The Effect of Savings Annuitization on Elderly Outcomes

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Abstract

I estimate the causal effects of annuitization on old-age outcomes. Annuities insure retirees against outliving their savings by providing income for life, yet whether annuitizing pension wealth improves retirees' well-being remains an open empirical question. I exploit a rule in Colombia's private pension system under which retirees whose savings exceed an individual-specific threshold receive a life annuity and those below it receive a lump sum. Using administrative records covering the universe of retirees from 2008 to 2023, linked to death records and the social-assistance registry, I implement a regression discontinuity design at this threshold. Annuitization increases life expectancy within the 17-year observation window by 1.3 years and reduces social-assistance registration by 23 percentage points. Mortality effects emerge two to three years after retirement and are concentrated among lower-wage retirees, consistent with annuitization easing financial constraints later in retirement. These results provide some of the first causal evidence on the effects of annuitization and show that pension payout design has lasting consequences for longevity and reliance on the social safety net.

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1 Introduction

Population aging places growing pressure on retirement systems to provide income security in old age while remaining fiscally sustainable. The number of individuals aged 65 and older is projected to more than double, from 761 million in 2021 to 1.6 billion in 2050 (Wilmoth et al., 2023). Pensions already account for a substantial share of public spending, yet coverage remains incomplete and old-age poverty persists (OECD, 2025; ILO, 2024).[2] At the individual level, longevity risk remains a central concern. Retirees must finance consumption from a finite stock of wealth without knowing how long they will live.

Annuities are the canonical instrument for insuring against longevity risk. They convert a stock of savings into a stream of income guaranteed for life. Economic theory, starting with Yaari (1965), has consistently shown that retirees should annuitize all or most of their wealth, a result that holds under general conditions (Davidoff et al., 2005).[3] Yet despite this theoretical case, there is little empirical evidence on whether annuitization actually improves retirees' well-being.

Understanding these consequences has become increasingly important as pension systems shift from defined-benefit (DB) arrangements, such as traditional employer pensions and social security, to defined-contribution (DC) arrangements, such as 401(k)s and funded individual-account systems.[4] The shift has been substantial. DB plans now account for less than one-third of pension assets in OECD countries, down from nearly 40 percent a decade ago, while in the United States the coverage of DC plans rose from 55 percent to 83 percent over the past two decades (Bennett et al., 2025). Because retirees rarely convert accumulated savings into private annuities, perhaps due to failures in annuity markets or behavioral biases (Finkelstein and Poterba, 2004; Brown et al., 2008; Hershfield et al., 2026), a growing share of retirees bear longevity risk themselves (Poterba, 2005). Understanding the causal effects of annuitization is therefore critical for determining the consequences of this shift and for policy design.

In this paper I study whether retirees who receive a lifelong annuity, rather than a finite

[2] Across OECD countries, the average relative income poverty rate among people over age 65 is 15 percent, reaching 23 percent in the United States and 40 percent in Korea (OECD, 2025).

[3] Annuities pool longevity risk. The savings of those who die early finance the payments of those who live long, so the annuity pays more than a bond conditional on survival, a premium known as the mortality credit. Yaari (1965) shows that without a bequest motive and with fair pricing, full annuitization is optimal. Davidoff et al. (2005) relax fair pricing, separable utility, and exponential discounting, and find that most wealth should still be annuitized.

[4] A DB plan promises a lifelong pension, typically tied to earnings history, and leaves longevity risk with the employer or the government. A DC plan accumulates contributions in an individual account and delivers the balance as a stock at retirement, and the retiree decides how to convert it into income.

payout of the same accumulated wealth, live longer and fare better in old age. Answering this question is challenging: the decision to annuitize is generally endogenous. Annuitants differ systematically from other retirees—in wealth, education, health, expected longevity, financial sophistication, and preferences over risk and bequests (Finkelstein and Poterba, 2004; Inkmann et al., 2011)—and these same characteristics independently predict old-age outcomes. Determining the causal effects of annuitization therefore requires a unique combination of (i) a setting that provides quasi-random variation in annuity access versus other forms of payout, and (ii) detailed administrative data that follow individuals across time and are linked to old-age outcomes.

This paper overcomes both obstacles by studying Colombia's private pension system. For a group of retirees, a legally mandated savings threshold determines the form of the pension payout. Retirees with savings above the threshold must annuitize. Those below it receive their savings as a lump sum. The threshold is individual-specific and depends on policy parameters—most importantly the statutory minimum wage—that generate substantial variation in its level over time and are difficult for workers to anticipate. I exploit this rule in a regression discontinuity design, comparing retirees with savings just above and just below their threshold. Because I measure the distance of individuals' savings to the threshold with error, the main empirical strategy is a fuzzy regression discontinuity design, where I use being above the threshold as an instrumental variable that predicts the probability of annuitizing. I estimate the effects of annuitization on mortality and poverty using administrative records on the universe of private-pension affiliates, linked to individual death records and to the registry of the social-assistance system. To my knowledge, this is the first paper to provide quasi-experimental evidence on the causal impact of annuitization against a lump-sum counterfactual.

The analysis draws on confidential administrative records from Colombia's private pension system. Every year, each of the four pension funds reports the full stock of its affiliates. Together these records span 2008 to 2023, more than 8 million workers, and roughly 600,000 retirees. For each individual, I observe annual earnings, contributions, and savings, together with the date of retirement, the payout form, and demographic characteristics. I link these records to the national civil registry, which reports vital status as of June 2025 and the date of death for deceased retirees, and follow survival for up to 17 years after retirement. I also link retirees to the social-assistance registry and observe whether each retiree is registered at the same date.

A natural identification concern is sorting around the threshold. Institutional features make manipulation difficult. Contributions are a fixed share of taxable earnings, leaving

workers little discretion over accumulated savings, while investment decisions are managed by pension funds. The threshold is never disclosed to workers, and its year-to-year changes are large relative to annual contributions and difficult to predict. This makes it difficult for workers to sort around the cutoff by adjusting when they retire. I find no evidence that retirees use this margin to sort. Retirement remains concentrated at the statutory ages regardless of position relative to the threshold, the density of the running variable is smooth at the cutoff, and predetermined characteristics are balanced across it.

I find that annuitization lowers mortality. Among retirees near the threshold, annuitization lowers the annual death hazard by 0.96 percentage points. The effect builds over time. It is absent in the first year after retirement, emerges in the second and third, and peaks in the fourth and fifth. These hazard estimates translate into additional years of life. Within the 17 years of follow-up, annuitization adds 1.3 years of life. Extending the estimates beyond the observed window, annuitization raises remaining life expectancy at the statutory retirement ages of 57 for women and 62 for men by 2.5 and 2.3 years.

Annuitization also lowers poverty in old age. I measure poverty with registration in the means-tested social-assistance registry. Among retirees near the threshold, annuitization lowers registration by 23 percentage points. The effect grows with time since retirement, from zero for retirees observed 8 to 9 years after retirement to 37 percentage points for those observed 15 to 17 years after, although the estimates for the earliest cohorts are imprecise.

The timing of the effects and their concentration among lower earners offer suggestive evidence on mechanisms. Mortality differences emerge two to three years after retirement and persist, while registration in the social-assistance registry falls only among cohorts observed ten or more years after retirement. The mortality effect is concentrated among retirees at the bottom of the pre-retirement wage distribution, for whom the pension replaces nearly all earnings and who are most exposed to financial constraints late in retirement. Together these patterns are consistent with lump-sum recipients depleting their savings, although they do not establish this mechanism directly. Whether guaranteed income also changes health behavior or investment in health remains an open question.[5]

This paper's first contribution is to provide causal evidence on the consequences of annuitization. Despite a large theoretical literature on the value of lifetime income insurance, there is remarkably little empirical evidence on its effects due to the identification challenges this question imposes. To the best of my knowledge, the only exception is re-

[5]The next step of this project links the retirees in the sample to Colombian health records on health care use and spending, to directly pin down these mechanisms.

cent work by Larrain et al. (2026), who study the effect of annuitization on longevity in the Chilean private pension system. This paper complements theirs along four dimensions: First, the source of quasi-experimental variation differs. In Chile, identification relies on instrumenting retirees' individual choice between annuities and phased withdrawals with recent pension-fund returns. In Colombia, by contrast, identification comes from an institutional mandate that assigns the payout form based on a threshold whose parameters are determined independently of retirees' choices. Second, the relevant counterfactual differs substantially. In Chile, phased withdrawals already spread pension wealth over expected remaining life, so annuitization primarily adds insurance against the risk of living substantially longer than expected. In Colombia, by contrast, non-annuitized retirees receive a single lump sum and bear the full market and longevity risk of financing retirement, so the comparison captures the full annuitization margin studied in the theoretical literature. Third, the two designs identify effects for different populations: retirees choosing among payout products in Chile and retirees at a mandate boundary in Colombia. The fact that both settings yield similar effects on mortality provides evidence that the consequences of annuitization are not specific to a single institutional environment or source of identifying variation. Finally, I extend the analysis beyond mortality to reliance on means-tested social assistance and poverty, shedding light on additional margins through which annuitization may affect retirees and on potential mechanisms underlying the mortality effects.

Second, this paper contributes to the literature on adverse selection in annuity markets. A long line of work documents that mortality is systematically related to annuity purchase and contract choice, interpreting these relationships as evidence that longer-lived individuals select into more longevity insurance (Mitchell et al., 1999; Finkelstein and Poterba, 2002, 2004). Philipson and Becker (1998) further show theoretically that mortality-contingent claims create incentives to invest in longevity, implying that mortality may also respond to annuity receipt. Mortality differences between annuitants and non-annuitants may reflect both ex ante selection on longevity and an ex post effect of annuitization on survival, and quantifying the split between these channels in voluntary markets is an important open question. Finkelstein and Poterba (2004) explicitly consider this ex post channel but argue that it is unlikely to be quantitatively important.[6] My estimates suggest that this channel can be substantial.

Third, this paper speaks to the literature on the annuity puzzle. A central debate in this literature is whether low annuity demand reflects optimal responses to bequest

[6]However, Einav and Finkelstein (2011) note that the standard positive-correlation test between coverage and realized claims cannot separate selection from a behavioral response to the contract, since both generate the same correlation.

motives, liquidity needs, and existing sources of lifetime income, or instead institutional and behavioral barriers that prevent retirees from obtaining valuable longevity insurance (Inkmann et al., 2011; Hershfield et al., 2026). My results show that for retirees at the mandate margin—a clear institutional barrier—annuitization lowers mortality and reduces reliance on means-tested assistance. This suggests that if low annuity take-up reflects, at least partially, the impact of administrative barriers, those barriers have consequences for outcomes that matter to retirees and policymakers.

Fourth, this paper contributes to the literature on retirement income and old-age mortality. Existing causal evidence studies the level of resources in old age, and the effects are mixed. In the United States, lower Social Security benefits reduced mortality (Snyder and Evans, 2006). In Chile, a permanent increase in pension income reduced mortality among the low-income elderly (Miglino et al., 2023). In Germany, a pension subsidy for low earners raised age at death (Malavasi and Ye, 2024). Related work shows that mortality also responds to the retirement transition (Fitzpatrick and Moore, 2018) and to the timing of income receipt, with deaths rising immediately after Social Security payments arrive (Evans and Moore, 2011). This paper identifies a different margin: the structure through which retirement wealth is paid out. Retirees just above and below the threshold hold essentially the same pension wealth at retirement, but one group receives guaranteed lifetime income while the other receives an exhaustible lump sum. The results show that old-age mortality responds not only to how much retirement income individuals receive or when they receive it, but also to how retirement wealth is transformed into income over the remainder of life. How retirement resources arrive may matter as much as how much arrives.

Finally, this paper contributes to the literature on fiscal externalities across social programs. Work on interactions across social insurance programs shows that reforms to retirement, disability, and unemployment insurance can shift beneficiaries and expenditures across programs (Staubli and Zweimüller, 2013; Borghans et al., 2014; Inderbitzin et al., 2016). These spillovers can substantially alter the net fiscal cost of a policy (Hendren and Sprung-Keyser, 2020). In pensions, existing evidence focuses primarily on changes in eligibility ages and benefit generosity (Duggan et al., 2007; Staubli and Zweimüller, 2013). I show that the form in which pension wealth is paid out can generate a similar spillover. Funded systems shift longevity risk from the government to individuals, who may subsequently rely on other safety-net programs if their savings are depleted. Consistent with this mechanism, annuitization lowers registration in the means-tested social-assistance registry among retirees observed more than a decade after retirement. This interaction runs in

the opposite direction of Bütler et al. (2017), who show that means-tested benefits can reduce annuity demand: here, annuitization itself reduces subsequent reliance on means-tested benefits. The fiscal relief that funded systems provide may therefore depend on how accumulated savings are paid out.

The shift from DB to DC pensions has consequences beyond who bears financial risk. As retirement systems move away from lifetime income guarantees, policies governing how retirement savings are paid out become increasingly important. These policies are part of a broader debate over how much control individuals should have over the timing and form of benefits from pension systems and other social programs (Currie and Gahvari, 2008; Brown, 2003). My findings show that payout design can affect survival and economic vulnerability. The results also suggest that leaving retirees exposed to longevity risk may increase their later reliance on social assistance. The fiscal consequences of how savings are paid out may therefore extend beyond the pension system to other parts of the social safety net.

The remainder of the paper proceeds as follows. Section 2 describes the institutional setting, Section 3 the data, and Section 4 the empirical strategy. Section 5 presents the results, Section 6 interprets them, and Section 7 concludes.

2 Institutional Background

2.1 The Colombian Private Pension System

In 1994, Colombia replaced its existing public pension system with a dual system consisting of two mutually exclusive regimes: a reformed public defined-benefit system and a newly created private defined-contribution system. The private regime, known as the Régimen de Ahorro Individual con Solidaridad (RAIS), is a defined contribution system based on individual retirement savings accounts administered by four private pension fund administrators (Administradoras de Fondos de Pensiones, AFPs). Formal-sector workers, including the self-employed, must affiliate with one of the two systems. Retirements under the private system remained limited during its first decade but began to grow substantially around 2004. Since then, the system has expanded rapidly: as of 2021, it covered 73 percent of formal-sector workers and 15 percent of pensioners.

There are several key features to note regarding this system. First, contributions are non-discretionary and are levied as a constant payroll tax of 16 percent on taxable earn-

ings.[7] Of this amount, 11.5 percentage points are credited to the individual account, while the remainder finances administrative fees, a funeral benefit, and a solidarity component. Second, unlike in a U.S. 401(k), affiliates do not choose how their retirement savings are invested. Investments are managed by pension funds in regulated portfolios. Third, the public and private systems are mutually exclusive. While workers may switch across systems, they can only do so up to ten years before the statutory retirement age (57 for women and 62 for men).[8] Fourth, a subset of affiliates hold a *pension bond* (*bono pensional*) alongside their account balance. Bonds recognize contributions made before the system was created in 1994 by workers who then moved into the private regime. The Ministry of Finance issues them and pays them at retirement, and their value counts toward the savings against which the payout form is determined.[9]

2.2 Forms of Pension Payout

Old-age retirees in the private system receive their savings in one of three mutually exclusive payout forms: a lifetime annuity, a subsidized pension, or a lump-sum withdrawal.[10] These payout forms are governed by statutory criteria tied to accumulated savings, age, and contribution history. Importantly, *conditional* on these criteria, the form of payout is deterministic: individuals cannot choose among the payout forms.

Individuals receive a lifetime annuity if their savings at retirement reach a threshold, which the regulation calls the *Saldo de Pensión Mínima* (SPM). Savings are measured as the balance of the individual account plus the pension bond, capitalized and updated to the calculation date.[11] The threshold stems from a constitutional mandate that no pension may fall below one statutory monthly minimum wage at any point in time.[12] Although the RAIS is a privately administered defined contribution system, this mandate effectively imposes an individual-specific minimum level of savings required to finance an annuity. I

[7] Wage workers contribute 4 percent and employers 12 percent; self-employed workers pay the full 16 percent.

[8] Decreto 692 de 1994; Ley 797 de 2003. See Becerra (2024) for a description of the public defined-benefit regime and its retirement incentives in detail.

[9] The bond is valued at the date of transfer to the private regime and grows at inflation plus a real rate of 3 or 4 percent, depending on that date (Ley 100 de 1993, Art. 117; Decreto 1748 de 1995); I capitalize it to the retirement date at this rate. Forty-one percent of retirees in the analysis sample hold a bond, 88 percent of those within the estimation window, and among holders it accounts for 67 percent of savings on average.

[10] The private system also pays disability and survivorship pensions, which are governed by distinct eligibility and benefit rules and are outside the scope of this paper. In 2023, 55.1 percent of pensioners in the private system received an old-age pension, 27 percent a survivorship pension, and 17.9 percent a disability pension. Source: own computations from Superintendencia Financiera de Colombia (2026), based on monthly reports from the pension fund administrators (Formato 495). Appendix A.1 describes these pathways.

[11] Res. 1875 de 1997; Res. 3099 de 2015.

[12] Article 48 of the Political Constitution of Colombia, as incorporated by Legislative Act 01 of 2005.

discuss the calculation of the SPM in more detail in the following subsection.

For these individuals, pension benefits are determined actuarially based on accumulated savings and realized returns. Retirees may choose among annuity products that provide either constant or time-varying payments. Benefits in all these products are contingent on survival and cannot fall below one statutory monthly minimum wage. Near the savings threshold, therefore, all payout products converge to approximately the same benefit: a lifetime pension of one monthly minimum wage.[13]

Individuals whose savings are below the SPM can only retire with a pension when they reach the statutory retirement age: 57 for women and 62 for men. After these ages, the payout form is determined by the individual's lifetime contribution history. Those who have contributed for at least 1,150 weeks receive a subsidized minimum pension paying the monthly minimum wage. Those who have not contributed a sufficient number of weeks receive their accumulated savings, including accrued returns, as a lump-sum withdrawal. The pension pathways and these criteria are summarized in Figure 1.

The empirical analysis of this paper focuses on workers who have reached the statutory retirement age but do not satisfy the 1,150-week contribution requirement. Among this group, the payout form (annuity vs. lump sum) is entirely determined by whether accrued savings exceed the SPM threshold. There are several forms of selection into the subgroup for whom this threshold is marginal relative to the universe of retirees in Colombia. Individuals in this group decided to affiliate and contribute up to retirement to the private pension system, have relatively short formal sector work histories, and have relatively high levels of accumulated savings. In the following sections, I focus on obtaining internally valid estimates of the causal effects of savings annuitization for this subpopulation.

2.3 The Savings Threshold

The SPM translates the minimum-pension floor into a savings requirement for each retiree. It is the amount of savings needed at retirement to finance a minimum-wage pension over the expected lifetime of the retiree and her eligible survivors,[14] under regulated actuarial

[13] The principal modalities are an immediate life annuity (*renta vitalicia*), a phased withdrawal (*retiro programado*), and a phased withdrawal followed by a deferred life annuity. Under phased withdrawal, benefits are recalculated as the account evolves; once the remaining balance reaches the amount required to finance the minimum pension, it must be converted into a life annuity. Appendix A.1 describes these modalities in detail.

[14] When a pensioner dies, the pension passes to her eligible survivors: the spouse or permanent partner, children under 18 or under 25 if dependent students, disabled children for life, and in their absence dependent parents (Ley 797 de 2003). The original formula prices the pension over the joint lifetime of the retiree and one beneficiary, chosen in this order. The 2015 formula allows up to two.

assumptions. I first present in Equation (1) a simplified version of the original formula, established by Res. 1875 de 1997 and in force until 2015. The complete formula is presented in Appendix A.2:

$$\text{SPM}_{i,t} \;=\; 1.05 \cdot SM_t \cdot \left[\, 14 \sum_{s=1}^{T} {}_s p_i^{LS}\, v^s \;+\; 5F_i \,\right], \qquad v = \tfrac{1}{1.04} \tag{1}$$

where SM_t is the statutory monthly minimum wage at the calculation date, ${}_s p_i^{LS}$ is the probability that the retiree or her eligible survivor remains alive s years after retirement, v discounts future payments at the regulated real technical rate of 4 percent, T is the last age of the mortality table (110), and F_i is the expected discounted value of one minimum wage paid at the retiree's death, so $5F_i$ values a funeral benefit of five minimum wages. The formula values fourteen pension payments per year, twelve monthly payments and two additional ones. It weights future payments by the probability that the affiliate or her dependent stays alive, grows the payment by consumer-price inflation, and applies a 5 percent loading for insurer expenses. For simplicity, the expression abstracts from the month of retirement. The legal formula does not, so the threshold also changes from month to month within a year.

