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Universitat Autònoma de Barcelona

WORKING PAPER 24-01

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Abstract

This chapter studies social security reforms and trends in inequalities among older workers over the last decades in Spain. Its main goal is to analyze the redistributive impact of the various pension reforms on older income inequality.

Compared to the rules in 1985, recent pension reforms have led to an average increase on Social Security Wealth of approximately 18,000€ for men and 15,000€ for women. This represents a ten and eight percent increase, respectively. This effect is mostly driven by the mechanical or direct effect (e.g. via benefit adjustments), while changes in retirement probability (secondary or behavioral effect) are close to zero. Furthermore, we find striking differences across income quartiles, for both men and women. In both cases, there is a clear income gradient, where the richest quartile has benefitted the most with an increase close to twenty percent, or over €50,000, for both men and women. Conversely, the change for the poorest income quartile for men and the two poorest income quartiles for women is close to zero or even slightly negative. This is likely due to the effect of minimum benefits (that mark the generosity of the system, see Boldrin et al, 1999) that automatically absorb any other effect for low-income individuals.

Keywords: Social Security, Inequality, pension reforms, life expectancy, Spain.

JEL classifications: H55, D31

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1. Introduction

This chapter examines social security reforms and inequalities among older workers in Spain since 1985. It focuses on the redistributive effects of the different old-age pension reforms on older income inequality, especially social security wealth. While a large body of research examines the sustainability of the Social Security System (see for example Carlos Vidal-Melià, 2024, and the references therein) less attention has been given to the analysis of pension inequality (see for example Arghel et al, 2018) and, particularly it has been shaped by the different social security reforms.

We first review the major social security reforms implemented in Spain since 1985 and discuss their potential impacts on inequality. Next, we describe how inequalities in income and wealth among Spanish seniors have evolved over time. We then use a sample of working histories to evaluate how pension reforms over the last decades have affected the inequality in social security wealth across older workers in Spain. We further explore the drivers behind this change with a focus on changes in benefits and changes in retirement behaviour. In this context, a key contribution of this study is to disentangle to what extent this effect has operated directly, i.e. mechanically due to e.g. changes in pension benefits, or rather via a secondary or behavioral effect driven by changes in retirement or life expectancy patterns.

As a major contribution to previous studies, we incorporate existing differences in life expectancy by socio-economic status (SES) into social security wealth calculations. This is an important component. when assessing the redistributive effects of these pension reforms. Therefore, unlike previous studies for Spain, we compute cohort life expectancies (LE) instead of relying on period life tables. This is important because the systematic gains in LE in Spain may not be equally distributed across SES groups and cohorts.

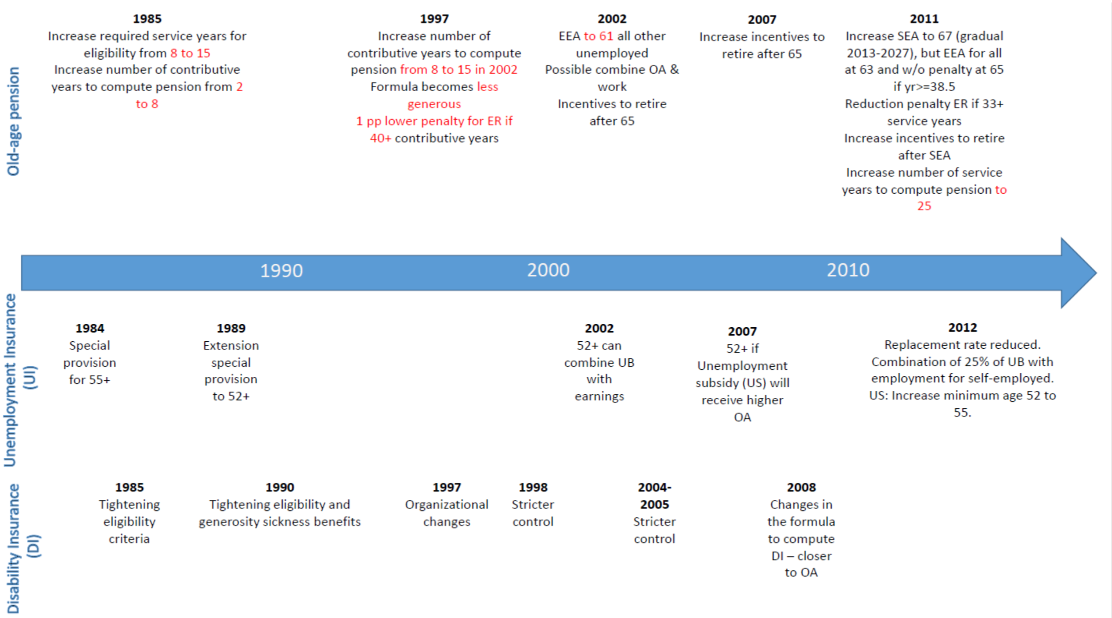
The rest of the paper is organized as follows. Section 2 reviews Spanish social security reforms since 1985. Section 3 introduces the data and social security measures used in the analysis. Section 4 presents the life expectancy analysis. Section 5 presents descriptive evidence on trends in inequality and evaluates the impacts of the recent reforms on inequality. Finally, Section 6 concludes.

2. Social security reforms

Figures 1 and 2[6], illustrate the main characteristics of the reforms in old-age pensions, disability insurance and unemployment insurance from 1985 until 2023. There have been six major old-age pension reforms (1985, 1997, 2002, 2007, 2011/13, 2021/23) and additional changes specific to the disability insurance (DI) or unemployment insurance (UI) programs.

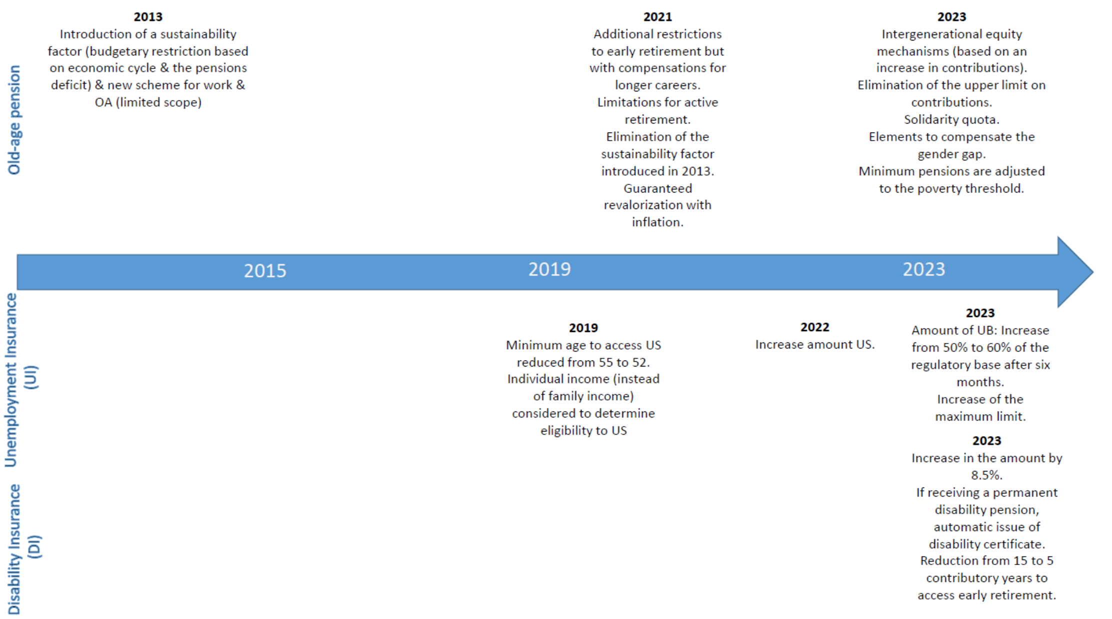
[6] Figures based on Figure A.1 in Garcia-Gómez et al (2024).

Figure 1. Timing and main characteristics of the social security reforms in the period 1985-2011.



Source: Own elaboration based on Garcia-Gómez et al (2024).

Figure 2. Timing and main characteristics of the social security reforms in the period 2012-2023.



Source: Own elaboration based on Garcia-Gómez et al (2024).

2.1. Old-age pension reforms

As described in Boldrin et al. (1999) the transition from the old Mutualidades system to a system of Social Security contributions was completed between 1979 and 1985, with the removal of "bases tarifadas" (fixed covered wages). The crucial ingredients of the system from 1985 were as follows[7]:

- The statutory eligibility age was 65, while the earliest eligibility age was 60 if the individual did not have any job that required an affiliation to the social security system.
- A minimum of 10 years of contributions were required to gain access to a contributive pension.
- The pension was calculated based on three elements: (1) the average of the contributions in the 24 months preceding retirement, (2) the penalty for early retirement (8 percent per year), and (3) the penalty for insufficient contributions (2 percent per year not contributed, full benefit reached with 35 contributory years).

The key elements of the Spanish pension system prevailing until 2011 were set in 1985. Eligibility for the old-age benefits increased from 10 to at least 15 years of contributions. The pension amount was calculated by multiplying a regulatory base by a percentage based on age and the number of years contributed to the system. Under the 1985 regime, the regulatory base was calculated as (wages of the last 96 months (8 years) before retiring)/112. This regulatory base was multiplied by a percentage based on the number of years of contributions (n).

The pension amount was capped from below by the minimum pension (see Jiménez-Martín 2014 for details) and the maximum benefit (between 4 and 5 times the minimum wage).

In 1997, the number of contributory years used to compute the benefit base was progressively increased from 8 to 15 years in 2002, and the formula for the replacement rate was made less generous. On the other hand, the 8% penalty applied to early retirees between the ages of sixty and sixty-five was reduced to 7% for individuals with at least 40 years of contributions at the time of early retirement.

In 2002, further changes were introduced. Before 2002, only individuals who had contributed to the system earlier than 1967 could benefit from early retirement at sixty, while the rest had to wait until the statutory eligibility age of sixty-five. In 2002, early retirement at sixty-one was made available for the rest of the population. At the same time, there was an impulse for partial and flexible retirement with the possibility of combining income from work with old-age benefits and the introduction of incentives for individuals to retire after the statutory eligibility age of 65[8]. At the same time, workers who have contributed for at least 30 years and have been registered in the employment office for the previous 6 months could now access early retirement at sixty-one if they were involuntarily unemployed.

In 2007, the incentives to retire later than 65 were further increased providing an additional three percent, instead of the two percent agreed in 2002. The 8% penalty applied to early retirees between the ages of sixty and sixty-five was reduced to 6-7.5%,

[7] See Boldrin et al. (1999, 2004) for other details regarding disability and survivor pensions.

[8] An additional two percent per additional year of contribution beyond the age of 65 for workers with at least 35 years of contributions on top of the 100% applied to the regulatory base.

depending on the number of years contributed, for those individuals with at least 30 years of contributions. In addition, the contributions for unemployed workers older than fifty-two were increased so that they would receive a higher old-age pension when retiring.

All these reforms aimed to increase labor supply of older male workers. However, existing evidence (see for example Cairó-Blanco 2010 and García-Pérez et al. 2013) suggests there was not a clear link between these reforms and the increased labor supply of older male workers (see for example García-Gómez et al., 2018).

The discouraging demographic and labor market scenarios during the first years of the Great Recession led the Spanish government, under pressure from the EU to reduce the future deficit, to undertake significant reforms of the pension system in 2011. Two main elements were targeted: (1) the number of contributory years entering the pension calculation was increased from 15 to 25, and (2) the statutory eligibility age was raised from 65 to 67, gradually. The latter was particularly relevant for Spain since the statutory eligibility age had not been modified since it was first established in 1979. These two changes severely cut the generosity of the pension system (see Sánchez 2017 for a recent evaluation). The reform also restricted the eligibility conditions for early retirement. The effect of this change on the system's generosity remains unclear because the reform barely changed the eligibility conditions to access the minimum pension. Consequently, workers expecting to receive the minimum pension, i.e. workers with low income and short contributive careers, were less affected by the reform (Jiménez-Martín 2014).

In 2013, the Spanish government amended the 2011 reform attempting to stabilize the short- and long-term financial sustainability of the Social Security system. In particular, this amendment introduced a sustainability factor (SF), which links the initial pension level to the evolution of life expectancy (Conde-Ruiz et al. 2013). This mechanism can be seen as transforming defined benefit schemes into defined contribution schemes.

The SF has two key components, the intergenerational equity factor (IEF) and the pension revaluation index (PRI). The IEF aims to provide equal treatment to those retiring at the same age with the same employment history but with different cohort-specific life expectancy. The introduction of this factor was not controversial, since it was perceived as reasonable that if pensioners were to receive the same total pension throughout their retirement, an individual with a greater life expectancy should receive a little less each year. The second factor, the PRI, fixes a budgetary constraint on the economic cycle and, as such, is relatively flexible in the short term. However, the discretionary rule chosen by the Government guarantees that, even if Social Security revenues are insufficient to cover pension costs, pensions rise each year by at least 0.25%, and by no more than the annual change in the CPI + 0.25%.

The last two reforms were introduced in 2021 and 2023. The 2021 reform introduced stricter restrictions on accessing early retirement while providing compensations for those with longer careers. It also included limitations on active retirement. The sustainability factor introduced in the 2013 reform was eliminated and substituted by a guaranteed revalorization of the amount of the pension benefits with inflation.

The 2023 reform established an intergenerational equity mechanism by increasing contributions and eliminating the upper limit on contributions. It also introduced elements to compensate for the gender gap in pensions and the minimum pensions were adjusted to the poverty threshold.

We expect the early reforms of the Spanish system that progressively increased the number of contributory years to compute the benefit base to be beneficial for individuals with long careers, which are typically those with low skill levels. At the same time, this change may be less beneficial for workers with steeper wage profiles, who are generally high-skill workers. Therefore, we expect those reforms to reduce income inequality. However, women in Spain during those years typically had fewer years of contributions. We would therefore expect an increase in gender inequalities as more women would be left out of the system (as they would not have the required number of years of contributions).

Similarly, we expect the 2021/2023 pension reform to have a positive impact (a reduction) on inequality in Spain by increasing the pension of some of the most disadvantaged individuals, such as those receiving the minimum pension.

2.2. Reforms in the Disability System

Here we focus on some distinctive features of the main reforms since the creation of the National Institute of Social Security (NISS) in 1979.

