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# THE EFFECT OF BENEFIT REDUCTIONS ON THE RETIREMENT AGE: THE HETEROGENEOUS RESPONSE OF MANUAL AND NON-MANUAL WORKERS

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I estimate 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. Responses to benefit reductions are elaborated separately for manual and non-manual workers. On average, individuals postpone retirement by 13.2 months if pension benefits are reduced by 3.6 percent for each year of early retirement. This result is in line with the previous quasi-experimental literature and suggests that people respond to the incentive of reducing the implicit tax on further periods of work. However, among men the response is about 50 percent 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 disability pensions are available at age 63—without benefit reductions.

**JEL Codes**: C41, H55, J26

**Keywords**: financial incentives, natural experiment, retirement age, expected pension wealth, effect heterogeneity

## 1. Introduction

Ageing populations and the rise of expected years in retirement are a challenge for retirement security.[1] To maintain the current level of pension income, workers need to exit the labor force later, contribute to social security longer, and claim benefits at higher ages. Benefit reductions are one possible way to set incentives for postponed retirement. However, workers in physically demanding occupations may lose their work capacity and may thus be distracted from continued work. The likelihood of early retirement is considerably higher for these people compared to those with physically less demanding jobs. This raises the question whether individuals differ in their occupation- and health-related ability to respond to financial incentives by postponed retirement.

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[1]In OECD countries, the expected years in retirement have risen from 15 to 22.5 years (women) and from 11 to 18 years (men) between 1970 and 2012 on average (OECD, 2014). For a recent overview of the challenge in the U.S., see Poterba (2014).

The purpose of this paper is to analyse the response in retirement timing with respect to a major reform that introduced benefit reductions into the German public pension system between 1997 and 2004. Pension benefits are reduced by 3.6 percent for each year (0.3 percent per month) by which an old age pension is claimed early.[2] Benefit reductions are permanent and prevail for all future periods of monthly pension benefits. 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 two research questions of this paper are, first, to what extent individuals postpone retirement (by how many months) in response to benefit reductions and, second, whether manual and non-manual workers respond differently.

To answer these questions, I estimate the retirement hazard between age 60 and 66 for the birth cohorts 1930–47. I use survey data from the German Socio-Economic Panel (SOEP) that include a wide range of socio-demographic variables to model the complex process of retirement decisions. Throughout the analysis, retirement is defined as self-reported retirement status from retrospective calendar questions. I estimate and predict the elapsed duration of non-retirement after age 60 when people become eligible for old age pensions in Germany. All estimates are obtained from a discrete time proportional hazards model in its most flexible version with duration indicators for each month at risk.

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. The gradual implementation of benefit reductions induces cohort discontinuities that I use to estimate the effect on the retirement timing. I provide comprehensive graphical evidence on the retirement distribution across birth cohorts, showing that there is substantial variation in the data. Exploiting this type of exogenous variation in pension benefits supports identification because the determinants of expected pension wealth such as previous earnings are highly correlated to labor market attachment, which may confound the estimated effect of financial incentives on retirement timing (see, e.g. Krueger and Pischke, 1992).

The results indicate that introducing benefit reductions has a substantial effect on the timing of retirement. The predicted response to benefit reductions is a delay in retirement by 13.2 months on average. The magnitude of the estimated effect differs considerably by population subgroups but is largely in line with previous findings for Germany (Börsch-Supan and Schnabel, 1999; Hanel, 2010, 2012). The novel feature of this paper is that I show how the response to benefit reductions differs by physical demands of occupations. Among men, the model predicts that manual workers postpone retirement by only 9.7 months on average, while non-manual workers postpone by 19.6 months. Thus, the response to benefit reductions is about 50 per cent lower among manual workers compared to non-manual workers. These findings are not only robust against alternative health measures (disability status, number of doctor visits, and self-reported health status) but also against household-related background variables (marital status) and aggregate labor market conditions (monthly unemployment rate). Among

[2]Claiming an old age pension early refers to ages previous to the normal retirement age, which was fixed to age 65 throughout the observation period (1990–2012).

women, however, I do not find a significant difference in the retirement response of worker groups that differ by physical demands of occupations.

While the difference between manual and non-manual workers is significant and considerably large among men, the only nearby explanation other than health is a specific retirement pattern at age 63. Disability pensions are available without benefit reductions at this age and I provide analytical and graphical evidence that indicates a direct link of male manual workers taking this pathway of retirement. Strikingly, manual workers retire significantly more often at age 63 after the introduction of benefit reductions and no other institutional program provides an alternative explanation for this exit route.[3] The rationale behind this behavior is to increase the expected pension wealth by claiming disability pensions at the earliest date without reductions. Thus, delays in retirement are lower among manual workers not only because formerly harsh occupations are correlated to poor health, but also because a specific retirement scheme sets incentives to retire at age 63.

The key finding of differential responses to retirement benefit reductions by manual and non-manual workers is, to the best of my knowledge, a novel one. Distinguishing worker types by physical demands of occupations adds to the literature on health-related aspects of retirement decisions. Existing evidence on German data shows that health is a much stronger predictor for disability retirement than expected pension benefits (Riphahn, 1999). The results that I derive here are largely in line with these findings, showing that the impact of different health measures on retirement timing is much larger than the expected pension wealth. Moreover, previous research has shown that the response to financial incentives among disability pensioners is more likely for individuals in good health (Hanel, 2012). I add to this literature by providing evidence that the retirement probability mass is shifted towards age 63 for male manual workers who are exposed to benefit reductions, indicating that these workers are sufficiently flexible to wait for reduction-free disability pensions. The induced kink in the cumulative distribution function of retirement is strongly suggestive of a large-scale program tradeoff in terms of disability retirement at this age.

The remainder of this paper is structured as follows. Section 2 provides institutional details, the conceptual framework, and an overview on related literature. Section 3 outlines the data, the empirical strategy, and identification issues. Section 4 presents results and Section 5 concludes.

## 2. Conceptual Framework and Related Literature

### 2.1. *Institutional Details*

The German public pension system is organized as an earnings-related pay-as-you-go system. Contributions are mandatory for all employees, who are automatically covered by social security. Public pensions are the single most important

[3]In fact, first eligibility of pensions for older workers with at least 35 contribution years (long-term insured persons) is reached at age 63. However, these pensions are subject to benefit reductions of 7.2 percent at this early retirement age and less attractive than disability pensions for which benefit reductions are zero at age 63 in any case (for more details on eligibility criteria, see Table 1).

TABLE 1
Old Age Pensions and Eligibility, 1990–2012

| Pension Type | Retirement Age: ERA | NRA | Eligibility Criteria: MCT (years) | Other | BRR (%) Min./Max. |
|---|---|---|---|---|---|
| Unemployed | 60 | 65 | 15 | 12 months of UE | 0/18 |
| Women | 60 | 65 | 15 | 10 contribution years after age 40 | 0/18 |
| Long-term insured | 63 | 65 | 35 | – | 0/7.2 |
| Severely disabled | 60 | 63 | 35 | 50% disability status | 0/10.8 |
| Regular pension | – | 65 | 5 | Available at NRA | – |

*Notes*: ERA, early retirement age; NRA, normal retirement age; MCT, minimum contribution time; BRR, benefit reduction rate; UE, unemployment. Eligibility criteria are valid for the observation period 1990–2012.
*Source*: German Social Security Code (SGB VI).

source of old age income in Germany, with a current net replacement rate of 55 percent (all pensions: 76 percent) (OECD, 2013). Pension claims are calculated from earnings points that reflect the relative income position in each contribution year to ensure that claims are proportional to contributions. Throughout the observation period (1990–2012), the average contribution rate was 19.3 percent of gross labor earnings. At that time, old age pensions were generally available at age 60 and eligibility varied by (i) employment status (pension for the unemployed), (ii) gender (women's pension), (iii) contribution time (pension for long-term insured persons), and (iv) health (disability pension[4]). All types of old age pensions and eligibility criteria are summarized in Table 1.

