The Effect of Removing Early Retirement on Mortality*

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Abstract

We examine the mortality effect of delaying retirement by exploiting the 1967 Spanish pension reform that exogenously removed the early retirement scheme for individuals that started contributing to the Social Security system after 1967. The reform delayed labour market exit by half a year and increased the probability that individuals take up disability pensions, partial pensions, and no pensions. Delaying exiting employment increases the hazard of dying between the ages of 60 and 69, especially for those employed in low-skilled, physically and psycho-socially demanding jobs. Allowing for flexible retirement schemes mitigates the detrimental effect of delaying retirement on mortality.

JEL Codes: I10, I12, J14, J26
Keywords: Delaying retirement, Mortality, Heterogeneity, Flexible retirement

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1 Introduction

Many countries have reformed their public pension systems to cope with their aging populations and to maintain financial solvency. One of the main policy tools used is to restrict access to early retirement schemes by increasing the minimum pension eligibility age. While there has been an extensive literature studying the labour supply responses to such pension reforms,[1] there are relatively few studies about the impact of delaying retirement on mortality. Moreover, the existing empirical evidence mostly draws lessons from policy experiments that allow for earlier retirement (Coe and Lindeboom, 2008; Hernaes et al., 2013; Hallberg et al., 2015; Bloemen et al., 2017; Kuhn et al., 2020). Because the effects on mortality from preponing and postponing the retirement age are not necessarily symmetric, these estimates might not generalize to today's policy world, where most policymakers aim to incentivize prolonged working lives. Therefore, it is policy-relevant to understand the impact of delaying retirement on mortality, particularly the effect of closing the early retirement options on mortality.

This paper provides novel empirical evidence on this important issue by investigating a Spanish pension reform. This 1967 reform exogenously changed the early retirement age depending on the date individuals started contributing to the Social Security system. Individuals who contributed to the pension system before 1 January 1967 could voluntarily claim a pension as early as 60 years of age. On the other hand, individuals who started contributing after 1967 could only voluntarily claim a pension at age 65.[2]

This reform has several advantages in answering our research question. First, the discontinuity change in retirement age, based on the year the individuals started contributing to the Social Security system, allows us to credibly identify causal effects. Second, in contrast to most of the previous literature, this reform creates a substantial increase (approximately four years) in the early retirement age and leads to a considerable delay in the exit time of the labour market. Third, the reform affects a more general population compared to existing studies (see, e.g., Hallberg et al. (2015); Bloemen et al. (2017); Hagen (2018), all of which study specific subsets of the population, such as military personnel and civil servants). This feature allows us to capture the mortality responses in the general population and examine the heterogeneous responses of subgroups. Lastly, the treatment was determined at the early stage of a worker's career, which provides a long-term horizon for the expected retirement age to impact mortality, if there are some anticipatory responses.

We use a novel version of the Spanish administrative Social Security panel data covering 10% of

[1] For example, see Coile and Gruber (2007), García-Pérez et al. (2013), Atalay and Barrett (2015), Manoli and Weber (2016), Blundell et al. (2016) and Geyer and Welteke (2021) for recent evidence on the direct effects of raising retirement ages.

[2] Individuals of certain cohorts can retire at age 61 through involuntary early retirement under certain conditions. See Section 2 for more details on the institutional setting.

the cohort of individuals born between 1938 and 1949 who are registered with the Social Security system at any point in time until 2020. We compare individuals who started contributing one year before 1 January 1967 with those who started one year after that date. Using within-cohort first-difference regression and controlling for a broad list of fixed effects, we find that the reform delays the age at last employment by around half a year.[3] Those who contributed in 1967 are also less likely to claim a regular pension and more likely to claim partial and disability pensions. This indicates that individuals have utilized other ways to leave the labour market earlier when the early retirement schemes are not available anymore. We also show that they have a higher probability of not claiming any pension, driven mainly by premature mortality. More specifically, individuals who started contributing after 1 January 1967 are 2.5 percentage points more likely to die before claiming any pension. To test the causality of our estimates, we use placebo cut-off dates and find no significant impacts on these placebo dates both before and after 1967.

To show the impact of delaying labour market exit on mortality, we instrument the age at last employment using the year individuals started contributing to the Social Security system. We examine the impact of age at last employment on the hazard of dying between different age brackets. We find that delaying labour market exit by one year increases the hazard of dying between the ages of 60 to 69 by 4.2 percentage points (equivalent to a relative increase of 43%). When we look more closely, we find that the mortality responses are the strongest between ages 60 and 64, when individuals have no longer access to regular pensions. This result indicates that the negative effect of delaying retirement on mortality is driven mainly by the immediate effect of losing access to early retirement schemes. Furthermore, we shed some light on the possible mechanisms behind the detrimental effect of delaying retirement on mortality. In particular, we focus on the heterogeneous effects of delaying the labour market exit by the individuals' labour market conditions prior to retirement. As the parameters of most jobs are multi-dimensional, we examine four dimensions of individuals' labour environments: the physical burden, psychosocial burden, self-value at work, and occupational skill level.

First, using registered workplace accidents at the industry level as a proxy for physical burden, we show that the increase in mortality is stronger for those who have worked in sectors with a very high number of workplace accidents. This finding is consistent with previous literature establishing that physically demanding occupations lead to adverse health effects. We also find that the mortality effect is stronger for individuals in high psychosocial burden jobs (with a high level of mental and social stress). We measure the psychosocial exposure in a job following the Job Exposure Matrices constructed by Kroll (2011). Delaying labour market exit by one year increases the hazard of dying between the ages of 60 and 69 by 5.7 percentage points for individuals in

[3] We show our results are robust using several robustness tests, including regression-based tests of the differences in covariates, and using within-age start contributing fixed effects analysis. For more details see Section 4.3.

high psycho-socially demanding jobs, while this number is 2.8 percentage points for those in low psychosocially burdensome jobs. Furthermore, we measure an individual's sense of achievement and recognition at their last job using the Occupational Information Network data (O*NET). We show that only individuals who work in low self-value industries are more likely to die when facing a one-year delay in the labour market exit. This result indicates that individuals who 'feel recognized' and have a sense of achievement in their work do not experience a negative mortality effect due to a delay in the labour market exit. Lastly, similar to previous literature, we find that delaying labour market exit by one year increases the hazard of dying by 5.4 percentage points for blue-collar workers, while this number is 2.7 percentage points for other workers. The heterogeneous results suggest that advocating for different ages to exit employment, depending on the working conditions of each individual's occupations, can mitigate the detrimental impacts of delaying retirement.

Our findings imply that losing access to early retirement can decrease life expectancy. One proposal to incentivize individuals to stay longer in the labour force without having such a negative impact on their health is to allow them to gradually reduce their working time towards the end of their careers. In Spain, some workers can access partial retirement by working part-time while claiming a partial pension. One of the eligible conditions is to have contributed to the Social Security system for at least for 33 years. Comparing individuals with and without access to partial retirement, we find that individuals who have no access to partial pension experience higher mortality rates when the retirement age is delayed. This finding highlights the importance of providing the opportunity for gradual retirement, which can smooth the adverse effects of delaying retirement.

Apart from contributing to studies on the impact of pension reforms on retirement decisions (e.g., Mastrobuoni (2009); García-Gómez et al. (2012); Manoli and Weber (2016); Geyer and Welteke (2021)), our paper relates to and completes papers studying the mortality effects of retirement. The existing well-identified empirical literature finds mixed results, and explore three types of policy experiments: allowing earlier retirement (Coe and Lindeboom, 2008; Hernaes et al., 2013; Hallberg et al., 2015; Bloemen et al., 2017; Kuhn et al., 2020), promoting later retirement (Zulkarnain and Rutledge, 2018; Hagen, 2018; Bozio et al., 2021) and switching to retirement at the statutory retirement age (Fitzpatrick and Moore, 2018).

The studies of earlier retirement overall find no significant or positive impacts on mortality. For example, Hernaes et al. (2013) find that accessing a pension two to five years earlier has no effect on the probability of dying by the ages of 67, 70, 74, and 77 for the entire population of Norway. Looking at some particular population groups, Hallberg et al. (2015) and Bloemen et al. (2017) find a positive impact of earlier retirement. Hallberg et al. (2015) show that a five-year early access to a retirement pension reduces the mortality of male army officers in Sweden. Bloemen et al. (2017)

find that male civil servants in the Netherlands who are entitled to claim a pension around eight years earlier have a lower mortality rate. The only paper that finds (earlier) access to a pension increases mortality is Kuhn et al. (2020). Using Austrian register data, they estimate the (very) short-term impact of a three-year early access to pension on mortality. Kuhn et al. (2020) find that early retirement increases male deaths before the age of 67.

Evidence on the impacts of later retirement is more scarce. Our paper directly contributes to this literature and is the first paper that provides a precisely estimated impact of later retirement induced by the delay of early retirement age. To the best of our knowledge, only three papers study the effect of delayed retirement. While Hagen (2018) studies the mortality effect of a two-year increase in the statutory retirement age, they find an imprecisely measured no effect on mortality by the age of 69.[4] Bozio et al. (2021) and Saporta-Eksten et al. (2021) show precisely estimated impacts on mortality by exploring pension reforms that change early retirement financial incentives. Bozio et al. (2021) find that delaying retirement in France has a zero effect on the probability of dying between the ages of 61 and 79. Saporta-Eksten et al. (2021) explore an exogenous decrease in the implicit working tax in Israel and show the impact of work on longevity. They find that later retirement increases mortality between the ages of 75 and 85 but that it has no impact on mortality between the ages of 65 and 74. Our paper differs from Bozio et al. (2021) and Saporta-Eksten et al. (2021), as we expect the response to pension reforms that incentivize retirement via financial incentives to be different from reforms that shut down early retirement schemes.

Our paper is the first one to provide empirical evidence that later retirement increases mortality. When we look at the literature on the health impacts of delayed retirement, it is not surprising to find a negative impact of delayed retirement. Many studies on the health impacts of retirement find adverse health effects through increased social isolation and depression (Atalay and Barrett, 2014; Eibich, 2015). Studies also find a positive impact of retirement on health outcomes due to the adoption of a healthier lifestyle (Insler, 2014; Celidoni and Rebba, 2017; Gorry et al., 2018). Therefore, it is reasonable to expect that later retirement might increase mortality rates.

Our findings have important policy implications. First, we show a large heterogeneity in the effect of delayed retirement on mortality, depending on the characteristics of jobs that the individuals held before retirement. Going beyond distinguishing between blue- and white-collar jobs, we show that other job dimensions (such as physical, psycho-social, and self-value) also matter. This finding implies that policies that remove access to early retirement for the general population can exacerbate the socio-economic disparities in life expectancy.

Second, we show the option of a gradual transition to retirement matters with regard to the

[4] Hagen (2018) studies the mortality effect of a two-year increase in the statutory retirement age of local government female workers in Sweden and finds that the reform had no impact on mortality and/or health care utilization. See Section 6.1 for a more detailed comparison with the existing literature.

impacts of retirement on mortality. Allowing older workers to gradually reduce their working time at the end of their careers can mitigate the adverse effects on mortality. Such mitigating effects can be made possible by promoting gradual retirement options. The results also speak to the recent public discussions on flexible retirement. This insight is also relevant for public policy and budgetary considerations, particularly when policymakers in many countries face long-term solvency challenges in both the pension and public healthcare systems.

This article proceeds as follows: Section 2 presents a brief description of the institutional setting in Spain and the 1967 pension reform. Section 3 describes the data and the empirical strategy. Section 4 and 5 present results on the labour supply responses, plus the instrumental variable estimates of the impact of age at last employment on mortality. We also discuss the heterogeneity and potential mechanisms that may explain the impact on mortality. We provide a discussion of the findings in Section 6. Section 7 concludes.

2 Institutional Setting

The current old-pension system for the elderly in Spain is a pay-as-you-go system, with an average replacement rate of around 80% (one of the highest in the European Union). The key elements of the existing Spanish pension system were established in 1967.[5] Prior to 1967, a fixed-amount pension (*SOVI*) financed by employers and the state was available for low-income or disabled workers. The SOVI pension, which was basic and insufficient, was complemented by the Mutual societies (*Mutualidades Laborales*), which were specific to each occupation/sector.

In 1967, the General Social Security Law (*Ley General de Seguridad Social*) unified the pre-existing insurance systems into a single institution, called 'Social Security'. In the new system, further modified by the 1985, 1997, and the 2002 reforms,[6] the statutory retirement age became 65 years of age. Initially, individuals needed a minimum of eight years of contributions to gain access to the pension, which gradually increased to 15 years after the 1997 reform. The pension benefits were calculated based on the average contributions during the 15 years preceding a claim. In addition, full benefits are given to individuals with 35 contribution years. Finally, the penalty for insufficient years of contributions is 2 percent per year.[7]

[5]It was then further developed in the 1970s. In the last four decades, the system has experienced six important reforms, in 1985, 1997, 2002, 2007, 2011, and 2013. See Boldrin et al. (1999), Boldrin et al. (2004) and García-Gómez et al. (2012) for a detailed explanation of all the reforms of the old-age pension system in Spain.

[6]Ley 24/1997, de 15 de julio, de Consolidación y Racionalización del Sistema de Seguridad Social, and Ley 35/2002, de medidas para el establecimiento de un sistema de jubilación gradual y flexible.

[7]It is important to note that in many cases, the claim of a regular retirement pensions is preceded by a period of unemployment that can last for a considerable time. To assist older workers in long unemployment spells, since 1989, those unemployed at age 52 or above who have exhausted their contributive benefits have been allowed to receive unemployment assistance benefits until their pension-claiming age. The only prerequisite is to reach the minimum contribution years to become eligible for an old-age pension. This unemployment assistance paid 75% of

The pension of all the individuals considered in our sample is regulated by the same pension law and face a statutory retirement age of 65 years, with a minimum contribution period of eight years (further modified to 15 years after the 1997 reform). However, individuals from the selected cohort (1938 - 1949) who contributed before 1 January 1967,[8] even by one day, maintained an indefinite right to early retirement from the old-age pension system. These individuals could freely retire early from age 60, though with some financial penalties.[9] Around 13% of the individuals who started contributing in 1966 claimed a regular pension at the age of 60.

On the other hand, individuals from the selected cohorts (1938 - 1949) who contributed after 1 January 1967 faced a statutory retirement age of 65. They can now only retire early under the involuntary early retirement scheme, set in the 2001 law, which allows individuals to retire as early as age 61 (again with some financial penalties, between 6 and 8 percent, depending on the years of contribution, per year of advancement) under certain conditions. These individuals need to have been unemployed (involuntarily) for at least six months and have contributed to the Social Security system for at least 30 years. Due to these stringent requirements, a very small proportion of workers have taken up this involuntary early retirement option.

Because the law was published on 30 December 1966, there is little room left to manipulate the date of the first social security contribution. This feature therefore allows us to compare individuals who started contributing before and after 1 January 1967. As we can see in Figure 1, individuals who contributed before 1967 (independently of their birth year) could voluntarily retire early at the age of 60. For those who contributed after 1967, the only other way to receive early retirement was to claim involuntary early retirement at the age of 61; otherwise, the earliest an individual can voluntarily claim a pension is at the age of 65. Therefore, we expect that individuals who started contributing after 1967 to increase their retirement age considerably.

In addition to the regular retirement pathway, there are two alternative pathways: permanent disability and partial retirement pensions. Permanent disability benefits have been used extensively in Spain as an early retirement mechanism (Boldrin et al., 1999; García-Gómez et al., 2012). This option has thus prompted several reforms since 1985 that have tightened the eligibility criteria in order to maintain a steady level of applications into the disability system henceforth. Nevertheless, disability insurance is an important way by which to exit the labour market. Additionally, from 2002, partial retirement options became available, allowing the combination of income from work with old-age pension benefits. The partial retirement option enables individuals aged 60 years and

the minimum wage. Moreover, a reform in 2002 also created the possibility of combining unemployment insurance claims with labour earnings. Older workers could receive 50% of their unemployment insurance entitlement and work simultaneously, with the employer paying the remaining wages.

[8]The January 1967 deadline was set at a later date for workers in specific sectors, such as construction, mining, fishing, and the railway. We control for these specific deadlines for workers in these sectors.

[9]The penalty for early retirement is 8 percent per year of early claim. After the 2001 reform, the yearly penalty for early retirement was reduced (up to 6 percent per year) as a function of the years contributed.

older, with at least 33 years of contribution and six years of tenure in the same company, to claim 85% pension while working 15% of the time (up to 75% of benefits after the 2011 reform). The partial retirement option requires the firm's agreement because the worker must be replaced with a new employee. In this paper, we investigate the impact of the reform on the age of the individual when claiming disability, partial and regular retirement pensions, and the probabilities of choosing these alternative exit routes from the labour market.

3 Data and Empirical Strategy

3.1 Data and Sample

This paper uses novel administrative data of an extended sample from the Continuous Sample of Working Histories (*Muestra Continua de Vidas Laborales* (MCVL)) provided by the Spanish Social Security system. The dataset contains a 10% random sample of individuals born between 1938 and 1949 who have registered with the Social Security (such as contributive workers and pensioners) at any point of their lives up until 2020.[10]

Therefore, we use a non-publicly available version of the MCVL provided by the Spanish Social Security system, which allows us to observe contributive workers and pensioners prior to 2005 (the starting time of the publicly available version). This data advantage makes it possible to explore a representative sample of workers affiliated with the Spanish Social Security at any point in their working lives and examine their mortality responses.

The MCVL includes time-invariant information, such as gender, birth month, and birth year. It also contains detailed labour market biographies from the date individuals started contributing to the Social Security system, up until their death.[11] Moreover, we observe their employment and unemployment spells, occupations, industry, and monthly contributions. The pension records from the MCVL contain accurate information on an individual's age at the time of claiming a pension, pension benefits, the type of pension they receive at each point in time, and the total number of contributive years before retirement. When individuals exit from the dataset due to death, we observe the exact date of their death, which helps us measure mortality accurately.

Our sample covers Spanish individuals born between 1938 and 1949 who started contributing to the Social Security system 12 months before and after 1 January 1967. We further restrict our sample to individuals still contributing at the age of 50, with at least eight years of contribution.