In 2015, the Ministry of Finance replaced the original formula (Res. 3099 de 2015). The reform preserved the same economic object—the savings required to finance a lifetime minimum-wage pension—but made the threshold responsive to current market conditions. In particular, the computation began to incorporate a market-linked real interest rate updated every six months, recent inflation, and an explicit productivity-growth assumption.[15] As a result, the value of the minimum pension is periodically recomputed to reflect changing economic conditions. The formula and its inputs, however, remain centrally determined and publicly available, allowing the threshold to be reconstructed for each calculation date.

A final change took effect in 2018 (Res. 3023 de 2017). Pension funds obtained the ability to raise the threshold above the SPM, within narrow regulatory guidelines. Since then, each fund applies its own threshold, computed from a technical note it submits to the financial regulator, the Superintendencia Financiera de Colombia, for approval. The note preserves the same actuarial structure but allows fund-specific parameter updates and finer differentiation by sex, age, and family structure—for example, probabilities of acquiring future beneficiaries, more frequently updated mortality tables, and separate

[15] The update also refreshed the mortality tables and expanded the survivor structure from one substitute to up to two beneficiaries. Appendix A.2 presents the 2015 formula in full.

short- and long-run assumptions on inflation and productivity growth. This flexibility is tightly bounded. Funds have no discretion over any individual's payout form, the threshold is fixed ex ante by the approved note, the resulting thresholds cannot vary by more than 5 percent across funds, and the formula may not use individual health, morbidity, or any other risk-classification information.[16] The threshold each retiree faces therefore remains deterministic, individual-specific, and regulator-approved; what changed in 2018 is that its exact level became private to each fund. Because the technical notes are not public, the post-2018 threshold cannot be reconstructed from published parameters alone. The estimation sample therefore covers retirees of 2008 through 2017.

The formula determines a threshold for each retiree at the time of calculation, based on individual characteristics and policy parameters that change over time. Holding policy parameters fixed, the threshold is higher for women, lower for individuals who retire at older ages, and higher for retirees with eligible beneficiaries, particularly younger ones. Figure 3 illustrates these differences for four representative profiles between 2008 and 2023. At statutory retirement age, the threshold is 18–24 percent higher for a single woman than for a single man, reflecting her younger retirement age and longer life expectancy. Retiring five years later lowers a man's threshold by 13–14 percent. A 62-year-old man with a 57-year-old wife faces a threshold 32 percent higher than an otherwise identical single man. Thus, two retirees with the same savings and otherwise similar characteristics can face different thresholds, and therefore receive different payout forms, simply because their spouses differ in age. For this reason, the empirical design in Section 4 measures each retiree's savings relative to her own threshold. Beyond these individual differences, thresholds also vary over time as the minimum wage and policy parameters change. The four series move in parallel, with gaps reflecting demographic differences and their common movement reflecting the calendar inputs described next.

The policy inputs used to calculate the threshold are set by different entities and updated at different frequencies. The minimum wage is the most important. Holding other inputs fixed, a given percentage increase in the minimum wage raises the threshold by the same percentage. Each December, the minimum wage is set for the following year.[17] Inflation in the preceding year sets a floor, but increases above that floor follow no fixed

[16] Personal health data are classified as sensitive under Colombian law, and their processing is prohibited when it may generate discrimination (Ley 1581 de 2012, Ley 2475 de 2025).

[17] Formally, the minimum wage is negotiated each December by a tripartite commission representing the government, employers, and workers. If the commission does not reach consensus, the national government sets the following year's minimum wage by decree. Between 2000 and 2023, the parties reached agreement in 9 of 23 negotiations; in the remaining 14, the adjustment was determined by the government. See Ley 278 de 1996, arts. 2 and 8.

rule. Figure 2 shows substantial variation in this component. Between 2006 and 2023, the gap between minimum-wage growth and lagged inflation ranged from near zero to 4.5 percentage points, with no stable pattern.

Other inputs also change over time. The technical interest rate, recent inflation, and mortality tables are updated at different intervals, some every six months and others irregularly. Appendix A.2 details each parameter, its source, and how often it is updated. Together, these changes make both the magnitude and direction of future threshold movements difficult for workers to anticipate.

Figure 4 compares the annual change in the threshold with one year of mandatory contributions. The bars show the year-to-year threshold change for a single man aged 62, and the dashed lines show a year of contributions at 2, 3 and 5 minimum wages. The threshold moves by about USD 1,750 a year on average in absolute value, four times the USD 420 that a retiree near the cutoff contributes in a year at the median wage. A year of contributions often does not cover the change even at higher wages. At twice the minimum wage, contributions fall short of the change in 10 of 15 years, and at five times the minimum wage in 6. The direction is not fixed either. The threshold rose sharply after the regulatory revisions of 2010 and 2015 and fell in three other years. A worker who delays retirement by a year therefore faces a shift in the threshold that is uncertain in sign and often larger than her added contributions. The threshold is also never disclosed. Affiliates learn only which payout form they qualify for when they apply to retire. Timing retirement to fall on one side of the threshold is therefore very challenging and unlikely. Section 4.3 presents empirical evidence on this.[18]

The threshold thus defines a deterministic assignment rule. Among workers at the statutory retirement age with fewer than 1,150 contribution weeks, those with savings at or above their threshold receive an annuity and those below it receive a lump sum. The empirical design compares retirees just above and just below this cutoff. Section 4 defines the running variable, states the identifying assumption, and presents the evidence that workers do not sort around the threshold.

[18]The benchmark counts only mandatory contributions credited to the account and excludes investment returns, which add further variation in savings.

3 Data

3.1 Universe Administrative Records

The primary data source consists of confidential administrative records covering the universe of affiliates to Colombia's private pension system. Each pension fund administrator generates an annual cross-section containing all individuals affiliated with the fund at the end of the year, including active contributors, inactive affiliates, and retirees. I harmonize these cross-sections across funds and years and link individuals over time using their national identification numbers. The resulting panel spans 2008–2023, except 2019, when the cross-sections were not generated. Individuals enter the panel when they first appear in a fund's records and remain observable while affiliated with the private system, including after they stop contributing or retire; individuals who transfer to the public pension system cease to appear.[19] These records therefore cover the universe of individuals observed as affiliated with Colombia's private pension system during the study period. To the best of my knowledge, this is the first paper to use these records for research.

The records report each affiliate's sex, birthdate, date of affiliation, and municipality of residence. Before retirement, the annual cross-sections additionally contain accumulated pension savings and contribution weeks, the amount and date of the most recent contribution, the earnings on which it was based, the type of affiliation (salaried, self-employed, or mixed), the investment portfolio, and the value of any pension bond.[20] These variables allow me to follow the evolution of affiliates' contributions, earnings, and pension wealth through retirement. At retirement, I observe the payout form—annuity, lump sum, or subsidized pension—and, depending on the payout form, either the annuity value or the amount withdrawn as a lump sum. The unit of observation throughout is the individual affiliate, and the records contain no household or marriage identifiers that would allow me to link spouses directly. Family pensions, in which spouses pool their savings to qualify for a pension, are observed as a separate payout form and excluded from the analysis sample.

I construct the analysis sample from the stock of retirees observed in the 2023 cross-section.[21] Because retirees remain affiliated with the private system after retirement, this cross-section identifies individuals who retired at any point between 2008 and 2023. I

[19] Transfers to the public system are allowed only until ten years before the statutory retirement age. Most individuals who disappear from the records do so as they approach that deadline, and the records identify their departure as a transfer.

[20] The records report the bond's nominal value at issuance, not its value at retirement.

[21] The retiree stock includes all individuals ever registered as retirees in the private system, including those who died before 2023. The sample therefore does not condition on survival to 2023.

first restrict the sample to ordinary old-age claims, excluding disability, survivorship, and family pensions because they are governed by different eligibility and financing rules.[22] Among the remaining retirees, I retain individuals who retired at or after the statutory retirement age—62 for men and 57 for women—with fewer than 1,150 weeks of contributions and who received either an annuity or a lump-sum refund. These restrictions isolate the population subject to the threshold rule studied in this paper: affiliates who have reached retirement age but do not qualify for the minimum-pension guarantee and whose accumulated savings therefore determine whether they receive an annuity or withdraw their balance as a lump sum. For each retiree in the resulting sample, I retain all available annual records, including the contribution and savings history observed before retirement.

For affiliates of the largest pension fund, which covers approximately 40 percent of retirees in the system, I additionally observe the family structure recorded at retirement. The records report whether the retiree has eligible dependents and, for each dependent, the relationship to the retiree (spouse, child, or parent), sex, birthdate, and disability status. Because the age, sex, and relationship of eligible dependents enter the threshold formula, these records give the family structure needed to compute the threshold for this fund's retirees. For retirees of the other funds, I use them to impute family structure, as Section 4.1 and Appendix B.2 describe.

3.2 Linked Outcomes

I link the universe of retirees by national identification number to two administrative sources, one for mortality and one for living standards.

Mortality. The death records of the National Civil Registry (*Registraduría Nacional*) report vital status as of June 30, 2025 and, for the deceased, the date of death. Because vital status is observed on a fixed date, follow-up varies by cohort: retirees of 2008 are observed for up to 17 years, retirees of 2023 for two.

Poverty. Sisben, maintained by the National Planning Department (*Departamento Nacional de Planeación*), is the registry Colombia uses to target social programs, known as the census of the poor (Camacho and Conover, 2011; Londoño-Vélez et al., 2020). Individuals enter or update their record by requesting a household survey, usually when applying

[22] Old-age pensions account for 70.2 percent of all retirees in the private system, while disability, survivorship, and family pensions account for 8.4, 20.3, and 0.1 percent, respectively.

for a social program, or through the door-to-door sweeps that municipalities conduct when a new version of the registry is rolled out. The survey records the household's living standards, including housing, assets, education, and composition, and summarizes them in a score. The score is the official instrument for classifying households as poor and for determining program eligibility. Richer households are not included in the sweeps and are unlikely to request the survey, so being in the registry is itself a proxy for poverty. From each wave I construct an indicator for registration in Sisben.

I observe two waves of the registry. Sisben III was collected mainly between late 2008 and 2010 and remained in use through 2020. Among retirees matched to it, 97 percent were surveyed before retirement, so it provides a pre-retirement measure. Sisben IV was collected mainly between 2017 and 2020, became the official classification system in 2021, and has been continuously updated since. I observe it as of June 2025, so it provides the post-retirement outcome.

Descriptive statistics. Table 1 describes the sample of interest: all 423,255 retirees of the private system between 2008 and 2023 who reached the statutory age with fewer than 1,150 weeks of contributions and received an annuity or a lump sum. Column (1) is the full sample; columns (2) and (3) split it by payout form. Lump sums are 97 percent of the sample and annuities 3 percent. The typical retiree is a woman who retired at 60, close to the statutory age, with 413 weeks of contributions, eight years, and a wage of about 1.5 minimum wages in the year before retirement. By June 2025, 5 percent had died, 60 percent were registered in Sisben, and 30 percent were classified as poor. Three-quarters of the sample retired after 2017. Annuitants differ from lump-sum retirees on every characteristic: they contributed for almost twice as many weeks, earned 4.2 minimum wages against 1.3, reached retirement with far larger savings, and are one-sixth as likely to be registered in Sisben.

4 Empirical Strategy

My empirical strategy exploits the discontinuity that the savings threshold creates in the payout of retirement savings. Among affiliates who reach the statutory retirement age with fewer than 1,150 weeks of contributions, those whose savings equal or exceed their threshold must annuitize them, whereas those below it can only withdraw their accumulated balance as a lump sum. The running variable is the difference between savings at retirement and the retiree's individual threshold, normalized to zero at the cutoff. Because

retirees cannot choose between the two payout forms, assignment is a sharp regression discontinuity in the true distance to the threshold.

I do not observe retirees' true distance to the threshold. Both savings at retirement and the threshold are reconstructed from administrative records, introducing error into the measured distance. This error can place retirees on the wrong side of the cutoff, so the sharp rule is fuzzy in the measured distance. Despite this error, the probability of annuitizing remains discontinuous at the measured cutoff. I use this discontinuity to estimate the effect of annuitization and state below the assumptions required for a causal interpretation.

4.1 Measurement Error and Identification

Let R_i^* denote retiree i's true distance to the threshold, defined as savings at retirement minus her individual threshold. The fund uses this distance to assign the payout form. I observe instead $R_i = R_i^* + e_i$, where e_i combines errors in reconstructed savings and the computed threshold. The payout form is recorded directly, so measurement error enters only through the running variable.

Measurement error arises in both the threshold and savings at retirement. The threshold depends on the retiree's family structure, since eligible dependents extend the pension's expected payment horizon. Because three of the four funds do not report dependents, I impute family structure from the largest fund's records within cells of sex, age, and retirement year. Savings at retirement must also be reconstructed. The records report year-end balances, up to twelve months before retirement. During that interval, savings change with contributions, investment returns, and, for bond holders, the crediting of the pension bond. I project the last recorded balance to the retirement month to account for these changes. Errors in either the threshold or savings can place a retiree on the wrong side of the measured cutoff. Appendix B details the computation of the threshold, the imputation of family structure, and the projection of savings to the retirement date.

This setting can be represented by a mixed sample, in which the distance to the threshold is measured correctly for some retirees and with error for others, as in Battistin et al. (2009). The observed first stage is consistent with this structure.[23] For correctly measured retirees, the measured cutoff coincides with the threshold used by the fund, preserving

[23] If every retiree were measured with a continuously distributed error, the probability of annuitizing would rise smoothly through the measured cutoff and there would be no discontinuity to divide by (Davezies and Le Barbanchon, 2017); the fuzzy estimand would be undefined (Pei and Shen, 2017). A discontinuity remains only if a share of retirees is measured correctly, which is the case studied by Battistin et al. (2009). The jump in annuitization at the measured cutoff is therefore itself evidence that such a share exists.

the discontinuity in annuitization. Under the conditions stated below, the fuzzy estimator identifies the effect of annuitization. A simulation in Appendix C.1 examines whether the estimated measurement errors can reproduce a first-stage estimate of the observed magnitude.

This discontinuity identifies the effect of annuitization under three conditions. First, among correctly measured retirees, expected potential outcomes under each payout form are continuous at the cutoff. Their average mortality and living standards would therefore not change discontinuously at the threshold if they all received the same payout form. Second, a positive share of retirees at the measured cutoff is measured correctly. For these retirees, the measured distance coincides with the distance used by the fund. Third, the measurement error varies smoothly through the threshold. Neither the share of mismeasured retirees nor their average annuitization and outcomes changes discontinuously at the threshold. Under these conditions, the usual fuzzy regression discontinuity estimator identifies the average effect of annuitization for correctly measured retirees at the cutoff. An additional assumption, that average treatment effects do not differ between correctly measured and mismeasured retirees, extends this interpretation to all retirees at the measured cutoff. Appendix C.2 provides the formal proof.

4.2 Estimating Equations and Implementation

I use a fuzzy regression discontinuity design to estimate the effect of annuitization on mortality and poverty. The running variable is the measured distance to the threshold, R_i, and an indicator for being at or above the measured cutoff, $Z_i = \mathbb{1}\{R_i \geq 0\}$, instruments for annuitization. To estimate the effect on mortality, I construct a person-year panel that follows each retiree from the year after retirement until death or the end of follow-up in 2025. The outcome is an indicator for dying in that year. Its mean among retirees alive at the start of the year is the annual probability of death.[24] Let i index retirees, j years since retirement, t calendar years, and $c(i)$ retirement cohorts. I estimate the following equations by two-stage least squares,

[24] The specification is similar to Sullivan and von Wachter (2009), who estimate a person-year mortality model and translate the estimates into life expectancy, and to Larrain et al. (2026), who estimate the effect of annuities on mortality in Chile with a linear person-year panel and two-stage least squares. Their outcome is cumulative mortality rather than the annual hazard. I estimate a linear probability model because its linear structure allows instrumental-variable estimation, which nonlinear duration models do not accommodate directly (Palmer, 2023).

$$A_i = \pi_0 + \alpha\, Z_i + f(R_i) + g(a_{ij}) + \gamma_{c(i)} + \phi_t + \mathbf{x}_i'\boldsymbol{\pi} + v_{ij}, \tag{2}$$

$$y_{ij} = \beta\, A_i + f(R_i) + g(a_{ij}) + \gamma_{c(i)} + \phi_t + \mathbf{x}_i'\boldsymbol{\delta} + \varepsilon_{ij}. \tag{3}$$

The annuity indicator A_i equals one if retiree i annuitized and zero if she received the lump sum. Both A_i and R_i are fixed at retirement. The function $f(R_i)$ is linear in the measured distance with separate slopes on each side of the cutoff. The remaining terms are controls. $g(a_{ij})$ is a quartic in current age, which captures the age profile of mortality. $\gamma_{c(i)}$ and ϕ_t are retirement-year and calendar-year fixed effects, which absorb differences across cohorts in the level of the threshold and in the composition of retirees, and aggregate mortality shocks common to all retirees alive in a given year. $\mathbf{x}_i$ is sex. I estimate the equations by two-stage least squares within a window of 90 minimum wages on each side of the cutoff, with a uniform kernel and standard errors clustered by retiree.[25] The coefficient α is the jump in the probability of annuitizing at the cutoff. The coefficient of interest, β, is the difference in the annual probability of death between annuitants and lump-sum retirees at the threshold, under the conditions of Section 4.1.

I next allow the effect to vary with time since retirement. I group the years after retirement into bins and estimate a separate effect in each. Let $B_{ij}^b = \mathbb{1}\{j \in b\}$ indicate that year j after retirement falls in bin b, for the bins $\{1\}$, $\{2,3\}$, $\{4,5\}$, $\{6,\ldots,10\}$, and $\{11+\}$. The dynamic specification is

$$y_{ij} = \sum_b \beta_b\, A_i B_{ij}^b + \sum_{b>1} \mu_b B_{ij}^b + f(R_i) + g(a_{ij}) + \gamma_{c(i)} + \phi_t + \mathbf{x}_i'\boldsymbol{\delta} + \varepsilon_{ij}, \tag{4}$$

where each interaction $A_i B_{ij}^b$ is instrumented by $Z_i B_{ij}^b$ and the remaining terms are as in Equation (3). The coefficient β_b measures the effect of annuitization on the annual probability of death in bin b, conditional on being alive at the start of each year, and μ_b allows the baseline hazard to differ across bins. The bins widen with time since retirement to maintain sufficient observations as fewer cohorts contribute to later years. The hazard is the main specification because its annual death probabilities combine into survival curves, which express the effects as years of life. As a complement, I also estimate cumulative mortality within two, five, and ten years of retirement, with one observation per retiree, among the cohorts observed for the full horizon. Appendix D.2 reports the specification

[25] The window is the mean-squared-error-optimal bandwidth of Calonico et al. (2014) for the reduced form, computed at the retiree level on each retiree's realized annual mortality rate, which is the quantity the hazard averages. Appendix D.1 reports the selection and the results for alternative windows.

and the results.

The outcomes on living standards are observed once, in June 2025, so the same design applies with one observation per retiree. The time index, the age polynomial, and the calendar-year fixed effects drop out of Equations (2) and (3), and the retirement-year fixed effects remain. I report a pooled estimate for all cohorts and estimates by groups of retirement cohorts. For the latter I follow the dynamic specification of Equation (4), with indicators for cohort groups in place of the bins, and let the slopes in the running variable vary by group, since the groups are different retirees. The groups are meant to trace the effect as time since retirement passes. Because they are also different cohorts, differences across them can reflect cohort differences as well as exposure. Registration before retirement is balanced in every group, which limits this concern.

