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Matthias Giesecke[1]

Actuarial Adjustments, Retirement Behaviour and Worker Heterogeneity

Abstract

The behavioural response with respect to actuarial adjustments in the German public pension system is analysed. The introduction of actuarial adjustments serves as a source of exogenous variation to estimate discrete time transition rates into retirement. The analysis is conducted on administrative data from social security records and on survey data in a comparative scenario. Probability mass points that occur for institutional reasons and due to social norms are controlled for. Moreover, worker heterogeneity is taken into account, which has not been addressed in the previous literature. The results show that on average retirement is postponed by five months due to financial incentives via actuarial adjustments. However, this response is about 40 per cent lower for manual workers compared to non-manual workers which indicates that their retirement income may deteriorate.

JEL Classification: C41, H55, J26

Keywords: Natural experiment; actuarial adjustments; retirement; worker heterogeneity

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1 Introduction

Demographic change jeopardises pay-as-you-go pension systems, as noted by numerous authors (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. Actuarial adjustments are one possible way to incentivise postponed retirement and thus to redesign a pay-as-you-go pension system towards more actuarial neutrality. The purpose of this paper is to analyse the response in retirement behaviour with respect to a major reform that introduced actuarial adjustments into the German pay-as-you-go pension system from 1997 to 2004. Consequently, retirement benefits are reduced by 3.6 per cent for each year by which an old age pension is claimed early.[1] Those benefit reductions are permanent, which means that they prevail for all periods of benefit receipt. Looking forward in terms of expected present discounted values therefore reveals that retirement incomes decrease remarkably once benefit reductions apply. The central research question of this paper is whether individuals postpone retirement and - if so - by how many months. Moreover, differences in this behavioural response are analysed for individuals with formerly harsh occupations compared to individuals with formerly soft occupations.

The existing literature finds a substantial response to actuarial adjustments in terms of postponed retirement. Börsch-Supan and Schnabel (1999) report an increase of about six months in mean retirement ages as reported from ex-ante simulations. Using data until 2002, Hanel (2010) finds that financial incentives as imposed by actuarial adjustments induce postponed benefit claims by up to 14 months. However, some puzzles remain to be resolved for new economic insight such that the main contributions of this study can be summarised as follows. First, this study covers the full implementation period of actuarial adjustments from January 1997 to December 2004, which has not been the case in previous work. Second, the differentiation between manual and non-manual workers allows to reveal how individuals that are heterogeneous with respect to the harshness of their occupation respond differently to financial incentives in their timing of retirement. If individuals with physically demanding occupations respond less to actuarial adjustments,

[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 2.

then they are subject to benefit reductions that are systematically higher. This aspect, while important and continuously present in the public debate about retirement policy, has not been analysed in the previous economic literature. Third, the results in this study are derived from two diametrically opposed data sources. Survey data are used to extend the analysis to relevant aspects of worker heterogeneity but also allow to confirm the robustness of specific empirical results as drawn from social security records.

In this paper, discrete time duration models are estimated where exits into retirement are defined as failure event. Starting at age 60, estimated hazard rates serve to predict the mean duration for exits into retirement. Identification is based on a natural experiment, where the intensity of actuarial adjustments (i.e. the magnitude of benefit reduction) is a function of the date of birth only. This 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). First, administrative data are used to exploit precise information on worker biographies where exact retirement entries are documented on a monthly basis. Second, survey data are used to draw on a rich set of individual socio-demographic information, which plays an important role in the rather complex decision process that underlies retirement entry behaviour. A baseline scenario compares estimates of the impact of financial incentives on the timing of retirement for social security records and survey data. In this baseline scenario, information that is available in both data sources is used to estimate similar models.[2] Second, survey data are exploited to draw on further determinants that may be crucial in the rather complex retirement decision process.[3] Most importantly, differences in the behavioural response of manual workers and non-manual workers are analysed. Finally, data patterns that are relevant for the age group between 60 and 65 are modeled explicitly. That is, probability mass points with respect to retirement entries that are due to the institutional setting or social norms are taken into consideration.

The results clearly indicate that introducing actuarial adjustments induces a behavioural response in terms of postponed retirement. On average, retirement is postponed by five

[2]Information that is available in both data sources comprises region, sex, duration of non-retirement and benefit entitlements.

[3]Further determinants include marital status as well as individual health and years of education.

months due to financial incentives, where the delay in retirement is about seven months for men and about 4.5 months for women. This finding is robust across data sources and distributional assumptions. However, the behavioural response to the reform differs substantially by the subgroups of manual and non-manual workers. Using additional information from survey data reveals this clear pattern of worker heterogeneity. Individuals with formerly harsh occupations postpone retirement only by about three months on average. Their response to actuarial adjustments is some 40 per cent lower compared to individuals with formerly soft occupations. Moreover, specific retirement patterns around age 63 suggest that manual workers seem to claim disability pensions at large scale as soon as this type of old age pension is available without reductions. This supports the view that formerly harsh occupations are correlated to poor health and delays in retirement are a rather improbable event for this group. The consequence from these findings is that retirement benefit reductions are somewhat higher for individuals with formerly harsh occupations which may translate into considerable inequality in social security wealth.

The remainder of this paper is structured as follows. Section 2 provides an overview on the institutional setting and the corresponding reform to be analysed. Section 3 reviews the relevant literature, illustrates the theoretical background and formulates hypotheses to be tested. In section 4, the econometric strategy and identification are outlined. Furthermore, the two data sources and the sample construction are described and some descriptive statistics are provided. Section 5 presents all results and section 6 concludes.

2 Institutional Setting

Germany has one of the oldest public pension systems which has been converted into a genuine pay-as-you-go pension scheme essentially without any privately funded pillar in 1969. As the result of an influential reform in 1972, during an era of prosperity and strong economic growth, the generosity of this pension scheme was increased dramatically. Consequently, the replacement rates were far above from what is known to be actuarially fair. Moreover, the design of the system imposed strong disincentives for labour supply at late stages of a working career and to retire early instead. Due to these disincentives and demographic change, the system inevitably ran into serious financing problems (see e.g.

