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Actuarial Fairness When Longevity Increases

An Evaluation of the Italian Pension System

Actuarial fairness when longevity increases: an evaluation of the Italian pension system

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Abstract

In this study, we analyse the actuarial features of the Italian pension system after the 1995 reform. We consider both the old defined benefit, the *pro-rata* and the new notional defined contribution pension rules applied to private sector employees born between 1945 and 2000. In the computations, we allow for dynamic mortality. To this aim, we project cohort- and gender-specific mortality rates based on a limit demographic scenario recently depicted by demographic experts. We compare findings for the current legislation with those from a *quasi*-actuarially fair scenario, where cohort- and gender-specific mortality rates are taken into account in the pension computation.

The old DB rules are extremely generous and offer strong incentives to early retirement. Due to dynamic efficiency, the new NDC scheme provides less than actuarially fair benefits. As a consequence of the rules adopted to compute coefficients used to convert the notionally accumulated sum at retirement into the annuity and to update them in response to increasing longevity, the NDC scheme is more than *quasi*-actuarially fair. Periodical *ex-post* adjustments of the coefficients generate big incentives to retire. The main cause for the actuarial unfairness embedded in the new Italian pension system is the use of cross-sectional mortality rates in the computation of conversion coefficients: retired cohorts will likely live longer than what accounted for in the computation of their pension benefits, since cohort effects in mortality are disregarded by the Italian law.

Keywords: social security, notional defined contribution pension systems, actuarial fairness, longevity, cohort-specific mortality forecasts.

JEL codes: H55, J11, J14.

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1 Introduction

In the last two decades, population aging and low economic growth have undermined the financial stability of several pay-as-you-go pension schemes. Some countries (including Italy, Latvia, Poland and Sweden) have reacted to the crises by replacing their defined benefit (DB) pension system with a notional defined contribution (NDC) system, keeping the previous pay-as-you-go financial architecture. An NDC scheme consists of an individual account system to which contributions are earmarked and interests notionally paid; at retirement, the notionally accumulated sum is converted into the pension taking into account life expectancy, i.e. incorporating actuarial fairness.

An ideal actuarially fair pension scheme is characterized by both actuarial fairness and actuarial fairness at margin. Actuarial fairness guarantees that, for each individual, the discounted sum of contributions paid during the working career is equal to the discounted sum of expected benefits. On the other hand, actuarial fairness at margin ensures that, for each individual, the discounted sum of expected benefits (net of expected contributions to pay in case of continued work) does not depend on the worker's retirement age (see e.g. Legros 2006, Gruber and Wise 1999).

An NDC scheme cannot guarantee the two above described "pure" actuarial conditions; however, it must satisfy the weaker conditions of *quasi*-actuarial fairness on average, and *quasi*-actuarial fairness at margin on average (Palmer 2006). These conditions first recognize that, in a dynamic efficient economy - i.e. an economy where the risk-adjusted return on assets is greater than the earnings growth rate - only a funded DC pension scheme can guarantee pure actuarial fairness. Since an NDC system is pay-as-you-go financed, its feasible (equilibrium) rate of return is approximately equal to the wage bill growth rate, i.e. the sum of earnings growth per head and the population growth (see Aaron 1966, Samuelson 1958). Moreover, these conditions allow for actuarial fairness (at margin) to only hold "on average", whereas in a pure actuarial system they must hold for each individual.¹

The implementation of a *quasi*-actuarially fair NDC system needs to handle increasing longevity, and to insure that life expectancy used to compute the annuity at retirement is as close as possible to actual, *ex-post*, residual life (Disney 2004).² Palmer (2006) has recently suggested three procedures to handle increasing longevity in NDC schemes. The first procedure consists in forming a committee of demographic experts in charge of the analysis of long-run scenarios and the publication of official cohort projections. Revisions would always apply to non-retired cohorts and would be increasingly small as the cohort approaches retirement age. The second procedure consists in estimating cohort-specific mortality tables based on historical cross-sectional survival data. In this case, periodical adjustments are necessary to avoid producing excessively generous, i.e. system-

¹ NDC systems should account for known differences in life expectancy related to individual characteristics such as gender, race (Sorlie, Rogot, Anderson, Johnson, and Backlund 1992), socioeconomic status (Kitagawa and Hauser 1973) and region (Caselli, Peracchi, Balbi, and Lipsi 2003).