4.3 Evidence Supporting Identification

The conditions of Section 4.1 require that neither retirees nor pension funds manipulate position relative to the threshold. Institutional features make such manipulation challenging and unlikely. This section discusses these features, then presents evidence on the one margin that remains open to workers, the timing of retirement, and on the continuity of the density of the running variable and of predetermined characteristics at the threshold.

4.3.1 Constraints to Manipulation

The threshold is never disclosed to affiliates. Affiliates can observe their accumulated savings at any time, since funds report balances online or on request, but they never observe the value against which that balance will be compared. Funds provide simulations of future accrual toward a pension, but these carry explicit disclaimers that the projections change with the legal parameters and the state of the economy, and they commit the fund to nothing. Even an affiliate who attempted to infer her threshold would be aiming at a moving target. As Section 2.3 showed, the threshold depends on parameters that are updated at different frequencies, and its annual change averages 4.1 times a year of contributions at the median wage of retirees near the threshold (Figure 4). The balance itself moves at an even finer frequency: it fluctuates daily with the market returns of the fund's portfolio, which change the distance to the threshold in either direction and are unknown in advance.

Workers also have limited control over pension savings. Contributions are a fixed share of taxable earnings, of which 11.5 percentage points are credited to the individual

account, and workers cannot alter it.[26] Nor can affiliates choose how their savings are invested. Portfolio allocation is largely determined by age, and affiliates within ten years of the statutory retirement age are placed in conservative portfolios whose realized returns are very similar across funds.

Pension funds have no discretion over assignment either. Through 2017 the threshold was a statutory formula, and since 2018 each fund applies the formula of a technical note approved by the financial regulator, which may not differ by more than 5 percent from the reference value or use any individual information beyond the sex and age of the affiliate and her beneficiaries. Balances and payouts are reported to the regulator at the individual level every month and audited against the aggregate figures. These constraints make manipulation of savings implausible, for workers and pension funds alike. The remaining margin is the timing of retirement, which I address next.

4.3.2 Retirement Timing

A worker who reaches the statutory age with savings below her threshold could postpone retirement in the hope that her balance eventually exceeds it. Delaying, however, raises savings but also changes the threshold, which is recomputed with the next year's parameters. Because the threshold's annual change can be comparable to a full year of contributions, delaying does not reliably close the gap. Retirement also remains concentrated at the statutory age. In the sample, 60 percent retire at exactly that age and three-quarters within a year of it; Figure 5 shows the retirement hazard by age, which spikes at the statutory age and halves within two years.

I test whether crossing the threshold affects retirement timing. If workers delay retirement to qualify for an annuity, a worker whose savings cross the threshold should be more likely to retire that year than an otherwise similar worker just below it, so the probability of retiring should jump at the threshold. I use a person-year panel that follows each worker from the year she reaches the statutory retirement age until retirement. In each year, I compute the distance between her savings, projected to that year, and the threshold in force that year, and estimate

$$\text{Retire}_{it} = \theta\, Z_{it} + f(R_{it}) + \phi_t + \varepsilon_{it}, \qquad \text{for } |R_{it}| < h, \tag{5}$$

where Retire_{it} equals one if worker i retires in calendar year t, R_{it} is her distance to the

[26] Affiliates may also hold voluntary savings in a separate account that counts toward the balance compared with the threshold, but fewer than 0.1 percent of affiliates held any in 2018.

threshold in force that year, $Z_{it} = \mathbb{1}\{R_{it} \geq 0\}$, $f(R_{it})$ is the linear function with separate slopes on each side of the cutoff as in Equation (2), and ϕ_t are calendar-year fixed effects. The coefficient θ is the jump in the probability of retiring at the threshold. I estimate Equation (5) by local linear regression within the same window as in Section 4.2, with standard errors clustered at the worker level.

The retirement probability is smooth through the threshold. Figure 6 plots the share of workers retiring in each bin of the distance to the threshold within the estimation window, together with the local linear fits on each side of the cutoff. The estimated jump is −0.011, with a robust bias-corrected standard error of 0.034, relative to an annual retirement rate of about 50 percent. A worker just above the threshold is therefore no more likely to retire than one just below it. I find no evidence that position relative to the threshold affects retirement timing.

4.3.3 Density and Predetermined Characteristics

If retirees sorted around the threshold through any channel, the density of the running variable would be discontinuous at the cutoff. I test this with the local polynomial density test of Cattaneo et al. (2020), which estimates the density on each side of the cutoff directly from the distance to the threshold and tests the equality of the two limits at zero with robust bias-corrected inference. Figure 8 shows the histogram of the distance together with the estimated density and its confidence bands on each side of the cutoff. The density is continuous through zero: the local quadratic test within the same window gives a statistic of 0.47 and a robust p-value of 0.64. The test speaks to sorting in the measured distance; sorting in the true distance would appear in it only to the extent that the measurement error does not blur it.

If potential outcomes are continuous at the threshold, characteristics determined before retirement should be continuous there as well. I test this by estimating Equation (2) with each predetermined characteristic in place of the annuity indicator, so that the coefficient on Z_i is the jump in the characteristic at the cutoff. Each characteristic is estimated within its own mean-squared-error-optimal bandwidth, with a triangular kernel, retirement-year fixed effects, and robust bias-corrected inference. Table 2 reports the results for sex, age at retirement, residence in a department capital, accumulated contribution weeks, the last wage observed before retirement, and savings at retirement.[27] I find no evidence that any of these characteristics is discontinuous at the threshold. The estimates are small

[27] The wage is the last one observed before retirement, available for 57 percent of the sample, rather than the wage in the year immediately before retirement, which is observed for far fewer retirees.

relative to the means on each side and none is statistically significant.

The largest fund's records provide a further check. Because dependents are recorded there, I can test the continuity of family structure, the input that the universe threshold imputes. Among retirees of the largest fund with a record, the estimated discontinuity in having any registered dependent is −0.008, with a p-value of 0.67. Table 2 also reports balance in Sisben registration before retirement, restricting the comparison to records dated before the retirement year; it is continuous as well. Figure 7 plots these estimates.

To summarize, affiliates do not observe their position relative to the cutoff, retirement timing does not seem to respond to it, and both the density of the running variable and predetermined characteristics are continuous through it. This evidence supports the continuity and smooth-measurement conditions of Section 4.1.

5 Results

This section presents the results in four parts. The first examines the first stage and whether its magnitude is consistent with the measurement error described in Section 4.1. The second evaluates the effect of annuitization on mortality, both as a constant effect on the annual probability of death and by years since retirement, and translates these estimates into years of life. The third turns to poverty, pooling retirement cohorts and then examining groups of cohorts separately. The fourth places the magnitudes in context.

For each outcome, I report reduced-form and 2SLS estimates. The reduced form captures the effect of crossing the measured threshold. The 2SLS estimate captures the effect of annuitization for correctly measured retirees at the cutoff and, if the effect of annuitization does not differ between correctly measured and mismeasured retirees, for all retirees at the cutoff.

5.1 First Stage

Figure 9 shows the share of retirees who annuitized against their distance to the threshold. Each marker represents the share of retirees who annuitize within a bin spanning 30 minimum wages (about USD 8,000 at 2023 prices)[28], with marker size proportional to the number of retirees in the bin. Far below the threshold no one annuitizes. The share rises in the last bins before the cutoff, jumps at zero, and keeps rising above it until it approaches one. Table 3 reports the formal first stage. It estimates Equation (2) within

[28]Peso amounts are converted to US dollars at 4,000 pesos per dollar.

the optimal bandwidth of the specification in Section 4.2. Being at or above the threshold raises the probability of annuitizing by 27.5 percentage points, with a first-stage F-statistic of 147.

The first stage is below one, even though the rule requires the probability of annuitizing to rise from zero to one at the true threshold. The estimate of 0.28 is therefore the share of retirees near the measured cutoff whose measured and true sides coincide. A share this large and precisely estimated is evidence that correctly measured retirees exist, the second condition in Section 4.1 for identification. The gradual rise in annuitization on either side of the cutoff is consistent with measurement error among the remaining retirees. To assess whether measurement error can explain the smaller first stage, I simulate errors using the distributions observed among retirees whose dependents and savings at retirement are known. Errors of this size explain 87 percent of the gap between one and the estimated first stage. Appendix C.1 provides the details of the simulation exercise, and Table C.1 and Figure C.1 report its results.[29]

5.2 The Effects of Annuitization on Longevity

Table 4 reports the effect of annuitization on the annual probability of death from Equations (2) and (3). Column (1) reports the reduced form. Being at or above the measured threshold lowers the annual probability of death by 0.30 percentage points, against a hazard of 0.95 percent per year among retirees just below the cutoff. Column (2) reports the 2SLS estimate. Annuitization lowers the annual probability of death by 0.96 percentage points, against a baseline hazard of 1.26 percent per year. Both estimates are negative and significant. The 2SLS estimate is the reduction in the annual mortality hazard for correctly measured retirees at the cutoff, who are the retirees for whom crossing the threshold changes the payout form.

Figure 10 plots the effect of annuitization on the annual probability of death by years since retirement, from Equation (4), which lets the effect differ across five groups of years. There is no effect in the first year after retirement. The effect appears in years 2 to 3, at −0.97 percentage points, and is strongest in years 4 to 5, at −1.10 percentage points. It then remains close to one percentage point in years 6 to 10 and after year 10. The point estimate therefore increases over the first three groups, although the effects from

[29]The size of the first stage depends on how I impute family structure. With the dependents recorded by the largest fund in place of the imputation, the estimate rises from 0.21 to 0.28. Pricing every retiree as single lowers it to 0.17, and computing the cell median over retirees with a priceable spouse gives 0.20. Appendix D.3 reports these estimates. The imputation changes how many retirees are priced correctly and so the size of the first stage, but a strong discontinuity remains in every case.

year 2 onward are not statistically different from one another. The reduced form follows a similar pattern, with estimates between −0.31 and −0.39 percentage points from year 2 onward. The results are similar under an alternative specification that estimates cumulative mortality within fixed horizons after retirement, detailed in Appendix D.2. Annuitization lowers the probability of dying within two and five years of retirement by 3.0 and 6.0 percentage points, on baselines of 2.3 and 7.2 percent, and the ten-year estimate is of the same order but imprecise (Table D.1). Table 5 reports the estimates of the hazard by years since retirement.

To express the mortality effects in years of life, I convert the estimated annual hazards into survival probabilities.[30] For each year after retirement, I compute the probability of surviving that year with and without the annuity, using the baseline hazard and the estimated effect, either the constant effect of Table 4 or the effect by years since retirement of Table 5. Rolling these probabilities forward gives a survival curve under each payout form, and summing each curve gives expected years lived. The difference between the two sums is the gain in years of life from annuitization. Over the 17 years observed in the data, the calculation uses only the estimated hazards. To extend it to remaining life expectancy, I take sex-specific mortality tables beyond the observed period and continue rolling the survival probabilities forward to age 110. I report the gains for both the reduced form and the 2SLS estimates.

Table 6 reports the results. Over the 17 observed years, crossing the threshold adds 0.4 years of life and annuitization adds 1.3 years, under both the constant and the bin-specific effects. In remaining life expectancy at the statutory retirement age, which is 28.3 years for a woman retiring at 57 and 20.0 years for a man retiring at 62, crossing the threshold adds 1.1 and 0.7 years and annuitization adds 2.5 and 2.3 years under the constant effect, with nearly identical gains under the bin-specific effects. These figures apply the estimated effect over the observed years and leave the mortality tables unchanged afterwards. Keeping the effect through the end of life instead raises the gains from annuitization to 4.0 years for women and 2.8 for men.

5.3 The Effects of Annuitization on Living Standards

Figure 11 shows the effect of annuitization on registration in the social-assistance registry (Sisben). Panel A reports reduced-form estimates. Red diamonds report the effect on registration after retirement, measured in June 2025 among retirees alive at that date, and

[30] This follows Sullivan and von Wachter (2009), who estimate the effect of job displacement on mortality in a person-year hazard model, and translate the estimated hazards into years of life lost.

black circles report the balance test on registration before retirement. The first pair of estimates uses the full sample and the remaining pairs split it into groups of retirement cohorts, labeled by years since retirement in June 2025. Crossing the threshold lowers registration after retirement by 5.5 percentage points, significant at the 5 percent level, and has no effect on registration before retirement. The effect is concentrated in the earlier cohorts, those with more time between retirement and the observed outcome: −0.084, −0.055, and −0.051 for the cohorts retired 15–17, 12–14, and 10–11 years, respectively, and zero for the latest cohorts. Registration before retirement is balanced in every cohort. Because the outcome is measured only once, I cannot disentangle cohort effects from the effects of time since retirement. Panel B reports the 2SLS estimates, which follow the same pattern. Annuitization lowers registration by 23 percentage points in the full sample; the effect is larger for the earlier cohorts and zero for the most recent group. Table 7 reports the estimates.

These estimates are based on survivors. Registration is measured among retirees alive in June 2025, but Section 5.2 shows that annuitization increases survival. This means that some retirees above the cutoff would not have survived had they received a lump sum. If these retirees differ in their likelihood of registering, the estimates reflect both the effect of annuitization on registration and differences in who survives. To address this, I follow Lee (2009) and bound the estimates. I trim the excess share of survivors above the cutoff, first from those most likely to register and then from those least likely, and re-estimate the effects. This gives lower and upper bounds. Table 8 reports the results. The reduced-form bounds range from −0.067 to −0.051, and the 2SLS bounds from −0.28 to −0.21. Across groups, the bounds remain negative for those retired ten or more years and reach zero for those retired eight to nine years. Although the confidence intervals include zero in every group, the bounds suggest that differences in survival do not explain the negative estimates.

Taken together, registration is balanced at the threshold before retirement but lower after retirement, particularly among cohorts observed ten or more years later. The pooled estimate is statistically significant, and the pattern across cohorts remains after accounting for survival selection, although the bounds are imprecise. Section 6 examines this pattern alongside the timing of the mortality effects.

5.4 Magnitudes in Context

The estimated effects are large relative to baseline outcomes. Crossing the threshold reduces the annual probability of death by 0.30 percentage points, 31 percent of the hazard

below the cutoff, and annuitization reduces it by 0.96 percentage points, 76 percent of the baseline hazard. Among survivors, crossing the threshold lowers registration in the social-assistance registry by 5.5 percentage points, 16 percent of registration below the cutoff, and annuitization lowers it by 23 percentage points, 62 percent of the baseline. These are large effects, and there is no direct evidence on annuitization against which to assess them. I therefore use two imperfect benchmarks. The first is the OLS difference between annuitants and lump-sum retirees at every distance from the threshold, which combines the effect of annuitization with the differences between retirees far above and far below the cutoff and serves as a descriptive reference. The second is the closest causal evidence on annuitization and mortality, from Larrain et al. (2026) in Chile. Their counterfactual is a phased withdrawal rather than a lump sum, which limits the comparison.

The first benchmark is a naive OLS regression of the annual hazard of death on annuitization. It uses all retirees in my sample, with the same controls as the main specifications but no restriction on distance to the threshold. Annuitants and lump-sum recipients in this sample are far apart in distance to the threshold and may differ substantially in pension wealth, earnings histories, and other characteristics. The OLS estimate therefore combines the effect of the payout form with pre-existing differences between the two groups. It shows that annuitants have an annual hazard of death 89 percent lower than lump-sum recipients (Table 9), compared with reductions of 76 percent from 2SLS and 31 percent from the reduced form. The larger OLS gap is consistent with evidence from voluntary annuity markets that longer-lived individuals select into annuities (Finkelstein and Poterba, 2002). However, this selection mechanism does not necessarily apply here. In Colombia, payout form is assigned by an institutional rule rather than chosen by retirees. Annuitants may nevertheless have better survival prospects because they have higher pension wealth, earnings, and potentially other favorable characteristics. The OLS gap may therefore serve as an approximate ceiling for the proportional effect of annuitization, although not a formal upper bound.

The second benchmark is Larrain et al. (2026), the closest causal evidence on annuitization and mortality. In Chile, retirees choose between an annuity and a phased withdrawal, and the authors instrument this choice with pension-fund returns in the six months before retirement. Their outcome is cumulative mortality in each year after retirement, averaged over the first five or ten years, and their 2SLS estimates are -2.55 and -3.62 percentage points over these horizons. These estimates come from a different specification and outcome, so they are not comparable with my estimates of the annual hazard. I therefore replicate their specification within my design, keeping their panel structure,

outcome, and horizons, but instrumenting annuitization with the threshold indicator and restricting the sample to the optimal bandwidth. Appendix E.1 describes the implementation. My 2SLS estimates are −3.08 percentage points over five years and −4.79 over ten years, 1.2 and 1.3 times their estimates (Table E.1).

The two sets of estimates are of the same order of magnitude, with larger point estimates in Colombia. One reason to expect larger effects in Colombia is the difference in the counterfactual. In Chile, retirees who do not annuitize receive phased withdrawals, which spread the balance over remaining life expectancy in annual installments. In Colombia, these retirees receive their entire balance as a lump sum and must manage its drawdown themselves. The Colombian setting therefore captures a larger change in the structure of retirement income, which could explain the larger effects. However, two differences limit the comparison. First, the estimates identify effects for different populations: retirees whose choices respond to pension-fund returns in Chile and retirees near the threshold in Colombia. Second, the annuity contracts and the sources of identifying variation differ across settings.

Two conclusions follow. First, the reduced form is a lower bound for the effect of annuitization, since crossing the measured threshold changes the payout form for only a share of retirees. On this bound, annuitization reduces the annual probability of death by at least 31 percent, which is 0.4 years of life over the 17 observed years, and reduces registration in the social-assistance registry by at least 16 percent. Second, the 2SLS estimate is large, but it is below the OLS gap of 89 percent and within 1.2 to 1.3 times the only other causal estimate of annuitization on mortality.

6 Interpretation of Results

The previous sections show that annuitization reduces mortality and registration in the social-assistance registry. This section considers which mechanisms are consistent with these effects and their timing. The reduction in mortality is absent in the first year after retirement, appears in the second and third years, is largest in the fourth and fifth, and persists for as long as the data follow retirees. The reduction in registration is zero among cohorts retired eight to nine years and present among those retired ten or more years. A natural benchmark for this timing is the standard model of annuitization (Yaari, 1965; Davidoff et al., 2005). In this model, the payout form matters through consumption. An annuity pays a return conditional on survival that exceeds the return on conventional

assets, so the same savings fund a higher consumption level at every age.[31] If the main outcomes respond to consumption, the model implies an effect that is present from retirement and grows with time since retirement. The prediction fits the data only in part. The registration effect appearing after a decade is consistent with a gap that widens. But the model implies a gradual onset. The mortality effect is instead absent in the first year and large soon after.

I consider three potential explanations, which are not mutually exclusive. First, the pension may be worth more than the lump sum, because payments rise with the minimum wage, so crossing the threshold changes not only the form of the payout but also its value. Second, lump-sum recipients may spend their savings faster than the standard model predicts, leaving them with limited resources within a few years. Third, guaranteed lifetime income may affect retirees even before savings run out, by reducing financial stress or by changing how much they invest in health and when they seek care. The data do not allow me to test these mechanisms directly or to separate their contributions to the estimated effects. Instead, I provide suggestive evidence on the plausibility of each.

6.1 The Value of the Pension at the Threshold

The first explanation is that the pension is worth more than the lump sum it replaces. A retiree at the threshold gives up her savings and receives a pension of one minimum wage for life. By law, the pension is indexed to the minimum wage, which grew faster than prices over the period. The regulator's original formula, however, prices the pension as constant in real terms.[32] As a result, the pension may be worth more than the lump sum it replaces.