The first large disability insurance reform took place in 1997 and included 4 main points:

- Sickness benefits: stricter control of the sickness status by Social Security physicians, a reduction of the level of long-term sickness benefits, and the replacement of the old job assessment by a more objective definition of the usual occupation of the individual.
- Permanent disability pensions of individuals aged at least 65 were automatically transferred to the old-age pension system. This was just a change in the classification within the pensions system.
- Organizational reform: all the issues related to disability insurance were transferred to the NISS. In the past, the permanent disability status was assessed and granted by local GP's. This reform established a group of experts (the disability assessment team inside the NISS) responsible for assessing applicants' ability to work based on the available medical files and a medical assessment from NISS physicians.
- The claimant did no longer lose entitlement to non-contributory disability benefits if they started working. They would remain entitled to receive non-contributory disability benefits in case of job loss.

In addition to this major reform in 1997, the 1998 budget law introduced the possibility for NISS physicians and mutual insurance companies to review the health situation and status of beneficiaries. Effectively, only very few claimants in the permanent disability system effectively exit the program.

In 2004 and 2005, monitoring of the use of sick leave was tightened with the creation of a new sub-department at the NISS and a new monitoring tool to reduce absence rates. In 2005, a general absence control was introduced for cases in which absenteeism took longer than six months.

In 2008, the minimum contributory period to access permanent disability pensions was reduced for young workers to adjust for their later entrance into the job market. At the same time, the formula to calculate the regulatory base of the benefit was slightly modified: the regulatory base of permanent disability due to a common illness has since then decreased by 50% if the individual had not contributed at least 15 years and it is lower the further the individual is from age 65.

Last, in 2023 the amount of disability benefits was increased by 8.5%. In addition, those receiving a permanent disability benefit now automatically receive a disability certificate, which grants their other type of rights, like tax breaks. Before 2023, individuals with disabilities had to navigate these two different processes: applying for disability benefits and obtaining a disability certificate. Finally, there was also a reduction from 15 to 5 years of contribution to access early retirement benefits from disability benefits.

All these reforms ensured the financial stability of the disability system in Spain as inflow rates have remained stable, at odds with the dramatic increase experienced by other industrialized countries (see Jiménez-Martín et al. 2018).

The overall effect of these reforms on income inequality is unclear. On one hand, features like the increase in the amount of disability benefits have most likely reduced inequalities in Spain, while those that have tightened the entrance into the system, like in 1997, have probably increased inequalities.

2.3. Reforms in the Unemployment Insurance Scheme

In 1984, the government introduced unemployment benefits for workers employed in temporary contracts, and non-contributory unemployment benefits (also called unemployment assistance benefits). In addition, it established a special provision for workers aged over 55 who were allowed to receive unemployment assistance benefits until the claiming age. To receive these benefits, individuals had to satisfy the entitlement requirements of the retirement pension, except for age. The subsidy paid 75% of the minimum wage until reaching the age to be transferred to the old-age pension system. Furthermore, the years in unemployment under this special scheme were counted as contributive years towards an old-age benefit.

In 1989, the special provision of unemployment assistance benefits until the statutory eligibility age of 65 for individuals aged at least 55 was extended to individuals aged 52, thus increasing the incentives for older workers to leave the labor market at younger ages.

The reform in 2002 opened the possibility for individuals aged at least 52to combine the unemployment insurance benefits with earnings. They could receive 50% of their previous unemployment insurance entitlement, and the employer would pay the remaining amount in wages. The reform also introduced an extension of the program to help vulnerable groups to integrate in the labor market. These groups included individuals aged 45+ who have been unemployed for one month and people with disabilities, among others.

In 2012, the amount an individual receives from unemployment insurance after the first six months was reduced from 60 to 50 percent of previous earnings. This applied to all unemployment spells starting after the 15th of July 2012. For the first six months, it was kept constant at 70%. Individuals receiving unemployment benefits could combine 25% of the benefits with self-employment. Finally, the last element in this reform package was the increase in the minimum age at which individuals could receive unemployment subsidies until retirement (absorbing state) from 52 years old to 55.

In 2019, the government reduced again the minimum age to get access to unemployment subsidies (absorbing state) from 55 to 52. Additionally, eligibility for the unemployment subsidy scheme was determined based on individual income rather than family income.

In 2022, there was an increase in the unemployment subsidies, and in 2023 the unemployment benefit scheme was increased from 50% to 60% of the regulatory base after the first six months of benefits. As part of the same package of reforms, there was also an increase in the maximum limit for those benefits.

3. Data and incentive measures description

In this paper, we use three datasets. First, we use individual-level data from the Spanish sample of the Survey of Health and Retirement in Europe (SHARE), a multidisciplinary and representative cross-national panel of the European population aged 50 and older[9]. We use data from the first eight waves, collected in approximately biannual periods between 2004/2005 and 2019 (see Börsch-Supan et al., 2013, for a detailed description of SHARE). We exclude waves 3 and 7, as these are retrospective waves. The data includes information on sociodemographic background characteristics, current health, and socioeconomic status. We keep individuals aged between 50 and 70. This selection yields 8,375 observations for 3,337 respondents. We use imputations and unfolding bracket values whenever possible to preserve the sample size. Nevertheless, we lose 1,046 observations corresponding to 211 respondents due to missing values in our analytical variables (mostly in income). Our final sample includes 7,329 observations for 3,126 respondents. Despite the lower sample size, SHARE allows us to estimate trends in inequality on variables like wealth, assets, or total household income.

The second dataset is the Muestra Continua de Vidas Laborales (MCVL), an administrative dataset from the Spanish Social Security Administration. It includes 4% of all individuals who have contributed for at least one day in their careers to the Social Security in Spain. For those individuals, there is retrospective information on the entire labor market career, including a proxy for wages: social security contributions, and personal information such as sex and month and year of birth. 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

[9] SHARE is the European equivalent of the Health and Retirement Survey, a panel dataset of interviewees born in 1960 or earlier and their partners. The use of individual survey data is especially important given that common administrative data lacks important controls for detailed socio-economic and demographic characteristics available in survey data. SHARE interviews all persons aged 50 years and over at the time of sampling[9] who have their regular residence in the respective SHARE country. Individuals are excluded from baseline or refreshment samples if they are incarcerated, hospitalized or out of the country during the entire survey period, unable to speak the country's language(s) or have moved to an unknown address.

each point in time, and the total number of contributive years before retirement[10] (for more information on this database, see Garcia-Perez 2008).

We use this second dataset to estimate the direct and indirect effects of pension reforms on SSW by quartiles of LTE. To do so, we construct a panel of 55–70-year-olds that we use to estimate the retirement hazards in section 6.1. and to compute key pension variables such as SSW and the Implicit Tax Rate (ITR) under the different pension systems. For our panel to be representative of the individuals who retired under the different systems, we pool the 2007-2022 waves of MCVL and drop the information of Social Security contributions records before 1970 because they are not accurate. In each wave, we first select our target population (55–70-year-olds) and then select a random sample of up to 5,000 individuals. Our final panel contains 9,763 individuals with a total 115,536 observations.

Third, we use an extended and restricted sample from the MCVL provided by the Spanish Social Security system as in Belles et al (2022). This dataset contains a 10% random sample of individuals born between 1935 and 1949 who have registered with the Social Security at any point in their lives up until 2020. We further restrict the sample to individuals who survived to at least age 50. Note that the sampling differs from the publicly available version of the MCVL, described above. This allows us to observe contributive workers and pensioners prior to 2005 (the starting time of the publicly available version). In addition to the information available in the public version, this restricted version also contains detailed information on the date of death before 2005. This allows us to observe differences in mortality across cohorts and socioeconomic groups. Our final dataset contains 600,063 individuals.

3.2. Incentive measures

As in Gruber and Wise (1999), we use social security wealth (SSW) as our main measure of financial incentives. It measures the present discounted value of lifetime social security benefits as follows:

$$SSW_{k,t}(R,i) = \sum_{a=R}^{T} B_{k,t,a}(R,i)\sigma_{t,a}\beta^{a-R} \qquad (1)$$

Where i is the type of individual defined by their cohort, gender, skill level, region, and marital status, starting to claim benefits B from program k at age R; $\sigma_{t,a}$ is the survival probability at age a in year t T is the maximum length of life, and β^{a-R} is the discount factor set at a rate of 3%. As described in the next section, we allow $\sigma_{t,a}$ vary by age, sex, and lifetime earnings quartile.

All calculated magnitudes are net of Social Security contributions and personal income taxes. Calculating after-tax social security wealth is complex due to the high number of bend points in the Spanish marginal tax schedule, which has decreased over time (from thirty-four in 1985 to five in 2016). As an approximation, we use the 1995 tax schedule

[10] 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).

to establish the relationship between the average tax rate (net of standard deductions) and income (net of social security contributions paid by the worker). We then estimate this relationship fitting a fourth-order polynomial using ordinary least squares. The estimated coefficients from this model are subsequently used to determine after-tax earnings and benefits for all prior and subsequent years. Section 4 describes how we estimate mortality by sex and income quartile. In the analyses using administrative data from MCVL, we approximate individuals' lifetime income level by their level of contributions. In particular, we sum all the Social Security contributions, excluding self-employment, up to age 50.

In addition, we compute the implicit tax rate on working longer and claiming benefits later (ITAX) as:

$$ITAX_{k,t}(R,i) = \frac{SSW_{k,t+1}(R+1,i) - SSW_{k,t}(R,i)}{Y_{t+1,i}} \qquad (2)$$

where $Y_{t+1,i}$ is after-tax earnings during the additional year at work.

4. Differences in life expectancy by socioeconomic group

In 2023, Spain ranked among the European Union countries with the highest life expectancy, at 84 years (Eurostat, 2023[11]). This increase has been driven by steadily declining mortality rates across all age groups over the past few decades. However, overall changes can obscure differences in mortality trends among various socioeconomic groups, making it essential to examine these disparities. In this section, we first document the existing evidence for Spain on the gradient in life expectancy. We then explore the relationship between socioeconomic status and mortality in Spain using administrative data from the extended sample from the MCVL. Finally, we estimate different mortality rates by sex and income quartile to allow for differences in life expectancy across these groups in the estimated SSW.

For our analysis, we use a representative sample of individuals born between 1935 and 1949 who had at least one interaction with the Social Security system up to September 2023 and survived to at least age 50. Consequently, our mortality and survival measures are conditional on individuals being alive at age 50.

4.1 Previous evidence for Spain

Previous literature has documented the relationship between socioeconomic status and mortality in Spain. For instance, Redondo-Sánchez et al. (2022) use administrative data for the years 2011-2013 to calculate sex and age-specific life tables according to socioeconomic status at the census tract level. The authors combine data from the mortality and population registers provided by the Spanish Statistical Office and. They combine information on high school dropout rates, unemployment rates and share of dwellings in buildings in bad conditions to construct a deprivation index by census tractThey find that life expectancy at birth is 3.2 (men) and 3.8 (women) years shorter for those living in the most deprived census tract compared to their neighbors in less deprived regions.

[11] Eurostat database: https://ec.europa.eu/eurostat/web/population-demography/demography-population-stock-balance/database?node_code=demo_r_mlifexp

González and Rodríguez-González (2021) use Spanish data from the death and population registers and combine it with census data for the years 1990-92, 2000-02, and 2016-2018 to construct mortality rates at the municipality of residence level. Death rates are calculated by sex, five-year age group and calendar year. They construct a deprivation index at the municipality level using same characteristics as Redondo-Sánchez et al (2022) for each year. Their results show that mortality rates have decreased over the period of study for all age groups. These reductions have been stronger for men and children. The authors provide evidence that inequality in mortality has increased over the period for adult and older men and this increase is explained by stronger declines in cancer-related mortality in less-deprived areas.

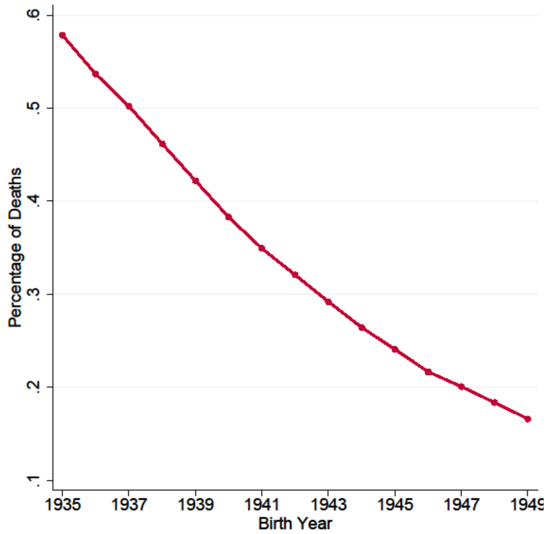
Bilal et al. (2019) use a different dataset to estimate the association between individual-level socioeconomic status, on the one hand, and life expectancy and mortality, on the other. They use data from the Catalan Health Surveillance System (CCHS), which includes information on sociodemographic characteristics, medical diagnoses and mortality for all residents in Catalonia during 2016. The authors create four categories of socioeconomic status based on annual income and calculate age-adjusted mortality by sex and socioeconomic status. Their results show that men and women of the lowest socioeconomic status have 12.0 and 9.4 years lower life expectancy when compared to their counterparts in the highest socioeconomic category.

These studies collectively underscore the persistent and substantial differences on life expectancy by socioeconomic status in Spain. Our main contribution in this respect is to use a cohort life table stratified by a measure of lifetime earnings not yet influenced by retirement decision.

4.2 Descriptive Evidence

We first describe the probability of dying after the age of 50 up to 2023 for the different cohorts in our dataset. Figure 3 illustrates the percentage of individuals who had died between the age of 50 and 2023 for each cohort. Notably, nearly 60% of individuals from the 1935 cohort had died by 2023, when they were 88 years old. In stark contrast, only 17% of the 1949 cohort had passed away by 2023, when they were 74 years old.