² DG-ECFIN (2006) projections state that life expectancy at age 65 for the EU-25 countries will increase by about 4 years by 2050.

atically more-than-actuarially fair, annuities. The third procedure makes use of recent cross-sectional survival data to compute annuities at retirement; it regularly adjusts the pensions of all retirees on the basis of more updated statistics. According to Palmer (2006), the first and the third procedures have the advantage of explicitly fulfilling the criterion of financial equilibrium, whereas the second and the third are, in principle, less discretionary than the first since they are based on historical statistics.

In 1995, the Italian pension system was reformed, and an NDC scheme was introduced to replace the previous unsustainable DB scheme. The reform set up a transitional phase lasting almost thirty years toward the new rules. Therefore, in the coming two decades, pensions will still be computed according to the old rules (the so-called *pro-rata* system); consequently, their actuarial features will still depend on the characteristics of the old DB scheme. New DC pensions, on the other hand, will be computed on the basis of updated cross-sectional mortality tables, while benefits of retired individuals will be kept untouched regardless of longevity changes. From a comparison with the third procedure suggested by Palmer (2006), it is clear that the Italian NDC rules will systematically violate actuarial principles if longevity continues to rise.

Various studies have analysed the actuarial characteristics of the Italian pension system after the 1995 reform. Ferraresi and Fornero (2000) and Fornero and Castellino (2001) analyse a set of workers which is representative of future cohorts of retirees. Relying on a static mortality assumption, these scholars provide empirical evidence suggesting that the DB scheme is extremely generous, whereas the NDC is almost actuarially fair on average and at margin.³ While Ferraresi and Fornero (2000) and Fornero and Castellino (2001) focus on differences across cohorts, others (e.g. Caselli, Peracchi, Balbi, and Lipsi 2003, Borella and Coda 2006) look more within cohorts. Caselli, Peracchi, Balbi, and Lipsi (2003) find evidence of a sizable redistribution generated by the DC scheme which occurs across genders and Italian regions and which is induced by differential mortality. Borella and Coda (2006) develop a microsimulation model to study the redistributive impact of the Italian pension system both between and within cohorts; simulations are all based on cross-sectional mortality tables (ISTAT 2000).

In this study, we follow the representative agents approach proposed by Ferraresi and Fornero (2000) and Fornero and Castellino (2001). We analyse the actuarial features of the Italian pension system after the 1995 reform. We consider both the DB, the *pro-rata* and the (N)DC steady state rules applied to INPS-FPLD (i.e. private sector) employees. We compute two actuarial indicators, namely the present value ratio (PVR) and the implicit tax rate (TAX). Differently from previous studies, we allow for dynamic mortality in their computation. In particular, we investigate the actuarial properties of the rules designed by the 1995 reform to handle increasing longevity. To this aim, we build projected cohort- and gender-specific mortality tables based on a limit demographic scenario recently depicted by Robine, Crimmins, Horiuchi, and Zeng (2006). We first use the projected life tables in a baseline scenario (scenarios “B95” and “B07”, see section 4) to simulate pension benefits

³ These results are qualitatively confirmed by other studies (see e.g. Brugiavini and Peracchi 2004, Brugiavini and Peracchi 2003) which quantified the degree of actuarial fairness at margin of the Italian pension system in order to gauge its effects on retirement choices.

according to the current legislation. We then compare these results with alternative findings from a *quasi*-actuarial fair scenario (“AB”) where conversion coefficients are cohort- (and gender-)specific, as indicated by the first procedure in Palmer (2006). Although our analysis focuses on actuarial differences across cohorts, in section 7, we also provide some evidence of within cohorts redistribution (as in Caselli, Peracchi, Balbi, and Lipsi 2003) by distinguishing workers by gender and occupation.

The paper proceeds as follows. Section 2 illustrates the institutional framework, section 3 describes the projected cohort- and gender-specific mortality tables. Sections 4 and 5 describe the model and the actuarial indicators. Section 6 shows the main results, section 7 provides a sensitivity analysis with respect to some key model assumptions and section 8 concludes. Three appendixes describe (A) the methodology used to forecast mortality, (B) the formulas used to predict conversion coefficients for future retirees and (C) detailed simulation results.