To assess this possibility, I compare what the pension is worth with what it costs, the savings threshold, following Mitchell et al. (1999) and Poterba (2001). The threshold is the regulator's estimate of the pension's expected present value, so under the regulator's assumptions the pension is worth exactly its price. The exercise recomputes the pension's value under more realistic assumptions, holding its price fixed at the threshold, and asks whether the pension is then worth more than it costs. The regulator relies on three assumptions. First, the original formula assumes that the pension remains constant in real

[31] Without annuities and without a bequest motive, the retiree sets consumption against the risk of a long life, so it declines with the survival probability. With an annuity, mortality risk is pooled and consumption is constant.

[32] From September 2015, Res. 3099 de 2015 introduces in the formula an additional parameter based on historical productivity growth, as defined in Decreto 36 de 2015. The parameter is fixed in advance rather than based on actual minimum-wage increases.

terms; I let it grow with the minimum wage, at the average real rate of the previous ten years. Second, payments are discounted at a fixed real interest rate of 4 percent; I use the market real interest rate. Third, mortality follows regulatory tables that date from 1994 and were updated only once over the period, in 2010; I use population cohort tables. I change each assumption separately and then all three together. Figure E.1 shows the results by retirement cohort, and Appendix E.2 describes the calculations.

Minimum-wage growth alone raises the value of the pension by 17 to 25 percent for the 2008–2015 cohorts, with a median of 19 percent (about USD 10,500 at 2023 prices). This gap falls to 5 to 9 percent in 2016–2017, after the 2015 revision of the formula. The other assumptions partly offset this gain. With all three together, the pension is worth up to 19 percent more than the threshold, in 2010, and 5 to 10 percent more for the other 2009–2013 cohorts. It is worth about the same as the threshold in 2008, 2014, and 2015, and 10 to 15 percent less from 2016. The transfer therefore varies across cohorts and is not always positive.

To assess how much of the mortality effect this transfer could explain, I use the median transfer of 19 percent under the most generous scenario, which accounts only for minimum-wage growth. This is equivalent to receiving a pension about one-fifth larger for life. Appendix E.3 scales the estimated effects of lifetime pension increases on old-age mortality in Germany and Chile (Malavasi and Ye, 2024; Miglino et al., 2023) to a transfer of this size. It would add 1.2 to 1.9 months of life over the 17 observed years, which is 23 to 38 percent of the reduced-form gain and 10 to 16 percent of the 2SLS gain. Two other studies would imply none of it. In the United States, higher Social Security benefits raised mortality (Snyder and Evans, 2006), and in Sweden, lottery wealth had no effect on mortality (Cesarini et al., 2016). The comparison is an approximate benchmark for how much of the effect a transfer of this size could drive, since the studies cover different populations and different transfer settings.

These results suggest that the higher value of the pension may contribute to the mortality effect, and to the large magnitudes discussed in Section 5.4, but is unlikely to explain most of it. Beyond any transfer, the pension also provides protection against inflation, interest-rate risk, and outliving one's savings, as well as survivor benefits. The following sections examine how the payout form may affect survival through savings depletion and responses to guaranteed lifetime income.

6.2 Do Lump Sums Leave Retirees Exposed?

A second explanation is that lump-sum recipients spend their savings faster than an annuity would pay them out, leaving them with fewer resources later in retirement, when medical needs and expenses rise with age (De Nardi et al., 2010; Fahle et al., 2016). The data do not record spending or remaining savings, so I cannot test this mechanism directly. Instead, I examine whether the mortality effects are larger among retirees who depend most on their pension and draw on survey evidence from Peru on how retirees use lump sums.[33]

If rapid savings depletion contributes to mortality, the effects should be larger among retirees who rely most on their pension. At the threshold, every annuitant receives one minimum wage for life, which replaces nearly all of a low earner's wage but only a fraction of a high earner's. I test this prediction by splitting retirees into tertiles of their wage before retirement. Table 10 shows that the mortality effect is concentrated in the bottom tertile, where crossing the threshold reduces the annual hazard of death by 0.96 percentage points. This suggests that the benefits of annuitization are largest among retirees with lower previous earnings, who may be more financially constrained later in retirement.

Evidence from Peru suggests that lump-sum recipients may spend their savings quickly. In 2016, Peru allowed retirees to withdraw up to 95.5 percent of their pension savings. A survey of 1,816 retirees in Lima and Callao found that 86 percent of those retiring after the reform withdrew their entire balance (Bosch et al., 2020). By the time of the survey, one to two years after withdrawing, retirees had spent on average 29 percent of their lump sum, against the 17 percent an annuity would have paid out over the same period. Retirees with low balances had spent on average 34 percent of their lump sum, and 16 percent of them had spent all of it. The most common uses were paying debts, home improvements, and medical expenses. The evidence has limits. Withdrawals were voluntary, the sample is not representative and spending is self-reported over a short period.

These patterns are consistent with fast savings depletion as a potential mechanism. In Colombia, the mortality effect is concentrated among lower-wage retirees, who may face greater financial constraints later in retirement. Evidence from Peru, despite its limitations, suggests that lump sums can be spent relatively quickly, particularly among retirees with less savings. Depletion would also explain why the reduction in registration in the social-assistance registry appears only among cohorts retired ten or more years. Neither fact

[33]There is limited evidence on how retirees use lump-sum payouts in settings like Colombia. Future work on this project could produce such data by surveying the retirees in my sample, to examine how recipients spend or invest their savings and how these decisions affect their financial security in retirement.

singles out depletion, however: lower-wage retirees are also the ones for whom a guaranteed income, or a pension worth more than their savings, would matter most. Section 6.3 examines how guaranteed income may matter even before savings run out.

6.3 Does Guaranteed Income Matter before Savings Run Out?

A third explanation is that guaranteed lifetime income matters even while lump-sum recipients still hold savings. Two channels could produce this effect. The first is financial security. A guaranteed income removes uncertainty about whether resources will last, and financial stress can harm health. This channel may be especially relevant for a lump sum, since the retiree bears both the investment risk and the longevity risk of her savings on her own. The second is health care use and investment. A stable income makes recurring health spending easier to sustain. In addition, because payments stop at death, an annuity raises the value of each additional year of life, which strengthens the incentive to invest in health (Philipson and Becker, 1998).

Evidence from Chile is consistent with both channels. Larrain et al. (2026) find that the effect of annuities on mortality is stronger when retirees experience poor fund returns after retirement, consistent with annuities reducing financial stress. In survey data, they also find that annuitants are more likely to have a routine health examination and report less disability.

Taken together, the evidence suggests that both the availability of resources later in retirement and the security of lifetime income may matter, although the current data cannot distinguish their contributions. The next step of this project is to link the retirees in my sample to Colombian health records on health care use and spending. These data would allow me to examine whether annuitization changes how retirees use health care and help assess the pathways through which it may affect survival.

7 Conclusion

In this paper I provide causal evidence on the effects of annuitizing pension savings, relative to withdrawing them as a lump sum, on outcomes in old age. I exploit a legally mandated savings threshold in Colombia's private pension system that determines the payout form and compare retirees with savings just above and just below it. Using novel administrative records of the private pension system, linked to mortality records and to the social-assistance registry, I follow retirees for up to 17 years after retirement. I show

that annuitization improves survival and reduces poverty. It adds 1.3 years of life over the observed period and lowers registration in the social-assistance registry among survivors by 23 percentage points.

The timing of the effects, their concentration among lower earners, and the value of the annuity relative to the savings it replaces speak to potential mechanisms. The annuity bundles longevity insurance with protection against inflation and investment risk, and a pension that may be worth more than the savings it replaces. Benchmarks suggest this higher value contributes to the effect but is unlikely to explain most of it. The delayed emergence of the mortality effect, its concentration among lower-wage retirees, and the late decline in social-assistance registration are consistent with lump-sum recipients depleting their savings. The same patterns are also consistent with guaranteed income mattering before savings run out, through financial security or through sustained investment in health. To pin down the mechanisms, the next step of this project links the retirees in the sample to records on health care use.

These findings speak to the widespread shift from defined-benefit to defined-contribution pensions, which increasingly places longevity risk on retirees. As pension systems have moved toward defined-contribution arrangements, retirees have gained greater control over their accumulated wealth but also bear more of the risk of outliving it, and policies governing decumulation determine how much of this risk remains with them and how much is insured. The results show the cost of leaving retirees uninsured against this risk, in higher mortality and poverty in old age. The estimates in this paper are local to retirees near the annuitization threshold and to a setting in which the alternative is a lump sum, and therefore do not imply that annuitization is optimal for all retirees or in all settings. They do show, however, that payout design has consequences that extend beyond the allocation of consumption over retirement.

The social-assistance results suggest that the consequences of pension payout design extend beyond the pension system. Even among retirees with similar accumulated pension wealth, receiving a lump sum rather than lifetime income increases subsequent registration in the means-tested social-assistance registry. This identifies a margin through which the payout rules of one program can shift demand toward another. Although registration does not measure benefit receipt or public spending, the findings suggest that fiscal relief sought through pension reform may be partly offset by greater pressure on the safety net. Whether these spillovers also extend to the health system remains an open question. Pension payout rules therefore matter both for retirees' well-being and for the demands placed on other public programs.

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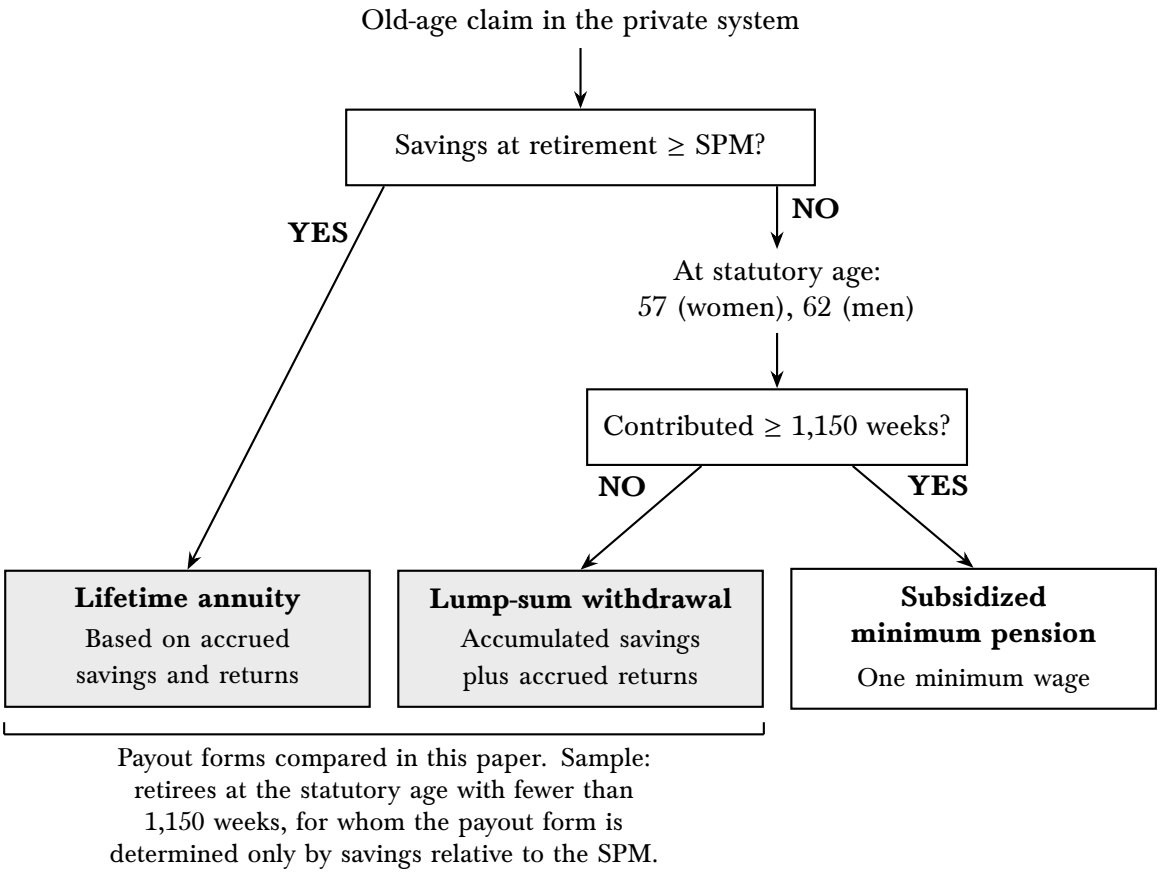
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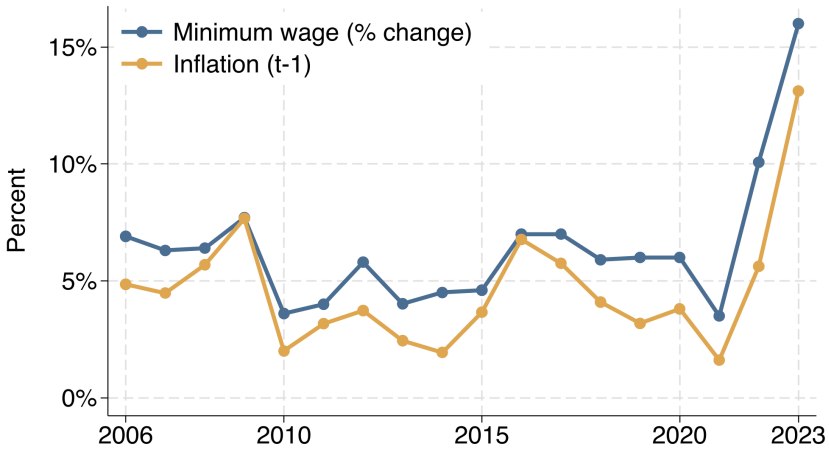
Figures

Figure 1: Pension pathways and payout forms in the private system



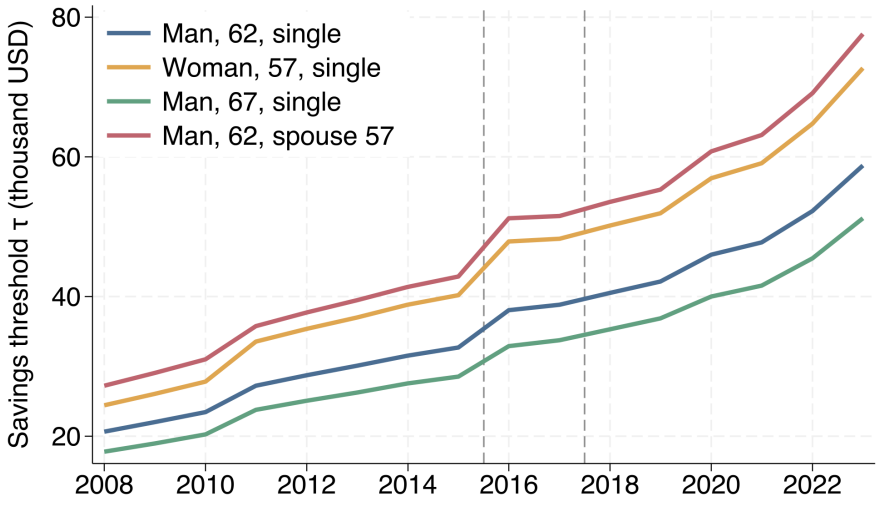
Notes: Old-age retirement pathways in the private pension system. The analysis focuses on affiliates with fewer than 1,150 contribution weeks, whose payout is an annuity if savings reach the threshold τ and a lump sum otherwise. τ is the *Saldo de Pensión Mínima* described in Section 2.3.

Figure 2: Minimum-wage growth and lagged inflation, 2006–2023



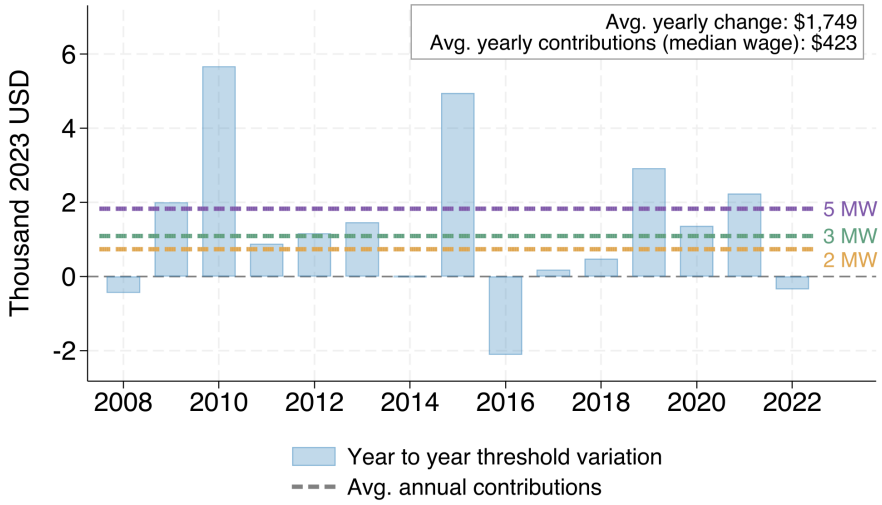
Notes: Minimum-wage growth is the annual statutory increase, and lagged inflation is the preceding year's CPI inflation, which serves as a floor for the adjustment. Their difference captures the increase above inflation and, because the savings threshold moves one for one with the minimum wage, the corresponding component of threshold growth. Sources: Banco de la República and DANE.

Figure 3: The savings threshold for representative retirees, 2008–2023



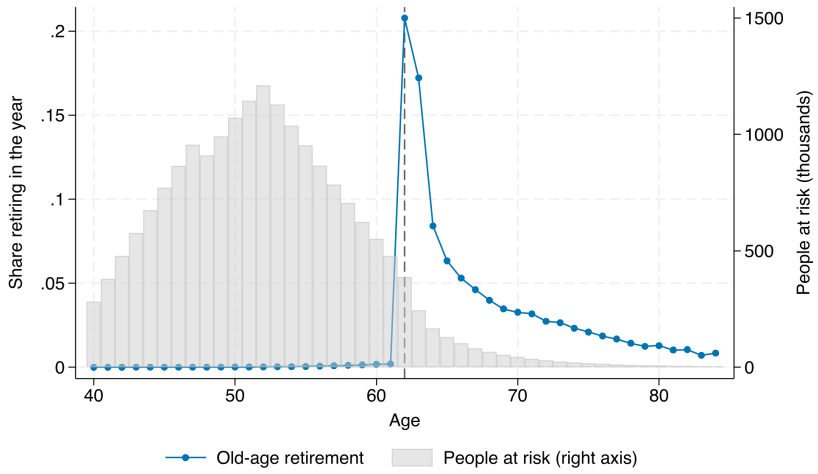
Notes: The figure plots the January statutory threshold, in thousands of 2023 USD, for four representative profiles. Dashed lines mark the 2015 and 2017 regulatory changes; after 2017, thresholds became fund-specific, so the statutory reference value is shown.

Figure 4: Year-to-year change in the threshold and one year of contributions



Notes: Bars show the change from one January to the next in the savings threshold of a single man age 62, in thousands of 2023 USD. Dashed lines show one year of mandatory contributions credited to the account (11.5 percent of earnings) at 2, 3 and 5 monthly minimum wages, averaged over 2008–2022 in 2023 USD. The box reports the average absolute annual change and one year of contributions at the median wage of retirees within 90 minimum wages of the threshold. Investment returns are excluded.