To reduce early retirement incentives and to promote labor supply among older workers, benefit reductions were enacted in 1996,[5] and effectively implemented between January 1997 and December 2004.[6] An adjustment factor reduces pension claims by 0.3 percent for each month of early retirement relative to the normal retirement (NRA) age of 65.[7,8] For a whole year of early retirement, the reduction amounts to 3.6 percent of monthly retirement benefits and the

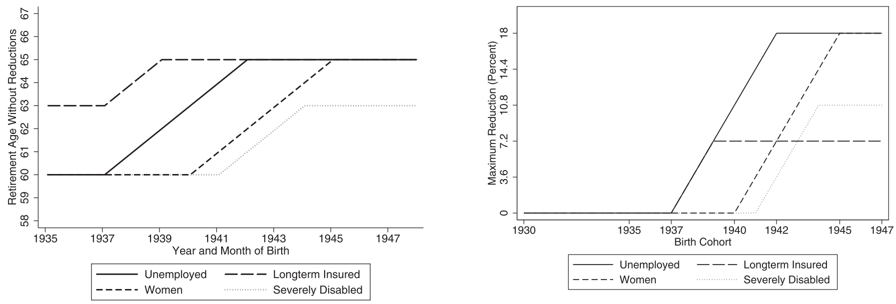
[4]Eligibility for disability pensions is conditional on an assessed degree of disability of at least 50 percent and a minimum of 35 contribution years (German Social Security Code, §37 SGB VI). The degree of disability is a medical indication (German social security code, §2 SGB IX Abs. 2) and its diagnosis conforms to the principles of medical care. The assigned degree of disability is subject to unobserved physician-specific variation that cannot be captured and is thus not fully objective. This includes the (positive) probability that disability status is assigned to people with disability below 50 percent, if desired.

[5]Prior versions of the reform were discussed in the early 1990s and finalized for the unemployed, women and long-term insured in 1996 (corresponding law: Wachstums- und Beschäftigungsförderungsgesetz, 1996) and for the severely disabled in 1997 (corresponding law: Rentenreformgesetz 1999, 1997) with slight changes in 1998 (corresponding law: Korrektur des Rentenreformgesetzes 1999, 1998).

[6]Börsch-Supan and Schnabel (1998, 1999) and Börsch-Supan (2000a) show how, prior to the introduction of benefit reductions, the German pension system imposed strong incentives to retire early.

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

[8]Throughout the observation period (1990–2012), the NRA was fixed at age 65. A current reform gradually raises the NRA from age 65 to 67 between 2012 and 2029, but has no relevance for this study.




(a) Retirement Age Without Reductions

(b) Maximum Reduction Rate

Figure 1. The Rising Treatment Intensity of Benefit Reductions across Birth Cohorts (by Eligibility Type)

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

maximum reduction is 18 percent if early retirement takes place five years previous to the NRA.

Individuals who reach the early retirement age and fulfill the eligibility criteria (see Table 1) can claim an old age pension prior to the NRA, but only at the reduced rate. Figure 1 shows how benefit reductions phased in gradually in monthly steps for the birth cohorts 1937–44 with different timing across eligibility types.

Panel (a) illustrates how the reduction-free retirement age has been raised by pension type, that is, for the unemployed (cohort 1937–41), for the women's pension (cohort 1940–4), for long-term insured persons (cohort 1937–8) and for disability pensions (cohort 1941–3). Panel (b) shows the corresponding maximum reduction rate that applies to each type of old age pension according to year of birth. For example, an unemployed person born in January 1937 who claims an old age pension for the unemployed in the first month of eligibility (after the 60th birthday) exhibits a benefit reduction of 0.3 percent. This is because the reduction-free retirement age for this person has been raised by one month. In contrast, an unemployed person born in January 1942 who also retires just after the 60th birthday faces the maximum reduction rate of 18 percent.

## 2.2. *Expected Pension Wealth and Benefit Reductions*

Structural models explain retirement decisions by individual preferences over consumption and leisure in combination with external incentives that are set by the social security system. Individuals maximize utility over the life cycle and chose to retire exactly when benefits and costs for this decision are balanced (Gordon and Blinder, 1980; Crawford and Lilien, 1981; Gustman and Steinmeier, 1986). The optimal retirement age is determined when a change in the utility from leisure is just offset by the change in utility from consumption. Individuals postpone retirement if a later retirement date corresponds to a utility gain. Benefit reductions as analysed in this paper reduce the expected pension wealth (the present discounted value), *ceteris paribus*, in the case of early retirement and raise the value of the future stream of retirement benefits at higher retirement ages.

To transmit this view into the empirical framework of this study, I compute the expected pension wealth as a discrete sum over all future pension benefits, $B$, for each individual. The expected pension wealth is computed in each month of the relevant age window of retirement between age 60 and 66 (72 months), starting in the month after an individuals' 60th birthday when old age pensions become available. The resulting measure is the expected present discounted value (EPDV),

$$\text{EPDV}_t(R) = \sum_{t=s}^{T} \pi_t(s)\delta^{t-s}B_t(R), \tag{1}$$

as a function of time $t = 1, \ldots, 72$ and the retirement age $R = 1, \ldots, 72$ (measured in months), discounted by the rate $\delta$ to time $t$ and weighted by conditional survival probabilities $\pi_t(s)$ from official mortality tables (Federal Statistical Office, 2012). I assume that individuals discount at a rate of 3 percent per year (0.25 percent for each month) and that individuals do not live beyond age 100 (mortality tables end at that age). Financial incentives from the benefit reduction rate (BRR) are incorporated into equation (1) by supplementing an adjustment factor $(1-\text{BRR}(R))$, which can be written as

$$\text{EPDV}_t(R) = (1-\text{BRR}(R)) \sum_{t=s}^{T} \pi_t(s)\delta^{t-s}B_t(R), \tag{2}$$

where $0 \leq \text{BRR}(R) \leq 0.18$ is the benefit reduction rate as implied by the German social security legislation. Whether benefit reductions apply (i.e. $\text{BRR}(R) > 0$) and to what extent depends on (i) the exact retirement age (R), (ii) the year and month of birth, and (iii) the type of old age pension. Since I cannot distinguish pension types in the data used for this study (the German SOEP, details in Section 3) the assignment of the benefit reduction rate to each individual person–month in the sample is based on assumptions about pension types. These assumptions are derived from eligibility criteria of old age pensions in Germany (see Table 1). First, pensions for the unemployed are assumed only for those persons who have contributed at least 15 years and have unemployment experience. Second, old age pensions for long-term insured are assumed for those persons who contributed for at least 35 years and are only applied after reaching age 63. Third, women's pensions are assumed for women with at least 15 contribution years and, finally, disability pensions are assumed only for those persons who are assigned to disability and have contributed for at least 35 years.[9]

[9]The assumptions are straightforward because they are directly derived from eligibility criteria in Table 1. Subsequent estimates are very stable against assuming equal benefit reductions for all persons, that is, 18 percent at age 60 and then gradually declining in 0.3 percent steps until age 65, phasing in for birth cohort 1937–41 (men) and 1940–4 (women) (see Table S5).

The relevant variable to identify the effect of benefit reductions on the timing of retirement is the percentage difference between the EPDV (equation (1)) and the adjusted EPDV (equation (2)). This difference is the benefit reduction rate and to isolate the BRR from the EPDV I include two separate regressors in all subsequent estimations.[10]

### 2.3. *Related Literature*

Many empirical studies on retirement decisions have pointed out the importance of forward-looking incentive measures (see, e.g. Fields and Mitchell, 1984; Mitchell and Fields, 1984; Samwick, 1998; Coile and Gruber, 2000; Coile *et al*., 2002). In contrast to the wealth level or current income at a given point in time, the forward-looking perspective supports identification of the impact of social security benefits on the retirement age. The present paper relates to this literature because the EPDV is exactly such a measure that discounts the sum of all future pension benefits and takes uncertain lifetimes into account. 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, 2000a, b; Coile *et al*., 2002; Hanel, 2010).[11]

The relationship between financial incentives and retirement timing has been analysed in the German context before. The closest studies are probably those by Börsch-Supan and Schnabel (1999) and Hanel (2010), which examine actuarial adjustments in the German public pension system in different settings. Börsch-Supan and Schnabel (1999) provide an *ex-ante* simulation of the reform that introduced benefit reductions at a rate of 3.6 percent and they report an increase of about six months in the mean retirement age. Hanel (2010) provides an *ex-post* analysis based on social security records from insurance accounts, using benefit reductions as a natural experiment. She finds that benefit reductions induce postponement in benefit claiming by up to 14 months (depending on subgroups). This paper proceeds on these findings in first resembling the principal findings of Hanel (2010) on a different database (survey data) and then showing the differential response of worker types.