2 Institutional framework

Until the 1990s, the Italian pension system was financially unsustainable. Pension expenditure grew from 7.4 % of GDP in 1970 to 14.9 % in 1992 (Brugiavini and Galasso 2004). To improve the social security budget, an impressive sequence of reforms was introduced during the 1990s. Most of them (such as the 1992 and 1997 reforms) were designed to be effective in the short run by acting on specific parameters of the existing system, such as the minimum retirement age. Law n.335/1995, which was designed instead to improve the budget in the long run, replaced the existing DB scheme with an (N)DC one. In line with its long run view, the law classified workers into three groups: the oldest, the middle-aged and the youngest. The oldest are those workers who, at the time of the reform, had accrued more than 18 years of seniority. They were left totally untouched by the reform and the old DB system is still applied to them. The middle-aged are those who had accrued less than 18 years of seniority by the end of 1995. Their pension is computed according to a mixed (*pro-rata*, PR henceforth) system where old and new rules are combined in proportion to the number of years worked prior to and after 1995. The youngest, to whom the new system fully applies, are those who started working after the 1995 reform.

In this study, we focus on private sector employees enrolled in the FPLD (*Fondo Pensioni Lavoratori Dipendenti*) fund. This fund enrolls almost all employees in the private sector; in 2008, it paid around 10 millions of pensions (INPS 2009). It is managed by INPS (*Istituto Nazionale della Previdenza Sociale*), the most important social security institution in Italy. The main retirement options for FPLD workers are the old-age and the seniority pensions. Their access has been progressively tightened by various reforms during the 1990s and the 2000s. Currently, male (female) FPLD workers can claim an old-age pension at age 65 (60) provided that they have accrued 20 years of seniority. They can claim a seniority pension at age 60, provided that 35 years of seniority have been accrued and only if the sum of seniority plus age is greater than or equal to 96 (called “quota 96”). Alternatively, they can claim a seniority pension if 40 years of seniority have been accumulated. In the future, DC benefits could be claimed when 5 years of senior-

ity plus a pension benefit equal to at least 1.2 times the social assistance benefit will be accumulated.⁴

The DB benefit is computed as the product of three factors: pensionable earnings, seniority and annual return. Pensionable earnings are the average wage of the last years of work. The number of years to include in the computation of pensionable earnings was progressively increased from 5 to 10 by the reforms of the 1990s. Seniority includes the number of years of regular contribution to the scheme, as well as years of notional contribution spent during out-of-work periods (e.g. unemployment spells, maternity leaves and military service); total seniority is topped at 40 years. Annual return is a decreasing function of pensionable earnings, equal to 2 % for a large part of the earnings distribution. Payroll tax rates grew dramatically in the last decades: from 19 % in 1967 (the first relevant year for our simulation, see section 4) to 33 %; they are paid for one third by the employee and for two thirds by the employer. Since 1992, pensions have been price-indexed.⁵

The DC pension for a worker retiring at age x is computed as:

$$P(x) = \left[c_a + \sum_{i=1}^{a-1} c_i \prod_{j=i}^{a-1} (1 + \bar{g}_j) \right] \delta_x \quad (1)$$

where c_i are the contributions paid when seniority is i , a is seniority at retirement and \bar{g}_j is the geometric mean of nominal GDP growth rate in the 5 years preceding the year in which seniority is j . The amount in squared brackets is the nominally accrued fund at retirement. δ_x is the conversion coefficient for retirement at age x , where $x \in [57, 65]$, defined as⁶:

$$\delta_x = \left(\frac{\sum_{s=m,f} dir_{x,s} + ind_{x,s}}{2} - k \right)^{-1} \quad (2a)$$

$$dir_{x,s} = \sum_{t=0}^{\Omega-x} \frac{\ell_{x+t,s}}{\ell_{x,s}} (1 + g_f)^{-t} \quad (2b)$$

$$ind_{x,s} = \theta \sum_{t=0}^{\Omega-x} \frac{\ell_{x+t,s}}{\ell_{x,s}} \left(1 - \frac{\ell_{x+t+1,s}}{\ell_{x+t,s}} \right) (1 + g_f)^{-(t+1)} a_{x+t+1}^W \quad (2c)$$

where $\frac{\ell_{x+t,s}}{\ell_{x,s}}$ is the gender- s -specific conditional survival probability at age $x + t$; g_f is the long-run expected GDP growth rate; a_{x+t+1}^W is the expected present value of a unitary annuity paid to the widow(er) at time $x + t + 1$; θ is the fraction of the annuity paid to the widow(er) and k is an actuarial adjustment factor that takes into account different frequencies in pension payments (set to 0.4615 by law to account for anticipated monthly payments). Notice that (equation 2a) differences in death probabilities between genders

⁴ The requirement on the minimum pension is not needed if retirement occurs at age 65.

⁵ A formal description of the DB and PR pension formulas applied to FPLD workers is provided by Fornero and Castellino (2001) and Ferraresi and Fornero (2000). Brugiavini and Galasso (2004) provide a comprehensive description of social security reforms in Italy.