Figure 5: Retirement hazard by age, men



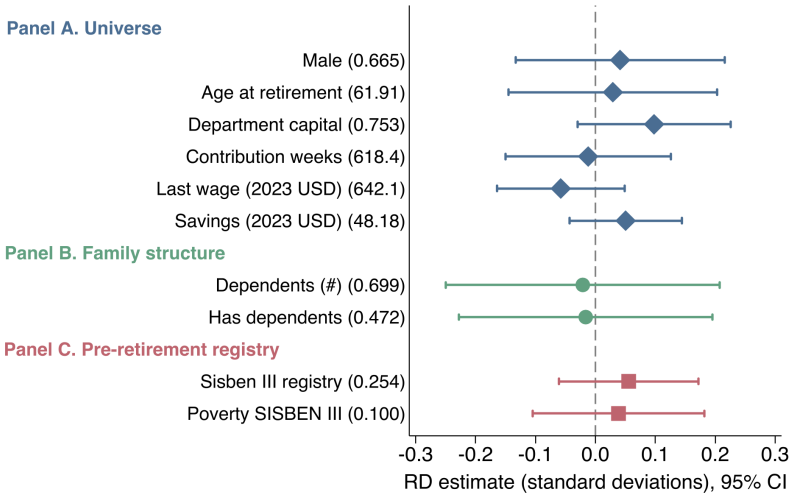
Notes: Male affiliates of the private pension system, 2008–2023. The line shows the annual retirement hazard at age a, defined as the share of affiliates age a at December 31 who retire that year with an old-age pension. Bars show the number at risk, in thousands (right axis). The dashed line marks the statutory retirement age of 62. Appendix Figure C.2 shows women.

Figure 6: Retirement timing and position relative to the threshold



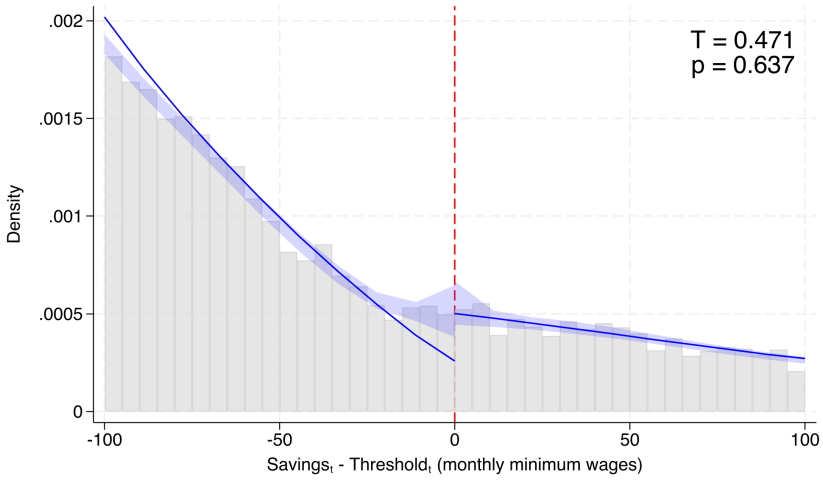
Notes: Person-year panel of the estimation sample. Affiliates contribute one observation per year from the statutory retirement age until retirement; the outcome equals one in the year they retire. The running variable is projected savings minus the threshold in force each year, in monthly minimum wages of that year. Markers are binned means and lines are linear fits within 90 minimum wages of the cutoff. $\hat{\theta}$ estimates Equation (5) with calendar-year fixed effects; the robust bias-corrected standard error, clustered by affiliate, is in parentheses.

Figure 7: Continuity of predetermined characteristics at the threshold



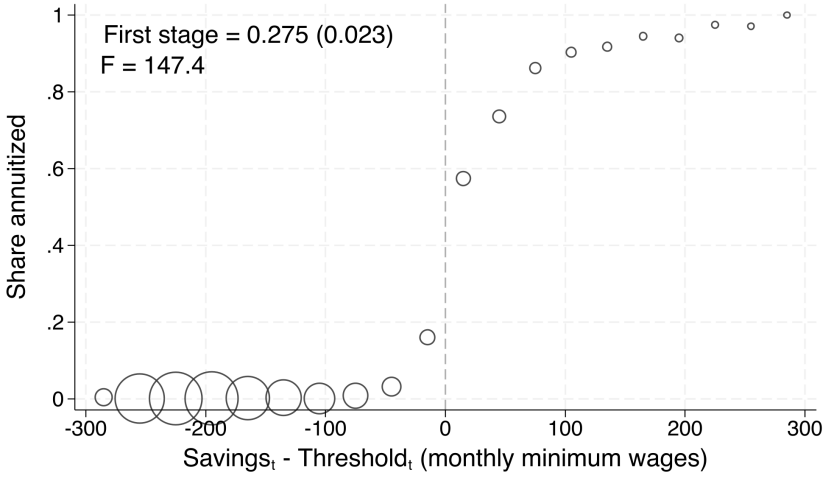
Notes: Each row reports the RD estimate in Table 2 for the indicated covariate, divided by the standard deviation of the covariate within its bandwidth, with conventional 95 percent confidence intervals. Means below the cutoff within the bandwidth are in parentheses. Estimates are local linear with a triangular kernel, retirement-year fixed effects, and covariate-specific MSE-optimal bandwidths, in monthly minimum wages. Wage is in 2023 USD and savings in thousands of 2023 USD. Panel B uses the retirees of the largest fund with a record; Panel C excludes retirees whose Sisben III record is dated in or after the retirement year.

Figure 8: Density of the running variable at the threshold



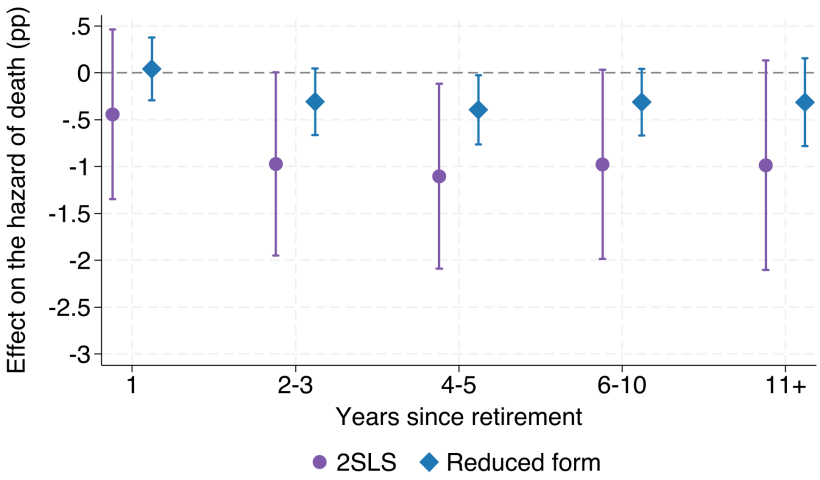
Notes: The running variable is savings at retirement minus the threshold, in monthly minimum wages of the retirement year. Bars show its density in bins of 5 minimum wages and the solid line the local polynomial density estimate of Cattaneo et al. (2020), with 95 percent confidence intervals. T and p report the robust bias-corrected test of continuity of the density at the cutoff, local quadratic within 90 minimum wages.

Figure 9: Discontinuity in annuitization at the savings threshold



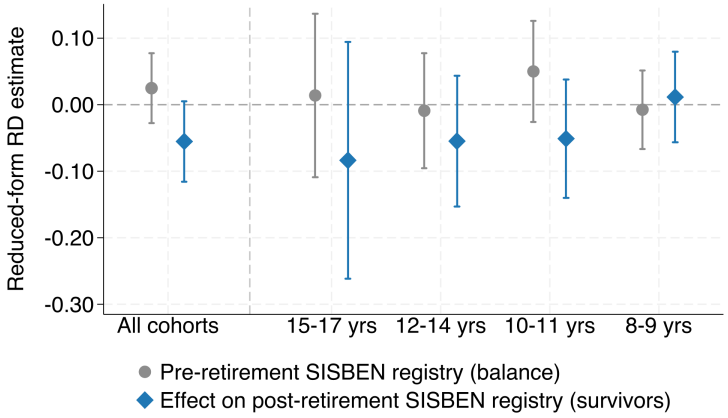
Notes: Markers show the share annuitized in bins of 30 minimum wages of the running variable, within 300 minimum wages of the cutoff, with size proportional to the number of retirees in each bin. Bins do not cross the cutoff. The running variable is savings at retirement minus the threshold, in monthly minimum wages of the retirement year. The first stage is the estimate of column (1) of Table 3, with its conventional standard error in parentheses.

Figure 10: Effect of annuitization on the annual hazard of death, by years since retirement

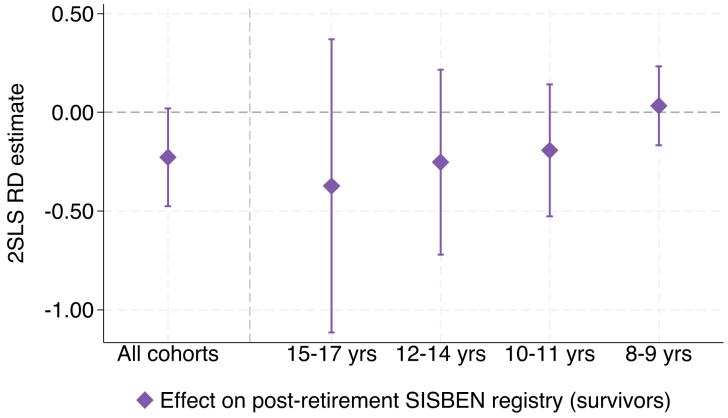


Notes: 2SLS and reduced-form estimates of Table 5, with 95 percent confidence intervals. Coefficients are in percentage points of the annual probability of death. Standard errors are clustered by retiree.

Figure 11: Sisben registration before and after retirement, by retirement cohort



Panel A. Reduced form: balance before retirement and effect after retirement



Panel B. 2SLS: effect on registration after retirement

Notes: Estimates of Table 7. The first group pools the complete sample; the other four are retirement cohorts labeled by years since retirement in June 2025. Each estimate is a separate RD with a triangular kernel within 90 minimum wages and retirement-year fixed effects. Panel A reports reduced-form estimates: black circles are the balance test on registration in Sisben III before retirement and red diamonds the effect on registration in Sisben IV among survivors. Panel B reports 2SLS estimates of the effect on registration in Sisben IV among survivors. Markers are conventional estimates with conventional 95 percent confidence intervals.

Tables

Table 1: Descriptive statistics of the analysis sample

	All (1)	Lump sum (2)	Annuity (3)
Panel A. Predetermined characteristics			
Male (%)	43.1	42.6	58.9
Age at retirement	60.5	60.5	61.7
	(3.5)	(3.5)	(3.3)
Retired at the statutory age (%)	60.5	60.9	46.5
Contribution weeks	413.0	401.9	767.7
	(308.0)	(302.2)	(281.6)
Wage in the year before retirement (2023 USD)	443.1	401.2	1244.5
	(762.2)	(677.8)	(1494.5)
Holds a pension bond (%)	41.1	39.5	92.4
Lives in a department capital (%)	66.0	65.6	81.7
Panel B. Savings and threshold (thousand 2023 USD)			
Savings at retirement	16.3	12.5	141.4
	(31.7)	(15.7)	(98.5)
Threshold, 2008–2017	61.2	61.3	59.6
	(7.9)	(7.9)	(7.8)
Above the threshold, 2008–2017 (%)	8.9	1.8	93.9
Panel C. Outcomes			
Deceased by June 2025 (%)	5.1	5.1	5.5
Age at death, among the deceased	67.3	67.2	70.8
	(5.7)	(5.7)	(6.2)
Registered in SISBEN (%)	60.3	61.8	9.8
Classified as poor (%)	30.3	31.2	1.4
Panel D. Retirement year (%)			
2008–2012	4.0	3.5	17.3
2013–2015	8.8	8.4	21.0
2016–2017	10.3	10.0	20.2
2018–2023	76.9	78.0	41.5
Panel E. Pension fund (%)			
Fund A	54.2	54.8	36.1
Fund B	30.5	30.0	45.0
Fund C	14.9	15.0	13.3
Fund D	0.3	0.2	5.6
Retirees	423,255	410,416	12,839

Notes: Universe of retirees, 2008–2023, at or above the statutory retirement age with fewer than 1,150 contribution weeks who received an annuity or lump sum. Column (1) reports the full sample; columns (2) and (3) split it by payout form. Entries are means, with standard deviations in parentheses for continuous variables, and percentages for indicators. Wages are contribution-base earnings in the year before retirement, when observed. Savings and thresholds are in thousands of 2023 USD; threshold statistics use the predicted threshold and are available for 2008–2017. Mortality and Sisben are measured in June 2025.

Table 2: Continuity of predetermined characteristics at the threshold

	Mean within h		RD	Robust	Bandwidth	Effective
Covariate	Below	Above	estimate	p-value	h (min. wages)	obs.
Panel A. Universe						
Male	0.665	0.577	0.020	0.468	49.8	2,470
	(0.012)	(0.015)	(0.043)			
Age at retirement	61.91	61.56	0.09	0.563	51.6	2,578
	(0.09)	(0.09)	(0.29)			
Department capital	0.753	0.807	0.041	0.191	91.2	5,602
	(0.007)	(0.010)	(0.027)			
Contribution weeks	618.4	668.2	-3.8	0.975	86.5	5,185
	(5.2)	(7.7)	(22.0)			
Last observed wage before retirement (2023 USD)	642.1	806.5	-69.6	0.353	93.8	3,822
	(17.2)	(52.4)	(65.6)			
Savings at retirement (thousand 2023 USD)	48.18	67.82	0.70	0.494	65.7	3,526
	(0.22)	(0.28)	(0.67)			
Panel B. Family structure						
Number of registered dependents	0.699	1.004	-0.016	0.630	69.2	1,572
	(0.024)	(0.034)	(0.092)			
Any registered dependent	0.472	0.705	-0.008	0.667	77.0	1,835
	(0.014)	(0.019)	(0.054)			
Panel C. Pre-retirement registry						
Registered in Sisben III	0.254	0.144	0.023	0.363	102.8	6,557
	(0.006)	(0.008)	(0.025)			
Poor: score below the subsidy cutoff	0.100	0.056	0.011	0.521	82.7	4,770
	(0.005)	(0.006)	(0.021)			

Notes: Each row reports an RD balance test using the specification in Equation (2), replacing the outcome with the indicated covariate. Estimates are local linear with a triangular kernel, retirement-year fixed effects, and covariate-specific MSE-optimal bandwidths, in monthly minimum wages. Means are computed within the bandwidth; conventional standard errors are in parentheses and p-values are robust bias-corrected. Wage and savings are in 2023 USD. Panel B uses the retirees of the largest fund with a record; Panel C excludes retirees whose Sisben III record is dated in or after the retirement year. *, **, and *** denote significance at the 10, 5, and 1 percent levels.

Table 3: First stage: annuitization at the savings threshold

	Annuitization (A_i)
	(1)
Above threshold (Z_i)	0.275***
	(0.023)
	[0.034]
Robust p-value	<0.001
First-stage F	147.4
Bandwidth (minimum wages)	90
Effective obs., below	7,040
Effective obs., above	2,866
Total individuals	87,906

Notes: Local linear estimate of Equation (2) with a uniform kernel within 90 minimum wages of the cutoff. Conventional standard errors in parentheses and robust bias-corrected standard errors in brackets; the p-value is robust bias-corrected. Effective observations are those within the bandwidth. *, **, and *** denote significance at the 10, 5, and 1 percent levels.

Table 4: Effect of annuitization on the annual hazard of death

	Reduced form	2SLS
	(1)	(2)
Above threshold (Z_i)	-0.30**	
	(0.15)	
Annuitization (A_i)		-0.96**
		(0.48)
Baseline hazard	0.95	1.26
First stage (α)		0.310***
		(0.026)
First-stage F		137.8
Person-years	63,356	63,356
Individuals	5,512	5,512
Deaths	508	508

Notes: Person-year panel of retirees, followed from the year after retirement until death or 2025; the outcome equals one in the year of death. Estimates of Equations (2) and (3) within 90 minimum wages of the cutoff, with a uniform kernel. Coefficients, standard errors, and baseline hazards are multiplied by 100 and expressed in percentage points. Standard errors are clustered by retiree. *, **, and *** denote significance at the 10, 5, and 1 percent levels.

Table 5: Effect of annuitization on the annual hazard of death, by years since retirement

Years since retirement	2SLS (1)	Reduced form (2)
1 year	-0.44	0.04
	(0.46)	(0.17)
2–3 years	-0.97*	-0.31*
	(0.50)	(0.18)
4–5 years	-1.10**	-0.39**
	(0.50)	(0.19)
6–10 years	-0.98*	-0.31*
	(0.51)	(0.18)
11+ years	-0.99*	-0.31
	(0.57)	(0.24)
Person-years	63,356	63,356
Deaths	508	508

Notes: Column (1) reports 2SLS estimates of Equation (4) and column (2) the reduced form, with effects specific to each group of years since retirement and a common savings slope on each side of the cutoff. Coefficients are in percentage points. Standard errors are clustered by retiree. *, **, and *** denote significance at the 10, 5, and 1 percent levels.

Table 6: Years of life gained from annuitization

		Gain from annuitization		
	Baseline	Constant effect	By bin	By bin, extrapolated
Panel A. Expected years lived over the 17 observed years				
2SLS	15.24	1.35	1.32	
Reduced form	15.67	0.42	0.40	
Panel B. Remaining life expectancy at the statutory retirement age				
Women, retiring at 57	28.33			
2SLS		2.51	2.54	4.00
Reduced form		1.07	1.06	1.47
Men, retiring at 62	20.02			
2SLS		2.25	2.24	2.76
Reduced form		0.67	0.66	0.80

Notes: Years of life implied by the hazard estimates of Tables 4 and 5, over the 17 observed years after retirement (Panel A) and in remaining life expectancy from the official pensioner mortality tables (Panel B). Effects apply over the first 17 years. The extrapolated column keeps the effect for 11+ years through age 110.

Table 7: Effect of annuitization on Sisben registration, by years since retirement

	All cohorts	2008–2010 (15–17 years)	2011–2013 (12–14 years)	2014–2015 (10–11 years)	2016–2017 (8–9 years)
	(1)	(2)	(3)	(4)	(5)
Panel A. After retirement: SISBEN IV registration among survivors					
Above threshold (Z_i), reduced form	-0.055**	-0.084	-0.055	-0.051	0.011
	(0.031)	(0.091)	(0.050)	(0.045)	(0.035)
	[0.048]	[0.472]	[0.189]	[0.241]	[0.354]
Annuitization (A_i), 2SLS	-0.227***	-0.373	-0.252*	-0.192	0.033
	(0.127)	(0.379)	(0.239)	(0.171)	(0.102)
	[0.007]	[0.213]	[0.066]	[0.135]	[0.310]
First stage (α)	0.243	0.224	0.217	0.266	0.342
	(0.032)	(0.090)	(0.053)	(0.043)	(0.034)
First-stage F	58.8	6.3	16.9	38.1	98.7
Effective obs., below / above	3,466 / 1,538	374 / 217	1,209 / 534	1,883 / 787	2,900 / 1,183
Panel B. Before retirement: balance in Sisben III registration					
Above threshold (Z_i), reduced form	0.025	0.014	-0.009	0.050	-0.008
	(0.027)	(0.063)	(0.044)	(0.039)	(0.030)
	[0.638]	[0.538]	[0.844]	[0.475]	[0.292]
Effective obs., below / above	3,809 / 1,590	382 / 207	1,342 / 557	2,085 / 826	3,155 / 1,239
Bandwidth (minimum wages)	90	90	90	90	90

Notes: Each column is a separate RD, local linear with a triangular kernel within 90 minimum wages and retirement-year fixed effects. Column (1) pools all cohorts in the sample. Panel A is registration in Sisben IV in June 2025 among survivors. Panel B is a balance test on registration in Sisben III before retirement. Conventional standard errors in parentheses and robust bias-corrected p-values in brackets. *, **, and *** denote significance at the 10, 5, and 1 percent levels of the robust p-value.