[10]Note the different sampling criteria our sample has with respect to the publicly available version of the MCVL (see Garcia-Perez (2008)).

[11]Note that the date that individuals started contributing to the Social Security system coincides with the date at which they started their first formal job. It is important to emphasize that, for some individuals, this date does not correspond to the date they started working (for example, for those who switch from the informal sector to the formal sector).

We drop 20% of observations with this restriction. In Table A1, we verify that our sample is not selected. First, we check if the reform has impacted the probability of not being in the main sample due to either having stopped contributing to the Social Security system before the age of 50 or to not having at least eight years of contribution. We find no significant differences. Moreover, we also show no significant mortality differences among individuals not included in the main sample. The final sample contains 25,903 individuals, of whom 27% are female.

To identify the treatment status, we need information about the exact date individuals started contributing. One caveat of the dataset is that the exact date of the first contribution is poorly recorded for some individuals, especially those who started contributing around 1967. We partially correct this measure by using the number of contribution years, which are better documented, as this information is used to calculate the pension benefit.[12] Figure A1 in the Appendix shows the distribution of years the individuals in our sample started contributing, before and after the adjustment. This limitation is the reason we cannot use a Regression Discontinuity Design. In Table A13 in the Appendix, we perform a robustness check for our main results using instead the reported date without making any correction and show that the results are similar.

3.2 Empirical Strategy

Estimating the causal effect of the retirement age on mortality is difficult because many unobserved factors can influence both retirement age and mortality. To deal with the endogeneity in retirement behaviour, we exploit an exogenous variation in retirement age provided by the 1967 Spanish pension reform. We first provide causal estimates of the reform on retirement outcomes using a within-cohort OLS regression with a list of fixed effects and controls. We then use an instrumental variable (IV) approach to estimate the causal impact of age at last employment on mortality.

Within-cohort OLS Strategy

First, we estimate the following equation, where $Treated_{it}$ is a dummy that takes the value of one for individuals who started contributing to the Social Security system in 1967 and zero for those that started contributing in 1966. The treated group can claim regular pensions voluntarily at age 65 (involuntarily at 61), while the control group can claim them as early as 60 years of age.

$$R_{icmt} = \beta_0 + \beta_1 \delta_c + \beta_2 \mu_m + \beta_3 Treated_{it} + \gamma X_{icmt} + U_{icmt} \quad (1)$$

[12] To correct the reported date of the first contribution, we subtract the total number of years contributed from the year they claim a regular pension. If the corrected year of starting contributions is before the reported date of the first contribution, we make this correction. This correction is only possible for individuals who have claimed a regular pension, as only for them is the total number of years contributed reported.

R_{icmt} represents the outcome variable of individual *i* born in year *c* and month *m* who started contributing in year *t* (1966 or 1967). The outcome variables include the age at last employment, the probability, and the age at which individuals claim the different pensions and pension benefits. δ_c is the year of birth, and μ_m is the month of birth fixed effects. β_3 measures the average treatment effect of the reform on the different outcomes.

X_{imct} includes a list of fixed effects, such as highest level of occupation and industry sector between the ages of 30 and 40, and a list of other predetermined covariates, including individuals' mean monthly contribution to the Social Security system, the fraction of days active and employed, and the fraction of time self-employed between the ages of 30 and 40. We cluster the standard errors at the birth year level and report the wild-bootstrap p-values in brackets in all tables.

The estimates from Equation 1 provide us with the reform's reduced-form effects, plus the first stage estimates for the IV regression.

Instrumental Variable Strategy

The causal effect of age at last employment on mortality is estimated by the following equation, where age at last employment (R_{icmt}) is replaced by the predicted value ($\widehat{R_{icmt}}$) from Equation 1:

$$M_{icmt} = \alpha_0 + \alpha_1 \delta_c + \alpha_2 \mu_m + \alpha_3 \widehat{R_{icmt}} + \gamma X_{icmt} + U_{icmt} \tag{2}$$

M_{icmt} represents the probability of dying of individual *i* born in year *c* and month *m* that started contributing in the year *t*. We also include the same list of controls used in Equation 1 (δ_c, μ_m and X_{icmt}). Additionally, we control for the individuals' proxy of their mean pension benefit.[13] The coefficient α_3 captures the local average treatment effect of age at last employment on mortality among individuals who delayed their retirement because they were not able to claim a regular pension at age 60 (compliers). In Section 5, we discuss who the compliers are in our estimation.

Assumptions

The critical assumption for the treatment status to be a valid instrument for access to early retirement is that the year individuals started contributing to the Social Security system is independent of unobserved characteristics that affect the age at last employment and mortality. The following steps support the validity of our instrument.

First, we restrict our sample to those who started contributing in 1966 and 1967. The treated

[13]We do not have information on pension benefits for individuals who have never claimed a pension of any kind. Therefore, for all individuals in our sample, we construct a proxy of the mean pension benefit using monthly contributions and years of contribution (or years of employment and unemployment) using the Social Security formula to calculate pension benefits. The correlation between this proxy and the actual mean pension benefit is 0.93 for individuals who claim a regular pension, indicating that it is a good proxy. Moreover, in Table A9, we show the effect of the reform on this proxy.

and control group individuals had similar labour market conditions when they began working: they were born in the same year and started working only one year apart. Second, we include their highest occupation, industry, birth year, and month of birth fixed effects, which allows us to estimate variations within occupation, industry, and birth year.

Furthermore, we check whether the characteristics of the treated and control groups are similar when they are between 30 and 40 years old.[14] Table A2 shows the impact of the treatment on a list of predetermined variables, including the fraction of time spent in employment, activity, and self-employment between the ages of 30 and 40; the probability of working in a blue-collar occupation and industry sectors; and average monthly contributions between the ages of 30 and 40. The estimates are obtained from estimating Equation 1. Except for the fraction of time spent in self-employment, there are no significant impacts.[15] This suggests that there is no manipulation of the treatment status and that our control and treatment groups are very similar.

To further establish causality of the first stage estimates, we perform placebo tests using other years to define treatment status and a robustness test using age at first contribution fixed effect instead of birth month fixed effects. These tests rule out the possibility that other confounding factors drive our reduced-form estimates. For more details, see Section 4.

To fulfil the exclusion restriction, we need to ensure that the treatment status only affects mortality through its impact on age at last employment. The only possible alternative channels through which the year individuals started contributing can affect mortality are changes in labour market outcomes close to retirement and changes in pension benefits. We always control for the proxy of the mean of the pension base to wash out any possible income effect. In Table A16, we also show that controlling for the labour market decisions before retirement (between 45 and 55 years of age) does not affect our IV estimates.

Finally, the monotonicity assumption requires that contributing to the Social Security system in 1967 instead of 1966 would not lead to earlier retirement. The monotonicity assumption is also satisfied in our context because the treated individuals do not have the option to retire as early as 60 years of age.

[14] Ideally, we would like to check whether the characteristics of individuals in the treatment and control groups differ at the beginning of their career (before the age of 30). However, the data quality was not particularly good when our individuals were that young so the labour market characteristics during the first years of their career might be wrongly recorded for some individuals. We, therefore, look at their characteristics between the ages of 30 and 40.

[15] In Table A12, we show that the effect of the reform on our main outcomes is robust to excluding from our sample individuals in one of the self-employed pension regimes.

4 The Reform Effect on Retirement Outcomes

4.1 Descriptive Evidence

Table A3 provides summary statistics for the main outcomes used in our analysis. There are three different pensions that individuals can claim. Table A3 shows that 47% of individuals claim a regular pension (old-age pension), while 33% claim a disability pension, and 4% of individuals choose a partial pension. Some individuals in our sample never claim any pension due to reasons such as a period of prolonged inactivity ($\sim$ 6%), dying before a claim can be made ($\sim$ 8%), and still being active in the labour market in 2020 ($\sim$ 0.3%). Figure A2 compares the share of different types of pensions by treatment status. Compared with those who started contributing in 1966 (control, light green bars), individuals who started contributing in 1967 (treated, darker green bars) have a lower likelihood of claiming a regular pension and are more likely to claim a disability pension, a partial pension, or claim no pension.

On average, individuals leave the labour market at 59.6 years old and claim regular pensions at 63.6 years of age. Figure 2 shows the age distribution at last employment for individuals who started contributing in 1966 and 1967. As expected, we see a distinct difference. Figure 2 shows that around 8% of individuals who started contributing in 1966 (control group, solid red line) leave the labour market at the age of 60, while this percentage is almost zero for those individuals that started contributing in 1967 (treated group, green dashed line). More than 22% of the treated individuals exit the labour market at the age of 65, while this number is only 17% for the control group. We see the same pattern regarding the age of claiming a regular pension. Figure 3 a) shows that 28% of individuals who started contributing in 1966 (control group, solid red line) claim a regular pension at the age of 60, and 32% of them claim at the age of 65. We also see some claims at the ages of 61 to 64. However, for those individuals who started contributing in 1967 (treated group, green dashed line), almost no one claims a regular pension at any age other than 65 years, whilst almost 70% claim a regular pension at 65 years of age. These figures provide visual evidence that the reform is binding, and that individuals affected by it delayed their retirement.

In our sample, individuals, on average, claim a disability pension at the age of 57 and a partial pension at 61 years of age. Figure 3 b) and Figure 3 c) show that the distribution of these ages by treatment status. We observe that individuals who started contributing in 1967 (green dashed line) claim more disability insurance between the ages of 60 and 65 than those who started contributing in 1966. Moreover, individuals who started contributing in 1967 (green dashed line) claimed partial pensions at slightly earlier ages.

Finally, regarding the mortality measure, conditional on being alive at the age of 50, 32% of our sample died between the ages of 50 and 86. The hazard rate of dying between the ages of 50 and 59 years and the hazard rate of dying between the ages of 80 and 86 are low, at 7% and 2%,

respectively. The highest mortality occurs between 60 and 79 years of age. The hazard of dying between the ages of 60 and 69 is 11%, and the hazard of dying between the ages of 70 and 79 is 16%.

4.2 Regression Results

Table 1 examines the impact of the reform on the different types of pensions that individuals have claimed. We find that individuals who started contributing to the Social Security in 1967 are less likely to claim a regular pension by 10 percentage points ($\sim$ 19%), yet their probability of claiming a disability insurance increases by 5.8 percentage points ($\sim$ 19%). In Table A4, we further show that the reform equally impacted the probability of claiming a severe or absolute disability and a partial or professional disability pension (by a 3.1 and 2.7 percentage point increase, respectively).[16] They are also 1.9 percentage points ($\sim$ 54 %) more likely to claim a partial pension. These results indicate that individuals did not fully comply with the rise in statutory retirement age and have utilized other ways to leave the labour market before claiming a regular pension, by either claiming disability insurance or a partial pension.

We also observe that individuals contributing to the pension system in 1967 are 2.5 percentage points ($\sim$ 18 %) more likely to leave the labour market without any pension. In Table A5 in the Appendix, we further explore three reasons why individuals might not claim any pension: first, they were still working in 2020; second, they became inactive; third, they died before claiming any pension. Table A5 indicates that the reform has only a minimal impact (an increase of 0.2 percentage points) on the probability of continuing to work up until 2020 and has no impact at all on the probability of becoming inactive. Interestingly, individuals who started contributing in 1967 have a 1.6 percentage point ($\sim$ 22%) higher probability of dying before claiming any pension. This finding implies that premature death is the main driver for not claiming any pension. We further explore this effect in Section 5.

Table 2 examines the impact of the reform on the ages at which individuals leave the labour market and claim different types of pensions. The 1967 reform resulted in the treated individuals delaying their labour market exit by almost half a year[17] and delaying claiming their first pension

[16]There are four types of disability pensions. First, *partial disability* pensions are for individuals who have seen their functional capacity reduced by at least 33 percent. These individuals can continue working, even in the same jobs they had before applying for the pension. Second, *professional disability* is assigned to those workers who cannot resume their work activity but could carry out a different occupation. Third, *absolute pensions* are thought for individuals who cannot carry out any type of work due to physical or mental deterioration. Finally, *severe disability* occurs when the worker needs the support of another person to carry out their daily subsistence tasks.

[17]Table A6 in the Appendix shows the reform's effect on the probability of exiting the labour market in different age brackets. The reform decreases the probability of leaving the labour market between the ages of 55 to 63. As expected, the reform most significantly decreased the most the probability of exiting the labour market at 60 (by 4.1 percentage points or 37%). Individuals who started contributing in 1967 also have a higher probability of exiting the

(regardless of the type) by 0.28 years (four months). We find that early retirement is reduced by 1 year and four months for individuals who claim a regular pension.[18] The ages at claiming a disability pension and a partial pension are also affected. Individuals who contributed after 1967 delay claiming disability by around four months but anticipate claiming a partial pension by around two months. Table A8 in the Appendix shows that the reform significantly increased the probability of claiming a disability pension between the ages of 60 and 65 only. This result suggests that individuals affected by the reform use disability pensions as an early retirement scheme between the ages of 60 and 65, ages at which these individuals would have been able to retire with a regular pension if they had contributed in 1966. Moreover, it also indicates that the reform does not capture differential ex-ante health conditions of individuals. If this were the case, the reform should have significantly increased disability pensions before the age of 60.

In Table 3, we examine the reform impact on the pension benefit amount. We expect the pension benefits to be affected because the reform incentivizes individuals to work longer (as shown in Table 2), which increases the pension base and decreases the penalty for early retirement. Moreover, as more individuals claim disability insurance due to the reform, we expect the overall pension benefits to be lower as disability pension benefits are typically less generous. We find that the total pension benefit of individuals who started contributing in 1967 increased by 43€ ($\sim$ 4%). The increase in the pension benefit is driven by an increment in the base pension (without any financial adjustments) of 19€ ($\sim$ 1.6%) and an increase in the pension adjustment (due to later claiming) of 5 percentage points ($\sim$ 6%).[19]

4.3 Robustness and Placebo Tests

In this section, we perform several robustness checks on the reduced form effects of the reform. Moreover, we test the causality of our estimates by using placebo cut-off dates from both before and after 1967.

labour market after the age of 64. Again, the reform most significantly increased the probability of exiting the labour market at the age of 65 (by 7 percentage points or 42%).

[18] We observe in Table A7 that the reform decreased by 10 percentage points ($\sim$ 67 %) the probability of claiming a regular pension at age 60 and between 2 and 3 percentage points ($\sim$ 48 % to 76%) the probability of claiming it between the ages of 61 and 64. On the other hand, the reform increased the probability of claiming a regular pension at age 65 by 9.5 percentage points ($\sim$ 50 %) and 3.6 percentage points after the age of 65 ($\sim$ 32 %).

[19] It is important to note that the positive effects on pension benefits that we observe for all the sample where individuals claimed any pension is driven mainly by those individuals who claimed a regular pension, as Table A9 shows. In particular, individuals who claimed a regular pension and started contributing in 1967 received, on average, 73€ higher monthly pension benefit, which is driven by an increase of 31€ of the pension base and an 8.5 percentage point in the pension adjustments. We can also observe that the mean monthly pension benefit decreases by 26€ for affected individuals claiming a disability pension, while partial pension benefits are not significantly affected by the reform.

Within-Age at First Contribution Fixed Effects Model

The baseline analysis compares individuals born in the same year (along with their highest occupation level and industry sector fixed effects), who started contributing to the system one year apart (1966 vs. 1967). One potential confounding factor of this specification is the age at which individuals started contributing. These individuals were born in the same year but started contributing in 1966 and 1967 and were at different ages when they started contributing. One reason for starting at different ages could be differences in educational attainment. Unfortunately, we do not have information on the education level of individuals in our database. Therefore, to test that the reform is not capturing differences in educational attainment, we use age at first contribution fixed effect instead of birth year fixed effects in Table A10. This robustness check estimates the impact of losing access to early retirement for individuals who start working at the same age but were born one year apart. These estimates should be similar to the main estimates unless the different starting age is a confounding factor. Compared with the baseline results in Tables 1 and 2, the magnitudes of the estimates in Table A10 are very similar.

Cohorts Born between 1941 and 1949

In the baseline sample, we consider individuals born between 1939 and 1949. A law in 2002 introduced the possibility of retiring early via the involuntary pathway. As a result, cohorts born from 1938 to 1940 can claim a pension at the ages of 64, 63, and 62, respectively, while cohorts born after 1941 can claim at the age of 61 (see Figure 1).

Therefore, we perform a robustness check, dropping the cohorts that were only partially affected by the law of 2002. Table A11 reports the main reduced form results for the cohorts born between 1941 and 1949. If we compare them with the baseline results in Tables 1 and 2, we can see that the magnitude of the estimates is quite similar. These results demonstrate that our reduced form effects are not driven by the older cohort of individuals with later access to involuntary early retirement.

Dropping Self-Employed Individuals

Table A2 shows the impact of the treatment on a list of labour market variables when the individuals were between 30 and 40 years old. Except for the fraction of time spent in self-employment, we do not find significant impacts, suggesting that there is no manipulation of the treatment status. A potential reason for finding significant effects on individuals' fraction of time spent in self-employment is that self-employed individuals might have more flexibility in deciding when they want to start contributing to the Social Security system. In this robustness check, we want to ensure that our main baseline results are not driven by these individuals.

Therefore, we perform a robustness check dropping those individuals who received a pension under the self-employed regime (see Table A.14). If we compare them with the baseline results in

Tables 1 and 2, we can see that the magnitude of the estimates is quite similar, indicating that our baseline reduced form effects are not driven by those individuals who were self-employed.

No Correction for the Starting Year of Contribution

As mentioned earlier, one caveat of the Spanish administrative Social Security database that we are using is that the exact date of the first contribution is poorly recorded for some individuals, especially those who started contributing around 1967. In fact, Figure A1 a) in the Appendix shows that in the distribution of the start year of contribution, there is some bunching around 1965, 1966, and 1967. Fortunately, we have excellent information for many individuals about the number of years of contribution and the exact date they claimed a pension. Therefore, we correct the reported date of the first contribution by subtracting the total number of years of contribution from the date they claimed a pension for those who reported having started contributing in 1965, 1966, and 1967. If the corrected year of starting contributions is before the reported date of the first contribution, we make this correction. Figure A1 b) shows how the distribution of the starting year of contribution looks after we make this correction, and we see that the bunching has been dramatically reduced.

