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Unfair Redistribution in Actuarially Unfair Pension Schemes: Evidence from Italy

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Abstract

Recent pension reforms in many countries, including Italy, have adopted annuity conversion factors that link key pension parameters to average life expectancy, in line with the principle of actuarial fairness. However, when life expectancy systematically differs across socioeconomic groups, relying on uniform factors can lead to *perverse redistribution*—transferring resources from individuals with shorter lives (typically lower-income) to those with longer lives (typically higher-income). Using rich administrative microdata from a representative sample of private-sector employees in Italy, we quantify the extent of this perverse redistribution by comparing pension wealth distributions under uniform versus differentiated (heterogeneous) annuity factors. We focus on cohorts born between 1960 to 1971 who retire with an old-age or seniority pension between 2021 to 2038. Results show that redistribution is both perverse (from poor to rich) and increasing over time. This trend is driven primarily by widening income-related longevity gaps and further reinforced by Italy's ongoing transition from a defined benefit to a notional defined contribution scheme.

Keywords: pension, redistribution, annuity, life expectancy

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1. Introduction

Pension systems are designed to insure individuals against the uncertainty of death—that is, longevity risk. By incorporating life expectancy into the computation of benefits, the system implies a form of ex-post redistribution where individuals who die earlier cross-subsidize those who live longer. Such redistribution is intentional and typically regarded as socially desirable, as it helps to ensure that retirees have a stable income throughout their retirement years (Diamond, 2004).[1]

However, when systematic longevity differentials exist across socioeconomic groups and are not reflected in the annuities calculation, pension system may unintentionally generate *ex-ante* redistribution. In particular, uniform annuity factors—based on average life expectancy—implicitly transfer resources from those with shorter expected lifespans (typically low-income earners) to those expected to live longer (typically higher-income earners). This mechanism gives rise to what we refer to as *perverse redistribution*, which is neither actuarially fair nor socially desirable (Brown, 2003; Caselli and Lipsi, 2018; Jijiie et al., 2021; Lallo and Raitano, 2018).

In this paper, we measure the magnitude and evolution of such perverse redistribution in the Italian pension system. We focus on the period of transition from the traditional Defined-Benefit (DB) scheme to the Notional Defined-Contribution (NDC) system. Our analysis compares the distribution of pension wealth under the current uniform annuity factor, with a counterfactual distribution based on annuity factors diversified by socioeconomic position (SEP), as proxied by earnings quantiles. By estimating the Reynolds-Smolensky progressivity index, we measure the redistribution originating from the application of uniform annuity factors in the presence of

[1] A progressive pension system, by violating actuarial fairness at the individual level, allows to move part of the redistribution from the working life, where tax progressivity applies to annual earnings (tax base) and is immediately paid, to the elder life, where redistribution is based on lifetime earnings and paid after retirement.

significant longevity differentials by earnings.

Over the past decades, concerns over population ageing and fiscal sustainability have prompted widespread pension reforms. These reforms have typically tighten the link between contributions and benefits, in line with the principle of *actuarial fairness*[2]. As such, *annuitization* and the use of annuity factors have become central components of public pension benefit formulas (Holzmann and Palmer, 2006).

Despite robust empirical evidence documenting substantial longevity differentials by income, education, and occupation (Ardito et al., 2024; Chetty et al., 2016; Fedotenkov and Derkachev, 2020; Mackenbach et al., 2008; Nelissen, 1999), no public pension system, to our knowledge, currently applies differentiated annuity factors. Several factors may explain this inertia, including limited financial literacy, administrative complexity, legal concerns about discrimination, and the political invisibility of deceased individuals (Auerbach et al., 2017). Nonetheless, ignoring mortality differentials implicitly legitimates redistribution from poorer to richer retirees, a dynamic that violates both actuarial and distributive principles.[3] Worryingly, this redistribution is expected to become increasingly regressive, as socioeconomic mortality differentials continue to widen (Ardito et al., 2024; Bosworth and Burke, 2014; Chetty et al., 2016; Haan et al., 2020). It is therefore unsurprising that such issues have began to attract growing attention in policy debates in both the United States and Europe (Auerbach et al., 2017; OECD, 2023).

To study this phenomenon, we use administrative microdata from the Work Histories Italian Panel (WHIP), a 7% representative sample of employment and pension records from Italy's National Institute of Social Security (INPS). This database allows us to (i) estimate life expectancy differentials

[2]For major details on actuarial fairness see Belloni and Maccheroni (2013).

[3]The convergence of actuarial and distributive unfairness is particularly evident in income-based longevity gaps, where shorter life expectancy aligns with lower income. A similar alignment often holds for education and occupation, but may be less direct for dimension like genders.

by income groups, and (ii) simulate prospective pension wealth for male private sector employees born between 1960 to 1971 and retiring between 2021 to 2038. While most working career data are observed, we simulate end-of-career employment and earnings trajectories using recursive logistic employment model and a wage equation. The analysis focuses only on individuals eligible for old-age or seniority pension—so excluding social assistance mechanisms—thereby isolating the insurance component of the pension system and limiting sample heterogeneity; this is key to identify the effect of an undiversified annuity factor.

The initial pension benefits are computed under legislation in force as of 2024, distinguishing between three regimes: Fornero pro-rata (DB with NDC after 2011), Dini pro-rata (DB with NDC after 1995), and fully NDC. Using both the expected pension stream and projected life expectancy, we calculate factual pension wealth (under uniform life expectancy), and counterfactual pension wealth (under SEP-specific life expectancy) for each retiree. By constructing distributions by year of birth (cohort) and retirement year, we are able to disentangle the redistributive effects of both (i) increasing longevity differentials and (ii) the transition from DB to NDC.

To quantify redistribution, we estimate the Reynolds-Smolensky progressivity index (Reynolds and Smolensky, 1977) to both factual and counterfactual distributions. We find that uniform annuity factors generate substantial perverse redistribution, which increases significantly from 2030 onward. The dynamic is driven by both widening longevity gaps and the full implementation of the NDC regime.

Our contribution is mainly empirical: we quantify the extent of perverse redistribution induced by the Italian public pension system in the presence of longevity differentials and we propose a measurement methodology inspired by tools used in tax progressivity analysis, adapted to pension wealth. Crucially, our work offers several key advantages over existing studies. First, it is prospective—we compute pension wealth based on observed and simulated career paths of individuals approaching retirement. Second, it focuses solely on the contributory (insurance) component of the pension system, abstracting from means-tested or assistive mechanisms. Finally, as far as we reconstruct—almost entirely—real working careers of individuals, we are able to identify the distinct contributions of two critical forces: (i) the gradual transition from the DB scheme to the NDC system, and (ii) the projected increase in longevity gaps by income. From a policy perspective, this distinction is critical for understanding the future dynamics of redistribution and evaluating potential reforms to mitigate inequity in pension outcomes.

The rest of the paper is organized as follows. Section 2, reviews the literature on longevity differentials and their implications for pension systems, with a focus on the Italian case. Section 3 presents data and methods, including the estimation and projection of life expectancies, the simulation of end-of-life working careers, and the quantification of the redistributive impact of longevity differentials based on a counterfactual approach. Section 4 reports our findings and their policy implications. Section 5 concludes.

2. State of art and institutional background

The issue of unintended redistribution resulting from longevity differentials in pension systems, has been acknowledged since the earliest studies on longevity differences across socioeconomic groups (Antonovsky, 1967; Kitagawa and Hauser, 1973).

Pioneering contributions such as Garrett (1995) and Nelissen (1999) established that income-based mortality differentials can offset the progressivity of pension systems. More recently, evidence from both U.S. and European countries is pointing to a large and widening life expectancy gaps across socioeconomic groups (Ardito et al., 2024; Chetty et al., 2016; Fedotenkov and Derkachev, 2020; Mackenbach et al., 2008)[4].

[4]Longevity may differ across socioeconomic groups depending on several variables,

In what follows, we focus on studies that examine the distributive implications in pension systems of male longevity differentials by income (or earnings), consistent with the scope of our analysis. The focus on males is justified by the methodological challenges associated with estimating mortality differentials among female workers. In particular, these estimates tend to be less stable due to historically lower and more variable labor force participation rates among women, as well as limited sample sizes (Waldron, 2007; Bosworth and Burke, 2014; Auerbach et al., 2017).

Moreover, when Socioeconomic position (SEP) is measured using individual characteristics such as income or occupation, the resulting longevity differentials for women are often less meaningful. This is especially true for cohorts with low female labor market attachment, where women's health and mortality outcomes are more strongly influenced by household or spousal socioeconomic conditions than by their own individual status (Ardito et al., 2024; Bosworth and Burke, 2014; Ghislandi and Scotti, 2022; Haan et al., 2020). Finally, we conclude this section with an overview of the main reforms to the Italian pension system, focusing on those that have been most relevant to the gradual transition from a Defined Benefit (DB) to a Notional Defined Contribution (NDC) scheme.