Several types of discontinuities have been used to identify retirement responses in the empirical literature. For example, Mastrobuoni (2009) uses

[10]Note that both the EPDV and the BRR depend on the retirement date $R$ independently. The BRR is a function of the retirement age as determined by the rules of the policy change. The EPDV is a function of the retirement age that is determined by total pension claims (i.e. contributions) and the remaining life expectancy. The mechanics of the system usually imply that a higher retirement age coincides with higher monthly benefits (more contributions) that are received for a shorter period (fewer years to live), *ceteris paribus*.

[11]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 utility function (constant relative risk aversion). Estimating a fully structural model is usually circumvented by assuming parameter values for the preferences over risk and leisure (Samwick, 1998; Börsch-Supan, 2000a, b; 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 behavior (for a discussion, see Coile and Gruber, 2000).

cohort discontinuities from social security benefit cuts in the U.S. that are linked to an increase of the normal retirement age. Each increase of the normal retirement age by two months roughly translates into a 1 percentage point reduction of benefits, and both men and women postpone retirement by about 50 percent on average (i.e. by one month if the rise of the normal retirement age is two months). Liebman *et al*. (2009) exploit existing discontinuities that are generated from benefit rules within the U.S. social security system. They find that a 10 percent increase in the net-of-tax share on labor earnings (nominal tax rate minus marginal social security tax) reduces the two-year retirement hazard by 2 percentage points from a base of 15 percent. Recent non-parametric evidence employs an Austrian rule in employer-provided retirement benefits that creates discontinuities in the incentives to delay retirement (Manoli and Weber, 2016). They find a considerable response in terms of increasing labor supply (average participation semi-elasticities between 0.14 and 0.28), but conclude that for large retirement delays one needs disproportionately high financial incentives.

The mentioned types of discontinuities differ very much by institutional details but have proven to be a reasonable source of variation to identify retirement and labor supply responses. The present paper adds to this literature in using cohort discontinuities to identify the effect of benefit reductions on the timing of retirement, with a focus on the heterogeneous response by occupation. Here, discontinuities are induced by a differential treatment intensity from benefit reductions across birth cohorts.

## 3. Data and Empirical Strategy

### 3.1. *Data*

I use survey data from the German Socio-Economic Panel Study (SOEP), including the panel waves 1990–2013.[12] The SOEP is a sample of the German population including about 20,000 individuals in 11,000 households who are repeatedly interviewed over several years (Haisken-DeNew and Frick, 2005). This study is based on the SOEP because it is the only panel dataset covering Germany that combines retrospective employment histories to a rich set of individual- and household-level socio-demographic variables. This combination is important to identify worker heterogeneity in the response to financial incentives, which is the focus of this paper.

Throughout all estimations, I use longitudinal weights (inverse staying probabilities) that are provided by the SOEP to account for panel attrition (for details on longitudinal weights in the SOEP, see Kroh, 2009). These weights match the margins of the panel with respect to the reasons for leaving the sample in each wave. Panel attrition is a specific problem of the application in this paper, where survey respondents are followed over several years to construct individual spells of non-retirement.

[12]Note that the observation period is 1990 to 2012, but the analysis includes retrospective questions from 2013 which correspond to the year 2012.

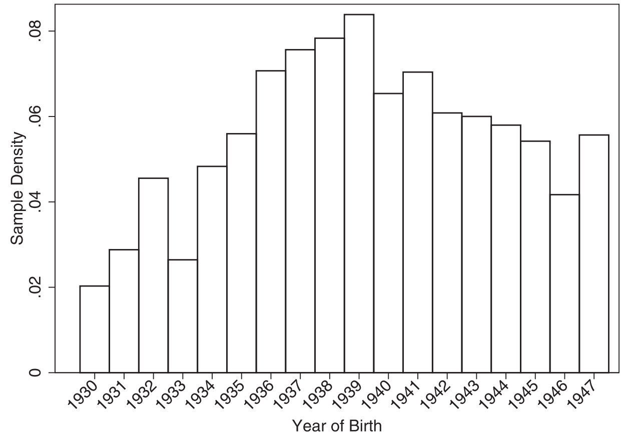



Figure 2. Sample Composition by Year of Birth

*Source*: Own calculation based on the SOEP, 1990–2013.

## 3.2. *Final Sample and Selection Criteria*

The empirical analysis is based on spells of non-retirement until the event of retirement takes place. Retirement is defined as self-reported retirement status from retrospective employment questions. To obtain the final estimation sample, I use the following selection criteria.

First, the sample is restricted to person–month-observations between age 60 and age 66 because old age pensions are available no earlier than age 60.[13] Moreover, the majority of the retirement entries are observed no later than age 66 because the normal retirement age was uniformly fixed at age 65 throughout the observation period.[14] Only 0.4 percent (17 out of 4,350) of the individuals in the estimation sample have right-censored spells due to non-retirement until age 66.

Second, the sample includes individuals from the birth cohorts 1930–47, as shown in Figure 2. The particular choice of these cohorts is based on the timing of the introduction of benefit reductions. They gradually phased in for the birth cohorts 1937–44 and were fully implemented thereafter (1945–ongoing). Birth cohorts previous to the implementation (1930–6) did not face any benefit reductions whatsoever. To obtain a sample that includes all treatment intensities (not affected [1930–6], partially affected [1937–44], fully affected [1945–7]), the birth cohort choice is very much predetermined.

The final estimation sample consists of 4,350 individual spells that correspond to 80,017 person–month observations. Of these, 471 spells (11 percent) are

[13]See Section 2 for details. Note that reduced earnings capacity pensions (“Erwerbsminderungsrenten”) are available before age 60, but this study is solely focused on old age pensions.

[14]Birth cohort 1947 is an exception because for this cohort, the normal retirement age has been raised by one month. The corresponding reform started in the last observation year (2012) and raises the normal retirement age by two years, from 65 to 67, between 2012 and 2029. However, this does not confound the estimates but, rather, implies that the expected reaction will be even stronger in the future. Early retirement is then even more costly due to the greater distance to the normal retirement age (higher benefit reductions at a given early retirement age).

right-censored, which can occur for two different reasons.[15] First, spells are right-censored if no retirement entry is observed within the age limit. This only occurs for few people (17 out of 4,350) and is explicitly modeled in the likelihood function of the subsequent duration model. The second type of right-censoring, however, is caused by panel attrition, which is a specific type of sample selection. This is a potential source of bias, especially if reasons for leaving the sample early are systematically correlated with the retirement age. To overcome this problem, I weight all regressions by inverse staying probabilities that are available for each panel wave.[16]

### 3.3. *Descriptive Statistics*

Descriptive statistics for all estimation variables are summarized in Table 2. The mean retirement age in the full sample is 61.3 years. The corresponding retirement indicator is the dependent variable of the duration model and its mean indicates that, on average, 4.9 percent of the persons at risk leave the sample through retirement in each period, conditional on not having left the sample before.

The right panel of Table 2 shows differences by occupation type. Manual workers retire earlier, have lower expected pension wealth, are more likely to be disabled, are in worse health, include a higher share of male workers, and have fewer years of education. However, their benefit reduction rate at retirement does not significantly differ from that of non-manual workers.[17]

The conditioning variables shown in Table 2 are essential to model retirement decisions. The empirical literature has shown that individual health plays an important role in retirement decisions (see, e.g. Berkovec and Stern, 1991; McGarry, 2004). I include different health measures because previous studies have raised concerns about using subjective and objective health measures in retirement models (e.g. Bound, 1991). Since the SOEP provides both subjective (self-reported health) and objective (number of doctor visits) measures, I check the sensitivity of the results against alternating measures. I also make use of the disability status (yes/no) to show that the disability status is strongly correlated to physical demands of occupations (manual workers are more likely disabled). Second, disability pensions are an important pathway of retirement, which I discuss subsequently. Finally, I use family-related background variables to take marital status into account as a driving factor in retirement decisions (e.g. Blau and Riphahn, 1999).