In Table A13, we perform a robustness check for our main reduced form results using the reported date of starting contributions without making any correction. We show that, if anything, the results are slightly stronger than in our main specification. First, individuals in the uncorrected sample seem to react more to the reform by using different ways to leave the labour market before claiming a regular pension. The reform increased by 7.5 percentage points (instead of 1.9) the probability that the individuals claimed a partial pension, and by almost the same amount (7.4 percentage points instead of 5.8) the probability of claiming a disability pension. They also delayed leaving the labour market by 0.67 years instead of 0.4 years and by 1.43 years instead of 1.31 years for individuals who claimed a regular pension.

Placebos

A concern for causality is that our results could be potentially biased by unobserved characteristics that affect both the date of starting contributions and labour supply decisions. To test this possibility, we perform several placebo tests where we assign placebo treatment status to the individuals using other dates at first contribution. Figure 4 plots the estimated coefficients of the different placebos, where the treatment is defined as having started making contributions the year after the one indicated in the y-axis (from 1959 to 1976).[20] The placebo estimates are labelled in black,

[20] We do not perform the placebo test on years that are too close to the actual treatment years (1965, 1966, and 1968), given that the start year of contribution is not well defined around that time, and we use the years of contribution to correct for it.

while our baseline estimates are labelled in red. We can see that almost all placebo estimates are close to zero. This suggests that the estimated changes in our baseline analysis result from the exogenous increase in early retirement age rather than from other confounding factors.

5 Right to Retire Early and Mortality

5.1 The Effect of Age at Last Employment on Mortality

In this section, we examine the impact of retiring later in life on mortality using the instrumental variable method. Table 4 reports the effects of age at last employment on mortality at different age brackets (conditional on having survived until that age). Panel 1 reports the simple OLS estimation where we regress age at last employment over mortality. For all our estimations, delaying retirement is negatively and significantly correlated with mortality. This correlation likely captures the fact that less healthy workers tend to retire early. Panel 2 shows the reduced form effect of the reform on mortality. We find that individuals who contributed in 1967 have a 3 percentage point ($\sim$ 10%) higher probability of dying between the ages of 50 and 86. When we examine the reform's impact on mortality at different age brackets, we observe that the increase in mortality is concentrated between the ages of 60 and 69. In particular, individuals who contributed in 1967 die between those ages (that is, ages 60 and 69) with a 2.1 percentage points higher probability ($\sim$ 22%). We also find a minor increase in mortality after 80 (0.4 percentage points). Figure 5 shows that all the placebo estimates for overall mortality after the age of 50 and between the ages of 60 and 69 are insignificant and close to zero. This confirms that our reduced-form estimates result from the exogenous increase in early retirement age due to the reform rather than from other confounding factors.

The IV estimates in Panel 3 of Table 4 indicate that delaying the age at last employment by one year increases the probability of dying between the ages of 50 and 86 by 7.4 percentage points ($\sim$ 24%), 4.2 percentage points ($\sim$ 43%) between the ages of 60 and 69, and 0.6 percentage points ($\sim$ 26%) after 80. In Table A14, we also report the effect of age at last employment on mortality in five-year age brackets. We observe that the mortality responses are the strongest between ages when public pensions are not accessible (between the ages of 60 and 64). Delaying leaving the labour market by one year increases mortality in that age bracket by 3.8 percentage points ($\sim$ 67%). This result indicates that the negative effect of delaying retirement on mortality is driven mainly by the short-term effect of losing access to early retirement schemes.

In Column 6 of Table 4, we also examine the effect of delaying retirement on age at death (in years). This measure will capture both the extensive margin (the effect of delaying retirement on premature death) and the intensive margin (the length of life). We censor the age of death at 71

years old for those individuals still alive at that age (as the younger cohort, born in 1949, are 71 years old at the end of our database in 2020). We find that delaying the age at last employment by one year reduces individuals' age at death by 0.75 years.

It is important to note that the F-statistic of the first stage regression for all our IV estimates in Table 4 are above the rule-of-thumb threshold of 10. Our instrument (the year individuals started contributing) is relevant and correlated with the endogenous variable we are instrumenting (age at last employment).

When we compare the OLS results with the IV estimates, we can see that the IV strategy does a good job controlling for the negative bias present in the correlation between age of last employment and mortality. Moreover, compared with the reduced-form estimates, the IV results are more than double. This is consistent with the almost half a year increase in age at last employment (as estimated in Table 2). In addition, the IV estimates control for other effects of the reform that could potentially impact mortality through different channels. In particular, in the IV strategy, we control for the positive effect of the reform on pension income. As we do not have information about the potential pension benefits for those individuals who never claimed any pension, we control for a proxy of the monthly pension base. We constructed this proxy using the individual's history of monthly contributions and the formula used by the Social Security system to calculate the pension base.[21] Another potential channel through which the reform could be affecting mortality is the labour market outcomes of individuals before retirement (between the ages of 45 and 55).[22] Table A16 shows the IV estimates of the effect of age at last employment on mortality between the ages of 60 and 69 with different control variables. We observe that adding the labour market outcomes of individuals before retirement as controls in the IV estimation does not change the estimates to any great extent. Thus, our baseline IV estimations will only control for the proxy of the pension base.

Finally, it is important to understand who the compliers are. In our specification, the compliers are those individuals who, before the reform, when they had the chance, claimed a regular pension early at the age of 60. Table A17 compares the labour market conditions at the end of the working careers of compliers and non-compliers.[23] Not surprisingly, compliers have less attach-

[21]The correlation between this proxy and the actual mean pension base is 0.9 for the whole sample and 0.93 for individuals who claim a regular pension, indicating that it is a good proxy. Moreover, in Table 3, we show that the reform impacted this proxy similarly to the actual pension base.

[22]Table A15 shows that the reform had an impact on the labour market outcomes of individuals between the ages of 45 and 55. We observe that individuals that started contributing in 1967 spent 1.85 percent more time employed during these years, while they spent 0.67 percent less time self-employed. We also find that the individuals affected by the reform have a 1.5 percentage point higher probability of having a blue-collar occupation. They also have a 1.9 percentage point lower probability of working in the trade or transportation sector, 2.7 percentage points of working in the public, health, or education sectors, and 0.8 percentage points of working in the services, hotel, and housekeeping sectors.

[23]For this comparison, we only consider individuals who were not affected by the reform (who started contributing

ment to the labour market at the end of their working career (between the ages of 45 and 55), have a lower probability of employment in a white-collar occupation, and work in industries with more workplace accident incidences and higher psychosocially burdensome environment. This result suggests that, before the reform, those claiming retirement earlier worked in occupations and sectors with a higher health burden.

5.2 Mechanisms

This section attempts to shed light on some of the potential mechanisms explaining why losing access to early retirement increases mortality. We focus on two types of heterogeneities: labour market conditions before retirement and the possibility of flexible retirement.

Labour Market Conditions Prior to Retirement

Delaying retirement can have very different effects on an individual's life expectancy, depending on the working conditions experienced by the individuals during their last years of employment (Mazzonna and Peracchi, 2017).[24] In this paper, we acknowledge that the burden of a job may be multi-dimensional, and we examine four characteristics of the individuals' labour environment before retirement: first, physical burden; second, psychosocial burden; third, self-value at work; and last, the skill level of their last occupation before retirement. Table 5 reports the heterogeneity results for the probability of dying between the ages of 60 and 69 (conditional on surviving to age 60)[25] by all four measures. In the first panel, we report the reform's effect on the age at last employment for each subgroup, which will be the first stage of the IV estimation. In the second and third panels, we report the effect of delaying retirement on mortality between the ages of 60 and 69. First, we report the reduced-form effect of the reform and then the IV estimates, which capture the effect of delaying by one year the exit of the labour market on mortality.

to the Social Security in 1966). The difference between the two groups of individuals is estimated by running a regression on a dummy variable for compliers while controlling for gender, month, and year of birth fixed effects.

[24] One of the reasons we expect to see heterogeneity in mortality by labour market conditions is because harsher working conditions are more likely to trigger mortality due to specific causes, which are predominant during the ages of 60 to 69. For instance, the medical literature has long established that circulatory system diseases can often be correlated to work-related stress (Kivimäki et al., 2002). In fact, both Bloemen et al. (2017) and Hallberg et al. (2015) report that retirement reduces the risk of heart-related mortality. For the cohort considered in our sample, circulatory diseases are the second cause of mortality (after tumours) between the ages of 60 and 69. Moreover, the third cause of mortality for our cohort of individuals is due to respiratory diseases. Important risks for respiratory conditions include smoking and lack of physical activity (Godtfredsen et al., 2008; Lee et al., 1999). Both factors can be affected by working status and, ultimately, retirement (Falba et al., 2005; Black et al., 2015; Evenson et al., 2002; Barnett et al., 2014). Fitzpatrick and Moore (2018) find that mortality due to two lung-related conditions (COPD and lung cancer) statistically increases immediately after retirement at the age of 62.

[25] As the reform has no impact on the probability of dying before age 60, the sample used in this regression is not selected.

Physical and Psychosocial Burden

Retirement enables individuals to enjoy more leisure time and eliminates work-related stress and exposure to job-specific accidents, potentially positively impacting individuals' mental and physical health, and their wellbeing. Thus, retirement may be particularly beneficial for those who work in strenuous occupations, either physically or mentally. Indeed, labour unions have used this argument heavily in their opposition to increases in the statutory retirement age. Therefore, we first classify individuals' last industry depending on their physical and psychosocial burden in order to analyse if the adverse effects of delaying retirement on mortality differ by these characteristics.

Previous literature has already established that physically demanding occupations lead to adverse health effects (see Case et al. (2005) and Ravesteijn et al. (2013) for a summary). To measure physical burden at work, we use the Spanish Register of Workplace Accidents between 2003 and 2019, which has information on the total number of workplace accidents that individuals in our sample (cohorts born between 1938 and 1949) experience in different industry sectors. Figure A3 shows the distribution of industry sectors depending on their incidence of workplace accidents. We link individuals' last industry to this aggregate industry-level data and divide our sample by the median of the workplace incidence. After this division, the manufacturing, energy, water, sanitation, and construction sectors are considered to have a high incidence of workplace accidents, and the rest are included in the low incidence group.

Columns 1 and 2 of Table 5 show that the increase in mortality seems stronger for those individuals who worked in sectors with a higher incidence of workplace accidents before retirement. Delaying the age at last employment by one year increases the probability of dying between the ages of 60 and 69 by 6.8 percentage points ($\sim 57\%$) in sectors with a high incidence of workplace accidents. At the same time, the effect is only 2.5 percentage points in sectors with a low incidence of workplace accidents. However, the p-value of the difference between these two groups is 0.13. We further divide our sample into three groups (see Table A18) and find that the impact of delaying retirement is significantly different between the highest and lowest groups. In particular, we find no effect on mortality for individuals working in sectors with a very low incidence of workplace accidents, while mortality increases by 6.8 percentage points for those in sectors with a very high incidence of workplace accidents. This heterogeneity confirms that individuals in more physically demanding jobs will benefit the most from having access to early retirement.

Next, we examine the heterogeneity effect of delaying retirement on mortality by the mental and social stress that individuals have experienced before retirement. Unfortunately, we do not have a good measure of occupations or industries by this measure for the Spanish context. Thus, we measure psycho-social exposure by adopting occupational indexes based on the Job Exposure Matrices constructed by Kroll (2011) using a large-scale representative survey of working conditions of approximately 20,000 employees in Germany. In particular, we use their measure of 'psycho-

social burden', which is based on mental stress, social stress, and temporal loads. Figure A4 shows a distribution of industry sectors by this psychosocial exposure index. We link individuals' last industry with this aggregate occupation-level data[26] and divide our sample by the median of this index.

Columns 3 and 4 of Table 5 report that a delay of one year increases the probability of dying by 5.7 percentage points ($\sim 50\%$) for individuals with occupations in industries with a high psychosocial burden. In contrast, the increase is smaller (2.8 percentage points) for those with occupations in industries with fewer psychosocial burderns. Again, in Table A18, we show that the impact on the highest and lowest groups are statistically different. These results imply that losing the right to retire early can lead to the death of individuals who were not only in physically demanding jobs but who also had a high exposure to psychosocial burdens in their workplace.

Self-value at Work

Previous literature has pointed out that retirement can negatively impact individuals' wellbeing, as they often lose the social network of their co-workers and may feel less valuable to society (Szinovacz et al., 1992). Therefore, we want to test this hypothesis by looking at the heterogeneous effect of delaying retirement on mortality based on how and whether individuals felt useful in their job before retirement.

As we do not have a good proxy of usefulness at work in the Spanish context, we utilize the Occupational Information Network (O*NET) collected by the US Department of Labor. We use the work value classification to measure self-value at the workplace, which includes two elements: a sense of achievement and recognition within the workplace. Figure A5 shows the distribution of industry sectors by this self-value index. In our sample, we link individuals' last industry with this aggregate occupational-level data,[27] and divide the sample by the index's median.

In columns 5 and 6 of Table 5, we find strong evidence that the mortality effects between the ages of 60 and 69 are driven by individuals working in low self-value industries. Delaying the labour market exit by one year increases the probability of dying between the ages of 60 and 69 by 6.4 percentage points ($\sim 60\%$) for individuals working in these sectors, while the impact is small and insignificant for individuals working in sectors with high self-value. The impacts on these two subgroups are statistically different. Therefore, this result indicates that individuals who

[26]The psychosocial burden occupational index elaborated by Kroll (2011) is linked to individuals' last industry following these steps. First, we group all the industries defined in CNAE09 into 21 different groups. Using the Labour Force Survey of 2011, we observe which occupations (defined by CNO11) are most often performed in each of the 21 industries groups and with what frequency. Finally, we link the psychosocial index with each industry depending on which occupations are usually performed within each industry, using the frequencies as weights to calculate the mean psychosocial burden in each sector.

[27]We link the occupational index of self-value with individuals' last industry following the same steps as for the psychosocial burden index.

feel a sense of achievement and recognition within their workplace do not experience a negative mortality effect due to a delay in their exit from the labour market.

Skill Level

Finally, previous literature has relied heavily on heterogeneity differentiating between blue- and white-collar jobs, typically based on each occupation's assumed skill level (Coe et al., 2012). Following this previous literature, we also look at the differential effect of age at last employment on mortality for individuals working in white- and blue-collar occupations in columns 7 and 8 of Table 5. Contrary to Mazzonna and Peracchi (2017), we find this heterogeneity very similar to that based on the physical burden. Delaying retirement by one year increases the probability of dying between the ages of 60 and 69 by 5.4 percentage points ($\sim 68\%$) for individuals with a blue-collar job, while it is 2.7 percentage points for the rest. Even though this difference in results is not statistically different, it indicates that, in this context, skills capture differences in physical burden across occupations.

Possibility of Gradual Retirement

Reducing the possibility of early retirement, as examined here, appears to be a good strategy to cope with an aging population, as it prolongs the working careers of older workers. However, we have shown that this type of policy leads to serious adverse effects on individuals' life expectancy. A potential solution to incentivize workers to stay longer in the labour force without negatively impacting their health is to allow these workers to gradually reduce their working time at the end of their careers.

As already explained in Section 2, from 2002, the Spanish pension system introduced the possibility of individuals partially retiring after the age of 60, allowing them to combine income from work with old-age pension benefits. They were allowed to claim up to 85% of their pension while reducing employment time from 85% to 15% of the original contract. However, this option, which is also subject to the agreement of the firm, was only available for workers with at least 33 years of contribution and six years of tenure in the same company.

In this section, we analyse whether having the option to claim a partial pension can mitigate the negative impact of delaying the age at which individuals leave the labour market on mortality. As we observe that the reform affected the probability of individuals claiming a partial pension, we cannot simply directly look at the mortality effect of those individuals who chose this retirement scheme. Therefore, we take advantage of the fact that only individuals with at least 33 years of contribution have access to this scheme.

Table 6 shows that an increase of one year in the age at last employment increases mortality between the ages of 60 and 69 by 6.8 percentage points ($\sim 62\%$) for individuals with less than 33

years of contributions, who could not access to partial retirement. On the other hand, the effect is much smaller (2 percentage points or 23%) for individuals with more than 33 years of contributions who could potentially access the partial retirement scheme. We also observe that the differential impact of the reform on both subgroups is statistically different.

This result indicates that introducing the possibility of reducing the working time for older workers at the end of their careers can help mitigate the adverse effects on health of delaying retirement.

6 Discussion

6.1 Comparison with Existing Studies

We find that individuals who contributed in 1967 (a delay of five years in statutory retirement age) have a 2-percentage point higher probability of dying between the ages of 60 and 69 (22% increase). The IV estimates indicate that delaying the age at last employment by one year increases the probability of dying between the ages of 60 and 69 by 4.2 percentage points (43%). This may seem quite a large effect; however, our estimates are comparable in magnitude with studies showing that early retirement reduces mortality (Hallberg et al., 2015; Bloemen et al., 2017).

Hallberg et al. (2015) find that by offering a five-year reduction of the statutory retirement age from the age of 65 to 60 reduces the probability of dying by the age of 70 by 26 percent. Using the same measure of mortality, we find that a five-year increase in the statutory retirement age from the age of 60 to 65 increases the probability of dying before age 70 by 3 percentage points, which is equivalent to a 17 percent increase. Additionally, Hallberg et al. (2015) show that the mortality effects are driven by those who are more exposed to workplace hazards; that is, those with low pre-retirement incomes and those without a college education. Their finding is consistent with our heterogeneous results. Bloemen et al. (2017) also find estimates of a similar magnitude. They find that retirement induced by a temporary decrease in the retirement eligibility age (from the age of 65 to 61 or 62) for male Dutch civil servants decreased the probability of dying within five years by 47 percent (2.6 percentage points).[28] Although our prior is that the effect of delaying retirement is not necessarily symmetric with the impact of early retirement, our estimates suggest that the effect on mortality has a similar magnitude when the nature of the reform and affected age ranges are comparable.