2.1. Related literature

To estimate longevity differentials, Whitehouse and Zaidi (2008) use data from three sources, namely, the British Household Panel Survey (BHPS), the US Panel Study on Income Dynamics (PSID) and the German SOcio-Economic Panel (GSOEP). Focusing on males aged 40, they group individuals into tertiles based on equivalized household net income and find that the life expectancy (hereafter, LE) gap between the top and bottom income

e.g., education, income, occupation, health (Cannari and D'Alessio, 2016; Fedotenkov and Derkachev, 2020; Lallo and Raitano, 2018; Mackenbach et al., 2008; Preston and Elo, 1995; Mackenbach et al., 2019; Katikireddi et al., 2017) tertiles is approximately 5 years in both the UK and the US, and 3.7 years in Germany.

To measure the implications of these LE-gaps for pension systems, they combine OECD earnings-distribution data with official OECD replacement rates for the three countries. Assuming that longevity differentials by household incomes proxy those by individual earnings, they simulate the distribution of pension wealth as a function of the earnings distribution.[5] Progressivity is then calculated as 100 minus the ratio between the Gini of pension wealth and the Gini of earnings. This metric is calculated under two scenarios: (i) a *factual* scenario using uniform annuity factor based on average LE and (ii) the *counterfactual* scenario using differentiated annuity factors by lifetime income. Their findings show that ignoring income-based longevity differentials reduces the progressivity of pension systems by 4 to 7 percentage points across the three countries.

Bosworth and Burke (2014) analyze socioeconomic disparities in life expectancy (LE) using the Health and Retirement Study (HRS), defining SEP by educational attainment and mid-career average earnings (ages 41–50). For males aged 55, they find an average LE-gap between the bottom and the top SEP-decile that increases from 4.1 years in the 1920 cohort to 8.2 years in the 1940 cohort.[6] To assess distributional effects of longevity inequality, they compare the distribution of career earnings across SEP-deciles with that of both annual and lifetime OASDI benefits (Old Age, Survivors, and Disability Insurance). Assuming actuarial fairness, annual OASDI benefits distribution should mirror career earnings, while lifetime benefits should match annual benefits in the absence of LE-gaps. Bosworth and Burke (2014)'s findings show that although the US social security system is broadly progressive, this

[5]For household income, a 10-year moving average is used in the construction of diversified longevity.

[6]The analysis is based on Health and Retirement Study (HRS) and mortality differences are estimated using a proportional hazard model.

is partly offset by income-related LE-gaps, which reduce lifetime benefits for lower-SEP individuals.

Auerbach et al. (2017) compute LE-gaps using HRS data from 1992 to 2008 for cohorts born between 1912 and 1957. SEP is defined by household mid-career earnings (ages 41-50), capturing lifetime resources and smoothing issues linked to female intermittent work histories.

Longevity differentials are obtained from a probit model estimating two-year mortality probabilities by earnings quintile. Among 50-year-old males, the life expectancy gap between the bottom and top income quintiles widens from 5.1 years in the 1930 cohort to 12.7 years in the 1960 cohort. As for the implications of these longevity differentials for the progressivity of the U.S. Social Security system, the authors show that—holding the earnings distribution and pension rules constant—the widening LE-gap raises the disparity in average lifetime benefits received by males in the top and bottom income quintiles by about $103,000.

Haan et al. (2020) analyze German administrative data from the Versicherungskontenstichprobe (VSKT), a random sub-sample of complete monthly employment biographies for cohorts born 1935-1985. Focusing on men born between 1935 and 1949, they estimate earning decile and age-specific mortality rates. They find that (i) life expectancy at age 65 strongly varies between earnings deciles and (ii) LE-gaps by earnings increase across cohorts—from 4 years for those born in 1926–28 to nearly 7 years for those born in 1947–49.

To assess distributional effects, they compute Internal Rates of Return (IRR) and Gini coefficients of pension wealth distributions under two different scenarios: one with estimated heterogeneous mortality (factual), and one with homogeneous, cohort-specific mortality (counterfactual).

Surprisingly, the German pension system appears regressive. For the 1935–1937 cohorts, the IRR is 1.2% in the bottom earnings decile compared with 1.5% in the top, and this disparity widens markedly for the 1947–1949 cohorts (0.38% vs. 1.16%). In monetary terms, top-decile earners gain approximately €80,000–€110,000 in pension wealth due to differential mortality, while bottom-decile earners lose €30,000–€50,000, indicating a substantial perverse redistribution.

For Italy, Ghislandi and Scotti (2022) draw on INPS (Italian Social Security Administration) data covering the universe of employment spells in the Italian private sector (1975–2017) and pension benefits (1995–2017) to examine the distributional impact of longevity differentials. SEP is defined in terms of both mid-career employment income and pension income. For males aged 67, they estimate that the LE-gap between the top and bottom pension quintiles increased from one year in 1995 to nearly two years in 2017.

To evaluate the distributional implications of LE-gaps in the Italian pension system, they consider recipients of old-age, seniority and early-retirement pension benefits and focus on individuals born between 1910 and 1950 who retired by 67.[7] By using a counterfactual approach, they compare the distribution of PW and IRR based on cohort-, sex- and lifetime income quintile-specific mortality profiles previously estimated, with the distribution of PW and IRR based on cohort- and sex-specific mortality profiles, averaged across all lifetime income quintiles. They find that among males born in 1930-1934, those in the top quintile gain about 58,000€ (+5.3%) in PW, while those in the bottom quintile lose around €13,000 (-5.7%). These disparities widen for the 1945-1950 cohorts, with gains reaching €110,000 (+10.8%) for the top quintile and losses of €38,000 (-10.1%) for the bottom. Similarly to Haan et al. (2020), they also track changes in the cohort-specific Gini coefficients of PW. Overall, their results indicate regressive redistribution and that widening disparities in life expectancy among males magnify the erosion

[7] To analyze the distributional impact of longevity differentials, they compute sex-, age-, and pension–income–quintile–specific mortality rates for ages 67 to 85 in each calendar year, defined as the ratio of death counts to population exposure (person–years lived). Mortality rates at older ages (85+) are extrapolated using the Kannisto model, which is particularly effective for approximating mortality at advanced ages (Thatcher et al., 1998). This yields mortality estimates for ages 86 to 120.

of progressivity in the Italian pension system—particularity as reforms have increasingly tied contributions to pension benefits.

Ardito et al. (2024) use INPS administrative data on private sector employees for the period 1990–2019, considering two indicators of SEP: average full-time equivalent weekly wage and occupational social class. They find that the LE-gap for males aged 65 in 1990 is 1.09 years and it increases to 1.79 years in 2000. To assess the distributional implications of these LE-gaps, they estimate that males aged 65 retiring between 1995 and 1997 (with at least 10 years of contributions) in the bottom SEP-quartile would gain approximately €20,000 in pension wealth if they had the same LE of the top SEP-quartile, a figure which increases to €24,000 for retirees between 2001 and 2003.

2.2. Institutional background: progressive transition from DB to NDC

During the nineties, due to financial sustainability problems, the Italian pension system underwent a major reform process, moving from a DB to a NDC pension system, while maintaining the Pay-As-You-GO (PAYGO) financing mechanism.

Until 1992, according to the DB formula, pension benefits were determined by *pensionable earnings*, defined as the average wage earned in the last years before retirement,[8] by the *number of years of contribution at retirement* (topped at 40 years), and by the *accrual rate* of 2% for each year of contribution. The main idea was that of preserving, to some extent, the living standard achieved during the working life.

Workers became eligible for retirement if they paid either at least 35 years of contribution (known as seniority pension), or they were 55/60 years old (female/male) with at least 15 years of paid contributions (known as old-age

[8]The number of years to be included in the computation of pensionable earnings has been progressively increased by reforms in the nineties from 5 up to 10 for employees and from 10 up to 15 for self-employed workers.

pension). In addition, an indexation mechanism to both price inflation and real wages was implemented for the adjustment of the first pension benefit in later life, i.e. after retirement.