Börsch-Supan and Schnabel, 1998, 1999; Börsch-Supan, 2000a). Without considerable adjustments, either the replacement rate would have dropped or the contribution rate would have increased remarkably or both. Several reforms starting in the early 1990's eventually introduced mechanisms to countervail this process.

The relevant reform was implemented between January 1997 and December 2004 for the corresponding birth cohorts from 1937 to 1944. As of 2005, all individuals in the respective age are fully affected by actuarial adjustments, if they retire early and have claims from employment contracts that are subject to social security contributions. 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).[4] The so called entry factor ("Zugangsfaktor") is part of the pension formula and equal to one, if old age pensions are claimed at the NRA. This entry factor reduces by the value 0.003 for each month of early retirement, which technically implements the actuarial adjustment (§ 77 SGB VI, German Social Security Code). 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 year previous to the NRA, i.e. at age 60. However, the early receipt of an old age pension is subject to some restrictions in the German social security legislation and can be described as follows (for the relevant period). 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 severely disabled.[5] That is, only individuals who fulfill the aforementioned requirements can receive an old age pension before the NRA of 65.[6] As indicated, the reforms phased in gradually from 1997 to 2004, but with slightly different timing for the four different types of eligibility.

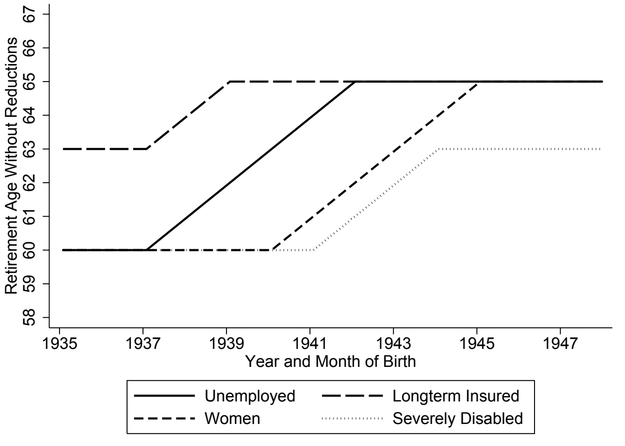
Figure 1 describes the gradual increase in the retirement age without reductions by

[4] The pension formula is used to calculate the monthly public pension entitlements of each individual when entering retirement in Germany. It takes into account individual aspects such as years and amounts of contribution but also aggregate aspects such as the current annuity value.

[5] The final rules for old age pensions due to unemployment, for women and for longterm 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).

[6] 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.

Figure 1: Gradual Increase of Retirement Age Without Reductions across Eligibility Types.



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

types of eligibility. The reform raised the reduction-free retirement age of an old age pension *(i)* due to unemployment for the birth cohorts 1937 to 1941, *(ii)* for women born between 1940 and 1944, *(iii)* for long-term insured in the birth cohorts 1937 to 1938 and *(iv)* for severely disabled individuals born between 1941 and 1943.

3 Literature, Theory and Hypotheses

3.1 Literature Review

The timing of retirement is an issue of labour supply, as is the trade-off between consumption of goods and leisure which is central for retirement decisions (Hurd, 1990). An early contribution by Weiss (1972) elaborates the optimal lifetime pattern of labour supply, which has some relevance for older workers when deciding upon their retirement age. However, this model is limited to the extent, that it does not take into account institutional and social constraints as imposed by social security design and unobserved social norms. Sheshinski (1978) motivates his model of social security and retirement decisions by the finding, that the presence of retirement benefits crucially influences consumption paths and retirement ages. This latter model is one of intertemporal utility maximisation over the life cycle. Further seminal papers, such as Gordon and Blinder (1980), Crawford and Lilien (1981) and Gustman and Steinmeier (1986) all have in common to establish models, where individuals maximise utility over the life cycle and chose to retire exactly when the benefits and costs for this decision equalise. The challenge in these models is to precisely pin down the interdependence between life cycle preferences and external incentives from the policy design, which is a complicated exercise - but is of very much importance for this present study. In a world of perfect capital markets, actuarial fairness and certain lifetimes, Crawford and Lilien (1981) find that there is no effect of social security on individual retirement decisions, which is the reason why they relax these assumptions each in turn. One important finding is that deviating from actuarial fairness can induce earlier retirement. Gordon and Blinder (1980) estimate wage equations, where individuals retire, once their market wage is below their reservation wage. Using U.S. data, they find that individual retirement decisions are much more influenced by age effects and private pension plans compared to social security, which does not induce a noteworthy amount of early retirement. The major contribution of Gustman and Steinmeier (1986) is that they implement a structural life cycle model which allows individual responses in labour supply to changes in their budget constraints that may impose different incentives at different ages.

Fields and Mitchell (1984b) provide simulation results of a change in the early retirement reduction factor for the U.S. in the 1980s, where the reduction is increased from 6.66 per cent to 15 per cent for each year of early retirement. The finding is that individuals postpone retirement by some three months as a response to such a (simulated) policy change. Further influential studies by Fields and Mitchell (1984a) and Mitchell and Fields (1984) point out an important aspect in the framework of retirement decision modeling: the expected present discounted value (EPDV). In contrast to the wealth level or income at a given point in time, this measure discounts all future income streams while taking into account uncertain lifetimes. The accrual rate is the relative change of the EPDV and indicates the expected gain from postponing retirement by one period. In slightly modified versions, this approach has established some prominence in several empirical studies, such as Samwick (1998), Börsch-Supan and Schnabel (1999), Börsch-Supan (2000b), Coile et al. (2002) or Hanel (2010) among many others.