Our paper is the first to find that retirement reduces mortality by exploring quasi-experiments that delay retirement. Existing papers find no effect of delaying retirement on mortality. Bozio et

[28] Although Hallberg et al. (2015) study male military officers in Sweden and Bloemen et al. (2017) focus on Dutch male civil servants, both papers point out that the working environment of these subgroups of males is not more demanding than that for the general population.

al. (2021) and Saporta-Eksten et al. (2021) are the only two papers we know of that have the statistical power to conclusively estimate the mortality impacts, and they find precisely zero effects of delaying retirement.[29] One common feature of these two papers is that they explore reforms that increase the financial incentives to delay retirement while keeping the statutory retirement age unchanged. Bozio et al. (2021) find a precisely zero impact of delaying retirement on the probability of dying between the ages of 61 and 79 for private-sector workers in France. Saporta-Eksten et al. (2021) find no effect of delaying retirement on mortality between the ages of 65 and 74 by exploring a reform that reduced the implied tax of working for married males in Israel.[30] One possible reason that we find an adverse impact of delayed retirement (while their study find no effect) is that the groups of compliers differ. Those who retire later as a response to a pension reform incentivizing later retirement via financial incentives differ from those who retire later because the early retirement possibility is not available. Therefore, it is reasonable to expect a more harmful impact on mortality when the early retirement option is removed than when early retirement is financially less attractive. Moreover, we find that workers entitled to gradual retirement suffer less from the reform (see Table 6). This finding indicates that delaying retirement is less harmful when pension reform provides a flexible choice, rather than a paternalistic policy that prohibits workers from retiring earlier.

6.2 Policy Discussion

The heterogeneous mortality impacts of delayed retirement suggest important distributional consequences of raising the statutory retirement age. In particular, the socio-economic disparities in lifespans are large and have increased in recent decades (OECD, 2016).[31] One possible contributing factor might be the heterogeneous mortality responses to pension reforms, which could exacerbate the disparity. Moreover, individuals who survive longer receive more years of pen-

[29] Hagen (2018) explores a reform that increases the statutory retirement age from the age of 63 to 65 for Swedish public sector workers born since 1938. They find an imprecisely measured no effect on mortality by the age of 69. Their IV estimates show that a one-year increase in retirement age results in a 0.34% increase in morality by the age of 69 (insignificant).

[30] It is important to note that Saporta-Eksten et al. (2021) show a decline in the probability of survival of the affected men between the ages of 75 and 85 due to later retirement. Overall, they find that one additional year of employment decreases longevity by 9 to 12 months. Our result is in line with this finding.

[31] We acknowledge that life expectancy also differs largely by gender. In Spain, in 2021, men live on average until age 80.2, while women live on average until age 85.8 (Spanish National Institute of Statistics). In Table A19, we examine if the reform had differential effects across gender. We show that a year delaying retirement increases mortality more for men than women. In particular, a one-year delay in the age at which men exit the labour market increases by 7.7 percentage points ($\sim$ 59%) the probability of dying between the ages of 60 and 69. The same delay for women increases mortality by 1.9 percentage points ($\sim$ 48%). Factors influencing gender differences in mortality include biological factors (genetics and hormones) and behavioural and environmental factors. One behavioural factor that explains part of the mortality gender gap is that women and men select different occupations. Therefore, the differential effect of delaying retirement on mortality by gender may be partly driven by men and women selecting occupations and sectors with different degrees of health burden (DeLeire and Levy, 2001).

sion. The resulting gaps in life expectancy will affect the actuarial fairness and progressivity of public pension systems (Sanchez-Romero et al., 2020). Specifically, individuals from lower socio-economic groups (typically those who are more exposed to workplace hazards) spend fewer years in retirement than the rest of the population due to the pension reform.

A reduction in the duration of claiming a pension is composed of two factors: delayed claiming a pension and earlier death. While the welfare impact of delayed labour market exit can be positive, earlier death is harmful. Table 5 shows that individuals with strenuous employment (both physically and psychosocially), low-self-value and who work in blue-collar jobs experience a greater increase in mortality between the ages of 60 and 69 due to the reform. In comparison, the reform impact on the individual's age at last employment is relatively similar between the different subgroups. If anything, individuals with better jobs delay their exit from the workplace for a longer period. This comparison implies that the mortality impact plays an important role in explaining the shortened pension claiming duration for workers with worse working conditions. One possible policy recommendation would be to consider reforms that link retirement age to changes in life expectancy. It might be worthwhile to consider a target retirement age based on the years a person is expected to claim a pension rather than a uniform nationwide retirement age.

7 Conclusion

This paper studies the effect of delaying retirement on mortality. We exploit the 1967 Spanish reform that removed access to voluntary early retirement for individuals who had not contributed since that year. Individuals who started contributing to the pension system before 1 January 1967 maintained the right to retire early at the age of 60. However, individuals who have not contributed by that date can only retire voluntarily at the statutory retirement age of 65 (although, under certain circumstances, some individuals can involuntarily retire early at the age of 61).

Focusing on cohorts born between 1938 and 1949, we use Spanish administrative Social Security data and compare individuals who started contributing 12 months before and after 1 January 1967. We first show the reform effect on labour supply outcomes using a within-cohort OLS regression controlling for gender and individuals' employment history between the ages of 30 and 40. We find that individuals who started contributing after 1967 delayed their labour market exit by almost half a year. The reform not only modified the age at last employment but also changed the age of claiming a pension and the types of pensions claimed. We find a decrease in the probability of claiming a regular pension by 19%, an increase in the probability of claiming a partial pension by 54%, and an increase in the probability of claiming disability insurance by 19%. This indicates that individuals did not fully comply with the rise in the statutory retirement age and utilized other ways to leave the labour market before claiming a regular pension. Moreover, the results suggest

that treated individuals are more likely to claim no pension, driven mainly by premature death.

Furthermore, we estimate the effect of age at last employment on mortality using the instrumental variable method. We find that delaying labour market exit by one year increases the hazard of dying between the ages of 60 and 69 by 4.2 percentage points (43%). The mortality responses are the strongest between the ages of 60 and 64 (67%), when public pensions are no longer accessible for individuals who started contributing after 1967. This suggests that the effect of delaying retirement on mortality is driven mainly by the immediate effect of losing access to early retirement schemes.

We explore several mechanisms to explain the detrimental effects of delaying retirement on health. First, we show that individuals' workplace conditions before retirement are an essential factor. Moreover, we show that allowing workers to gradually reduce their working time towards the end of their career and making partial retirement an option can incentivize workers to stay longer in the labour force without negatively affecting their health.

The applicability and relevance of our findings extend further than the Spanish setting. Delaying statutory retirement and closing early retirement options is a pertinent policy agenda in many countries. However, the existing empirical evidence on the mortality effects of retirement rests almost exclusively on the estimates of policy experiments that have allowed for earlier retirement. Given that it is unclear if there is a symmetry impact of advancing or delaying retirement age, our findings on the mortality effect of delaying retirement are particularly relevant.

Additionally, the heterogeneous mortality impacts of delaying retirement raise discussions on the distributional consequences of raising the statutory retirement age. We find that individuals who have high physically and/or psychosocially burdensome jobs are those who suffer the most from a delay in retirement. Furthermore, the reform has a more substantial effect on individuals in jobs where they feel they have achieved less and received less recognition for their contributions. Combining the results on partial retirement, our findings suggest that it is crucial to provide options for gradual and flexible retirement while raising the age of statutory retirement.

References

Atalay, Kadir and Garry F Barrett, "The causal effect of retirement on health: New evidence from Australian pension reform," *Economics Letters*, 2014, *125* (3), 392–395.

_ and _, "The impact of age pension eligibility age on retirement and program dependence: Evidence from an Australian experiment," *Review of Economics and Statistics*, 2015, *97* (1), 71–87.

Barnett, Inka, Esther van Sluijs, David Ogilvie, and Nicholas J Wareham, "Changes in household, transport and recreational physical activity and television viewing time across the transition to retirement: longitudinal evidence from the EPIC-Norfolk cohort," *J Epidemiol Community Health*, 2014, *68* (8), 747–753.

Black, Sandra E, Paul J Devereux, and Kjell G Salvanes, "Losing heart? The effect of job displacement on health," *ILR Review*, 2015, *68* (4), 833–861.

Bloemen, Hans, Stefan Hochguertel, and Jochem Zweerink, "The causal effect of retirement on mortality: Evidence from targeted incentives to retire early," *Health economics*, 2017, *26* (12), e204–e218.

Blundell, Richard, Eric French, and Gemma Tetlow, "Retirement incentives and labor supply," in "Handbook of the economics of population aging," Vol. 1, Elsevier, 2016, pp. 457–566.

Boldrin, Michele, Sergi Jiménez-Martín, and Franco Peracchi, "Micro-modeling of retirement behavior in Spain," *NBER Chapters*, 2004, pp. 499–578.

_ , _ , _ et al., "Social security and retirement in Spain," *Social Security and Retirement around the world*, 1999, pp. 305–53.

Bozio, Antoine, Clémentine Garrouste, and Elsa Perdrix, "Impact of later retirement on mortality: Evidence from France," *Health economics*, 2021, *30* (5), 1178–1199.

Case, Anne, Angus Deaton et al., "Broken down by work and sex: How our health declines," *Analyses in the Economics of Aging*, 2005, *1*, 185–214.

Celidoni, Martina and Vincenzo Rebba, "Healthier lifestyles after retirement in Europe? Evidence from SHARE," *The European Journal of Health Economics*, 2017, *18* (7), 805–830.

Coe, Norma and Maarten Lindeboom, "Does retirement kill you? Evidence from early retirement windows," 2008.

Coe, Norma B, Hans-Martin von Gaudecker, Maarten Lindeboom, and Jürgen Maurer, "The effect of retirement on cognitive functioning," *Health economics*, 2012, *21* (8), 913–927.

Coile, Courtney and Jonathan Gruber, "Future social security entitlements and the retirement decision," *The review of Economics and Statistics*, 2007, *89* (2), 234–246.

DeLeire, Thomas and Helen G Levy, "Gender, occupation choice and the risk of death at work," 2001.

Eibich, Peter, "Understanding the effect of retirement on health: Mechanisms and heterogeneity," *Journal of health economics*, 2015, *43*, 1–12.

Evenson, Kelly R, Wayne D Rosamond, Jianwen Cai, Ana V Diez-Roux, and Frederick L Brancati for the Atherosclerosis Risk in Communities Study Investigators, "Influence of retirement on leisure-time physical activity: the atherosclerosis risk in communities study," *American journal of epidemiology*, 2002, *155* (8), 692–699.

Falba, Tracy, Hsun-Mei Teng, Jody L Sindelar, and William T Gallo, "The effect of involuntary job loss on smoking intensity and relapse," *Addiction*, 2005, *100* (9), 1330–1339.

Fitzpatrick, Maria D and Timothy J Moore, "The mortality effects of retirement: Evidence from Social Security eligibility at age 62," *Journal of Public Economics*, 2018, *157*, 121–137.

García-Gómez, Pilar, Sergi Jiménez-Martín, and J Vall Castelló, "Health, disability, and pathways into retirement in Spain," *Social security programs and retirement around the world*, 2012, pp. 127–174.

Garcia-Perez, J, "La muestra continua de vidas laborales (MCVL): una guía de uso para el análisis de transiciones," *Revista de Economia Aplicada*, 2008, *16* (8), 5–28.

García-Pérez, J Ignacio, Sergi Jiménez-Martín, and Alfonso R Sánchez-Martín, "Retirement incentives, individual heterogeneity and labor transitions of employed and unemployed workers," *Labour Economics*, 2013, *20*, 106–120.

Geyer, Johannes and Clara Welteke, "Closing Routes to Retirement for Women How Do They Respond?," *Journal of Human Resources*, 2021, *56* (1), 311–341.

Godtfredsen, Nina S, Tai H Lam, Trevor T Hansel, ME Leon, N Gray, C Dresler, DM Burns, E Prescott, and J Vestbo, "COPD-related morbidity and mortality after smoking cessation: status of the evidence," *European Respiratory Journal*, 2008, *32* (4), 844–853.

Gorry, Aspen, Devon Gorry, and Sita Nataraj Slavov, "Does retirement improve health and life satisfaction?," *Health economics*, 2018, *27* (12), 2067–2086.

Hagen, Johannes, "The effects of increasing the normal retirement age on health care utilization and mortality," *Journal of Population Economics*, 2018, *31* (1), 193–234.

Hallberg, Daniel, Per Johansson, and Malin Josephson, "Is an early retirement offer good for your health? Quasi-experimental evidence from the army," *Journal of health economics*, 2015, *44*, 274–285.

Hernaes, Erik, Simen Markussen, John Piggott, and Ola L Vestad, "Does retirement age impact mortality?," *Journal of health economics*, 2013, *32* (3), 586–598.

Insler, Michael, "The health consequences of retirement," *Journal of Human Resources*, 2014, *49* (1), 195–233.

Kivimäki, Mika, Päivi Leino-Arjas, Ritva Luukkonen, Hilkka Riihimäi, Jussi Vahtera, and Juhani Kirjonen, "Work stress and risk of cardiovascular mortality: prospective cohort study of industrial employees," *Bmj*, 2002, *325* (7369), 857.

Kroll, Lars Eric, "Construction and validation of a general index for job demands in occupations based on ISCO-88 and KldB-92," *methods, data, analyses*, 2011, *5* (1), 28.

Kuhn, Andreas, Stefan Staubli, Jean-Philippe Wuellrich, and Josef Zweimüller, "Fatal attraction? Extended unemployment benefits, labor force exits, and mortality," *Journal of Public Economics*, 2020, *191*, 104087.

Lee, I-Min, Howard D Sesso, and RS Paffenbarger Jr, "Physical activity and risk of lung cancer.," *International Journal of Epidemiology*, 1999, *28* (4), 620–625.

Manoli, Day and Andrea Weber, "Nonparametric Evidence on the Effects of Financial Incentives on Retirement Decisions," *American Economic Journal: Economic Policy*, November 2016, *8* (4), 160–82.

Mastrobuoni, Giovanni, "Labor supply effects of the recent social security benefit cuts: Empirical estimates using cohort discontinuities," *Journal of Public Economics*, dec 2009, *93* (11-12), 1224–1233.

Mazzonna, Fabrizio and Franco Peracchi, "Unhealthy retirement?," *Journal of Human Resources*, 2017, *52* (1), 128–151.

OECD, "Fragmentation of retirement markets due to differences in life expectancy," 2016.

Ravesteijn, Bastian, Hans van Kippersluis, and Eddy van Doorslaer, "The contribution of occupation to health inequality," in "Health and inequality," Emerald Group Publishing Limited, 2013.

Sanchez-Romero, Miguel, Ronald D Lee, and Alexia Prskawetz, "Redistributive effects of different pension systems when longevity varies by socioeconomic status," *The Journal of the Economics of Ageing*, 2020, *17*, 100259.

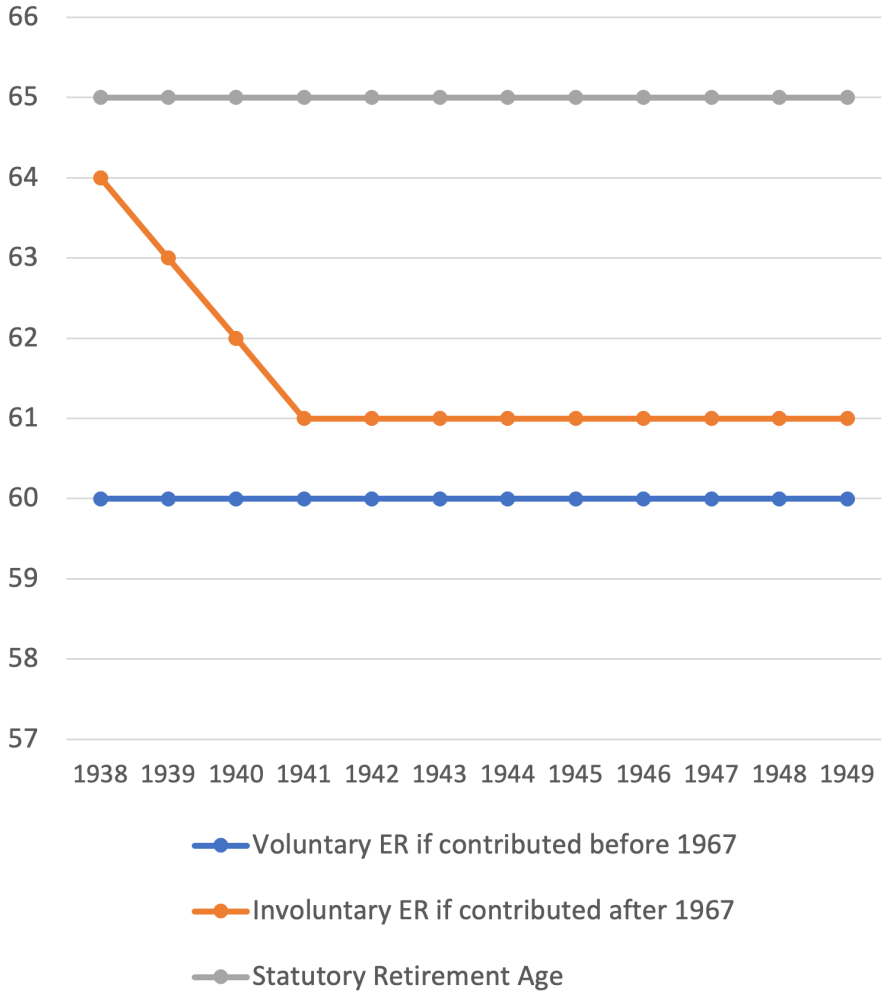
Saporta-Eksten, Itay, Ity Shurtz, and Sarit Weisburd, "Social Security, Labor Supply, and Health of Older Workers: Quasi-Experimental Evidence from a Large Reform," *Journal of the European Economic Association*, 2021, *19* (4), 2168–2208.

Szinovacz, Maximiliane, Maximiliane Szinovacz, David J Ekerdt, and Barbara H Vinick, *Families and retirement*, Sage, 1992.