Most importantly, the old DB scheme was characterized by several redistributive tools like floors, ceilings, and redistributive accrual rates. These have been slightly amended over time beginning in the eighties. Floors were established to raise pension benefits to a minimum amount, if they were lower. On the other hand, ceilings were used to identify an upper bound for pensionable earnings to be accounted for in the calculation of the first pension. The accrual rate was—and still is—redistributive; it was set at 2% for pensionable earnings below a reference level legislated annually (inflation-adjusted), at 1.50% for pensionable earnings between the reference level and 1.33 times the reference level, at 1.25% between 1.33 and 1.66 times the reference level, and at 1% above 1.66 times the reference level (Brugiavini and Peracchi, 2003). In a sense, the design of the old DB pension scheme was intended to be progressive.[9]

In 1992, to account for inter-generational inequity originating from the financial burden caused by the excessive generosity of the pre-1992 system, the Amato Reform (Law 503/1992) sensibly modified the DB scheme by setting different and more stringent eligibility rules and by gradually increasing the number of years over which pensionable earnings were to be computed. Specifically, while preserving the old DB scheme for years of contribution until 1992 (so-called "Quota A"), a modified DB scheme has been introduced for years of contribution after 1992 (so-called "Quota B"). As for the latter, pensionable earnings are calculated up to the average wage over the last 10 years (15 for self-employed workers) for those with at least 15 years of con-

[9]Even if the benefit formula of the Italian DB scheme is redistributive (Mazzaferro, 2023), evidence of perverse redistribution has been reported in some cases (Abatemarco, 2006; Borella, 2004; Borella and Coda Moscarola, 2006; Castellino, 1995; Fornero and Castellino, 2001; Leombruni and Mosca, 2016).

tributions in 1992. For workers with fewer than 15 years of contributions in 1992, the reference period for pensionable earnings is gradually extended to include all contribution years, starting in 2001. The extension of the averaging period for the computation of pensionable earnings was intended (i) to lower the amount of pensionable earnings on average, and (ii) to reduce (undesired) perverse redistribution originating from greater end-of-career income growth in top income classes (Leombruni and Mosca, 2016). Finally, it is worth noting that the Amato reform also changed the indexation mechanism of the initial pension benefit by removing the adjustment to real wage growth, while retaining only the link to price inflation (still in force).[10]

The most significant reform affecting the design of the pension benefit formula was enacted in 1995. The Dini reform (Law 335/1995) modified the structure of the Italian pension scheme by initiating a gradual transition from a DB to a NDC formula. Under the NDC system, the first pension is calculated directly through the annuitization of accrued contributions paid in the working life. To enable a gradual implementation of the new scheme across different cohorts, the reform established distinct eligibility rules and benefit formulas based on each worker's years of contributions as of 1995. In practice, a grandfathering clause—still in effect today—was introduced, drawing a dividing line between *elder*, *middle-aged*, and *younger* workers. This clause ensures that those with sufficient contribution seniority at the time of the reform could retain, fully or partially, the more generous DB rules, while newly entered workers were entirely subject to the new NDC formula.

For newly entered workers (those entering the labor market after 1995), only the NDC benefit formula applies. Under this system, the first pension benefit at retirement is calculated from the total amount of contributions paid during the entire working life, notionally capitalized at the GDP

[10]In addition, the set of accrual rates for "Quota B" was modified through the introduction of a new earnings bracket for very high pensionable earnings.

nominal growth rate, and converted into an actuarially fair annuity using an age-dependent annuity factor (known as Transformation Coefficients)[11]. Specifically, the conversion into a pension annuity is determined by the expected number of years for which pension benefits are to be paid (Coppola et al., 2006), accounting for the life expectancy of both retirees and survivors (see Section C in the appendix for details). According to current legislation, annuity factors must be updated every two years based on demographic tables provided by the Italian National Institute of Statistics (ISTAT).

For workers with at least 18 years of contribution as of 1995 (elder workers), the DB formula continues to apply. However, pensionable earnings after 1992 are calculated as an average over a longer reference period, according to the modified DB formula set up by Amato reform.

Finally, for workers with less than 18 years of contribution in 1995 (middle-aged workers), pension benefits are determined according to a pro-rata mechanism (hereafter, Dini's pro-rata). In this case, the portion of the pension corresponding to pre-1995 accrued contribution is computed using the modified DB formula, while the contributions paid after 1995 are subject to the NDC scheme until retirement.

After the Dini reform, several legislative acts intervened to introduce more stringent rules concerning both eligibility requirements and pensionable earnings.[12] Among these, it is worth recalling the Monti-Fornero Reform (Decree-Law 201/2011), as it introduces several measures aimed at tightening, while

[11] Transformation Coefficients vary from 57 to 71 years old according to the worker's retirement age.

[12] In 1997, to further reduce the tax burden of the pension system, the Prodi government raised the retirement age, which was increased again a few years later by the Berlusconi government (Law 243/2004). Another step was taken in 2007, when the Prodi–Damiano Reform (Law 247/2007) amended the previous reform by introducing the quota system—age and contribution requirements for claiming a pension—aimed at progressively increasing the retirement age. Regarding retirement age adjustments, the Sacconi Reform (Law 102/2009) introduced, for the first time, an indexation mechanism whereby the retirement age is automatically adjusted to changes in life expectancy.

also making more flexible, retirement eligibility requirements.[13] For the purposes of our analysis, the Monti-Fornero reform specifically amended the modified DB scheme established by the Dini reform for workers with more than 18 years of contributions as of 31 December 1995 (elder workers). More specifically, under this reform, the NDC scheme has to be applied from 1 January 2012 to older workers still employed as of 31 December 2011—despite this possibility having been originally excluded by the Dini reform. As a result, for these elder workers, a pro-rata mechanisms (hereafter, Fornero's pro-rata) now applies, which should not be confused with the Dini's pro-rata that concerns middle-aged workers with fewer than 18 years of contributions as of 1995.

To sum up, given our focus on old-age and seniority pensions for Italian workers retiring after 2020—excluding transitional pro-quota (mixed) mechanisms that combine both age and contribution requirements[14]—we consider three main benefit formulas: the Fornero's pro-rata mechanism (retirees with at least 18 years of contribution as of 1995 but retiring after 2011), the Dini's pro-rata mechanism (for retirees with fewer than 18 years of contributions as of 1995), and the fully NDC benefit formula (for retirees whose contributions

[13]As for the standard old-age pension, provided that (i) contributions have been paid for at least 20 years and (ii) the pension benefit is no less than 1.5 times the old-age social assistance payment, under the new system the age of retirement has been gradually increased for men and women, until 2021, when no categories of workers are able to retire before the age of 67. Similarly, the years of contribution needed to be eligible for a seniority pension (early retirement) has increased gradually for men and women, until 2023, when eligibility was fixed at 42 years and 10 months. Moreover, in line with previous legislative actions, the Monti-Fornero reform established that both age and seniority requirements must be periodically adjusted according to life expectancy coefficients published by ISTAT. Finally, the possibility of a *fully-contributive early retirement* at age 64 has been introduced for early workers with at least 20 years of contributions who have accumulated sufficient contributions to obtain a pension benefit greater than 2.8 times the social assistance pension.

[14]For instance, according to quota 100—in force from 2019 to 2021—eligibility was achieved by workers aged at least 62 years old with 38 years of contributions. We omit these eligibility requirements, as they have been substantially weakened by the Meloni Reform (Law 197/2022).

begin in 1996 or later).[15]

3. Data and Methodology

We use administrative microdata from the Work Histories Italian Panel (WHIP), a 7% representative sample of employment and pension records from the Italian National Institute of Social Security (INPS). Three key preliminary choices are specified at the outset.

First, our study population consists of Italian-born male individuals with a career as dependent employees. The exclusion of foreign-born individuals is motivated by their relatively small share of the Italian population during the period under consideration—below 5% at the turn of the century—which limits the availability of data and the statistical power needed to account for country of origin in the estimation of longevity differentials.

Second, for women, we cannot detect a correlation between mortality rates and the available measure of SEP, which is in line with previous findings for Italy using the same source of data (e.g. Ghislandi and Scotti, 2022). The WHIP section on dependent workers earnings is actually driven from the Italian Social Security Administration (INPS) files, which collect individual earnings and not households' ones. It turns out that the former are a good proxy of SEP only for males, since in the years we are considering women were less attached to the labour market and frequently played a role of second earners in the family.

Finally, we exclude self-employed individuals, as the income information reported in the Social Security data is a poor proxy for their actual earnings and, consequently, for their SEP.[16]

[15]To focus on the insurance mechanism of the Italian pension system, we disregard social assistance pensions. Moreover, due to the limited number of retirees observed (and expected) in our analysis, we exclude uncommon eligibility requirements and benefit formulas allowed by Italian Law like "Opzione Donna", "fully-contributive early retirement", and "early-starting workers pensions".

[16]In Italy, self-employed workers are allowed to pay social security contributions con-

Two different sub-samples of the study population have been considered: a larger one for the estimation of longevity differentials by SEP and their evolution in time; a smaller one for the reconstruction of working careers and factual/counterfactual PW distributions, as detailed in the following sections.

3.1. Life expectancies: differentials and projections

To estimate life expectancies and life expectancy differentials by using period life tables, we considered all Italian born male individuals aged 20-90 in the years from 1990 to 2019, either retired or employed as dependent workers.[17] For all individuals, we compute the Socio-Economic Position (SEP) by classifying them using age-specific tertiles of weekly earnings in all the years available. On average, the dataset consists of about 3 million private sector employees per-year. This number is sensibly increased from 1990 (1,698,199 records) to 2019 (4,149,936 records). Due to a limited number of deaths at initial and final ages (respectively, 20-23 and 87-90 years old), mortality rates—for each year and age—have been calculated by using a 5-yearly moving average by year and age (separately). This procedure is expected to provide more robust estimates of mortality rates in poorly populated cells, at the cost of reduced variability across years and age.

In the actuarial literature, a variety of alternative methodologies for projecting mortality have been proposed (Booth and Tickle, 2008; Lee and Carter, 1992; Russolillo et al., 2011). We adopt the Lee–Carter (LC) model, consistent with the methodology employed by ISTAT to produce the official life expectancy estimates and forecasts, which the Government currently uses to determine various pension system parameters, such as the annuity

sidering a minimum legal income instead than their actual one. It turns out that, for the cohorts we are considering, about two thirds of individuals opted for the minimum level, meaning that, for the vast majority of them we do not observe their actual SEP.