Descriptive statistics by gender are provided in Table S1. Women have a lower average expected pension wealth and fewer years of education, indicating a

[15]In contrast to right-censoring, there are also about 3 percent of left-censored spells in the final sample. These individuals enter the sample not directly after the 60th birthday but, instead, at higher ages (e.g. if their first SOEP interview is at age 61). All estimates obtained in Section 4 are robust against dropping the left-censored observations (see Table S6).

[16]I also provide unweighted estimates that ignore selection due to panel attrition (Table S7) and baseline estimates that include the predicted (first-stage) attrition probability (Table S8) to account for sample selection from panel attrition. None of the coefficients concerning the main variables (BRR, Manual, BRR X Manual, and EPDV) is sensitive to these changes, substantiating the robustness of the results.

[17]The mean benefit reduction rate (BRR) of 5.6 percent in the full sample is measured at retirement for a total of 3,879 non-censored observations (89 percent of the sample). Including censored spells (no retirement until age 66), the mean reduction rate is slightly lower (5.0 percent).

TABLE 2

Descriptive Statistics

| | Full Sample | | By Occupation Type | | | |
|---|---|---|---|---|---|---|
| | Mean | Min./Max. | Mean | | Difference | $t$-statistic ($p$-value) |
| | | | Non-manual | Manual | | |
| *Dependent Variable* | | | | | | |
| Retirement age | 61.26 | 60/66 | 61.54 | 61.10 | 0.44 | 7.46 (0.000) |
| Retirement (indicator) | 0.049 | 0/1 | 0.038 | 0.062 | 0.024 | 15.98 (0.000) |
| *Main Variables* | | | | | | |
| BRR | 0.056 | 0/18 | 5.8 | 5.5 | 0.3 | 1.22 (0.223) |
| EPDV | 204,460 | 1,225/642,962 | 253,820 | 167,181 | 86,639 | 16.47 (0.000) |
| Manual | 0.528 | 0/1 | – | – | – | – |
| BRR * Manual | 3.1 | 0/18 | – | – | – | – |
| *Health-Related Variables* | | | | | | |
| Disability status | 0.112 | 0/1 | 0.093 | 0.131 | 0.038 | 17.28 (0.000) |
| Annual doctor visits | 11.26 | 0/300 | 9.77 | 13.04 | 3.27 | 26.48 (0.000) |
| Good health | 0.420 | 0/1 | 0.485 | 0.336 | 0.149 | 42.96 (0.000) |
| Moderate health | 0.402 | 0/1 | 0.379 | 0.434 | 0.055 | 15.61 (0.000) |
| Poor health | 0.178 | 0/1 | 0.136 | 0.230 | 0.094 | 35.02 (0.000) |
| *Other Control Variables* | | | | | | |
| Male | 0.537 | 0/1 | 0.519 | 0.552 | 0.033 | 2.18 (0.022) |
| West Germany | 0.750 | 0/1 | 0.792 | 0.713 | 0.079 | 6.09 (0.000) |
| German | 0.905 | 0/1 | 0.947 | 0.867 | 0.08 | 9.07 (0.000) |
| Married | 0.854 | 0/1 | 0.840 | 0.871 | 0.031 | 12.46 (0.000) |
| Separated | 0.013 | 0/1 | 0.017 | 0.009 | 0.008 | 8.92 (0.000) |
| Single | 0.028 | 0/1 | 0.034 | 0.020 | 0.014 | 12.12 (0.000) |
| Divorced | 0.061 | 0/1 | 0.067 | 0.054 | 0.013 | 7.60 (0.000) |
| Widowed | 0.044 | 0/1 | 0.042 | 0.046 | 0.004 | 2.16 (0.031) |
| Years of education | 11.89 | 7/18 | 13.41 | 10.54 | 2.87 | 38.83 (0.000) |
| $N$ (person–months) | 4,350 (80,017) | | 2,303 (44,188) | 2,047 (35,829) | | |

*Notes*: Reported values are from individual spells for time-constant variables (manual, sex, region, nationality, education) and from person–month observations for time-varying variables (health and marital status). Incentive variables and pension wealth (BRR, BRR*Manual, EPDV) are measured in the month of retirement (excluding censored observations). EPDV, expected present discounted value; BRR, benefit reduction rate.

*Source*: Own calculations based on the SOEP, 1990–2013.

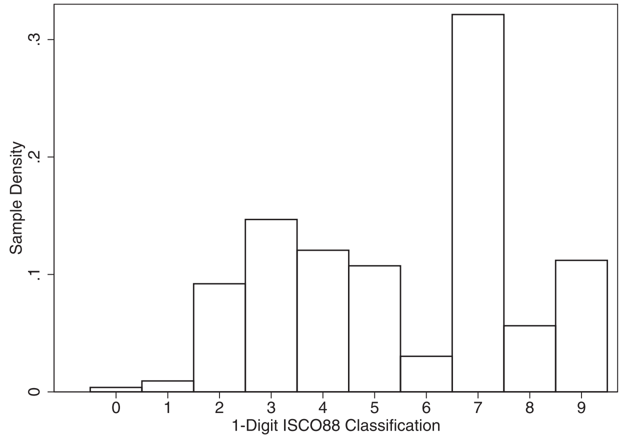



Figure 3. International Standard Classification of Occupations (ISCO 88): One-Digit-Level

*Notes*: The coding is as follows: 0, armed forces; 1, legislators, senior officials, and managers; 2, professionals; 3, technicians and associated professionals; 4, clerks; 5, service and sales workers; 6, agricultural workers) 7, craft and trade workers; 8, machine operators and assemblers; 9, elementary occupations.

*Source*: Own calculation based on the SOEP, 1990–2013.

lower labor force attachment in the observed cohorts. Moreover, the retirement age does not significantly differ between men and women in the overall sample. I subsequently discuss differences in retirement timing between men and women in more detail and show that the initially lower retirement age among men compensates through a stronger delay over time.

### 3.3.1. Physical Demands of Occupations: The Definition of Manual and Non-Manual Workers

The estimation sample consists of 53 percent manual workers, as indicated by the corresponding dummy variable (Manual) that equals one for manual workers and zero for non-manual workers (Table 2). The definition of this variable follows the International Standard Classification of Occupations from 1988 (ISCO 88). The sample distribution of occupations on the one-digit level is shown in Figure 3. Workers in coding groups 5–9 are characterized by physically demanding occupations and thus defined as manual workers. The predominant share of this group consists of craft workers, followed by elementary occupations, service workers, machine operators, and assemblers. In contrast, the coding groups 1–4 are defined to be non-manual workers, including professionals, clerks, legislators, and managers (more details are available in Table S2). I exclude people who work for the armed forces (one-digit code: 0) from the estimation sample because here I cannot distinguish between manual and non-manual workers.

### 3.3.2. Expected Pension Wealth

The expected pension wealth (EPDV) is the present value of all future income streams from pension benefits, taking uncertain lifetimes into account.

The EPDV is a forward-looking measure of pension wealth for each point in time and thus an important determinant of retirement decisions that needs to be included in a model of retirement. To compute the EPDV, I average all available gross annual labor earnings for each individual earnings biography and determine the relative income position of an individual for a given year. This is essential to calculate individual pension entitlements, which are based on the concept of earnings points. A person with average labor earnings in a given year receives one earnings point and a person with twice the average earnings receives two earnings points. When claiming an old age pension, the individual sum of earnings points collected during the entire earnings biography captures the amount of contributions. These are evaluated at the current annuity value to finally obtain individual pension entitlements.[18] For the purpose of interpretation of the regression results, the EPDV is transformed into its real euro value of the last observation year (2012).

The sample mean of the EPDV (equation (1)) and the adjusted EPDV (equation (2)) are plotted in Figure 4. The difference between the mean EPDV and its counterfactual situation (the mean adjusted EPDV) is the hypothetical difference imposed by the benefit reduction.[19] Figure 4 and corresponding values in Table 2 illustrate the incentive structure as faced by the decisionmaker between the two situations of being subject to benefit reductions or not. The adjusted EPDV is initially lower for those who are subject to benefit reductions, but converges to the EPDV towards age 65 when the normal retirement age is reached and no further benefit reductions apply. Another interesting feature of Figure 4 is how the slope of the EPDV changes over the retirement age. After a rather steep increase between age 60 and 62, it declines slightly towards age 63 and eventually flattens out after age 64. This pattern of expected pension wealth over age results from the interplay between the payoff structure (monthly pension benefits) and the life expectancy (remaining years to live).