An alternative approach as elaborated in a seminal study by Stock and Wise (1990) is the so-called option value. In this model, an individual irreversibly retires if there is no expected gain from future work (as measured in utility) or chooses to continue work and retire in any 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”). The option value approach is related to the EPDV or its accrual rate as it incorporates the future stream of income from work and retirement, but it differs (among other aspects) in that it weights future utility as indirectly derived from earned income and pension benefits. Despite the theoretical appeal of this model, empirical implementation is difficult for several reasons. First, it requires data that allow to precisely estimate the parameters of a CRRA utility function. Numerous empirical studies such as Samwick (1998), Börsch-Supan (2000b), Blundell et al. (2002) or Asch et al. (2005) have applied the option value and frequently the estimation of a fully structural model is circumvented by assuming parameter values for the preferences over risk and leisure. Second, as noted by Coile and Gruber (2000), variation in the option value predominantly originates from variation in wages, which is problematic when identifying the impact of social security benefits on retirement behaviour. Especially if wages are correlated to preferences for retirement, identification of the impact of social security benefits on retirement is difficult.[7]

Another relevant study is from Blau (1994), who reveals important patterns of labour force dynamics of older male workers by discriminating between the states of full-time employment, part-time employment and out of labour force in a discrete time hazard model. One key finding of his study is that the level of social security benefits substantially incentivises labour force exits.

Previous studies for Germany suggest that early retirement was prevalent in the German pay-as-you-go pension system before some major reforms phased in during the late 1990s. Siddiqui (1997) reveals that the German pension system provides financial incentives to retire early, using survey data for the years 1984 to 1990. Very much in line with this result is a study of Börsch-Supan and Schnabel (1999), who find negative accrual rates for individuals who do not retire early. Moreover, Börsch-Supan (2000b) finds strong incentives for early retirement in Germany, using survey data for the years 1984 to 1996. Further evidence suggests that individuals substantially postpone retirement as a

[7] See Coile and Gruber (2000) for a detailed discussion of this aspect.

response to benefit reductions on early retirement (Hanel, 2010). However, little is known on heterogeneous effects of benefit reductions, which are globally imposed on early retirement ages for old age pensions, irrespective of differences over subgroups such as manual and non-manual workers.

3.2 Theoretical Considerations and Hypotheses

The decision to retire is a complex phenomenon, as it 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 subject to a lifetime budget constraint, can explain retirement decisions while taking into account the interdependence between preferences and incentives. The starting point here is an intertemporal model of consumption and retirement as in Samwick (1998), where lifetime utility can be formulated as

$$U = \int_t^R e^{-\delta(s-t)} u(c_s, 0) ds \int_R^T e^{-\delta(s-t)} u(c_s, 1) ds \tag{1}$$

and $u(\cdot)$ is utility from consumption and leisure, δ is the individual discount rate of future utility and leisure from retirement (the second argument in $u(\cdot)$) switches from 0 to 1 as soon as transitions into retirement are observed. Maximisation of $u(\cdot)$ is with respect to consumption c_s and the retirement date R and is subject to the lifetime budget constraint

$$\omega_t + \int_t^R e^{-r(s-t)} y_s ds + \int_R^T e^{-r(s-t)} B_s(R) ds - \int_t^T e^{-r(s-t)} c_s ds = 0 \tag{2}$$

where r is the interest rate, ω_t is wealth at time t, the second term corresponds to exogenous lifetime earnings (the sum of y_s), the third term corresponds to the sum of all future retirement benefits $B_s(R)$ as a function of the retirement date R and the fourth term is lifetime consumption. It is important to note that R crucially influences social security wealth in the third term of equation (2), which acknowledges the design of the German public pension system. Differentiating with respect to the retirement date R

yields the first order condition

$$e^{-\delta(R-t)}[u(c_R,1)-u(c_R,0)]=\lambda e^{-r(R-t)}\left[(y_R-B_R(R))+\int_R^T e^{-r(s-R)}B_s'(R)ds\right] \quad (3)$$

The result in equation (3) is important as it illustrates how retirement decisions are driven in the lifetime utility framework. 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. According to this result, an individual will postpone retirement if an increase of the retirement date corresponds to a utility gain from the future stream of retirement benefits. This argument is central for the present study, because actuarial adjustments are benefit reductions that may reduce present discounted values if exits into retirement take place before the normal retirement age.

To transmit this result into the empirical framework of this paper, few modifications have to be made. First, the focus in this paper is on the impact of the change in retirement wealth on retirement age. Following the notion of Samwick (1998), the "actuarial present value" (APV) of all future retirement benefits of an individual is given by

$$\text{APV}(R)=\int_R^T e^{-r(s-R)}B_s(R)ds \quad (4)$$

which is conditional on retirement at age R. Second, 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 quantity is the "expected present discounted value" (EPDV)

$$\text{EPDV}(R)=\sum_{s=R}^{T}\pi(s)\delta^{s-t}B_s(R) \quad (5)$$

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 (Statistisches Bundesamt) (2012). Note that values for EPDV are calculated at each month between age 60 and 66, i.e. $t = 1, ..., 72$ for each individual.[8] The existing literature points out that the use of forward-looking incentive measures supports identification of the impact of social security benefits on retirement behaviour (see e.g. Samwick (1998), Coile and Gruber (2000), Coile et al. (2002)). Third, financial incentives from the introduction of actuarial adjustments need to be measured. This is incorporated into equation (5) by supplementing an adjustment factor $(1-\tau(R))$, which can be written as

$$\text{EPDV}_{\tau(R)}(R) = (1-\tau(R))\,\text{EPDV}(R) = (1-\tau(R))\sum_{s=R}^{T}\pi(s)\delta^{s-t}B_s(R) \tag{6}$$

where $0 \leq \tau(R) \leq 0.18$ is the adjustment rate as implied by the German social security legislation. Whether actuarial adjustments apply (i.e. $\tau(R) > 0$) and to what extent depends on the exact retirement age as well as year of birth and month of birth, and the type of old age pension.[9] Finally, subgroup heterogeneity for manual and non-manual workers and differences in their behavioural response to financial incentives is analysed. Such heterogeneity is revealed by computing mean durations as predicted by estimated models across subgroups for manual and non-manual workers, which may differ in their response to actuarial adjustments.

In the empirical analysis, two fundamental hypotheses are tested.

Let a denote the normal retirement age and let $a-1$ denote some early retirement age. Suppose that actuarial adjustments reduce monthly retirement benefits by 3.6 per cent if retirement takes place at age $a-1$ compared to retirement at age a.

