incentives, and Mortality Expectations

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Name:	Matthias Nicolas Giesecke
Geburtsort:	Berlin

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Erstgutachter:	Prof. Dr. Reinhold Schnabel
Zweitgutachter:	Prof. Dr. Thomas K. Bauer

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Preface

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Chapter 1

Introduction

1.1 The Problem

Demographic change jeopardises pay-as-you-go pension systems (PAYG) (Hairault et al., 2010). The essence of a PAYG is that contributions are more or less directly transmitted to benefit recipients at a given point in time. Consequently, any event that imposes shifts on the side of contributors, benefit recipients or both may unbalance the system. The reasons for such temporary shifts are manifold but are currently driven by decreasing birth rates, changes in labour force participation of older workers and increasing life expectancy.[1]

The particularly large "baby boom" cohorts born between 1955 and 1970 move towards retirement as of the year 2015, once they successively become eligible for old age pensions; at the same time, younger birth cohorts are much smaller (Federal Statistical Office, 2014). Moreover, life expectancy has increased remarkably in recent years, which is evident from life tables as provided by Federal Statistical Office (2012b). These changes do have an impact on the ratio of contributors and benefit recipients, if there is no response in terms of a postponed retirement age. The consequences of population ageing for the PAYG are easily summarised: Allowing for longer lives and decreasing birth rates while holding everything else constant either raises the contribution rate or alleviates benefit entitlements or both.

[1] Importantly, changes in the labour force participation of other groups such as women indeed strongly influence the PAYG, but those patterns are beyond the scope of this thesis.

Irrespective of the normative question of how to allocate this burden, reduced benefits may come along with old age poverty and increased contribution rates may set disincentives to supply labour at the margin. The most notable way to confront this challenge is an increasing effective retirement age, i.e. more years of contribution and relatively fewer years of benefit receipt.[2]

The traditional explanation for retirement decisions is a combination of individual preferences for consumption and leisure and incentives that are set by the social security system (see e.g. Weiss, 1972; Sheshinski, 1978, for early contributions). Clearly, the consequence of early retirement is more leisure (less work) accompanied by less consumption (lower income from labour and/or social security) and thus retirement is an issue of labour supply (Hurd, 1990). The economic theory is very clear in explaining retirement decisions; studies such as Gordon and Blinder (1980), Crawford and Lilien (1981) and Gustman and Steinmeier (1986) all have in common that they establish models where individuals maximise utility over the life cycle and chose to retire exactly when benefits and costs for this decision are balanced.

The way in which these models differ is the challenge to define a precise mechanism for the interdependence between life cycle preferences and external incentives from social security. Retirement incentives result from a mix of rules that determine benefit entitlements and this is mainly driven by contributions, the statutory retirement age and financial incentives from actuarial adjustments.[3] A convenient way to summarise these aspects are expected present discounted values, which provide a measure of the total sum of all future pension claims, discounted to a given point in time while taking into account uncertain lifetimes (see e.g. Fields and Mitchell, 1984a; Mitchell and Fields, 1984). In the option value approach as proposed by Stock and Wise (1990), an individual irreversibly retires if there is no expected gain from future work as mea-

[2]Self-evidently, this view abstracts from changes in productivity and labour force participation. Also, this view does not exclude the possibility of a growing average of total years of benefit receipt.

[3]Financial incentives in terms of actuarial adjustments refer to reductions or increments of retirement benefits in order to adjust the pension system towards more actuarial neutrality. In Germany, the current reduction rate is 3.6% for each year of retirement previous to the normal retirement age and the increment is 6% for each year of retirement after the normal retirement age.

sured in utility, or chooses to continue work and retire in some later period otherwise. The key feature in this model is the possibility to reevaluate the retirement decision in each period, as long as the transition into retirement has not been materialised (i.e. the “option value of work”). All these approaches have in common that they take a forward-looking perspective on future streams of income from work and retirement benefits. Beyond individual preferences for leisure and consumption, retirement may coincide with subsequent phenomena that either support or prevent an individual to live longer. Retirement may relieve individuals from work-related stress with a positive impact on remaining years to live but empirical evidence suggests that cognitive decline sets in after retirement (see e.g. Rohwedder and Willis, 2010; Bonsang et al., 2012).

Explaining retirement decisions is an intricate task. It is the ultimate aim of the present dissertation to examine the role of factors that are likely to be influential for retirement decisions in terms of the retirement age. Specifically, this thesis is concerned with labour supply (i.e. reemployment prospects) of older workers. Broadly speaking, this is a question of labour force participation of older workers which is directly linked to the timing of retirement. Another issue to be raised are financial incentives and their impact on the timing of retirement where responses are examined for workers that differ by physical demands of occupations. Finally, the relationship between differential mortality expectations and the retirement age are explored. All studies of this thesis take a perspective that combines individual preferences to incentives from the social security system. For this purpose, the German social security system serves as reference point.

The German literature on the topic is multifaceted and has brought up a large body of research. The relationship between social security and labour force participation and corresponding (dis-)incentives from the pension system have been described in numerous studies (see e.g. Börsch-Supan and Schnabel, 1998, 1999; Börsch-Supan, 2000a,b; Börsch-Supan and Berkel, 2004; Börsch-Supan et al., 2004; Börsch-Supan, 2005; Hanel, 2010, 2012). The articles of this dissertation carry on and extend this literature in several aspects.

The need for further research in this field is manifold. First, current developments in terms of demography, as outlined above, may impose a fair amount of challenge on the PAYG in the near future. The present work attempts to capture developments that may have important policy implications and may contribute to find solutions for problems of the social securtiy system.

Second, new administrative data sources of exceptional quality are available from the German federal pension insurance (Himmelreicher and Stegmann, 2008). Chapter 3 of this dissertation contributes to a strand of the literature that examines the relationship of retirement decisions and financial incentives (Börsch-Supan and Schnabel, 1999; Hanel, 2010, 2012). This study is the first one to cover the full implementation period of benefit reductions between January 1997 and December 2004. Furthermore, research on differential mortality and the retirement age in chapter 5 is based on a unique data source from social security records (Scholz, 2005). The data provide precise information on entry into retirement and the shortfall of pension benefit recipients and thus allow to determine the duration of benefit receipt. The present work adds to the literature that examines the relationship between retirement age to mortality (see e.g. Kühntopf and Tivig, 2012; Hernaes et al., 2013).

Third, the pension system has evolved over time as several modifications to the social security legislation in Germany have been implemented recently. All articles in this thesis either examine the impact of such policy modifications directly or refer to aspects of retirement behaviour that are closely related such changes. This research is important in terms of programme evaluation and informative about current policy debates. One example is chapter (2) which is concerned about the question as to whether raising the early retirement age has a causal effect on reemployment. Moreover, the effect of benefit reductions on the retirement age is examined (chapter 3). The remaining studies refer to policy changes indirectly and discuss retirement behaviour as motivated from social security reforms. Examples are the increase of the normal retirement age from 65 to 67 between 2012 and 2029 (chapter 4) and the reduction of the retirement age for long-term-insured individuals from 65 to 63 without benefit reductions in 2014 (chapter 5).

Finally, this dissertation adds to the literature in terms of methodology. The estimation of a treatment effect in chapter (2) is based on a combination of selection on observables and selection on unobservables. The idea is to improve the potentially poor properties of a non-experimental estimator (Heckman et al., 1997). The study makes use of a very recent matching procedure (Hainmueller, 2012; Hainmueller and Xu, 2013) in the first step and then estimates differences-in-differences. This is an example of how this dissertation combines well-established empirical methods to state of the art techniques in order to improve the quality of results obtained.

1.2 Some Descriptives

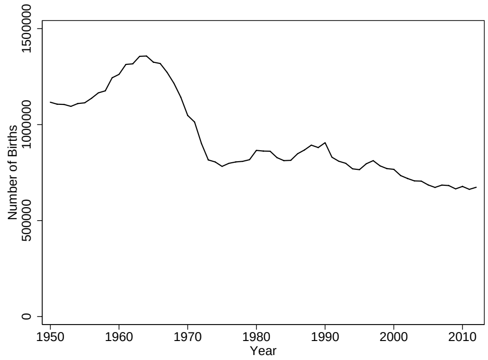
In 2012, the total revenue of the German federal pension insurance (Deutsche Rentenversicherung Bund, DRV-Bund) was about 260 billion Euros of which roughly 194 billion Euros were due to contributions (German Federal Pension Insurance, 2013). At the same time, the total expenditure amounted to 255 billion Euros of which about 229 billion Euros were paid to different types of pensions. In relation to the federal government budget, this is the second largest public budget in Germany and is a manifestation of the monetary dimension as managed by DRV-Bund. Clearly, very small changes in the ratio of contributors to benefit recipients potentially imply large deficits or surpluses in the accounting balance.

Several factors may induce such changes in the ratio of contributors to benefit recipients and thus challenge the financial stability of a PAYG. In a broader sense, these are demographic aspects which are subject to constant changes, and therefore it is problematic to speak of a steady state or any kind of long-run stability condition. Instead, a PAYG is naturally subject to permanent adjustments and in the following I illustrate the German case for some recent phenomena that jeopardise the system.

First, birth rates have been decreasing in Germany over decades. Figure 1.1 indicates an overall decrease in the absolute number of births between 1950 (about 1,100,000) and 2012 (about 670,000), which has two important implications for the PAYG: On the one hand, fewer individuals enter the labour market and contribute to the PAYG.

At the same time, figure 1.1 shows that relatively large "baby boom" cohorts move towards retirement. Precisely, the cohorts born between 1955 and 1970 become eligible for different types of old age pensions between 2015 and 2035, which means that an exceptionally large proportion of contributors will turn into benefit recipients within the next 20 years.

Figure 1.1: Absolute Number of Births in Germany: 1950 - 2012.



Source: Federal Statistical Office (2014).

Second, recent data indicate that individuals live longer and that life expectancy increases impressively. Over the ten-year period from 2002 to 2011, the expected remaining years to live for male (female) individuals, conditional on reaching age 60, has increased from 19,68 (23,84) years to 21,31 (24,96) years, which is a difference of 1,63 (1,12) years (Federal Statistical Office, 2012b). This means that retirement benefit receipt largely extends to higher ages on average. Holding everything - and specifically the retirement age - constant, benefits are received for a longer period.

Third, taking into account recent demographic trends as outlined above, the critical question is how the retirement age has evolved in the recent past. Figure 4.1 displays the development of the average retirement age for different types of pensions in West Germany ("Alte Bundesländer") separately for the two sexes.
There is a remarkable drop in the average retirement age, following the change in social security legislation from 1972. The trend towards early retirement after 1972 has been discussed elsewhere and this literature is very clear in revealing major disincentives with respect to the choice of retirement ages between the 1970s and 1990s (see e.g. Börsch-Supan and Schnabel, 1998; Börsch-Supan, 2000a).

Figure 1.2: Average Retirement Age in West Germany: 1960 - 2012.

(a) Male (b) Female

Source: German Federal Pension Insurance (2013). *Note:* Retirement ages are restricted to West Germany ("Alte Bundesländer") due to data limitations. OA-Pensions abbreviate old age pensions; REC-Pensions abbreviate reduced earnings capacity pensions.

Consequently, a major public debate in the late 1980s and early 1990s has initiated extensive reforms. As an example, the introduction of actuarial adjustments for different types of old age pensions between 1997 and 2004 has imposed financial incentives on the timing of retirement. Indeed, we observe an increase of the average retirement age in old age pensions starting in the late 1990s for both men and women. Whether this increase in the average retirement age can be assigned to a response to financial incentives is an empirical question to be analysed in chapter 3.

Another pattern that appears remarkable is the strong decline in retirement age for reduced earnings capacity pensions. While the observed pattern is certainly subject to compositional effects as accompanied by a decline in absolute numbers (see German Federal Pension Insurance, 2013), it raises questions about the retirement age of individuals with specific health indications as a prerequisite for eligibility. In the light of recent reforms, where actuarial adjustments have been introduced and the normal retirement age is raised, it is of some interest whether these rules impose benefit reductions on a group that may have difficulties to respond. Similarly, chapter 4 contributes to this debate by an extension of the framework of "equality of opportunity" to retirement age, with a specific focus on physical demands of occupations.

The future development of the average retirement age is uncertain. Whether the recent trend of an increasing average retirement age for old age pensions as apparent from figure 4.1 will continue is strongly driven by retirement policy. Recent reform plans to increase the attractiveness of early retirement for longterm-insured individuals, for example, are likely to hamper the recent upward trend. As outlined above, this thesis contributes to the discussion about the design of retirement policies that are closely related to the development of the retirement age. A short overview and summary of studies is outlined below.

1.3 Overview and Summary of Studies

The first study **"Does Raising the Early Retirement Age Increase Reemployment of Older Unemployed Workers?"** (chapter 2) examines an upward shift 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.

The second study **"The Effect of Benefit Reductions on the Retirement Age: The Heterogeneous Response of Manual and Non-Manual Workers"** (chapter 3) estimates the effect of benefit reductions on the timing of retirement. The introduction of actuarial adjustments in the German public pension system serves as a source of exogenous variation to estimate discrete time transition rates into retirement for individuals of age 60 - 66. The heterogeneous response to benefit reductions is elaborated separately for manual and non-manual workers. On average, individuals postpone retirement by five months if pension benefits are reduced by 3.6 per cent for each year of early retirement. This confirms the common finding that people respond to incentives if the implicit tax on further periods of work is reduced. However, the response is about 60 per cent lower for manual workers compared to non-manual workers. Surprisingly, this does not necessarily indicate that retirement incomes of manual workers deteriorate. The explanation is that large scale retirement into disability pensions takes place at age 63 - without benefit reductions.

The third study **"Retirement Age and Inequality of Opportunity: The Role of Physical Job Demands"** (chapter 4) quantifies differences in the retirement age between manual and non-manual workers. The difference is evaluated in the context of the literature on equality of opportunity. The focus is on the question how individual background during childhood transmits through physical demands of occupations on retirement ages. Individual retrospective data from the German Socio-Economic Panel are used to analyse labour force dynamics over the years 1984 to 2011. Discrete time duration models suggest that retirement ages differ substantially between manual and non-manual workers. To elaborate how such differences are explained by individual background characteristics on the one hand and effort and luck on the other hand we make use of tests for stochastic dominance and a Blinder-Oaxaca decomposition. The main result is that individual background characteristics explain a share of about one third of inequality in retirement ages as transmitted through physical demands of occupations.

Finally, the fourth study **"Redistribution in a Pay-As-You-Go Pension System: The Relationship between Differential Mortality and the Retirement Age"** (chapter 5) reveals the relationship between the age of retirement and differential mortality for the birth cohorts 1912 - 1943 in Germany. The analysis is based on a unique data source from social security records which contain precise information retirement entries and the shortfall of pension benefit recipients. Central interest is on the average remaining life expectancy as a function of retirement age, which is expected to be monotonically decreasing. Our results indicate, however, that this is not the case among male workers; especially towards age 63 the slope is positive for men and thus they receive old age pensions for more periods on average compared to those who retire at lower ages. This pattern is due to subgroup heterogeneity for individuals who select themselves into specific retirement ages. That is, specific types of workers retire at certain ages and these groups differ in remaining life expectancy by relevant aspects such as health-related behaviours and wealth. This finding has important implications for redistribution in a pay-as-you-go pension system: Abolishing benefit reductions for early retirees (age 63) who are particularly long-living seems not to be a good idea.

Chapter 2

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

2.1 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.

Until 2006, the minimum age for early retirement due to unemployment was 60 years in Germany. Then, the social security legislation was modified such that the ERA for unemployed individuals was increased to age 63 in monthly steps for the birth cohorts 1946-1948; the reform fully phased in between January 2006 and December 2011. As of January 2012, no old age pension due to unemployment is available at ages below 63.

A recent body of literature has investigated the relationship between the timing of retirement and late-career employment patterns (see e.g. Chan and Stevens, 2001; Tatsiramos, 2010; Hairault et al., 2010; Staubli and Zweimüller, 2013). This paper carries on research of Staubli and Zweimüller (2013), who investigate employment patterns of older workers after the ERA has been raised in Austria. In contrast to their study, we explicitly take the perspective of older *unemployed* individuals to shed light on the effect of raising the ERA on reemployment. We consider only those individuals with a job loss such that they face a situation of unemployment by definition. Once individuals enter 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 policy change of raising the ERA restricts the availability of old age pensions due to unemployment to higher ages. In this sense it may change the trade-off between relevant alternatives after a late-career job loss. The previous literature provides arguments for the existence of a positive impact of raising the ERA on employment. The distance to retirement does matter for employment decisions. Hairault et al. (2010) find that the likelihood of employment of older workers is significantly affected by the distance to retirement. In the context of this paper, incentives for reemployment may arise if the distance to retirement is larger because early retirement is only available after three additional years. Furthermore, raising the ERA prolongs labour market careers. Staubli and Zweimüller (2013) show that one consequence of raising the ERA in Austria is an increase in employment for the relevant age group.

However, the group of older individuals has specific characteristics with respect to job prospects and labour force participation which may challenge the positive impact of raising the ERA on reemployment. First, empirical evidence suggests a negative relationship between age and reemployment (see e.g. Chan and Stevens, 2001). Second, returning to a job may be an increasingly burdensome event as individuals grow older (Hernaes et al., 2013). And third, 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). It is ultimately an empirical question whether raising the ERA affects the reemployment probability and we aim to identify size and direction of the potential impact which is not clear a priori. Raising the ERA is a source of exogenous variation and our empirical strategy allows to identify a causal effect on reemployment following a late-career job loss. The methodology of this paper is indebted to a range of identification issues with respect to observable and unobservable heterogeneity. Therefore, 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 (see e.g. Marcus, 2013, for a recent application). 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. We use data from the German Socio-Economic Panel (SOEP) for the years 1991 - 2012 which contain rich information on recently unemployed individuals. Unemployment spells are subsequently followed to analyse whether transitions into employment occur.

Our results suggest that raising the ERA does have a positive causal impact on the probability of reemployment for 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.

The remainder of this paper is structured as follows. Section 2.2 provides an overview on recent employment patterns and the institutional setting. Section 2.3 describes data, identification, and empirical strategy. Section 2.4 provides empirical results and section 2.5 concludes.

2.2 Recent Employment Trends and Institutional Setting

2.2.1 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, 2012a), 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.

2.2.2 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.

Unemployment Insurance

The duration and level of unemployment benefit entitlements influences the incentives for reemployment (Hunt, 1995; Lalive and Zweimüller, 2004; Kyyrä 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.[1] 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.[2] 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.[3]

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, 2000a).

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[4] that reorganises the specific retirement rules for unemployed individuals.[5] The reform was originally enacted in March 2001.

[1]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.

[2]See *§ 147 SGB III Grundsatz* of the German social security legislation for details about benefit duration and age.

[3]For a discussion of the common trend assumption, see section 2.3.4.

[4]The legislative change was part of the so-called *Altersvermögensergänzungsgesetz* (AVmEG) from March 2001, as published in *BGBl. 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 *BGBl. I 2002, Nr. 12, p. 754.*

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

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 2.4 in appendix 2.A 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 2.3.2 for a detailed discussion).

2.3 Data and Identification Strategy

2.3.1 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).[6] Furthermore, civil servants are excluded from the analysis due to different legislative rules. After conditioning on a set of control variables (see table 4.3), the final estimation sample consists of 971 individuals (19,795 person-months at risk).

2.3.2 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.

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 2.2) 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.[7] The explanation is

[6] 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). 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’.

[7] For this reason we exclude the dummy that indicates sex from the estimation and let the treatment

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).

2.3.3 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.

Table 2.1: 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.

As apparent from table 4.3, 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 (2012); Hainmueller and Xu (2013). This method allows to achieve exact balancing

indicator capture variation between male and female individuals.

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 4.3 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 4.3). 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.

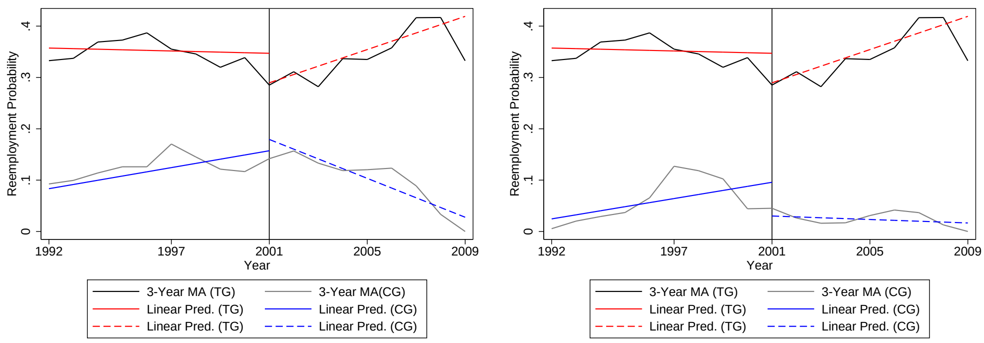
Pre-retirement agreements are a tool to downsize the workforce. Contracts are canceled by mutual agreements and employees receive a compensation. On this pathway, individuals enter unemployment using the compensation and unemployment benefits to bridge the gap until they reach the ERA. The take-up of 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.

2.3.4 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 2.1 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.

Figure 2.1: Reemployment across Control- and Treatment Group.