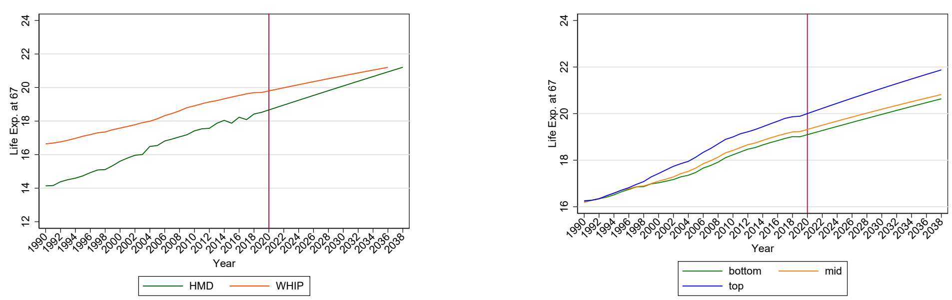
[17]Even if data for the biennium 2020-2021 are available, we omit this information due to the impact of COVID-19 on observed mortality rates.

conversion factors (Transformation Coefficients).

We use the LC method to extrapolate age-specific mortality rates, along with survival and death probabilities, based on historical mortality data from WHIP sample. These projections extend 19 years beyond the base year of 2020, covering the period up to 2038, and include life expectancy estimates for the overall population, as well as stratified analyses by socioeconomic position SEP-tertiles (details in Appendix B).

To assess the representativeness of our projected mortality rates derived from WHIP, which focuses on Italian private sector employees, we compared the resulting life expectancies at retirement with estimates based on data from the Human Mortality Database (HMD), encompassing the entire national population.

Figure 1: WHIP validation and LE by SEP-tertiles



(a) Fitted and projected male LE at 67

(b) Fitted and projected male LE at 67

In Fig.(1a), we observe, on average, a one-year life expectancy gap between WHIP and HMD data, a gap that narrows over time. Given that the WHIP sample historically displayed life expectancy levels similar to the overall population (Leombruni et al. 2010), this pattern may reflect emerging disparities—such as the impact of worsening job conditions or increasing socioeconomic polarization—affecting specific segments of the private sector workforce (Fletcher et al., 2011; Pförtner and Demirer, 2022).[18]

In Fig. (1b), we present our LE estimates (at age 67) based on WHIP data partitioned by SEP-tertiles, which, as expected, reveal a significant socio-economic gradient in life expectancies. Consistent with Leombruni et al. (2010), the differentials are minimal between the first and second tertile, while they become more pronounced for the top SEP-tertile, whose members are expected to live, on average, one year longer than the rest of the population. In the early 1990s, the LE-gap between tertiles appears negligible, although the estimates for these initial years are less robust due to data limitations prior to 1993. Overall, the long-term trend shows a clear widening of the differentials, in line with the recent findings of Ardito et al. (2024), who used INPS social security data covering the entire population.

3.2. Sample selection and working careers

As for the analysis of PW distributions, we restricted the sample to individuals born between 1960 and 1971, corresponding to cohorts transitioning from the DB to the NDC system. We further limited the sample to private-sector employees, since data on work histories for public-sector workers is incomplete. Employment status—whether as a dependent worker or as an employee in the private or public sector—is, in fact, a time-varying characteristic. The vast majority of Italian workers in the selected cohorts have mixed careers, meaning they switched between the public and private sectors and/or between dependent employment and various forms of self-employment. To focus on more stable careers, which better reflect the "normal" functioning of pension legislation, we selected only workers who spent either their entire career or at least 75% of it in the private sector, representing 42.9% and 10.7% of the full sample, respectively. Finally, we retained only individuals who, in 2020 (the last year available in our dataset), were not yet eligible for

[18]This finding is also consistent with what one would expect as a result of an increasing participation in the labor market over time.

retirement under the eligibility rules considered—namely, the seniority and old-age pension requirements set by the 2024 legislation.[19]

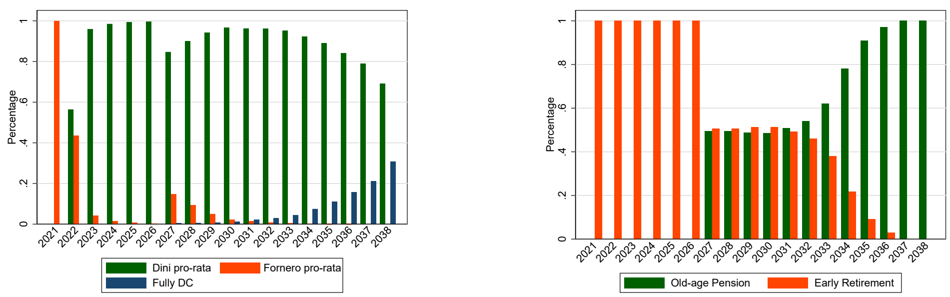
We then simulate their working career and retirement from 2020 to 2038. To complete the working career of individuals, we first simulate the employment status—whether they held a job if employed, or they find a job if not employed—according to a logistic model of employment. Details are reported in Table (1) in Appendix A. Then, for those individuals who are (or should be) still employed in 2020, we estimate their salary, according to the wage equation reported in Table (2) in Appendix A. As regards retirement, for every simulated year we check workers' eligibility to a pension and assume they will retire when they first meet minimum eligibility requirements for an old-age or seniority pension. Projected mortality rates are also considered in order to account for the possibility of pre-retirement death.

3.3. First pension and annuity factors projection

To compute initial pension benefit for retirees from 2021 onward, we consider three main pension benefit regimes established under current legislation: the Fornero pro-rata mechanism, the Dini pro-rata mechanism and the fully NDC scheme. The Fornero pro-rata applies to individuals with at least 18 years of contributions as of 1995 who were still employed in 2020. The Dini pro-rata applies to individuals who began working before 1995 but who had accumulated fewer than 18 years of contributions by that year. Finally, and fully NDC for scheme concerns individuals with no contributions history before 1995. We excluded minor regimes and social assistance mechanisms from our main analysis to focus on these primary structures (see footnote 15).

[19] A small fraction of the selected individuals (1.5%) are pension recipients who would not be eligible under the rules considered, due to the existence of other minor retirement pathways besides seniority and old-age schemes (see footnote 15). To avoid discontinuities around 2024, we disregarded their pensioner status and simulated their careers in the years prior to 2020.

Figure 2: Retirees by pension benefit formula and eligibility rule



(a) Benefit formula by cohort

(b) Eligibility rule by year of retirement

For retirees under Fornero pro-rata mechanism, pension benefits are calculated by combining three components: Quota A, corresponding to contributions accrued before 1992; Quota B, covering contributions from 1992 to 1995; and Quota C, which accounts for contributions paid after 2011 until retirement. Quota C is calculated by using the standard NDC formula, whereby total accrued contributions are multiplied by the Transformation Coefficients (TC), which vary according to age and year.

For retirees falling under the Dini pro-rata mechanism, pension benefits similarly consider Quota A for pre-1992 contributions and Quota B for contributions made between 1992 and 1995. In this case, however, contributions paid from 1 January 1996 onward are calculated according to the NDC formula.

Finally, for individuals entering the workforce after 1995, only the fully NDC benefit formula is applied.

To calculate both actual and virtual pension benefits under the three pension benefit regimes defined by current legislation—for individuals retiring from 2021 to 2038—it is necessary to update TCs. These coefficients, which convert accumulated pension wealth into annuities, are officially revised by the government every two years based on the observed change in life expectancy. At present, official TC are available up to 2024.

To project future TCs for the period beyond 2024, we adopt the same official methodology used by the Italian Government, applying the Lee-Carter model to mortality data from the Human Mortality Database (HMD). Specifically, we used data for Italian male and female populations ranging from 1970 to 2019, for ages from 50 up to 119 years, disaggregated by single year of age and calendar year. Hence, survival and death probabilities are derived from 2020 up to 2036. In turn, TCs are obtained for the bienniums 2023–2024, 2025–2026, 2027–2028, and continuing up to 2037-2038 by using as baseline years 2022, 2024, 2026 and so forth up to 2036, respectively. The estimated and projected TCs are presented in Table (3) in Appendix C, together with a detailed description of the computational procedure.

Once the initial annual pension income has been calculated for each retiree, subsequent annual pensions incomes are adjusted for inflation in accordance to Italian laws in force in 2024 (Law 213/2023). This is done using the forecasting scenario published by the European Commission's Working Group on Ageing Populations and Sustainability (AWG) up to the expected year of death (both factual and counterfactual). The indexation follows the progressive scheme currently active in Italy: pensions are fully indexed to inflation up to approximately €2,100 gross per month (four times the minimum pension benefit), while there is only a partial inflation indexation in higher pension brackets (85% between 4 and 5 times the Minimum Pension Benefit (MPB); 53% between 5 and 6; 47% between 6 and 8; 37% between 8 and 10; 22% above 10 times the MPB).