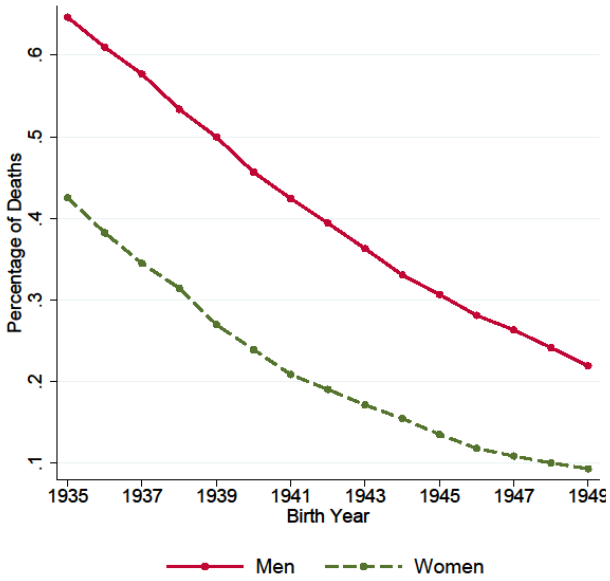
Figure 3. Fraction of individuals that have died per cohort between the age of 50 and September 2023.



Source: Own elaboration using data from MCVL, cohorts 1935-1949.Notes: This figure reports the percentage of individuals in each cohort that died between the age of 50 and September 2023.

However, this figure masks significant gender differences. Figure 4 reveals that the probability of dying is consistently higher for men than for women across all cohorts. For individuals born in 1935, 65% of men had died by the end of our observation period, compared to only 42% of women. Additionally, the gender gap in mortality rates decreases for younger cohorts. While the differential was as high as 23 percentage points for the 1935 cohort, it narrowed to 12 percentage points for the 1949 cohort.

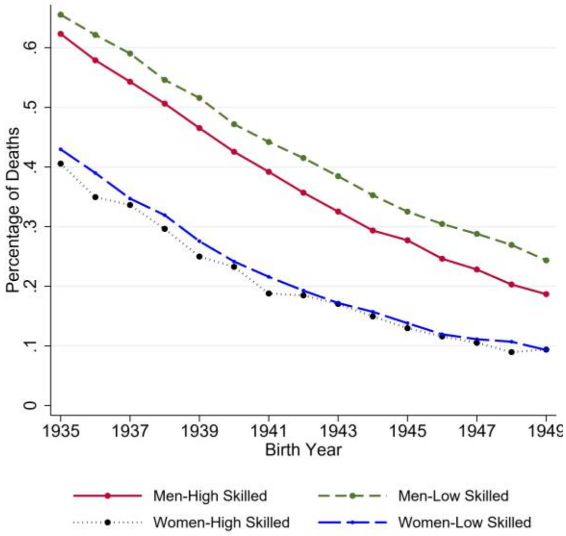
Figure 4. Fraction of individuals that have died per cohort and gender between the age of 50 and September 2023.



Source: Own elaboration using data from MCVL, cohorts 1935-1949. Notes: This figure reports the percentage of men and women in each cohort that died between the age of 50 and September 2023.

There are also important differences by skill level and income. We define skill level based on the highest occupation an individual eld before any decision regarding permanent exit from the labor market is taken , that is between ages 40 and 50. High-skilled individuals are those whose highest occupation was a white-collar job (including engineers, graduates, senior management, technical engineers, experts, administrative and workshop managers, and various administrative roles), while low-skilled individuals are those whose highest occupation was a blue-collar job (including first and second-class officers, third-class officers and specialists, and unqualified workers over 18). As expected, Figure 5 shows that both men and women with high-skilled occupations before retirement have a lower probability of dying. Importantly, this effect is more pronounced for men than for women.

Figure 5. Fraction of individuals that have died per cohort and occupational skill level between the age of 50 and September 2023.

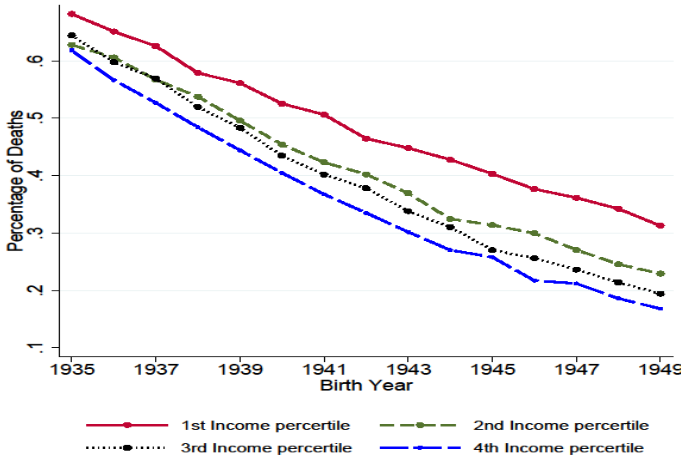


Source: Own elaboration using data from MCVL, cohorts 1935-1949.
Notes: This figure reports the percentage of men and women in each cohort that died between the age of 50 and September 2023.

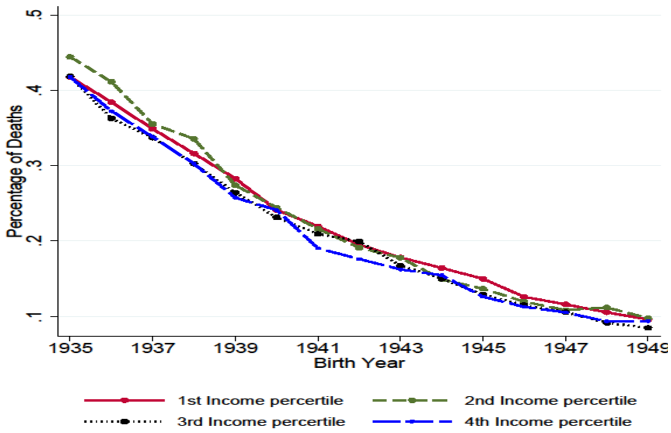
A similar pattern emerges when examining differences by income, measured by the sum of Social Security contributions up to age 50 and divided into four quartiles, rather than occupation. Figure 6 displays the percentage of men and women who died during our observation period by income quartile. The first quartile represents individuals with the lowest contributions, while the fourth quartile corresponds to those with the highest contributions. The data reveal larger differences on mortality by income for men than for women. Men with the lowest contributions before retirement have a higher probability of dying across all cohorts. This income gradient is particularly steep for younger cohorts, indicating that income-related health disparities among men have increased over time. In contrast, we do not observe such disparities for women. This may be due to lower income variability among working women in these cohorts compared to men.

Figure 6. Fraction of individuals that have died per cohort and contribution quartile between the age of 50 and September 2023.

a) Men



b) Women



Source: Own elaboration using data from MCVL, cohorts 1935-1949.
Notes: This figure reports the percentage of men and women in each cohort that died between the age of 50 and September 2023, categorized by their level of contributions (sum of contributions up to age 50). The first income percentile represents individuals with the lowest level of contributions, while the fourth income percentile represents those with the highest level of contributions.

4.3. Regression analysis

We analyze the conditional probability of mortality using a Gompertz distribution within a parametric proportional hazard model framework. The Gompertz distribution, widely used by medical researchers and biologists for modeling mortality data, is a two-parameter function (Lee and Wang, 2013). The hazard and survivor functions of the Gompertz distribution are defined as follows:

$$h(t) = \lambda \, exp(\gamma t) \tag{3}$$

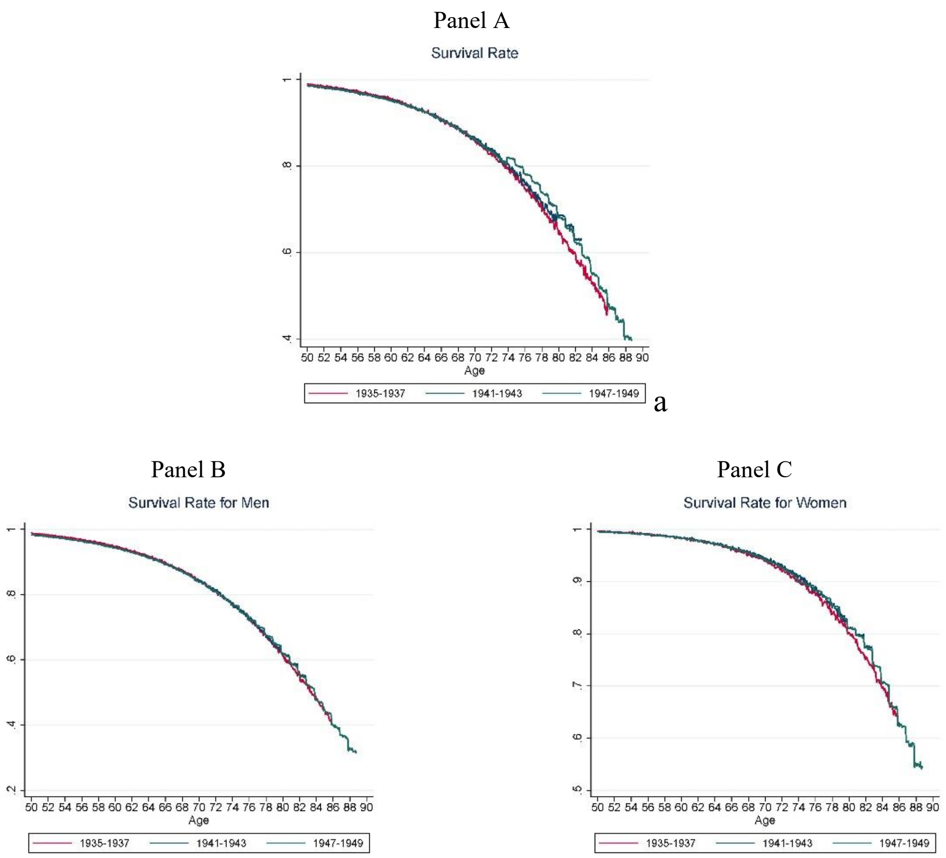
$$S(t) = exp\big(-\lambda\gamma^{-1}(exp(\gamma t) - 1)\big) \tag{4}$$

with λ = exp(x'β) and γ being the ancillary parameter estimated from the data. The vector x includes the following covariates: cohort dummies, sex, first year of contribution, minimum and maximum group of contribution, number of days active in the labor market between ages 30 and 50, number of days employed between ages 30 and 50, fraction of time worked in the construction and mining sectors between ages 30 and 50, and fraction of time worked as self-employed between ages 30 and 50.

In the following figures, we present the predicted survival functions for three birth cohorts: 1935-37, 1941-43, and 1947-49. Consistent with our descriptive analysis, the results are disaggregated by sex, occupational skill level, and income quartiles.

Figure 7a displays the predicted survival rates for all individuals across the three cohort groups. The results corroborate our findings from the descriptive analysis: even when controlling for various labor market factors, younger cohorts exhibit higher survival probabilities, particularly after age 70. Interestingly, as shown in Figures 7b and 7c, these differences across cohorts are more pronounced for women than for men. This suggests that changes in the labor market may have a greater explanatory power for the increased survival rates observed in younger male cohorts compared to their female counterparts.

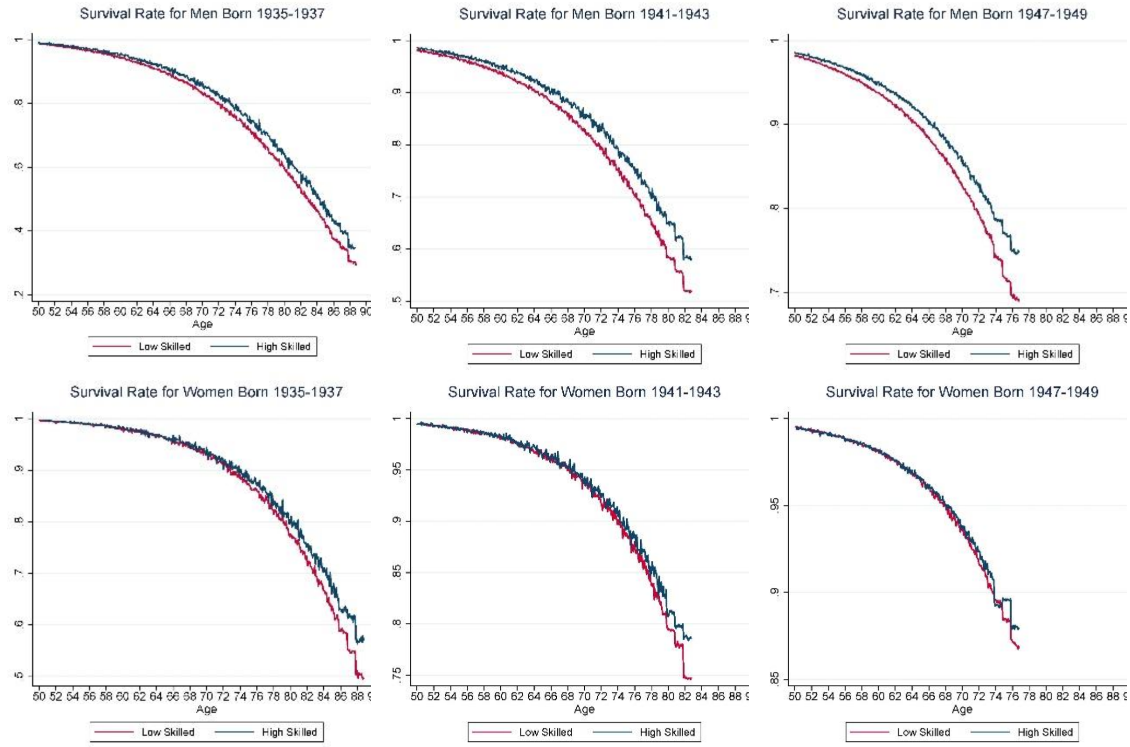
Figure 7. Survival rate (after age 50) by cohort and gender.



Source: Own elaboration using data from MCVL, cohorts 1935-1949.
Notes: This figure reports the predicted survival rate following a Gompertz distribution after the age of 50 for three birth cohort groups (1935-37, 1941-43, and 1947-49) and for men and women separately.

Figure 8 shows the predicted survival rates by occupational skill level, using the same definitions as in the descriptive analysis. For all three cohort groups, individuals holding high-skilled occupations before retirement exhibit higher survival rates, and this holds true for both men and women. However, the trend across cohorts varies by sex. Among men, the survival gap between high-skilled and low-skilled occupations appears to widen in younger cohorts. In contrast, among women, this difference seems to diminish over time.