(a) Unbalanced Sample (b) Balanced Sample

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.

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 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).[8] 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 2.1 are not conditional on observed variables such as the regional unemployment rate.

2.3.5 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, to other undefined states[9], 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 4.4).

Table 2.2: Transitions out of Unemployment and Number of Spells.

	Treatment Group		Control Group		
Transition into...	BC $<$ 1946	BC $>$ 1945	BC $<$ 1946	BC $>$ 1945	Total Spells
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.

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

[9] 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.

The treatment group accounts for a total number of 594 observations (about 61%), while the control group consists of 377 observations (roughly 39%). Table 4.4 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.[10] To overcome these 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

[10] Although the probit framework may condition on the length of unemployment spells which could be included as a regressor, it does not provide appropriate mechanisms to account for right censoring.

can be written as

$$\lambda(t) = \lim_{h \to 0} \frac{\text{Prob}[t \leq T < t + h\,, \varepsilon = 1 | T > t\, or (T \leq t\,, \varepsilon \neq 1)]}{h} \tag{2.1}$$

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

$$\lambda(t|x, \varepsilon) = \lambda_{\varepsilon}(t) \exp[\text{X}\beta_1 + \beta_2\text{Treat} + \beta_3\text{AddMonths} + \beta_4\text{Treat} * \text{AddMonths}] \tag{2.2}$$

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 (2.2), the treatment effect of the policy change is estimated by the coefficient β_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.

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.[11]

2.4 Results

2.4.1 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 2.3 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

[11]The results from this exercise are available from the authors upon request.

term “Treat * AddMonths” identifies the treatment effect of the raised ERA on the reemployment probability.

Table 2.3 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%.[12] The probability, that these results occur by chance 1.1% (Continuous Treatment, CR Balanced) and 1.9% (Binary Treatment, CR Balanced) and thus the estimated treatment effects are statistically significant assuming conventional error probabilities.

Figure 4.3 plots cumulative incidence functions as predicted from the competing risks model for both the unbalanced and balanced sample (table 2.3, 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.[13] In this sense, the difference between the cumulative incidence functions between groups in figure 4.3 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

[12]Note that the interpretation is restricted to estimations as based on the balanced control group. The hazard ratios are $\exp(0.035) = 1.0356$ (Continuous Treatment, CR Balanced), $\exp(1.029) = 2.798$ (Binary Treatment, CR Balanced).

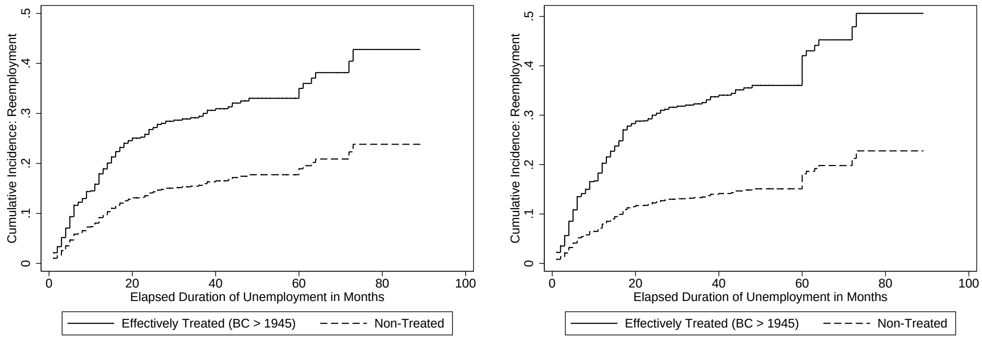
[13]Effectively treated individuals thus belong to the treatment group and are born after 1945, while non-treated individuals are all remaining observation of the estimation sample, i.e. the full control group and those individuals of the treatment group who were born before 1946.

Table 2.3: Reemployment Probability: DiD considering “Other Exits” as Competing Risk.

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

Figure 2.2: 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.

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%).

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 2.5 and 2.6 in appendix 2.B). 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.[14] 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 2.6). 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.[15] 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).

2.5 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 increased due to the reform. Using data from the German Socio-Economic Panel allows to control for a rich set of socio-demographic factors, which 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.

[14] 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).

[15] 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).

The reason is that raising the ERA expands the time horizon until an old age pension is available in total by three years. Holding everything else constant, we investigate whether this exogenously prolonged unemployment period stimulates reemployment. The reasoning is that, if the state of unemployment becomes less attractive, 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.

Appendix

2.A Reform Steps

Table 2.4: Reform Steps of Raising the ERA (Excerpt from Social Security Code).

	NRA		ERA	
Year of Birth	Years	Months	Years	Months
Before 1946	65	0	60	0
1946				
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
1947				
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
1948				
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
1949 - 1951	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.

2.B Results: Sensitivity Analysis

Table 2.5: Reemployment Probability: Continuous Treatment DiD

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

Note: Own calculations based on the SOEP (1991-2012). Robust standard errors in parentheses. α is the additional parameter of the Weibull distribution that allows for duration dependence, where α = 1 implies the exponential distribution. θ 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 2.6: Reemployment Probability: Binary Treatment DiD

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

Note: Own calculations based on the SOEP (1991-2012). Robust standard errors in parentheses. α is the additional parameter of the Weibull distribution that allows for duration dependence, where α = 1 implies the exponential distribution. θ 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.

Chapter 3

The Effect of Benefit Reductions on the Retirement Age: The Heterogeneous Response of Manual and Non-Manual Workers

3.1 Introduction

Demographic change jeopardises pay-as-you-go pension systems (see e.g. Börsch-Supan, 2000a; Hairault et al., 2010). Increasing life expectancy and lower birth rates may require individuals to contribute longer and claim benefits at higher ages (Federal Statistical Office, 2012b, 2014). Benefit reductions in the case of early retirement are one possible way to incentivise postponed retirement and thus to redesign a pay-as-you-go pension system towards more actuarial neutrality. However, the loss of work capacity may distract individuals from continued work and induce early retirement. Thus, questions arise as to whether individuals are heterogeneous in terms of their ability to postpone retirement.

The purpose of this paper is to analyse the response in retirement behaviour with respect to a major reform that introduced benefit reductions into the German public pension system between 1997 - 2004. Pension benefits are reduced by 3.6 per cent for each year (i.e. 0.3 per cent/month) by which an old age pension is claimed early.[1] Those benefit reductions are permanent, which means that they prevail for all periods of one's benefit receipt. Looking forward in terms of expected present discounted values therefore reveals that, ceteris paribus, retirement incomes are remarkably lower once benefit reductions apply in case of early retirement. The central research question of this paper is to what extent individuals postpone retirement (by how many months) in response to benefit reductions. Most importantly, do manual and non-manual workers respond heterogeneously?

To answer these questions, discrete time duration models for exits into retirement are estimated for the birth cohorts 1935 - 1945. Starting at age 60, hazard rates are estimated to predict the mean duration for exits into retirement. Identification is based on a natural experiment, where the intensity of benefit reductions (i.e. the magnitude of benefit reduction) is a function of the date of birth only. Exploiting this type of exogenous variation rules out a common critique that factors to determine social security wealth such as previous earnings are highly correlated to labour market attachment which may confound the estimated effect of financial incentives (see Krueger and Pischke, 1992). Important retirement patterns for the age group 60 - 65 are taken into consideration by modeling age-in-months using respective dummies. Thus, age enters the model in the most flexible way and therefore allows to control for retirement due to institutional reasons or unobserved social norms.
Results are derived separately on two diametrically opposed data sources. First, we use administrative data of exceptional quality which contain worker biographies with exact retirement entries on a monthly basis (Himmelreicher and Stegmann, 2008). The definition of retirement for these social security records is claiming a pension as documented from the administrative process. Second, survey data are used to draw on a rich set of individual socio-demographic information which is essential to model the complex decision process that underlies retirement behaviour. Here, retirement is de-

[1]In most cases, claiming an old age pension "early" refers to ages previous to the normal retirement age of 65, which corresponds to the legal rule for the observation period between 1995 and 2010. For more institutional details, see section 3.2.

fined as respondents' self-reported retirement status.

The results clearly indicate that introducing benefit reductions induces postponed retirement. On average, retirement is postponed by 5 months due to financial incentives, where the delay in retirement is about 7 months for men and about 4.5 months for women. This finding is robust across data sources and distributional assumptions. The magnitude of the estimated effect is slightly lower than previous findings for Germany but by and large in line with these studies (Börsch-Supan and Schnabel, 1999; Hanel, 2010).

The response to benefit reductions does differ to some extent by manual and non-manual workers. Manual workers postpone retirement only by about 3 months on average while non-manual workers are found to postpone by 8 months. Thus, the response to benefit reductions is some 60 per cent lower for manual workers compared to non-manual workers. The absolute difference of 5 months between the two groups is equivalent to 1.5 per cent of benefit reductions. For a German retiree with average pensions of 13440 Euros per year (40 contribution years at the average earnings level, evaluated at the annuity value of 2014) the difference in benefit reductions amounts to roughly 200 Euro for each remaining year to live.

While the difference between manual and non-manual workers is significant and considerably large its proximate cause is a specific retirement pattern at age 63. Old age pensions due to disability are available without benefit reductions at this age and the results show that manual workers take this option at large scale. The striking feature is that this has not been the case for manual workers previous to the introduction of benefit reductions. Clearly, manual workers tend to maximise the sum of future benefits by claiming disability pensions at the earliest date without reductions. Delays in retirement are less probable for manual workers not simply because formerly harsh occupations are correlated to poor health but also because a specific retirement scheme sets incentives to retire at age 63. Non-manual workers are indeed more flexible on retirement timing but benefit reductions are not necessarily larger for manual workers if they take reduction-free disability pensions to retire early. Consequently, benefit reductions seem not to be a considerable source of inequality in terms of social security wealth across workers that differ by physical demands of occupations. The bottom line is that the impact of financial incentives on manual workers is lower because alternative pathways into retirement do exist.

This paper is related to several strands of the existing literature. The closest studies are probably those by Börsch-Supan and Schnabel (1999) and Hanel (2010). The former report an increase of about six months in mean retirement ages from ex-ante simulations of the reform. The latter finds that financial incentives as imposed by benefit reductions induce postponed benefit claims by up to 14 months. For the U.S., an early contribution simulates an increase in the reduction factor from 6.66 per cent to 15 per cent for each year of early retirement (Fields and Mitchell, 1984b) and finds that individuals postpone retirement by some three months as a response to this policy change. More recent studies on the impact of financial incentives on the retirement age also find postponed retirement in response to financial incentives in terms of raised employment Pingle (2006); Blau and Goodstein (2007). The present paper adds to this literature by confirming the common finding that people respond to incentives if the implicit tax on further periods of work is reduced.

Second, influential studies point at the importance of an income measure that discounts all future streams of benefits while taking into account uncertain lifetimes when examining the relationship of financial incentives and retirement. (Fields and Mitchell, 1984a; Mitchell and Fields, 1984). In contrast to the wealth level or current income at a given point in time, the forward-looking perspective is much more valuable in terms of an incentive measure. This study is based on the expected present discounted value (EPDV hereafter) of all future retirement benefits and thus relates to this strand of the literature. A large body of literature has evolved using the EPDV or accruals of it to measure the expected gain from postponing retirement by one period (see e.g. Samwick, 1998; Börsch-Supan and Schnabel, 1999; Börsch-Supan, 2000b; Coile et al., 2002; Hanel, 2010).[2]

[2]Several studies apply versions of the option value (Stock and Wise, 1990). While theoretically appealing, the empirical implementation is difficult because it requires data to estimate the parameters of a CRRA utility function. Estimating a fully structural model is usually circumvented by assuming parameter values for the preferences over risk and leisure (Samwick, 1998; Börsch-Supan, 2000b; Blundell et al., 2002; Asch et al., 2005). Moreover, variation in the option value is predominantly determined by variation in wages which may cause problems when identifying the impact of social security benefits on retirement behaviour (see Coile and Gruber, 2000, for a discussion).

Finally, the key finding of a heterogeneous response of manual and non-manual workers to benefit reductions is, to the best of my knowledge, a novel one. Differentiating these worker types adds to the literature on health-related aspects of retirement decisions. An important finding from the previous literature is that the response to financial incentives among disability pensioners is more likely for individuals in good health (Hanel, 2012). The results from this paper suggest that non-manual workers are in fact more flexible when taking the aggregate view on delaying retirement. However, it also highlights the likelihood that manual workers are more flexible than they seem to be. It is important to take a close look at alternative pathways into retirement (here: disability pensions) that may incentivise specific behaviours.

The remainder of this paper is structured as follows. Section 3.2 provides an overview on the institutional setting and the corresponding reform. Section 3.3 provides a theoretical background and formulates hypotheses to be tested. Section 3.4 outlines the data, econometric strategy and identification issues. Section 5.3 presents results and section 5.4 concludes.

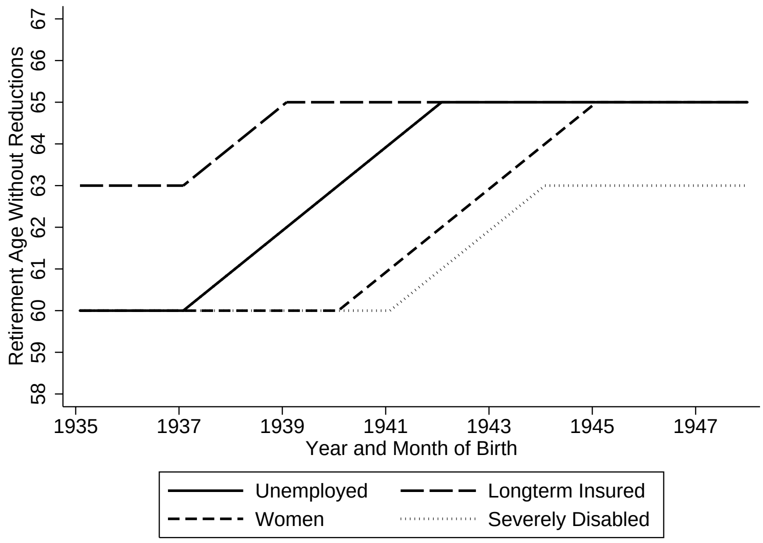
3.2 Institutional Setting

The German public pension system is organised as a pay-as-you-go system. Due to generous early retirement rules from the 1970s to the 1990s, the pension system imposed strong disincentives to supply labour after age 60 (see e.g. Börsch-Supan and Schnabel, 1998, 1999; Börsch-Supan, 2000a). To remove disincentives and to stabilise both the contribution- and replacement rate has led to several reforms starting in the early 1990's.

One influential reform introduced financial incentives to postpone retirement. Benefit reductions for early retirement were implemented between January 1997 and December 2004 for birth cohorts born from 1937 to 1944. As of birth cohort 1945, all individuals are fully affected by benefit reductions if they retire early. The reform imposes an adjustment factor to the pension formula, which effectively reduces pension claims by 0.3 percentage points for each month of early retirement relative to the normal retirement age (NRA). The NRA was fixed at 65 for the relevant period.[3] For a whole year of early retirement the reduction thus amounts to 3.6 percent of monthly retirement benefits and the maximum reduction is 18 per cent if retirement takes place five years previous to the NRA, i.e. at age 60.

The minimum age to receive an old age pension early is 60 years and an individual is eligible, if she is *(i)* unemployed, *(ii)* a woman, *(iii)* has contributed for at least 35 years or *(iv)* is disabled.[4] Only those individuals who fulfill the aforementioned requirements can receive an old age pension previous to the NRA of 65.[5] The reforms phased in gradually from 1997 to 2004 but the timing differed by types of eligibility (figure 4.2).

Figure 3.1: Gradual Increase of Retirement Age Without Reductions across Eligibility Types.



Source: Wachstums- und Beschäftigungsförderungsgesetz (1996); Rentenreformgesetz 1999 (1997); Korrektur des Rentenreformgesetzes 1999 (1998).

The gradual increase in the reduction-free retirement age by types of eligibility is shown in figure 4.2. The reform raised the reduction-free retirement age of an old age pension

[3]The adjustment factor is administered in § 77 SGB VI, German Social Security Code.

[4]The final rules for old age pensions due to unemployment, for women and for long-term insured individuals were enacted in the Wachstums- und Beschäftigungsförderungsgesetz (1996), while the legislative change for old age pensions for severely disabled individuals was enacted in the Rentenreformgesetz 1999 (1997) and slightly changed in the Korrektur des Rentenreformgesetzes 1999 (1998).

[5]Note that the legislative rules which are relevant for both the reform (1997-2004) and the observation period (1995-2010) implied an NRA of 65. A recent reform that raises the NRA from 65 to 67 is implemented between 2012 and 2029, but has no relevance for this study.

(i) due to unemployment for the birth cohorts 1937 - 1941, *(ii)* for women born between 1940 - 1944, *(iii)* for long-term insured persons born between 1937 - 1938, and *(iv)* for disabled individuals born between 1941 - 1943. For example, an unemployed individual born in January 1937 who retires in the first month of eligibility for an old age pension due to unemployment (i.e. in January 1997) exhibits a benefit reduction of 0.3 per cent. This is because the reduction-free retirement age for this person has been raised by one month.

3.3 Conceptual Framework and Hypotheses

The decision to retire crucially relies on individual preferences in combination to external incentives that are set by the social security system. Models where utility is maximised over the arguments of consumption and leisure and subject to a lifetime budget constraint can explain retirement decisions while taking into account the interdependence between preferences and incentives (see e.g. Samwick, 1998). The optimal retirement age is determined exactly by a point in time, where an infinitesimal change in the retirement date induces a change in the utility from leisure which is just offset by the change in utility from consumption as indirectly derived from financial resources. Individuals postpone retirement if a later retirement date corresponds to a utility gain from the future stream of retirement benefits. This argument is central for the present study, because benefit reductions may reduce present discounted values if exits into retirement take place previous to the normal retirement age.

To transmit this view into the empirical framework of this study, the notion of Samwick (1998) is followed where the "actuarial present value" (APV) of all future retirement benefits of an individual is given by

$$\mathrm{APV}(R) = \int_R^T e^{-r(s-R)} B_s(R) ds \tag{3.1}$$

and is conditional on retirement at age R. Time is measured discretely in months for the relevant age group. That is, net present values are discrete sums and computed for individuals of age 60 to 66, i.e. for 72 months, starting in the month after an individuals' 60th birthday. The resulting measure is the "expected present discounted value" (EPDV)

$$\text{EPDV}(R) = \sum_{s=R}^{T} \pi(s)\delta^{s-t} B_s(R) \tag{3.2}$$

which is the future stream of all future retirement benefits B, discounted by the rate δ to time t and weighted by conditional survival probabilities π as provided by the Federal Statistical Office (2012b). Note that values for EPDV are calculated at each month between age 60 and 66, i.e. $t = 1, ..., 72$ for each individual.[6] The existing literature points out that the use of forward-looking incentive measures supports identification of the impact of social security benefits on the retirement age ((see e.g. Samwick, 1998; Coile and Gruber, 2000; Coile et al., 2002). Third, financial incentives from the introduction of benefit reductions need to be measured. This is incorporated into equation (3.2) by supplementing an adjustment factor $(1-\tau(R))$, which can be written as

$$\text{EPDV}_{\tau(R)}(R) = (1 - \tau(R))\,\text{EPDV}(R) = (1 - \tau(R)) \sum_{s=R}^{T} \pi(s)\delta^{s-t} B_s(R) \tag{3.3}$$

where $0 \leq \tau(R) \leq 0.18$ is the reduction rate as implied by the German social security legislation. Whether benefit reductions apply (i.e. $\tau(R) > 0$) and to what extent depends on *(i)* the exact retirement age, *(ii)* year and month of birth, and *(iii)* the type of old age pension.[7] Finally, subgroup heterogeneity in the response to benefit reductions is revealed for manual and non-manual workers by computing mean durations as predicted from estimated models across subgroups.

We are primarily interested in the identification of the impact of benefit reductions on the timing of retirement. The relevant variable is the percentage difference between

[6] It is assumed, that individuals do not live beyond age 100, as survival probabilities are not available for older individuals (see Federal Statistical Office (2012b)).

[7] See section 3.2 for a detailed description of the reform.

EPDV (equation (3.2)) and the adjusted EPDV (equation (3.3)) which is the adjustment rate $\tau(R)$ since we have $\text{BRR}(R) = \frac{\text{EPDV}(R) - \text{EPDV}_{\tau(R)}(R)}{\text{EPDV}(R)} = \frac{\text{EPDV}(R) - [1-\tau(R)]\text{EPDV}(R)}{\text{EPDV}(R)} = \tau(R)$. Consequently, all subsequent empirical models include two separate main regressors: "BRR" and "EPDV".[8] Precisely, this involves the assumption that the unobserved counterfactual for the treatment group is the observed factual of the control group ($\tau(R) = 0$).

Two fundamental hypotheses are tested. Let a denote the normal retirement age and let $a - x$ denote some early retirement age. Suppose that monthly retirement benefits are reduced by 3.6 per cent if retirement takes place at age $a-x$ compared to retirement at age a.

Hypothesis 1 *Financial incentives from benefit reductions lower the relative attractiveness of early retirement at age $a - x$. Increasing the reduction rate $\tau(R)$ raises the probability that transitions into retirement are postponed (e.g. from $a - x$ to a).*

Let π_A be the survival probability for non-manual workers and let π_B be the survival probability for manual workers. Further suppose that $\pi_A > \pi_B$.