Hypothesis 1 *Financial incentives as imposed by those actuarial adjustments lower the relative attractiveness of early retirement at age $a-1$. Increasing the adjustment rate $\tau(R)$ raises the probability that transitions into retirement are postponed (e.g. from $a-1$ to a).*

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

[9] See section 2 for a detailed description of the reform.

Let π_A be the survival probability for individuals with soft occupations and let π_B be the survival probability for individuals with harsh occupations. Further suppose that $\pi_A > \pi_B$.

Hypothesis 2 *The relative change in the hazard rate for exits into retirement is larger for individuals with formerly soft occupations ("non-manual" workers) compared to individuals with formerly harsh occupations ("manual" workers). That is, the degree of postponement for transitions into retirement is larger for individuals with formerly soft occupations as compared to individuals with formerly harsh occupations.*

4 Conceptual Framework and Data

4.1 The Two Datasets

The first data source contains waves from 2002 to 2010 from the Insurant Account Sample (Versicherungskontenstichprobe, VSKT). The VSKT are administrative data provided by the German Federal Pension Insurance (DRV-Bund) and their primary purpose is to serve for internal calculations of DRV-Bund and for political consulting. Altogether, the VSKT is a sample of about 240,000 individuals of age 14 to 66, regarding their insurance accounts. This study makes use of a 25% subsample of the VSKT, which is provided as Scientific-Use-File and includes some 60,000 individual observations. The observed entries into the VSKT have in common that they are any employee-related information which is subject to social security contributions. Thus, sample selection is present to the extent, that civil servants and self-employed individuals are ruled out from the analysis.[10] Outstanding features of the dataset are the large number of observations and the high reliability of data that are process-produced and therefore do not suffer from typical problems of survey data (Himmelreicher and Stegmann, 2008). Information is available on 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

[10] The German public pension systems offers the possibility to self-employed individuals, to contribute voluntarily and therefore receive pension benefits after retirement. In fact, there is a small number of voluntarily insured self-employed individuals who do participate. As this group is a somewhat specific group and a very small minority, they are ruled out from further analysis. Further note that civil servants can be part of the sample, if they formerly contributed as employees.

(SOEP) are used from the panel waves 1995 to 2011.[11] 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. The resulting activity spells are on a monthly basis. In contrast to the administrative data, a rich set of socio-demographic information on the individual and the household level is available which is used to identify worker heterogeneity in retirement entry behaviour. Most notably, marriage status, health status and occupational information to distinguish manual workers from non-manual workers are available. Generally, the SOEP allows to pin down the socio-demographic situation of individual observations in a much more detail compared to the VSKT data. This aspect is of particular importance, as the determinants for retirement entry decisions have a somewhat complex relationship to characteristics that vary on individual and household level. While administrative data provide very precise information on retirement entries and entitlements, the analysis may suffer from omitted variables.

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 VSKT data, retirement is defined as benefit claim, which is documented without measurement error. For SOEP data, retirement is defined as self-reported retirement status in retrospective questions, which are subject to some 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 the early retirement age (ERA) if eligibility is achieved.[12] 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 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 three per cent of the total spells are right-censored, which means that no exit into

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

[12] See section 2 for the four relevant types of old age pensions.

retirement has taken place until age 66.[13]

Table 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 1 reports 1497 observed individuals (i.e. spells) for survey data (SOEP), which amounts to a total of 42353 person-month-observations. There are 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,

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

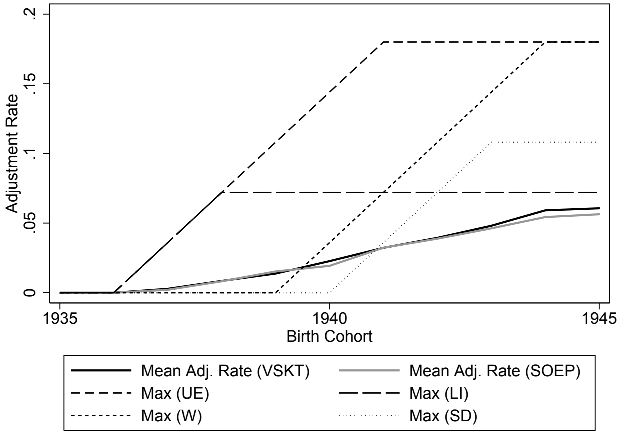
which implies that birth cohorts previous to the reform serve as control group, i.e. those cohorts in the sample that are not affected by actuarial adjustments (1935 and 1936). During the implementation period, the treatment intensity differs by month and year of birth on the one hand and the type of 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 actuarial adjustments 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 2 illustrates the gradual increase of actuarial adjustments 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 shows the share of retirement entries for age 60 and age 65 across birth cohorts

Figure 2: Adjustment Rates at Retirement across Birth Cohorts.



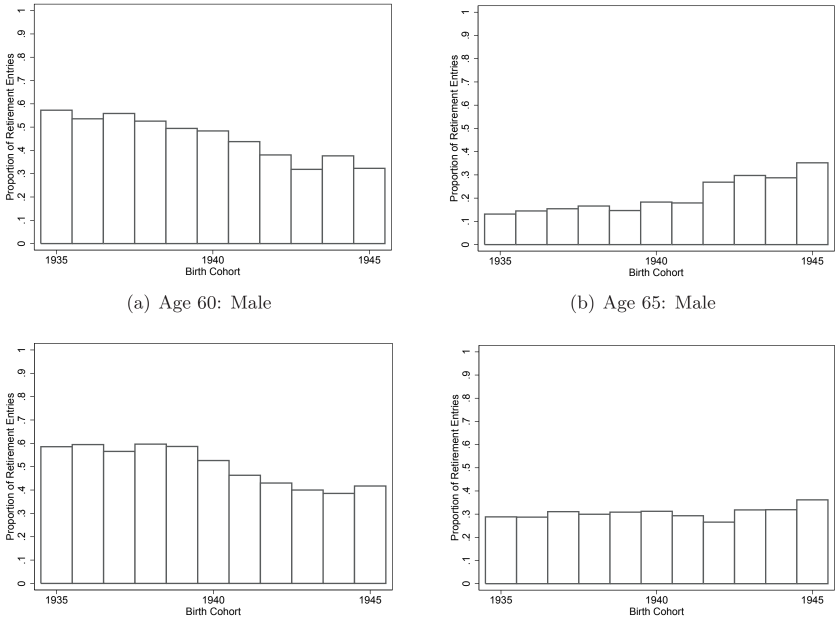
Source: Own calculations based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011).
Note: Mean adjustment rates are arithmetic means at retirement entry, plotted across birth cohorts for both social security records (VSKT) and survey data (SOEP). Maximum adjustment rates are for different types of eligibility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

for males (panel a and b) and females (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 males and females. This 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.[14] Clearly, these patterns are purely descriptive and do not help to establish any causal relationship. However, they do indicate that retirement ages increase by birth cohorts during the reform period.