Zulkarnain, Alice and Matthew S Rutledge, "How does delayed retirement affect mortality and health?," *Center for retirement research at Boston College, CRR WP*, 2018, *11*.

8 Figures and Tables

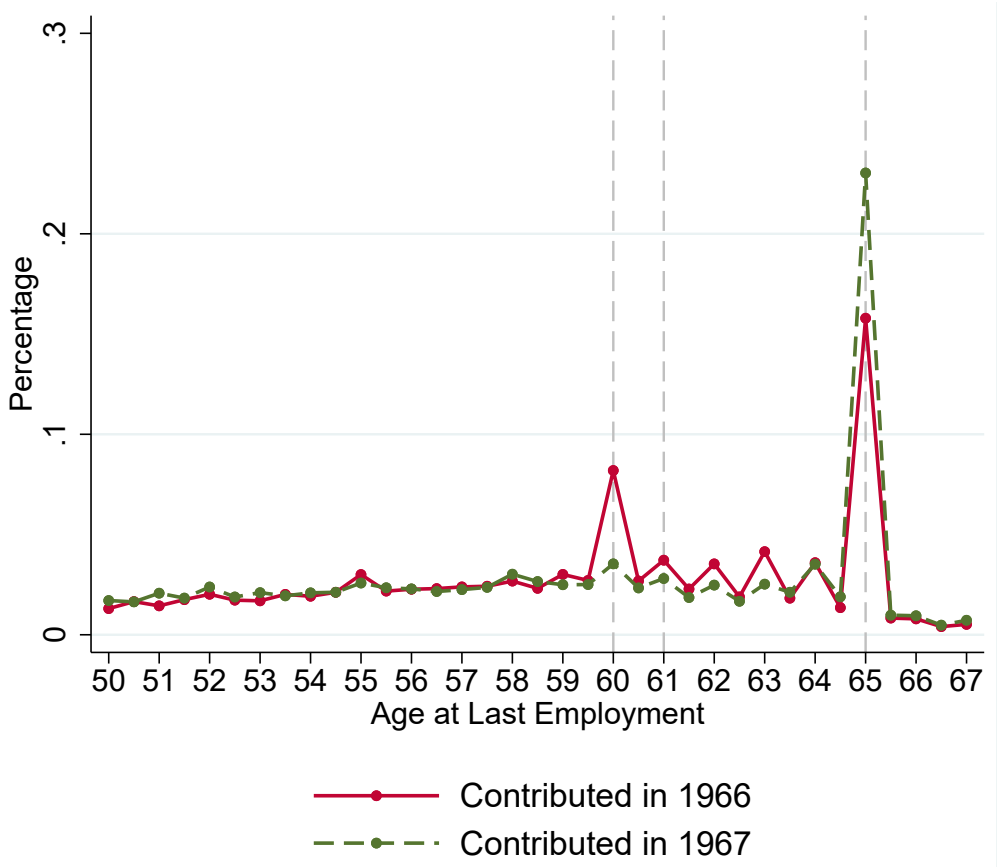
Figure 1: Retirement Age by First Year of Contribution and Cohort



Source: Authors' own construction according to the pension laws.

Notes: This figure plots the statutory retirement age and the earliest possible early retirement age for individuals that contributed before and after 1 January 1967 as a function of their birth year. The blue line shows that individuals who start contributing before 1 January 1967 can voluntarily retire after age 60, independently of their birth year. The orange line shows that those who start contributing after 1967 can only involuntary retire after 64 to 61, depending on their birth year. The grey line shows that the statutory retirement age remains at age 65 for all cohorts independently from the moment they started contributing.

Figure 2: Density of Age at Last Employment by Treatment Status

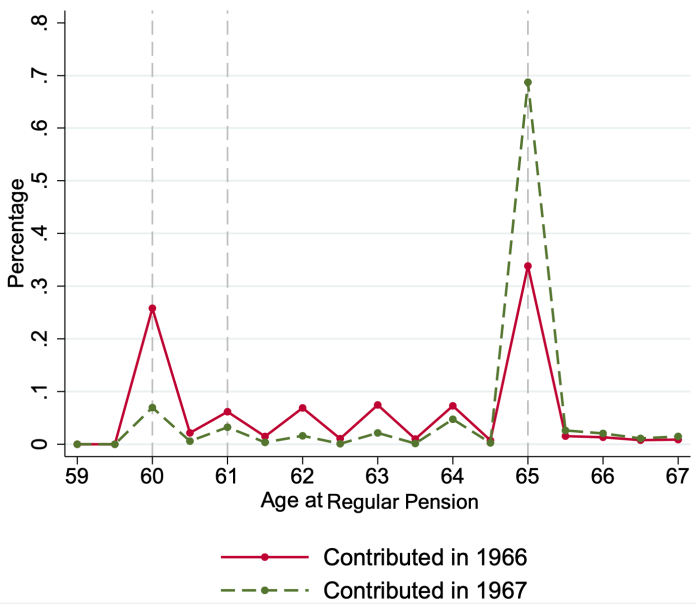


Source: MCVL, cohorts 1938-1949.

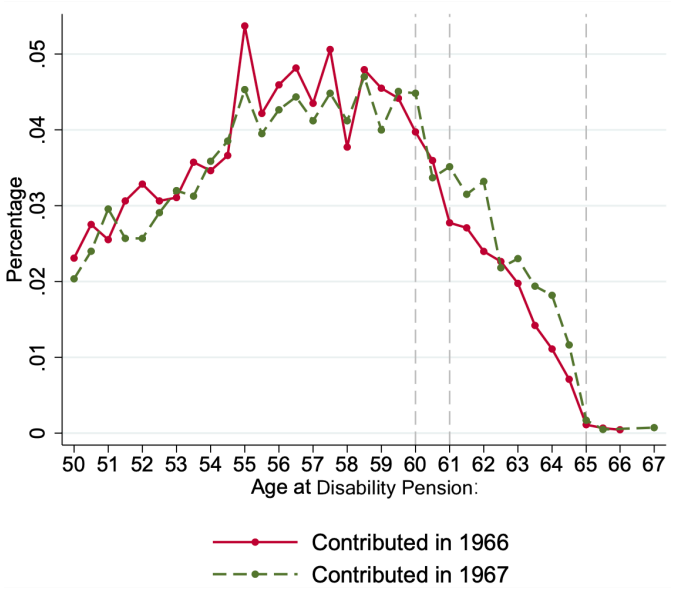
Notes: This figure plots the percentage of individuals by the age at which they finished their last employment. The solid red line shows the density for individuals who started contributing in 1966, while the green dashed line shows those who started contributing in 1967.

Figure 3: Density of Pension Ages by Treatment Status

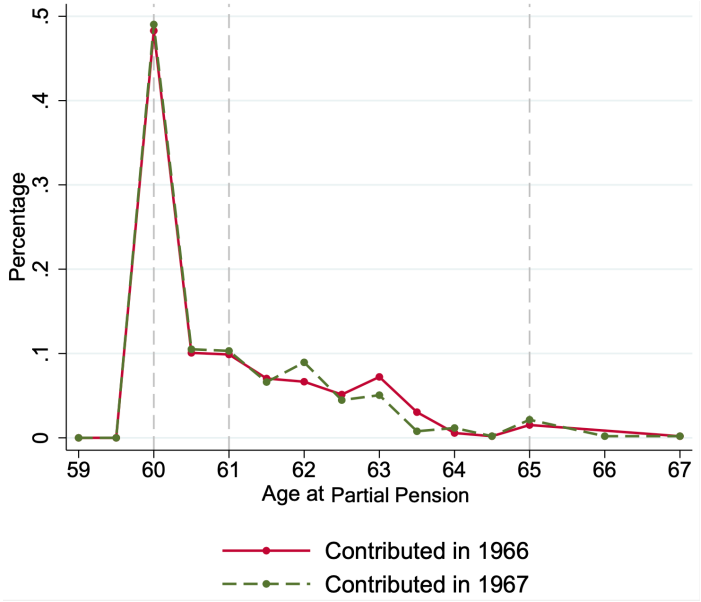
(a) Age at Regular Pension



(b) Age at Disability Pension



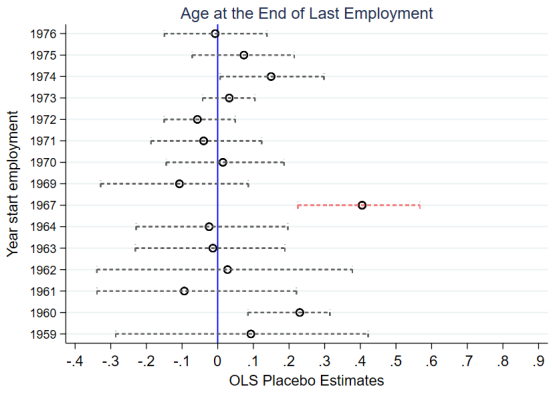
(c) Age at Partial Pension



Source: MCVL, cohorts 1938-1949.

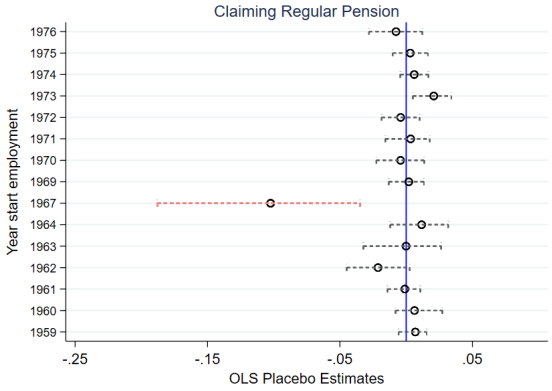
Notes: This figure plots the percentage of individuals by the age at claiming regular pension (Graph a), age at claiming disability pension (Graph b), and age at claiming partial pension (Graph c). The solid red lines show the density for individuals who started contributing in 1966, while the green dashed lines show those who started contributing in 1967.

Figure 4: Placebo Tests: Using Other Cutoffs

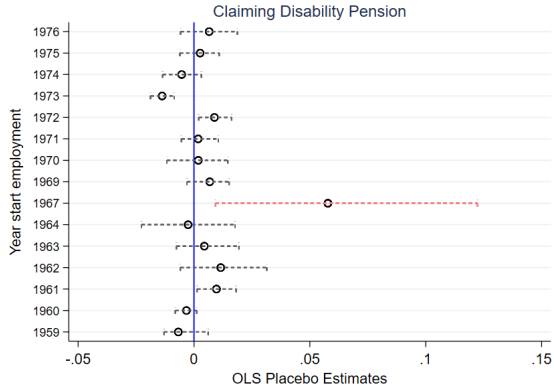


(a) Age at last employment

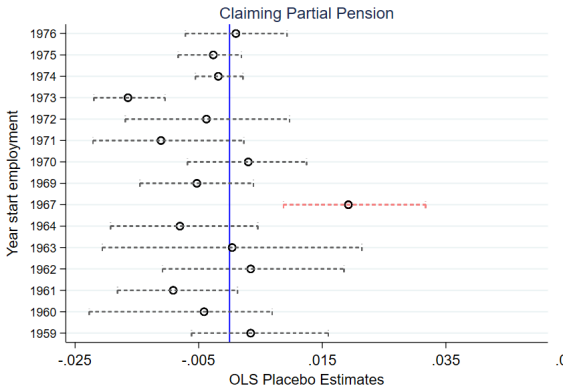
(b) Age at claiming regular pension



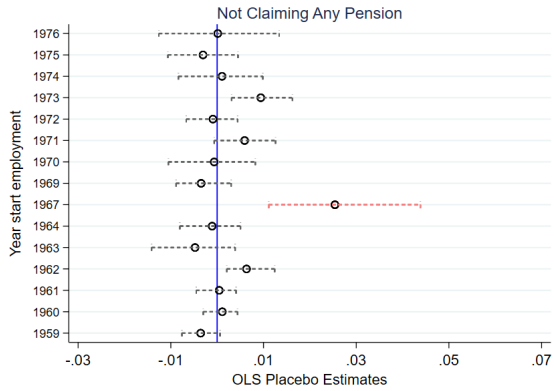
(c) Probability of claiming regular pension



(d) Probability of claiming disability pension



(e) Probability of claiming partial pension



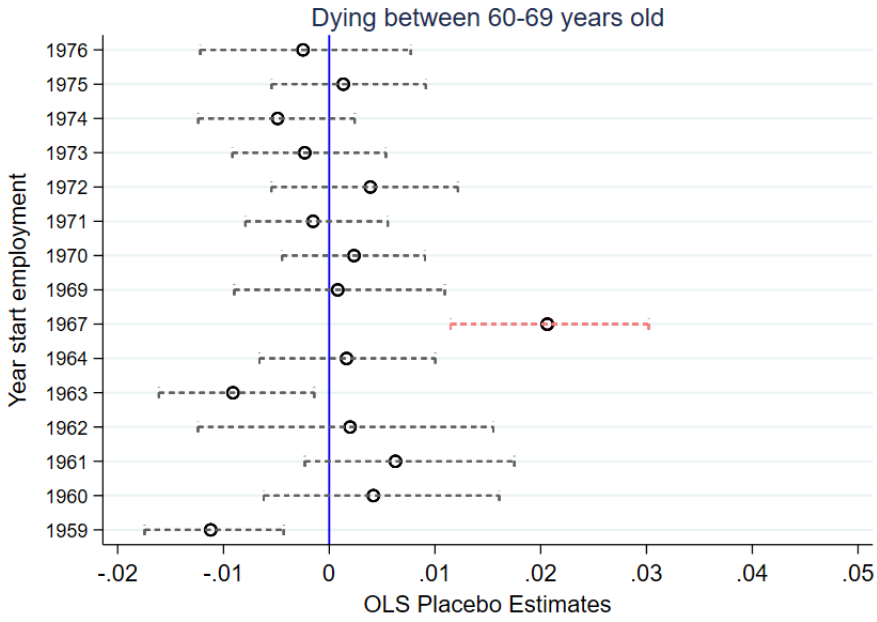
(f) Probability of not claiming any pension

Source: MCVL, cohorts 1938-1949.

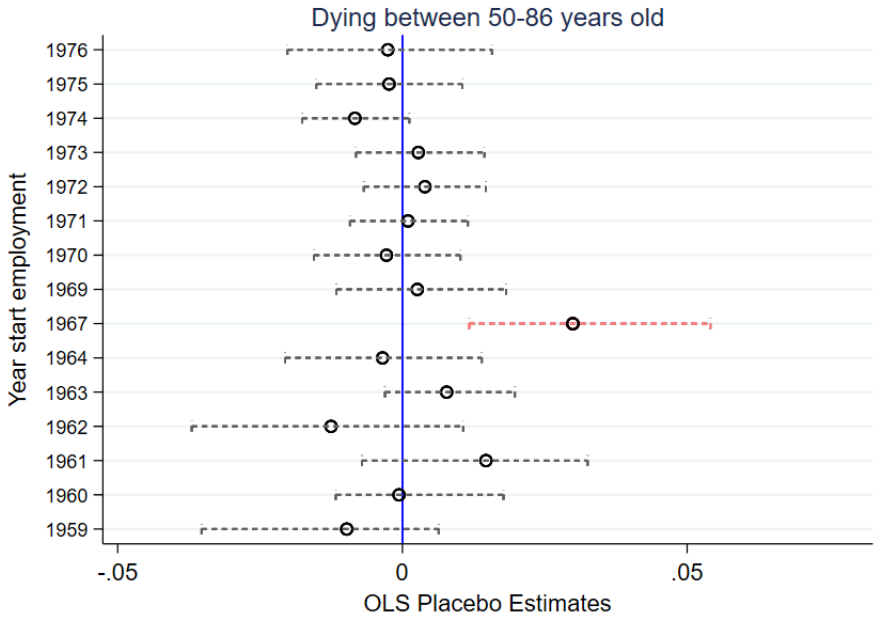
Notes: This figure shows the estimates and the 95 percent confidence intervals of a list of placebos, estimating regression 1 using the years of the y-axis as the cutoff. The red estimate corresponds to the estimation of the regression 1 on the real cutoff: 1967. The outcomes considered are displayed on top of each figure.

Figure 5: Placebo Tests for Mortality: Using Other Cutoffs

(a) Dying between 60 and 69 years old



(b) Dying between 50 and 86 years old



Source: MCVL, cohorts 1938-1949.

Notes: This figure shows the estimates and the 95 percent confidence intervals of a list of placebos, estimating regression 1 using the years of the y-axis as the cutoff. The red estimate corresponds to the estimation of the regression 1 on the real cutoff: 1967. The outcomes considered are displayed on top of each figure.