Hypothesis 1a *The relative change in the hazard rate for exits into retirement is larger for non-manual workers compared to manual workers. That is, the degree of postponement for transitions into retirement is larger for non-manual workers as compared to manual workers.*

3.4 Data and Empirical Strategy

3.4.1 The Two Datasets

The study uses two very different data sources. First, administrative data of exceptional quality are used. The Insurant Account Sample (Versicherungskontenstichprobe, VSKT) is provided by the German Federal Pension Insurance (Himmelreicher and Stegmann, 2008). Altogether, the VSKT is a sample of about 240,000 individuals of

[8]Please note that both EPDV and BRR depend on the retirement date R independently. For BRR, this is determined by the rules of the policy change and for EPDV this is determined by total pension claims.

age 14 to 66 regarding their insurance accounts and randomly drawn in annual waves from the population of employees who are subject to social security contributions. This study draws on the waves 2002 - 2010, using a 25% subsample of the VSKT which is provided as scientific use file and includes some 60,000 individual observations per wave. The sample is selective to the extent, that it only collects entries for employment that is subject to social security contributions. [9] An outstanding feature is that social security records are process-produced and thus do not suffer from typical problems of survey data. However, available information is limited to contribution time, monthly amounts of contribution (which allows to calculate benefit entitlements), retirement entry date and a few socio-demographic variables such as age, sex and region.

In a comparative scenario, survey data from the German Socio-Economic Panel Study (SOEP) are used from the panel waves 1995 to 2011.[10] The SOEP is representative for the German population and includes some 11.000 households and about 20.000 individuals (Haisken-DeNew and Frick, 2005). Subjects are repeatedly interviewed over several years. The SOEP contains retrospective calendar data on employment and retirement and these activity spells are on a monthly basis. In contrast to administrative data, a rich set of individual- and household level socio-demographic variables is available to identify worker heterogeneity in the response to financial incentives; marriage status, health status and occupational information to distinguish manual workers from non-manual workers are available.

3.4.2 Sample Construction and Descriptive Statistics

A similar data structure is constructed for both data sources, where the focus is on duration times until exits into retirement are observed. For the VSKT, retirement is defined as benefit claim without measurement error. For SOEP data, retirement is defined as self-reported retirement status in retrospective questions which are likely to exhibit measurement error. Both samples are restricted to person-month-observations in the relevant age range from 60 to 66. Precisely, old age pensions are available as of

[9] Civil servants and self-employed individuals are ruled out from the analysis. While the German pension systems offers the possibility to contribute voluntarily and to accumulate benefit entitlements this is a minor group. Further note that civil servants can be part of the sample, if they formerly contributed as employees.

[10] Note that the observation period is 1995 to 2010, but the analysis includes retrospective questions from 2011 which correspond to the previous year.

the early retirement age (ERA) if eligibility is achieved.[11] The sample includes spells for the birth cohorts 1935 to 1945. For 14660 observed individuals (i.e. spells) and 407663 person-month-observations in social security records (VSKT), table 3.1 reports that roughly 37 per cent claim regular old age pensions (available at age 65), which is the largest group among old age pensions. The second largest group are old age pensions for women (28 per cent) as followed by old age pensions due to unemployment (16 per cent). About 3 per cent of the total spells are right-censored, which means that no exit into retirement has taken place until age 66.[12]

Table 3.1: Observations across Birth Cohorts.

	VSKT (Recorded Eligibility Type)						
Cohort	ROAP	UE	W	SD	LI	Censored	Total
1935	456 (19027)	232 (1519)	342 (1721)	100 (570)	105 (4106)	48 (3456)	1283 (30399)
1936	454 (20350)	234 (1684)	380 (1700)	94 (642)	132 (4989)	39 (2808)	1333 (32173)
1937	497 (22126)	264 (2109)	370 (1970)	81 (592)	107 (4103)	47 (3384)	1366 (34284)
1938	509 (21268)	225 (2376)	365 (2054)	98 (756)	88 (3197)	38 (2736)	1323 (32387)
1939	504 (22169)	239 (2937)	417 (2325)	94 (1046)	83 (3162)	44 (3168)	1381 (34807)
1940	483 (23334)	251 (3147)	399 (3486)	121 (1050)	88 (3342)	32 (2304)	1374 (36663)
1941	449 (21414)	225 (4244)	414 (4833)	149 (1087)	85 (3260)	35 (2520)	1357 (37358)
1942	486 (23104)	164 (2834)	406 (6734)	170 (1563)	83 (3418)	45 (3240)	1354 (40893)
1943	507 (25693)	151 (2727)	368 (6896)	159 (1794)	91 (3603)	39 (2808)	1315 (43521)
1944	497 (25593)	158 (2160)	357 (7758)	187 (2345)	102 (3686)	37 (2664)	1338 (44206)
1945	536 (27273)	169 (2879)	265 (3953)	164 (2430)	98 (4186)	4 (251)	1236 (41089)
Total	5378 (251351)	2312 (28616)	4083 (43430)	1417 (13875)	1062 (41052)	408 (29339)	14660 (407663)
Share	0.367 (0.617)	0.158 (0.070)	0.278 (0.107)	0.097 (0.034)	0.072 (0.101)	0.028 (0.071)	1(1)

	SOEP (Retrospective Retirement Status)		
Cohort	Retired	Censored	Total
1935	43 (1167)	2 (83)	45 (2062)
1936	56 (1491)	3 (216)	59 (1707)
1937	66 (1578)	3 (184)	69 (1762)
1938	88 (1893)	14 (838)	102 (2731)
1939	101 (2062)	9 (554)	110 (2616)
1940	184 (4051)	29 (1402)	213 (5453)
1941	128 (3217)	19 (775)	147 (3992)
1942	180 (4805)	43 (1523)	223 (6328)
1943	135 (4019)	43 (1370)	178 (5389)
1944	165 (4882)	37 (1218)	202 (6100)
1945	95 (2911)	54 (2114)	149 (5025)
Total	1241 (32076)	256 (10277)	1497 (42353)
Share	0.829 (0.757)	0.171 (0.243)	1(1)

Source: Own calculation based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011). *Note:* Reported values by birth cohort and type of old age pension. Person-Month-Observations in parentheses. Abbreviations are: Regular old age pension (ROAP), old age pension due to unemployment (UE), old age pension for women (W), old age pension for severely disabled (SD) and old age pension for longterm insured (LI).

The lower part of table 3.1 reports 1497 observed individuals (i.e. spells) for survey data (SOEP), which amounts to a total of 42353 person-month-observations. There are

[11]See section 3.2 for the four relevant types of old age pensions.

[12]Right-censoring is taken into account as it is explicitly modeled in the likelihood function of subsequent duration models.

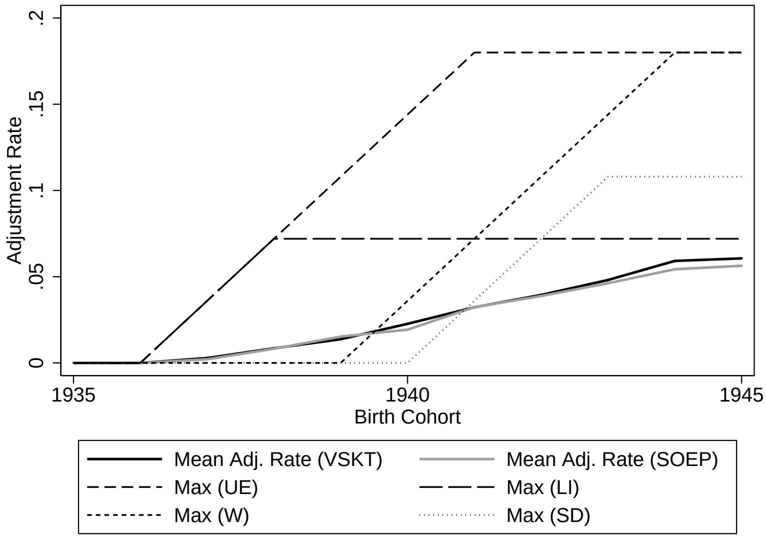
some key differences between the two data sources with respect to sample construction. First, the number of spells and person-month-observations is much larger for social security records. Second, eligibility types of old age pensions are observed for the VSKT, but cannot be discriminated for the SOEP (empty columns). Third, right-censoring is much more prevalent in the SOEP (about 17 per cent). These findings indicate that retirement entries are reported with less precision in the SOEP.

The choice of particular birth cohorts from 1935 to 1945 is important and needs some explanation. First, financial incentives were introduced for the birth cohorts 1937 to 1944, which implies that birth cohorts previous to the reform serve as control group, i.e. those cohorts in the sample that are not affected by benefit reductions (1935 and 1936). During the implementation period, the treatment intensity differs by month and year of birth on the one hand and the type of old age pension on the other hand. By the end of 2004, all reduction free age limits are completely raised, such that all individuals that claim an old age pension are fully affected if they retire early. Second, the analysis in this study is based on a quasi-experimental setting. The only reason why individuals face benefit reductions is their date of birth. However, unobserved birth cohort heterogeneity may confound identification of the effects of financial incentives on the timing of retirement. For this reason the number of birth cohorts must be kept small such that no systematic differences in retirement behaviour materialise.

A variable that indicates duration time is equal to one in the first month of eligibility (i.e. the month after the 60th birthday) and then counts each subsequent month. Individuals are not allowed to enter the sample after age 60, which holds for social security records as well as for survey data. This restriction is important as it rules out left-censoring and allows to observe all individuals at risk to retire starting from the same age. Each individual is observed as long as no exit into retirement has taken place. In the specific month, where retirement is observed, the spell ends and the corresponding individual is not observed for further periods. The dependent variable takes the value zero for all months that are previous to the retirement entry and is equal to one in the month, where retirement takes place.

Figure 3.2 illustrates the gradual increase of benefit reductions across birth cohorts. The solid lines (VSKT: black; SOEP: gray) illustrate how the mean adjustment rate at retirement entry gradually increases from zero (birth cohorts 1935 and 1936) to roughly 6% (birth cohort 1945). Mean adjustment rates are slightly lower in the SOEP but follow very much the same pattern as for the VSKT. As eligibility types cannot be discriminated in the SOEP and no information on reasons for retirement is observable, all individuals are assigned the average adjustment rate that is prevalent at some given calendar time. Maximum adjustment rates for all four types of old age pensions illustrate how the treatment intensity increases across birth cohorts as implied by the corresponding legislative rules.

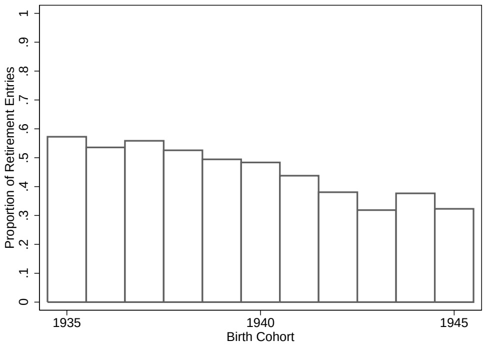
Figure 3.2: Adjustment Rates at Retirement across Birth Cohorts.



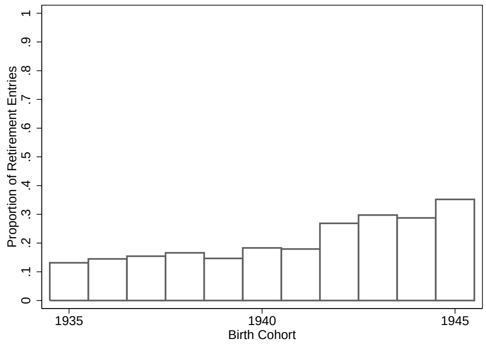
Source: Own calculations based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011).
Note: Mean adjustment rates are arithmetic means at retirement entry, plotted across birth cohorts for both social security records (VSKT) and survey data (SOEP). Maximum adjustment rates are for different types of eligibility for old age pensions. Abbreviations are UE (Unemployed), LI (Longterm Insured), W (Women) and SD (Severely Disabled).

Figure 4.3 shows the share of retirement entries for age 60 and age 65 across birth cohorts for men (panel a and b) and women (panel c and d). A clear pattern suggests that younger birth cohorts retire at higher ages. Retirement at age 60 reduces by roughly 20 percentage points from almost 60 per cent (cohort 1935) to about 40 per cent (cohort 1945) for both men and women. The decrease is particularly large for the birth cohorts 1939 to 1943, i.e. for those who retired during the implementation period of the reform when the treatment intensity successively increased. Contrarily, retirement at age 65 increases across birth cohorts for both males and females. This increase is more pronounced for male cohorts, where the difference between birth cohorts 1935 and 1945 is about 23 percentage points from 12 per cent (1935) to about 35 per cent (1945). For female cohorts, the share of retirees at age 65 is initially higher at about 30 per cent and increases by roughly 5 percentage points to about 35 per cent.[13] Clearly, these patterns are purely descriptive but they do indicate that retirement ages increase by birth cohorts during the reform period.

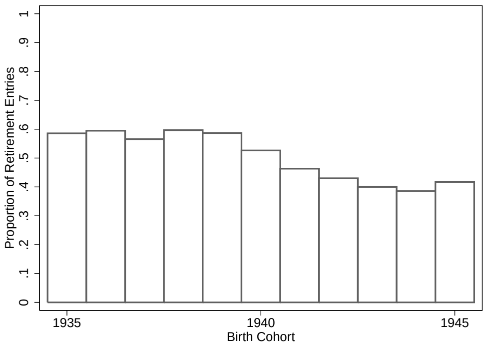
Figure 3.3: Share of Retirees within Birth Cohorts for Selected Ages.



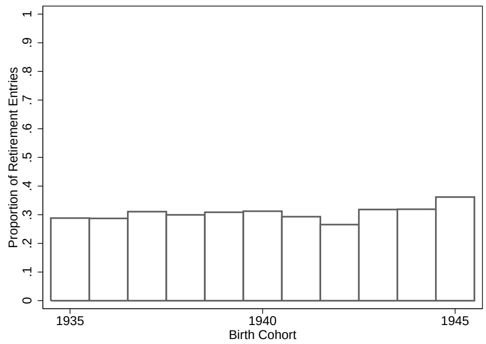
(a) Age 60: Male



(b) Age 65: Male



(c) Age 60: Female



(d) Age 65: Female

Source: Own calculations based on SUFVSKT2002-SUFVSKT2010.

Table 4.3 provides an overview on all variables that are part of the analysis.[14]

The variable "BRR" has a somewhat larger mean for survey data, which is due to the assumptions to assign reduction rates to individuals when eligibility types are not known. The variable "Manual" is a dummy that equals one if individuals are manual workers and zero if they are non-manual workers. The recoding of this variable follows

[13] Those differences are in line with official statistics as reported in German Federal Pension Insurance (2014), where average retirement ages for men are initially (i.e. in 1995) lower but eventually align to those of women.

[14] For brevity, the table does not report duration-dummies, year-dummies and eligibility-type-dummies which are available from the author upon request.

Table 3.2: Descriptive Statistics.

	VSKT			
	Mean	Std.Dev.	Min.	Max.
Dependent Variable				
Benefit Claim	.0348	.1833	0	1
Reform Variables				
BRR	.0133	.0362	0	.18
EPDV	8023.1	4557.4	6.1	23510.9
Adjusted EPDV	7902.3	4477.3	6.1	23510.9
Control Variables				
Male	.3900	.4877	0	1
West Germany	.7284	.4448	0	1
N (Person-Month)	14660(407663)			

	SOEP			
	Mean	Std.Dev.	Min.	Max.
Dependent Variable				
Retirement	.0291	.1682	0	1
Reform Variables				
BRR	.0600	.0585	0	.18
EPDV	8348.4	4069.6	261.3	22078.9
Adjusted EPDV	7807.5	3735.7	214.2	22078.9
Manual	.5155	.4998	0	1
BRR * Manual	.0287	.0502	0	.18
Control Variables				
Male	.4690	.4990	0	1
West Germany	.7959	.4030	0	1
Married	.8535	.3536	0	1
Years of Education	11.54	2.69	7	18
Poor Health	.3664	.4818	0	1
Moderate Health	.4290	.4949	0	1
Good Health	.2046	.4034	0	1
N (Person-Month)	1497(42353)			

Source: Own calculation based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011).
Note: Reported values are for monthly records, i.e. person-month observations, to take into account variation over time. EPDV is the expected present discounted value and BRR is the benefit reduction rate.

the International Standard Classification of Occupations from 1988 (ISCO 88). Finally, table 4.3 reports the interaction term "BRR * Manual", which is given by the product of the two respective variables. This interaction term is crucial for the analysis with survey data, as it reflects to what extent the response of manual and non-manual workers differs. The remaining variables in table 4.3 are control variables in subsequent regressions. For survey data, the two sexes more or less balance out with slightly fewer male individuals. This is a good representation when taking into account that male individuals have somewhat lower life times. However, female individuals seem to be somewhat over-represented in social security records.

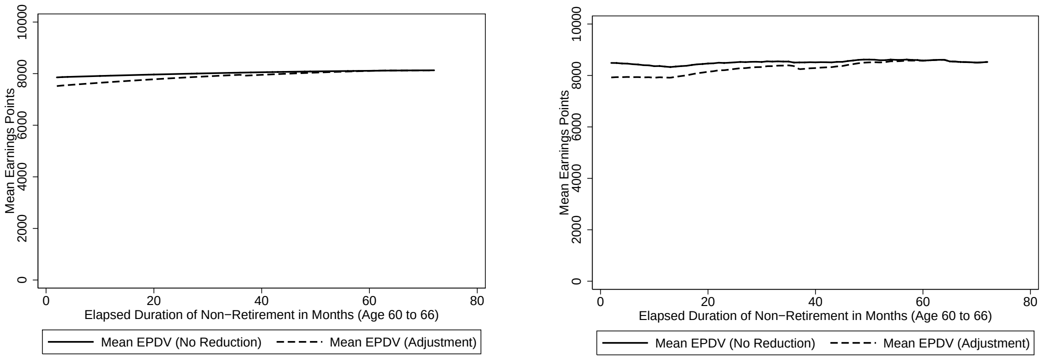
The computation of the EPDV (and the corresponding incentive measure) is central because the estimation of the effect of benefit reductions on the retirement age is based on it. For the computation of the EPDV, information on pension entitlements is essential and the two data sources provide different information in this respect. For social security records (VSKT) pension claims are measured in so-called earnings points (EP). EP reflect the relative income position of an individual for a given year, which implies that an individual with average earnings yields exactly one EP. Correspondingly, an individual with twice the average earnings yields two EP. The EPDV is calculated using EP without transforming them into real units.[15] For survey data (SOEP), access to pension entitlements is not as straightforward. Here, gross labour income is averaged over earnings biographies for all observed values in a first step. In a second step, the relative income position for each individual and year is calculated, which directly yields EP as defined above. Since earnings biographies are not completely observed, it is assumed that individuals have contributed for 35 years when reaching age 60. Thus, average earnings points are multiplied by the factor 35 in a final step to generate pension entitlements in the SOEP.

One further challenge for the computation of the EPDV in social security records (VSKT) is that no information on further accumulation of pension claims is available, once an individual has retired previous to age 65. The most straightforward assumption is that for each individual, the observed average monthly pension claim is extrapolated to higher ages. For survey data (SOEP), average earnings points as indirectly derived from labour income are extrapolated from age 60 to 66.[16] Under the assumption, that individuals never grow older than 100 years (i.e. 480 months from the 60th birthday), and using conditional survival probabilities, the sample mean of the EPDV is plotted in figure 3.4. Evidently, EPDVs follow similar patterns for both data sources but are measured with less precision for the SOEP.

[15] As of 2013, one EP is worth between 26 and 28 Euros, depending on the region where pensions are claimed. This means that an individual that has accumulated 40 EP, receives a pension of about $40 \times 28 = 1120$ Euros per month (assuming a regular old age pension without reductions). Each year, the annuity value for pension claims is set according to a formula of the German social security legislation (§ 68 SGB VI) which takes into account wage growth and changes in the share of retirees.

[16] An alternative assumption was tested and led to decreasing EPDVs over duration time but did not change any estimation results. In this scenario, pension claims are held constant (i.e. no further accumulation) if an exit into retirement takes place before age 65.

Figure 3.4: Expected Present Discounted Value (Sample Mean).



(a) VSKT (b) SOEP

Source: Own calculations based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011).

3.4.3 Econometric Strategy

To identify the impact of benefit reductions on the timing of retirement decisions, discrete time duration models are estimated. Such models are commonly used in the literature to analyse transition behaviour (see e.g. Lancaster, 1979; Meyer, 1990). Individuals at risk are followed until they exhibit the failure event retirement or censoring otherwise. Individuals are censored if they do not exit into retirement before the observation period ends. The discrete time duration framework is advantageous in this context as it *(i)* allows to control for right-censored spells, *(ii)* explicitly takes into account the discrete measurement of time in months, *(iii)* allows for a large number of transitions at particular points in time. Such probability mass points are controlled for by implementing the most flexible version of a duration model with duration dummies for each point in time. Discrete time proportional hazards are modeled by assuming type-I-extreme-value distributed spell lengths (complementary log-log model).[17] The choice of this distribution is motivated by the fact that rare events (i.e. retirement entries) are analysed, that accumulate at very few duration times (i.e. individual ages).