⁶ Retirement at ages greater than 65 implies the application of δ_{65} and is thus highly discouraged. Retirement before age 57 is allowed only if 40 years of seniority are accrued; δ_{57} is applied in this case.

are averaged-out. Further computational details on the conversion coefficients are provided in Appendix B.

Conversion coefficients enclosed to the 1995 law incorporated the ISTAT (Italian National Institute of Statistics) 1990 mortality tables and a value for g_f equal to 1.5 %. The 1995 reform established that conversion coefficients had to be revised every ten years according to updated mortality tables and projected GDP growth rates. However, the first update of the coefficients was delayed: the new values - expected in 2005 - were legislated in 2007 (law n.247/2007) and applied in 2010. In addition to updating conversion coefficients, the 2007 law reduced their temporal validity from ten to three years. Current values, based on the ISTAT 2002 mortality tables and $g_f = 0.015$, are thus valid for the period 2010-2012.

3 Projected cohort-specific mortality tables

Starting from the 1970s, a sizable downward trend in old-age mortality has taken place in Italy.⁷ Between 1970 and 1990, male (females) life expectancy at age 60 (e_{60}) gained 2.4 (3.2) years (Caselli, Peracchi, Balbi, and Lipsi 2003) while, between 1990 and 2007, it earned an additional 3.4 (3) years (ISTAT 2010). In 2007, e_{60} was 21.89 years for males and 25.98 for females.

Positive past data on mortality has brought about an optimistic view of its future evolution. The greatest improvements are expected for the oldest-old, i.e. individuals aged 80 and over. The twenty-first century is considered by the experts as characterized not only by further progress in the prevention and cure of disease but also by firmly-rooted life styles able to promote more general “successful ageing”. In this context, younger cohorts are expected to live longer than older ones, i.e. the “cohort” effect (Caselli 1990) will continue to play its positive role.

In this study, we build projected life tables for the Italian cohorts who will face the transitional phase from DB to NDC pension rules (i.e cohorts 1945-2000). We rely on the theory of mortality expansion (Myers and Manton 1984, Olshanski, Carnes, and Cassel 1993) according to which the average life span will continue to rise in the future. In order to build “limit” life tables set in the distant future, we follow the indications provided by recent interdisciplinary studies which have depicted the most important traits for human survival (see Robine, Crimmins, Horiuchi, and Zeng 2006)⁸. We then connect current tables with limit mortality tables; therefore, in the first years of the projections, our tables reflect the characteristics of current mortality, while for farther projection years, our tables progressively come closer to those of the limit scenario. The projected period is extremely long (the limit scenario is set beyond year 2100) and thus our age-times-time matrix includes all the information related to the future process of mortality of the

⁷ In the 1970s, this phenomenon was a novelty for males, since their life expectancy at age 60 had only increased by 0.25 years in the previous forty years.

⁸ Robine, Crimmins, Horiuchi, and Zeng (2006) collects studies from biology, medicine, epidemiology, demography, sociology, and mathematics. We also refer to the proceedings of the conference “Health, ageing and work. Strategies for the new welfare society in the larger Europe”, The Geneva Association, Trieste, 2004.

Table 1. predicted period life expectancies at birth: our tables *versus* official tables

| Forecast/Year: | Males | | Females | |
|------------------------------------|-------|------|---------|------|
| | 2030 | 2050 | 2030 | 2050 |
| Our tables | 81.8 | 83.2 | 86.1 | 88.7 |
| ISTAT^a - central | 82.2 | 84.5 | 87.5 | 89.5 |
| ISTAT^b - low | 80.2 | 81.9 | 85.7 | 87.2 |

^a Source: ISTAT (2008)

^b Source: ISTAT (2008)

cohorts 1945-2000. Of course, for the oldest cohorts, the forecasts refer only to mortality at older ages, while for the youngest they cover their entire life. Consequently, projection uncertainty is higher for younger cohorts than for those approaching retirement in the coming years. Unfortunately, the method followed in the projections is deterministic and thus it does not allow for a quantification of projection uncertainty.⁹ Appendix A provides more methodological details.