### 3.4. *Econometric Strategy*

To estimate the effect of benefit reductions on the retirement hazard, I use a discrete time proportional hazards model. Individuals are observed repeatedly from the first month of eligibility (age 60) to a maximum of age 66. The dependent variable is zero for each month of non-retirement, equal to one in the month of retirement and not observed thereafter. Spells are right-censored if individuals leave the sample or do not retire until reaching age 66.

[18]In 2012, the annuity value (German *Aktueller Rentenwert*) of one earnings point was €28.07 in the former West Germany and €24.92 in the former East Germany. The annuity value is set according to a formula of the German Social Security Code (§68 SGB VI), which takes into account wage growth and changes in the share of retirees. For example, a person who has contributed to the public pension system for 40 years at average earnings has collected 40 earnings points. In the former West Germany, this person would receive monthly pension benefits of $40 \times 28.07 =$ €1,122.80.

[19]Departing from the EPDV that reflects individual earnings biographies, the adjusted EPDV is calculated by applying the benefit reduction rate according to the assumptions in Section 2.2. As counterfactual outcome, the adjusted EPDV assumes that benefit reductions are in place holding everything else constant.

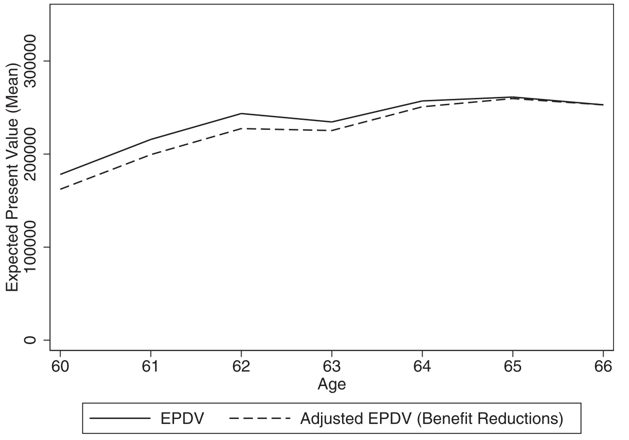



Figure 4. Expected Pension Wealth: Counterfactual Situation

*Notes*: Computed values are weighted at each age using conditional survival probabilities and assuming that individuals never grow older than 100 years (i.e. 480 months from the 60th birthday). Future streams of pension benefits are assumed to be discounted at a rate of 3 percent.

*Source*: Own calculation based on the SOEP, 1990–2013.

Duration models are useful to model transition behavior (see, e.g., Lancaster, 1979; Meyer, 1990). The discrete time duration framework is advantageous in this context as it (i) allows us to control for right-censored spells, (ii) explicitly takes into account the discrete measurement of time in months, and (iii) allows for a large number of transitions at particular points in time. I control for these probability mass points by implementing the most flexible version of a duration model with duration dummies for each point in time, assuming type-I-extreme-value distributed spell lengths (complementary log–log model). The choice of this distribution is motivated by the fact that retirement entries are rare events that accumulate at few duration times. The estimation is based on a sample likelihood function as proposed by Prentice and Gloeckler (1978) and Meyer (1990). The final estimation equation (the log-likelihood function) is as follows:

$$\ln L(\gamma,\beta)=\sum_{i=1}^{N}\left[\delta_i \ln\left[1-\exp\left[-\exp\left(z_i(k_i)'\beta+\gamma(k_i)\right)\right]\right]-\sum_{t=1}^{k_i-1}\exp\left(z_i(t)'\beta+\gamma(t)\right)\right], \tag{3}$$

where $z_i(\cdot)$ is a vector of time-varying explanatory variables (containing the main regressors BRR, EPDV, and all conditioning variables) for individual $i$ and $\beta$ is the vector of corresponding unknown parameters.[20]

[20]In this notation we have that $k_i = min(int(T_i; C_i)$, where $T_i$ is the length of individual $i$'s spell of non-retirement and $C_i$ is the censoring time. $\delta$ is the corresponding censoring indicator, where $\delta_i = 1$ if $T_i \leq C_i$ and $\delta_i=0$ otherwise. The last term including the sum over $t=1$ to $k_i-1$ is the probability that a spell lasts at least until $k_i$. For the full derivation of the log likelihood function, please see Meyer (1990).

### 3.5. *Identification of the Effect of Benefit Reductions on the Retirement Age*

Identification is based on estimating the difference in the outcome (retirement age) from cohort discontinuities in the benefit reduction rate, conditional on a set of observable characteristics. Discontinuities are determined by the date of birth only and in that sense the benefit reduction rate is a source of exogenous variation. The identifying assumption is that the individual choice of the retirement age would not have changed in the treatment group in the absence of the introduction of benefit reductions. The unobserved counterfactual outcome of the treatment group (retirement age without reductions) is assumed to be equal to the observed (factual) outcome of the control group.

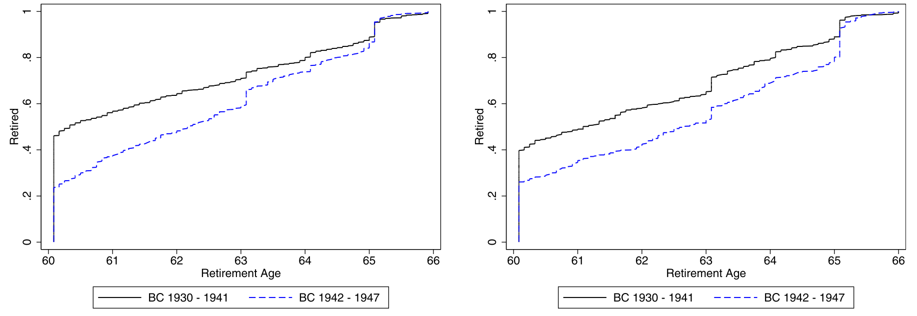
What is important about this assumption is that birth cohorts are not allowed to differ by unobserved factors. To make this assumption as realistic as possible, I keep the window of birth cohorts (1930–47) narrow. This reduces the likelihood of systematic differences from unobserved cohort characteristics that may confound the estimates. To further disentangle the policy effect from secular employment trends, I condition all regressions on the monthly unemployment rate in Germany.

Another potential challenge for identification of the effect is adaptive behavior ahead of time. This is unlikely to confound the estimates, because in the short run (between the announcement in 1996 and implementation starting in 1997) treatment cohorts can only avoid benefit reductions by postponing retirement and this is exactly the difference that I want to measure (i.e. the causal effect). In the long run, related studies on retirement policies (e.g. Mastrobuoni, 2009) point out that people may adapt consumption and savings, and in that case the estimated effect of financial incentives on the retirement age would be smaller. One could think of people who learn about the policy well ahead of time and react by lowering consumption or increase hours worked in order to still be able to retire at the desired early retirement date at reduced benefits. Only including early treatment cohorts (1937–47) does not entirely rule out this behavior, but ensures that long-run adaptations in consumption and savings plans are either small or non-existent.

Shifts in the retirement distribution in Figure 5 illustrate the type of exogenous variation that I use to identify the effect of benefit reductions on the retirement age. Cohorts exposed to a higher treatment intensity retire later irrespective of their subgroup. The cutoff point between cohorts 1941 and 1942 distinguishes retirement profiles for groups with low and high treatment intensity. I define low and high treatment intensity at this particular cutoff point because it divides the sample into the two groups (zero/low vs. high benefit reductions) while keeping the size of the two respective samples sufficiently large.[21]

The kinks at age 60, 63, and 65 are induced by early and normal retirement ages when old age pensions become available, highlighting the importance of social security legislation as a driving force for retirement decisions. Retirement just after the 60th birthday is considerably lower among cohorts

[21]Since this choice is rather arbitrary, I provide similar graphs for other birth cohort cutoffs in Figures S1 and S2. These graphs are in fact very useful to illustrate how the retirement distribution shifts across cohorts and thus treatment intensity.