3.4. Measuring redistribution by counterfactuals

Let $\bar{y}_{i,h,t} = (y_{i,h,t}, ...y_{i,h,t+e(h,t)})$ denote the vector indicating the stream of pension benefits paid to ith individual from cohort h retiring at time t with LE equal to $e(h,t)$, based on uniform LE estimates. Similarly, let $\hat{\bar{y}}_{i,h,t} = (y_{i,h,t}, ...y_{i,h,t+e(h,t,s)})$ denote the vector representing the stream of pension benefits paid to ith individual from cohort h retiring at time t with a LE equal to $e(h,t,s)$, where non-uniform LE estimates are used, with s

indicating the individual's SEP.

We define the factual and counterfactual PW for the ith as follows:

$$\begin{aligned} PW_{i,h,t} &= \sum_{j=0}^{e(h,t)} y_{i,h,t+j} \quad \forall\ i, \\ \widehat{PW}_{i,h,t} &= \sum_{j=0}^{e(h,t,s)} y_{i,h,t+j} \quad \forall\ i. \end{aligned} \tag{1}$$

Conditional on cohorts and years of retirement, we can derive the factual and counterfactual PW distributions, denoted as $PW_{h,t}$ and $\widehat{PW}_{h,t}$, respectively. In Appendix D, we also explore an alternative definition of PW, which is based on the unweighted aggregation of pension benefits over time *discounted* by projected death probabilities (e.g. Haan et al., 2020).

The monetary value of the gap between factual and counterfactual PW serves as a measure of the impact of uniform annuity pricing on each retiree. Furthermore, the inequality gap among the two PW distributions quantifies the redistribution (progressive or regressive, i.e. "perverse") induced by the application of a uniform annuity factor when LEs differ significantly across SEP.

To assess the redistribution induced by uniform annuity pricing for retirees from cohort h retiring in year t, we utilize the two definitions of PW in eq. (1). Let $L\left(PW_{h,t}\right)$ and $L^C\left(\widehat{PW}_{h,t}\right)$ represent, the Lorenz curve of the factual PW distribution and the Concentration curve of the counterfactual PW distribution ordered by factual PW, respectively. Specifically, the factual distribution is what one would observe under the hypothesis—currently in force—that life expectancy does not significantly vary by SEP. In contrast, the counterfactual distribution accounts for the systematic LE-gaps by SEP, as estimated in Section 3.1).

Drawing from the existing literature on fiscal progressivity measurement (Reynolds and Smolensky, 1977), the redistribution between the factual and the counterfactual distribution—which corresponds to the "unintended" redistribution induced by the pension scheme—can be quantified by the Reynolds-Smolensky index, as follows

$$RS_{h,t} = G\left(PW_{h,t}\right) - C\left(\widehat{PW}_{h,t}\right) \quad \forall\ h,t \tag{2}$$

where $G\left(\cdot\right)$ and $C\left(\cdot\right)$ indicate the Gini and the Concentration index respectively. The Reynolds-Smolensky (RS) index ranges from -1, indicating maximal perverse (poor-to-rich) redistribution, to +1, indicating maximal progressive (rich-to-poor) redistribution. A value of 0 indicates the counterfactual hypothesis has no effect on overall inequality, implying no redistribution.

From eq. (2), it is possible to derive semi-aggregate measures of redistribution by conditioning either on cohort h or on the year of retirement t. These measures are defined as:

$$\begin{aligned} RS_t &= G\left(PW_t\right) - C\left(\widehat{PW}_t\right) \quad \forall\ t, \\ RS_h &= G\left(PW_h\right) - C\left(\widehat{PW}_h\right) \quad \forall\ h \end{aligned} \tag{3}$$

where RS_t captures redistribution across all retirees in year t, and RS_h captures redistribution within cohort h. These semi-aggregate indicators may yield more stable and reliable estimates than the fully disaggregated measures in eq. (2), particularly in context with limited data availability or small sample size.

Finally, one can compute an aggregate measure of redistribution for the entire population, irrespective of retirement year or cohort:

$$RS = G\left(PW\right) - C\left(\widehat{PW}\right) \tag{4}$$

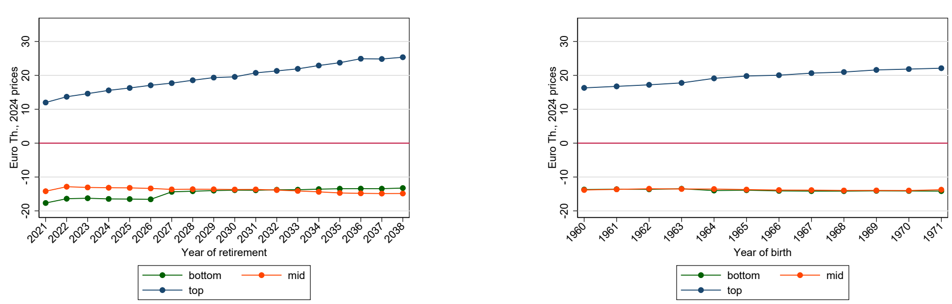
While this aggregate indicator provides a broad summary of overall redistribution, it is less directly applicable for targeted policy evaluation. Nonetheless, it can serve as a useful benchmark for interpreting cohort- and period-specific redistribution measures.

4. Results and Discussion

It is worth noting that the findings in this section are based on a highly homogeneous population consisting solely of private sector employees who meet the minimum eligibility requirements for a fully insured pension, thereby excluding any form of social assistance and minimizing the potential for substantial heterogeneity in outcomes.

Figure 3a and 3b report gains and losses in pension benefits—expressed in thousand euros at 2024 prices—arising from applying uniform versus non-uniform LEs, by year of retirement and by cohort, across SEP tertiles. As expected, retirees in the top SEP-tertile systematically gain from LE differentials. Remarkably, these gains increase over time, more than doubling from approximately €12,000 in 2021 to around €25,000 in 2038. In contrast, losses for the mid and bottom SEP-tertile are similar in magnitude and remain relatively stable over time, with the two curves largely overlapping.

Figure 3: Gains and losses (Th. Euros, prices 2024)



(a) By year of retirement

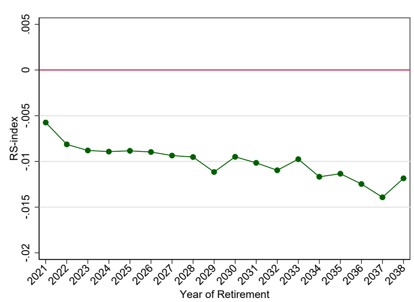
(b) By cohort

The similarity of losses between the two lower SEP-tertiles reflects two factors: first, their LE patterns evolve closely (see Figure 1b); second, the greater money value of contributions paid in the mid SEP-tertile with respect to the bottom SEP-tertile amplifies the money value of gains and losses in the LE gamble. For the same motivation, we observe a reduction in the money loss of the bottom SEP-tertile from 2026 to 2027, that is, when the average pension in the bottom SEP-tertile falls due to the massive entry of old-age retirees in 2027 (see Figure 2b).

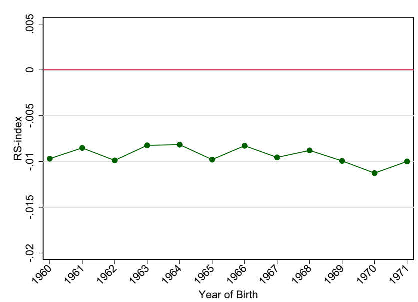
To assess the redistributive impact of LE-gaps, we first compute the RS-index for the entire population using eq. (4), pooling all cohorts (1960–1971) and retirement years (2021–2038). We find $RS = -0.011$, indicating the existence of perverse (i.e. poor-to-rich) redistribution. Once again, while this figure is relatively small, it is not negligible considering that our focus is on a highly homogeneous population. In addition, using only SEP-tertiles rather than finer stratification likely attenuates the full redistributive effect.[20]

Figure 4a and 4b display the RS-index by year of retirement and by cohort, respectively, following eq. (3).[21]

Figure 4: RS-index



(a) By year of retirement



(b) By cohort

When plotted by year of retirement (Figure 4a), the RS-index is consis-

[20]Moreover, the Gini index is known to exhibit little variability for a sufficiently large number of observations (Milanovic, 1997).

[21]Section D in Appendix, provides robustness checks using alternative definition of PW based on mortality discounting.

tently negative and shows a decreasing trend, with a sharper drop in 2028 and a partial rebound in 2038. This reflects two overlapping dynamics: an increasing share of workers retiring under the old-age pension rule, typically with lower seniority, and a growing prevalence of pensions calculated under the Dini's pro-rata benefit formula, which incorporate the less generous and less redistributive NDC component.