The derivation of the sample likelihood function as proposed by Prentice and Gloeckler (1978) and Meyer (1990) is moved to appendix 3.A. The final estimation equation is

[17] As for robustness of the results, complementary log-log models are supplemented by a logit model (assuming a logistic distribution, i.e. proportional odds) and a probit model (assuming a normal distribution).

the log-likelihood function

$$\ln L(\gamma,\beta)=\sum_{i=1}^{N}\left[\delta_i \ln\left[1-\exp\left[-\exp\left(z_i(k_i)'\beta+\gamma(k_i)\right)\right]\right]-\sum_{t=1}^{k_i-1}\exp\left(z_i(t)'\beta+\gamma(t)\right)\right] \tag{3.4}$$

where $z_i(\cdot)$ is a vector of time-varying explanatory variables (containing the main regressors "BRR", "EPDV" and other explanatory variables) for individual i and β is the vector of corresponding unknown parameters. The log-likelihood is maximised using the "cloglog"-procedure as implemented into the software STATA.

In the presence of unobserved heterogeneity, however, the results from this model may be biased. Selection bias may be present, if individuals have high hazard rates that are correlated to unobserved characteristics. These individuals leave the sample systematically earlier than the selected pool of survivours. Abbring and Van den Berg (2007) show that the distribution of unobserved heterogeneity among survivours often rapidly converges to a gamma distribution. Following this logic, the hazard function from equation (3.7) (appendix 3.A) must be rewritten as

$$\lambda_i(t)=\theta_i\lambda_0(t)\exp(z_i(t)'\beta) \tag{3.5}$$

where θ_i is a random variable that captures unobserved heterogeneity which is assumed to be multiplicatively linked to the hazard function and to be independent from observed characteristics $z_i(t)$. If θ_i follows a gamma distribution with mean one and variance σ^2 finally yields the log-likelihood function

$$\begin{aligned}\ln L(\gamma,\beta,\sigma^2)=\sum_{i=1}^{N}\ln\Bigg[&\Big[1+\sigma^2\sum_{t=0}^{k_i-1}\exp\left(z_i(t)'\beta+\gamma(t)\right)\Big]^{-\sigma^{-2}}\\ &-\delta_i\Big[1+\sigma^2\sum_{t=0}^{k_i}\exp\left(z_i(t)'\beta+\gamma(t)\right)\Big]^{-\sigma^{-2}}\Bigg]\end{aligned} \tag{3.6}$$

which is maximised using the STATA procedure "pgmhaz8" (Jenkins, 2004).

3.4.4 Identification of the Effect of Benefit Reductions on the Retirement Age

Hazard rates of entering retirement are estimated using the aforementioned methodology. Starting in the month after the 60th birthday (first eligibility), individuals are observed repeatedly until they enter retirement. The dependent variable takes value zero for each month of non-retirement and the value one if an individual enters retirement.[18]

Identification is based on a differences estimator by discriminating between control- and treatment group. Control and treatment group are determined by the date of birth and the eligibility type. Clearly, the introduction of benefit reductions must be identified as a source of exogenous variation in order to be interpreted as a causal effect. The identifying assumption is that the individual choice of retirement age would not have changed in the treatment group (as compared to the control group) in absence of the introduction of benefit reductions. This involves the assumption that no birth cohort heterogeneity is present and under such circumstances the reform can be interpreted as a natural experiment. The narrow window of birth cohorts (1935 - 1945) supports the view that these cohorts do not considerably differ by unobserved factors. Moreover, adaptive behaviour previous to the reform is unlikely to confound the estimates. There are no incentives to change the retirement decision for individuals who are not affected. Finally, secular time trends are controlled in the regression framework.

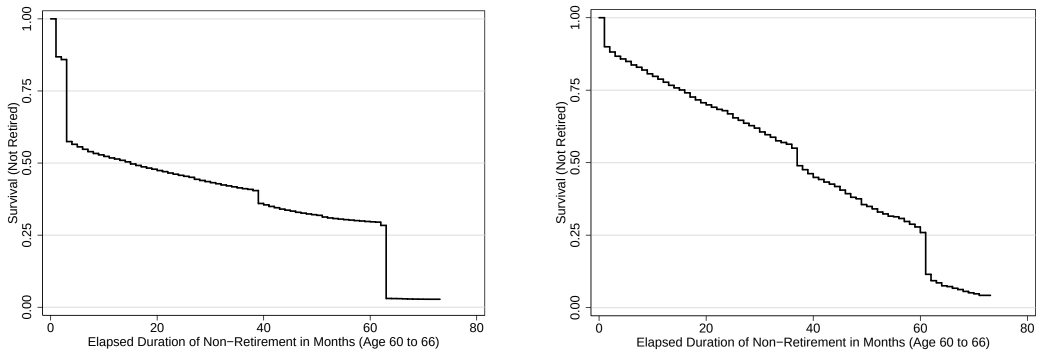
3.4.5 Specific Data Patterns

Specific patterns with respect to the retirement age are shown in figure 3.5 and occur for the following reasons. First, the institutional setting strongly influences the timing of retirement through the channel of early and normal retirement ages. That is, availability for different types of old age pensions is a driving force for retirement decisions. Second, occupational agreements with employers that do not necessarily comply with applicable law may influence retirement patterns. Third, social norms may as well be

[18] An important difference arises with respect to social security records and survey data. While panel attrition does not take place in social security records besides few exceptions, it does so in survey data. In the former, leaving the state "employee" to self-employed or civil servant may cause an exit from the sample but is a rare event. However, in survey data panel attrition frequently occurs (e.g. refused further participation).

influential. While the first aspect is easily traced back, the other two are difficult to measure.

Figure 3.5: Retirement Entry and Probability Mass Points.



(a) VSKT (b) SOEP

Source: Own calculations based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011).

From the Kaplan-Meier survival estimates in figure 3.5 we can infer that a large fraction of retirement entries takes place exactly in the month of first eligibility, i.e. at age 60, 63 and 65. Those probability mass points cause a considerable amount of variation that needs to be taken into account when identifying the causal effect of benefit reductions. The subsequent analysis employs the most flexible version of a discrete time duration model in which duration dummies (i.e. age) are included for every month and thus control for cumulative retirement entries at specific ages.

3.5 Results

3.5.1 Baseline Estimation

Discrete time transition rates into retirement for both social security records and survey data are estimated. All relevant information from social security records is used while the resulting models are replicated using survey data. Baseline estimations serve for comparative purposes and it is important to keep in mind that these data sources differ by the type of information that they provide. Thus, the definition of retirement differs across data sources (VSKT: benefit claiming; SOEP: self-reported retirement) and different assumptions for the calculation of benefit entitlements are necessary. Moreover, eligibility types are precisely identified in the VSKT but not for the SOEP. The results are provided in table 3.3.[19]

Table 3.3: Baseline Estimation: Benefit Reductions and Retirement Transitions.

Variable	Logit		Probit		Compl. Log-Log	
	m.eff.	s.e.	m.eff.	s.e.	m.eff.	s.e.
Social Security Records (VSKT)						
BRR	−.068	(.007)	−.079	(.007)	−.067	(.007)
EPDV	.000	(.000)	.000	(.000)	.000	(.000)
Male	−.008	(.001)	−.007	(.001)	−.006	(.001)
West	−.009	(.001)	−.009	(.001)	−.009	(.001)
+ Eligibility-Type-Dummies						
+ Year-Dummies						
+ Duration-Dummies						
Mean Transit. Rate (%)	3.48		3.50		3.47	
Obs.(Person-Month-Obs.)	14660(407663)		14660(407663)		14660(407663)	
Survey Data (SOEP)						
BRR	−.190	(.030)	−.198	(.030)	−.183	(.030)
EPDV	.000	(.000)	.000	(.000)	.000	(.000)
Male	.008	(.002)	.009	(.002)	.007	(.002)
West	−.023	(.003)	−.021	(.002)	−.023	(.003)
+ Year-Dummies						
+ Duration-Dummies						
Mean Transit. Rate (%)	2.92		2.92		2.92	
Obs.(Person-Month-Obs.)	1527(43245)		1527(43245)		1527(43245)	

Source: Own calculation based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011). *Note:* Reported values are average marginal effects. For factor variables, reported values are the discrete change corresponding to the reference category. Standard errors in parentheses. Mean transition rates are predicted from respective models and reported in per cent. EPDV is the expected present discounted value and BRR is the benefit reduction rate.

The key regressor BRR reflects individual adjustment rates, which are exogenously determined by year and month of birth. This variable varies between 0 for individuals at ages where no adjustment applies up to 0.18 (i.e. 18%) for individuals at ages where the maximum adjustment applies. For social security records, the marginal effect for BRR suggests, that increasing the adjustment rate by one percentage point reduces the probability to retire by 0.07 percentage points on average for a given point in time. While this effect seems small in absolute terms, evaluated at the predicted mean transition rate of 3.47%, this is an average decrease of 2% in the probability to observe a transition into retirement in a given period at risk. For survey data, the marginal effect for “BRR” indicates that increasing the adjustment rate by one percentage point reduces the probability to retire by 0.18 percentage points on average for a given point

[19] All binary choice models as presented across columns, i.e. logit, probit and complementary log-log, are robust over corresponding distributional assumptions. For this reason, the subsequent discussion concentrates on results as obtained from the discrete time proportional hazard model (complementary log-log).

in time. Evaluated at the predicted mean transition rate of 2.92%, this is an average decrease of 6% in the probability to observe a transition into retirement in a given period. For all estimated models in the baseline scenario, the probability that this result occurs by chance is very small (<0.001). Thus, in terms of conventional error probabilities, the null hypothesis of no impact of benefit reductions is rejected. This is in favour of Hypothesis 1, that benefit reductions lower the attractiveness of early retirement and thus induce postponed retirement. While the results are very similar with respect to their sign, they do differ by magnitude across data sources.

Differences in magnitude of estimated coefficients and average marginal effects are not surprising when taking into account that the underlying data sources differ substantially by quality and quantity of information that they provide.
Omitted variable bias in regressions using the VSKT may attenuate estimates towards zero and explain the smaller marginal effects. Retirement decisions are outcomes from a complex mix of determinants. Aside from financial resources as discussed in section 3.3, the literature is clear about other influential determinants such as marital status and health. Married individuals may condition their retirement entry decision on their spouses retirement behaviour (see e.g. Blau and Riphahn, 1999). Individual health may play an important role in the timing of retirement, which has been subject to many previous studies (see e.g. Berkovec and Stern, 1991; McGarry, 2004).
For survey data (SOEP), measurement error in regressors may attenuate estimated coefficients towards zero. For both the VSKT and the SOEP the estimated response in retirement timing is a lower bound for the true (unknown) response.

3.5.2 The Heterogeneous Response of Manual and Non-Manual Workers

Estimating a richer model with further information on worker heterogeneity shows that the marginal effect in the baseline estimation on survey data seems to be biased towards zero to some extent (table 3.4). Consequently, the marginal effect of benefit reductions as measured by "BRR" is larger in absolute terms (i.e. more negative).
Raising the adjustment rate by one percentage point results in a decrease of the probability to retire by 0.22 percentage points on average for a given point in time. Evaluated

Table 3.4: Benefit Reductions, Retirement Transitions, and Worker Heterogeneity.

Variable	Logit		Probit		Compl. Log-Log	
	m.eff.	s.e.	m.eff.	s.e.	m.eff.	s.e.
Survey Data (SOEP)						
BRR	–.226	(.034)	–.234	(.033)	–.220	(.034)
EPDV	.000	(.000)	.000	(.000)	.000	(.000)
Manual	.003	(.002)	.002	(.002)	.002	(.002)
BRR X Manual	.063	(.026)	.068	(.026)	.067	(.026)
Male	.006	(.002)	.007	(.002)	.005	(.002)
West	–.022	(.003)	–.020	(.003)	–.022	(.003)
Married	–.010	(.003)	–.010	(.003)	–.009	(.003)
Years of Education	.000	(.000)	.000	(.001)	.001	(.000)
Moderate Health	.001	(.002)	.001	(.002)	.002	(.002)
Good Health	.001	(.002)	.001	(.002)	.001	(.002)
+ Year-Dummies						
+ Duration-Dummies						
Mean Transit. Rate (%)	2.91		2.91		2.92	
Obs.(Person-Month-Obs.)	1497(42353)		1497(42353)		1497(42353)	

Source: Own calculation based on SOEP (1995-2011). *Note:* Reported values are average marginal effects. For factor variables, reported values are the discrete change corresponding to the reference category. Standard errors in parentheses. Mean transition rates are predicted from respective models and reported in per cent. EPDV is the expected present discounted value and BRR is the benefit reduction rate.

at the predicted mean transition rate of 2.92%, this is an average decrease of 7.5% in the probability to observe a transition into retirement in a given period. However, the central finding in this model is the positive marginal effect for the interaction "BRR * Manual". This result indicates, that manual workers respond to a much lower degree to benefit reductions compared to non-manual workers. For both, "BRR" and "BRR * Manual", the probabilities that these results occur by chance are not larger that 0.01 and thus the null hypothesis is rejected supporting Hypothesis 1a. Further results suggest that living in West Germany and being married are in favour of retiring at higher ages.

Explicitly assuming and modelling an unobserved heterogeneity distribution following Prentice and Gloeckler (1978); Meyer (1990) yields coefficient estimates that are robust in magnitude and sign to all previous results (see appendix 3.B) suggesting that these findings are robust unlikely to be confounded by unobserved factors.

3.5.3 Predictions

Expected duration times from first eligibility (age 60) until retirement are predicted to quantify the response to benefit reductions by subgroups. All values as reported in table

3.5 are predicted from the complementary log-log model. On average, individuals who are exposed to benefit reductions postpone retirement by about five months compared to the control group. The response for men is somewhat larger which is induced by lower duration times before the reform. This result is stable across all models and it does coincide with officially reported retirement ages that were lower for male individuals in the mid-1990s but aligned to those of women during the observation period, indicating a larger response for men (German Federal Pension Insurance, 2014).[20] Most importantly, manual workers postpone retirement only by about three months which implies that their behavioural response is some 60 per cent lower compared to non-manual workers who postpone by 8 months.

Interestingly, predicted differences in retirement postponement with respect to benefit reductions are more or less identical across data sources. This indicates that retirement decisions as documented in VSKT and SOEP are very similar. Despite this similarity, duration times *in levels* are somewhat lower for the VSKT (table 3.5). The previous discussion on marginal effects suggests that part of the explanation are omitted variables (VSKT) and measurement error (SOEP) while the true size of the bias is likely to be small but unknown. Moreover, sample selection may play a minor role but imposing similar sampling restrictions should make this difference rather small. However, the most relevant explanation here is right censoring in the SOEP. While the VSKT comprises a negligible share of right-censored observations (3 per cent) the share in the SOEP (17 per cent) is much larger (see section 3.4). This may contribute to an over-estimation of the total duration until retirement and thus explain the larger level in duration times as documented in the SOEP.

Predicted survival functions for baseline estimations are shown in figure 3.6. The predictions highlight the general finding that benefit reductions induce postponed retirement. The response for male individuals is larger, starting from lower survival rates in the control group. For social security records, the 95% confidence bands do not overlap in the relevant interval roughly between the fifth and sixtieth month of eligibility. Generally, the 95% confidence bands are much more narrow for social security records (VSKT),

[20]Differentials in retirement ages (old age pensions) between men and women depend on the region, with some variation across former east and west Germany. In 1995 and across all regions, old age pensions were claimed at age 62.5 (women) and 62.3 (men), indicating that men retired two months earlier on average. This difference successively disappeared and as of 2004, the retirement age of men even slightly surmounted the one of women.

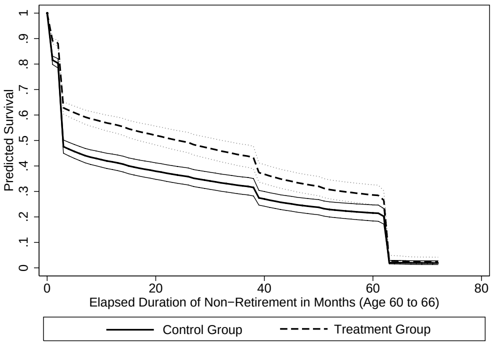
Table 3.5: Expected Duration of Non-Retirement (from First Month of Eligibility).

	Baseline: VSKT		
	Before Reform	During/After Reform	Difference
Full Sample	23.2	28.5	5.3
Male	20.9	27.9	7.0
Female	24.7	28.9	4.2
	Baseline: SOEP		
	Before Reform	During/After Reform	Difference
Full Sample	27.3	32.4	5.1
Male	24.9	32.1	7.2
Female	28.3	32.8	4.5
	Further Information: SOEP		
	Before Reform	During/After Reform	Difference
Full Sample	27.2	32.5	5.3
Male	25.3	32.5	7.2
Female	27.8	32.6	4.8
Manual	26.6	29.7	3.1
Non-Manual	28.1	36.0	7.9

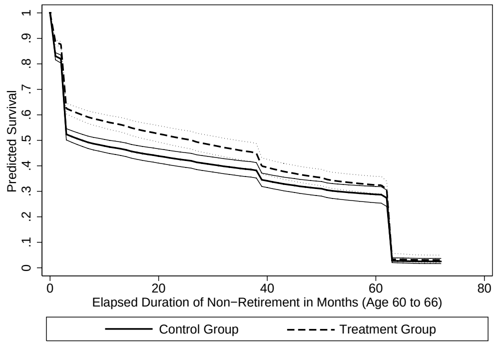
Source: Own calculation based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011). *Note:* Reported values are computed from complementary-log-log models, i.e. discrete time proportional hazard models.

indicating that these results are measured with much more precision. The low number of observations in the SOEP induces 95% confidence bands to be wide and to overlap; however, the overall pattern is very similar to the one predicted from the VSKT. Predicted survival functions for the estimation including further information from the SOEP are shown in figure 3.7. The focus is on differences in the behavioural response manual versus non-manual workers (pooled for men and women). The plotted lines confirm that the response to benefit reductions is much larger for non-manual workers. Moreover, their survival rates are considerably larger in absolute terms. The striking result is, however, that predicted survival rates of manual workers in the treatment group exhibit a remarkable drop at age 63. At this age, old age pensions due to

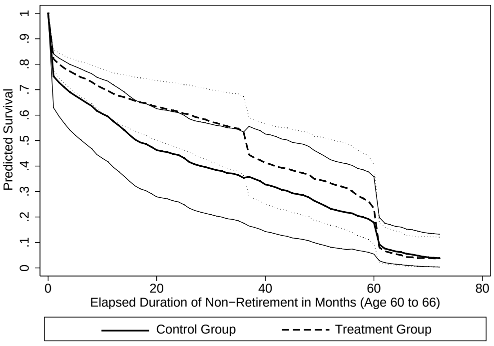
Figure 3.6: Predicted Survival Rates: Baseline Estimation.



(a) VSKT: Men



(b) VSKT: Women



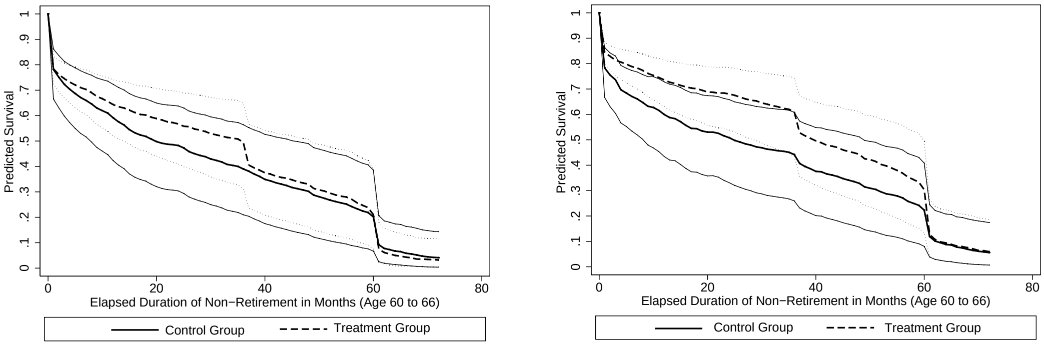
(c) SOEP: Men

(d) SOEP: Women

Source: Own calculation based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011). *Note:* Thin lines are 95% confindence bands (solid: control; dotted: treatment).

disability are available without reductions after full implementation (recall figure 4.2). Manual workers who are affected by the reform seem to choose this alternative as soon as it is available without benefit reductions. No such drop occurs for manual workers in the control group which clearly indicates that other types of old age pensions served as pathways into retirement before benefit reductions became effective (e.g. old age pensions due to unemployment). The high take-up rate of disability pensions among manual workers is in accordance to a presumably higher prevalence of poor health compared to non-manual workers. Again, it must be noted that 95% confidence bands indicate some imprecision in measured differences as group sizes are initially small and decrease over duration time, when fewer individuals are still at risk.

Figure 3.7: Predicted Survival Rates for Manual versus Non-Manual Workers.



(a) Manual Workers (b) Non-Manual Workers

Source: Own calculation based on SOEP (1995-2011). *Note:* Thin lines are 95% confindence bands (solid: control; dotted: treatment).