All regressors are allowed to vary over time, if such variation is present in the data;

[14]Those differences are in line with official statistics as reported in German Federal Pension Insurance (2013), where average retirement ages for men are initially (i.e. in 1995) lower but eventually align to those of women.

Figure 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.

data structure and subsequent models allow for time-varying regressors, such that any relevant variation is captured. Table 2 provides an overview on all variables that are part of the analysis.[15]

Information that is of central interest for the identification of the impact of actuarial adjustments on the timing of retirement are reported as “Reform Variables”. The variable “Incentive” is the percentage difference between the expected present discounted value (variable “EPDV” as defined in equation (5)) and the adjusted expected present discounted value (variable “Adjusted EPDV” as defined in equation (6)). This quantity, however, is nothing but the adjustment rate $\tau(R)$, since we have

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

Table 2: Descriptive Statistics.

	VSKT			
	Mean	Std.Dev.	Min.	Max.
Dependent Variable				
Benefit Claim	.0348	.1833	0	1
Reform Variables				
Incentive (Adj. Rate)	.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				
Incentive (Adj. Rate)	.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
Incentive * 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.

$$\text{Incentive}(R) = \frac{\text{EPDV}(R) - \text{EPDV}_{\tau(R)}(R)}{\text{EPDV}(R)} = \frac{\text{EPDV}(R) - [1-\tau(R)]\text{EPDV}(R)}{\text{EPDV}(R)} = \tau(R) \tag{7}$$

The variable "Incentive" has a somewhat larger mean for survey data, which is due to the assumptions to assign adjustment rates to individuals when eligibility types are not known. The variable "Manual" is a dummy that equals one, if individuals are or were formerly manual workers ("harsh occupations") and zero if they are non-manual workers ("soft occupations"). The recoding of this variable follows the International Standard

Classification of Occupations from 1988 (ISCO 88). Finally, table 2 reports the interaction term “Incentive * 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 behavioural response of manual and non-manual workers differs. The remaining variables in table 2 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.

4.3 Econometric Strategy

To identify the impact of actuarial adjustments 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).[16] 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).

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

[16] 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).

$$\lambda_i(t) = \lim_{h \to 0} \frac{\text{Prob}[t+h > T_i \geq t | T_i \geq h]}{h} = \lambda_0(t) \exp(z_i(t)'\beta) \tag{8}$$

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 (8) 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 $S(t)$)

$$\begin{aligned} \text{Prob}[T_i \geq t+1 | T_i \geq t] = S(t) &= \exp\left[-\int_t^{t+1} \lambda_i(u) du\right] \\ &= \exp\left[-\exp(z_i(t)'\beta) \int_t^{t+1} \lambda_0(u) du\right] \end{aligned} \tag{9}$$

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 (9) makes use of the proportional hazards specification in equation (8) and exploits the fact that $z_i(t)$ is constant between t and $t+1$. Reformulating the survival function yields

$$\text{Prob}[T_i \geq t+1 | T_i \geq t] = S(t) = \exp\left[-\exp(z_i(t)'\beta + \gamma(t))\right] \tag{10}$$

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

$$L(\gamma,\beta)=\prod_{i=1}^{N}\left[1-\exp\left[-\exp\left(z_i(k_i)'\beta+\gamma(k_i)\right)^{\delta_i}\right]\times\prod_{t=1}^{k_i-1}\exp\left[-\exp\left(z_i(t)'\beta+\gamma(t)\right)\right]\right] \tag{11}$$

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. Taking logs finally yields the log-likelihood function

$$\ln L(\gamma,\beta)=\sum_{i=1}^{N}\left[\delta_i\ln\left[1-\exp\left[-\exp\left(z_i(k_i)'\beta+\gamma(k_i)\right)\right]\right]-\sum_{t=1}^{k_i-1}\exp\left(z_i(t)'\beta+\gamma(t)\right)\right] \tag{12}$$

which is maximised using the “cloglog”-procedure as implemented into the statistical software STATA. In the presence of unobserved heterogeneity, however, the results from this model may be incorrect. 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, unobserved heterogeneity is taken into account again making use of the model proposed by Prentice and Gloeckler (1978) and Meyer (1990). The hazard function from equation (8) must be rewritten as

$$\lambda_i(t)=\theta_i\lambda_0(t)\exp(z_i(t)'\beta) \tag{13}$$

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

$$
\begin{aligned}
\ln L(\gamma,\beta,\sigma^2) = \sum_{i=1}^{N} \ln\Bigg[& \Big[1+\sigma^2 \sum_{t=0}^{k_i-1} \exp\big(z_i(t)'\beta+\gamma(t)\big)\Big]^{-\sigma^{-2}} \\
& - \delta_i \Big[1+\sigma^2 \sum_{t=0}^{k_i} \exp\big(z_i(t)'\beta+\gamma(t)\big)\Big]^{-\sigma^{-2}} \Bigg] \qquad (14)
\end{aligned}
$$

which is maximised using the "pgmhaz8"-procedure as implemented into the statistical software STATA.