In Figure 4b, the RS-index is plotted by cohort. Here, the pattern remains negative but is flatter. This stability is attributable to the relatively uniform distribution of Dini's pro-rata pensions across years of birth[22], which tends to smooth redistributive effects over cohorts. A more granular disaggregation—simultaneously by year of birth and year of retirement—cannot be systematically performed, as numerous cells in the cross-classification contain either very few or no observations, preventing reliable or even feasible calculation of the RS-index in many combinations[23]

In summary, the persistence of negative RS values confirms that redistribution driven by longevity differentials is a systematic and structural feature of the Italian pension system. Notably, the downward trend in the RS-index observed in recent years suggests that, as longevity gaps widen and the transition from the DB to the NDC scheme progress, the continued use of uniform annuity factors will become increasingly inequitable. This evolution poses significant challenges for the future design and sustainability of pension policies.

4.1. Extensions

4.1.1. Disentangling redistributive drivers

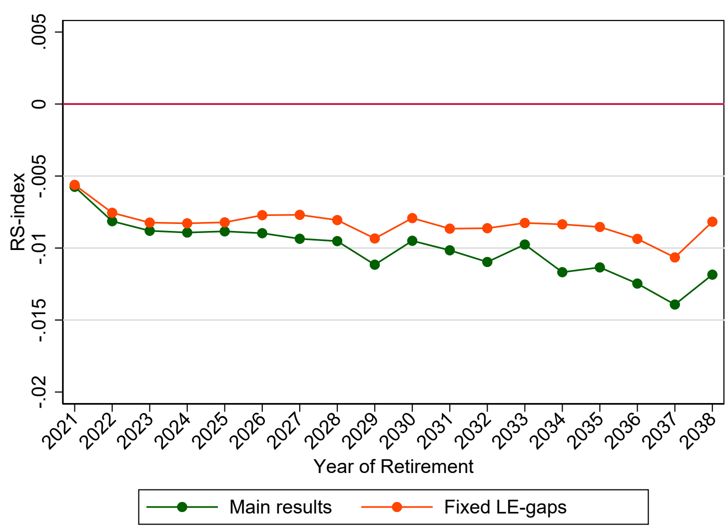
To isolate the effect of increasing LE-gaps from that of the progressive transition from DB to NDC schemes, we calculate the RS-index assuming

[22]Detailed cohort-level distributions are available upon request.

[23]Cross-tabulations reveal several empty cells and many others with insufficient sample size for reliable estimates. Tables are available upon request.

fixed LE at 2020 levels. Specifically, in Figure 5, we compare the baseline RS-index (as shown in Figure 4a) with a counterfactual scenario that incorporates projected increases in overall longevity while holding fixed the LE-gaps among SEP tertiles at their 2020 levels, thus abstracting from the widening inequality in LE depicted in Figure 1b. This counterfactual is obtained by applying the overall projected increase in LE over time while keeping constant the LE-gaps among SEP-tertiles observed in 2020. As expected, the extent of perverse redistribution is lower in the absence of the projected growing of LE disparities. Nevertheless, the DB-to-NDC transition continues to contribute positively to redistribution in favor of higher LE groups, and this effect increases slightly over time, except in 2038.

Figure 5: RS dynamics by year of retirement with fixed LE-gaps

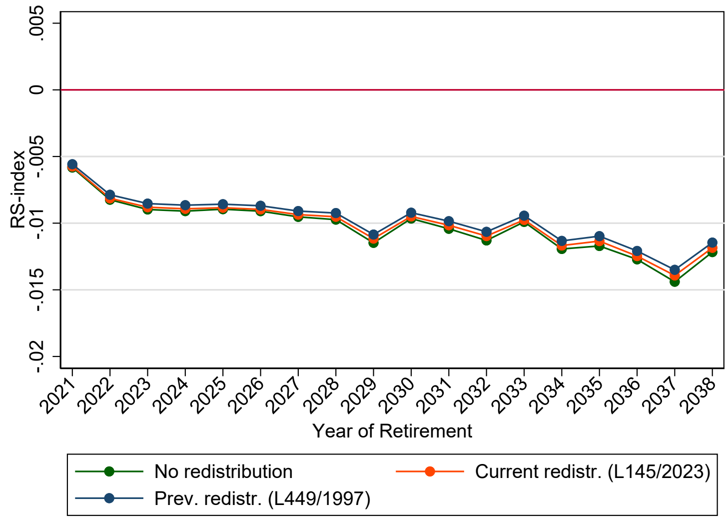


4.1.2. Inflation indexation

As discussed in Section 2.2, once the initial pension benefit is determined based on the applicable benefit formula, the Italian pension system indexes future benefits to inflation. This indexation mechanism is designed to be mildly progressive (see Section 3.3) which may, to some extent, mitigate the perverse redistribution effects stemming from LE-gaps.

Figure 6 shows the RS-index under a counterfactual scenario without any redistributive indexation. The two curves—baseline and non-redistributive—almost entirely overlap, indicating that the current indexation mechanism has only a minimal effect in reducing perverse redistribution.

Figure 6: RS-index with alternative indexation



To explore the potential role of a more redistributive indexation scheme, we also plot the RS-index assuming the reintroduction of the more progressive inflation indexation applied under Law 449/1997.[24]. While this alternative does reduce perverse redistribution stemming from LE-gaps, the overall impact remains modest. This suggests that even a more progressive indexation rule may be not an effective tool for mitigating the perverse redistribution caused by increasing LE-gaps.

[24]According to Law 449/1997, the system provided for a three-tier adjustment mechanism: (i) full adjustment (100%) for pensions up to five times the minimum INPS benefit; (ii) partial adjustment (30%) for pensions between five and eight times the minimum benefit; (iii) no adjustment for pensions over eight times the minimum benefit

4.1.3. Gains and losses: regional analysis

To further investigate the redistributive effects of LE-gaps, we disaggregate gains and losses by geographic area, focusing on the prevalent region of work. This territorial analysis is particularly relevant in the Italian context, where persistent North–South disparities characterize the socioeconomic landscape. Northern regions consistently exhibit higher levels of economic development, employment and wages compared to the Center and South (Felice, 2019). Life expectancy at birth and at age 65 is also higher in the North, reflecting more widespread access to healthcare and healthier living conditions (ISTAT, 2024).

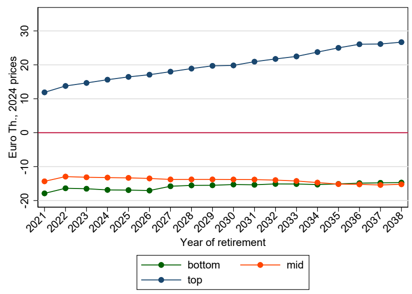
However, the pattern of inequality within Italian regions is complex. While income disparities are generally greater in the North—driven by broader wage dispersion and occupational stratification—regional inequalities in health outcomes depend on both the type of disease and the population subgroup under consideration (Petrelli et al., 2019). For instance, socioeconomic gaps in morbidity and mortality are more pronounced in the South when analyzing the general population (Petrelli et al., 2019), whereas among employed individuals, disparities in life expectancy tend to be wider in the North (Ardito et al., 2024). These structural asymmetries suggest that the redistributive implications of a life expectancy–based pension system may not be geographically neutral, but instead magnified by persistent territorial inequalities.

The results, presented in Figures 7a, 7b, 7c, show that gains for top SEP-tertiles are indeed greater in the North and lower in the South, consistent with the well-known socioeconomic gradient. This is because average earnings are higher in the North, leading to greater contributions and, consequently, larger gains and losses from the LE gamble in absolute terms. Notably, the increase in gains for the top SEP-tertile is steeper in northern regions.

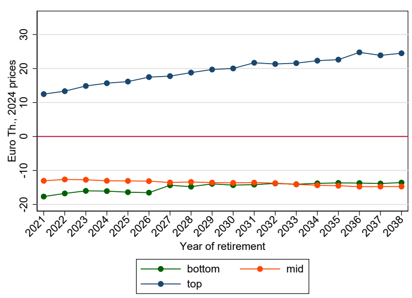
In contrast, income losses for individuals in the lower SEP-tertiles are remarkably similar across all regions. This pattern likely reflects a compressed earnings distribution at the bottom of the SEP scale and a less pronounced

North–South gradient in this group. Accordingly, the burden of perverse redistribution appears uniformly distributed among low earners, regardless of geography.

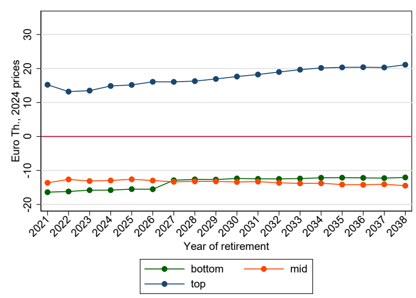
Figure 7: Gains and losses (Th. Euros, prices 2024)



(a) North-Italy



(b) Central-Italy



(c) South-Italy and Islands

Taken together, these findings suggest that the LE gamble may be perpetuating a cycle of inequality, where perverse redistribution penalizes poorer individuals evenly across Italy, but disproportionately rewards high-earning individuals in the North.