Table 8: Lee bounds for survivor selection, Sisben registration after retirement

Cohorts	Alive in 2025			Reduced form		2SLS	
	Below	Above	Trimming	Survivors	Lee bounds	Survivors	Lee bounds
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
All cohorts	0.906	0.921	0.016	-0.055	[-0.067, -0.051]	-0.227	[-0.277, -0.211]
					(-0.133, 0.010)		(-0.556, 0.047)
2008–2010 (15–17 years)	0.839	0.844	0.006	-0.084	[-0.088, -0.082]	-0.373	[-0.394, -0.365]
					(-0.290, 0.113)		(-1.723, 0.903)
2011–2013 (12–14 years)	0.902	0.963	0.064	-0.055	[-0.109, -0.041]	-0.252	[-0.503, -0.190]
					(-0.205, 0.041)		(-1.166, 0.313)
2014–2015 (10–11 years)	0.928	0.913	0.000	-0.051	[-0.051, -0.051]	-0.192	[-0.192, -0.192]
					(-0.136, 0.032)		(-0.521, 0.128)
2016–2017 (8–9 years)	0.928	0.942	0.015	0.011	[0.001, 0.016]	0.033	[0.002, 0.046]
					(-0.065, 0.077)		(-0.189, 0.227)

Notes: Cohorts as in Table 7. Columns (1)–(2) are the shares alive in June 2025 below and above the threshold, and column (3) is the trimming share. Columns (5) and (7) report Lee (2009) bounds in square brackets.

Table 9: OLS and RD estimates of the effect of annuitization

	OLS All retirees (1)	Reduced form $\|R_i\| < 90$ (2)	2SLS $\|R_i\| < 90$ (3)
Panel A. Annual hazard of death (pp)			
Annuitization (A_i)	-0.72***		-0.96**
	(0.03)		(0.48)
Above threshold (Z_i)		-0.30**	
		(0.15)	
First stage			0.310***
			(0.026)
Baseline hazard	0.81	0.95	1.26
Percent of baseline	-88.9	-31.2	-75.8
Person-years	515,095	63,356	63,356
Individuals	45,553	5,512	5,512
Panel B. SISBEN registration among survivors			
Annuitization (A_i)	-0.456***		-0.227*
	(0.006)		(0.128)
Above threshold (Z_i)		-0.055*	
		(0.031)	
First stage			0.243***
			(0.032)
Baseline registration	0.576	0.344	0.369
Percent of baseline	-79.1	-16.1	-61.7
Individuals	41,596	5,004	5,004

Notes: Column (1) is OLS on all retirees. Columns (2) and (3) are the reduced form and 2SLS within 90 minimum wages of the cutoff. Controls, kernels and standard errors are those of the main results. Baselines are lump-sum retirees in column (1), retirees below the cutoff in column (2), and the untreated mean at the cutoff in column (3). *, **, and *** denote significance at the 10, 5, and 1 percent levels.

Table 10: Effect of annuitization on the annual hazard of death by pre-retirement wage

	All	Wage tertiles		
	retirees	Bottom	Middle	Top
	(1)	(2)	(3)	(4)
Reduced form (ρ)	-0.30**	-0.96***	0.37	-0.04
	(0.15)	(0.33)	(0.33)	(0.27)
Baseline hazard below the cutoff	0.95	1.00	0.86	0.71
First stage (α)	0.310***	0.306***	0.324***	0.275***
	(0.026)	(0.058)	(0.056)	(0.055)
First-stage F	137.8	28.3	34.0	24.7
2SLS (β)	-0.96**	-3.15***	1.13	-0.13
	(0.48)	(1.20)	(1.06)	(0.98)
Complier baseline hazard	1.26			
Equal reduced forms across tertiles, p			0.023	
Individuals	5,512	1,232	1,170	1,199
Deaths	508	114	107	79

Notes: Person-year panel and specification of Table 4, by tertile of the last contribution wage before retirement. Tertiles cover retirees with a recorded wage. Coefficients and baseline hazards are in percentage points. The equality test comes from a pooled regression with tertile-specific intercepts, jumps, and slopes. Standard errors clustered by retiree in parentheses. *, **, and *** denote significance at the 10, 5, and 1 percent levels.

A Institutional Details

A.1 Pension Pathways Outside the Analysis Sample

Old age. This is the pension route studied in this paper. It covers affiliates who reach retirement without qualifying for a disability pension. Figure 1 summarizes the eligibility criteria and payout rules. There are three mutually exclusive payout forms in this pathway: annuity, subsidized pension, and lump sum.

First, an affiliate qualifies for an old-age pension if her accumulated savings, including any pension bond, are sufficient to finance a lifetime annuity of at least one monthly minimum wage. The expected present value of this pension defines the savings threshold (SPM). Affiliates who meet this criterion may retire at any age. Their savings are then used to finance a pension, and they can choose among the following annuity options.[34]

- *Renta vitalicia (life annuity).* The fund transfers her balance to a life insurer she chooses, which pays a monthly pension for life. Payments are indexed to inflation, or to the minimum wage when the pension equals it. The contract is irrevocable and not inheritable. At her death, the insurer pays only survivors' pensions to eligible beneficiaries.

- *Retiro programado (phased withdrawal).* The savings balance is kept in the retiree's account and managed by the pension fund. The pension the retiree receives is withdrawn from that balance and recomputed each year. The recomputation takes into account the remaining balance, including returns, and the life expectancy of the retiree and her beneficiaries. The balance has a floor. If it falls to the savings amount needed to buy a life annuity of one minimum wage, the fund must buy that annuity on her behalf. If she dies without eligible beneficiaries, the remaining balance passes to her estate.

- *Other modalities.* The law also allows a phased withdrawal followed by a deferred life annuity, and other combinations authorized by the financial regulator, the Superintendencia Financiera de Colombia. These are rarely used. Between 1998 and 2017, *renta vitalicia* and *retiro programado* together accounted for more than 95 percent of old-age pensions in the private system in every year. *Retiro programado* alone accounted for close to 90 percent by 2017.[35]

[34] Ley 100 de 1993, art. 64, governs early retirement. Arts. 79–82 define the payout modalities, and art. 14 sets their indexation. Decreto 832 de 1996, art. 12, sets the floor on the *retiro programado* balance.

[35] Superintendencia Financiera data reported in Villar et al. (2017).

Second, affiliates who reach the statutory retirement age (57 for women and 62 for men) without sufficient savings (their accumulated pension savings are insufficient to finance an annuity paying one monthly minimum wage) face two alternatives. Those with at least 1,150 weeks of contributions qualify for the minimum-pension subsidy. The government provides the additional funds needed to finance a pension of one monthly minimum wage, effectively providing a subsidized annuity. Those with fewer than 1,150 weeks of contributions may continue contributing or receive a *devolución de saldos*, a lump-sum payment equal to their accumulated savings, including investment returns and any pension bond.[36]

Disability. This pathway covers affiliates who lose at least 50 percent of their capacity to work from a cause unrelated to work.[37] Under statutory rules, the insurer covering the pension fund's disability risk assesses the degree of disability using a national rating manual. The worker must also have contributed for at least 50 weeks in the three years before the disability began. Otherwise she can withdraw her accumulated savings as a lump sum.

The value of the disability pension corresponds to a share of the worker's average past earnings, rising with the degree of disability and weeks contributed to the system. The share ranges from 45 to 75 percent, and the pension cannot fall below one minimum wage. The pension is financed by her accumulated savings. If these are insufficient, the fund's disability insurer covers the remaining balance. The pension is paid for life under the same payment options as an old-age pension, but eligibility is reviewed every three years and payments stop if the disability ends. Upon death, eligible beneficiaries receive a survivors' pension.[38]

Survivors' benefits. Eligible dependents receive a survivors' pension in two cases. The first is the death of an affiliate who has not yet retired. Her dependents then receive a new pension if she contributed at least 50 weeks in the three years before her death. As in the disability pathway, its amount is a share of her average past earnings, between 45 and 75 percent. It is financed by her account, any pension bond, and a top-up from the fund's disability and survivors' insurer. If she did not meet the contribution requirement, her dependents receive her balance as a lump sum. The second case is the death of a pensioner, known as *sustitución pensional.* Her dependents then receive the same pension

[36] Ley 100 de 1993, arts. 65–66.

[37] Disability from an occupational accident or disease is covered by a separate occupational risk insurance system (*Sistema General de Riesgos Laborales*), which pays its own disability pensions (Ley 776 de 2002).

[38] Ley 100 de 1993, arts. 38–41, 44, 69–72 and 79.

she was receiving. In both cases, the pension cannot fall below one minimum wage. Spouses aged 30 or more are eligible dependents who receive the pension for life. Children until the age of 18 (or 25 if they are students) also qualify as eligible dependents. In cases without eligible dependents, the remaining balance passes to the worker's estate.[39]

Family pensions. Under *pensión familiar* a married or cohabiting couple may combine their pension savings and/or contribution weeks to receive a single old-age pension. In this case, the pension payout options are those in the old-age pathway.[40]

Sample of interest. The main analysis of this paper focuses on the old-age pension pathways. In particular, the sample of interest is restricted to retirees with fewer than 1,150 contribution weeks who retired at or above the statutory retirement age. For this group, accumulated savings at retirement determine the payout form. Those whose savings exceed the threshold (SPM) receive an annuity, and the rest receive a lump sum. The sample excludes disability and survivors' pensions, whose eligibility rules and payout forms differ. It also excludes family pensions, which pool savings within couples, and pensions granted by court order rather than through the standard administrative process.

A.2 The Savings Threshold: Original Formula and Reforms

The savings threshold was introduced by *Ley 100 de 1993*, which created the private pension system. The law requires affiliates who reach the statutory retirement age to have enough savings to finance a lifetime pension of at least one minimum wage, but does not specify how to calculate the required savings. *Decreto 832 de 1996* assigned this calculation to the Ministry of Finance and defined the threshold as the *Saldo de Pensión Mínima* (SPM). The Ministry established the first formula in *Resolución 1875 de 1997* and replaced it in 2015 with *Resolución 3099*, which introduced time-varying financial parameters. In 2017, *Resolución 3023* retained the formula but made it a reference value rather than a binding threshold. This section describes the original formula, the changes introduced in 2015, and the implications of the 2017 reform.

[39] Ley 100 de 1993, arts. 46–48 and 73–78. Spouses under 30 years receive a pension for at most 20 years.
[40] Ley 1580 de 2012, arts. 1–2, which added arts. 151A–151B to Ley 100 de 1993. Family pensions were 0.3 percent of old-age pensions in the private system in 2023.

A.2.1 The Original Formula: Resolución 1875 de 1997

The original formula values a lifetime pension of one statutory monthly minimum wage. The pension makes fourteen monthly payments a year, twelve regular and two additional, and continues to one dependent after the affiliate's death. The formula adds a funeral benefit of five minimum wages paid at the affiliate's death and loads the total by five percent for insurer expenses. Resolución 1875, issued by the Ministry of Finance on 15 September 1997, sets the SPM threshold at

$$\text{SPM} \;=\; 1.05 \cdot SM \cdot \Big[\big((12 f_{12} + 2 f_2)\, a \;+\; 6 \;+\; 5A\big)\,(1 + K^*)\,U^{\,n} \;+\; C\Big]. \tag{A.1}$$

Each term in Equation (A.1) is defined in the following table, together with its source:

	Definition	Source
SM	Monthly minimum wage at the calculation date.	National government, end of year
K	Consumer-price inflation in the calendar year before the calculation date.	National statistics office (DANE)
v	Annual discount factor at a real interest rate of 4 percent, $v = 1/1.04$.	Fixed
U	Monthly nominal discount factor, $U = [\, v/(1+K)\,]^{1/12}$.	Computed from v and K
a	Expected present value, at the 4 percent real rate, of one peso a year paid while the affiliate or her dependent is alive, $a = a_x^{(\alpha)} + a_y^{(\beta)} - a_{xy}^{(\alpha\beta)}$. Here x and y are the ages of the affiliate and the dependent, and α and β their mortality types. Without a dependent, $a = a_x^{(\alpha)}$.	Mortality tables of the financial regulator (SFC)
A	Expected present value of one peso paid at the affiliate's death. The term $5A$ values a funeral benefit of five minimum wages, written F_i in Equation (1).	Mortality tables of the financial regulator
f_{12}, f_2	Adjustments for the timing within the year of the twelve regular and the two additional payments, $f_{12} = \big(1 + \frac{11}{24}K\big)/(1+K)$ and $f_2 = \big(1 + \frac{1}{4}K\big)/(1+K)$.	Computed using K
m, n, K^*, C	Terms for the first, partial year. m is the month of the first payment and $n = 13 - m$ the months left in that year. C is the present value of the payments made in that year, and K^* indexes the minimum wage to the next year. n, K^*, and C are zero when the first payment falls in January.	Month of the first payment

Mortality tables enter through a and A. The tables were updated in 2010. The regulator publishes separate tables for non-disabled men and women ($T = 1, 2$) and for disabled men and women ($T = 3, 4$). Sons and daughters use the non-disabled tables, cut at age 25 ($T = 5, 6$). Ages are measured on 1 January after the calculation date, adding one year when the birthday falls before 1 July. The formula covers a single dependent, chosen in this order: spouse, disabled children (youngest to oldest), non-disabled children (youngest to oldest) and parents.

A.2.2 The 2015 Revision: Resolución 3099 de 2015

Resolución 3099, issued by the Ministry of Finance on August 19, 2015, replaced the original savings threshold formula. While the original formula used fixed financial parameters, the new formula updates them over time to reflect market conditions. The real interest rate is set every six months by the Ministry of Finance, and inflation is measured as a three-year average. The formula also introduces two new parameters: productivity growth, which accounts for expected increases in minimum-wage pensions above inflation, and a safety factor that accounts for changes in beneficiaries. It also allows for up to two beneficiaries instead of one. In real terms, the threshold depends on the difference between the real interest rate and expected productivity growth. Under this revision, the overall value of the SPM is

$$\text{SPM} \;=\; \frac{VA + AF}{1 - g}, \tag{A.2}$$

where VA is the expected present value of the pension and AF that of a funeral benefit of five minimum wages paid at the affiliate's death. For a calculation on 1 January,

$$VA \;=\; SM \left[\sum_{k=1}^{\omega(\Omega)} \sum_{t=1}^{12} \frac{M_t \, (1+f)^{k-1} (1+\Delta)^{k-1}}{(1+i)^{\,k-1+t/12}} \; {}_{k-1+t/12}p_{\Omega} \right] (1+\mu), \qquad 1 + i = (1+f)(1+r). \tag{A.3}$$

The sum runs over each month t of each year k after the calculation. Each payment grows with inflation f and productivity growth Δ, is discounted at the nominal rate i, and is weighted by the probability that the affiliate or a beneficiary is alive to receive it. The parameters in the formula are defined as:

	Definition	Source
SM	Monthly minimum wage at the calculation date.	National government, each December
M_t	Number of payments in month t, two in June and December and one otherwise, so fourteen a year.	Fixed in the resolution
Ω	Group formed by the affiliate, of age x, and up to two beneficiaries, of ages y and z. The resolution prices the beneficiaries that give the highest value.	Affiliate's records
${}_{k-1+t/12}p_{\Omega}$	Probability that at least one member of Ω is alive $k-1$ years and t months after the calculation. $\omega(\Omega)$ is the number of years until the last member reaches the end of the mortality table.	Mortality tables of the financial regulator
r	Real interest rate, 3.81 percent through December 2015.	Ministry of Finance, every six months
f	Inflation, a weighted average of the last three years, $f = (3f_{A-1} + 2f_{A-2} + f_{A-3})/6$.	Computed from DANE's consumer-price index, each year
Δ	Productivity growth, the expected real increase in the minimum wage (Decreto 36 de 2015).	National government, each year
μ	Safety factor for changes in the beneficiaries, 0.6 percent.	National government, reviewable every six months
g	Expense load, 5 percent.	National government, reviewable each year

A.2.3 The 2018 Change in Legal Status: Resolución 3023 de 2017

From 2018 onward, each pension fund sets its own annuitization threshold. Resolución 3023 de 2017 requires each fund to compute the savings needed to finance a *retiro programado* of one minimum wage (Appendix A.1), which I call the RP threshold. The SPM formula does not change, but it becomes only a reference value. The resolution then sets what happens to an affiliate at the statutory retirement age according to where her savings fall relative to the two thresholds.

- *Savings at or above the RP threshold.* The affiliate receives a pension and chooses

among all annuity modalities.

- *Savings between the SPM and the RP threshold.* The fund asks insurers for quotes on a life annuity and the other modalities. If an insurer makes an offer, the affiliate receives a pension under it. If no insurer does, she is treated as if her savings were below the SPM (and if contribution weeks are fewer than 1,150 the lump-sum payout may take place).

- *Savings below the SPM.* If weeks of contributions are below 1,150, then the lump-sum payout takes place.

The RP threshold remains tightly regulated. Each fund computes it with a technical note that it files with the regulator for approval. The note must use, at a minimum, the parameters listed in art. 4 of Resolución 3099. These are the mortality tables, the real interest rate, and the inflation of the SPM formula, the growth of the minimum wage above inflation over the previous ten years, and assumptions on the beneficiaries and on the fund's administration fee. Funds have no discretion over individual cases and cannot use health information. The notes are proprietary, however, so I cannot compute the RP threshold from published parameters. My estimation sample therefore ends in 2017, the last year in which the SPM formula determined annuitization.

B Construction of the Running Variable

B.1 Savings Threshold Computation (SPM)

I compute the statutory threshold for each retiree between 2008 and 2017, using the corresponding formula. Through August 2015, I use the original formula (Equation (A.1)); from September 2015 through December 2017, I use the revised formula (Equation (A.2)).

For retirements between 2008 and 2015, each year the formula updates two inputs, the minimum wage and the previous year's inflation. Inflation only adjusts for the timing of payments within the year, while the real discount rate is fixed at 4 percent. Thus, for a given year and month of retirement, the threshold depends only on the affiliate's sex, age, and dependents. Table B.1 reports the parameters I use for the threshold computation of retirements in this period.

Table B.1: Parameters of the original formula, 2008–2015

Year	Minimum wage SM (COP)	Inflation K	Mortality tables
2008	461,500	5.69	RV89
2009	496,900	7.68	RV89
2010	515,000	2.01	RV89
2011	535,600	3.16	RV08
2012	566,700	3.73	RV08
2013	589,500	2.44	RV08
2014	616,000	1.94	RV08
2015 (Jan–Aug)	644,350	3.66	RV08

For retirements between 2016 and 2017, I apply the revised formula with the parameters that the financial regulator reports for each semester (II-2015 through II-2017 in Table B.2). The real rate r changes every semester, so the threshold shifts within a calendar year for otherwise identical affiliates. The expense load and the beneficiary safety factor are fixed at 5.00 and 0.60 percent.

Table B.2: Published parameters of the 2015 formula, 2015–2023

Semester	Minimum wage SM (COP)	Real interest rate r	Inflation f	Productivity growth Δ
II-2015	644,350	3.81	2.88	0.92
I-2016	689,455	3.87	4.93	0.71
II-2016	689,455	3.99	4.93	0.71
I-2017	737,717	3.86	5.74	0.27
II-2017	737,717	3.66	5.74	0.27
I-2018	781,242	3.77	5.09	0.00
II-2018	781,242	3.89	5.09	0.00
I-2019	828,116	4.02	3.91	0.00
II-2019	828,116	4.06	3.91	0.00
I-2020	877,803	4.03	3.64	0.28
II-2020	877,803	4.04	3.64	0.28
I-2021	908,526	4.01	2.60	0.24
II-2021	908,526	4.03	2.60	0.24
I-2022	1,000,000	4.01	3.98	0.26
II-2022	1,000,000	4.04	3.98	0.26
I-2023	1,160,000	4.12	8.70	0.31
II-2023	1,160,000	4.16	8.70	0.31

Notes: Rates in percent. Source: Superintendencia Financiera de Colombia (SFC).

I simplify the family structure in two ways. First, I price at most one dependent, although the 2015 formula allows two. Second, the dependent is always a spouse of the opposite sex whose age is within 20 years of the affiliate's. A retiree whose only dependents are children or parents, or whose spouse falls outside this range, is priced as single. The records report every other input. Dependents are recorded only by the largest fund, and Appendix B.2 describes how I impute them for the other funds.