3.6 Conclusion

I estimate the effect of benefit reductions on the timing of retirement separately for men and women of age 60 - 66. Identification is based on a natural experiment where the intensity of benefit reductions (i.e. their magnitude) is a function of the date of birth only. Specific interest is on differences in the response of manual and non-manual workers.

I show that, on average, individuals postpone retirement by five months if pension benefits are reduced by 3.6 per cent for each year of early retirement. This confirms the common finding that people respond to incentives if the implicit tax on further periods of work is reduced. However, manual workers postpone retirement only by 3 months on average and thus their response is about 60 per cent lower compared to non-manual workers who postpone by 8 months.

Surprisingly, this does not indicate that retirement incomes of manual workers deteriorate. The explanation is that large scale retirement into disability pensions takes place at age 63 - without benefit reductions. This alternative pathway into retirement is attractive for manual workers because claiming disability pensions at the earliest date without reductions allows them to maximise the sum of future benefits. The clear message from this result is that manual workers do postpone retirement but only until they become eligible for a specific alternative that is financially attractive. Consequently, delays in retirement are less probable for manual workers not simply because formerly harsh occupations are correlated to poor health but also because a specific retirement scheme sets incentives to retire earlier (at age 63). Non-manual workers are indeed more flexible on retirement timing but this does not necessarily mean that retirement incomes of manual workers deteriorate.

The results suggest a sizable difference for the delay in retirement between men (7 months) and women (4.5 months). By coincidence, officially reported retirement ages were lower for men in the mid-1990s but aligned to those of women during the observation period (German Federal Pension Insurance, 2014) which helps to explain the larger response for men. Moreover, the analysis is conducted on social security records (VSKT) and on survey data (SOEP) and therefore allows to obtain separate results that differ by data quality and the definition of retirement. While the predicted differences in the duration until retirement (after age 60) are surprisingly similar these predictions differ to some degree by their level. This difference is mainly due to right-censored spells which are much more prevalent in the SOEP and are likely to cause an overestimation of retirement timing relative to the VSKT.

From the perspective of actuarial fairness, there is no rationale against benefit reductions. Manual workers receive benefits earlier at a higher reduction rate if they delay retirement to a lower extent. Redistribution in the PAYG system from manual to non-manual workers depends on the total duration of benefit receipt and is thus related to differential mortality expectations which is beyond the scope of this study. The central finding is that manual workers are more flexible than they seem to be because they wait for disability pensions until they are available without benefit reductions. This strongly suggests that the institutional setting is a relevant determinant for retirement patterns once it imposes financial incentives.

Appendix

3.A Derivation of the Sample Likelihood Function

Let T_i denote a random variable for the duration of individual i in a spell of non-retirement and let t denote an arbitrary point in time where a failure takes place. Let the instantaneous probability to exhibit a failure event at time t conditional on survival until time t be represented by the hazard function

$$\lambda_i(t) = \lim_{h \to 0} \frac{\text{Prob}[t + h > T_i \geq t | T_i \geq h]}{h} = \lambda_0(t) \exp(z_i(t)'\beta) \tag{3.7}$$

where h is an infinitesimal instant of time, $\lambda_0(t)$ is the unknown baseline hazard, $z_i(t)$ is vector of time-varying explanatory variables for individual i and β is a vector of unknown parameters. The hazard function in equation (3.7) is parameterised assuming proportional hazards. Maximum likelihood estimation of the parameters of this hazard function requires the construction of a sample likelihood function as proposed by Prentice and Gloeckler (1978) and extended by Meyer (1990). In this framework, no specific assumptions about the functional form of the baseline hazard need to be made. Following Meyer (1990), the discrete time proportional hazards model reduces to a complementary-log-log model with extreme value distributed spell lengths. The probability that a spell lasts until time $t + 1$, given that it has lasted until time t can be written as a function of the hazard (i.e. the survival function $S(t)$)

$$\begin{aligned}\text{Prob}[T_i \geq t + 1 | T_i \geq t] = S(t) &= \exp\left[-\int_t^{t+1} \lambda_i(u) du\right] \\ &= \exp\left[-\exp(z_i(t)'\beta) \int_t^{t+1} \lambda_0(u) du\right]\end{aligned} \tag{3.8}$$

where the otherwise continuous hazard function $\lambda_i(\cdot)$ is integrated over a discrete interval from t to $t + 1$. The second line of equation (3.8) makes use of the proportional hazards specification in equation (3.7) and exploits the fact that $z_i(t)$ is constant between t and $t + 1$. Reformulating the survival function yields

$$\text{Prob}[T_i \geq t+1 | T_i \geq t] = S(t) = \exp\left[-\exp(z_i(t)'\beta + \gamma(t))\right] \tag{3.9}$$

where $\gamma(t) = \ln\left[\int_t^{t+1} \lambda_0(u)du\right]$ such that the corresponding cumulative distribution function is given by $F(t) = 1 - S(t) = 1 - \exp\left[-\exp(z_i(t)'\beta + \gamma(t))\right]$. Note that this is exactly the cumulative distribution function of the type-I-extreme value distribution (Gumbel distribution). Estimation of the hazard for a sample of N individuals is achieved by estimating the parameters of the likelihood function

$$L(\gamma, \beta) = \prod_{i=1}^{N} \left[1 - \exp\left[-\exp\left(z_i(k_i)'\beta + \gamma(k_i)\right)^{\delta_i}\right] \times \prod_{t=1}^{k_i-1} \exp\left[-\exp\left(z_i(t)'\beta + \gamma(t)\right)\right] \right] \tag{3.10}$$

where $\delta_i = 1$ if an individual spell exhibits a transition into retirement and $\delta_i = 0$ if individual duration times T_i are censored.

3.B Unobserved Heterogeneity

Unobserved heterogeneity is examined following (Prentice and Gloeckler, 1978; Meyer, 1990) as outlined in section 3.4. A reduced model without full flexibility with respect to the baseline hazard is estimated because the fully flexible model lacks convergence. The reduced model includes three dummies to capture variation at entry ages where probability mass points occur but requires less computational power. The respective dummy variables (PM) summarise duration time (i.e. age in months) for the subsequent three months after age 60, 63 and 65 and calendar time enters the model linearly (variable “Year”). The baseline results for VSKT and SOEP are reported in table 3.6 and the results for the model including further information are reported in table 3.7.[21] All models are estimated as a complementary log-log model without frailty and in a version, where gamma frailty is assumed. Corresponding likelihood ratio tests indicate that the null hypothesis of a gamma variance equal to zero is rejected under conventional error probabilities, which means that unobserved heterogeneity is likely to be

[21]For computational reasons, these models report estimated coefficients and no average marginal effects as in previous estimations.

Table 3.6: Baseline Estimation: Benefit Reductions and Retirement Transitions (Frailty).

	Compl. Log-Log without Frailty		Mixed Gamma (Frailty)	
	coeff.	s.e.	coeff.	s.e.
Social Security Records (VSKT)				
BRR	–2.256	(.251)	–2.424	(.263)
EPDV	.000	(.000)	.000	(.000)
Male	–.253	(.029)	–.273	(.030)
West	–.388	(.021)	–.403	(.022)
Year	–.069	(.003)	–.069	(.003)
PM 60 + 3	1.285	(.023)	1.247	(.024)
PM 63 + 3	1.444	(.038)	1.469	(.038)
PM 65 + 3	4.995	(.032)	5.040	(.033)
+ Eligibility-Type-Dummies				
Gamma Variance			0.041 (.009)	
LR Test: Gamma Variance = 0			p < 0.001	
Obs.(Person-Month-Obs.)	14660(407663)		14660(407663)	
Survey Data (SOEP)				
BRR	–6.976	(.659)	–7.710	(.876)
EPDV	.000	(.000)	.000	(.000)
Male	.216	(.064)	.231	(.073)
West	–.708	(.011)	–.792	(.090)
Year	.050	(.011)	.058	(.014)
PM 60 + 3	2.293	(.084)	2.248	(.090)
PM 63 + 3	.904	(.113)	.905	(.113)
PM 65 + 3	2.565	(.097)	2.642	(.111)
Gamma Variance			0.151 (.110)	
LR Test: Gamma Variance = 0			p < 0.001	
Obs.(Person-Month-Obs.)	1527(43245)		1527(43245)	

Source: Own calculation based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011). *Note:* Reported values estimated coefficients. Standard errors in parentheses. Unobserved heterogeneity is assumed to follow a gamma mixed distribution. EPDV is the expected present discounted value and BRR is the benefit reduction rate. Probability mass points are abbreviated by PM, where "+ 3" indicates that duration times (months) after the respective birthday are summarised in each dummy.

present. However, estimated coefficients are robust in magnitude and sign for all models suggesting that previous findings are robust and are unlikely to be confounded by unobserved factors.

Table 3.7: Benefit Reductions, Retirement Transitions, and Worker Heterogeneity (Frailty).

	Compl. Log-Log without Frailty		Mixed Gamma (Frailty)	
	coeff.	s.e.	coeff.	s.e.
Survey Data (SOEP)				
BRR	–8.525	(.890)	–9.356	(1.027)
EPDV	.000	(.000)	.000	(.000)
Manual	.105	(.082)	.170	(.099)
BRR X Manual	2.630	(1.006)	2.448	(1.086)
Male	.157	(.067)	.170	(.077)
West	–.691	(.072)	–.816	(.102)
Married	–.333	(.074)	–.408	(.092)
Years of Education	.224	(.081)	.234	(.092)
Moderate Health	.065	(.065)	.064	(.071)
Good Health	.081	(.079)	.095	(.087)
Year	.046	(.011)	.057	(.014)
PM 60 + 3	2.279	(.085)	2.212	(.091)
PM 63 + 3	.907	(.113)	.909	(.114)
PM 65 + 3	2.564	(.097)	2.669	(.111)
Gamma Variance			0.211 (.113)	
LR Test: Gamma Variance = 0			$p < 0.001$	
Obs.(Person-Month-Obs.)	1497(42353)		1497(42353)	

Source: Own calculation based on SOEP (1995-2011). *Note:* Reported values estimated coefficients. Standard errors in parentheses. Unobserved heterogeneity is assumed to follow a gamma mixed distribution. EPDV is the expected present discounted value and BRR is the benefit reduction rate. Probability mass points are abbreviated by PM, where "+ 3" indicates that duration times (months) after the respective birthday are summarised in each dummy.

Chapter 4

Retirement Age and Inequality of Opportunity: The Role of Physical Job Demands

4.1 Introduction

In 2007, the German Federal Parliament approved a law to gradually increase the normal retirement age from 65 to 67. The decision was accompanied by a public debate on the justness of a legal retirement age fixed at a high level towards workers being exposed to different levels of physical job demands. In particular, high physical demands of job duties may force a worker to retire early because work-related health impairments accumulate over time and human physical capacity declines naturally with increasing age. This may disadvantage respective occupations since retirement previous to the normal retirement age reduces benefit entitlements for two reasons. First, fewer years of work imply fewer years of contribution and thus lower pension claims, since their overall amount depends on the duration of preceding contributions. Second, early retirement is subject to actuarial adjustments, which additionally reduces pension claims by 3.6 per cent for each year by which an old age pension is claimed early.[1] In addition, occupations with high physical job demands tend to be low-wage professions, which

[1] Actuarial adjustments have been introduced in the German public pension system between 1997 and 2004.

restricts the potential for private pension provision.

The purpose of this paper is to quantify differences in the retirement age between manual and non-manual workers and to evaluate these differences with respect to equality of opportunity. The focus is on the question how individual background during childhood transmits through physical demands of occupations on retirement ages. Our study contributes to the existing literature in several respects. First, we provide a precise empirical description of labour force dynamics of older manual and non-manual workers with a particular focus on retirement patterns. Second, we contribute to the literature on equality of opportunity (EOP hereafter) as prominently discussed by Roemer (1993, 1998), by distinguishing between individual background beyond individual influence on the one hand and effort and luck on the other hand. This framework is useful to structure thoughts in a debate, where early retirement of manual workers is frequently declared as “unfair” because this usually implies a reduction in social security wealth. To the best of our knowledge, this is the first study on EOP in retirement age. Finally, we use data for Germany which provides an eminent case study for analyses of an ageing population. Demographic change will continue to impose a fair amount of pressure on the pay-as-you-go pension system. The baby boom cohorts born between 1955 and 1970 move towards retirement as of the year 2015, once they successively become eligible for old age pensions. At the same time, younger birth cohorts are much smaller (Federal Statistical Office, 2014). Severe population ageing will either induce raising contributions, alleviate benefit entitlements or both.

Our analysis departs from a description of labour force dynamics of older workers beginning at age 40 to elaborate differences in the retirement age for manual workers compared to non-manual workers. In a first step, discrete time duration models are used to estimate the hazard rate for transitions out of full-time employment, part-time employment or unemployment into retirement. To distinguish manual from non-manual occupations, we make use of a well-defined measure for the degree of physical demands on respective jobs (Kroll, 2011). In a second step, the question of EOP in retirement is elaborated. We begin with a non-parametric test for stochastic dominance at first order, which is applied to the EOP framework as in Lefranc et al. (2009, 2008); Trannoy et al.

(2010). This approach compares the cumulative distribution of the outcome retirement age, conditional on specific individual background characteristics or "circumstances" in the terminology of Roemer (1998). We then proceed to a decomposition as established by Blinder (1973) and Oaxaca (1973). This technique allows us to infer on how much of the difference in retirement ages between manual and non-manual workers is due to circumstances. Finding an answer to this question is crucial when evaluating policy design that may involve benefit reductions for early retirees.

Our results indicate that the estimated hazard profile of non-manual workers is about 20% lower compared to individuals with physically demanding occupations for the age group 55 to 65. Moreover, non-parametric tests for stochastic dominance at first order indicate that the distribution of retirement age differs significantly between individuals across circumstances. Most importantly, the Blinder-Oaxaca decomposition suggests that circumstances explain at least one third of the observed differences in the retirement age between workers of different degrees of physical job demands. This finding is important because it indicates that a considerable part of differences in retirement age are predetermined and thus not subject to individual choice.

The remainder of this paper is structured as follows. Section 4.2 reviews previous research on EOP, discusses ambiguity in the evaluation of early retirement as a "good" or a "bad" and provides an overview on the employment behaviour of older workers. Section 4.3 describes the data and sample construction. Section 4.4 provides the empirical analysis, quantifies differences in retirement age between manual and non-manual workers and attributes these differences to circumstances and effort/luck in a corresponding decomposition. Section 5.4 concludes.

4.2 Equality of Opportunity and Retirement

Modern egalitarian views such as expressed in Rawls (1971); Cohen (1989); Fleurbaey (1995a,b) postulate that, instead of equality in outcomes, distributive justice only requires equality of opportunity in achieving those outcomes. The recent economic literature usually follows the terminology as introduced by Roemer (1998), according to which individual outcomes are generated by two fundamental determinants: "Circumstances" and "effort", which are defined to be orthogonal. While circumstances reflect background characteristics for which an individual cannot be held responsible, differences in outcome due to effort are considered a legitimate source of inequality. Consequently, given equal circumstances, all remaining differences in outcomes are subject to personal responsibility. Lefranc et al. (2009) state that no consensus has been reached so far on how opportunities are precisely defined. They provide an extension of the EOP framework by introducing luck as an additional determinant of individual outcome. Lefranc et al. (2009) conclude that luck is a legitimate source of inequality in outcomes, as long as it is not correlated to circumstances and is thus "even-handed".

A large body of recent economic literature on EOP has emerged, with numerous applications to income distributions (e.g. Devooght, 2008; Lefranc et al., 2008, 2009; Aaberge et al., 2011) and health (e.g. Fleurbaey and Schokkaert, 2009; Trannoy et al., 2010; Jusot et al., 2013). In the present paper, we provide the first application of the EOP framework to inequality in retirement age. Specifically, we investigate the extent to which circumstances are mediated through physical demands of occupations to the age of retirement. This question is often raised in the public debate, where early retirement of manual workers is frequently considered to be "unfair" because this usually implies a reduction in social security wealth.

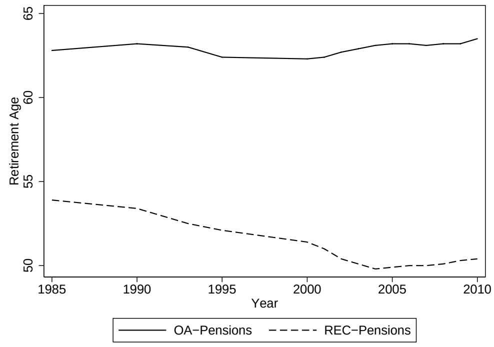
Whether early retirement is a "good" or a "bad" is not unambiguous and deserves some further discussion. In previous applications of EOP, a natural ordering of the outcome of interest is straightforward, i.e. "more income is better than less income" or "good health is better than poor health". Our research question differs from former applications in a sense that such an ordering is not as straightforward for the retirement age; it is difficult to say whether early retirement is good or bad. We take this puzzle as a motivation to briefly outline the view to be taken in this paper. The traditional view on retirement decisions are individual preferences for consumption and leisure in a combination to incentives that are set by the social security system (see e.g. Weiss, 1972; Sheshinski, 1978, for early contributions). Clearly, the consequence of early retirement is more leisure (less work) accompanied by less consumption (lower income from labour and/or social security), and thus retirement is an issue of labour supply (Hurd, 1990). Beyond individual preferences for leisure and consumption, retirement may coincide with subsequent phenomena that either support or prevent an individual to live longer. Retirement may relieve individuals from work-related stress with a positive impact on the remaining years to live, but empirical evidence suggests that cognitive decline sets in after retirement (see e.g. Rohwedder and Willis, 2010; Bonsang et al., 2012). Moreover, no causal link between retirement age and mortality can be established according to Hernaes et al. (2013). Aside from labour supply, employer behaviour may be responsible for the early termination of employment contracts if demand is weak and layoffs are necessary (Hutchens, 1999). In this context, demand-sided factors may induce early retirement even if employees wish to retire later (Hakola and Uusitalo, 2005).

The evaluation of early retirement is obviously an intricate task because a complex mix of "goods" and "bads" needs to be taken into account. For example, an individual with a strong preference for consumption may want to work longer and retire later but may be forced to behave differently due to plant closure or health issues. In the context of this paper, our interest in early retirement refers to the case where individuals retire early for reasons that are correlated to physical demands of occupations. Specifically, we calculate the difference in mean retirement age between manual and non-manual workers to assess the proportion of this difference which is due to individual background characteristics. Therefore, our view on early retirement focuses on its adverse effects, because it reduces retirement benefits and thus social security wealth in a situation where postponed retirement would avoid a decline in retirement income but is either difficult or impossible to realise.

Recent retirement patterns for Germany suggest that the average retirement age increases. However, figure 4.1 indicates that this trend is not quite the same for old age pensions (OA-pensions) and reduced earnings capacity pensions (REC-pensions).[2] Apparently, there is a large difference in the average retirement age between OA-pensions and REC-pensions. Most notably, while the average retirement age for OA-

[2]In contrast to old age pensions which are available after age 60 depending on the eligibility type, reduced earnings capacity pensions are available at any age before 60, once the corresponding medical indication has been assigned.

Figure 4.1: Development of Average Retirement Age in Germany.



Source: German Federal Pension Insurance (2013). *Note:* Retirement ages previous to 1993 are for former West Germany only, while all subsequent values are reported for reunified Germany.

pensions exhibits an upward trend especially after the introduction of actuarial adjustments in the late 1990s, the retirement age for REC-pensions declined. These differences are substantial, which needs some explanation. First, while REC-pensions do not necessarily coincide with physically demanding occupations, eligibility for REC-pensions is usually due to poor health, which itself is expected to be positively correlated to physically demanding occupations. Thus, manual workers should be largely over-represented in the group of individuals that receive REC-pensions. Second, selection into REC-pensions may have changed in the course of a large decrease in the total number of entries into REC-pensions (German Federal Pension Insurance, 2013), leaving a sample of the "worst cases" who retire earlier on average. Consequently, the diverging pattern of average retirement ages for OA-pensions and REC-pensions in figure 4.1 has important implications for the present study as it provides a first hint of differences in the retirement age between manual and non-manual workers; these differences will subsequently be analysed in further detail.

4.3 Data and Sample Construction

The empirical analysis of this paper is based on data from the German Socio-Economic Panel (SOEP) for the waves 1984 to 2012. The SOEP is a representative panel study for Germany which annually interviews households and its individual members since 1984. The survey started with about 6,000 interviewed households per year and comprises about 12,000 households per year since 2000 (see Haisken-DeNew and Frick, 2005). As the focus of this study is on transitions into retirement, we use individual retrospective calendar data on employment spells as provided by the SOEP.[3]

In order to analyse labour force transitions of older workers we restrict our sample to individuals aged 40 and older. For those individuals, who meet this restriction in the observation period from 1984 to 2011, we construct spells with respect to four defined states of labour force participation. Specifically, we distinguish spells of (i) full-time employment, (ii) part-time employment, (iii) unemployment and (iv) retirement. Our primary sample provides 13,304 total transitions from 17,594 individual spells as reported in tables 4.4 and 4.5 in the appendix.[4] Central to our analysis is the number of total transitions into retirement from all other states, which amounts to 3,036.