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

Figure 5. Cumulative Distribution of Retirement by Treatment Intensity [Colour figure can be viewed at wileyonlinelibrary.com]

*Notes*: Treatment intensity is defined at the birth cohort cutoff between 1930–41 (low) and 1942–7 (high). BC, birth cohort.

*Source*: Own calculation based on the SOEP, 1990–2013.

with high treatment intensity for both manual and non-manual workers and the corresponding distribution profiles evolve on a lower level. Among non-manual workers (Figure 5, panel b), the kink at age 63 is induced by claimants of long-term insured persons and this behavior seems relatively unaffected by the introduction of benefit reductions. What is remarkable, however, is the newly induced kink at age 63 among manual workers. This kink strongly suggests that manual workers more often retire at age 63, after the treatment intensity has increased (Figure 5, panel a).

## 4. Results

The estimates I present in this section are intended to show what part of the observed variation in the retirement age is a consequence of benefit reductions. Based on this quasi-experimental setting, I estimate the difference in the retirement age as the causal effect of benefit reductions, conditional on various individual- and household-specific characteristics. Furthermore, I quantify the difference in this response for manual and non-manual workers.

### 4.1. *Baseline Estimation*

Average marginal effects from the complementary log–log model are reported in Table 3. The retirement hazard is estimated separately for men and women because they differ considerably by labor force participation and retirement rules. Specification (1) includes the financial incentive variables (BRR and EPDV) only, but these results are robust against adding the manual job indicator, the interaction term,[22] and further conditioning variables. The interpretation is focused on specification (7), which includes all conditioning variables and the

[22]The interaction term (BRR * Manual), defined as the product of the two respective variables, identifies the differential response to benefit reductions between manual and non-manual workers.

TABLE 3
BASELINE ESTIMATION: BENEFIT REDUCTIONS AND RETIREMENT HAZARD

| | Men | | | | | | |
|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
| BRR | −0.002*** | −0.002*** | −0.002*** | −0.002*** | −0.002*** | −0.002*** | −0.002*** |
| EPDV/20,000 | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** |
| Manual | | 0.021*** | 0.017*** | 0.016*** | 0.021*** | 0.021*** | 0.021*** |
| BRR * Manual | | | 0.001** | 0.001** | 0.001** | 0.001** | 0.001** |
| Disability status | | | | 0.022*** | 0.021*** | | 0.021*** |
| Annual doctor visits | | | | | 0.000* | | 0.000* |
| Moderate health | | | | | | 0.005*** | |
| Poor health | | | | | | 0.010*** | |
| German | | | | | 0.014*** | 0.014*** | 0.013*** |
| Region (West = 1) | | | | | 0.000 | 0.003 | 0.000 |
| Separated | | | | | −0.012 | −0.012 | −0.012 |
| Single | | | | | −0.005 | −0.004 | −0.005 |
| Divorced | | | | | −0.000 | 0.001 | −0.000 |
| Widowed | | | | | 0.003 | 0.002 | 0.003 |
| Years of education | | | | | 0.001*** | 0.001*** | 0.001*** |
| Unemployment rate | | | | | | | −0.001** |
| Duration dummies | + | + | + | + | + | + | + |
| Base retirement hazard | 5.2% | 5.2% | 5.2% | 5.2% | 5.2% | 5.2% | 5.2% |
| *N* | 40,508 | 40,508 | 40,508 | 40,508 | 40,508 | 40,508 | 40,508 |
| | **Women** | | | | | | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
| BRR | −0.002*** | −0.002*** | −0.002*** | −0.002*** | −0.002*** | −0.002*** | −0.002*** |
| EPDV/20,000 | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.000*** | −0.000*** | −0.000*** |
| Manual | | −0.004** | −0.006** | −0.006*** | −0.003 | −0.002 | −0.003 |
| BRR * Manual | | | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| Disability status | | | | 0.012*** | 0.012*** | | 0.012*** |
| Annual doctor visits | | | | | 0.000* | | 0.000* |
| Moderate health | | | | | | 0.003 | |
| Poor health | | | | | | 0.006*** | |
| German | | | | | 0.010*** | 0.010*** | 0.010*** |
| Region (West = 1) | | | | | −0.030*** | −0.029*** | −0.030*** |
| Separated | | | | | 0.007 | 0.007 | 0.007 |
| Single | | | | | 0.016*** | 0.016*** | 0.016*** |
| Divorced | | | | | 0.008** | 0.009*** | 0.008** |
| Widowed | | | | | 0.018*** | 0.018*** | 0.018*** |
| Years of education | | | | | −0.000 | −0.000 | −0.000 |
| Unemployment rate | | | | | | | 0.000 |
| Duration dummies | + | + | + | + | + | + | + |
| Base retirement hazard | 4.5% | 4.5% | 4.5% | 4.5% | 4.5% | 4.5% | 4.5% |
| *N* | 39,509 | 39,509 | 39,509 | 39,509 | 39,509 | 39,509 | 39,509 |

*Notes*: Reported values are average marginal effects from the complementary log–log model. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Estimates are weighted by inverse staying probabilities for each panel wave of the SOEP. All specifications include 71 duration dummies (first month at risk is reference). Married is reference group for marital status; good health is reference group for self-reported health. The base retirement hazard is calculated as the sample mean retirement rate over 72 months at risk, conditional on not having left the sample.

*Source*: Own calculations based on the SOEP, 1990–2013.

objective health measure. All models include duration dummies (i.e. age in months) to control for age-specific ties in retirement that are induced by institutional features, social norms, and further unobserved reasons for retirement such as mutual agreements between employers and employees.

The results for men (upper panel) indicate that raising the benefit reduction rate by 1 percentage point reduces the retirement hazard by 0.2 percentage points from a base of 5.2 percent. Raising the expected pension wealth (EPDV) by €20,000 is associated with a 0.1 percentage point decline of the retirement hazard, indicating that those with higher expected pension wealth (higher previous labor earnings) retire later on average.

Among men, the estimated retirement hazard of manual workers is about 2 percentage points larger compared to non-manual workers from a base of 5.2 percent. This result suggests that manual workers retire substantially earlier than non-manual workers. Further, the positive sign of interaction term also indicates that their response to benefit reductions is significantly lower compared to that of non-manual workers. To quantify this key result from the interplay between the BRR and the manual-job indicator, I provide predictions for the expected retirement age in the next subsection. Men in worse health retire much earlier than those who are in good health and this result is robust against different health measures (disability status, number of annual doctor visits, and self-reported health). Moreover, the retirement decision for men does not correlate significantly with marital status.

In contrast, the retirement behavior of women is much more related to family background (Table 3, lower panel). Single, divorced, and widowed women have a significantly higher retirement hazard than married women (the reference). Among women, the manual versus non-manual worker distinction does not seem to play an important role, either in terms of the base comparison (manual dummy) or their response to benefit reductions. Generally, these results show that women's retirement patterns are much less related to labor market variables such as years of education and the unemployment rate. The finding that women's retirement behavior is rather more determined by family background than by job-related variables points to a low female labor force participation in the observed cohorts.