Table 1 compares the estimated life expectancy at birth with the corresponding statistics provided by the most recent ISTAT projections (ISTAT 2008).¹⁰ The table indicates that our forecasts lie between ISTAT-low and ISTAT-central mortality scenarios. For instance, according to our tables, in 2030 male e_0 is expected to equal 81.8 years, while official statistics predict either 80.2 (low scenario) or 82.2 (central scenario). Figures 1 and 2 show the forecasted survival curves by cohort for females and males respectively. They highlight two typical characteristics of the mortality process of younger cohorts: the “rectangularization” of life tables (i.e. deaths are concentrated around a narrower interval of age) and the increase in life span. As a consequence of these improvements, younger cohorts are expected to experience sizable improvements in life expectancy: males e_{60} will increase from 23.1 years (cohort 1945) to 27.1 years (cohort 1970), up to a striking 31.4 years (cohort 2000). Corresponding values for females are 28.3, 31.6 and 38.1 years.

No official cohort life tables for the overall population exist in Italy. There exist two cohort life tables (called “RG48” and “IP55”, pertaining to the 1948 and 1955 cohorts) used by Italian insurance companies to compute premiums. They refer to the annuitant population and correct for self-selection. Due to self-selection of healthier individuals, the annuitants’ expected life expectancy is longer than the whole population’s.

⁹ Maccheroni and Barugola (2010) project mortality for the Italian birth cohorts 1950-2005 using a Lee Carter (i.e. stochastic) model. They show that uncertainty is pretty small if one concentrates on mortality at older ages, especially for cohorts born prior to 1975.

¹⁰ At an intermediate step, our approach produces a forecast for fictitious cohorts, i.e. calculated year by year, as in the case of most forecasts. The results obtained in this phase can thus be compared with those obtained by other recent official forecasts. It should be pointed out, however, that the methodological approaches used in a very long-term forecast (like those adopted for the cohort tables we provide) and in standard demographic forecasts (which usually cover a period of 30-50 years) are extremely different. Therefore, a formal comparison between the two tables cannot be conducted.

Figure 1. Projected survival curves by cohort: females

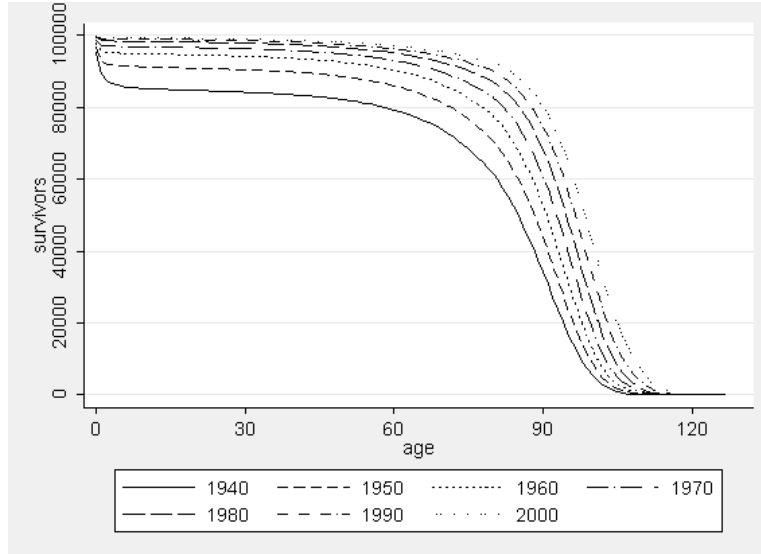
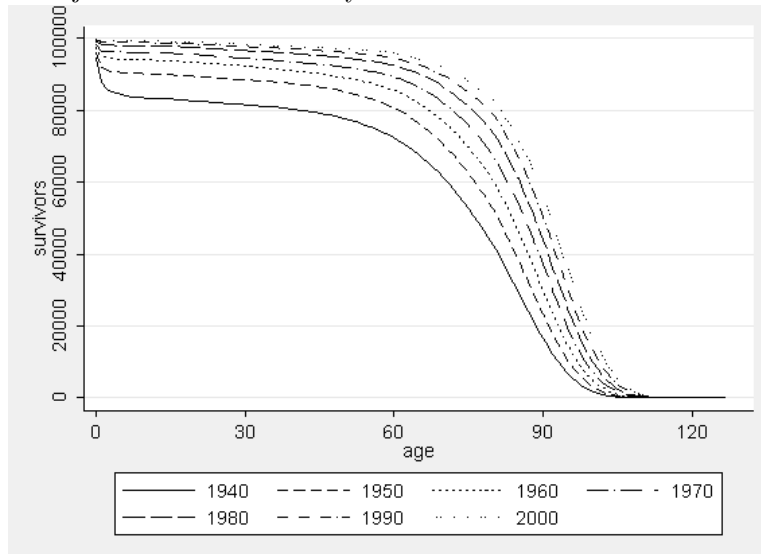


Figure 2. Projected survival curves by cohort: males



4 Representative agents

In this exercise, we evaluate the actuarial features of the Italian pension system for a set of representative individuals subject to DB, PR or (N)DC pension rules. Each analysed individual represents a typical INPS-FPLD (i.e. private sector) employee born in a given year, of a given gender and occupation (blue and white-collar). Agents are characterized by a stylized working career, described by an age of enrolling, a lifetime wage profile and a set of alternative retirement ages. The main features of the working career (e.g. age

of entry into the labour market) are kept constant across cohorts to better highlight the actuarial impact of both normative and mortality changes.