5. Concluding Remarks

Our analysis provides robust evidence of *perverse redistribution* in the Italian pension system, driven by the use of uniform annuity pricing. This finding is particularly striking in Italy, as it contradicts the Constitutional mandate of the public sector, which is instead systematically transferring financial resources from poorer to richer workers. Moreover, this result has far-reaching implicational for many other counties, as recent pension reforms worldwide have been strongly inspired by the ideal of actuarial fairness, while neglecting the impact of longevity differentials observed on real-life mortality rates.

Notably, our study demonstrates that LE-gaps generate perverse redistribution in Italy, even among a homogeneous population of private sector employees with stable working careers. Furthermore, we anticipate that this phenomenon will intensify in the near future, driven primarily by increasing LE-gaps over time. This trend is exacerbated by the gradual transition from the DB to the NDC scheme, which disproportionately penalizes the poorest segment of the population.

Our findings also indicate that the progressivity of the indexation mechanism is insufficient to offset perverse redistribution. Instead, this phenomenon is widespread across the country, with richer individuals in northern regions benefiting more from the higher accumulation of contributions on average.

While the increasing pattern of LE-gaps poses a significant political challenge, the possibility of *direct* intervention in pension schemes—such as introducing diversified annuity factors or retirement ages (Holzmann et al. 2019)—or *indirect* intervention—including measures like diversified retirement ages or redistributive accrual rates (Diamond 2004)—remains an open debate that falls beyond the scope of this paper.

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Appendix A Predictive models for the simulation of end-of-career earnings

For every individual in the sample we first simulate their probability of being employed using the estimates of a logistic model regressing the employment status (equal to 1 for those employed, equal to 0 for non employed), as a function of the age of the individual (with also a quadratic term), their age of entry in the labour market (age at first job), the total seniority (the sum of all years spent in employment during their work career), a measure of employment intensity in a lagged 4-years window (the sum of employment level in the years from t-4 to t-1) and their Socio Economic Position (SEP), as measured by their tertile of income at mid career (ages 45-49) and using the second tertile as a reference group (see Table 1). Note that the employment level used as a predictor is a continuous variable going from 0 (no employment) to 1 (full-year employment), since the administrative information available is very detailed (number of total worked weeks in the year). For the sake of , we dichotomize this information into the employment status used as the outcome variable and for the simulated careers, using 75% as a threshold to convert the continuous into a dichotomous variable: i.e., we set the employment status equal to 1 for those with an employment level in the year greater or equal to 0.75 (the threshold has been calibrated to maximize the accuracy of the average employment level prediction). As regard the estimates, the overall accuracy of the model is pretty high (95% concordant); the main predictor is the overall employment in the preceding 4 years window; as expected, those in a higher (lower) SEP have a higher (lower) probability of being employed (although we do not interpret these estimates in a causal way since the purpose of the model is strictly predictive).

Table 1: Logistic Regression Analysis

(a) Analysis of Maximum Likelihood Estimates

Parameter	DF	Estimate	Std. Error	Wald χ^2	Pr > χ^2
Intercept	1	-15.1114	1.9863	57.8765	<.0001
Age	1	0.5693	0.0673	71.6049	<.0001
Age_2	1	-0.00704	0.000586	144.3627	<.0001
Age at First Job	1	0.0534	0.00325	270.6444	<.0001
Total Seniority	1	-0.00572	0.00293	3.7941	0.0514
Employment Intensity	1	1.8313	0.0136	18121.0614	<.0001
SEP 1	1	-0.1039	0.0277	14.0545	0.0002
SEP 3	1	0.4412	0.0343	165.4678	<.0001

(b) Odds Ratio Estimates

Effect	Point Estimate	95% Confidence Limits: Lower	Upper
Age	1.767	1.549	2.016
Age_2	0.993	0.992	0.994
Total Seniority	0.994	0.989	1.000
Employment Intensity	6.242	6.078	6.411
Separation 1	0.901	0.854	0.952
Separation 3	1.555	1.454	1.663
Age at First Job	1.055	1.048	1.062

(c) Association of Predicted Probabilities and Observed Responses

Concordant Percentage	95.0
Discordant Percentage	5.0
Tied Percentage	0.0
Pairs	5138728932
Somers' D	0.900
Gamma	0.900
Tau-a	0.398
c	0.950

For those who are employed, we simulate their earnings using the estimates of a wage equation regressing the average yearly income in t on the average early income in a lagged 4-years window, total seniority (see above), employment intensity (see above), and a dummy identifying those who are re-entering into employment, i.e. those who were not employed in *t-1* (see Table 2). Indeed, lagged income is the main predictor in this model (every euro of average income is turned into 0.98 euros in time t), while no other predictors that have been tested besides those listed (such as age or age squared) add to the predictive power of the model, which is already pretty high (R-squared equal to almost 75%).

Table 2: Regression Results

(a) Model Fit Statistics

Statistic	Value
Root MSE	18491
R-squared	0.7450
Mean Dependent Variable	38556
Adjusted R-squared	0.7450
Coefficient of Variation	47.95967

(b) Parameter Estimates

Variable	DF	Estimate	Std. Error	t Value	Pr > $\|t\|$
Intercept	1	498.8443	482.93287	1.03	0.3016
Lagged Mean Wage	1	0.97672	0.00173	563.04	<0.0001
Total Seniority	1	-93.87127	8.87283	-10.58	<0.0001
Employment Intensity	1	961.45024	106.17915	9.05	<0.0001
Not Employed in t-1	1	-5101.92682	695.31951	-7.34	<0.0001

Appendix B Uniform vs non-uniform life expectancy

In our analysis we compute both uniform and non-uniform life expectancy: the former refers to life expectancy for the total population, while the latter accounts for heterogeneity by income, dividing individuals into tertiles based on SEP. To estimate mortality and life expectancy, we employ the LC model, applying it to age-specific mortality rates.

As concerns the computation of uniform life expectancy, we collect historical age-specific mortality data for both males and females, covering ages 20 to 90 from years 1990 to 2019, using data downloaded from the HMD.

We then fit the standard LC model to the selected dataset:

$$\ln\big(m_{x,t}\big) = \alpha_x + \beta_x \, \kappa_t + \varepsilon_{x,t}. \tag{5}$$

where $m_{x,t}$ is the mortality rate at age x in year t, α_x captures the average age-specific mortality pattern, β_x describes how mortality at each age responds to changes in the time-varying index κ_t, and $\varepsilon_{x,t}$ are the residuals. To ensure a unique solution, identifiability constraints such as $\sum_x \beta_x = 1$ and $\sum_t \kappa_t = 0$ are imposed.

Once the LC model has been fitted through Singular Value Decomposition, the subsequent step consists in projecting the time-varying index κ_t using a univariate ARIMA approach. The idea is that κ_t can be modelled as a stochastic process. To this end, we adopt the standard Box–Jenkins methodology, comprising identification, estimation, and diagnostic checking, to specify a suitable ARIMA(p, d, q) model for the mortality index κ_t. This approach enables a flexible time-series representation that captures both the stochastic trend and short-term dynamics in κ_t, thereby allowing for a coherent projection of future mortality patterns. In our empirical analysis, we rely on the `forecast` package (available on CRAN) and its `auto.arima()` procedure to determine the ARIMA parameters automatically. This approach often leads to a random walk with drift or a similar equivalent

$$\kappa_{t+1} = \kappa_t + c + \eta_t,$$

where c is a constant drift term and η_t is white noise, in line with the standard LC assumption of a random walk with drift. We then forecast κ_t over the horizon 2020–2038 using the chosen ARIMA model. These projected κ_t values are then substituted back into the LC equation to obtain forecasts of $m_{x,t}$. The next step is to convert the resulting rates into probabilities of death,

$$q_{x,t} = 1 - e^{-m_{x,t}},$$

and survival probabilities,

$$p_{x,t} = 1 - q_{x,t}.$$

Finally, assuming a fixed initial population, we construct life tables to compute life expectancies at various ages.

Similarly, to compute non-uniform life expectancy, we use the WHIP dataset, which aligns with the HMD in terms of age groups and time coverage. To assess the validity of the WHIP data, we compare it against the HMD, evaluating consistency and reliability across the two sources (see Fig.(1a)).

To account for systematic longevity differentials across socioeconomic groups, the same mortality analysis described previously was replicated by partitioning the WHIP dataset into SEP-tertiles based on age-specific weekly earnings. Specifically, the LC model was applied separately to each tertile to estimate mortality rates, survival probabilities, and life expectancies over time. The fitting period spanned the years from 1990 to 2019, while the forecasting window extended from 2020 to 2038, aligning with the projection horizon considered for pension benefit simulations in the main analysis. Mortality data for each SEP group were smoothed using five-year moving averages by age and calendar year, ensuring robustness in sparsely populated cells, particularly at the extremes of the age distribution. Time-varying mortality index forecasts were obtained via an ARIMA model fitted independently for each tertile and the projected κ_t values were used to estimate age-specific mortality rates. These were subsequently transformed into death probabilities and survival functions, from which period life expectancies at age 67 were derived.