4.4 Identification of the Effect of Actuarial Adjustments on Retirement Age

Making use of the abovementioned methodology, hazard rates of entering retirement are estimated. Starting in the month after the 60th birthday (first eligibility), individuals are observed repeatedly until they enter retirement. In the estimation framework, the dependent variable takes the value zero for each month of non-retirement and the value one if an individual enters retirement.[17] Identification is based on a differences estimator by discriminating between a treatment group that is affected by the reform ("after") and a control group that is not affected ("before"). Clearly, the introduction of actuarial adjustments must be identified as a source of exogenous variation in order to be interpreted as a causal effect. This is supported by the fact, that the reform affects individuals only by their month and year of birth. 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 actuarial adjustments. This requires the assumption that no birth cohort heterogeneity is present and under such circumstances the reform can be interpreted as a natural experiment (see Hanel, 2010). Moreover, adaptive behaviour previous to the reform that may counfound estimates is very unlikely because from a decision-theoretical point of view 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.

[17] 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, several reasons for sample drop-outs, mainly due to refused further participation, occur frequently.

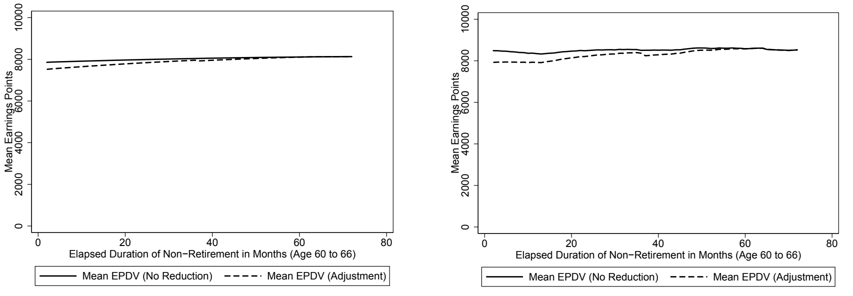
In the estimation procedure, the EPDV is central for identification. For the computation of EPDVs it is essential to have access to information on pension entitlements. The two data sources provide very different information on these aspects. For social security records (VSKT) pension claims are measured 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.[18] 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 before reaching age 65. The most straightforward assumption is that for each individual, the observed average monthly pension claim is extrapolated to higher ages if retirement is earlier. For survey data (SOEP), average earnings points as indirectly derived from labour income are extrapolated from age 60 to 66.[19] 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 4. Evidently, EPDVs follow similar patterns for both data sources but are measured with less precision for the SOEP.

[18] 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 in § 68 SGB VI, which takes into account wage growth and changes in the share of retirees.

[19] 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 4: Expected Present Discounted Value (Sample Mean).



(a) VSKT (b) SOEP

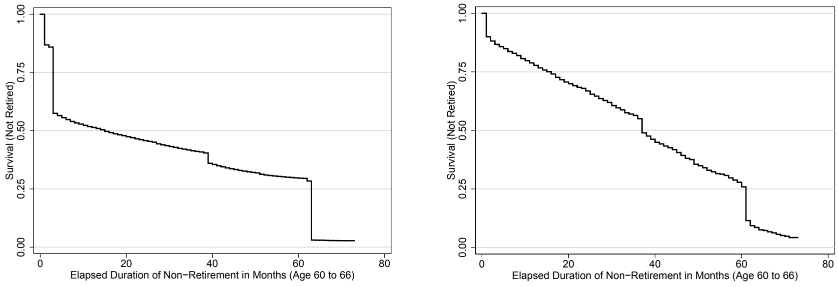
Source: Own calculations based on SUFVSKT2002-SUFVSKT2010 and SOEP (1995-2011).

4.5 Specific Data Patterns

The observed cohorts in this study follow very specific patterns with respect to their retirement age as shown in figure 5. These patterns occur for different 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. The corresponding patterns are induced by both early retirement ages and the normal retirement age. Second, occupational agreements with employers that do not necessarily comply with applicable law, may influence retirement patterns. Third, social norms may as well have an impact. While the first aspect is easily traced back, the other two are difficult to measure.

From the Kaplan-Meier survival estimates in figure 5 we can infer that a large fraction of retirement entries takes place exactly in the month of first eligibility. Those probability mass points cause a considerable amount of variation that needs to be taken into account when identifying the causal effect of actuarial adjustments. To sufficiently control for cumulative retirement entries, the subsequent analysis employs the most flexible version of a discrete time duration model, where duration dummies (i.e. age) are included for every month.

Figure 5: Retirement Entry and Probability Mass Points.



(a) VSKT (b) SOEP

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

4.6 Measurement Error and Omitted Variables

Since two different data sources are used, some specific characteristics should be kept in mind when discussing results subsequently. For social security records (VSKT), the omitted variable bias may attenuate estimated coefficients towards zero. For survey data (SOEP), measurement error in regressors may attenuate estimated coefficients towards zero. While the magnitude of these biases is unknown, the estimated response in retirement behaviour is a lower bound for the true (but unknown) response in both cases.

5 Results

In the baseline scenario, 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. The baseline scenario serves for comparative purposes to evaluate diverging results from the two data sources. It is important to note that these models are not perfectly similar for two reasons. First, the data quality differs very much by the characteristics of the two data sources (e.g. how benefit entitlements are calculated), which requires different assumptions. Second, eligibility types are precisely identified in the VSKT, but not so for the SOEP. The results are provided in Table 3.[20]

[20] 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

Table 3: Baseline Estimation: Actuarial Adjustments 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)						
Incentive	–.068	(.007)	–.079	(.007)	–.067	(.007)
Adjusted 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)						
Incentive	–.190	(.030)	–.198	(.030)	–.183	(.030)
Adjusted 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. Expected present discounted value is abbreviated by EPDV. Incentive is the relative change between EPDV and adjusted EPDV for each duration time.

The key regressor "Incentive" 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 "Incentive" 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. For survey data, the marginal effect for "Incentive" 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 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

log-log).

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 actuarial adjustments is rejected. This is in favour of Hypothesis 1, that actuarial adjustments lower the attractiveness of relatively 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. Those differences in magnitude of the estimated coefficients (and respective average marginal effects) seem substantial. This result is not surprising, when taking into account that the underlying data sources differ fundamentally by the quality and quantity of information that they provide. Moreover, the estimated coefficient for male individuals is negative for social security data which is contradicted by a positive sign in the baseline estimation on survey data. However, this result is very likely to be driven by the sample composition (see table 2).