Based on the data provided in the Bank of Italy’s “Survey of Household’s Income and Wealth” (SHIW), our study assumes that white-collar workers enrol to the pension scheme at age 24, while blue-collar workers enrol at age 22 if female or at age 21 if males. We further assume that, once enrolled, agents keep contributing to the same scheme until retirement.¹¹ Retirement occurs between age 60 and 65. Given our assumptions, all agents become eligible to claim pension benefits at age 60 (see section 2).

Lifetime wage profiles are obtained as predictions from a random effects model for individual wages estimated on the Italian administrative data “Estratti Conto INPS”, a panel data set that covers the period 1985-1997 and includes 1/365 of the Italian private sector workforce. In the wage model, the log of wages is regressed against an age spline, a variable capturing the cohort effect, a set of year, sector and area of work dummies.¹² The model is estimated separately by gender and occupation (blue and white collars). Due to self-selection problems, the estimation samples only include individuals younger than age 60. Between ages 60 and 65, we assume that wages are constant. Figure 3 displays the estimated lifetime wage profiles: among the four groups, male white-collar workers have the steepest and highest profile while female blue-collar workers are the poorest and have the less dynamic careers.

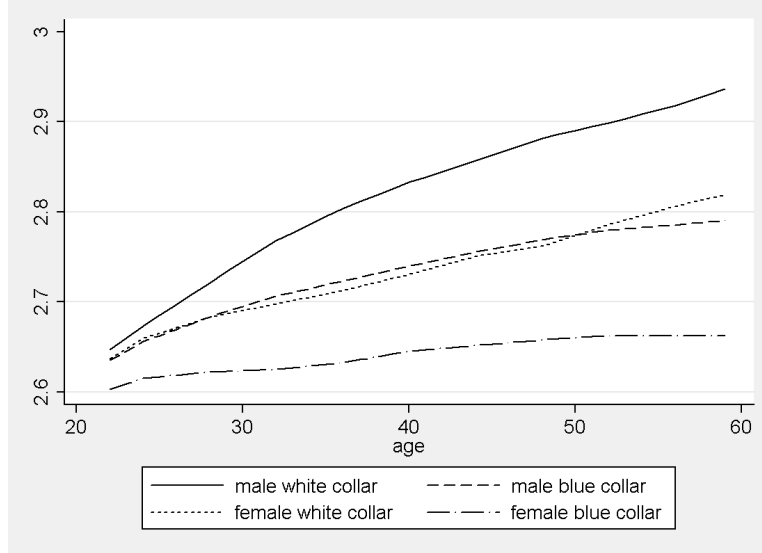
Given these stylized careers, each representative agent is assigned to a specific pension regime (DB, PR or NDC). Table 2, for example, reports assigned pension regime and retirement years by cohort for white-collar employees. It shows that agents born between 1945 and 1953 are subject to the DB rules, since in 1995 they had accumulated 18 years of contributions or more. Depending on retirement age and cohort, they claim pension benefits between year 2005 and 2018. Employees born between 1954 and 1970 are subject to the PR rules and retire between 2014 (cohort 1954, retiring at age 60) and 2035 (cohort 1970, retiring at age 65). Finally, employees born in 1971 or later are subject to the DC rules and retire in 2031 or later. A similar categorization of workers (not shown in the table) is obtained for male blue-collar and female blue-collar workers.

For each agent and retirement age between 60 and 65, we compute the accrued pension benefit and corresponding actuarial indicators (see section 5). In these computations, we consider both historical and projected values of key macroeconomic indicators such as GDP growth rates, inflation rates and long-run (10 years’ maturity) government bonds interest rate. Historical values are taken from various official sources: the OECD, the Bank

¹¹ Using the SHIW data (various cross-sections) we compute gender- and occupation-specific average ages of entry into the labour market. To avoid overestimation of accrued seniority at retirement, we account for an average period of 3 years spent out of the labour market and not compensated by notional contributions (see section 2).