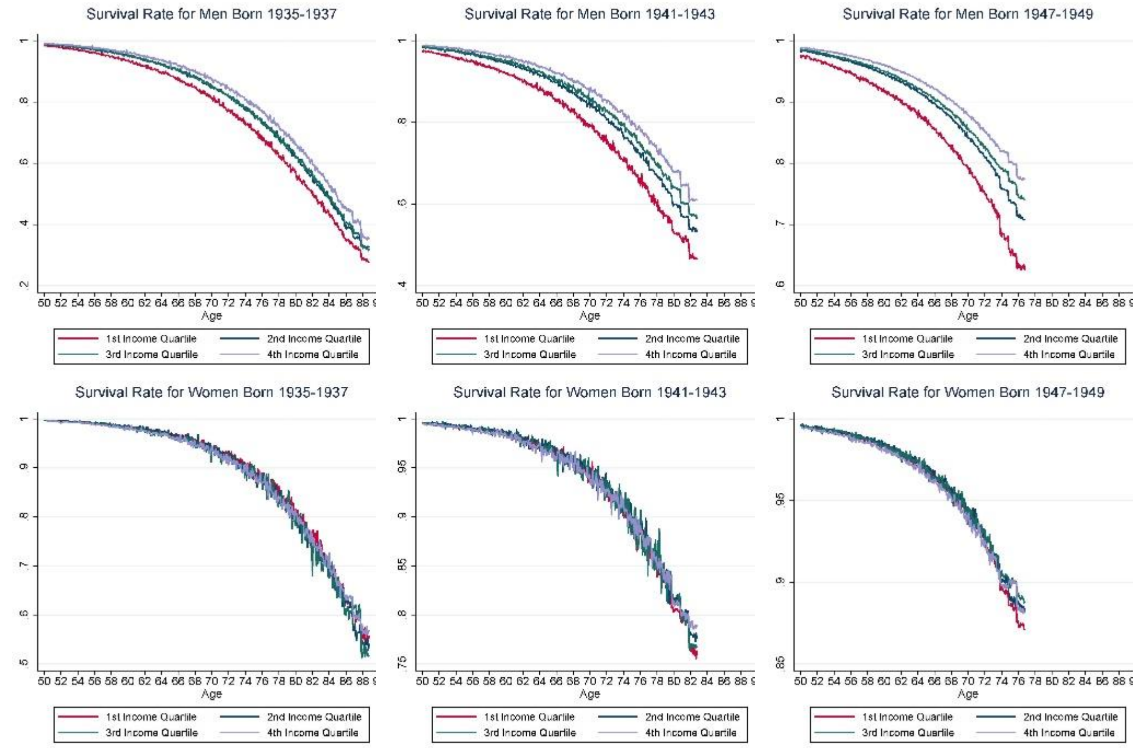
Figure 8. Survival rate (after age 50) by cohort, occupational skill level and sex.



Source: Own elaboration using data from MCVL, cohorts 1935-1949.
Notes: This figure reports the predicted survival rate following a Gompertz distribution after the age of 50 for three birth cohort groups (1935-37, 1941-43, and 1947-49), occupational skill level, and women and men separately.

Finally, we examine how the predicted survival rates differ by income quartiles, proxied by the sum of contributions before age 50. Figure 9 confirms the patterns observed in the descriptive analysis. Among men, survival rates are highly stratified by income percentile, with higher survival rates at all ages for those with higher contributions before retirement. Additionally, this income-related survival inequality is increasing in the younger cohorts. Conversely, for women, there is less of an income gradient in survival rates. However, this appears to be changing in younger cohorts, particularly at older ages.

Figure 9. Survival rate (after age 50) by cohort, contribution quartile and sex.



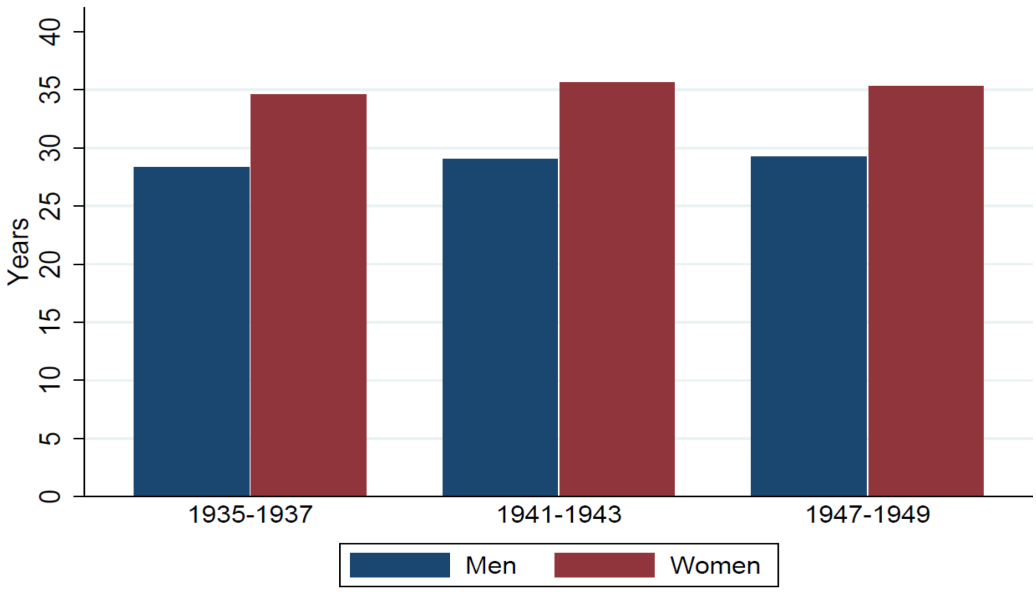
Source: Own elaboration using data from MCVL, cohorts 1935-1949.

Notes: This figure reports the predicted survival rate following a Gompertz distribution after the age of 50 for three birth cohort groups (1935-37, 1941-43, and 1947-49), women and men separately, categorized by their level of contributions (sum of contributions up to age 50). The first income percentile represents individuals with the lowest level of contributions, while the fourth income percentile represents those with the highest level of contributions.

4.4 Summary Results

In this section, we summarize the results by cohort and sex as well as level of contribution. Figure 10 presents life expectancy at age 55 by cohort and sex. Life expectancy increases with the cohort, especially for men. Figures 11 and 12 present life expectancy at age 55 by income quartile for the 1941-43 cohort. We define income quartiles based on cumulated contributions up to age 50. For both men and women, higher income quartiles experience higher life expectancy.

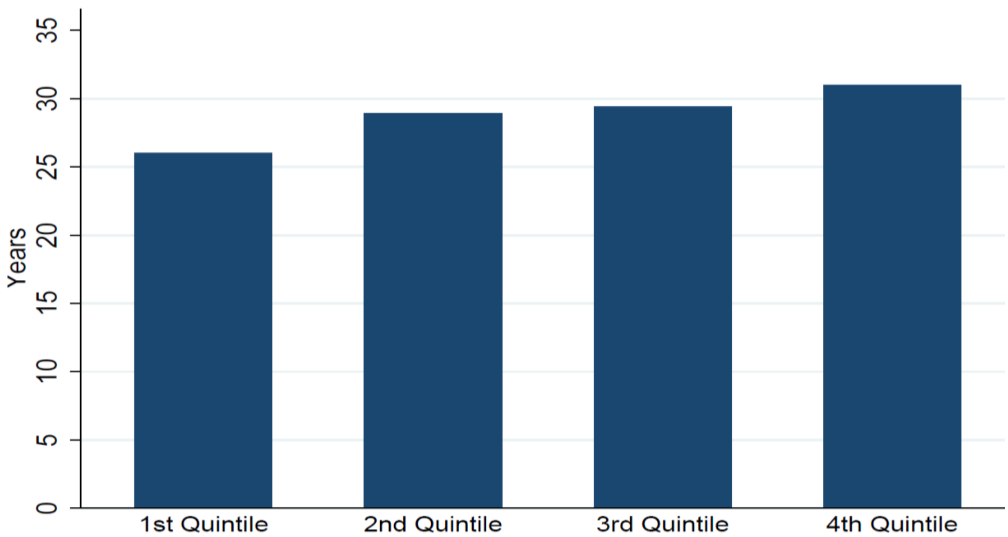
Figure 10. Life expectancy at age 55 by cohort and sex.



Source: Own elaboration using data from MCVL

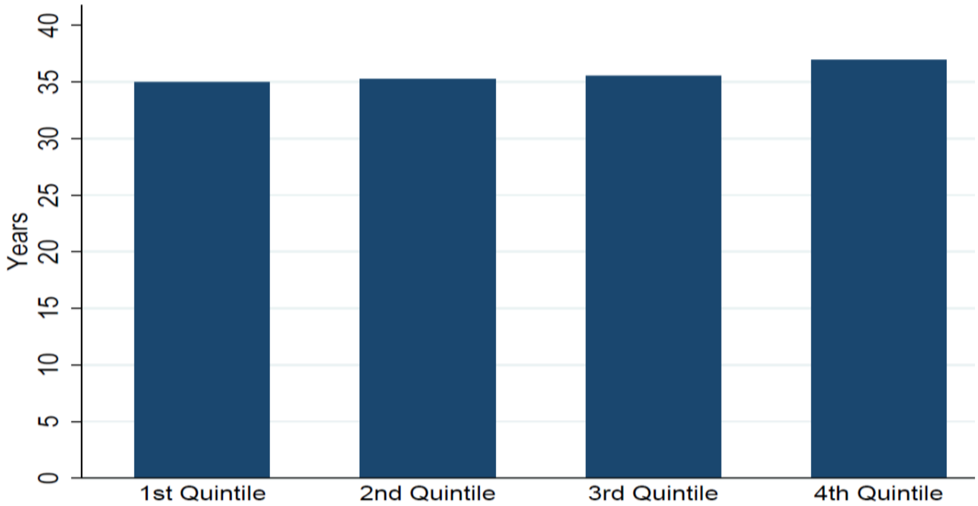
Figure 10 also shows that average life expectancy at age 55 did not vary much across cohorts. Therefore, we use the estimated values for the intermediate cohort (1941-1943) in our analysis of the effects of pension reforms on SSW across the income distribution (see Section 6). We distinguish by income quartile (based on the sum of the total contributions before age 50) and sex. Figures 11 and 12 plot these estimates for age 55.

Figure 11. Life expectancy at age 55 of men born 1941-1943 by income quartile (contribution).



Source: Own elaboration using data from MCVL, cohorts 1935-1949.

Figure 12. Life expectancy at age 55 of women born 1941-1943 by income quartile (contribution).



Source: Own elaboration using data from MCVL, cohorts 1935-1949.

5. Trends in older workers income and wealth inequality

In this section, we first present evidence on the evolution of income and wealth inequality for the population aged fifty to seventy in Spain using survey and administrative data. Then, we complement this evidence with trends in inequality in pension benefits from administrative data.

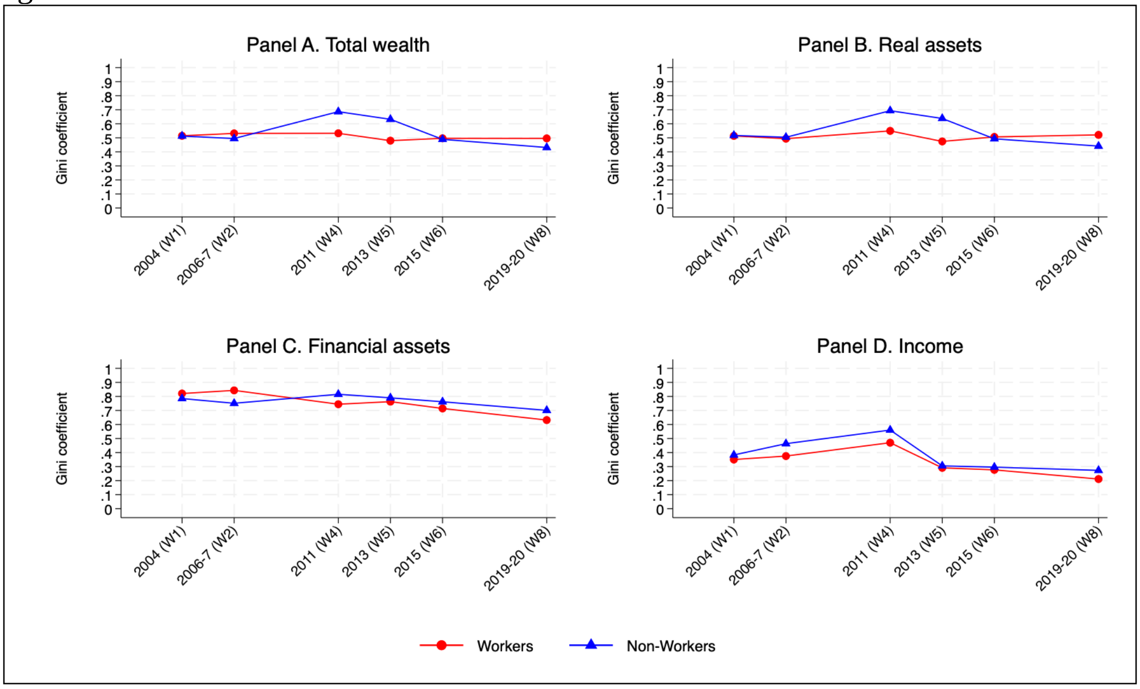
We use the Gini index to measure inequality. The figures are shown separately by workers and non-workers aged between fifty and seventy. The group of non-workers includes retired, and those who are unemployed or on disability. Some analyses are performed further by educational attainment.

5.1. Evolution of wealth and income inequality over time

We measure inequality trends using the Gini coefficient for total wealth, real assets, financial assets, and income. Figure 13 shows that the Gini coefficients for workers remained relatively constant for three of the four variables; total assets, real assets, and financial assets. On the other hand, for non-workers, the Gini increased substantially, from 0.5 to 0.7, in waves 4 and 5 (2011-2013), while in waves 6 and 8 (from 2015) the Gini went back to the level of wave 1 and wave 2 (pre-2008). Thus, during the worst moment of the 2008 financial and economic crisis, inequality in total wealth and real assets increased substantially among older non-working individuals in Spain.

In Panel D of Figure 13, we observe that inequality in income was reduced for both workers and non-workers from wave 5 (2013) and remained at this lower level during the rest of the observational period. Thus, the crisis reduced inequality in income for older individuals in Spain while it increased inequality for total wealth and real assets. The reduction in income inequality was permanent while the increase in wealth and real assets inequality was temporary for two waves, going back to pre-crisis levels in waves 6 and 8 (from 2015).

Figure 13. Gini coefficients of wealth and income by wave for workers and non-workers aged 50-70.