### 4.2. *The Distinction between Manual and Non-Manual Workers*

The results in Table 4 show estimates for two subsamples, stratified by manual and non-manual workers. A 1 percentage point increase of the BRR reduces the retirement hazard by 0.2 percentage points from a base of 6.2 percent among manual workers (upper panel) and by 0.1 percentage points from a base of 3.8 percent among non-manual workers (lower panel). Considering the base retirement hazard in the two respective groups suggests that, in relative terms, their reaction is about the same size (roughly 3 percent). However, splitting the two groups by disability status shows that manual workers with assigned disability status react to neither benefit reductions nor the EPDV (Table S3, upper panel). A significant reaction to benefit reductions (but not to the EPDV) only remains for manual workers without disability status (Table S3, lower panel). Among non-

TABLE 4

Stratified Estimation: Benefit Reductions and Retirement Hazard of Manual versus Non-manual Workers

| | Manual Workers | | | | | | |
|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
| BRR | −0.002*** | −0.002*** | −0.002*** | −0.002*** | −0.002*** | −0.002*** | −0.002*** |
| EPDV/20,000 | | −0.000** | −0.000*** | −0.001*** | −0.000 | −0.000 | −0.000 |
| Male | | | 0.026*** | 0.026*** | 0.021*** | 0.022*** | 0.021*** |
| Disability status | | | | 0.017*** | 0.020*** | | 0.020*** |
| Annual doctor visits | | | | | −0.000 | | −0.000 |
| Moderate health | | | | | | 0.006*** | |
| Poor health | | | | | | 0.011*** | |
| German | | | | | 0.009*** | 0.010*** | 0.009*** |
| Region (West = 1) | | | | | −0.022*** | −0.019*** | −0.022*** |
| Separated | | | | | −0.002 | −0.001 | −0.002 |
| Single | | | | | 0.003 | 0.004 | 0.003 |
| Divorced | | | | | 0.005 | 0.007 | 0.005 |
| Widowed | | | | | 0.019*** | 0.018*** | 0.019*** |
| Years of education | | | | | 0.003*** | 0.003*** | 0.003*** |
| Unemployment rate | | | | | | | −0.000 |
| Duration dummies | + | + | + | + | + | + | + |
| Base retirement hazard | 6.2% | 6.2% | 6.2% | 6.2% | 6.2% | 6.2% | 6.2% |
| *N* | 35,829 | 35,829 | 35,829 | 35,829 | 35,829 | 35,829 | 35,829 |
| | Non-manual Workers | | | | | | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
| BRR | −0.002*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** |
| EPDV/20,000 | | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** | −0.001*** |
| Male | | | 0.001 | 0.001 | 0.001 | 0.001 | 0.001 |
| Disability status | | | | 0.017*** | 0.016*** | | 0.016*** |
| Annual doctor visits | | | | | 0.000*** | | 0.000*** |
| Moderate health | | | | | | 0.002 | |
| Poor health | | | | | | 0.007*** | |
| German | | | | | 0.009** | 0.008** | 0.009** |
| Region (West = 1) | | | | | −0.007*** | −0.006*** | −0.007*** |
| Separated | | | | | −0.011 | −0.010 | −0.011 |
| Single | | | | | 0.008** | 0.009** | 0.007* |
| Divorced | | | | | 0.005* | 0.006* | 0.005 |
| Widowed | | | | | 0.008** | 0.008** | 0.008** |
| Years of education | | | | | −0.000 | −0.000 | −0.000 |
| Unemployment rate | | | | | | | −0.001** |
| Duration dummies | + | + | + | + | + | + | + |
| Base retirement hazard | 3.8% | 3.8% | 3.8% | 3.8% | 3.8% | 3.8% | 3.8% |
| *N* | 44,188 | 44,188 | 44,188 | 44,188 | 44,188 | 44,188 | 44,188 |

*Notes*: Reported values are average marginal effects from the complementary log–log model. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Estimates are weighted by inverse staying probabilities for each panel wave of the SOEP. All specifications include 71 duration dummies (first month at risk is reference category). Married is reference category for marital status; good health is reference category for self-reported health. The base retirement hazard is calculated as the sample mean retirement rate over 72 months at risk, conditional on not having left the sample.

*Source*: Own calculation based on the SOEP, 1990–2013.

TABLE 5

Further Results on Disability and Retirement of Manual Workers

| | Men | | | |
|---|---|---|---|---|
| | Disability (0/1) | Number of doctor visits | Self-Reported Health (1–3) | Retirement 63 (0/1) |
| Manual | 0.021*** | 1.690*** | 0.256*** | 0.002*** |
| | (0.003) | (0.156) | (0.008) | (0.001) |
| Constant | 0.098*** | 9.006*** | 1.594*** | 0.003*** |
| | (0.002) | (0.088) | (0.004) | (0.000) |
| | 40,508 | 40,508 | 40,508 | 40,508 |
| | **Women** | | | |
| | Disability (0/1) | Number of Doctor Visits | Self-Reported Health (1–3) | Retirement 63 (0/1) |
| Manual | 0.053*** | 4.154*** | 0.201*** | −0.000 |
| | (0.003) | (0.195) | (0.008) | (0.000) |
| Constant | 0.091*** | 10.786*** | 1.723*** | 0.002*** |
| | (0.002) | (0.105) | (0.005) | (0.000) |
| | 39,509 | 39,509 | 39,509 | 39,509 |

*Notes*: Reported values are coefficients from OLS regressions. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Estimates are weighted by inverse staying probabilities for each panel wave of the SOEP.
*Source*: Own calculation based on the SOEP, 1990–2013.

manual workers, both groups (disabled and non-disabled) do react to the EPDV, but only non-manual workers without disability show a significant and relatively large response to benefit reductions (Table S4).[23] Non-manual workers in good health are, in fact, the worker subgroup with the largest response to benefit reductions.

So far, the results have shown that manual and non-manual workers differ in their retirement response to benefit reductions and that disability and health status are key factors in explaining differential retirement responses. The particular challenge is to show the link between these results and the induced retirement kink at age 63 among manual workers. For this purpose, I shed light on the role of the disability status, aiming to show that manual workers are more likely to be disabled (assigned disability status) and that they take the pathway of disability pensions more often after the implementation of benefit reductions.

I do not directly observe disability pensions and thus they cannot be distinguished from other pension types in the SOEP data.[24] However, SOEP respondents are asked about their disability status as a medical indication in each year of the survey and I argue that this measure strongly correlates to the take-up rate of disability pensions. Manual workers are significantly and considerably more likely to be assigned to disability status, as shown in Table 5 (first column). The

[23]Note that for both manual and non-manual workers (Tables S3 and S4), the base retirement hazard differs tremendously by disability status.

[24]Disability retirement benefits were only surveyed in the first two waves (1984–5), which are not part of this study.

difference between manual and non-manual workers is 2 percentage points from a base of about 10 percent for men and is even larger for women (5 percentage points from a base of about 9 percent). The general finding of worse health among manual workers resembles those for other health measures, such as the number of annual doctor visits and self-reported health status (Table 5, columns 2 and 3). For men, the higher prevalence of poor health among manual workers is in accordance with the finding that, significantly, these workers more often retire at age 63 compared to non-manual workers. The probability of retirement at age 63 differs by 0.2 percentage points from a base of 0.3 percent (Table 5, column 4). Although the treatment status does not enter into these results, they do suggest a strong and direct link between the take-up rate of disability pensions at age 63 and the treatment intensity among men. This becomes clear when considering the findings in Table 5, in combination with the graphical evidence on retirement patterns in Figure 5 (panel a): among manual workers under high exposure to benefit reductions (cohorts 1942–7), retirement entries bunch at age 63. However, retirement at 63 is a rare event among manual workers under low exposure to benefit reductions (cohorts 1930–41). Despite the fact that there is no link with retirement at 63 for women (Table 5, column 4, lower panel), this does not imply that the disability pension route is not attractive for them. Rather, it is explained by low eligibility rates for disability pensions due to low labor force attachment of women in the observed cohorts.[25]

As mentioned in Section 2, eligibility for disability pensions is conditional on an assessed degree of disability as a medical indication (at least 50 percent), but this indication is not fully objective. Leaving physicians with a considerable margin of discretion, the system seems flexible enough that people who desire disability status can obtain it to claim the corresponding pension. The pension for long-term insured persons does generally qualify as an alternative explanation, but is financially less attractive because at age 63 it is only available with reductions of 7.2 percent (both pensions require a minimum of 35 contribution years). The bottom line is that those manual workers who want to retire as soon as possible at a zero reduction rate can take the disability pathway, conditional on assignment.

### 4.3. *Predictions*

The expected retirement age by subgroups is reported in Table 6.[26] All predictions are based on the birth cohort cutoff between 1941 and 1942 to divide the sample into a group with low (or zero) treatment intensity (1930–41) and one with high treatment intensity (1942–7).[27]

In the full sample, those individuals who are exposed to a high treatment intensity postpone retirement by 13.2 months on average. This prediction is

[25]Disability pensions require at least 35 contribution years (Table 1). More details on female labor force participation are provided in Section 4.3.

[26]Predictions from the complementary log–log quantify the elapsed duration of non-retirement after first eligibility at age 60, based on the model with the full set of regressors including disability status and number of doctor visits as health indicators (column (7) in Table 3).