¹² Given the long-run horizon of our analysis, we also need to predict wages for out-of-sample cohorts. Traditional econometric models, which use dummies to capture cohort effects, require additional assumptions for out-of-sample predictions. Following Heckman and Robb (1985) and Kapteyn, Alessie, and Lusardi (2005) we assume that wages differ across cohorts due to the macroeconomic conditions when individuals enter into the labour market. These conditions are summarized by productivity growth and are approximated by GDP per capita. Econometric details are provided upon request.

Figure 3. Estimated age-wage profiles by gender and occupation



Note: log-wages in €10,000 (2009 euros).

of Italy and ISTAT. In our main macroeconomic scenario, we assume a projected long-run real GDP growth rate (g_f , see section 2) equal to 1.5 % and a long-run riskless interest rate (r , see section 5) equal to 2 %. A very similar value for the GDP growth rate has been adopted in recent long-run EPC-WGA pension projections (EPC-WGA 2008) and in most of the Italian pension expenditure projections (MEF 2009, MEF 2011).¹³ The implied spread ($r - g_f = 0.5$ percentage points) is coherent with a dynamic efficient economy although it does not overestimate the dominance of the financial market with respect to the pay-as-you-go system. In a sensitivity analysis we consider alternative macroeconomic scenarios where the spread between the two key macroeconomic variables is wider.¹⁴

Actuarial fairness is evaluated in three different normative scenarios: “baseline 1995 (B95)”, “baseline 2007 (B07)” and “*quasi*-actuarial benchmark (AB)”. The first two incorporate actual pension rules. Scenario B95 considers the 1995 pension rules, whereas scenario B07 takes into account the reduced validity (from 10 to 3 years, as modified by the 2007 law) of the conversion coefficients. The AB scenario considers a hypothetical *quasi*-actuarial fair pension system, where conversion coefficients are cohort- and gender-specific and computed according to equation (B 2) in appendix B.¹⁵

¹³ See also MLSP (2002a) and MLSP (2002b). As already mentioned in section 2, a value of 1.5 % for the GDP growth rate has been used to compute conversion coefficients in 1995 and 2007.

¹⁴ Recent long-run EPC-WGA pension projections (EPC-WGA 2008) consider an interest rate equal to 3 %, which implies a wider spread $r - g_f$.

¹⁵ Consistently with current rules, in the B95 and B07 scenarios, conversion coefficients for the DC and PR benefits incorporate cross-sectional mortality rates, which we obtain by looking at the appropriate column of the gender-specific age-times-year projected life table (see equation (B 1)

Table 2. pension regime and retirement years by cohort: white-collar workers

| cohort | start ^a | sen.1995 ^b | regime ^c | retirement years: | | |
|--------|--------------------|-----------------------|---------------------|-------------------|-----|--------|
| | | | | age 60 | ... | age 65 |
| 1945 | 1969 | 26 | DB | 2005 | | 2010 |
| ... | ... | ... | ... | ... | | ... |
| 1953 | 1977 | 18 | DB | 2013 | | 2018 |
| 1954 | 1978 | 17 | PR | 2014 | | 2019 |
| ... | ... | ... | ... | ... | | ... |
| 1970 | 1994 | 1 | PR | 2030 | | 2035 |
| 1971 | 1995 | 0 | DC | 2031 | | 2036 |
| ... | ... | ... | ... | ... | | ... |
| 2000 | 2024 | 0 | DC | 2060 | | 2065 |

^a enrollment year.

^b seniority accrued at the end of 1995.

^c DB=defined benefit, PR=pro-rata, DC=notional defined contribution.

5 Actuarial indicators

We evaluate the actuarial characteristics of the Italian pension system by means of two social security money's worth measures (Geanakoplos, Mitchell, and Zeldes 2000): the present value ratio (PVR) and the implicit tax/subsidy rate (TAX). The former is used to evaluate actuarial fairness while the latter measures actuarial fairness at margin. Both of them require the computation of social security wealth (SSW). The SSW for retirement at a , computed at $t_1 \in [1, a + 1]$ and evaluated at t_2 , where a, t_1, t_2 define years of seniority, is given by:

$$SSW_{t_1, t_2}^a = - \sum_{j=t_1}^a c_j^* (1+r)^{t_2-j} + \left[P(e+a) \frac{1}{\bar{\delta}_{e+a, s}^{co}} \right] (1+r)^{t_2-(a+1)} \quad (3)$$

where c_j^* are the contributions at constant prices paid by the worker to the fund when accrued seniority is j and r is the time-constant riskless interest rate.¹⁶ $P(e+a)$ is the pension benefit associated with retirement at age $e+a$, where e is the age at which the employee starts contributing to the scheme and a is the number of years of seniority accrued at retirement. $\bar{\delta}_{e+a, s}^{co}$ is equal to $\delta_{e+a, s}^{co}$ as described in equation (B 2) but with g_f replaced by r . The sum in square brackets is therefore the present value of expected

and section 3). In the AB scenario, mortality rates are selected from the appropriate diagonal of the gender-specific age-times-year projected life table.