4.4 Empirical Analysis

4.4.1 Retirement across Physical Job Demands: Evidence from Duration Models

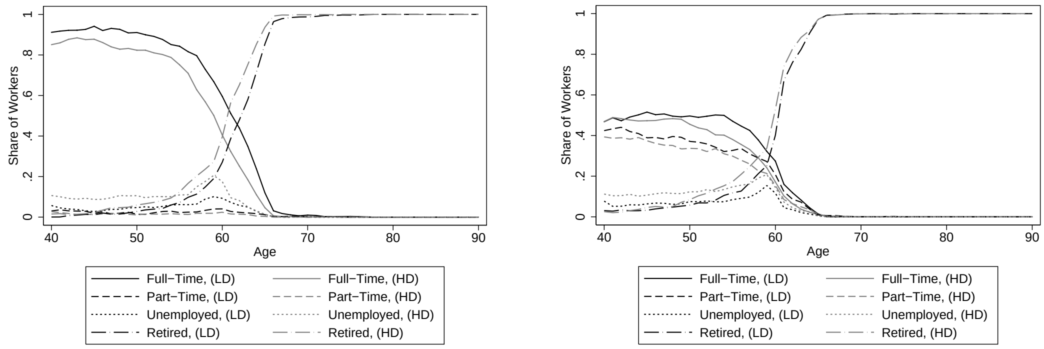
To classify individuals by the physical demands of their reported occupation, we make use of the International Standard Classification of Occupations from 1988 (ISCO 88). This classification serves to categorise physical demands on a 1-10 ordinal scale for physical job demands as constructed by Kroll (2011).[5] Using this index, we categorise individuals into a group of low physical job demands (index values 1-5, abbreviated as LD and also referred to as manual workers), and a group of high physical job demands (index values 6-10, abbreviated as HD and also referred to as non-manual workers). Figure 4.2 distinguishes between low physical demands (LD) and high physical demands (HD) in occupations and shows how shares of individuals in defined labour force states evolve over age separately for the two sexes.

[3]Taking into account retrospective calendar records where individuals are asked to report their labour force participation from the previous year, we effectively draw on information until 2011.

[4]Note that depending on the state of departure and the state of destination for specific transitions, the number of observations varies considerably in subsequent empirical investigations and is thus lower compared to the primary sample.

[5]The scale was developed using data from a large-scale representative survey for Germany from 2006, which collected information on workplace characteristics such as job requirements, main tasks, working conditions and job demands.

Figure 4.2: Distribution of Labour Force States across Age by Occupational Types.



(a) Male

(b) Female

Source: Own calculations based on SOEP (1984-2012). *Note:* High and low (physical) demands are abbreviated by HD and LD respectively.

According to figure 4.2, manual workers both exit full-time employment and enter retirement at lower ages compared to non-manual workers. Exits from full-time employment increase substantially after age 55. Moreover, retirement predominantly takes place between age 55 and 65, which is strongly driven by social security legislation. Aside from social security legislation, retirement patterns as displayed in figure 4.2 capture a rather broad situation where potentially unobserved factors such as mutual agreements or social norms do play a role; such aspects are taken into account in the subsequent duration model. Finally, the figure indicates that full-time employment is more prevalent among male individuals while female individuals work more frequently in part-time employment. As we are primarily interested in labour force dynamics that document transitions into retirement, the observed patterns in figure 4.2 justify that we restrict our analysis to individuals aged 55 to 65 (equivalent to 11 years or 132 months) in subsequent regressions, while capturing all relevant transition dynamics.

Further descriptive evidence for systematic differences in retirement across physical demands is attained from discrete time duration models.
Table 4.1 reports the results for two discrete time duration models, i.e. a binary complementary log-log model assuming discrete time proportional hazards and a multinomial logit assuming discrete time proportional odds. The binary model describes transitions into retirement from any of the other labour force states, while the multinomial logit

Table 4.1: Discrete Time Duration Models: Differences Estimation for Transitions into Retirement by Physical Demands.

	Compl. Log-Log		Multinomial Logit	
	coeff.	s.e.	coeff.	s.e.
Binary Case				
Manual	.247	(.041)	–	–
MNL (Coefficient w.r.t. Outcome):				
Manual (Part-Time)	–	–	–.417	(.098)
Manual (Unemployment)	–	–	.765	(.075)
Manual (Retirement)	–	–	.277	(.063)
+ Duration-Dummies	Yes		Yes	
Predicted Transition Rate (%)	3.00		3.19	
Obs.(Person-Month-Obs.)	5146(210770)		3630(134960)	

Source: Own calculations based on SOEP (1984-2012). *Note:* Reported values are estimated coefficients. For the multinomial logit, coefficients are estimated with respect to full-time employment as base category. Precise definition of labour force states for all ages requires more information and thus the multinomial logit has more missings, i.e. fewer observations.

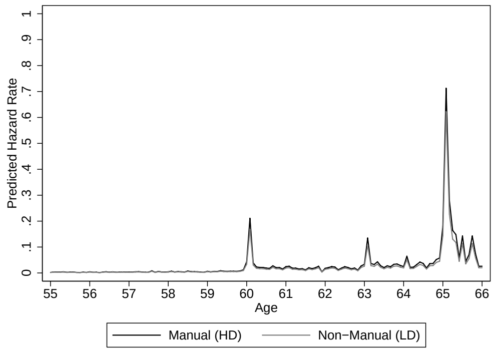
model allows for transitions into all different states, i.e. full-time employment, part-time employment, unemployment and retirement. Duration time enters both models in its most flexible form, where 132 dummies capture variation for each month (i.e. age) and thus no functional form assumption on the baseline hazard is imposed. The variable "Manual" discriminates between occupations with high physical demands (Manual = 1) and low physical demands (Manual = 0). The estimated coefficient on "Manual" is positive for the binary proportional hazards model, which indicates that manual workers have a larger hazard to enter retirement on average. In the multinomial logit model, the coefficient for "Manual" is as well positive and of similar magnitude for transitions into retirement (i.e. the respective outcome). This result is important to underline what we have found in the binary case, because the multinomial model allows for transitions into other states and is thus more general.

Figure 4.3 reports hazard profiles for exits into retirement as predicted from the complementary log-log model. Hazard profiles in figure 4.3 show that retirement entries accumulate around age 60, 63 and 65, as suggested by the respective peaks in predicted hazard rates. This pattern is a perfect projection of the German social security legislation, where eligibility for specific types of old-age pensions are achieved at these specific ages.[6] Similar patterns for Germany with spikes at age 60, 63 and 65 have been recog-

[6]Note that eligibility refers to both early and normal retirement ages. The normal retirement age for a regular old-age pension was 65 until December 2011, where our observation period ends. Besides regular old-age pensions, other types are available such as pensions for unemployed, women, long-

nised in previous studies, such as Börsch-Supan and Schnabel (1998); Börsch-Supan (2000a). Moreover, the hazard profile for non-manual workers (average hazard to enter retirement is 2.5%) is systematically lower compared to manual workers (average hazard to enter retirement is 3.2%) for the observed ages. Thus, the hazard profile of non-manual workers is about 20% lower compared to individuals with physically demanding occupations.

Figure 4.3: Predicted Hazard Profile for Retirement Entries by Physical Demands.



Source: Own calculations based on SOEP (1984-2012). *Note:* High and low (physical) demands are abbreviated by HD and LD respectively.

4.4.2 Retirement and Individual Background: Evidence from Non-Parametric Tests

A simple test of equality of opportunity is to check whether the retirement age distributions differ between individuals with different circumstances. If so, this points at differences in the timing of retirement which are beyond individual responsibility. Assume two retirement age distributions A and B and their cumulative distribution functions (CDF) $F_A(r)$ and $F_B(r)$. Then, A dominates B at first order if and only if $F_A(r) \leq F_B(r)$ for any retirement age $r_j = \{r_1, r_2, ..., r_k\}$. We apply the first order stochastic dominance concept using three different categories of circumstance variables to divide our sample, namely, personal characteristics, socio-economic background, and urbanisation of area of residence during childhood. We test for equality of distributions conducting Kolmogorov-Smirnov tests of equality of distributions.

term insured and severely disabled individuals. These types of old-age pensions are typically available "early", i.e. at age 60 or 63.

Dominance According to Socio-Economic Background

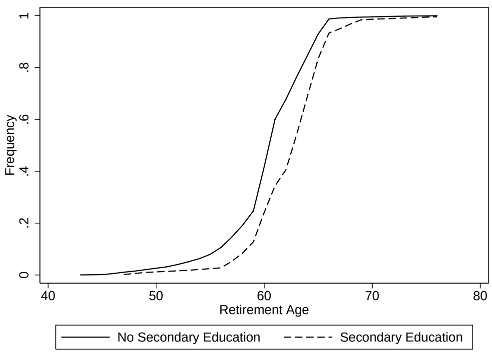
We use parental educational attainment as a proxy for socio-economic background. Figure 4.4 plots the CDFs of retirement age for individuals whose fathers achieved a secondary education degree (i.e. German “Abitur” or “Fachoberstufe”) and individuals whose fathers did not achieve such a degree (panel (a)). With an exception for the tails, the CDF of individuals born to highly educated fathers lies below the CDF of individuals born to less educated fathers across the complete distribution of retirement age. The largest difference between highly and less educated fathers is 0.286, the largest difference between less and highly educated fathers is zero. The distributions are tested to be significantly different from each other. Hence, the CDF of individuals born to highly educated fathers first order dominates the CDF of individuals born to less educated fathers.

Figure 4.4 illustrates the analogous CDFs with respect to maternal educational attainment (panel (b)). The overall picture is similar to the one based on paternal education. Again, the CDFs are significantly different with the CDF of individuals born to mothers with a secondary degree lying below the CDF of individuals born to mothers without such a degree. Both the findings based on paternal and maternal education indicate that equality of opportunity is violated.

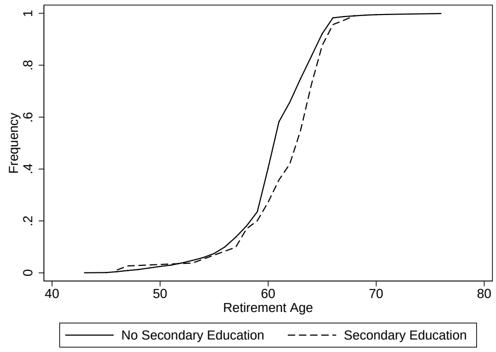
Dominance According to Personal Characteristics

We consider sex and body height as personal characteristics, which are attributes that individuals clearly cannot be held responsible for. Figure 4.4 illustrates CDFs of age at retirement by specific circumstances, where sex is displayed in panel (c). While the overall shapes of the distributions appear to be quite similar, the lines diverge between ages 60 and 65, implying that females enter retirement earlier than males. Table 4.2 reports all results of according stochastic dominance tests for equality of distribution by specific circumstances. The largest difference between the distribution functions is 0.215. The maximum difference between females and males is negative (-0.004), which indicates that the CDF of males lies below the CDF of females at any possible retirement age. The small p-value for the combined test indicates that the distributions of males and females are significantly different. Hence, the CDF of males first order

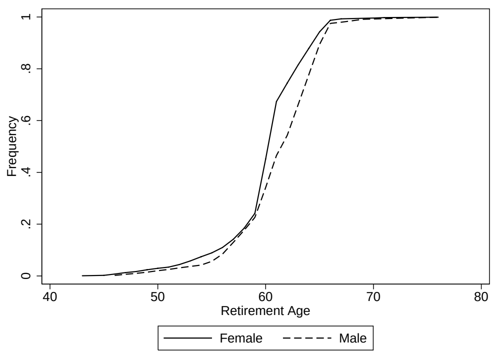
Figure 4.4: Cumulative Distribution Function of Retirement Age by Circumstances.



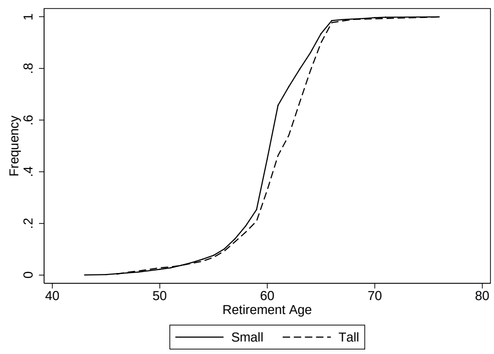
(a) Father's Schooling



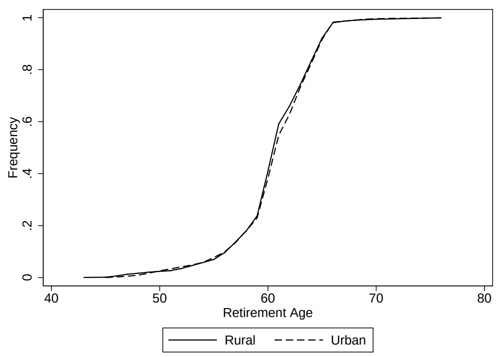
(b) Mother's Schooling



(c) Sex



(d) Body Height



(e) Urbanisation

Source: Own calculations based on SOEP (1984-2012). *Note:* Graphs are smoothed by averaging cumulative distributions on the level of age in years.

dominates the CDF of females. However, differences in retirement age by sex cannot generally be considered as illegitimate inequalities since males and females are subject to different retirement rules and social norms and are therefore of limited comparability. In contrast, retirement rules are independent of body height. Height is a proxy for childhood health as well as height-related differences in self-perception and the percep-

Table 4.2: Tests for Stochastic Dominance at First Order: Differences in Distribution of Retirement Age by Circumstances.

Test	Maximum Difference	p-value	Corrected p-value
By Father's Schooling			
No Secondary Education	0.286	0.000	
Secondary Education	0.000	1.000	
Combined	0.286	0.000	0.000
By Mother's Schooling			
No Secondary Education	0.273	0.000	
Secondary Education	-0.024	0.000	
Combined	0.273	0.000	0.000
By Sex			
Female	0.215	0.000	
Male	-0.004	0.185	
Combined	0.215	0.000	0.000
By Body Height			
Small	0.209	0.000	
Tall	-0.009	0.001	
Combined	0.209	0.000	0.000
By Urbanisation			
Rural	0.078	0.000	
Urban	-0.016	0.000	
Combined	0.078	0.000	0.000

Source: Own calculation based on SOEP (1984-2012).

tion by others which may both provoke differences in career chances. In a comparison of CDFs by body height, first order stochastic dominance would clearly imply the presence of inequality of opportunity in the timing of retirement. Figure 4.4 (panel d) illustrates the distribution functions for individuals who differ by body height, where the sample mean of body height has served as a threshold to assign individuals to the two groups. Again, the distributions appear very similar with a divergence between ages 60 and 65; the maximum difference amounts to 0.209. The largest difference between small and tall individuals is negative (-0.009), which implies that the CDF of tall individuals is below the CDF of small individuals at any retirement age. According to the combined test, the CDFs are significantly different from each other. The conditions for first order stochastic dominance are fulfilled, suggesting the presence of inequality of opportunity in favour of tall people.

Dominance According to Urbanisation in Area of Residence During Childhood

In contrast to personal characteristics and socio-economic background, visible differences in retirement age are not as distinctive when individuals who grew up in areas of different degrees of urbanisation are compared. Figure 4.4 suggests that the CDFs about coincide at most retirement ages (panel (e)). There is a slight divergence between ages 60 and 63. The stochastic dominance test reveals that at any retirement age the CDF of individuals who grew up in urban areas lies below the CDF of individuals who grew up in rural areas. The largest difference amounts to 0.078 and the CDFs are estimated to be significantly different. Hence, although the visible differences appear rather small when comparing CDFs by urbanisation, the urban-CDF dominates the rural-CDF at first order, suggesting inequality of opportunity in age at retirement as well for this characteristic.

4.4.3 Decomposition of the Difference in Retirement Age by Physical Demands: Circumstances versus Effort and Luck

Up to this point, we focused on retirement inequalities between different degrees of physical job demands (duration models) and inequality of opportunity in retirement (tests for stochastic dominance) separately. The ultimate aim of this study is, however, a combination of the two, i.e. an evaluation of the proportion of differences in retirement age between manual and non-manual workers that are attributed to circumstances as distinguished from effort and luck. In order to estimate this proportion, we conduct the decomposition method proposed by Blinder (1973) and Oaxaca (1973) based on the following linear model:

$$RA_g = \mathbf{X_g}'\beta_g + \varepsilon_g \,, \tag{4.1}$$

where $g = (N, M)$ denotes the groups of non-manual workers N and manual workers M, respectively, RA denotes individual retirement age expressed in years whereas varying by month, $\mathbf{X}$ contains a constant and a range of circumstance variables, and ε is an error term. The mean difference in retirement age, which is given by

$$\Delta_{NM} = E(RA_N) - E(RA_M)\,, \qquad (4.2)$$
$$= E(\mathbf{X}_N)'\beta_N - E(\mathbf{X}_M)'\beta_M\,,$$

where $E(\beta_g) = \beta_g$ and $E(\varepsilon_g) = 0$ by assumption, can generally be decomposed into

$$\{E(\mathbf{X}_N) - E(\mathbf{X}_M)\}'\beta^* + \{E(\mathbf{X}_N)'\{\beta_N - \beta^*\} + E(\mathbf{X}_M)'\{\beta^* - \beta_M\}\}\,, \qquad (4.3)$$

such that the first summand refers to the "explained" part and the second summand represents the "unexplained" part of the outcome difference between the two groups. The literature proposes several variants of the decomposition in equation (4.3) by determining β^* in different ways. Specifically, β^* can be defined as a weighted average of the group coefficient vectors:

$$\beta^* = \mathbf{\Omega}\beta_N + (\mathbf{I} - \mathbf{\Omega})\beta_M\,, \qquad (4.4)$$

where $\mathbf{\Omega}$ denotes a weighting matrix and $\mathbf{I}$ is an identity matrix. $\mathbf{\Omega} = \mathbf{I}$ and $\mathbf{\Omega} = \mathbf{0}$ represent the special cases proposed by Oaxaca (1973) and Blinder (1973). These two decompositions provide the extreme cases of assigning the complete interaction effect between endowments and coefficients either to differences explained by endowments ("explained" part) or to differences explained by coefficients ("unexplained" part), respectively. Several authors have suggested alternatives leading to decompositions in between. Neumark (1988) suggests to estimate a pooled model over both groups to infer β^*. Cotton (1988) proposes to choose $\mathbf{\Omega} = s\mathbf{I}$, where s denotes the sample fraction of group N.

In Section 4.4, we report estimates of all four described variants of equation (4.3), while our preferred decompositions are those proposed by Neumark (1988) and Cotton (1988) since they provide convincing strategies of achieving a result between the extreme cases. As the circumstances we include in our model are only a subset of all relevant circumstances (we do not observe talent, for example), the estimate of the explained part of equation (4.3) must be interpreted as a lower bound of the contribution of circumstances to the retirement differential between manual and non-manual workers. The unexplained part is to be interpreted as arising from differences in the coefficients as well as differences in unobserved predictors, such as effort and luck (Jann, 2008).

4.4.4 Decomposition Results

Table 4.3 reports the results from the Oaxaca-Blinder decompositions. While on average both considered occupation types retire in their early 60s, it is predicted that manual workers retire 1.1 years earlier than non-manual workers. The decomposition results shown in the lower panel of table 4.3 suggest that endowments explain between 0.25 and 0.46 years of this differential, depending on the choice of $\mathbf{\Omega}$, which is equivalent to a contribution of between 23.2% and 42.3%. When considering the results based on the pooled model over groups, 0.44 years of the differential (39.8%) are explained, while 0.66 years (60.2%) remain unexplained. Finally, the results for the decomposition proposed by Cotton (1988), suggest that 0.37 years of the difference (33.5%) are attributed to differences in endowments.

The findings from the latter two decompositions indicate that circumstances account for at least 33.5% to 39.8% of the differences in retirement age between manual and non-manual workers, while at most 60.2% to 66.4% can be attributed to effort and luck. The estimated contribution of circumstances can be considered as a lower bound since the range of circumstances accounted for is unlikely to be complete (e.g., we do not observe innate talents) and, in addition, circumstance effects are likely to be mediated partly through characteristics not included in the model which are potentially influenced by individual choices, such as educational attainment, occupational choice, or health behaviours, which cannot be distinctively classified as circumstance or effort variables. In sum, the decomposition results suggest that circumstances explain at least one third of the observed differences in the retirement age between workers of different degrees of physical job demands, which indicates that a considerable part of differences in retirement ages are predetermined and thus not subject to individual choice.

Table 4.3: Oaxaca-Blinder Decomposition.

Non-Manual Workers:				
Average retirement age			61.42	
Obs.(Person-Month-Obs.)			499(82185)	
Manual Workers:				
Average retirement age			60.33	
Obs.(Person-Month-Obs.)			539(97363)	
Difference			1.099	
Decomposition:				
	$\mathbf{\Omega = I}$	$\mathbf{\Omega = 0}$	Pooled	Cotton
Explained	0.255	0.465	0.437	0.369
	23.18%	42.30%	39.77%	33.55%
Unexplained	0.844	0.634	0.662	0.730
	76.82%	57.70%	60.23%	66.45%

Source: Own calculation based on SOEP (1984-2012).