Table 1: Impact of the Reform on the Type of Pension

	First Pension Claimed			
	Regular Pension	Partial Pension	Disability Pension	No Pension
Contributed in 1967	-0.102***	0.019***	0.058**	0.025***
	(0.032)	(0.005)	(0.024)	(0.007)
	[0.006]	[0.002]	[0.029]	[0.003]
Month-Year Birth FE	✓	✓	✓	✓
Controls	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓
Observations	25,903	25,903	25,903	25,903
R^2	0.140	0.067	0.091	0.054
Mean Dep. (Treated)	0.398	0.048	0.387	0.167
Mean Dep. (Control)	0.533	0.035	0.296	0.137

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the probability of leaving the labour market through regular pension (Column 1), partial pension (Column 2), disability pension (Column 3), and not claiming any pension (Column 4), obtained from the estimation of regression 1. The estimation sample includes individuals that started contributing 12 months before and after January 1st, 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table 2: Impact of the Reform on the Age at Claiming Pension

	Age of the Individual at				
	Last Employment	First Pension	Regular Pension	Disability Pension	Partial Pension
Contributed in 1967	0.405***	0.279**	1.310***	0.288**	-0.153***
	(0.073)	(0.107)	(0.207)	(0.067)	(0.054)
	[0.003]	[0.041]	[0.001]	[0.023]	[0.005]
Month-Year Birth FE	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓
Observations	25,903	22,040	12,367	8,633	1,040
R^2	0.083	0.106	0.220	0.035	0.245
Mean Dep. (Treated)	59.858	61.049	64.641	57.352	61.101
Mean Dep. (Control)	59.401	60.889	63.036	56.986	61.134

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the age at last employment (Column 1), at claiming first pension (any type) (Column 2), regular pension (Column 3), disability pension (Column 4), and partial pension (Column 5), obtained from the estimation of regression 1. The estimation sample includes individuals that started contributing 12 months before and after January 1st, 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table 3: Impact of the Reform on Pension Benefit

	Pension Benefit	Base Pension	Percent of Base Pension
Contributed in 1967	43.777***	19.112**	5.039***
	(10.913)	(6.570)	(1.014)
	[0.003]	[0.015]	[0.001]
Month-Year Birth FE	✓	✓	✓
Controls	✓	✓	✓
Contributed 1966-1967	✓	✓	✓
Observations	22,040	22,040	22,039
R^2	0.377	0.410	0.155
Mean Dep. (Treated)	1177.066	1203.640	86.099
Mean Dep. (Control)	1089.755	1150.027	78.796

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on monthly pension benefit (Column 1), pension base (Column 2), and pension adjustment factor (Column 3), obtained from the estimation of regression 1 for those individuals in our sample that claim any pension. The estimation sample includes individuals that started contributing 12 months before and after January 1st, 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table 4: Impact of Age at Last Employment on Mortality

	Probability of Dying between the Ages					
	50-86	50-59	60-69	70-79	80-86	Age at Death Censored 71
OLS:						
Impact of Age at Last Employment	-0.013***	-0.012***	-0.004***	-0.003***	-0.000	0.207***
	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.015)
	[0.001]	[0.001]	[0.001]	[0.001]	[0.622]	[0.001]
Reduced Form:						
Contributed in 1967	0.030***	0.010	0.021***	0.001	0.004*	-0.305***
	(0.009)	(0.006)	(0.004)	(0.007)	(0.002)	(0.092)
	[0.002]	[0.115]	[0.001]	[0.840]	[0.059]	[0.001]
IV:						
Impact of Age at Last Employment	0.074**	0.025	0.042***	0.003	0.006**	-0.757**
	(0.031)	(0.015)	(0.010)	(0.012)	(0.003)	(0.278)
	[0.019]	[0.147]	[0.003]	[0.834]	[0.036]	[0.012]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓	✓
Observations	25,903	25,903	23,987	21,009	17,516	25,903
Mean Dep. (Treated)	0.369	0.085	0.131	0.184	0.036	68.859
Mean Dep. (Control)	0.297	0.066	0.098	0.154	0.023	69.384
F-stat FS	35.762	35.762	59.364	56.135	68.987	35.762

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of age at last employment on the probability of dying between the ages of 50-86 (Column 1), 50-59 (Column 2), 60-69 (Column 3), 70-79 (Column 4), and 80-86 (Column 5). Column 6 reports the impact of age at last employment on age at death censored at 71 years old. The first panel reports the correlation of age at last employment on mortality (OLS), and the second panel shows the effect of the reform on mortality (reduced form effect using regression 1). The IV estimates, obtained from the estimation of regression 2, are reported in the third panel. The estimation sample includes individuals that started contributing 12 months before and after January 1st, 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. The IV estimation also controls for a proxy of the pension base. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

Table 5: Impact on Mortality by Labour Market Conditions Before Retirement

	Last Industry						Last Occupation	
	Workplace Accidents		Psychosocial Exposure		Self-value		Blue-collar	
	High	Low	High	Low	High	Low	No	Yes
	Age at Last Employment							
First Stage:	0.435***	0.682***	0.472***	0.671***	0.962***	0.420***	0.613***	0.437***
Contributed in 1967	(0.116)	(0.108)	(0.104)	(0.127)	(0.216)	(0.092)	(0.070)	(0.102)
	[0.006]	[0.001]	[0.001]	[0.002]	[0.001]	[0.001]	[0.001]	[0.005]
	Probability of Dying between 60 and 69							
Reduced Form:	0.029***	0.017***	0.027***	0.018**	0.011	0.027***	0.016**	0.024***
Contributed in 1967	(0.008)	(0.005)	(0.006)	(0.007)	(0.008)	(0.006)	(0.007)	(0.005)
	[0.016]	[0.007]	[0.002]	[0.028]	[0.150]	[0.003]	[0.040]	[0.001]
IV:	0.068**	0.025***	0.057***	0.028**	0.012	0.064***	0.027**	0.054***
Impact of Age at Last Employment	(0.028)	(0.007)	(0.016)	(0.011)	(0.009)	(0.019)	(0.011)	(0.017)
	[0.043]	[0.000]	[0.001]	[0.023]	[0.181]	[0.007]	[0.036]	[0.007]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓	✓	✓	✓
Observations	7,941	12,991	10,580	10,352	6,025	14,907	8,058	15,929
Mean Dep. (Treated)	0.158	0.122	0.148	0.126	0.112	0.146)	0.129	0.133
Mean Dep. (Control)	0.117	0.094	0.105	0.100	0.095	0.106	0.102	0.096
F-stat FS	15.848	46.688	22.664	24.433	18.253	23.120	72.882	20.457
P-value Difference (IV Est.)	0.136		0.131		0.016		0.197	

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of age at last employment on the probability of dying between the ages of 60-and 69 by the labour market conditions experienced by the individual just before retirement. Individual's last industry is classified depending on their workplace accident incidence for our cohorts between 2003 and 2019 (Columns 1 and 2), by the psychosocial exposure (mental stress, social stress, and temporal load) following Kroll (2011) (Columns 3 and 4), and by their self-value index (sense of achievement and recognition) constructed using O*NET (Columns 5 and 6). We also differentiate if individuals' last occupation pertains to a white or a blue-collar occupation (Columns 7 and 8). The first panel reports the first stage of the IV estimation (the reform's effect on the age at last employment, using 1). The second panel shows the second stage; the effect on the probability of dying between 60 and 69 years old. First, we report the reduced form effect of the reform on mortality using regression 1. After that, we report the IV estimates obtained from the estimation of regression 2. The estimation sample includes individuals that started contributing 12 months before and after January 1st, 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. The IV estimation also controls for a proxy of the pension base. At the bottom, we report the First Stage F-statistic and the p-value of the differences between groups in the IV estimation. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table 6: Impact on Mortality by Availability of Flexible Retirement

	More 33 Years of Contribution	Less 33 Years of Contribution
	Age at Last Employment	
First Stage:	0.825***	0.316***
Contributed in 1967	(0.155)	(0.079)
	[0.002]	[0.005]
	Probability of Dying between 60 and 69	
Reduced Form:	0.016**	0.022***
Contributed in 1967	(0.007)	(0.004)
	[0.019]	[0.005]
IV:	0.020***	0.068**
Impact of Age at Last Employment	(0.007)	(0.022)
	[0.010]	[0.013]
Month-Year Birth FE	✓	✓
Controls	✓	✓
Contributed 1966-1967	✓	✓
Observations	11,782	12,205
Mean Dep. (Treated)	0.108	0.155
Mean Dep. (Control)	0.087	0.109
F-stat FS	27.534	15.888
P-value Difference (IV Est.)	0.050	

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of age at last employment on the probability of dying between the ages of 60-and 69 for individuals with less (Column 1) or more than 33 years of contribution (Column 2). Only individuals with more than 33 years of contribution when claiming a pension can access the partial retirement scheme. The first panel reports the first stage of the IV estimation (the reform's effect on the age at last employment, using 1). The second panel shows the second stage; the effect on the probability of dying between 60 and 69 years old. First, we report the reduced form effect of the reform on mortality using regression 1. After that, we report the IV estimates obtained from the estimation of regression 2. The estimation sample includes individuals that started contributing 12 months before and after January 1st, 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. The IV estimation also controls for a proxy of the pension base. At the bottom, we report the First Stage F-statistic and the p-value of the differences between groups in the IV estimation. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Online Appendix

The Effect of Removing Early Retirement on Mortality

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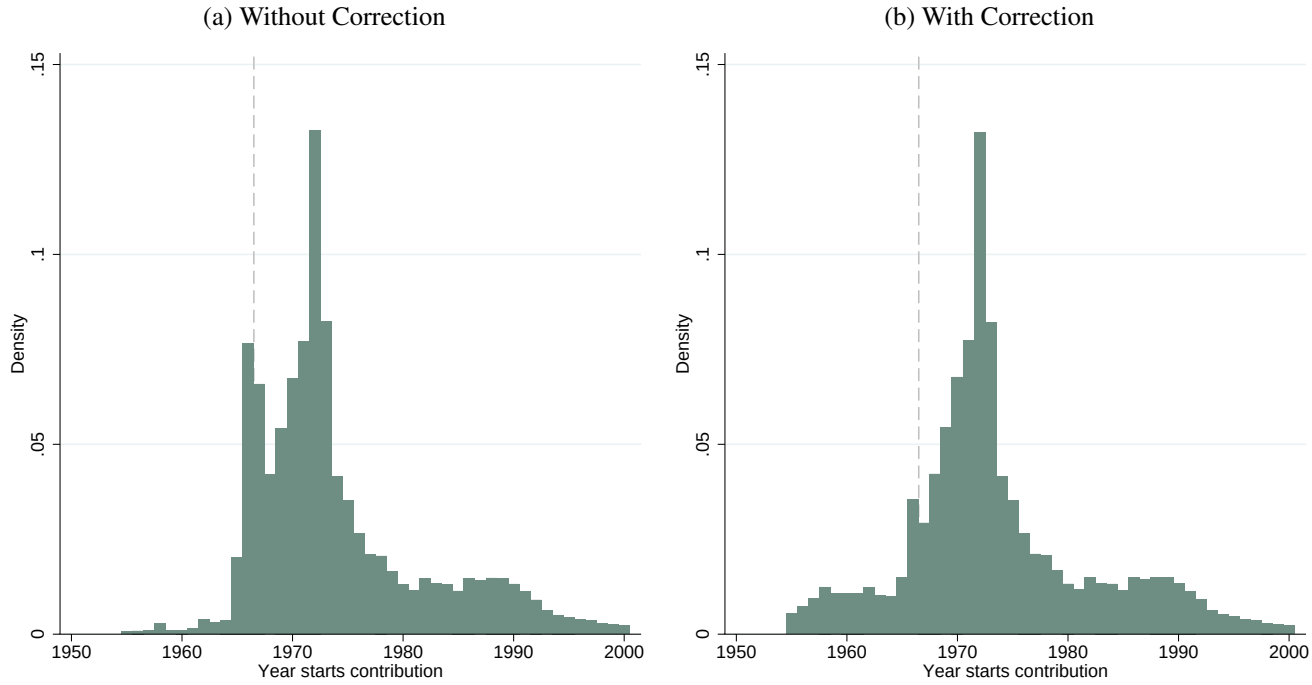
A Appendix Tables and Figures

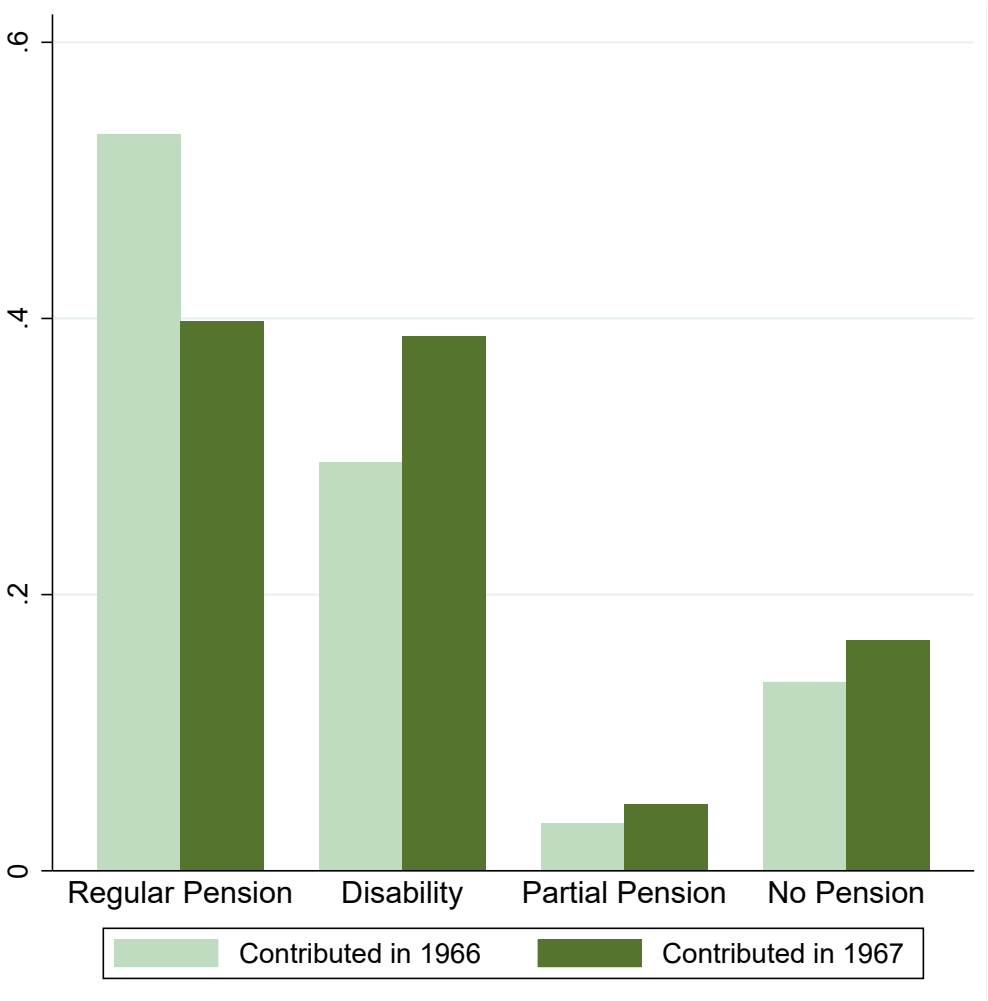
Figure A1: Correction of Year Started Contributing



Source: MCVL, cohorts 1938-1949.

Notes: This figure plots the density of date started contributing without correction (Graph a) and with correction (Graph b). The correction uses the number of years of contribution and the date starting a regular or partial pension (years of contribution are not available for individuals that claim a disability pension) to correct for the date of starting contributing for those whose year of started contributing was between 1965 and 1967.

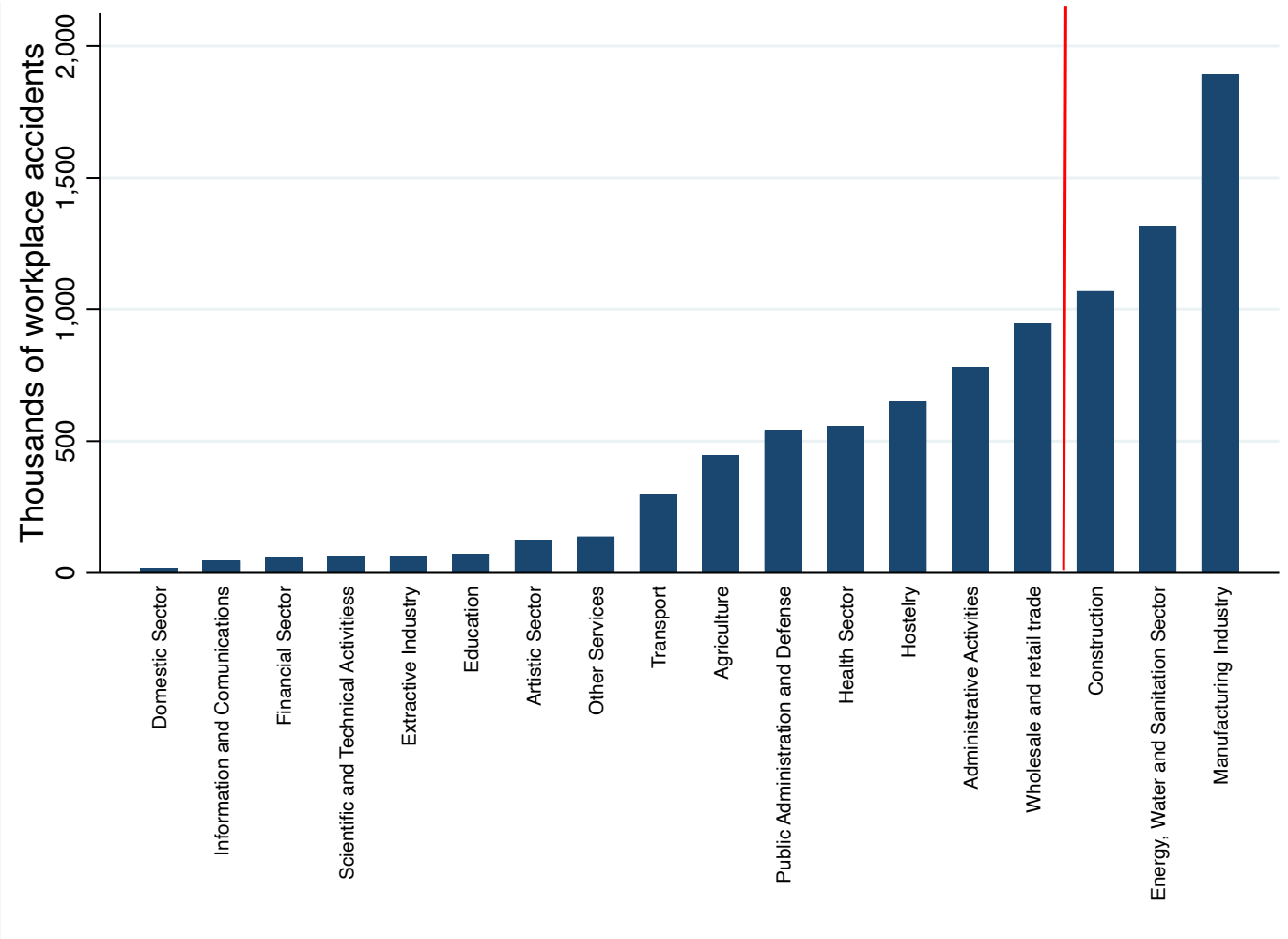
Figure A2: Types of Pension by Treatment Status



Source: MCVL, cohorts 1938-1949.