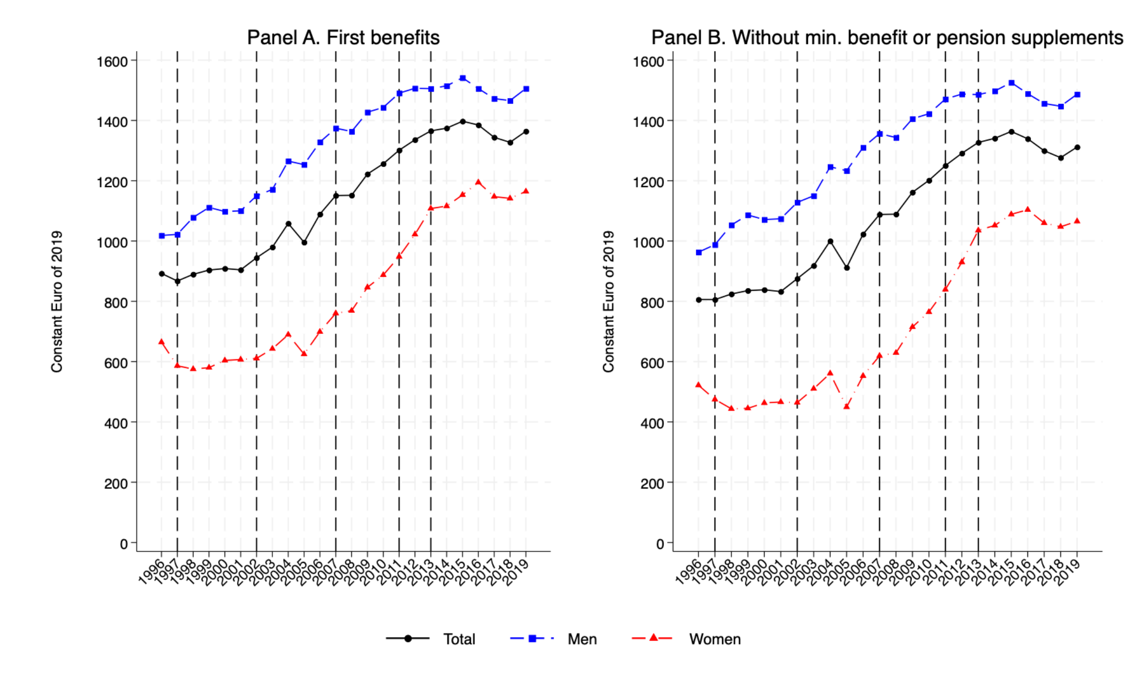
Source: Own elaboration using data from SHARE.

5.2. Trends in pension benefits inequality over time

In this section, we focus on first pension benefits. We plot trends in the average amounts, and a on the Gini coefficient. This will help us explore the extent of any reduction or increase in inequality of the first pension benefits and their potential link with the different old-age pension reforms. We calculate these numbers with and without adding minimum pensions and the complements individuals can receive on their pension benefits.

Figure 14 shows the evolution of first pension benefits using administrative data (MCVL). Initial pension benefits have been increasing over the last decades in Spain, for both men and women. The growth of in first pension benefits starts with the 2002 reform, being the level considerably higher for men and the slope mildly higher for women (Panel A). We see that this last trend remains when minimum pensions or supplements are removed (Panel B).

Figure 14. Evolution of first pension benefits over time, total and without complements

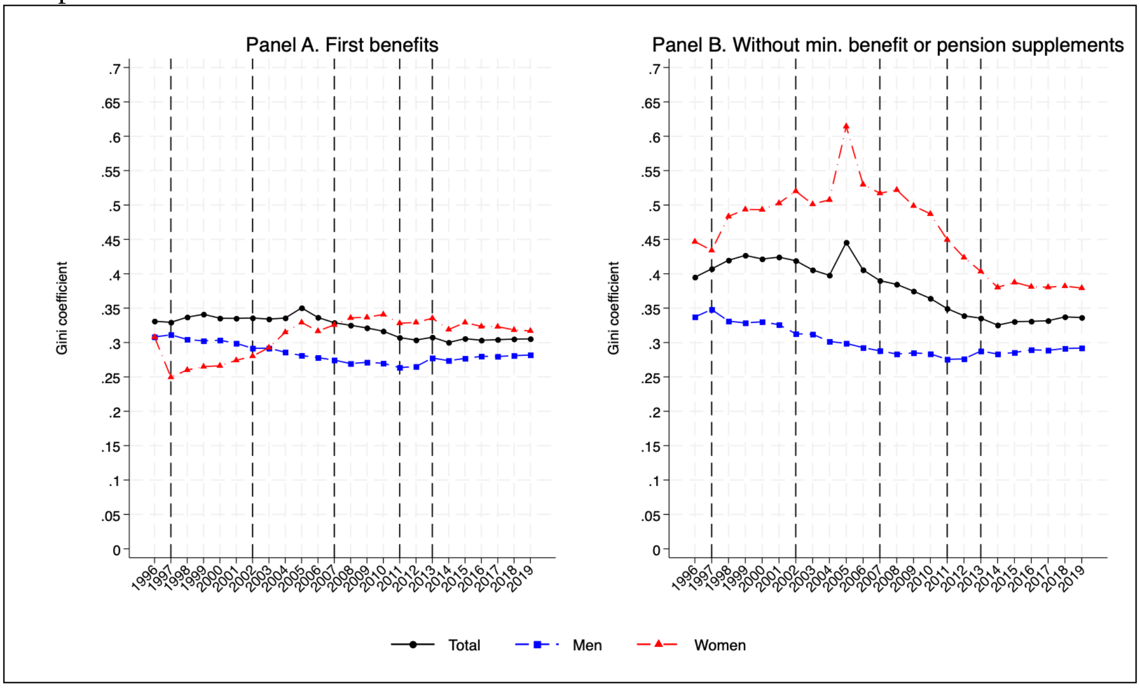


Source: Own elaboration using Continuous Sample of Working Histories (*Muestra Continua de Vidas Laborales*, MCVL).
Notes: Pension benefits are measured in constant Euro of 2019. The dashed vertical lines highlight the years of major pension reforms in Spain during the sample period, 1996-2019.

Figure 15 Panel A shows the evolution of inequality in the first pension benefits using administrative data (MCVL). Total inequality in the amount of the first benefits remained stable until 2005 when the Gini index peaked at 0.35. Inequality started then decreasing until 2013 when it stabilized at around 0.30. The trends in overall inequality mask two different trends and levels for female and male pensioners. We observe a large increase in inequality in first benefits among females between 1997 (Gini index =0.25) and 2010 (Gini index=0.34) followed by a mild decrease. In contrast, inequality in first benefits decreased among males between 1996 (Gini index=0.31) and 2011 (Gini index=0.26) and slightly increased thereafter. Inequality in pensions is larger among females than males at the end of our observation period.

Figure 15 Panel B illustrates the importance of minimum benefits and pension supplements in reducing inequalities in old-age pensions, specially among females. While the inequality trends in first benefits are similar when we exclude minimum benefits and pension supplements, the level of inequality is much higher. For example, Gini index for females would be over 0.45 until 2011 (except in 1997) and only slightly below 0.4 at the end of our observational period.

Figure 15. Gini coefficient of first pension benefits over time, total and without complements



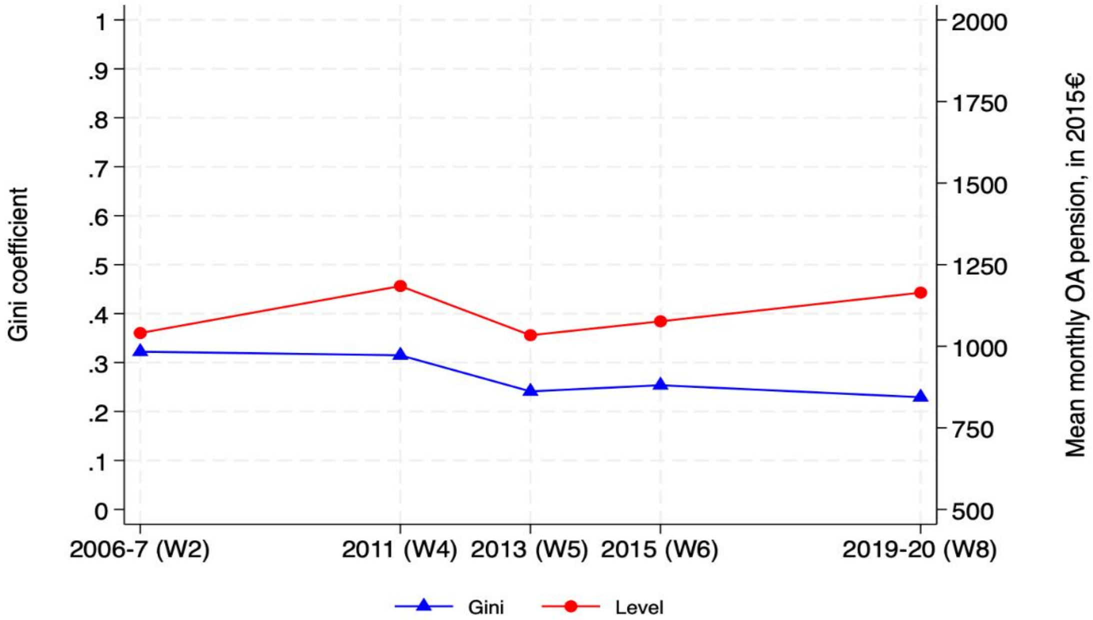
Source: Own elaboration using Continuous Sample of Working Histories (*Muestra Continua de Vidas Laborales*, MCVL).

Notes: Pension benefits are measured in constant Euro of 2019. The dashed vertical lines highlight the years Tof major old-age pension reforms in Spain during the sample period, 1996-2019.The 2005 peak in the women's series is due to an undetected error in women's first benefits data.

Next, we focus on the level and the Gini coefficient of net old-age pension income for each of the SHARE waves included in our sample for non-worker individuals aged between 50 and 70 years old. Figure 16 shows that net old-age pension income increases in wave 4 and wave 8, but always remains between 1000 and 1250 euros of 2015. At the same time, the Gini coefficient slightly decreases at wave 5, and this decrease is persistent in wave 6 and 8. Thus, the Gini coefficient of net old-age pension income decreased from slightly above 0.3 in wave 2 (2006-2007) to at around 0.2 at the end of the period, in wave 8 (2019-2020).

In our discussion of the old-age pension reforms in Section 2,, we anticipated early reforms of the Spanish system would benefit individuals with long careers, who are typically those with low skill levels. As shown in Figure 16, the Gini coefficient of net old-age pension income for non-workers aged 50 to 70 decreased after 2011. The 2011 reform progressively increased the number of contributory years required to calculate the benefit base. This change, thus, decreased income inequality among non-workers. However, during this period, women in Spain typically had fewer years of contributions. As we expected, Figure 14 illustrates an increase in inequalities for women, as they often did not meet the required number of years of contributions.

Figure 16. Level and Gini coefficient of Net OA pension income by wave for non-workers aged 50-70.



Source: Own elaboration using SHARE.

Notes: Individual sampling weights are used. Pension income is measured in constant Euro of 2015. The top and bottom 5% of OA pensions have been trimmed. "UB" uses unfolding brackets to retrieve missing values (imputed values for the single component of OA pensions are not available). W1 is excluded because OA pension income is available in gross terms only.

Although we still have no data available, we also expect that the 2021/2023 pension reform will have a positive impact (a reduction) on inequality in Spain due to the increase in the pension amount of some of the most disadvantaged individuals, such as those receiving the minimum pension.

6. Effects of pension reforms on Social Security Wealth across the income distribution

In this section, we analyze the effect of pension reforms on inequality (SSW) across the income distribution. We first evaluate the impact of the different pension reforms on the probability of retirement, and then on the average SSW by income quartile.

6.1. Retirement probability

We use our panel of 55–70-year-olds (see section 3) to estimate the retirement hazards under an individual's *actual* and *counterfactual* pension system. These systems are defined, correspondingly, as the one in which the individual has or will retire, and the system of 1985. All individuals in our analysis are initially working at age 55 and they exit the panel as they retire.

To obtain the retirement hazards, we estimate a discrete-time proportional hazard model using the complementary log-log link. The model is estimated by maximum likelihood. We include a fully non-parametric baseline hazard that is allowed to differ by sex. The key explanatory variables are the SSW and Implicit Tax Rate (ITAX) under either individuals' *actual* or *counterfactual* pension system. The effects of SSW and ITAX are allowed to vary by pension regime. We further include dummies for quartiles of lifetime earnings, highest skill level, and region. We then use the retirement hazards to compute the unconditional retirement probabilities in our counterfactual analyses in section 6.2.

Figure 17 plots the probability of retiring at specific ages by sex for both the actual $(p_{R,k,A})$ and counterfactual $(p_{R,k,CF})$ regime. We adjust both the SSW and the ITAX to either the *actual* or the *counterfactual* system to compute these predicted probabilities, while all the other variables remain as observed. When comparing individuals under their actual system and that of 1985, for both men and women, we observe a similar change in their distribution of retirement ages, namely a decrease in the probability of retiring between 60-62 and an increase in retiring at 65-66. Figure 21 shows this pattern is common across income groups, although it is most pronounced for men and women in the highest quartile.

Figure 17. Retirement probability at specific ages under Actual and Counterfactual systems, by sex.