Appendix C Biannual transformation coefficients

According to the Italian Legislation, the first pension benefit is calculated by multiplying the amount of accrued contribution by a Transformation Coefficient (TC_x) that reflects the worker's age x at retirement. This coefficient, periodically revised by the Ministry of Labor in cooperation with the Ministry of Economy and Finance, incorporates key demographic measures reported by ISTAT, such as probability of death, spousal survival probabilities and life-expectancy indicators.

TC_x (or rather its reciprocal, i.e., the Divisor) is expected to ensure the actuarial fairness between the amount of accrued contributions and the annuity paid to the retiree until the retiree or their eligible survivors pass away (depending on life expectancy).

Technically, let

$$y_x = (TC_x)(MC_x)$$

be the annual pension benefit for a worker retiring at age x, with MC_x being the total amount of accrued contributions at retirement age x, and

$$TC_x = \frac{1}{\Delta_x}$$

the Transformation Coefficient at retirement age x.

Given a specific retirement age x, a single divisor Δ_x is computed by averaging:

(i) $a_{x,s}^{v(t)}$: the present value of the direct pension itself (i.e., the benefit the worker receives personally);

(ii) $A_{x,s}^{v(t)}$: the present value of the survivor's pension (i.e., the portion transferred to the spouse, if eligible, after the pension holder's death).

Formally, the TC_x's computation merges men and women into a single demographic pool and employs unified mortality and remarriage data to determine a single divisor Δ_x for each permissible retirement age. From 2023 onward, these ages span from 57 to 71, as established by current law. Under existing legislation, the resulting TC_x's—computed at discrete retirement ages (e.g., 57, 58, ..., up to 71)—are published in ministerial decrees and revised every two years to incorporate updated LE estimates. More specifically, the Ministry of Labor, together with the Ministry of the Economy and Finance, is required to adopt the "coefficient tables" based on the latest ISTAT projections.

To represent these dynamics using actuarial notation, let s denote the gender of the individual (m or f, respectively male or female), and $\bar{s}$ the gender of the surviving spouse; let w be the ultimate age and $\varepsilon_{x,s}$ the gap between the retiree's age x and that of the spouse. The symbol $\frac{l_{x+t,s}}{l_{x,s}}$ indicates the probability of surviving from age x to $x+t$, while $q_{x+t,s}$ is the probability of dying between those same ages. The term $\Theta_{x+t,s}$ describes the probability of leaving the family at age $x+t$. Once the retiree dies, η (set to 0.6) is the survivor benefit rate reduced by an income-based means test δ_s, which is 0.9 for male retirees and 0.7 for female retirees. A discount rate of the form $\left(\frac{1+r}{1+\sigma}\right)^{-t}$ is applied, where $\frac{1+r}{1+\sigma} = 1.015$, and $k = 0.4615$ accounts for monthly payments made in advance. Finally, $l^{\text{ved}}_{x+t,s}$ is the probability that the surviving spouse either dies or remarries, affecting the present value of any continuation of pension benefits.

Within this framework, the divisor Δ_x combines the average present value of the worker's own pension, $a^{v(t)}_{x,s}$, and the corresponding present value of survivors' benefits, $A^{v(t)}_{x,s}$, for both male and female contributors:

$$\Delta_x = \frac{\sum_{s=m,f}\left(a^{v(t)}_{x,s} + A^{v(t)}_{x,s}\right)}{2} - k.$$

The direct pension value $a^{v(t)}_{x,s}$ is given by

$$a^{v(t)}_{x,s} = \sum_{t=0}^{w-x} \frac{l_{x+t,s}}{l_{x,s}} \left(\frac{1+r}{1+\sigma}\right)^{-t},$$

while the survivors' pension component $A^{v(t)}_{x,s}$ is

$$A^{v(t)}_{x,s} = \sum_{t=0}^{w-x} \frac{l_{x+t,s}}{l_{x,s}}\, q_{x+t,s} \left(\frac{1+r}{1+\sigma}\right)^{-t} \Theta_{x+t,s}\, \eta\, \delta_s \sum_{\tau=1}^{w-x-t+\varepsilon_{x,s}} \frac{l^{\text{ved}}_{x+t+\tau-\varepsilon_{x,s},\,\bar{s}}}{l^{\text{ved}}_{x+t+1-\varepsilon_{x,s},\,\bar{s}}} \left(\frac{1+r}{1+\sigma}\right)^{-\tau}.$$

Once Δ_x is determined, TC_x is simply its reciprocal, i.e. $\frac{1}{\Delta_x}$.

Official values of TC_x provided by ISTAT are currently available up to 2026. To extend these values beyond 2027, we project survival probabilities at retirement ages—ranging from 57 to 71 years under Italian legislation—using the Lee–Carter model (Lee and Carter, 1992), as detailed in Section B. In addition, for the biennia 2023–24 and 2025–26, we replaced the estimated TC_x values with the observed ones. This adjustment was introduced to mitigate the instability (spike) induced by the COVID-19 shock, particularly pronounced in the biennium 2023–24.

The computation of TC_x begins with the divisor Δ_x, from which we derive the expected present value of the direct pension $a_{x,s}^{v(t)}$, using the survival probabilities estimated in Appendix B. For $A_{x,s}^{v(t)}$—the expected present value of the survivor's pension—we adopt the values reported in the Ministerial Decree (January 2023), differentiated by gender. This choice is consistent with the fact that the first baseline year in the calculation of the transformation coefficients is 2022.

Table 3: Forecasted transformation coefficients

Age	**2023–24**	**2025–26**	**2027–28**	**2029–30**	**2031–32**	**2033–34**	**2035–36**	**2037–38**
57	4.203	4.172	4.142	4.113	4.085	4.058	4.031	4.006
58	4.308	4.275	4.244	4.214	4.184	4.156	4.128	4.101
59	4.419	4.385	4.352	4.320	4.289	4.259	4.230	4.202
60	4.537	4.501	4.467	4.433	4.401	4.369	4.339	4.309
61	4.662	4.625	4.589	4.553	4.519	4.486	4.454	4.423
62	4.796	4.757	4.719	4.682	4.646	4.611	4.577	4.544
63	4.939	4.898	4.857	4.818	4.780	4.744	4.708	4.673
64	5.092	5.048	5.005	4.964	4.924	4.886	4.848	4.812
65	5.255	5.209	5.164	5.121	5.079	5.038	4.998	4.960
66	5.430	5.382	5.334	5.289	5.244	5.201	5.159	5.118
67	5.618	5.567	5.517	5.468	5.421	5.376	5.332	5.289
68	5.822	5.768	5.715	5.663	5.614	5.565	5.519	5.473
69	6.042	5.984	5.928	5.874	5.821	5.770	5.720	5.672
70	6.280	6.219	6.160	6.102	6.046	5.992	5.939	5.888
71	6.538	6.473	6.410	6.349	6.290	6.232	6.176	6.122

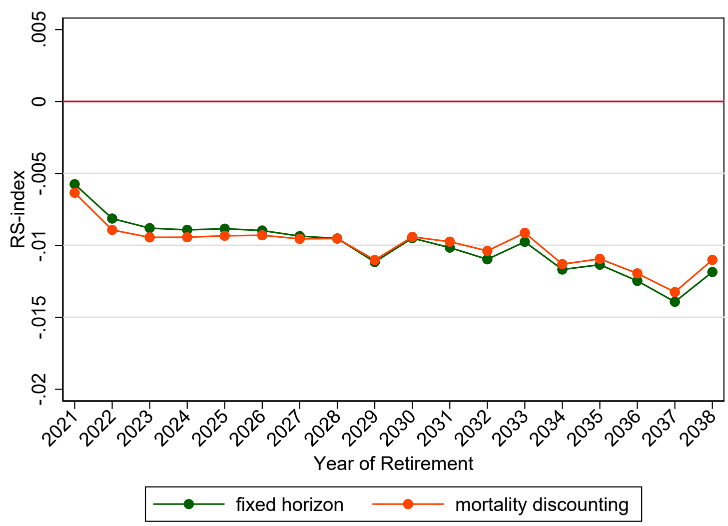
Appendix D Robustness check: MD-based PW

Let t_i be the year of retirement of ith individual, and let q_{i,t_i+k} and $\tilde{q}_{i,t_i+k}$ be the death probabilities of ith retiree at time $t_i + k$ according to, respectively, uniform and non-uniform mortality rates. We define the mortality-discounted pension wealth of ith retiree, $\widetilde{PW}_i$, by considering the probability of death in each year, that is,

$$\begin{aligned} \widetilde{PW}_i &= \sum_{k=0}^{M} q_{i,t_i+k} y_{i,t_i+k} \quad \forall\ i \\ \widehat{\widetilde{PW}}_i &= \sum_{k=0}^{M} \tilde{q}_{i,t_i+k} y_{i,t_i+k} \quad \forall\ i \end{aligned} \tag{6}$$

where M indicates the maximum survival age.

Figure 8: RS-index with MD-based PW



In Figure 8, we compare the RS-index by year of retirement as obtained from MD-based and LE-based PW distributions. The two patterns are found to be very similar proving that our results are robust to the definition of PW. A switching is observed in the 2029, which is what one may expect with MD-based estimates since workers in worse SEP-tertiles disappear gradually over time.