Notes: This figure plots the percentage of individuals by the different types of pension claimed (regular pension, disability insurance, partial pension, or no pension). The light green bars show the density for individuals that started contributing in 1966, while the dark green bars show the density for those who started contributing in 1967.

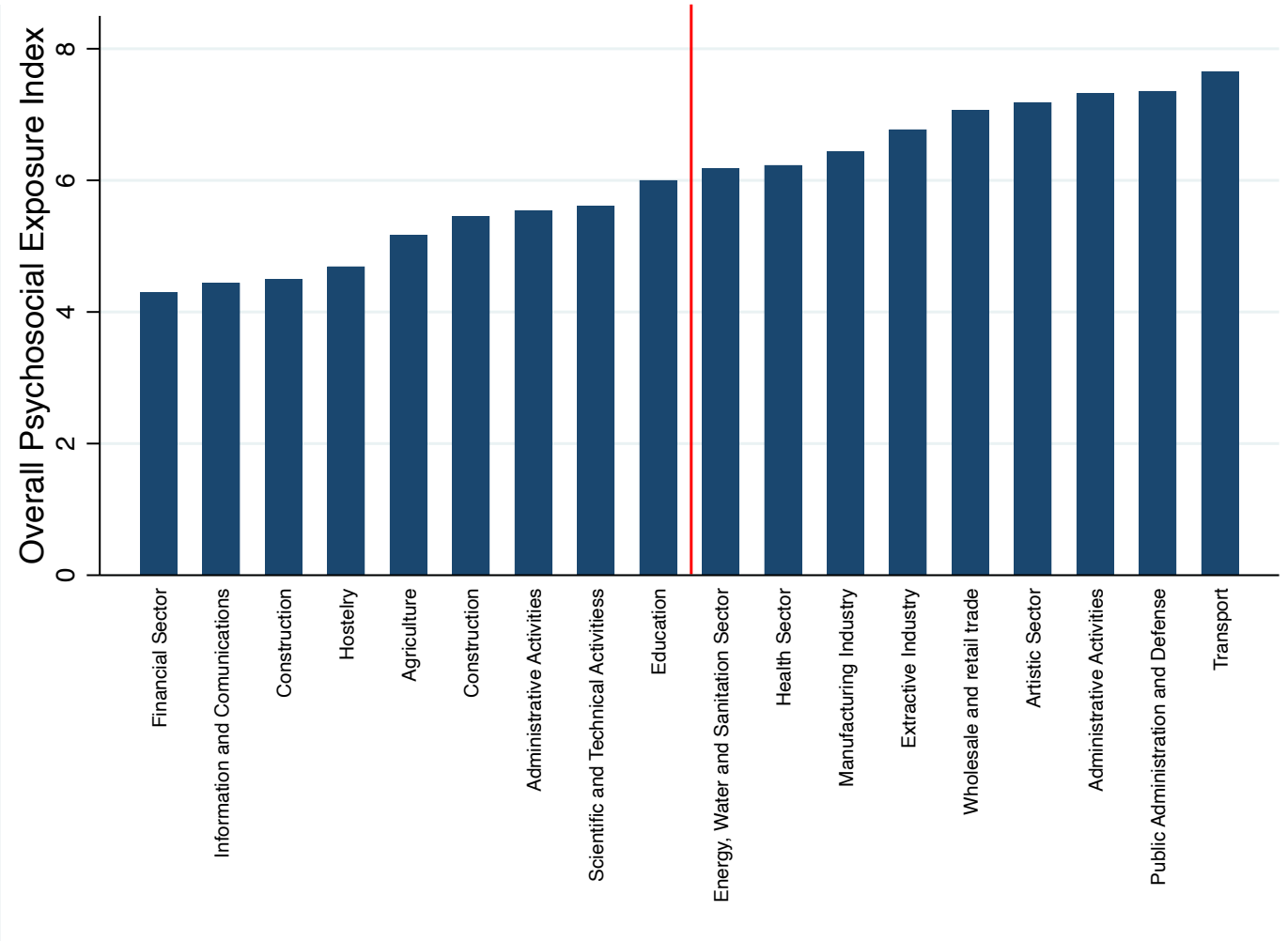
Figure A3: Classification of Industries by Incidence of Workplace Accidents



Source: Register of Workplace Accidents 2003-2019, cohorts 1938-1949.

Notes: This figure plots the total number of workplace accidents between 2003 and 2019 for workers born between 1938 and 1949 in the industry sector the workers were working at the moment of the accident.

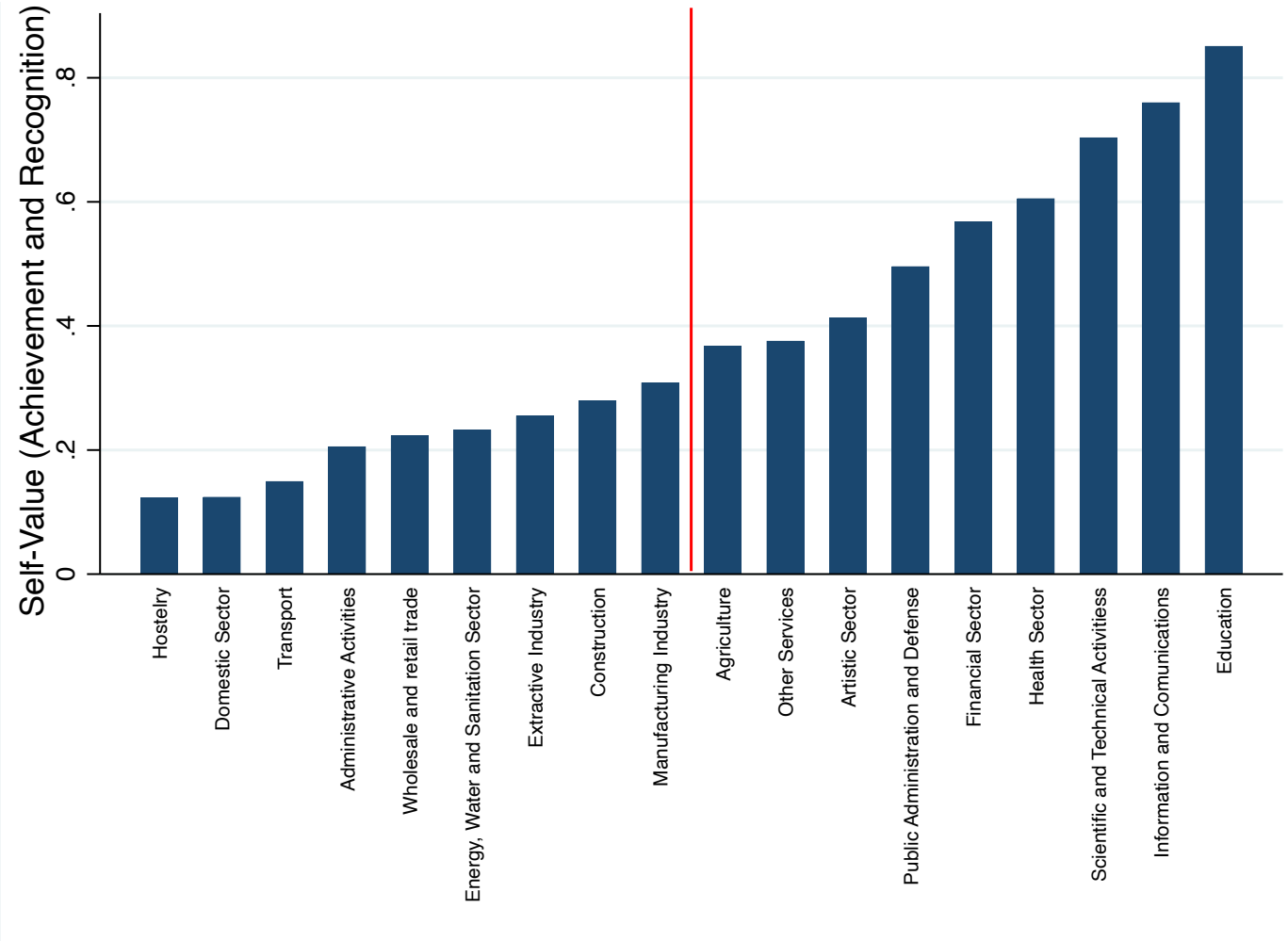
Figure A4: Classification of Industries by Psychosocial Exposure



Source: MCVL, cohorts 1938-1949.

Notes: This figure plots the different industry sectors classified by the degree of psychosocial pressure (mental, social stress, and temporal load) individuals working in these sectors are exposed to. We follow Kroll (2011) for the definition of psychosocial exposure.

Figure A5: Classification of Industries by Self-value Index



Source: MCVL, cohorts 1938-1949.

Notes: This figure plots the different industry sectors classified by the degree of self-value (sense of achievement and recognition) that individuals working in these sectors are exposed to. We follow the O*NET for the definition of the self-value index.

Table A1: Sample Selection

	Sample Selection	
	Observations dropped	Mortality Dropped Obs
Contributed in 1967	-0.021	0.016
	(0.012)	(0.013)
	[0.100]	[0.246]
Month-Year Birth FE	✓	✓
Controls	✓	✓
Contributed 1966-1967	✓	✓
Observations	30,888	6,385
R^2	0.030	0.203
Mean Dep. (Treated)	0.189	0.393
Mean Dep. (Control)	0.220	0.313

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the probability of not being in the main sample due to having stopped contributing to the Social Security system before age 50 or not having at least 8 years of contribution (Column1). Column 2 reports the effect of the reform on mortality for the sample of individuals dropped from the main sample, obtained from the estimation of regression 1. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. All standard errors are clustered at the birth year level, and wild-bootstrap p-values are reported in brackets.

Table A2: Smoothness of the Covariates

	Labor Market between the Ages of 30 and 40				
	Fraction active	Fraction employed	Blue-collar occ	Av. monthly contribution	Fraction selfempoyed
Contributed in 1967	0.108	0.581	-0.007	30.190*	2.602***
	(0.362)	(0.519)	(0.009)	(16.300)	(0.537)
	[0.754]	[0.344]	[0.493]	[0.098]	[0.002]
Month-Year Birth FE	✓	✓	✓	✓	✓
Observations	25,903	25,903	25,903	25,903	25,903
R^2	0.170	0.193	0.063	0.243	0.006

	Industry between the Ages of 30 and 40					
	Agriculture Minery Construction	Manufacturing	Trade Transportation	Public Health Education	Science Administrative	Services Hostelry Housekeeping
Contributed in 1967	0.002	-0.004	-0.002	-0.006	-0.000	0.000
	(0.007)	(0.002)	(0.002)	(0.007)	(0.001)	(0.003)
	[0.769]	[0.128]	[0.382]	[0.374]	[0.847]	[0.912]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓
Observations	25,903	25,903	25,903	25,903	25,903	25,903
R^2	0.045	0.012	0.008	0.044	0.004	0.004

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on a list of predetermined variables: fraction of time spent active (Column 1), the fraction of time spent employed (Column 2), probability of having been employed in a blue-collar occupation (Column 3), average monthly contribution (Column 4), the fraction of time self-employed (Column 5), and probability of being employed in the agriculture, minery or construction sectors (Column 6), manufacturing sector (Column 7), trade or transportation sectors (Column 8), public, health or educational sectors (Column 9), scientific or administrative sectors (Column 10), or services, hostelry or housekeeping sectors (Column 11). The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. All standard errors are clustered at the birth year level, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A3: Descriptive Statistics

Dependent Variables	N	Mean	SD	Min	Max
Regular Pension	25,903	0.47	0.49	0	1
Partial Pension	25,903	0.04	0.19	0	1
Disability Pension	25,903	0.33	0.47	0	1
No Pension	25,903	0.14	0.35	0	1
Age Last Employment	25,903	59.58	5.43	18.75	82.91
Age First Pension	22,040	60.95	4.25	50	79.41
Age Regular Pension	12,367	63.58	2.29	60	74
Age Disability Pension	8,633	57.16	3.76	50	79.41
Age Partial Pension	1,040	61.11	1.33	60	70.08
Dying 50-86 y.o.	25,903	0.32	0.46	0	1
Dying 50-59 y.o.	25,903	0.07	0.26	0	1
Dying 60-69 y.o.	23,987	0.11	0.31	0	1
Dying 70-79 y.o.	21,009	0.16	0.37	0	1
Dying 80-86 y.o.	17,516	0.02	0.16	0	1

Source: MCVL, cohorts 1938-1949.

Notes: This table reports summary statistics for the main outcome variables. The sample corresponds to individuals born between 1938 and 1949, registered in the Social Security (contributive workers and pensioners) at any point of their lives till 2020. We further restrict the same to individuals contributing to the Social Security system at age 50 with at least 8 years of employment.

Table A4: Impact of the Reform on the Type of Disability

	Type of Disability	
	Severe or Absolute	Partial or Professional
Contributed in 1967	0.031**	0.027*
	(0.012)	(0.013)
	[0.013]	[0.066]
Month-Year Birth FE	✓	✓
Controls	✓	✓
Contributed 1966-1967	✓	✓
Observations	25,903	25,903
R^2	0.040	0.051
Mean Dep. (Treated)	0.180	0.207
Mean Dep. (Control)	0.133	0.162

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the probability of claiming absolute or severe disability (Column 1) and partial or professional disability (Column 2), obtained from the estimation of regression 1. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets. *** p<0.01, ** p<0.05, * p<0.1.

Table A5: Impact of the Reform on Reason for No Pension

		Reason for No Pension		
	No Pension	Still Working	Became Inactive	Died before Pension
Contributed in 1967	0.025***	0.002*	0.007	0.016***
	(0.007)	(0.001)	(0.005)	(0.005)
	[0.003]	[0.091]	[0.141]	[0.007]
Month-Year Birth FE	✓	✓	✓	✓
Controls	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓
Observations	25,903	25,903	25,903	25,903
R^2	0.054	0.007	0.080	0.031
Mean Dep. (Treated)	0.167	0.004	0.065	0.098
Mean Dep. (Control)	0.137	0.003	0.064	0.070

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the probability of leaving the labour market without claiming any pension (Column 1), continuing working (Column 2), becoming inactive (Column 3), and dying before claiming a pension (Column 4), obtained from the estimation of regression 1. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A6: Impact of the Reform on Age at Last Employment (in Brackets)

	Last Employment at Age								
	50-54	55-59	60	61	62	63	64	65	After 65
Contributed in 1967	0.011	-0.011*	-0.041***	-0.011**	-0.011**	-0.011***	0.006**	0.070***	0.016***
	(0.008)	(0.005)	(0.013)	(0.005)	(0.004)	(0.003)	(0.002)	(0.012)	(0.004)
	[0.222]	[0.068]	[0.009]	[0.026]	[0.011]	[0.005]	[0.023]	[0.001]	[0.004]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓	✓	✓	✓	✓
Observations	25,903	25,903	25,903	25,903	25,903	25,903	25,903	25,903	25,903
R^2	0.033	0.041	0.020	0.007	0.007	0.012	0.009	0.076	0.049
Mean Dep. (Treated)	0.198	0.247	0.059	0.047	0.041	0.046	0.054	0.240	0.086
Mean Dep. (Control)	0.176	0.253	0.109	0.060	0.054	0.060	0.049	0.166	0.066

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the probability of leaving the labour market between the ages of 50-54 (Column 1), 55-59 (Column 2), at 60 (Column 3), at 61 (Column 4), at 62 (Column 5), at 63 (Column 6), at 64 (Column 7), at 65 (Column 8), and after age 65 (Column 9), obtained from the estimation of regression 1. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A7: Impact of the Reform on Age at Regular Pension (in Brackets)

	Regular Pension at Age						
	60	61	62	63	64	65	After 65
Contributed in 1967	-0.100***	-0.020***	-0.032***	-0.032***	-0.020***	0.095***	0.036***
	(0.024)	(0.005)	(0.005)	(0.006)	(0.005)	(0.014)	(0.008)
	[0.001]	[0.005]	[0.001]	[0.001]	[0.001]	[0.001]	[0.003]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓	✓	✓
Observations	25,903	25,903	25,903	25,903	25,903	25,903	25,903
R^2	0.088	0.029	0.023	0.028	0.019	0.107	0.139
Mean Dep. (Treated)	0.030	0.014	0.007	0.009	0.020	0.284	0.159
Mean Dep. (Control)	0.149	0.041	0.042	0.045	0.043	0.188	0.112

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the probability of claiming a regular pension between the ages of 50-54 (Column 1), 55-59 (Column 2), at 60 (Column 3), at 61 (Column 4), at 62 (Column 5), at 63 (Column 6), at 64 (Column 7), at 65 (Column 8), and after age 65 (Column 9), obtained from the estimation of regression 1. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A8: Impact of the Reform on Age at Disability Pension (in Brackets)

	Disability at Age			
	50-54	55-59	60-65	After 65
Contributed in 1967	0.012	0.017	0.028***	0.000*
	(0.009)	(0.012)	(0.005)	(0.000)
	[0.170]	[0.176]	[0.001]	[0.092]
Month-Year Birth FE	✓	✓	✓	✓
Controls	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓
Observations	25,903	25,903	25,903	25,903
R^2	0.026	0.045	0.017	0.006
Mean Dep. (Treated)	0.113	0.167	0.106	0.001
Mean Dep. (Control)	0.091	0.136	0.068	0.001

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the probability of claiming a disability pension between the ages of 50-54 (Column 1), 55-59 (Column 2), 60-56 (Column 3), and after age 65 (Column 4), obtained from the estimation of regression 1. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A9: Impact of the Reform on Pension Benefit by Type of Pension