B.2 Imputation of Family Structure

Table B.3 shows the number of registered dependents of retirees in the largest pension fund. For each affiliate, the records report the number of dependents registered and the relationship to the affiliate, their birthdate, sex, and vital status.

Table B.3: Registered dependents of the largest pension fund

	Men	Women	All
	(1)	(2)	(3)
Number of registered dependents	0.50	0.36	0.42
	(0.67)	(0.60)	(0.64)
Reports at least one dependent	0.40	0.29	0.34
	(0.49)	(0.46)	(0.47)
Reports a spouse	0.37	0.23	0.29
	(0.48)	(0.42)	(0.46)
Reports a child	0.11	0.11	0.11
	(0.31)	(0.31)	(0.31)
Reports a parent	0.00	0.00	0.00
	(0.03)	(0.05)	(0.04)
Age difference spouse (years)	6.04	-2.43	2.34
	(7.52)	(6.72)	(8.32)
Retirees	102,991	126,113	229,104

Notes: Family structure of retirees in the largest pension fund, 2008–2023. Means, with standard deviations in parentheses.

Imputation. For each cell of retirement year × sex × age of the affiliate, I impute the median number of dependents observed among retirees of the largest pension fund. When a retiree is imputed one dependent, the dependent's age is the median age of the spouses registered in her cell. Table B.4 shows the resulting family structure. Columns (1)–(3) report the cell medians of the observed family structure, and Columns (4)–(6) report the family structure imputed to the other pension funds.

Table B.4: Family structure imputation

	Largest fund			Other retirees		
	Men	Women	All	Men	Women	All
	(1)	(2)	(3)	(4)	(5)	(6)
Imputed spouse	0.29	0.07	0.17	0.27	0.05	0.14
	(0.45)	(0.26)	(0.37)	(0.45)	(0.23)	(0.35)
Imputed single	0.71	0.93	0.83	0.73	0.95	0.86
	(0.45)	(0.26)	(0.37)	(0.45)	(0.23)	(0.35)
Age difference spouse (years)	5.50	-2.83	3.52	5.65	-2.57	3.81
	(1.44)	(1.38)	(3.82)	(1.34)	(0.95)	(3.65)
Retirees	22,055	28,165	50,220	18,694	27,510	46,204

Notes: Columns (1)–(3) report, for retirees of the largest fund, the family structure that the cell-median rule would assign them, computed from their own cells. Columns (4)–(6) report the family structure imputed to retirees of the other funds. Medians, with standard deviations in parentheses. The age difference is the affiliate's age minus the imputed spouse's age.

B.3 Savings at Retirement

Accrued savings. The pension funds' records report each affiliate's account balance once a year, on December 31. Between that date and retirement, the balance changes for two reasons. Affiliates keep contributing to their accounts, and market returns move the value of the balance. To account for both, I assume that affiliates keep contributing at the same rate until retirement, and that the balance earns the returns of the pension funds published by the Superintendencia Financiera de Colombia.[41] The balance therefore moves with the reported annual returns.

Pension bond. Affiliates who contributed to the public pension system before the private system was introduced receive a pension bond (*bono pensional*) recognizing those contributions. The bond is valued when they enter the private system and grows annually with inflation plus 4 percent if they joined before 1999, or 3 percent thereafter (Decreto 1748 de 1995). I capitalize the bond to retirement and add it to savings, unless it is already included in the reported account balance. Pension bonds are particularly common near the threshold: Table B.5 shows that 41 percent of all retirees hold one, compared with 90 percent in the estimation sample. Among bondholders in the estimation sample, the median bond accounts for 78 percent of retirement savings.

[41] The Superintendencia Financiera de Colombia publishes the returns of the mandatory pension funds over different periods at https://www.superfinanciera.gov.co/powerbi/reportes/523/484/.

Table B.5: Timing of retirement and pension bonds

	Full sample (1)	Estimation sample (2)
Panel A. Timing of retirement		
Retires in December (percent)	8.9	8.7
Months from last observed balance to retirement	6.4	6.7
Panel B. Pension bond		
Holds a pension bond (percent)	41.1	51.8
Bond as percent of savings at retirement	81.9	84.1
Bond dated before 1999 (percent of holders)	66.8	67.1
Retirees	423,516	54,313

Notes: Column (1) shows the characteristics of retirees in the universe of retirees, 2008–2023. Column (2) shows the same characteristics for retirees in the estimation sample.

C Measurement Error and Identification

C.1 Simulation of the Measurement Error

This section describes a simulation exercise that asks whether the measurement errors that can be quantified in the data are large enough to attenuate the first stage to the size reported in Table 3.

I take the retirees of the estimation sample and treat each retiree's measured distance as her true distance, R_i^*. The payout form follows the sharp rule, $A_i = \mathbb{1}\{R_i^* \geq 0\}$, so the first stage equals one by construction. I then add an error to the distance in the following data generating process,

$$
\begin{aligned}
R_i &= R_i^* + \frac{T_i}{100}\left(u_i^T + u_i^S\right), \\
u_i^T &= W_i\,\delta_i, \qquad W_i \sim \text{Bernoulli}(p_s), \qquad \delta_i \sim F_\delta(\cdot \mid s), \\
u_i^S &\sim F_S(\cdot \mid b),
\end{aligned}
$$

where T_i is the threshold, and u_i^T and u_i^S are the errors in the threshold and in savings, in percent of the threshold. With probability p_s, which depends on sex s, the imputed family structure is wrong and the threshold is off by a gap δ_i. Otherwise the threshold is exact. The savings error depends on whether the retiree holds a pension bond, b. I estimate the first stage on R_i with the estimator and bandwidth of the main analysis and average it over 100 resamples.

I take p_s, F_δ, and F_S from the data.

- *Threshold.* For retirees of the largest fund, I price the threshold on both the imputed and the recorded family structure. The share of retirees near the cutoff whose two thresholds differ gives p_s, and the differences, in percent of the threshold, give F_δ, both by sex.

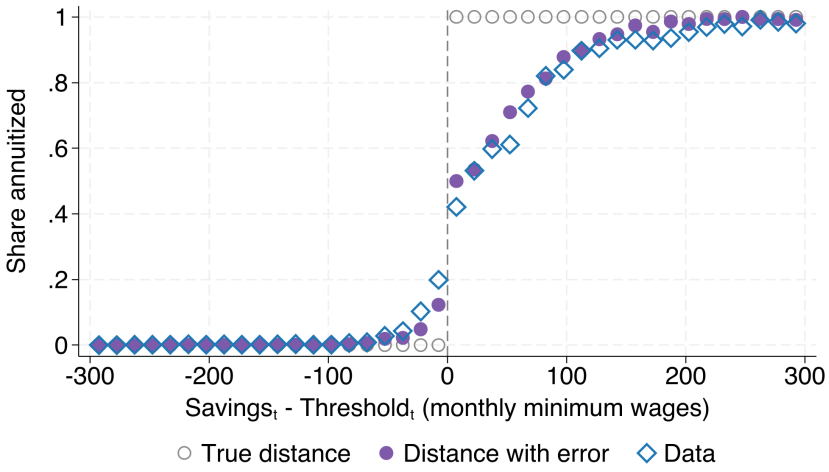
- *Savings.* For lump-sum recipients I observe the lump-sum transfer and compare the projected balance with it. The differences, in percent of the threshold, give F_S, by bond holding.

Table C.1 reports the results. The threshold error alone lowers the first stage from one to 0.42, and the savings error alone to 0.74. Together they lower it to 0.32, which accounts for 87 percent of the gap. Figure C.1 compares the share annuitized in the data with the share produced by the simulation with both sources of error.

Table C.1: First stage under the measured errors

		Simulated error			
	Data	None	Threshold	Savings	Both
First stage	0.21	1.00	0.42	0.74	0.32
Gap explained (%)		0	74	32	87

Figure C.1: Share annuitized in one simulated sample



Notes: Bins of 15 monthly minimum wages. Gray circles show the share annuitized by bin of the true distance in the simulated sample, where the payout form follows the sharp rule. Purple dots show the same retirees binned by the distance with both measured errors. Blue diamonds show the data.

C.2 Formal Result

Let R_i^* be the true distance to the threshold, $R_i = R_i^* + e_i$ the measured distance, and $A_i = \mathbb{1}\{R_i^* \geq 0\}$ the payout form. Write the outcome as

$$Y_i = \beta_0 + \beta_i A_i + u_i,$$

where (Y_i, A_i, R_i) is observed and R_i^* is not. The representation is without loss of generality. For any pair of potential outcomes $(Y_i(0), Y_i(1))$, define $Y_i(0) = \beta_0 + u_i$ and $\beta_i = Y_i(1) - Y_i(0)$, so that $Y_i = Y_i(0)(1 - A_i) + Y_i(1)A_i = \beta_0 + \beta_i A_i + u_i$. The parameter of interest is the average effect for retirees at the measured cutoff, $\mathbb{E}[\beta_i \mid R_i = 0]$. This is not the usual regression discontinuity estimand, because the measured cutoff is not the point at which the fund assigns the payout form.

Lemma. Suppose that

(a) $\mathbb{E}[u_i \mid R_i = r]$ and $\mathbb{E}[\beta_i \mid R_i = r]$ are continuous near $r = 0$;

(b) $\beta_i \perp e_i \mid R_i = r$ for r near 0;

(c) $\lim_{r\downarrow 0} \mathbb{E}[A_i \mid R_i = r] \neq \lim_{r\uparrow 0} \mathbb{E}[A_i \mid R_i = r]$.

Then

$$\mathbb{E}[\beta_i \mid R_i = 0] = \frac{\lim_{r\downarrow 0} \mathbb{E}[Y_i \mid R_i = r] - \lim_{r\uparrow 0} \mathbb{E}[Y_i \mid R_i = r]}{\lim_{r\downarrow 0} \mathbb{E}[A_i \mid R_i = r] - \lim_{r\uparrow 0} \mathbb{E}[A_i \mid R_i = r]}. \tag{C.1}$$

Condition (c) holds when e_i is independent of R_i^*, $e_i = 0$ with probability $\pi \in (0, 1)$ and e_i follows a continuous distribution G with probability $1 - \pi$, and R_i^* has a bounded continuous density f with $f(0) > 0$. In that case the jump in (c) equals $\pi f(0)/f_R(0)$, where f_R is the density of R_i.

Given the definitions of the potential outcomes, condition (a) is the usual continuity of $\mathbb{E}[Y_i(0) \mid R_i = r]$ and $\mathbb{E}[Y_i(1) \mid R_i = r]$ at the measured cutoff, the first condition of Section 4.1. The mixture structure that delivers (c) corresponds to the second and third conditions, a positive share of correctly measured retirees and an error that varies smoothly through the threshold. Condition (b) is the additional assumption that the effect does not differ between correctly measured and mismeasured retirees. Because e_i determines whether a retiree is measured correctly, (b) implies $\mathbb{E}[\beta_i \mid R_i = 0, e_i = 0] = \mathbb{E}[\beta_i \mid R_i = 0]$.

Proof. Since $A_i = \mathbb{1}\{R_i^* \geq 0\} = \mathbb{1}\{e_i \leq R_i\}$, the first stage on either side of the cutoff is

$$\lim_{r\downarrow 0} \mathbb{E}[A_i \mid R_i = r] = \lim_{r\downarrow 0} \mathbb{P}[e_i \leq r \mid R_i = r], \qquad \lim_{r\uparrow 0} \mathbb{E}[A_i \mid R_i = r] = \lim_{r\uparrow 0} \mathbb{P}[e_i \leq r \mid R_i = r].$$

For the outcome,

$$\begin{aligned}
\lim_{r\downarrow 0} \mathbb{E}[Y_i \mid R_i = r] &= \lim_{r\downarrow 0} \mathbb{E}[\beta_0 + \beta_i A_i + u_i \mid R_i = r] \\
&= \beta_0 + \lim_{r\downarrow 0} \mathbb{E}[\beta_i \mathbb{1}\{e_i \leq r\} \mid R_i = r] + \lim_{r\downarrow 0} \mathbb{E}[u_i \mid R_i = r] \\
&= \beta_0 + \lim_{r\downarrow 0} \mathbb{E}[\beta_i \mid R_i = r] \cdot \mathbb{P}[e_i \leq r \mid R_i = r] + \mathbb{E}[u_i \mid R_i = 0] \\
&= \beta_0 + \mathbb{E}[\beta_i \mid R_i = 0] \cdot \lim_{r\downarrow 0} \mathbb{P}[e_i \leq r \mid R_i = r] + \mathbb{E}[u_i \mid R_i = 0],
\end{aligned}$$

where the third equality uses (b) and the last uses (a). Likewise, $\lim_{r\uparrow 0} \mathbb{E}[Y_i \mid R_i = r] = \beta_0 + \mathbb{E}[\beta_i \mid R_i = 0] \cdot \lim_{r\uparrow 0} \mathbb{P}[e_i \leq r \mid R_i = r] + \mathbb{E}[u_i \mid R_i = 0]$. Subtracting the two outcome

limits and dividing by the difference of the two first-stage limits, which is nonzero by (c), gives Equation (C.1).

For the last statement, with probability π, $R_i = R_i^*$, and with probability $1-\pi$, $R_i = R_i^* + e_i$ with $e_i \sim G$ independent of R_i^*. The density of R_i is therefore $f_R(r) = \pi f(r) + (1-\pi)h(r)$, with $h(r) = \int f(r-s)\,dG(s)$, and

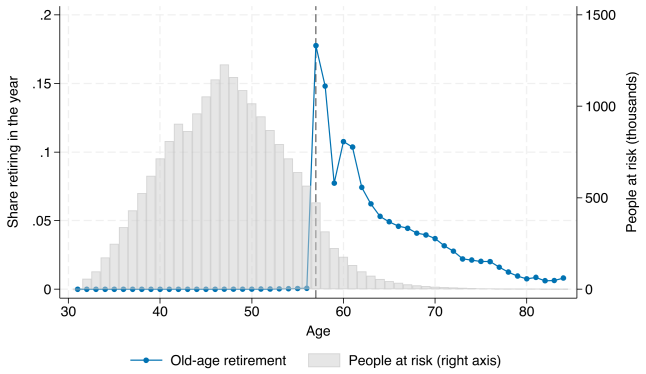
$$\mathbb{E}[A_i \mid R_i = r] = \frac{\pi f(r)\,\mathbb{1}\{r \geq 0\} + (1-\pi)k(r)}{f_R(r)}, \qquad k(r) = \int f(r-s)\,\mathbb{1}\{s \leq r\}\,dG(s).$$

Since f, h, and k are continuous at zero,

$$\lim_{r \downarrow 0} \mathbb{E}[A_i \mid R_i = r] - \lim_{r \uparrow 0} \mathbb{E}[A_i \mid R_i = r] = \frac{\pi f(0) + (1-\pi)k(0)}{f_R(0)} - \frac{(1-\pi)k(0)}{f_R(0)} = \frac{\pi f(0)}{f_R(0)} \in (0, 1]. \qquad \square$$

C.3 Age at Retirement in the Analysis Sample

Figure C.2: Retirement hazard by age, women



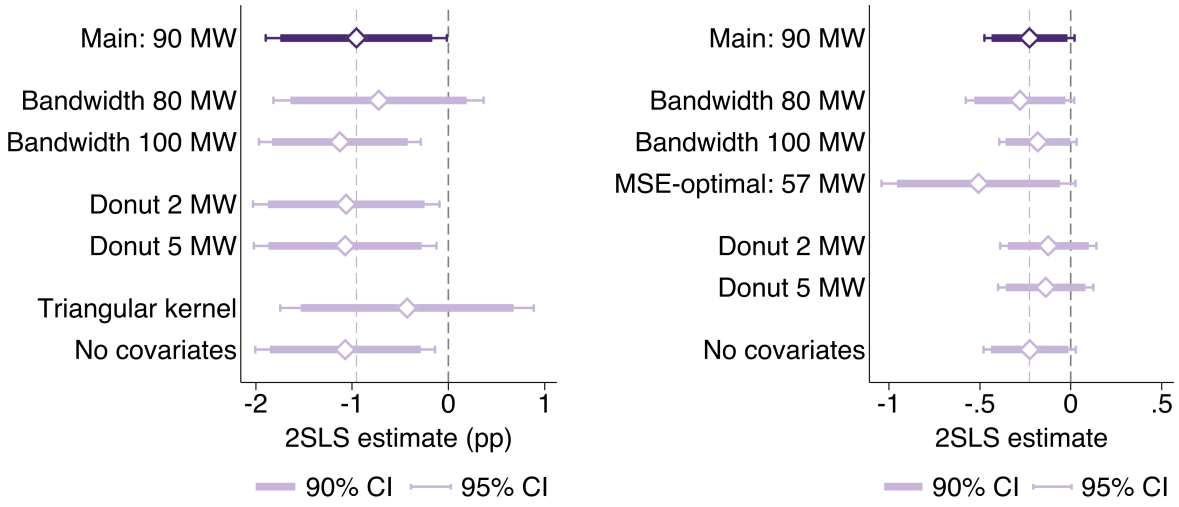
Notes: Female affiliates of the private pension system observed in the annual records, 2008–2023. Construction as in Figure 5: each marker is the share of women aged a at December 31 of the year, among those at risk, who retire during the year with an old-age pension of any payout form; bars are the number of women at risk at each age, in thousands (right axis). The dashed line marks the statutory retirement age of 57.

D Robustness and Additional Mortality Results

D.1 Robustness across Specifications

Bandwidth selection. The mean-squared-error-optimal bandwidth of Calonico et al. (2014) is defined for a single outcome per unit and assumes independent observations. The hazard model in Equation (3) is estimated on a person-year panel, in which each retiree contributes one row per year at risk at the same value of the running variable, so the selector has no direct counterpart for the hazard. Applications of regression discontinuity to survival outcomes address this by moving the bandwidth choice to the person level. Schuessler et al. (2026) fix a horizon, define one outcome per person, such as survival past the horizon or restricted mean survival time, and apply the standard selector to that outcome. Odden et al. (2023) estimate a Cox model within a window but choose the window with the selector applied to a patient-level outcome. I follow the same approach. I compute the bandwidth of Calonico et al. (2014) for the reduced form at the retiree level, using each retiree's realized annual mortality rate, the quantity that the hazard regression averages. The result is 90 minimum wages. This choice is not unique. Different person-level outcomes, such as death within 2, 5, or 10 years of retirement, yield different bandwidths, ranging from 70 to 120 minimum wages. For that reason Figure D.1 reports the main estimates at alternative bandwidths, together with the other specification checks.

Figure D.1: Robustness of the main estimates across specifications



Panel A. Annual hazard of death (pp)

Panel B. Sisben registration

Notes: The dark marker is the main estimate, using the MSE-optimal bandwidth of 90 MW (for the mortality outcome). Purple markers denote the estimates of alternative specifications: bandwidths of 80 and 100 MW, donuts of 2 and 5 MW around the threshold, a triangular kernel, and no covariates.

D.2 Cumulative Mortality by Horizon

This section estimates an alternative specification for the mortality results, as cumulative mortality. I estimate the effect of annuitization on the probability of dying within 2, 5, and 10 years of retirement. Each horizon is a separate regression with one observation per retiree. Retirement cohorts are 2008–2015, fixed for every horizon.

The first stage and the second stage are:

$$A_i = \pi_0 + \alpha\, Z_i + f(R_i) + g(a_i) + \gamma_{c(i)} + \mathbf{x}_i'\boldsymbol{\pi} + v_i, \tag{D.1}$$

$$y_i^k = \beta_k\, A_i + f(R_i) + g(a_i) + \gamma_{c(i)} + \mathbf{x}_i'\boldsymbol{\delta}_k + \varepsilon_i^k, \qquad k = 2, 5, 10, \tag{D.2}$$

where y_i^k equals one if retiree i dies within k years of retirement and zero otherwise. A_i, Z_i, R_i, $f(R_i)$, $\gamma_{c(i)}$, and $\mathbf{x}_i$ are as in Equations (2) and (3). With one observation per retiree, the calendar-year fixed effects drop out, and $g(a_i)$ is a quadratic in age at retirement instead of a quartic in current age. The coefficient β_k is the effect of the annuity on the probability of dying within k years of retirement for retirees at the threshold. I report it multiplied by 100, in percentage points.