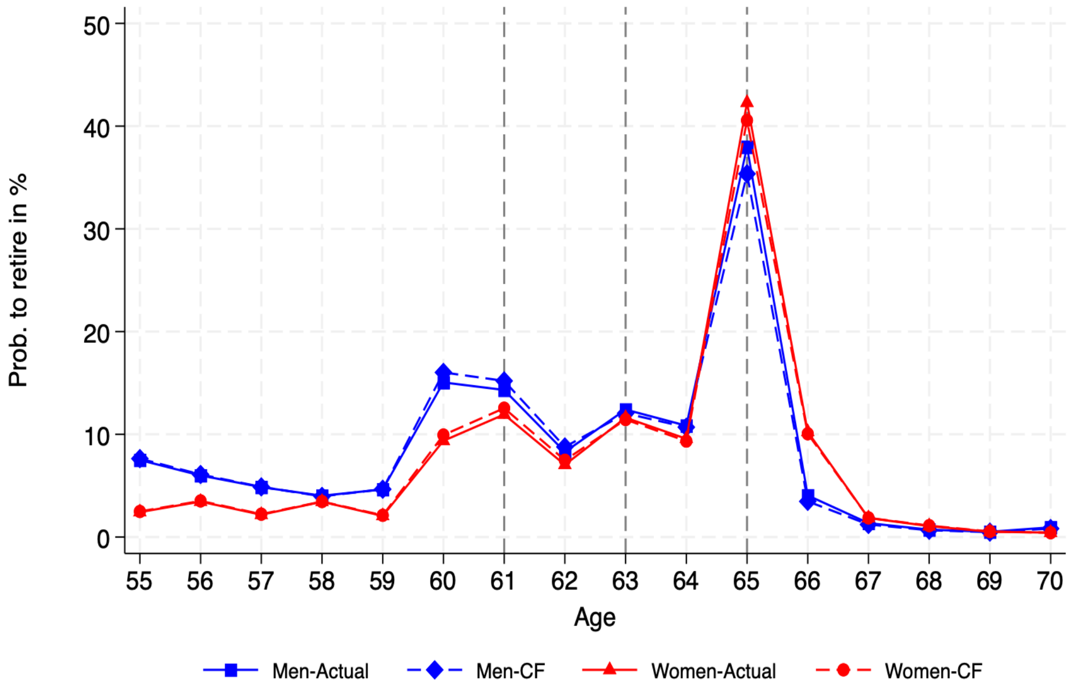
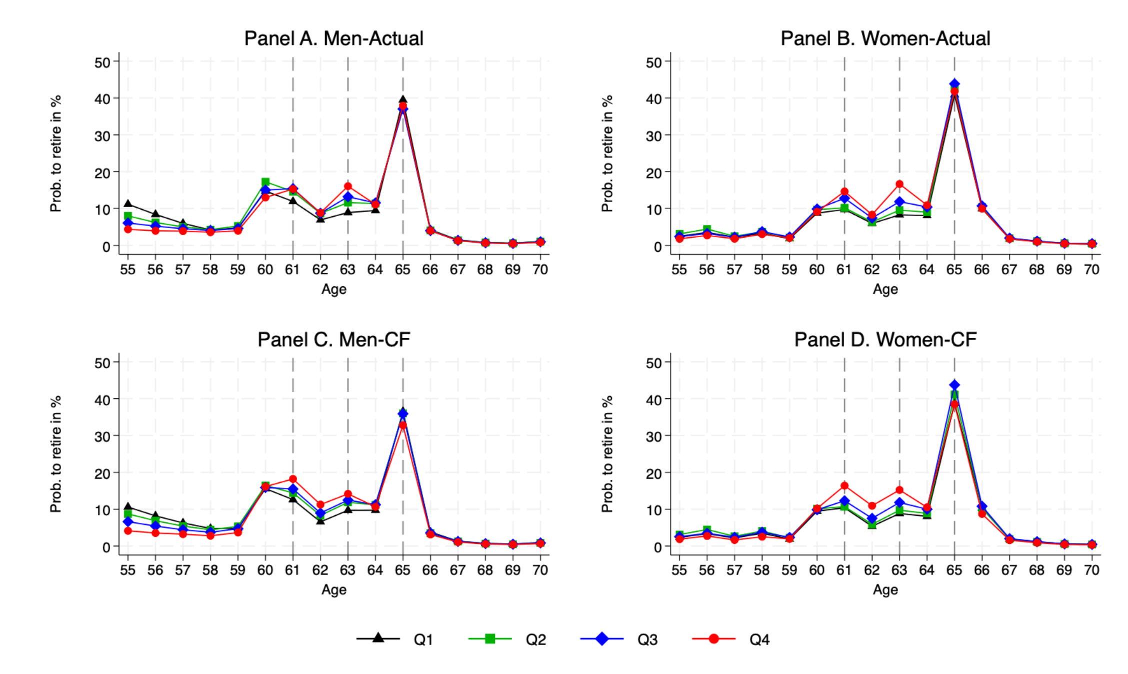


Figure 17. Retirement probability at specific ages under Actual and Counterfactual systems, by sex and income.



6.2. Counterfactual analysis

To estimate the effect of the pension reforms on the distribution of SSW, we perform a counterfactual analysis. In addition to the actual SSW measures, i.e. those that workers experienced at the time they retired, we compute counterfactual measures using the rules from the 1985 regime. Similarly, the probability of retiring would differ under the different regimes, so we also compute an actual and counterfactual probability using the estimates from 6.1. We see that most of our individuals (around 80%) retire under the 2013 regime as our panel pools the years 2007-2022. We allow for differences in life expectancy between individuals with different levels of lifetime earnings, as discussed in section 4.

In the analysis, we decompose the effects into a direct or mechanical effect and a secondary or behavioral effect. For the direct effect, we hold the retirement age distribution fixed, so it captures the mechanical effect on the distribution of benefits through changes in generosity but without behavioral changes. On the other hand, retirement behavior may change when rules change. The secondary or behavioral effect operates therefore through changes in retirement patterns. We simulate the retirement patterns and SSW of different individuals under the pension system in which they retired (post-1985) and that of 1985 as follows:

(5) $SSW(i) with\ reform = \sum_{R=55}^{70} SSW_{R,i,A} * p_{R,k,A}$

(6) $SSW(i) without\ reform\_mechanical\ effect = \sum_{R=55}^{70} SSW_{R,i,CF} * p_{R,k,A}$

(7) $SSW(i) without\ reform\_total\ effect = \sum_{R=55}^{70} SSW_{R,i,CF} * p_{R,k,CF}$

(8) $$SSW(i)\_direct\ effect = (5) - (6)$$

(9) $$SSW(i)\_secondary\ effect = (6) - (7)$$

(10) $$SSW(i)\_total\ effect = SSW(i)\_direct\ effect\ + SSW(i)\_secondary\ effect$$

Where R is the retirement age (which can be up to age 70, though most individuals retire earlier), p is the retirement probability at age R, A refers to the *actual* retirement regime while CF to the *counterfactual* (1985 regime), k is the retirement route.

In the appendix, we show estimates from two additional analyses. First, we perform a similar comparison without allowing for differences in life expectancy by income. Second, we restrict the analysis to compare the systems of 2013 and 1985, while allowing for differences in life expectancy by income. We use the same sample and estimate SSW and retirement probabilities under both systems for all the individuals.

Figure 18 plots SSW at different ages under different pension systems (actual and 1985 system), by age. We plot SSW as defined in equation (1), i.e., without adjusting for retirement probabilities and ignoring behavioral variability in the retirement age. We see that for both men and women there is an increase in SSW across pension systems, which widens after age 63. As reported in Table B1,the SSW is €14675 lower for women in the counterfactual system compared to their actual system, and the magnitude is even higher (€14675) for men. In addition, there is a gender gap in SSW under the counterfactual regime that disappears under the actual regime. Figure xx plots these trends across income quartiles. While we see similar trends across all quartiles, the increase in SSW across systems is largest, in absolute terms, for individuals in the highest income quartile.

Figure 18. SSW at specific ages under Actual and Counterfactual system, by sex.

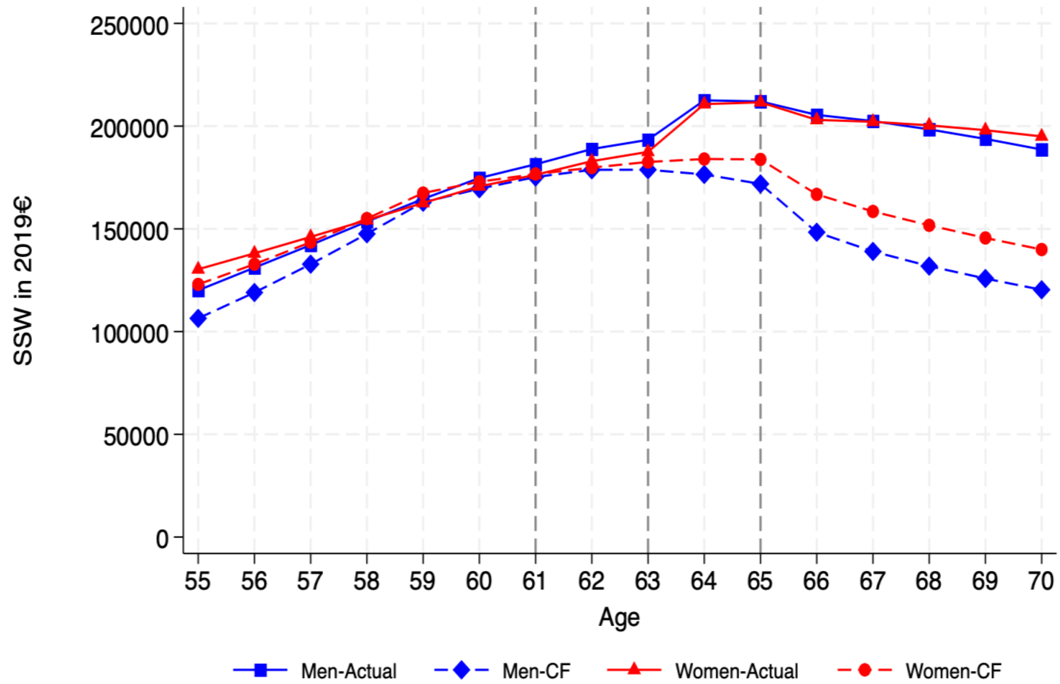
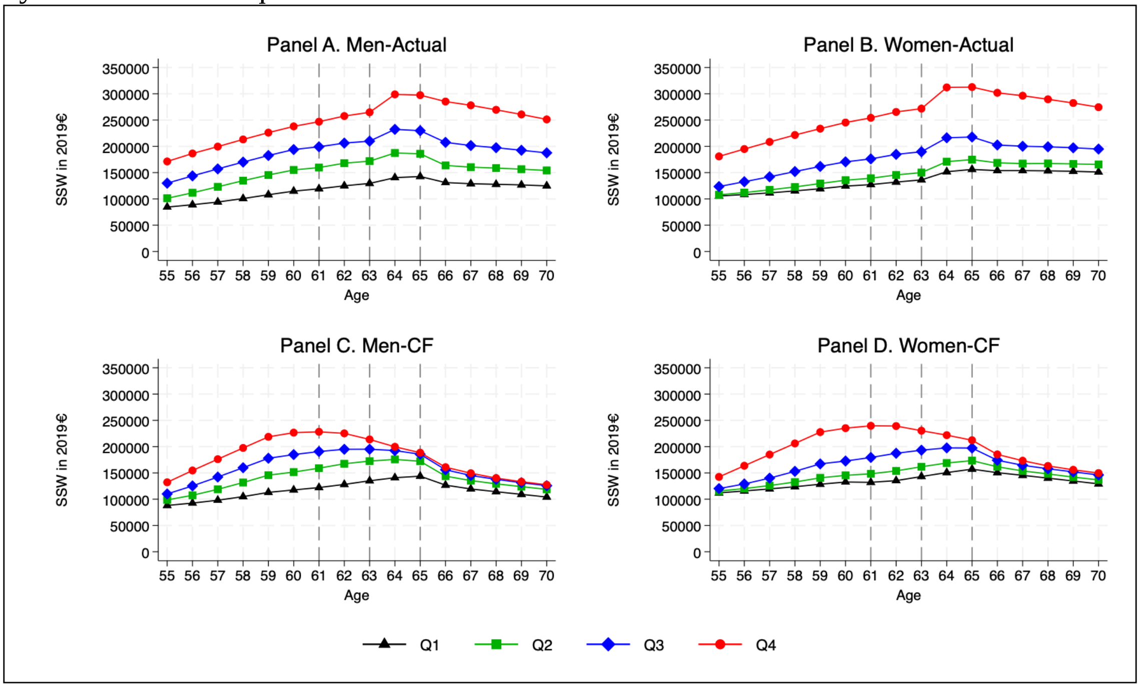


Figure 19. Unweighted SSW at specific ages under Actual and Counterfactual system, by sex and income quartile.



Previous figures show that changes in retirement probability and SSW across systems go in a similar direction. Figures 20-21 plot the changes in the retirement probability and the SSW between the actual and counterfactual systems. We find these changes reinforce each other, suggesting that older workers respond to financial incentives. This holds at least until age sixty-five. From age sixty-five, the new (actual) system becomes even more generous in terms of SSW compared to the 1985 system, but individuals modify their retirement behavior very little. These differences across systems are largest for individuals in the highest income quartile (see Figure 20).

Figure 20. Difference between Actual and Counterfactual systems in retirement probability and unweighted SSW at different ages, by sex.