	Regular Pensions				Disability Pensions				Partial Pensions				No Pension
	Mean Benefit	Base Benefit	Perc Base	Proxy Base	Mean Benefit	Base Benefit	Perc Base	Proxy Base	Mean Benefit	Base Benefit	Perc Base	Proxy Base	Proxy Base
Contributed in 1967	73.70***	31.30**	8.54***	46.192***	-26.28**	-5.04	0.97	-31.222**	-6.67	-5.04	-0.10	-40.421	-41.402
	(11.75)	(11.10)	(1.59)	(12.283)	(8.81)	(23.93)	(0.61)	(13.182)	(22.20)	(23.93)	(0.31)	(22.529)	(25.196)
	[0.00]	[0.02]	[0.00]	[0.012]	[0.02]	[0.81]	[0.13]	[0.063]	[0.74]	[0.81]	[0.76]	[0.112]	[0.159]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Observations	12,007	12,007	12,007	12,367	8,434	1,038	8,434	8,633	1,038	1,038	1,038	1,040	3,863
R^2	0.373	0.397	0.358	0.355	0.397	0.501	0.051	0.467	0.482	0.501	0.288	0.386	0.485
Mean Dep. (Treated)	1026.6	1041.36	88.30	1089.888	1252.7	1852.11	84.03	1235.633	1545.39	1852.11	81.72	1684.707	1231.647
Mean Dep. (Control)	937.0	1014.31	75.5	1028.816	1279.64	1856.49	84.01	1277.667	1544.27	1856.49	81.19	1715.096	1186.478

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on monthly pension benefit (Columns 1, 5, and 9), pension base (Column 2, 6, and 10), the pension adjustment factor (Column 3, 7 and 11), and the proxy of the pension base (calculated using years of contribution for those individuals that claimed regular pension and total years of activity for the rest) by type of pension claimed by the individual, obtained from the estimation of regression 1. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A10: Robustness: Age Start FE

	Type of Pension				Age at			
	Regular Pension	Partial Pension	Disability Pension	No Pension	Last Employment	Regular Pension	Disability Pension	Partial Pension
Contributed in 1967	-0.097***	0.025***	0.051**	0.020**	0.529***	1.362***	0.329***	-0.241**
	(0.032)	(0.006)	(0.021)	(0.009)	(0.165)	(0.154)	(0.109)	(0.103)
	[0.014]	[0.005]	[0.024]	[0.039]	[0.021]	[0.001]	[0.040]	[0.048]
Age Start Contributing FE	✓	✓	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓	✓	✓	✓
Observations	25,903	25,903	25,903	25,903	25,903	12,367	8,633	1,040
R^2	0.136	0.069	0.086	0.054	0.081	0.219	0.033	0.228
Mean Dep. (Treated)	0.398	0.048	0.387	0.167	59.858	64.641	57.352	61.101
Mean Dep. (Control)	0.533	0.035	0.296	0.137	59.401	63.036	56.986	61.134

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the probability of leaving the labour market through regular pension (Column 1), partial pension (Column 2), disability pension (Column 3), not claiming any pension (Column 4), age at last employment (Column 5), age at claiming regular pension (Column 6), age at claiming disability pension (Column 7), and age at claiming partial pension (Column 8), using age at first contribution fixed effects. This robustness check estimates the impact of losing access to early retirement for people that start working at the same age. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender and age at first contribution fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the age of the first contribution, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A11: Robustness: Cohorts 1941 to 1949

	Type of Pension				Age at			
	Regular Pension	Partial Pension	Disability Pension	No Pension	Last Employment	Regular Pension	Disability Pension	Partial Pension
Contributed in 1967	-0.156***	0.025***	0.095***	0.035***	0.395***	1.666***	0.187*	-0.163**
	(0.022)	(0.006)	(0.016)	(0.006)	(0.130)	(0.060)	(0.095)	(0.054)
	[0.002]	[0.002]	[0.000]	[0.004]	[0.006]	[0.000]	[0.262]	[0.035]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓	✓	✓	✓
Observations	18,288	18,288	18,288	18,288	18,288	9,605	5,340	995
R^2	0.130	0.061	0.087	0.063	0.074	0.269	0.042	0.200
Mean Dep. (Treated)	0.408	0.068	0.371	0.153	60.080	64.832	57.307	61.014
Mean Dep. (Control)	0.601	0.046	0.241	0.113	59.591	62.860	57.056	61.097

Source: MCVL, cohorts 1941-1949.

Notes: This table reports the impact of the reform on the probability of leaving the labour market through regular pension (Column 1), partial pension (Column 2), disability pension (Column 3), not claiming any pension (Column 4), age at last employment (Column 5), age at claiming regular pension (Column 6), age at claiming disability pension (Column 7), and age at claiming partial pension (Column 8), obtained from the estimation of regression 1 restringing the sample to cohorts born between 1941 and 1949. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A12: Robustness: No Selfemployed

	Type of Pension				Age at			
	Regular Pension	Partial Pension	Disability Pension	No Pension	Last Employment	Regular Pension	Disability Pension	Partial Pension
Contributed in 1967	-0.110***	0.024***	0.049**	0.037***	0.470***	1.563***	0.314***	-0.152**
	(0.030)	(0.006)	(0.021)	(0.007)	(0.098)	(0.169)	(0.077)	(0.053)
	[0.002]	[0.000]	[0.034]	[0.000]	[0.004]	[0.000]	[0.026]	[0.045]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓	✓	✓	✓
Observations	21,468	21,468	21,468	21,468	21,468	9,337	7,231	1,037
R^2	0.160	0.067	0.113	0.068	0.057	0.185	0.032	0.245
Mean Dep. (Treated)	0.338	0.060	0.394	0.208	59.141	64.334	57.141	61.099
Mean Dep. (Control)	0.499	0.041	0.299	0.161	58.797	62.618	56.792	61.131

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the probability of leaving the labour market through regular pension (Column 1), partial pension (Column 2), disability pension (Column 3), not claiming any pension (Column 4), age at last employment (Column 5), age at claiming regular pension (Column 6), age at claiming disability pension (Column 7), and age at claiming partial pension (Column 8), obtained from the estimation of regression 1 restringing the sample to individuals that are not in one of the self-employed pension regimes. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A13: Robustness: No Correction of the Year Start Contributing

	Type of Pension				Age at			
	Regular Pension	Partial Pension	Disability Pension	No Pension	Last Employment	Regular Pension	Disability Pension	Partial Pension
Contributed in 1967	-0.177***	0.075***	0.074***	0.029***	0.676***	1.432***	0.256**	-0.323***
	(0.028)	(0.013)	(0.012)	(0.005)	(0.201)	(0.269)	(0.108)	(0.069)
	[0.000]	[0.000]	[0.000]	[0.001]	[0.007]	[0.002]	[0.040]	[0.001]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓	✓	✓	✓
Observations	54,692	54,692	54,692	54,692	54,692	38,296	9,069	3,464
R^2	0.061	0.075	0.028	0.038	0.058	0.172	0.029	0.207
Mean Dep. (Treated)	0.600	0.097	0.214	0.088	61.149	64.000	57.496	61.207
Mean Dep. (Control)	0.759	0.043	0.137	0.060	60.437	62.534	57.189	61.400

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on the probability of leaving the labour market through regular pension (Column 1), partial pension (Column 2), disability pension (Column 3), not claiming any pension (Column 4), age at last employment (Column 5), age at claiming regular pension (Column 6), age at claiming disability pension (Column 7), and age at claiming partial pension (Column 8), obtained from the estimation of regression 1 without correcting for the year of starting contributing reported in the affiliation data. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A14: Impact of Age at Last Employment on Mortality at Five-year Intervals

	Probability of Dying between the Ages							
	50-86	50-54	55-59	60-64	65-69	70-74	75-79	80-86
OLS:								
Impact of Age at Last Employment	-0.013***	-0.007***	-0.006***	-0.002***	-0.002***	-0.002***	-0.002***	-0.000
	(0.001)	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.622]
Reduced Form:								
Contributed in 1967	0.030***	0.003	0.007	0.018***	0.007**	0.001	0.001	0.004*
	(0.009)	(0.002)	(0.005)	(0.004)	(0.002)	(0.006)	(0.005)	(0.002)
	[0.002]	[0.155]	[0.187]	[0.004]	[0.011]	[0.901]	[0.897]	[0.059]
IV:								
Impact of Age at Last Employment	0.074**	0.008	0.015	0.038***	0.013**	0.001	0.001	0.006**
	(0.031)	(0.006)	(0.011)	(0.009)	(0.006)	(0.010)	(0.008)	(0.003)
	[0.019]	[0.190]	[0.204]	[0.004]	[0.048]	[0.902]	[0.898]	[0.036]
Observations	25,903	25,903	25,223	23,987	22,383	21,009	19,174	17,516
Mean Dep. (Treated)	0.369	0.030	0.057	0.082	0.069	0.095	0.099	0.036
Mean Dep. (Control)	0.297	0.023	0.044	0.056	0.056	0.083	0.078	0.023
F-stat FS	35.762	35.762	42.753	59.364	55.275	56.135	65.193	68.987

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of age at last employment on the probability of dying between the ages of 50-86 (Column 1), 50-54 (Column 2), 55-59 (Column 3),60-64 (Column 4), 65-69 (Column 5), 70-74 (Column 6), 75-79 (Column 7), and 80-86 (Column 8). The first panel reports the correlation of age at last employment on mortality (OLS), and the second panel shows the effect of the reform on mortality (reduced form effect using regression 1). The IV estimates, obtained from the estimation of regression 2, are reported in the third panel. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. The IV estimation also controls for the proxy of the pension base. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A15: Impact of the Reform on Labour Market Outcomes between the Ages of 45 and 55

	Labor Market between the Ages of 45 and 55				
	Fraction active	Fraction employed	Blue-collar occ	Av. monthly contribution	Fraction selfempoyed
Contributed in 1967	0.410	1.858**	0.015***	5.109	-0.676*
	(0.230)	(0.647)	(0.004)	(13.201)	(0.364)
	[0.108]	[0.010]	[0.009]	[0.710]	[0.093]
Year Birth FE	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓
Observations	25,903	25,903	25,903	25,903	25,903
R^2	0.116	0.128	0.411	0.432	0.304
Mean Dep. (Treated)	95.656	86.537	0.469	1165.975	17.623
Mean Dep. (Control)	92.975	81.659	0.436	1099.951	15.984

	Industry between the Ages of 45 and 55					
	Agriculture Minery Construction	Manufacturing	Trade Transportation	Public Health Education	Science Administrative	Services Hostelry Housekeeping
Contributed in 1967	0.001	0.001	-0.019***	-0.027**	-0.006	-0.008**
	(0.005)	(0.004)	(0.004)	(0.008)	(0.003)	(0.003)
	[0.823]	[0.885]	[0.002]	[0.019]	[0.102]	[0.027]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓
Observations	25,903	25,903	25,903	25,903	25,903	25,903
R^2	0.249	0.078	0.046	0.094	0.042	0.068
Mean Dep. (Treated)	0.135	0.141	0.088	0.314	0.058	0.030
Mean Dep. (Control)	0.123	0.136	0.110	0.344	0.077	0.044

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of the reform on a list of labour market outcomes when the individual is between 45 and 55 years old: fraction of time spent active (Column 1), the fraction of time spent employed (Column 2), probability of having been employed in a blue-collar occupation (Column 3), average monthly contribution (Column 4), the fraction of time self-employed (Column 5), and probability of being employed in the agriculture, minery or construction sectors (Column 6), manufacturing sector (Column 7), trade or transportation sectors (Column 8), public, health or educational sectors (Column 9), scientific or administrative sectors (Column 10), or services, hostelry or housekeeping sectors (Column 11), obtained from the estimation of regression 1. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A16: Impact of Age at Last Employment on Mortality with Different Controls

	Probability of dying between the ages 60 and 69		
	(1)	(2)	(3)
IV:	0.042***	0.042***	0.049***
Impact of Age at Last Employment	(0.010)	(0.010)	(0.014)
	[0.001]	[0.003]	[0.009]
Month-Year Birth FE	✓	✓	✓
Controls	✓	✓	✓
Contributed 1966-1967	✓	✓	✓
Proxy Pension Base		✓	✓
LM Controls 45-55			✓
Observations	23,987	23,987	23,987
Mean Dep. Variable (Treated)	0.131	0.131	0.131
Mean Dep. Variable (Control)	0.098	0.098	0.098
F-stat FS	50.211	59.364	33.854

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of age at last employment on the probability of dying between the ages of 60 and 69 with no controls (Column 1), controlling for the proxy of the base of the pension benefit (Column 2), and also controlling for the labour market outcomes when the individuals were between 45 and 55 years old (Column 3), obtained from the estimation of regression 2. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.

Table A17: Descriptive Statistics of Compliers

Variables	**Compliers** Mean	**Non-compliers** Mean	**Difference** [P-value]
Frac. Active 45-55	89.78	93.00	2.169**
			[0.021]
Frac. Employed 45-55	62.66	84.43	-16.047***
			[0.000]
Frac. Self-employed 45-55	11.43	17.28	-4.609***
			[0.003]
Mean Contribution 45-55	819.94	1127.45	-189.401***
			[0.000]
Last Industry with High Workplace Accidents	0.33	0.37	0.031*
			[0.075]
Last Industry with High Psychosocial Exposure	0.49	0.51	-0.033*
			[0.079]
Last Industry with High Self-value	0.33	0.28	0.001
			[0.957]
Last Blue-collar Occupation	0.34	0.35	-0.036**
			[0.034]
More than 33 Years of Contribution	0.45	0.48	-0.002
			[0.892]
Observations	2,272	11,968	

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the summary statistics for individuals that are considered compliers and non-compliers. We define as compliers as those individuals that started contributing in 1966 and claimed a regular pension before age 61, while non-compliers are those that also started contributing in 1966 but did not claim a regular pension before 61. The third column shows the coefficient and wild-bootstrap p-value from the regression of all the variables on the dummy variable indicating if the individual is a complier. In this regression we also control for gender, year of birth, and month of birth fixed effects.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A18: Impact on Mortality by Labour Market Conditions Before Retirement: Division by 3 Groups

	Probability of Dying between 60 and 69								
	Workplace Accidents			Psychosocial Exposure			Self-value		
	High	Medium	Low	High	Medium	Low	High	Medium	Low
Reduced Form:									
Contributed in 1967	0.029***	0.025***	0.013	0.027***	0.030***	0.004	0.014	0.031***	0.020***
	(0.008)	(0.008)	(0.007)	(0.006)	(0.007)	(0.010)	(0.009)	(0.009)	(0.006)
	[0.016]	[0.012]	[0.076]	[0.003]	[0.003]	[0.755]	[0.089]	[0.016]	[0.012]
IV:									
Impact of Age at Last Employment	0.068**	0.050**	0.016	0.062***	0.040***	0.008	0.016	0.061**	0.049***
	(0.028)	(0.022)	(0.011)	(0.019)	(0.010)	(0.020)	(0.010)	(0.030)	(0.017)
	[0.043]	[0.026]	[0.139]	[0.002]	[0.000]	[0.731]	[0.110]	[0.059]	[0.010]
Month-Year Birth FE	✓	✓	✓	✓	✓	✓	✓	✓	✓
Controls	✓	✓	✓	✓	✓	✓	✓	✓	✓
Contributed 1966-1967	✓	✓	✓	✓	✓	✓	✓	✓	✓
Observations	7,941	5,133	7,858	9,756	6,736	4,440	5,152	5,993	9,787
Mean Dep. (Treated)	0.158	0.131	0.117	0.149	0.129	0.121	0.121	0.153	0.135
Mean Dep. (Control)	0.117	0.090	0.097	0.105	0.092	0.113	0.097	0.108	0.102
F-stat FS	15.848	13.012	26.495	17.207	24.195	14.418	17.007	10.935	11.446
P-value Diff High-Low		0.086			0.060			0.155	

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of age at last employment on the probability of dying between the ages of 60-and 69 by the labour market conditions experienced by the individual just before retirement. Individual's last industry is classified depending on their workplace accident incidence for our cohorts between 2003 and 2019 (Columns 1, 2 and 3), by the psychosocial exposure (mental stress, social stress, and temporal load) following Kroll (2011) (Columns 4, 5 and 6), and by their self-value index (sense of achievement and recognition) constructed using O*NET (Columns 7, 8 and 9). The first panel reports the effect of the reform on mortality (reduced form effect using regression 1). The IV estimates, obtained from the estimation of regression 2, are reported in the second panel. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. The IV estimation also controls for a proxy of the pension base. At the bottom, we report the First Stage F-statistic and the p-value of the differences between the high and the low groups in the IV estimation. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets. *** p<0.01, ** p<0.05, * p<0.1.

Table A19: Impact on Mortality by Gender

	Age at Last Employment	
	Men	Women
First Stage	0.285**	1.041***
Contributed in 1967	(0.106)	(0.108)
	[0.037]	[0.001]
	Probability of Dying between 60 and 69	
Reduced Form:	0.022***	0.019***
Contributed in 1967	(0.005)	(0.005)
	[0.003]	[0.007]
IV:	0.077***	0.019**
Impact of Age at Last Employment	(0.029)	(0.006)
	[0.005]	[0.017]
Month-Year Birth FE	✓	✓
Controls	✓	✓
Contributed 1966-1967	✓	✓
Observations	17,180	6,807
Mean Dep. (Treated)	0.150	0.057
Mean Dep. (Control)	0.129	0.039
F-stat FS	8.412	102.086
P-value Difference (IV Est.)	0.036	

Source: MCVL, cohorts 1938-1949.

Notes: This table reports the impact of age at last employment on the probability of dying between the ages of 60-and 69 for men (Column 1) and women (Column 2). The first panel reports the first stage of the IV estimation (the reform's effect on the age at last employment, using 1). The second panel shows the second stage; the effect on the probability of dying between 60 and 69 years old. First, we report the reduced form effect of the reform on mortality using regression 1. After that, we report the IV estimates obtained from the estimation of regression 2. The estimation sample includes individuals that started contributing 12 months before and after 1 January 1967. All specifications control for gender, year of birth, and month of birth fixed effects. Each regression also includes the following controls measured when the individuals were between 30 and 40 years old: average monthly contribution, fraction of time employed, fraction of time active, fraction of time in self-employment, and highest occupation and industry sector fixed effects. The IV estimation also controls for a proxy of the pension base. At the bottom, we report the First Stage F-statistic and the p-value of the differences between groups in the IV estimation. All standard errors are clustered at the year of birth, and wild-bootstrap p-values are reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1.