Table D.1: Effect of annuitization on cumulative mortality

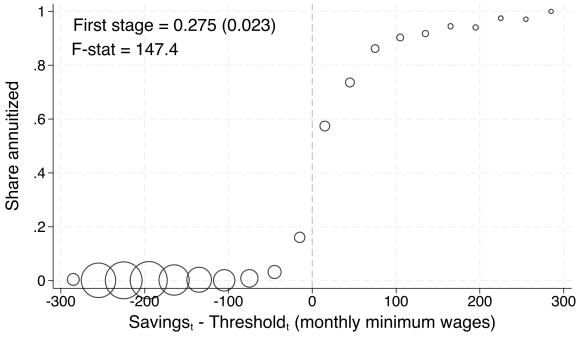
	Died within		
	2 years	5 years	10 years
	(1)	(2)	(3)
2SLS			
Annuitization (A_i)	-3.01**	-5.95**	-6.50
	(1.49)	(2.53)	(4.41)
Baseline lump-sum	2.27	7.20	9.49
Reduced form			
Above threshold (Z_i)	-0.98**	-1.93**	-2.11
	(0.48)	(0.81)	(1.43)
Baseline below	1.17	3.47	9.02
First stage			
Above threshold (Z_i)	0.324***	0.324***	0.324***
	(0.025)	(0.025)	(0.025)
First-stage F	168.5	168.5	168.5
Individuals	5,968	5,968	5,968
Deaths	56	166	456

Notes: Estimates of Equations (D.1) and (D.2). Baseline lump-sum is the mean outcome at the cutoff of compliers without the annuity. Baseline below is the mean outcome below the cutoff within the window. Robust standard errors in parentheses. Individuals are the retirees within the optimal window. Deaths are those within k years of retirement. *, **, and *** denote significance at the 10, 5, and 1 percent levels.

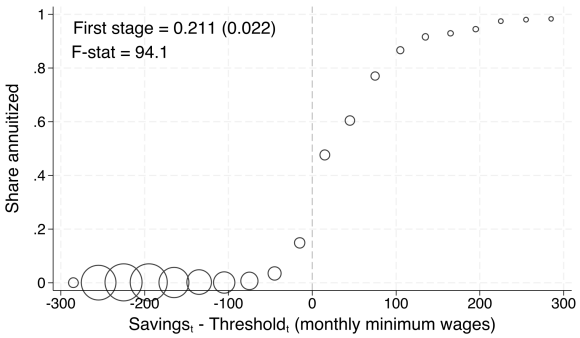
D.3 Alternative Imputations of Family Structure

This section tests the robustness of the first stage and the main results to the imputation of family structure. The main specification uses the observed family structure for the largest pension fund and the imputed structure for the other funds. I consider three alternative versions. The first assigns the cell median to every retiree, including those of the largest fund. The second computes the cell median counting only spouses within 20 years of the affiliate's age and ignores other types of dependents. The third prices every retiree as single. Figure D.2 presents the first stage, Table D.2 the effects on the hazard of death, and Table D.3 the effects on Sisben registration.

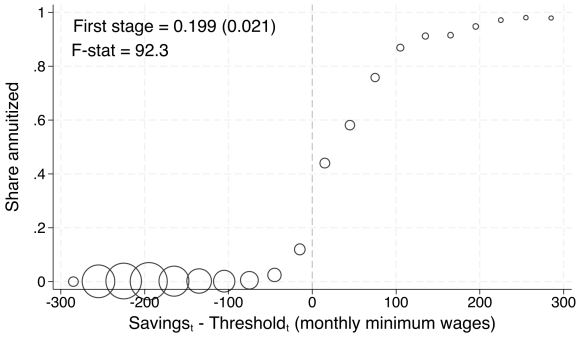
Figure D.2: Annuitization at the threshold under alternative imputations of family structure



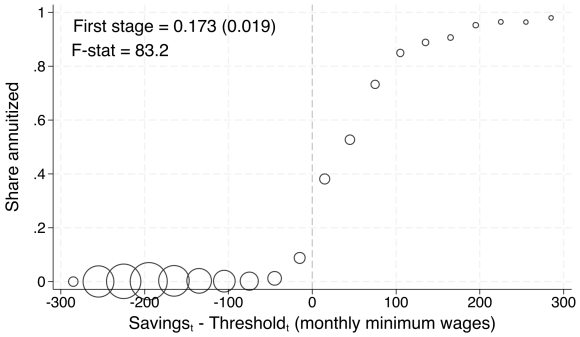
Panel A. Main specification



Panel B. Cell median – all funds



Panel C. Cell median, spouses only



Panel D. Everyone single

Notes: Markers show the share annuitized in bins of 30 minimum wages of the running variable, with size proportional to the number of retirees in each bin. The first stage in each panel is the local linear estimate with a triangular kernel within 90 minimum wages of the cutoff, with its conventional standard error in parentheses, and the F-statistic is its first-stage F. Panel A reproduces Figure 9.

Table D.2: Effect of annuitization on the annual hazard of death under alternative imputations of family structure

	Main	Cell median – all funds	Cell median, spouses only	Everyone single
	(1)	(2)	(3)	(4)
Annuitization (A_i), 2SLS	-0.96**	-1.10*	-1.20*	-0.96
	(0.48)	(0.61)	(0.72)	(1.08)
Above threshold (Z_i), reduced form	-0.30**	-0.26*	-0.25*	-0.13
	(0.15)	(0.14)	(0.15)	(0.15)
First stage (α)	0.310***	0.240***	0.209***	0.138***
	(0.026)	(0.027)	(0.025)	(0.020)
First-stage F	137.8	80.5	70.0	45.3
Person-years	63,356	62,223	68,653	78,304
Individuals	5,512	5,459	6,005	6,759
Deaths	508	473	580	745

Notes: Each column re-estimates the reduced form and 2SLS of Table 4 under each version of the threshold imputation, within 90 minimum wages of the cutoff with a uniform kernel. Column (1) reproduces Table 4. All specifications include the same controls as Equations (2) and (3). Coefficients and standard errors are in percentage points. Standard errors are clustered by retiree. *, **, and *** denote significance at the 10, 5, and 1 percent levels.

Table D.3: Effect of annuitization on Sisben registration under alternative imputations of family structure

	Main	Cell median – all funds	Cell median, spouses only	Everyone single
	(1)	(2)	(3)	(4)
Panel A. 2SLS				
Annuitization (A_i)	-0.227***	-0.229**	-0.068	-0.095
	(0.127)	(0.155)	(0.172)	(0.227)
	[0.007]	[0.017]	[0.241]	[0.203]
First stage (α)	0.243	0.198	0.177	0.125
	(0.032)	(0.031)	(0.029)	(0.025)
First-stage F	58.8	40.7	37.6	24.5
Panel B. Reduced form and balance				
Pre-retirement balance	0.025	0.003	0.010	0.012
	(0.027)	(0.027)	(0.027)	(0.025)
	[0.638]	[0.960]	[0.672]	[0.374]
Post-retirement registration	-0.055**	-0.045*	-0.012	-0.012
	(0.031)	(0.031)	(0.031)	(0.029)
	[0.048]	[0.051]	[0.295]	[0.255]
Observations, pre-retirement balance	5,399	5,358	5,887	6,555
Observations, post-retirement registration	5,004	4,986	5,425	6,014

Notes: Post-retirement registration is registration in Sisben IV in June 2025. Pre-retirement balance is registration in Sisben III. Each estimate is a separate RD within 90 minimum wages, with retirement-year fixed effects. Column (1) reproduces column (1) of Table 7. Conventional estimates and standard errors in parentheses. Robust bias-corrected p-values in brackets. *, **, and *** denote significance at the 10, 5, and 1 percent levels.

E Magnitudes in Context

E.1 Replicating the Specification of Larrain et al. (2026)

I replicate the results of Larrain et al. (2026), adapting their strategy to my regression discontinuity design. Following them, I build a panel with one observation per retiree and year after retirement, $j = 1, \ldots, T$, for horizons of $T = 5$ and $T = 10$ years, and estimate

$$\text{Dead}_{ij} = \beta A_i + f(R_i) + \gamma_{c(i)} + \Omega_i + \mathbf{x}_i'\boldsymbol{\delta} + \varepsilon_{ij}. \tag{E.1}$$

The outcome Dead_{ij} equals one if retiree i has died by year j after retirement, so β is the effect of annuitization on mortality averaged over the T years. They instrument A_i with pension fund returns in the six months before retirement. I instrument it with $Z_i = \mathbb{1}\{R_i \geq 0\}$, and the control $f(R_i)$ of Equation (2) replaces their control for savings. The terms $\gamma_{c(i)}$ are retirement-year fixed effects and Ω_i are retirement-month fixed effects and five birth-year bins. The vector $\mathbf{x}_i$ contains years of contributions and sex. Standard errors are clustered by birth year.

Table E.1 reports the results, with β multiplied by 100. The 2SLS estimate is -3.08 percentage points at five years and -4.79 at ten years, 1.2 and 1.3 times their estimates of -2.55 and -3.62.

Table E.1: Replication of Larrain et al. (2026) within my regression discontinuity setting

	Died within	
	5 years	10 years
2SLS: Annuitization	-3.08*	-4.79*
	(1.641)	(2.644)
Reduced form: Above threshold	-0.99*	-1.55*
	(0.496)	(0.826)
First stage: Above threshold	0.322***	0.322***
	(0.032)	(0.032)
First-stage F	102.1	102.2
Individuals	5,968	5,968
Deaths	166	456

Notes: Estimates of Equation (E.1). Standard errors clustered by birth year in parentheses. *, **, and *** denote significance at the 10, 5, and 1 percent levels.

E.2 The Value of the Annuity at the Threshold

Crossing the threshold changes the form of the payout. The question is whether it also changes the value of what the retiree receives. The guaranteed income comes with features the lump sum may lack, such as protection against inflation and market fluctuations. Most important in this setting, the annuity at the threshold pays one minimum wage and rises with it, so the growth of the minimum wage can change its value relative to the lump sum. This exercise therefore tests the extent to which the annuity is actuarially fair at the threshold.

Following Mitchell et al. (1999) and Poterba (2001), I compute the following ratio

$$\rho_i = \frac{\text{EPDV}_i(1 \text{ MW for life})}{\text{SPM}_i} - 1,$$

where the numerator is the expected present discounted value of the annuity at the threshold, one minimum wage for life, and the denominator is the premium the retiree pays for it, which at the threshold is the SPM itself. If $\rho_i = 0$, the annuity is exactly actuarially fair. If $\rho_i < 0$, it is actuarially unfair, since the retiree pays more than the annuity is worth. If $\rho_i > 0$, the annuity is worth more than its premium.

Under the regulator's assumptions ρ_i is zero by construction, because the SPM is the EPDV of one minimum wage for life computed with those assumptions. The value of ρ_i therefore depends on these assumptions, which need not be realistic. I relax three of them and recompute the EPDV in scenarios that replace each with a more realistic one.

- *Pension growth.* Until August 2015, the formula prices the pension as constant in real terms. From September 2015, it lets the pension grow at a pre-established productivity growth rate.
- *Discount rate.* The formula discounts at a fixed real rate of about 4 percent.
- *Mortality.* The formula uses regulatory mortality tables, which were updated in 2010.

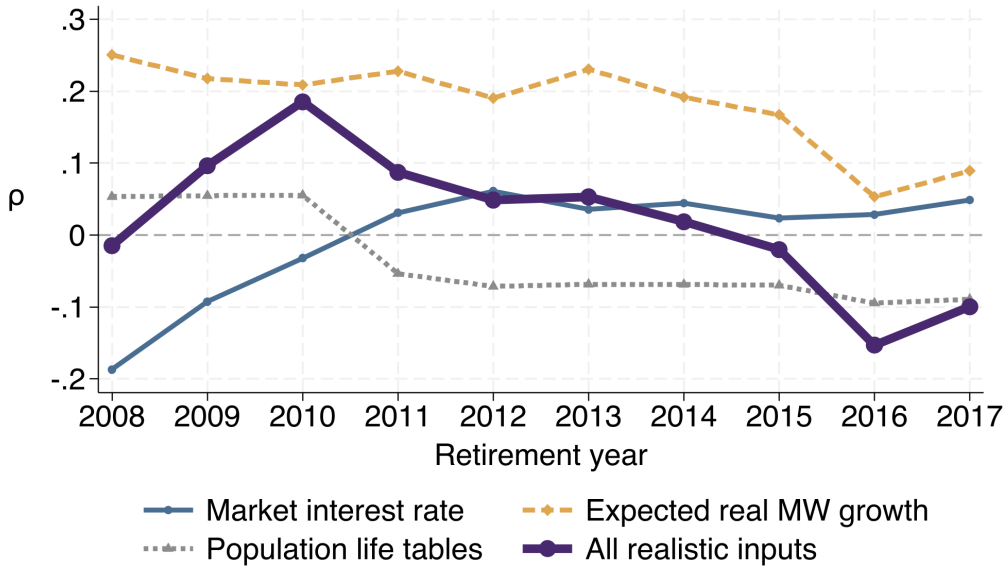
For each of the three assumptions, Table E.2 reports the legal value and the more realistic value I use instead. Figure E.1 shows the median ρ_i by retirement cohort in each scenario, changing one assumption at a time, and with all three changed together (dark purple line).

Table E.2: Inputs of the SPM formula: legal and realistic values

Parameter	Legal input	Realistic input
Pension growth	Constant in real terms until August 2015 (Res. 1875 de 1997), then a pre-established productivity growth rate (Res. 3099 de 2015).	Average real growth of the minimum wage over the ten years before retirement.
Discount rate	A real rate of about 4 percent (Res. 1875 de 1997 and Res. 3099 de 2015).	Market real rate, the 10-year zero-coupon yield on inflation-indexed government bonds (TES-UVR) in the month of retirement, from the yield curve of the central bank (Banco de la República).
Mortality	Regulatory mortality tables, updated in 2010 (Res. 585 de 1994 and Res. 1555 de 2010).	Population life tables of the national statistics office (DANE, *Departamento Administrativo Nacional de Estadística*), by sex and single age, following each birth cohort.

Pricing the growth of the minimum wage alone raises the value of the pension by 17 to 25 percent for the 2008–2015 cohorts. The market rate and the population tables move it less. With all three inputs, the pension is worth 5 to 19 percent more than the SPM for the 2009–2013 cohorts, about the same for 2008, 2014, and 2015, and 10 to 15 percent less for 2016 and 2017. These results hold at the threshold. Above it, retirees pay more than the SPM, and pensions above one minimum wage rise with prices.

Figure E.1: Ex-ante value of the 1-MW pension relative to the SPM, medians by cohort



Notes: Each line is the median ρ_i among retirees of the cohort retiring in that year, with one input of the legal formula replaced by its realistic value (blue: market interest rate, the TES-UVR 10-year yield; yellow: expected real MW growth; gray: population life tables, DANE cohort), and with all three replaced and the regulator's loads removed (purple). A value of zero means the annuity is actuarially fair.

E.3 Benchmarking the Wealth Transfer

The previous exercise suggests that, for some retirement cohorts, crossing the threshold may increase not only the security of retirement income but also its value. Because the annuity at the threshold is indexed to the minimum wage, accounting for its real growth makes the annuity worth 17 to 25 percent more than the lump sum for retirees in 2008–2015. This is a substantial difference, as it applies to lifetime pension wealth. In this section, I translate this additional value into dollars and benchmark it against estimates from the literature on wealth transfers and mortality among older adults to assess how much of the observed mortality effect it could plausibly explain.

I use the scenario that adjusts only expected minimum-wage growth (the yellow line in Figure E.1), which produces the largest positive valuation and therefore provides a generous benchmark for the wealth channel. The median ρ_i across the 2008–2015 cohorts is 0.19, implying that the annuity at the threshold is worth about 19 percent more than the lump sum. Because the EPDV is linear in pension benefits, this is equivalent to receiving a pension 19 percent larger for life. In monetary terms, the implied additional pension wealth, $\text{SPM}_i \times \rho_i$, has a median of USD 10,500, using an exchange rate of 4,000 pesos per

dollar.

To measure how much a 19 percent increase in lifetime income can change the years lived over my 17-year window, I start from the largest gain possible in that window. Retirees below the cutoff live 15.67 of the 17 observed years on average. The difference, 1.33 years, is the most they could gain within the observed period. A 19 percent increase in lifetime income reduces mortality by $-100\,\rho\,\varepsilon$ percent, where $\rho = 0.19$ and ε is the elasticity of mortality with respect to income. If this reduction applies to the probability of having died by each year of the window, retirees gain the same percent of the 1.33 years. The implied gain in years lived is

$$\Delta E \;=\; -\,\rho\,\varepsilon \times (17 - 15.67) \;=\; -\,\rho\,\varepsilon \times 1.33.$$

I compare ΔE with the reduced-form gain of 0.42 years, about five months (Table 6).[42]

I compute ΔE with the elasticity from four studies that estimate the effect of lasting increases in income or wealth on the mortality of older adults. The studies measure transfers in different currencies and settings, so I compare them through the elasticity, which expresses each effect relative to the size of the income change. The studies I benchmark with are the following.

- Malavasi and Ye (2024). A German pension top-up raises the pension of low earners by about 8 percent for life and lowers the probability of dying before 70 by 5.4 percent. The authors report $\varepsilon = -0.63$.[43] The implied gain is 1.9 months, 38 percent of the reduced-form gain.

- Miglino et al. (2023). Chile's basic pension raises the income of poor older adults by 72 percent and lowers four-year mortality by 28 percent. The authors report $\varepsilon = -0.386$. The implied gain is 1.2 months, 23 percent of the reduced-form gain.

- Snyder and Evans (2006). The US Social Security notch raised the income of men born in late 1916 by about 4 percent and their five-year mortality after 65 by about 2 percent. This implies $\varepsilon = +0.5$, the opposite sign of a wealth effect, and a loss of 1.5 months.[44]

[42] Expected years lived over the window are $E = 17 - \int_0^{17} F(t)\,dt$, where $F(t)$ is the probability of having died by year t. Scaling $F(t)$ by $(1 + \rho\,\varepsilon)$ at every t changes E by $-\rho\,\varepsilon \int_0^{17} F(t)\,dt = -\rho\,\varepsilon \times 1.33$.

[43] They also report elasticities for dying before 65 (-2.2) and before 75 (-0.19). I use the middle horizon.

[44] The study does not report an elasticity. I compute it as the 2 percent change in five-year mortality divided by the 4 percent change in family income.

- Cesarini et al. (2016). Swedish lottery prizes have no effect on mortality over ten years, so I set $\varepsilon = 0$ and the implied gain is zero.

Table E.3 summarizes the comparison. A transfer of this size implies at most two months of life, less than two fifths of the reduced-form gain. With the largest cohort median of ρ_i, 0.25, the largest benchmark rises to 2.5 months, half of the reduced-form gain. Applied to compliers, whose expected years lost are 1.76, and compared with the 2SLS gain of 1.35 years, the largest benchmark with $\rho = 0.19$ implies 2.5 months, 16 percent of that gain.

Table E.3: Income and wealth effects on mortality, scaled to the transfer at the threshold

	Transfer	Mortality, as reported	Share of the RF effect
Malavasi and Ye (2024)	German pension top-up: +8 percent for life	chance of dying before 70 down 5.4 percent	38 percent
Miglino et al. (2023)	Chilean basic pension: +72 percent income for life	4-year mortality down 28 percent	23 percent
Snyder and Evans (2006)	Social Security notch: +4 percent income for life	5-year mortality after 65 up 2 percent	wrong sign
Cesarini et al. (2016)	Swedish lottery windfalls	no effect on 10-year mortality	0 percent