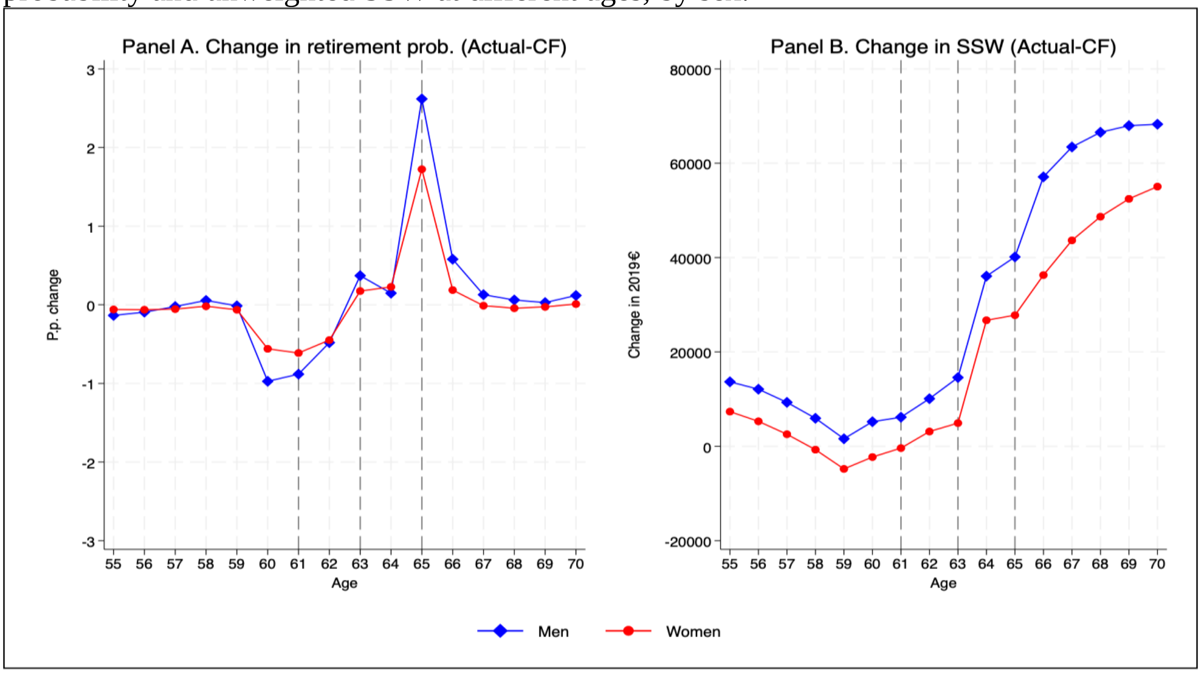
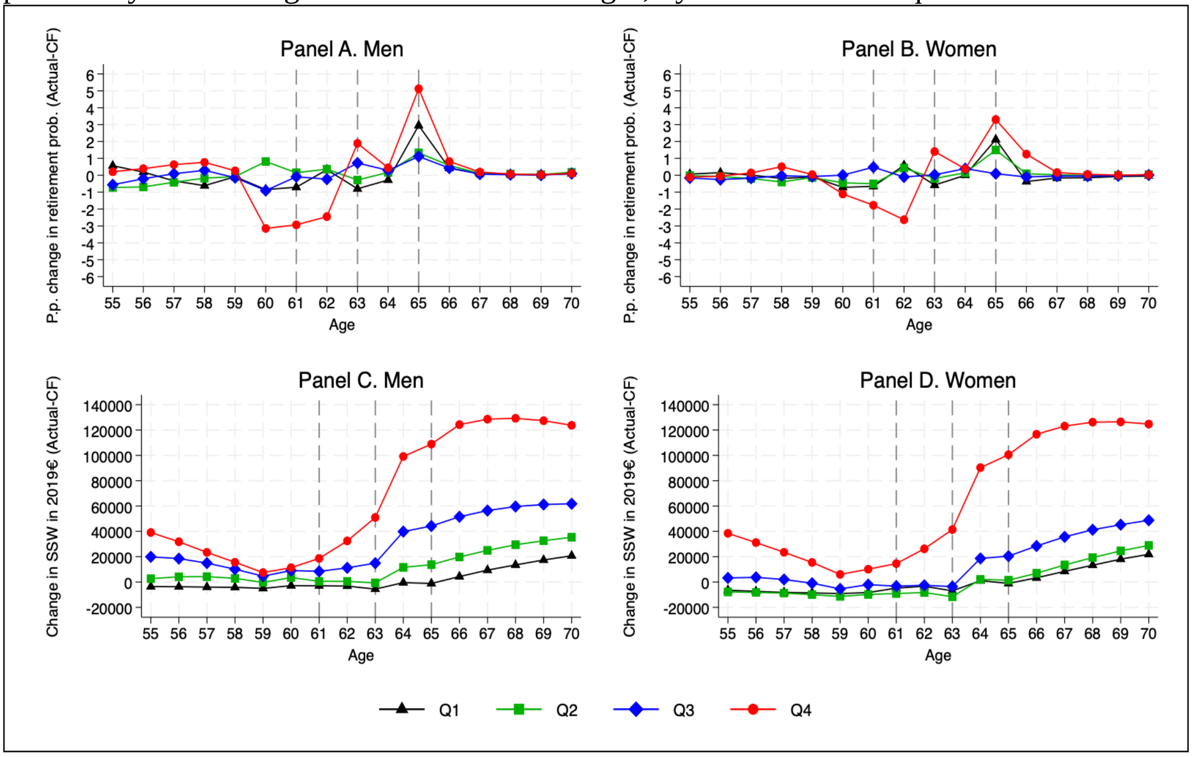


Figure 21. Difference between Actual and Counterfactual systems in retirement probability and unweighted SSW at different ages, by sex and income quartile.



Figures 22 and 23 plot the decomposition of overall effects into mechanical and behavioral effects (22 shows values in 2019 Euros and 23 in percentage with respect to the counterfactual levels), by sex and income quartile. Table B1 in the appendix provides all the numerical values and includes the estimated SSW with and without retirement probability weights.

Figure 22. Decomposition of overall effects into mechanical and behavioral effects, in 2019 Euros.

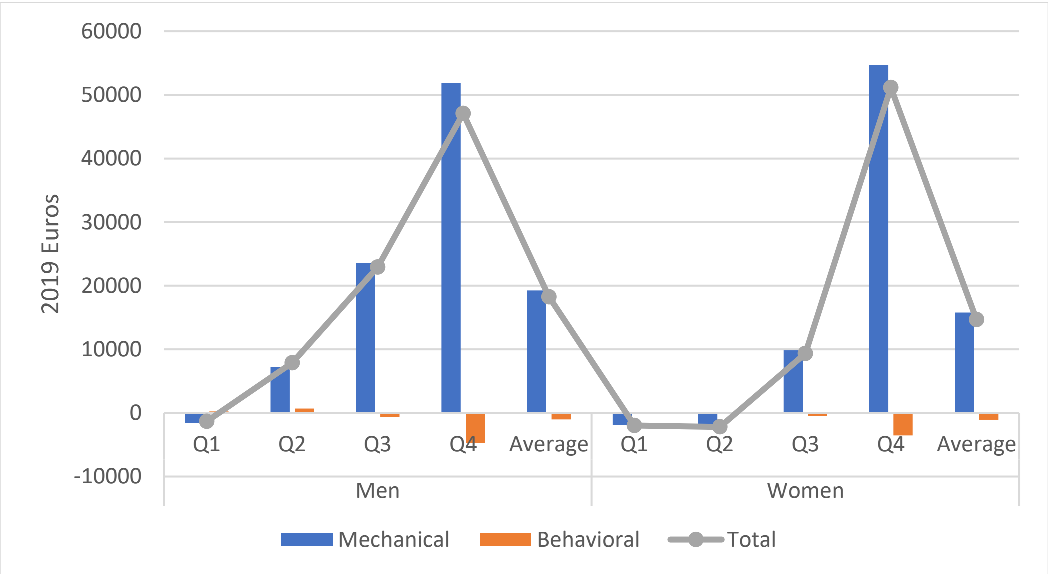
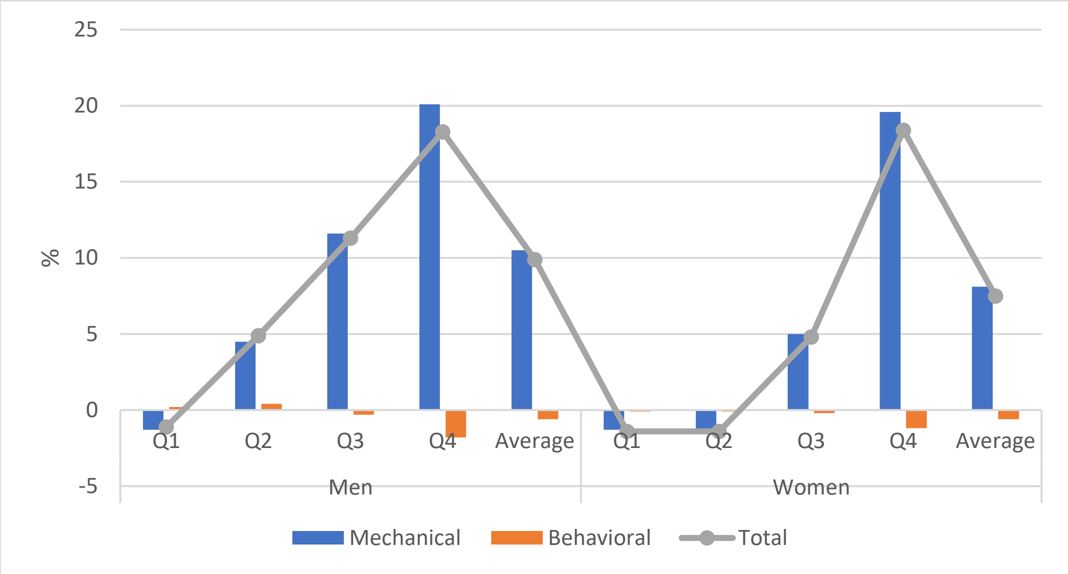


Figure 23. Decomposition of overall effects into mechanical and behavioral effects, in %.



Overall, compared to the system of 1985, SSW has increased on average by about 18,000€ for men and 15,000€ for women. This represents a ten and eight percent increase, respectively. This effect is mostly driven by the mechanical or direct effect (e.g. via benefit adjustments) as the part driven by changes in the retirement probability (secondary or behavioral effect) is close to zero.

We find striking differences across income quartiles, for both men and women. In both cases, there is a clear income gradient, where the richest quartile has benefitted the most with an increase close to twenty percent or over €50,000 for both men and women. In contrast, the total effect is close to zero, and even slightly negative, for the poorest income quartile for men and the two poorest income quartiles for women. Last, while there are small differences in the size of the behavioral effect, it is always close to zero.

7. Conclusion

In this chapter, we analyze the evolution of income inequality among the Spanish elderly and estimate the differential effect of old-age pension reforms across income quartiles using administrative data. We estimate the SSW of Spanish workers under the system in which they retired, and in a counterfactual scenario in which they would have faced the 1985 regime. We, then, decompose the total change in SSW between a direct or mechanical effect driven by changes in the rules to compute the benefits, and a secondary or behavioral effect due to induced changes in retirement behavior.

In this analysis, we allow for differences in life expectancy across income groups in our measures of financial incentives (SSW and ITAX). For this, we estimate the conditional probability of mortality using a Gompertz distribution within a parametric proportional hazard model framework. We find that younger cohorts experience higher survival probabilities, particularly after age 70, with more pronounced differences among women. Moreover, high-skilled occupations exhibit higher survival rates across all cohorts, with a widening gap among men but diminishing among women. Finally, income-related survival inequality is increasing among men, while for women, this gradient is less evident but emerging in younger cohorts at older ages.

Regarding the impact of recent pension reforms on social security wealth inequality, we find that compared to the 1985 rules, SSW has increased on average by about 18,000€ for men and 15,000€ for women. This represents a ten and eight percent increase, respectively. This effect is mostly driven by the mechanical or direct effect (e.g. via benefit adjustments) as the part driven by changes in the retirement probability (secondary or behavioral effect) is close to zero. Furthermore, we find striking differences across income quartiles, for both men and women. In both cases, there is a clear income gradient, where the richest quartile has benefitted the most with an increase close to twenty percent or over €50,000 for both men and women. In contrast, we find the total effect is close to zero, and even slightly negative, for the poorest income quartile for men and the two poorest income quartiles for women. This is likely due to the effect of minimum benefits (that mark the generosity of the system, see Boldrin et al, 1999) that automatically absorb any other effect for low-income individuals.

In this chapter, we have focused on only one dimension of inequality, SSW inequality at older ages, while old-age pension reforms may have consequences on other dimensions of individual well-being. For example, reforms of the Spanish system that increased the retirement age, by limiting early retirement and encouraging later retirement, may have adversely affect the health of low-skilled workers engaged in physically demanding jobs. These workers could face increased health risks and potentially even higher mortality rates. Therefore, future work should assess the differential impact of these reforms on the health of workers across socioeconomic groups.

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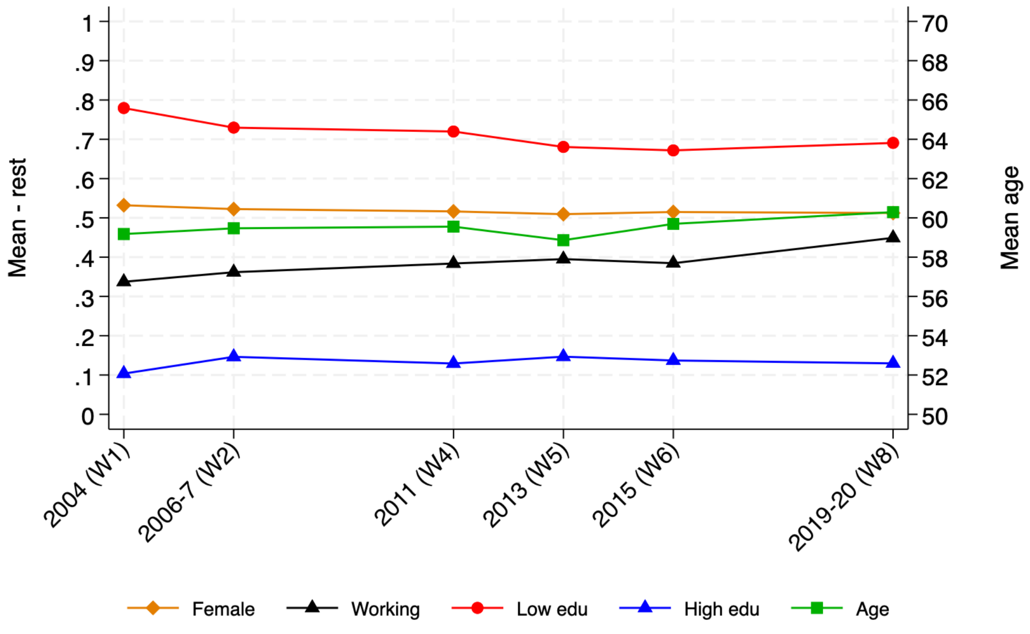
Appendix

This appendix provides additional analyses and results to the main paper.

A. Income and wealth in the Survey of Health, Aging and Retirement in Europe

Figure A1 plots descriptive statistics of the main socioeconomic characteristics of our sample of analysis using SHARE over time. Most variables are relatively constant over time, although we see a 10 percentage point decrease in the share of the population with low education and a small increase in the average age of the sample, in line with the observed aging of the Spanish population. Last, despite the aging of the population, we see a small increase in the share that is working from 0.39 in wave 1 (2004) to 0.45 in wave 8 (2019/2020).