Note: Decomposition of the difference in mean retirement age in years between non-manual and manual workers.

4.5 Conclusion

The purpose of this paper is to quantify differences in the retirement age between manual and non-manual workers and to evaluate these differences with respect to EOP. The focus is on the question how individual background during childhood transmits through physical demands of occupations on retirement ages.

Individual retrospective data from the SOEP are used to analyse labour force dynamics over the years 1984 to 2011. Discrete time duration models are estimated in the most flexible version, where age (i.e. duration time) enters the model on a monthly level and thus accounts for variation in the relevant range from age 55 to 65. The estimated hazard profile of non-manual workers is about 20% lower compared to individuals with physically demanding occupations. Non-parametric tests for stochastic dominance at first order indicate that the distribution of retirement age differs significantly between individuals across circumstances. However, the ultimate aim of this study is an evaluation of the proportion of differences in retirement age between manual and non-manual workers that are attributed to circumstances as distinguished from effort and luck. The result from a Blinder-Oaxaca decomposition suggests that circumstances explain at least one third of the observed differences in the retirement age between workers with different degrees of physical job demands. The result is a lower bound, as we do not observe the full set of individual circumstances. This finding is important because it indicates that a considerable part of differences in retirement age is predetermined and thus not subject to individual choice.

Retirement decisions are complex. Aside from general preferences for consumption and leisure, several aspects such as health-related behaviours, wealth and occupational sorting play a role in retirement choices, and most of these factors are not exogenously determined. Beyond individual choice, employer behaviour and the availability of retirement benefits (i.e. social security legislation) influence the observed outcomes for retirement ages. Thus, retirement decisions are influenced by a number of factors which in sum do not clearly indicate whether retiring early or working longer is more desirable from an individual point of view. In the present paper, we apply an approach that carries on the EOP literature and decomposes differences in the retirement age into individual responsibility (i.e. effort and luck) and personal background (i.e. circumstances). Thus, the relevant quantity is the share of individual background characteristics that transmits through physical demands of occupations to retirement ages. By nature, individual circumstances are predetermined to any endogenous decision which individuals could be held responsible for. Early retirement usually implies a reduction in social security wealth and to this end, differences in retirement age between subgroups do have economic relevance. When raising the normal retirement age, policy makers must be aware of potential disadvantages in terms of reduced benefit entitlements for manual workers. While the interpretation of our result is highly normative, it allows to focus thoughts in a debate, where early retirement of manual workers is often considered to be "unfair".

Appendix

4.A Transitions

Table 4.4: Distribution of Individual Transitions across States.

Origin	Transition into State				Total Net Transitions	Total
	(1)	(2)	(3)	(4)		
(1) Full-time Employment	559,944	1,656	2,977	1,323	5,956	565,900
	(98.95)	(0.29)	(0.53)	(0.23)	(1.05)	(100.00)
(2) Part-time Employment	1,595	177,439	727	583	2,905	180,344
	(0.88)	(98.39)	(0.40)	(0.33)	(1.61)	(100.00)
(3) Unemployment	2,294	738	94,057	1,130	4,162	98,219
	(2.34)	(0.75)	(95.76)	(1.15)	(4.24)	(100.00)
(4) Retirement	142	122	116	762,117	380	762,497
	(0.02)	(0.02)	(0.02)	(99.94)	(0.06)	(100.00)
Total					13,403	1,606,960

Source: Own calculation based on SOEP (1984-2012).
Note: Absolute transitions are reported; relative shares in parentheses. "Net transitions" refer to all transitions into respective other states.

4.B Spells

Table 4.5: Number of Spells per Individual.

Number of Spells	Number of Individuals	Per Cent
1	12,132	68.96
2	2,415	13.73
3	1,314	7.47
4	620	3.52
5	463	2.63
6	181	1.03
7	169	0.96
8	88	0.50
9	71	0.40
10 or more	141	0.80
Total	17,594	100.00

Source: Own calculation based on SOEP (1984-2012).

Chapter 5

Redistribution in a Pay-As-You-Go Pension System: The Relationship between Differential Mortality and the Retirement Age

5.1 Introduction

The purpose of this paper is to reveal important patterns of the relationship between the age of retirement and mortality. How does the remaining life expectancy at retirement evolve as a function of the retirement age? We link this question to a discussion of redistribution in a pay-as-you-go pension system (PAYG), because such a system redistributes towards those participants who receive benefits for more periods, ceteris paribus. Throughout this study, we measure the duration of benefit receipt as the remaining life expectancy after retirement in years.

This study is motivated from two main aspects. First, a recent German offers an exit from the labour force into early retirement to long-term-insured individuals at age 63 without benefit reductions from actuarial adjustments.[1] The argument in favour of this

[1]Individuals are considered to be "long-term-insured" if they contribute to the German pay-as-you-

policy change is that extraordinarily long working careers need advantageous treatment. It must be taken into account, however, that long-term-insured individuals usually exhibit a rather stable earnings history with strong labour market attachment and, qua construction of a PAYG, have relatively high benefit entitlements at their disposal. The critical question in this context is whether the average life expectancy of this group is systematically higher, i.e. whether there is subgroup heterogeneity in life expectancy with respect to selection into specific retirement ages; previous research for Germany shows that such selection does exist (Kühntopf and Tivig, 2012). Redistribution in a PAYG strongly depends on the life expectancy of its participants, which is a major determinant for net present values of retirement benefits and thus social security wealth. If the beneficiaries from the aforementioned reform are particularly long-living, then the discussed policy change will redistribute towards individuals who are privileged by the nature of the PAYG.

Second, there is no rationale for the retirement age to have a causal impact on life expectancy. No compelling reason is in favour of an event that condenses to a calendar date to influence the remaining length of life of an individual. Instead, the retirement age is likely to be correlated to factors which on their side are strongly correlated to mortality, such as health-related behaviours, wealth or physical demands of occupations. However, a recent body of literature has evolved on the topic, with a specific focus on the effect of retirement age on mortality (e.g. Hernaes et al., 2013; Bloemen et al., 2013; Coe and Zamarro, 2011; Kuhn et al., 2010; Anderson, 1985). Usually these studies circumvent endogenous retirement by instrumenting retirement decisions using exogenous changes for example in early retirement rules.

While methodologically interesting, these studies are in search of an effect that should not exist in the first place and hence it is not surprising that this strand of the literature is rather inconclusive. Using very reliable data and a convincing identification strategy, Hernaes et al. (2013) find that there exists no causal effect of retirement age on mortality and they finally conclude that mortality should not take a prominent place in discussions about the retirement age. Although their finding of "no effect" supports our logic from above, we provide evidence that suggests that mortality is indeed important in discussions about the retirement age. This stems from the fact that retirement ages

go pension system for at least 45 years.

are correlated to factors that are closely linked to mortality and therefore should be considered in policies that aim to redistribute within a PAYG.

The point of departure in this paper is an examination of the relationship between *overall* life expectancy and the retirement age, which are likely to exhibit a positive relationship. This involves the idea of benefit optimisation, where individuals who perceive to live longer tend to retire later (see e.g. Waldron, 2001).[2] Second and of most importance for this paper, we examine the relationship between *remaining* life expectancy and the retirement age. We expect these two quantities to exhibit a negative relationship because retiring at higher ages implies fewer remaining years to live compared to the counterfactual of earlier retirement. We shed light on this latter aspect and elaborate the potential for non-monotonicity in the average remaining years to live as a function of the retirement age.

Our results show that non-monotonicity does exist among male workers. The remaining life expectancy as an otherwise decreasing function of the retirement age is increasing towards age 63. This is of particular importance for the design of social security systems for two reasons. First, it clearly suggests that subgroup heterogeneity is present for individuals who select themselves into specific retirement ages. This heterogeneity refers to variables that are correlated to the retirement age and to mortality such as health-related behaviours and wealth. These are held to be responsible for the remaining length of life, supporting the view that the age of retirement itself has no impact on mortality whatsoever (Hernaes et al., 2013). Second, our result has important implications for redistribution in a PAYG. Offering early retirement at age 63 without imposing financial incentives (i.e. benefit reductions) to long-term-insured individuals is an example of redistribution towards long-term-insured individuals within the PAYG. This is problematic in the light of our result, because the beneficiaries of this policy change belong to a group whose average life expectancy after retirement is relatively large.

The remainder of this paper is structured as follows. Section 5.2 provides a brief overview on the institutional setting and describes the data source. Section 5.3 outlines the empirical analysis and presents the corresponding results. Section 5.4 concludes.

[2]For a discussion of subjective survival probabilities and their relationship to individual choice, see Hamermesh (1985); Hurd and McGarry (1995, 2002).

5.2 Institutional Setting and Data

The German public pension system is organised as a PAYG, where monthly contributions from insured employees are directly discharged to benefit recipients. Old age pensions are available for different types of eligibility, as summarised in table 5.1.

Table 5.1: Types of Eligibility for Old Age Pensions.

Eligibility Type	Retirement Age		Current Reform Status
	ERA	NRA	
Unemployed	63	65.33	not available for BC 1952+
Women	60	65.33	not available for BC 1952+
Longterm Insured (35 years)	63	65.33	
Longterm Insured (45 years)	65	65.33	ERA reduced to age 63 (2014)
Severely Disabled	60	63	

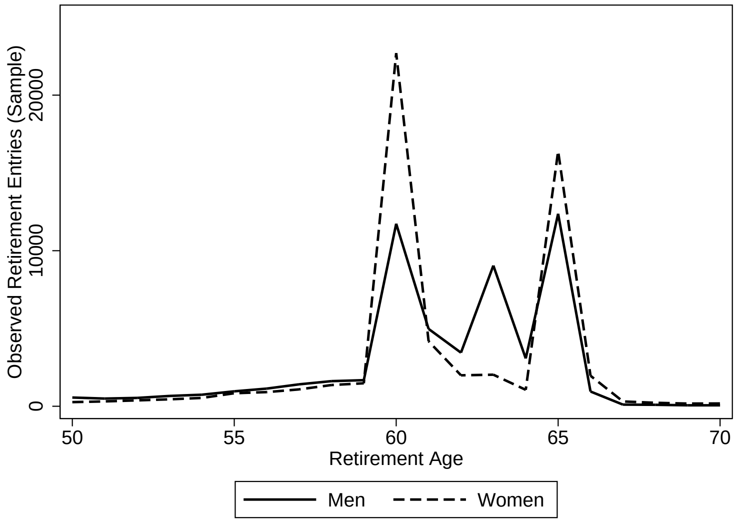
Source: Own illustration based on social security legislation (SGB VI). *Note:* ERA abbreviates early retirement age, NRA abbreviates normal retirement age and BC is for birth cohort. All types of old age pensions are subject to actuarial adjustments if applied before the NRA. All retirement ages are reported for the social security legislation as of the year 2015.

As is evident from table 5.1 and figure 5.1, early and normal retirement ages induce retirement patterns with mass points at age 60, 63 and 65. This pattern reflects an institutional setting that has been shaped from social security legislation as enacted in 1972, but prevails to be relevant in its essence until today (see Börsch-Supan, 2000a, for details). To ensure that individuals are homogeneous by their retirement rules (i.e. the 1972 legislation), we restrict our sample to birth cohorts 1912 - 1943. According to figure 5.1, the majority of female individuals retires at age 60 while for men the retirement ages 60 and 65 are equally prevalent. Moreover, a relevant share of men retires at age 63 which is not the case for women. This aspect is important because our main finding suggests that the group of men who retire at age 63 are those who are particularly long-living and thus receive pension benefits for more periods that other retirees.

Recent German reforms such as the introduction of actuarial adjustments (1997 - 2004) or the current raise of the normal retirement age from 65 to 67 (2012 - 2029) are likely to induce an upward shift of the real retirement age and thus introduce some variation in retirement patterns. However, a current reform intends to reduce the early retirement age for longterm-insured individuals with at least 45 contribution years from 65 to 63.[3]

[3]The reform was implemented in mid-2014 and has been subject to a controversial public debate in

Figure 5.1: Number of Observed Retirement Entries across Age: Birth Cohorts 1912 - 1943.



Source: Own calculations based on SUFRTWF(2006-2008). *Note:* Plotted lines restricted to old age pensions and reduced earnings capacity pensions.

The most remarkable detail with respect to this policy change is that pension benefits of eligible early retirees will not be subject to benefit reductions.[4] Thus, it is likely that the majority of individuals who fulfill eligibility will exercise this option because it is attractive. This is an exception from other types of old age pensions and sets strong financial incentives for long-term insured individuals to retire earlier. Consequently, the relevant group for this study is the group of long-term insured individuals with 45 contribution years (see table 5.1).

5.2.1 Data

The analysis is based on a unique source of micro-data from social security records, as provided by the research data centre of the German federal pension insurance (FDZ-RV). Specifically, we draw on scientific use files for the years 2006 to 2008 which contain a 10% sample of all shortfalls of pension benefit recipients in each respective year (SUFRTWF2006-SUFRTWF2008). Due to the high reliability of information, first studies found this data source to be valuable for mortality research (see Scholz, 2005,

Germany.

[4]One year of early retirement with reference to the normal retirement age is generally subject to a 3.6% reduction for all types of old age pensions.

2006). Moreover, the observed shortfall of benefit receipt is almost perfectly correlated to unobserved individual death and thus we have very reliable information on the remaining length of life after retirement entry.[5]

Our final sample consists of 56,507 men and 60,484 women. The final sample is shaped by specific restrictions with respect to birth cohorts and calendar years. First, we do not consider birth cohorts previous to 1912 in order to produce a homogeneous sample regarding the 1972 legislation (see section 5.2). Second, we make use of the most recent reporting years 2006 - 2008 in order to reduce right censoring in pension shortfall of relevant birth cohorts. The three-year bracket is used to smooth results and to reduce standard errors. The youngest birth cohort in the subsequent empirical framework is 1943 because we restrict pension shortfall to ages equal or above 65 (to rule out severe health cases). Thus, we are finally left birth cohorts 1912 - 1943.

5.2.2 Socio-demographic Variables

Socio-demographic information is important for mortality research as noted by Kreyenfeld and Scholz (2010). We have access to few such variables that are available in the data and are mainly used in subsequent regressions.

First, we are able to distinguish individuals by their (public) pension claims, i.e. so-called earnings points (EP). EP measure the relative individual income position for each contribution year.[6] We are able to draw on the individual sum of EP, which is a measure for lifetime contributions thus equivalent to total pension claims within the PAYG.[7]

Second, the data contain information on periods of illness that are creditable in the German public pension system. Such creditable periods are due to medical rehabilitation or severe health-issues that may prevent individuals from regular contributions to the public pension system (by working). We use these periods of illness as a proxy for

[5]Note that we can identify the last documented retirement spell previous to shortfall in the data. In this case it is very unlikely that shortfall is due to other reasons than death, such as returns to work in old age, for example.

[6]One EP is assigned to an individual who is located exactly at the mean of the earnings distribution for all of employees who are subject to social security contributions within a given year. Depending on regions, one EP is worth between 26 and 28 Euros as of 2014.

[7]While the data contain contribution time (in years) as well, unfortunately this measure is subject to a large number of missing values. Thus, we cannot construct an income measure that relates lifetime contributions to the total years of contribution and could be used as a proxy for annual income previous to retirement.

health status.

Third, just as for illness we can draw on periods of unemployment that are creditable in terms of pension claims as well. We use periods of unemployment to account for labour market histories and unobserved factors that may be correlated.

Finally, we can draw on various further conditioning variables such as birth cohort, marital status, region, and number of children.

The subsequent analysis distinguishes between men and women for two reasons. First, the two sexes differ fundamentally in mortality with a life expectancy that is much larger for women. Second, legislative rules that apply to the availability of old age pensions differ between men and women (see table 5.1 for old age pensions for women).

5.3 The Relationship between Differential Mortality and the Retirement Age

The average length of life as a function of the retirement age is plotted in figure 5.2. The corresponding lines are monotonically increasing, indicating a positive relationship between live expectancy and the retirement age. Clearly, individuals with a higher life expectancy retire later. Drawing on private information concerning health status and health-related behaviours, individuals who perceive to live longer tend to retire later (Waldron, 2001).

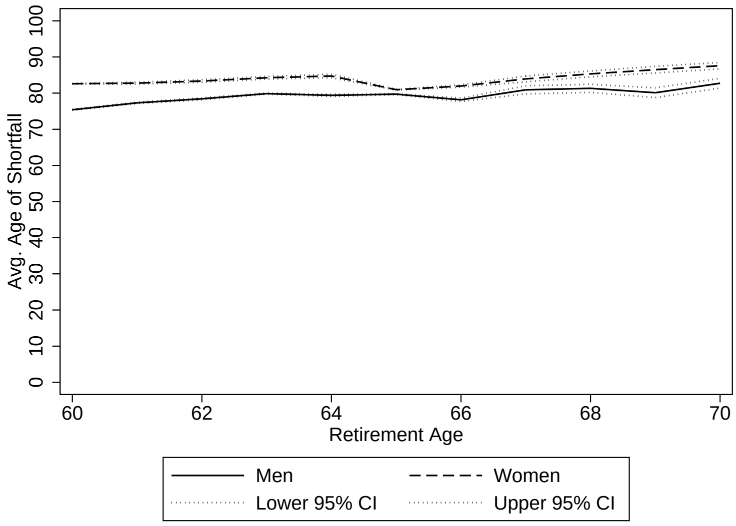
The figure also resembles the common phenomenon that women systematically live longer compared to men. For the year 2008, the absolute difference in life expectancy of newborn individuals in Germany amounts to roughly five years (77.2 years for men and 82.4 years for women, Federal Statistical Office (2012b)).

5.3.1 Remaining Life Expectancy and the Retirement Age

In contrast to the previous analysis, figure 5.3 shows how the average *remaining* years to live evolve as a function of the retirement age for old age pensions.[8] Retirement into old age pensions predominantly takes place between age 60 and 65 in Germany, such that the number of observations at this age range is large and 95% confidence bands are narrow.

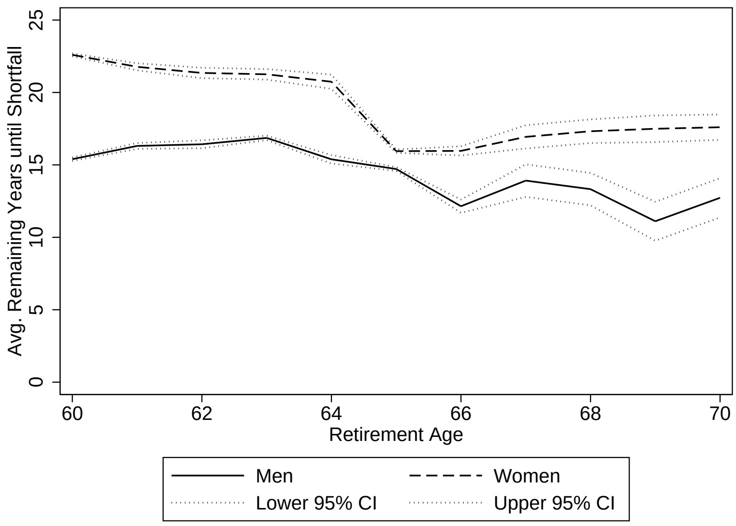
[8]See table 5.2 in the appendix for all underlying calculations that are plotted in figure 5.3.

Figure 5.2: Life Expectancy as a Function of the Retirement Age.



Source: Own calculations based on SUFRTWF(2006-2008). *Note:* The sample is restricted to individuals whose benefit shortfall is not before age 65. Plotted lines restricted to old age pensions and reduced earnings capacity pensions.

Figure 5.3: Remaining Life Expectancy as a Function of the Retirement Age.



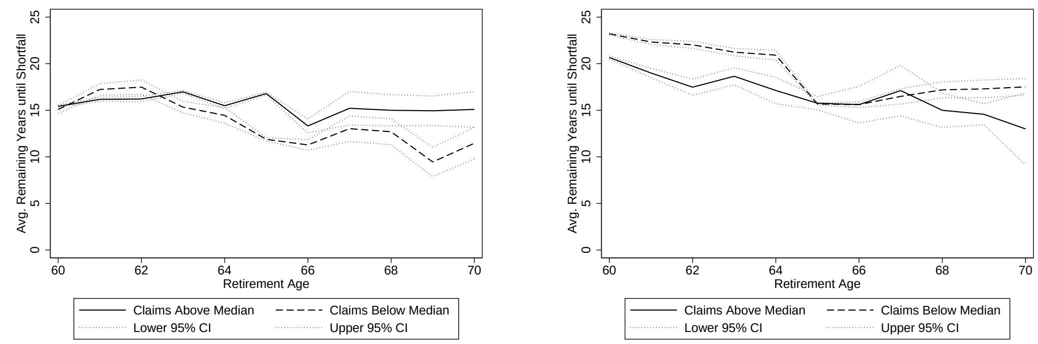
Source: Own calculations based on SUFRTWF(2006-2008). *Note:* The sample is restricted to individuals whose benefit shortfall is not before age 65. Plotted lines restricted to old age pensions and reduced earnings capacity pensions.

Remaining life expectancy as a function of the retirement age in figure 5.3 is decreasing over some interval between age 60 - 70 for male individuals but *not* monotonically. Towards age 63 the slope is positive for men, which means that we observe a group of individuals with higher remaining life expectancy at retirement compared to others who retire at lower ages. Male individuals retiring at age 63 systematically have more remaining years to live compared to those who retire between age 60 - 62. This subgroup heterogeneity indicates substantial redistribution within the PAYG system depending on the length of benefit receipt.