[27]This choice is similar to that for cumulative distribution functions in Figure 5.

TABLE 6
Expected Retirement Age by Treatment Intensity (TI)

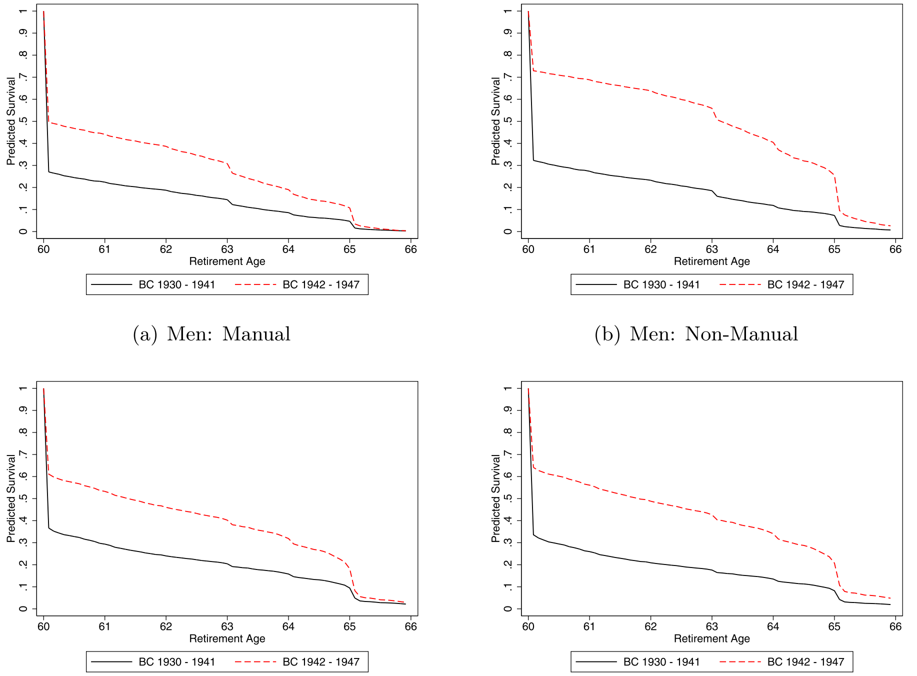
| | Predictions from Baseline Estimation | | |
|---|---|---|---|
| | Low TI | High TI | Difference (months) |
| | Full Sample | | |
| All | 61.45 | 62.55 | 13.2 |
| Manual | 61.38 | 62.15 | 9.2 |
| Non-manual | 61.54 | 62.86 | 15.8 |
| | Men | | |
| All | 61.27 | 62.55 | 15.3 |
| Manual | 61.16 | 61.97 | 9.7 |
| Non-manual | 61.44 | 63.07 | 19.6 |
| | Women | | |
| All | 61.72 | 62.64 | 11.0 |
| Manual | 61.71 | 62.55 | 10.1 |
| Non-manual | 61.74 | 62.70 | 11.5 |

*Notes*: Predicted values are computed from the complementary log–log model. The age distribution is truncated at age 66 (after 72 months at risk). Treatment intensity (TI) is defined at the birth cohort cutoff between 1930–41 (low) and 1942–7 (high).
*Source*: Own calculations based on the SOEP, 1990–2013.

slightly lower than the result from Hanel (2010), who predicts a postponement of 13.8 months on average. Effect heterogeneity is reflected by subgroup predictions, showing that the differential response between manual and non-manual workers is mostly driven by differences among men. Male manual workers postpone retirement by only 9.7 months, which implies that their reaction is some 50 percent lower compared to 19.6 months among male non-manual workers. Among women, the predicted difference between manual and non-manual workers is not only small (about 1.5 months) but also statistically insignificant, as is evident from the regression results in Table 3 (lower panel).

The predictions also suggest a sizable difference for the delay in retirement between men (15.3 months) and women (11.0 months). Differential retirement by gender carries forward throughout all results and coincides with official retirement statistics that document a lower retirement age among men during the 1990s and early 2000s that eventually exceeded the retirement age of women during the observation period (German Federal Pension Insurance, 2015, p. 138).[28] The key explanation for the differential response is the lower female labor force

[28]Examples of average retirement ages (old age pensions) in Germany are as follows: in 1995, men 62.3, women 62.5; in 2005, men 63.1, women 63.2; and in 2010, men 63.8, women 63.3. Note that differentials in the retirement age between men and women vary across the regions (the former East and West Germany). Note also that the definition of retirement in this paper differs from the official statistics, where retirement is defined as claiming pension benefits.




(a) Men: Manual

(b) Men: Non-Manual

(c) Women: Manual

(d) Women: Non-Manual

Figure 6. Predicted Survival Time: Elapsed Duration of Non-Retirement after Age 60 [Colour figure can be viewed at wileyonlinelibrary.com]

*Notes*: Predicted values are computed from the complementary log–log model. Treatment intensity is defined at the birth cohort cutoff between 1930–41 (low) and 1942–7 (high). BC, birth cohort.
*Source*: Own calculation based on the SOEP, 1990–2013.

participation. The baseline estimates have shown that labor-market related variables such as the manual versus non-manual distinction, years of education, and the unemployment rate do play a role in retirement decisions of men but not for women, whose retirement timing is more sensitive to family background. Recent statistics substantiate this finding by showing that women represent a share of 46 percent of employees (subject to social security contributions) in 2010 but, strikingly, only 65 percent of these female employees work full-time, compared to about 94 percent among male employees (Federal Employment Agency, 2012, p. 12).

Predicted survival functions in Figure 6 show how expected retirement evolves across age. The survival curves retrace, once again, that the delay in retirement as a response to benefit reductions differs substantially between manual and non-manual workers. While both groups do react to financial incentives, the predicted survival functions make clear that the response is much larger for non-manual workers. Their survival profiles evolve on a considerably higher level, implying a higher expected retirement age. Most of the difference occurs for changes in initial retirement just after first eligibility at age 60.

What is remarkable about the predicted retirement patterns is the bunching point at age 63 among male workers. As discussed above, at this age, old age pensions due to disability are available without reductions after full implementation (recall Figure 1). Among men, the bunching point is also observable for non-manual workers, but I have provided different sources of evidence to show that manual workers are significantly more likely to choose reduction-free disability pensions at age 63, both in comparison to non-manual workers and to the pre- and early reform period.

## 5. Conclusion

I have estimated the effect of benefit reductions on the timing of retirement for older workers between age 60 and 66. The focus of this study has been on differences in the response of manual and non-manual workers. Identification was based on cohort discontinuities where the intensity of benefit reductions was a function of the date of birth only.

I have shown that, on average, individuals postpone retirement by 13.2 months if pension benefits are reduced by 3.6 percent for each year of early retirement. This result is largely in line with the quasi-experimental retirement literature, which predominantly finds that people respond to incentives if the implicit tax on further periods of work is reduced. Among men, however, manual workers postpone retirement by only 9.7 months on average, and thus their response is about 50 percent lower compared to non-manual workers, who postpone by 19.6 months.

Surprisingly, this does not generally indicate that retirement incomes of manual workers deteriorate. The striking result is that a considerable share of manual workers substitute benefit programs. These workers postpone retirement, but only until disability pensions are available without benefit reductions. This alternative pathway into retirement is financially attractive for manual workers because claiming disability pensions at the earliest date without reductions allows them to increase their expected pension wealth. The lower response to benefit reductions among manual workers is thus not only explained by more physically demanding occupations and worse health. An important part of the explanation is the retirement scheme, which sets incentives to retire without benefit reductions at age 63. Strikingly, this finding suggests that manual workers are more flexible than they first seem to be, and it exemplifies the strong and encompassing impact of institutional details on retirement patterns. The availability of an attractive alternative incentivizes people to take it and causes patterns as revealed in this study.

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## Supporting Information

Additional Supporting Information may be found in the online version of this paper at the publisher’s website:

**Appendix A.** Further Descriptive Statistics
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**Table S1**: Descriptive Statistics by Gender.
**Appendix A.2** Cumulative Distribution of Retirement by Treatment Intensity
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**Figure S2**: Cumulative Distribution of Retirement by Treatment Intensity: Alternative Cut-off Points (Non-Manual Workers).
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