Figure A1. Sample composition by wave, 50–70-year-olds

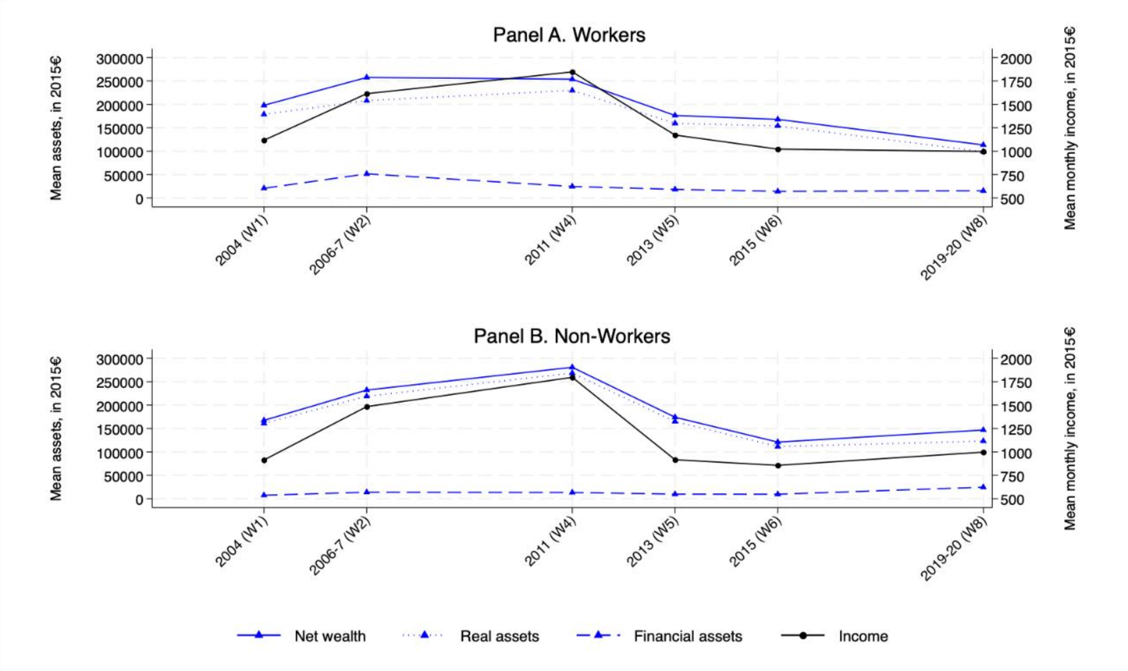


Source: Own elaboration using data from SHARE. Notes: Individual sampling weights are used.

Our two main variables of interest are income and wealth. We measure income at the household level using unfolded brackets and imputations. Similarly, wealth variables are measured at couple level and based on imputed data. The overall (net) wealth variable includes both real and financial assets. Real assets encompass the value of real estate (net of any debts, such as mortgages), cars, and businesses owned by the respondents and their spouses. Financial assets cover amounts in bank accounts, bonds, stocks, mutual funds, retirement accounts, and contractual savings. We set all negative income and wealth values to zero and trim the top and bottom 5%. Last, to ensure comparability across households and over time, we adjust income and wealth for economies of scale using the OECD Equivalence scale and constant 2015 Euros. Individual sampling weights are used.

Figure A2 plots household income and wealth for the population aged fifty to seventy by employment status over time. We use four measures: net wealth, real assets, financial assets, and income. Financial assets are close to zero for this group of older individuals in Spain for the entire period. For both working and non-working individuals at older ages, there is a steady increase in all the other three measures of income and wealth between 2004 and 2011 (wave 1, wave 2, and wave 4), followed by a substantial decline infrom 2013 (wave 5). This decline coincides with the worst moment of the financial and economic crisis in Spain; the unemployment rate in Spain reached 27% in 2013, the highest level since the onset of the 2008 crisis. Furthermore, although the unemployment rate started slowly decreasing right after the pick in 2013, the wealth and income levels for old age individuals in Spain stagnated and did not recover at the same pace.

Figure A2. Wealth and income by wave for workers and non-workers aged 50-70



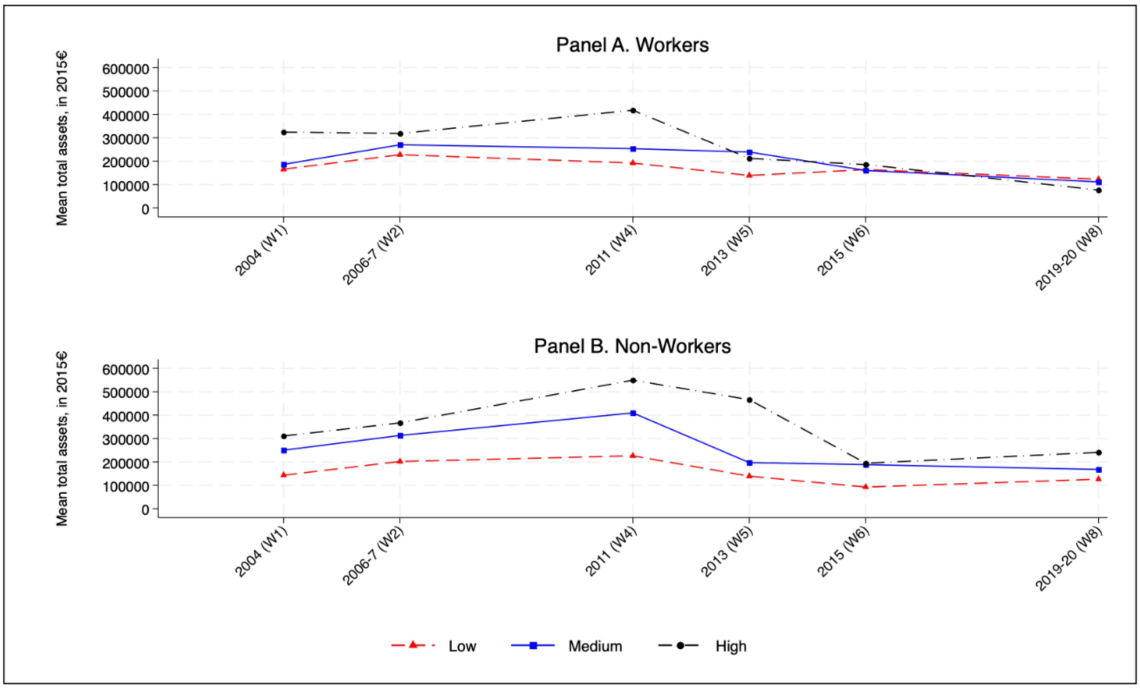
Source: Own elaboration using data from SHARE.

Notes: Individual sampling weights are used. All measures are measured in constant Euro of 2015. Wealth variables are measured at the couple level and based on imputed data. Negative values are set to zero. Income is measured at the household level and uses unfolding brackets and imputations. We trim the top and bottom 5% of income and use the OECD Equivalence Scale to enhance the comparability of wealth and income across households of different sizes. Non-workers include individuals who are retired, unemployed, or on disability.

Figure A3 plots our overall wealth variable, net wealth, and Figure A4 plots monthly income for both workers and non-workers aged 50-70 by level of education. Interestingly, for working individuals, the drop in net wealth in wave 5 observed in Figure A2 above is driven by high educated individuals, as individuals in the other two educational groups experienced a milder and more progressive drop. Conversely, for non-working individuals, we can see that the drop in net wealth after the financial crisis is driven by those with either high or middle education.

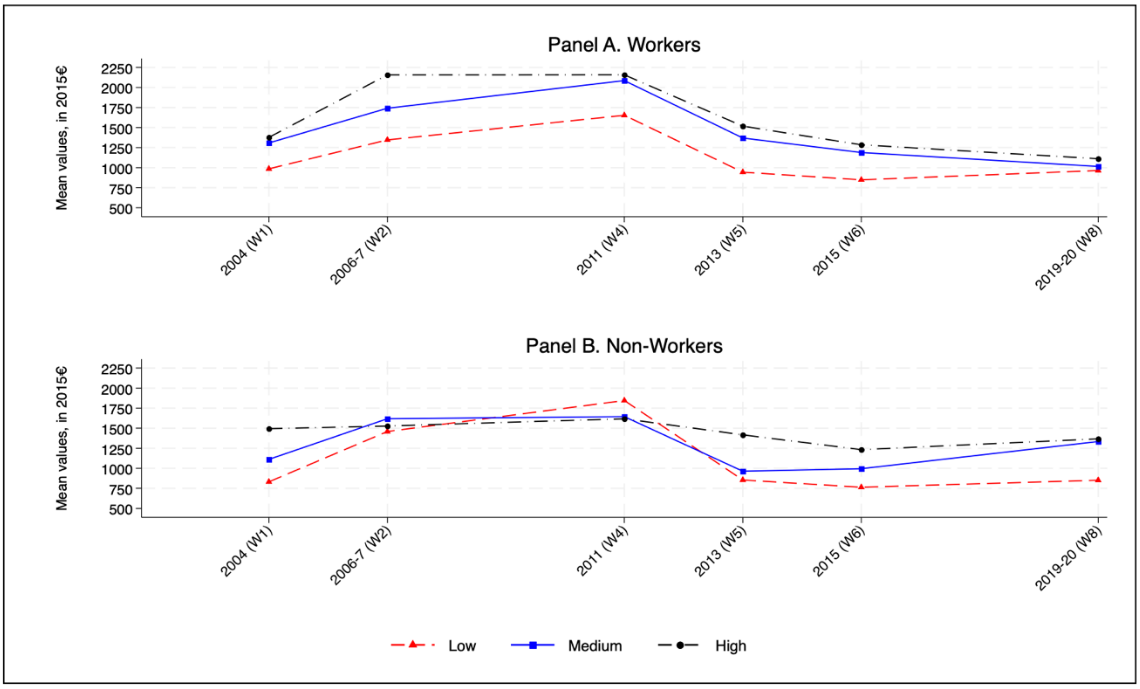
The picture is substantially different when we focus on income instead of wealth. Figure A4 shows a very similar drop in monthly income for the three educational groups for both working and non-working individuals.

Figure A3. Net wealth by wave and education level for workers and non-workers aged 50-70.



Source: Own elaboration using data from SHARE.

Figure A4. Net monthly income by wave and education level for workers and non-workers aged 50-70.



Source: Own elaboration using data from SHARE.

B. Decomposition of the impact of the reform. Detailed tables.

Table B1. Decomposition into direct and secondary effects - Differentiated mortality rates.

		SSW (in 2019€)					Direct effect		Secondary effect		Total effect	
	LTE	with reforms	without reforms – mechanical effect only	without reforms – total effect	with reforms (unweighted)	without reforms (unweighted)	€	%	€	%	€	%
Men	Q1	120364	121925	121667	1192052	1205996	-1560	-1.3	258	0.2	-1303	-1.1
	Q2	161394	154180	153490	1527008	1461480	7214	4.5	689	0.4	7904	4.9
	Q3	203952	180368	181004	2058394	1803772	23584	11.6	-636	-0.3	22947	11.3
	Q4	257868	206029	210800	2861334	2224128	51839	20.1	-4770	-1.8	47068	18.3
	Average	183481	164236	165240	1878368	1655312	19245	10,5	-1004	-0.6	18241	9.9
Women	Q1	142187	144110	144184	1891393	1906963	-1924	-1.3	-73	-0.1	-1997	-1.4
	Q2	156558	158604	158761	1840717	1869260	-2047	-1.3	-156	-0.1	-2203	-1.4
	Q3	196069	186234	186706	2168772	2051226	9835	5.0	-472	-0.2	9363	4.8
	Q4	278446	223778	227298	3192440	2532337	54668	19.6	-3520	-1.2	51149	18.4
	Average	195210	179445	180535	2287780	2096983	15765	8.1	-1090	-0.6	14675	7.5

C. Additional counterfactual analyses.

Table C2. Decomposition of overall effects into direct and secondary effects (comparing current vs system of 1985, without differentiated mortality rates).

		SSW (in 2019€)					Direct effect		Secondary effect		Total effect	
	LTE	with reforms	without reforms – mechanical effect only	without reforms – total effect	with reforms (unweighted)	without reforms (unweighted)	€	%	€	%	€	%
Men	Q1	126590	128116	127859	1253864	1266672	-1526	-1,2	257	0,2	-1269	-1,0
	Q2	159650	152442	151937	1504628	1440875	7208	4,5	505	0,3	7713	4,8
	Q3	196410	173882	174668	1973217	1734362	22528	11,5	-786	-0,4	21742	11,1
	Q4	242023	193931	198647	2671001	2090020	48091	19,9	-4716	-2,0	43375	17,9
	Average	179122	160992	162069	1823625	1617557	18130	10,1	-1077	-0,6	17053	9,5
Women	Q1	146198	148109	148145	1945662	1959444	-1911	-1,3	-35	0,0	-1947	-1,3
	Q2	156968	158960	159123	1845924	1874137	-1992	-1,3	-163	-0,1	-2154	-1,4
	Q3	193303	183649	184152	2137520	2023417	9654	5,0	-503	-0,3	9151	4,7
	Q4	270819	218144	221683	3103656	2468756	52675	19,4	-3539	-1,3	49136	18,1
	Average	193520	178301	179397	2270166	2086302	15219	7,9	-1096	-0,6	14123	7,3

Table C3. Decomposition of overall effects into direct and secondary effects (comparing system of 2013 vs that of 1985, with differentiated mortality rates).

		SSW (in 2019€)					Direct effect		Secondary effect		Total effect	
	LTE	with reforms	without reforms – mechanical effect only	without reforms – total effect	with reforms (unweighted)	without reforms (unweighted)	€	%	€	%	€	%
Men	Q1	118469	122127	97865	1077526	1184625	-3658	-3,1	24262	20,5	20604	17,4
	Q2	146214	144295	120748	1226906	1285931	1919	1,3	23547	16,1	25466	17,4
	Q3	188993	167497	149683	1714582	1582301	21496	11,4	17814	9,4	39310	20,8
	Q4	255373	200671	191537	2494782	1968617	54703	21,4	9133	3,6	63836	25,0
	Average	190210	166038	148806	1763691	1580790	24173	12,7	17232	9,1	41405	21,8
Women	Q1	147972	149124	112892	1304385	1425689	-1152	-0,8	36232	24,5	35080	23,7
	Q2	154490	156707	140136	1547420	1667415	-2217	-1,4	16571	10,7	14354	9,3
	Q3	186582	177500	163618	1790071	1787005	9082	4,9	13882	7,4	22964	12,3
	Q4	276235	218078	212226	2795152	2254595	58158	21,1	5851	2,1	64009	23,2
	Average	197638	178807	162116	1930902	1823400	18831	9,5	16692	8,5	35523	18,0