For women, remaining life expectancy is slightly decreasing until age 64 and drops substantially between age 64 and 65. Consistent to figure 5.2, women systematically live longer compared to men.

To avoid an underestimation of the remaining benefit receipt especially for those who retire between age 60 and 65, we restrict the sample to individuals who live at least until age 65 throughout the subsequent analysis. That is, we restrict the sample to individuals, where the shortfall of benefits is not documented before age 65.[9] Restricting the sample to cases of death for ages above 64 rules out cases of severe illness and is consistent with previous mortality research (Kühntopf and Tivig, 2012).

Figure 5.4: Remaining Life Expectancy as a Function of the Retirement Age: Differences by Total Pension Claims.



(a) Men (b) Women

Source: Own calculations based on SUFRTWF(2006-2008). *Note:* The sample is restricted to individuals whose benefit shortfall is not before age 65. Plotted lines restricted to old age pensions and reduced earnings capacity pensions. Individuals are distinguished by the median of total pension claims, i.e. the sum of earnings points.

Discriminating subgroups by total pension claims in figure 5.4 reveals that the relationship between retirement age and mortality depends on total pension claims. That is, benefit entitlements within the PAYG introduce systematic differences in remaining

[9]The results without this restriction are very similar and available from the authors upon request.

years to live after retirement. For male individuals, figure 5.4 (panel a) suggests that the remaining years to live are positively correlated to total pension claims. This means that men with benefit entitlements above the median generally tend to live longer (with an exception for retirement ages 61 - 62).[10] These results are in line with studies that document differential mortality with respect to economic welfare such as Attanasio and Hoynes (2000) or Reil-Held (2000) for Germany.

For women, the results in figure 5.4 (panel b) are remarkable. First, between age 60 - 64 women with low pension claims have more remaining years to live compared to those above the median. Second, the pattern from figure 5.3 with a sharp drop in life expectancy at age 64 seems to be strongly driven by women who have low pension claims. A potential explanation of this result is that women who retire between age 60 - 64 are in majority from traditional household constellations with low labour force attachment. This may well explain why women have low pension claims (i.e. low contributions) and at the same time systematically live longer (i.e. low physical/mental demands from work).

Generally, it must be noted that our measure total pension claims diverges from typical income measures and comparisons need to take this into account. The sum of benefit entitlements is problematic because it does not measure productivity. A given sum of pension claims may either be accumulated over a long period (low wage) or a short period (high contribution, i.e. high wage) and we are not able to distinguish the two.[11]

Regression Analysis

To examine the relationship between differential mortality and the retirement age in more detail, we make use of additional variables (see section 5.2). We regress remaining years to live (dependent variable) on the retirement age, conditional on a set of observable characteristics. This proceeding allows to capture a fair amount of variation and allows to obtain a clear picture of the relationship between differential mortality and the retirement age.

The regressions reveal important patterns (table 5.3, appendix). The first two columns report coefficients for the baseline regressions separately for men and women. In the

[10] These results are robust to changes in the rule of comparison, e.g. comparing top and bottom quintiles of the EP-distribution.

[11] Due to data limitations, we do not have access to a measure of contribution time.

baseline model, we regress the remaining years to live after retirement on a set of dummy variables that indicate retirement age. The purpose of this regression is to replicate the previous pattern; the results are very much similar to the previous finding where the remaining years to live are not decreasing linearly in the retirement age. Relative to the reference category (age 60), we observe a spike at age 63 for male individuals (column 1) as in the graphical analysis (figure 5.3).

The pattern changes fundamentally, however, once we include a set of control variables (columns 3 and 4, table 5.3, appendix). The remaining life expectancy decreases monotonically in annual steps. For example, at age 70 the remaining life expectancy is roughly 13 years lower compared to individuals who retire at age 60 (the reference). This has important implications. First, it shows that the previously observed pattern is due to underlying causes that are correlated to specific retirement ages. Second, while it is not possible to establish causality without an appropriate experiment, it seems that the predominant share of explanatory power is from birth cohort heterogeneity. The estimated coefficients from creditable periods of health and unemployment, total pension claims, region and household variables such as number of children are all small in magnitude. Thus, birth cohort heterogeneity seems to be important in explaining considerable differences in mortality.

5.4 Conclusion

Based on a unique data source from social security records we analyse the shortfall of pension benefit recipients for the years 2006 to 2008. The main result is clear and simple: At age 63, men receive old age pensions for more periods on average compared to those who retire at lower ages. We find non-monotonicity in average life expectancy after retirement as a function of the retirement age, where a monotonically decreasing function is expected. For male individuals, the average remaining years to live increase in retirement age towards age 63. In contrast, the remaining life expectancy after retirement slightly decreases between age 60 - 64 and drops sharply by about five years thereafter (i.e. moving from retirement age 64 to 65).

It is difficult to assess *why* individuals do retire at specific ages. The observed pattern is likely to be driven by the institutional setting of the German PAYG, where large-scale retirement takes place at age 60, 63 and 65. These subgroups may be heterogeneous by selection into specific retirement ages. That is, specific types retire at certain ages and these groups differ in remaining life expectancy by relevant aspects such as health-related behaviours and wealth. Our results indicate that birth cohorts do explain a considerable amount of this selection process. This birth cohort heterogeneity hints at the importance of existing retirement rules and that incentives from social security are important.

Alternative explanations may be linked to our result. For example, the question whether retirement coincides with subsequent phenomena that either support or prevent an individual to live longer is ambiguous. Individuals may live longer if they are relieved from work-related stress but on the other hand, cognitive decline may set in after retirement (see e.g. Rohwedder and Willis, 2010; Bonsang et al., 2012). Our result shows, however, that the "type" of individual to retire at a specific age differs substantially by characteristics that are strongly correlated to mortality.

Our main finding is that male workers who retire at age 63 receive pension benefits for more periods compared to other men who retire at lower ages. The explanation is that they live longer after retirement on average. This result is remarkable from the perspective of a PAYG pension system. Redistribution is critical if those who receive retirement benefits for a longer period are beneficiaries of a more generous benefit scheme. This has much relevance for the German example, where a new policy (as of mid-2014) offers an old age pension at age 63 without benefit reductions to the group of long-term insured persons. This policy change is not only targeted at a rather privileged group but also ignores the nature of a PAYG, where long time contributors accumulate higher benefit entitlements.

Appendix

5.A Descriptive Calculations

Table 5.2: Remaining Life Expectancy as a Function of the Retirement Age.

	Men				Women			
Retirement Age	Years until Shortfall Mean (s.e.)	Lower 95% CI	Upper 95% CI	N Obs.	Years until Shortfall Mean (s.e.)	Lower 95% CI	Upper 95% CI	N Obs.
60	15.40 (.062)	15.27	15.52	11492	22.60 (.049)	22.50	22.69	22552
61	16.31 (.104)	16.11	16.52	4884	21.77 (.125)	21.53	22.01	4143
62	16.43 (.139)	16.15	16.70	3323	21.35 (.184)	20.99	21.71	1951
63	16.87 (.080)	16.71	17.02	8917	21.26 (.186)	20.89	21.62	1992
64	15.39 (.148)	15.10	15.68	3030	20.74 (.252)	20.25	21.24	1047
65	14.72 (.067)	14.59	14.85	12374	15.95 (.057)	15.84	16.06	16404
66	12.15 (.233)	11.69	12.60	950	15.97 (.162)	15.65	16.28	1963
67	13.92 (.573)	12.79	15.04	108	16.94 (.410)	16.13	17.74	312
68	13.33 (.571)	12.21	14.45	95	17.32 (.417)	16.51	18.14	225
69	11.11 (.686)	9.77	12.46	63	17.49 (.468)	16.58	18.41	174
70	12.73 (.691)	11.38	14.08	63	17.60 (.447)	16.73	18.48	181

Source: Own calculations based on SUFRTWF(2006-2008). *Note:* Calculations for birth cohorts 1912 - 1943. Standard errors in parentheses. The sample is restricted to individuals whose benefit shortfall is not before age 65. All calculations restricted to old age pensions and reduced earnings capacity pensions.

5.B Regression Results

Table 5.3: Remaining Life Expectancy and the Retirement Age.

	Baseline		Further Controls	
	Men	Women	Men	Women
Retirement Age 61	-2.247	-1.889	-3.915	-2.862
	(0.126)	(0.128)	(0.055)	(0.053)
Retirement Age 62	-2.135	-2.294	-4.923	-3.865
	(0.149)	(0.181)	(0.065)	(0.074)
Retirement Age 63	-1.700	-2.397	-5.779	-4.926
	(0.100)	(0.179)	(0.045)	(0.074)
Retirement Age 64	-3.171	-2.921	-6.840	-5.934
	(0.155)	(0.243)	(0.068)	(0.100)
Retirement Age 65	-3.849	-7.712	-8.200	-7.302
	(0.090)	(0.075)	(0.042)	(0.037)
Retirement Age 66	-6.420	-7.696	-9.429	-8.242
	(0.265)	(0.180)	(0.116)	(0.076)
Retirement Age 67	-4.652	-6.727	-10.395	-9.130
	(0.772)	(0.440)	(0.335)	(0.180)
Retirement Age 68	-5.242	-6.339	-11.422	-10.057
	(0.822)	(0.518)	(0.357)	(0.211)
Retirement Age 69	-7.457	-6.169	-12.621	-10.891
	(1.009)	(0.588)	(0.437)	(0.240)
Retirement Age 70	-5.838	-6.061	-13.586	-11.951
	(1.009)	(0.577)	(0.438)	(0.235)
Retirement Age 71	-6.058	-7.323	-14.689	-12.968
	(1.144)	(0.646)	(0.496)	(0.263)
Retirement Age 72	-8.052	-7.213	-15.483	-13.882
	(1.438)	(0.708)	(0.622)	(0.288)
Retirement Age 73	-6.915	-12.123	-16.595	-14.979
	(1.570)	(0.901)	(0.680)	(0.365)
Retirement Age 74	-6.750	-10.680	-17.733	-15.982
	(1.706)	(1.017)	(0.739)	(0.412)
Retirement Age 75	-9.756	-10.642	-18.785	-16.582

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Table 5.3 – *Continued from previous page*

	Baseline		Further Controls	
	Men	Women	Men	Women
	(2.001)	(1.130)	(0.866)	(0.457)
Retirement Age 76	-7.807	-11.642	-19.483	-17.999
	(1.746)	(1.118)	(0.756)	(0.453)
Retirement Age 77	-8.069	-12.277	-20.423	-18.873
	(2.001)	(1.167)	(0.866)	(0.473)
Retirement Age 78	-8.881	-13.444	-21.729	-19.801
	(2.001)	(1.369)	(0.866)	(0.554)
Retirement Age 79	-11.369	-13.986	-22.813	-20.778
	(3.578)	(1.390)	(1.548)	(0.563)
Retirement Age 80	-9.569	-12.963	-23.616	-21.753
	(2.829)	(1.731)	(1.224)	(0.701)
Retirement Age 81	-12.444	-14.898	-24.799	-23.176
	(2.829)	(1.877)	(1.224)	(0.760)
Retirement Age 82	-11.319	-15.189	-25.814	-23.588
	(4.000)	(1.776)	(1.731)	(0.719)
Retirement Age 83	-14.569	-16.413	-28.067	-24.623
	(8.000)	(2.234)	(3.460)	(0.904)
Retirement Age 84	-13.569	-16.413	-27.563	-25.470
	(5.657)	(2.234)	(2.447)	(0.904)
Retirement Age 85	-11.569	-19.163	-28.938	-26.897
	(8.000)	(5.472)	(3.461)	(2.214)
Retirement Age 86	-13.069	-17.463	-29.706	-26.838
	(5.657)	(3.461)	(2.447)	(1.400)
Retirement Age 87	-11.569	-18.863	-30.849	-29.237
	(8.000)	(3.461)	(3.461)	(1.400)
Retirement Age 88	-12.569	-17.996	-33.195	-29.052
	(8.000)	(4.468)	(3.462)	(1.807)
Retirement Age 89	-13.569	-18.863	-32.855	-30.715
	(5.657)	(3.461)	(2.448)	(1.400)
Cohort 1913	–	–	30.319	28.524
			(0.253)	(0.248)
Cohort 1914	–	–	28.969	27.516

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Table 5.3 – *Continued from previous page*

	Baseline		Further Controls	
	Men	Women	Men	Women
			(0.243)	(0.245)
Cohort 1915	–	–	27.885	26.534
			(0.233)	(0.244)
Cohort 1916	–	–	26.976	25.567
			(0.233)	(0.246)
Cohort 1917	–	–	26.207	24.552
			(0.238)	(0.247)
Cohort 1918	–	–	25.055	23.578
			(0.240)	(0.248)
Cohort 1919	–	–	23.941	22.564
			(0.232)	(0.247)
Cohort 1920	–	–	22.970	21.669
			(0.218)	(0.242)
Cohort 1921	–	–	21.877	20.765
			(0.211)	(0.240)
Cohort 1922	–	–	21.135	19.827
			(0.210)	(0.239)
Cohort 1923	–	–	19.945	18.835
			(0.211)	(0.239)
Cohort 1924	–	–	19.128	17.908
			(0.211)	(0.240)
Cohort 1925	–	–	18.043	17.063
			(0.211)	(0.240)
Cohort 1926	–	–	17.056	16.136
			(0.209)	(0.240)
Cohort 1927	–	–	16.041	15.177
			(0.209)	(0.241)
Cohort 1928	–	–	14.853	14.114
			(0.208)	(0.242)
Cohort 1929	–	–	13.838	13.265
			(0.207)	(0.243)
Cohort 1930	–	–	12.897	12.344

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Table 5.3 – *Continued from previous page*

	Baseline		Further Controls	
	Men	Women	Men	Women
			(0.207)	(0.243)
Cohort 1931	–	–	11.919	11.352
			(0.207)	(0.244)
Cohort 1932	–	–	10.827	10.535
			(0.208)	(0.245)
Cohort 1933	–	–	9.771	9.339
			(0.209)	(0.248)
Cohort 1934	–	–	8.921	8.276
			(0.210)	(0.249)
Cohort 1935	–	–	7.789	7.688
			(0.207)	(0.247)
Cohort 1936	–	–	6.594	6.508
			(0.207)	(0.248)
Cohort 1937	–	–	5.539	5.697
			(0.208)	(0.249)
Cohort 1938	–	–	4.767	4.661
			(0.209)	(0.250)
Cohort 1939	–	–	3.936	3.833
			(0.209)	(0.250)
Cohort 1940	–	–	3.065	3.004
			(0.208)	(0.251)
Cohort 1941	–	–	2.143	2.054
			(0.209)	(0.253)
Cohort 1942	–	–	1.351	1.125
			(0.212)	(0.257)
Cohort 1943	–	–	1.010	0.591
			(0.228)	(0.281)
Married	–	–	-0.034	-0.031
			(0.034)	(0.033)
German	–	–	0.423	0.254
			(0.077)	(0.114)
Region: Schleswig-Holstein	–	–	-0.032	0.558

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Table 5.3 – *Continued from previous page*

	Baseline		Further Controls	
	Men	Women	Men	Women
			(0.114)	(0.095)
Region: Hamburg	–	–	0.058	0.526
			(0.131)	(0.110)
Region: Niedersachsen	–	–	-0.119	0.571
			(0.095)	(0.079)
Region: Bremen	–	–	-0.135	0.621
			(0.177)	(0.151)
Region: Nordrhein-Westfalen	–	–	0.281	0.583
			(0.088)	(0.073)
Region: Hessen	–	–	-0.092	0.428
			(0.099)	(0.083)
Region: Rheinland-Pfalz	–	–	0.028	0.680
			(0.105)	(0.091)
Region: Baden-Wuerttemberg	–	–	-0.056	0.659
			(0.093)	(0.077)
Region: Bayern	–	–	-0.109	0.594
			(0.092)	(0.075)
Region: Saarland	–	–	0.533	0.345
			(0.147)	(0.137)
Region: Berlin	–	–	0.007	0.419
			(0.112)	(0.091)
Region: Brandenburg	–	–	-0.108	-0.175
			(0.114)	(0.093)
Region: Mecklenburg-Vorpommern	–	–	0.045	0.025
			(0.127)	(0.106)
Region: Sachsen	–	–	-0.054	-0.455
			(0.100)	(0.081)
Region: Sachsen-Anhalt	–	–	-0.143	-0.173
			(0.112)	(0.091)
Number of Children	–	–	-0.120	-0.064
			(0.034)	(0.009)
Accumulated Pension Claims (EP)	–	–	-0.031	-0.027

Continued on next page

Table 5.3 – *Continued from previous page*

	Baseline		Further Controls	
	Men	Women	Men	Women
			(0.001)	(0.001)
Periods of Illness (Months)	–	–	0.118	0.179
			(0.004)	(0.005)
Periods of Unemployment (Months)	–	–	-0.015	-0.018
			(0.001)	(0.002)
Constant	18.569	23.663	8.877	7.748
	(0.053)	(0.044)	(0.227)	(0.268)
N	56507	60484	56507	60484
R^2	0.044	0.172	0.822	0.865

Note: Own calculations based on SUFRTWF(2006-2008). Reported values are coefficients from OLS regressions. Robust standard errors in parentheses. Retirement age in years and birth cohorts are modeled using binary indicators. Reference categories are age 60 and birth cohort 1912. All models estimated separately for men (column 1 and 3) and women (column 2 and 4) respectively.

Chapter 6

Concluding Remarks

This dissertation examines various aspects of retirement behaviour. The recurrent finding is that social security strongly influences retirement patterns. This result is central to all studies in this thesis and supports the general view that people respond to incentives.

Raising the distance to retirement in terms of the early retirement age stimulates reemployment among recently unemployed older workers (chapter 2). Individuals who are no longer eligible for an old age pension due to unemployment between age 60 - 63 are shown to be more likely to return to work. This reduces public costs from nonemployment and indicates that raising the early retirement age is a reasonable way to increase labour force participation of older workers and to induce a delay in retirement. Thus, the public pension insurance and the unemployment insurance are relieved from both ends because total contributions are raised and the total receipt of benefits is reduced.

Another general finding is that individuals respond to financial incentives in a sense that benefit reductions lead to postponed retirement (chapter 3). This has been shown elsewhere but the new finding is that manual workers postpone considerably less than non-manual workers. However, the surprising result is that manual workers are more flexible than they seem to be which contradicts the popular view that manual workers are unable to retire later. The striking feature is that manual workers do postpone retirement but only until alternative and financially attractive retirement pathways become available. This is another example of the tremendous influence of social security legislation on retirement timing.

A new perspective is taken in an evaluation of the difference in the retirement age between manual and non-manual workers (chapter 4). Research here is motivated from a public debate that is dominated by the view that involuntary early retirement (e.g. due to poor health) is financially disadvantageous and therefore concludes that this is "unfair". This type of argument is usually accompanied by the picture of workers in physically demanding occupations such as roofers or plumbers. Our approach attempts to move the discussion towards an economic one by embedding the subject into the literature of (in-)equality of opportunity. This framework allows to examine how predetermined background characteristics from the early childhood transmit through physical demands of occupations on the retirement age which is a determinant of social security wealth. The main result is that about one third of the difference is explained by individual background characteristics. Contrarily, this implies that about two thirds are due to effort and luck and thus "acceptable" if one only requires opportunities to be equal.

Finally, the present dissertation is concerned with selection into specific retirement ages that produces remarkable patterns with respect to mortality expectations (chapter 5). We examine the remaining years to live after retirement drawing on unique administrative data that contain information on both the beginning and end date of pension benefit receipt. Male individuals who retire at age 63 live systematically longer compared to those who retire between 60 - 62. Such differential mortality expectations must be considered when defining new retirement policies. Redistribution in the pay-as-you-go system depends on the number of years of benefit receipt where those who live longer receive more, ceteris paribus. Recent policy changes have been legislated in Germany without regarding this aspect.

The remarkable increase in life expectancy of the German population has been outlined in the introduction to this work. Such changes are prevalent in many industrialised countries and future research should find answers on how to redesign pension systems for more flexible retirement. Anchoring the “normal retirement age” at some flexible value as determined by the development of life expectancy combined to a well-defined plan of actuarial adjustments could help to establish a stable system in the long-run. Scientific guidance is important because legal rules strongly affect retirement decisions and policy makers exert a fair amount of influence on what people do.

One principal problem of empirical research is the quality of the underlying data. This dissertation highlights an important trade-off between different types of data sources. On the one hand, we have administrative data of exceptional quality but only few relevant variables are available. On the other hand, we have access to survey data that provide a huge amount of information which is, however, subject to measurement error. Once again this confirms the need for diverse empirical studies that investigate research questions on different data sources from different institutional backgrounds and thus allow to generate a more aggregate picture. State of the art techniques help to progress on the identification of causal effects but these methods are as well limited to the quality of available data. Future routes for research should proceed on linking different sources of administrative data or administrative data to survey data which is not yet fully satisfying. Not only would such combinations allow researchers to improve the efficiency of their estimates; they would allow to encounter research questions that have been impossible to answer to the present date.

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