When interpreting the results, it must be kept in mind that omitted variable bias may be present in analyses on VSKT data, if important variables are missing. Retirement decisions are outcomes of a rather complex relationship of a variety of aspects. Besides financial resources as discussed in section 3, other determinants such as marital status and health are important. 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). Central interest in this study is on the harshness of former occupations, which is likely to reflect health status in the context of retirement.

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. Consequently, the marginal effect of actuarial adjustments as measured by "Incentive" 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 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 "Incentive * Manual". This result indicates, that manual workers respond to a much lower degree to actuarial adjustments

Table 4: Actuarial Adjustments, 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)						
Incentive	−.226	(.034)	−.234	(.033)	−.220	(.034)
Adjusted EPDV	.000	(.000)	.000	(.000)	.000	(.000)
Manual	.003	(.002)	.002	(.002)	.002	(.002)
Incentive 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. Expected present discounted value is abbreviated by EPDV. Incentive is the relative change between EPDV and adjusted EPDV for each duration time.

compared to non-manual workers. For both, "Incentive" and "Incentive * Manual", the probabilities that these results occur by chance are not larger that 0.01 and thus the null hypothesis is rejected, which is in favour of Hypothesis 2. Further results suggest that living in West Germany and being married support retirement at higher ages. Moreover, self-reported health status seems to be fully captured by controlling for the harshness of occupations, as indicated by relatively large standard errors.

5.1 Predictions

Expected duration times from first eligibility (age 60) until retirement are used to quantify the behavioural response by subgroups. All values as reported in table 5 are predicted from the complementary log-log model. The levels of duration times are systematically lower for social security records (VSKT) which is a likely result from measurement error in the SOEP. The relevant finding from these figures, however, is that differences are very robust when comparing the two data sources. On average, individuals who are exposed to actuarial adjustments, postpone retirement by about five months compared to the control group. The response for male individuals 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, 2013).[21] Most importantly, individuals with formerly harsh occupations postpone retirement only by about three months which implies that their behavioural response is some 40 per cent lower compared to individuals with formerly soft occupations.

Table 5: Expected Duration of Non-Retirement (from First Month of Eligibility).

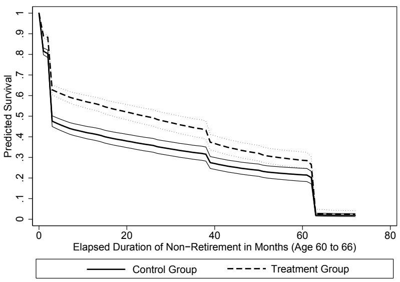
	Baseline: VSKT		
	Before Reform	During/After Reform	Difference
Full Sample	23.2	28.5	5.3
Male	20.9	27.9	7.0
Female	24.7	28.9	4.2
	Baseline: SOEP		
	Before Reform	During/After Reform	Difference
Full Sample	27.3	32.4	5.1
Male	24.9	32.1	7.2
Female	28.3	32.8	4.5
	Further Information: SOEP		
	Before Reform	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.

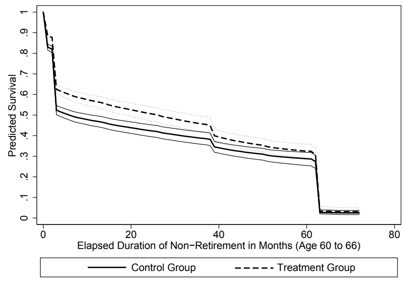
[21] Differentials in retirement ages (old age pensions) between men and women depend on the region, with some variation across former east and west Germany. In 1995 and across all regions, old age pensions were claimed at age 62.5 (women) and 62.3 (men), indicating that men retired two months earlier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

Predicted survival functions for baseline estimations are shown in figure 6. All predictions are in line with the general finding, that actuarial adjustments 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), 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.

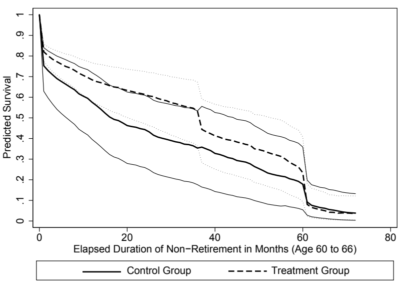
Figure 6: Predicted Survival Rates: Baseline Estimation.



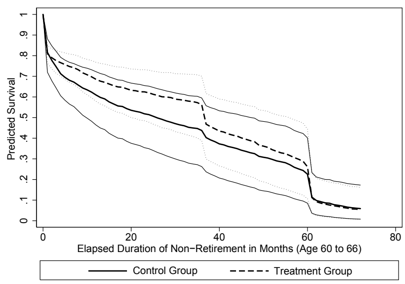
(a) VSKT: Men



(b) VSKT: Women



(c) SOEP: Men

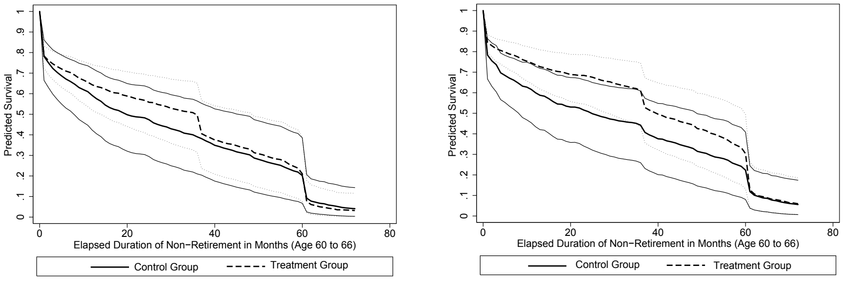


(d) SOEP: Women

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

Predicted survival functions for the estimation including further information from the SOEP are shown in figure 7. The focus is on differences in the behavioural response of individuals with formerly harsh occupations versus formerly soft occupations. The plotted lines confirm that the response to actuarial adjustments is much larger for individuals with formerly soft occupations; moreover, their survival rates are considerably larger in absolute terms. In this context, one further pattern for manual workers becomes evident and needs some attention. For predicted survival rates of manual workers in the treatment group, there is a remarkable drop at age 63, such that predicted values are very much the same thereafter. At this age, old age pensions for severely disabled individuals are available without reductions after full implementation (see figure 1). Manual workers who are affected by the reform seem to choose this alternative as soon as it is available without reductions. No such drop occurs for manual workers in the control group which clearly indicates, that other types of old age pension served as pathways into retirement before actuarial adjustments became effective (e.g. old age pensions due to unemployment). A high take-up rate of old age pensions due to disability of manual workers is in accordance to presumably higher prevalence of poor health in this group 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 7: Predicted Survival Rates for Manual versus Non-Manual Workers.



(a) Harsh Occupations (Manual) (b) Soft Occupations (Non-Manual)

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

5.2 Unobserved Heterogeneity

As outlined in section 4, unobserved heterogeneity is analysed in a framework proposed by Prentice and Gloeckler (1978) and Meyer (1990). The full model without any assumption on the baseline hazard, however, does not converge. To overcome this computational problem, a reduced model without full flexibility with respect to the baseline hazard is estimated. Instead, this model includes dummies to capture variation at entry ages where probability mass points occur. These three dummy variables summarise duration time (i.e. age in months) for the subsequent three months after age 60, 63 and 65. Calendar time enters the model linearly in the shape of the variable “Year”. The baseline models for VSKT and SOEP are reported in table 6, while the model including further information is in table 7 in the appendix.[22] 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 present. However, estimated coefficients are robust in magnitude and sign for all models, such that previous results should not be confounded by unobserved factors.

6 Conclusion

This study investigates the effect of the introduction of actuarial adjustments into the German public pension system from 1997 to 2004 for male and female individuals of age 60 to 65. The central question is to what extent individuals postpone retirement as a response to benefit reductions. Specific interest is on differences in the response behaviour of individuals with soft and harsh occupations. Moreover, results from two different data sources and their implications are critically acclaimed.

The specific data structure allows to estimate discrete time duration models, where exits into retirement are defined as failure event. Hazard rates that reflect the instantaneous probability of entering retirement in a given month, conditional on non-retirement until that month, are estimated. Starting at age 60, estimated hazard rates are used to predict the mean duration for exits into retirement. Identification is based on a natural experiment, where the intensity of actuarial adjustments (i.e. the magnitude of benefit reduction) is a function of the date of birth only. The baseline analysis is a comparative scenario for two different data sources with specific characteristics. On the one hand, social

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

security records provide very reliable information on retirement entries and contribution levels but contain only few explanatory variables. On the other hand, survey data provide a rich set of socio-demographic information but retirement behaviour is documented with less precision. Making use of survey data allows to control for worker heterogeneity and - most importantly - to discriminate between manual and non-manual workers.

The results clearly indicate that introducing actuarial adjustments causes postponed retirement. This finding is robust across data sources and distributional assumptions. However, estimated coefficients on the central reform variable vary considerably in magnitude between data sources. At the same time, these models predict more or less identical hazard rates and corresponding survival times of non-retirement. When taking into account that the underlying data sources differ fundamentally by the quality and quantity of information that they provide, this finding strongly supports the view that the introduction of actuarial adjustments is a source of exogenous variation.

On average, retirement is postponed by five months due to actuarial adjustments. However, the behavioural response to the reform differs substantially by subgroups of the population. Male individuals postpone retirement by about seven months on average while for women the delay in retirement is only about 4.5 months. This result is not surprising once the retirement age previous to the reform is taken into account, which is much lower for men. Using additional information from survey data reveals clear patterns of worker heterogeneity. Individuals with formerly harsh occupations postpone retirement only by about three months on average. Their response to actuarial adjustments is some 40 per cent lower compared to individuals with formerly soft occupations. Moreover, specific retirement patterns around age 63 suggest that manual workers seem to claim disability pensions at large scale as soon as this type of old age pension is available without reductions. This supports the view that formerly harsh occupations are strongly correlated to poor health and delays in retirement are a rather improbable event for this group. The general finding is that retirement benefit reductions are disproportionately high for individuals with formerly harsh occupations such that their retirement incomes may deteriorate.

In terms of policy design, this result requires flexibility in retirement ages for heterogeneous types of retirees under specific circumstances. Heterogeneous response behaviour for individuals that differ by physical demands from former occupations may translate into considerable inequality in social security wealth. From the perspective of actuarial fairness, there is no rationale against actuarial adjustments. However, the crucial question is whether expected life times of manual workers are lower and if so by how much. Manual workers receive benefits earlier if they delay retirement to a lower extent. Consequently and qua construction of the pay-as-you-go pension system, there may arise a substantial amount of redistribution from manual to non-manual workers, if they receive benefits for a shorter total period. The precise investigation of the relationship between differential mortality and retirement age is an open field for future research.

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A Appendix

Table 6: Baseline Estimation: Actuarial Adjustments, Retirement Transitions and Unobserved Heterogeneity.

	Compl. Log-Log without Frailty		Mixed Gamma	
	coeff.	s.e.	coeff.	s.e.
Social Security Records (VSKT)				
Incentive	–2.256	(.251)	–2.424	(.263)
Adjusted 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)				
Incentive	–6.976	(.659)	–7.710	(.876)
Adjusted 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. Expected present discounted value is abbreviated by EPDV. Incentive is the relative change between EPDV and adjusted EPDV for each duration time. Probability mass points are abbreviated by PM, where "+ 3" indicates that duration times (months) after the respective birthday are summarised in each dummy.

Table 7: Actuarial Adjustments, Retirement Transitions, and Worker Heterogeneity.

	Compl. Log-Log without Frailty		Mixed Gamma	
	coeff.	s.e.	coeff.	s.e.
Survey Data (SOEP)				
Incentive	–8.525	(.890)	–9.356	(1.027)
Adjusted EPDV	.000	(.000)	.000	(.000)
Manual	.105	(.082)	.170	(.099)
Incentive 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. Expected present discounted value is abbreviated by EPDV. Incentive is the relative change between EPDV and adjusted EPDV for each duration time. Probability mass points are abbreviated by PM, where "+ 3" indicates that duration times (months) after the respective birthday are summarised in each dummy.