¹⁶ In the empirical analysis, we approximate r with the long-run (10 years' maturity) government bonds interest rate. See Queisser and Whitehouse (2006) for a discussion of the appropriate discount rate. Notice that (see e.g. Coile and Gruber 2000) we assume that the individual is alive at retirement.

pension benefits. It is computed assuming cohort- and gender- specific mortality rates, a discount rate equal to r , price-indexation of pensions and survivors' benefits. Financial flows are annual and anticipated.

The PVR is given by the ratio between the present value of expected pension benefits and the present value of the contributions paid during the working career. This indicator shows how much the system returns to the worker for each euro paid. The PVR computed when seniority at retirement is equal to a is given by:

$$PVR^a = \frac{P(e+a) \frac{1}{\delta_{e+a,s}^{\overline{co}}}}{\sum_{j=1}^a c_j^* (1+r)^{a+1-j}} \quad (4)$$

A pension system is defined as actuarially fair if $PVR^a = 1$ (i.e. $SSW_{1,1}^a = 0$ in equation 3). A pension system is defined as *quasi*-actuarially fair if $PVR^a = 1$ when $r = g_f$.

The TAX computed at $t_1 \in [1, a+1]$, once a' years of seniority have been accrued and evaluated at t_2 , is given by:

$$TAX_{t_1, t_2}^{a'} = \frac{-acc_{t_2}^{a'}}{E_{t_1} [w_{a'+1}] (1+r)^{t_2-(a'+1)}} \quad (5)$$

where the numerator is called *accrual* and is defined as:

$$\begin{aligned} acc_{t_2}^{a'} &= SSW_{t_1, t_2}^{a'+1} - SSW_{t_1, t_2}^{a'} \quad (6) \\ &= -c_{a'+1}^* (1+r)^{t_2-(a'+1)} + \\ &\quad \left[P(e+a'+1) \frac{1}{\delta_{e+a'+1,s}^{\overline{co}}} - P(e+a') \frac{1}{\delta_{e+a',s}^{\overline{co}}} (1+r) \right] \\ &\quad (1+r)^{t_2-(a'+2)} \end{aligned}$$

Equation (6) highlights that, if retirement is postponed by one year, the SSW varies due to two reasons. First, it decreases because of additional contributions to pay. Second, it varies due to the difference in square brackets between the present values of pension benefits associated with the alternative retirement options a' and $a'+1$. The sign of this difference is undefined because a shorter retirement period is generally associated with a higher pension benefit. In equation (5), the accrual is normalized with respect to the expected wage for the additional year of work.

A pension system is defined as actuarially fair at margin if $TAX_{t_1, t_2}^{a'} = 0$. A pension system is defined as *quasi*-actuarially fair at margin if $TAX_{t_1, t_2}^{a'} = 0$ when $r = g_f$ in (5). If $TAX_{t_1, t_2}^{a'} > 0$ (i.e. $Accr_{t_2}^{a'} < 0$) the pension system imposes an implicit taxation on the continuation of the working activity, thus providing financial incentives to early retirement; if $TAX_{t_1, t_2}^{a'} < 0$ the pension system penalizes early retirement.

6 Results

6.1 Forecasted conversion coefficients

Table 3 shows legislated and forecasted conversion coefficients by retirement age (57-65) and selected retirement years. In the first two columns, it reports values fixed by the 1995

Table 3. Legislated and forecasted conversion coefficients by age and retirement year: current legislation

| Age | Selected retirement years (<i>life tables</i>) | | | | | |
|-----|---|-------------------------------|----------------------------|----------------------------|----------------------------|----------------------------|
| | 1995-09 (<i>ISTAT90</i>) | 2010-12 (<i>ISTAT02</i>) | 2019-21 (<i>2018</i>) | 2028-30 (<i>2027</i>) | 2040-42 (<i>2039</i>) | 2049-51 (<i>2048</i>) |
| 57 | 4.720 | 4.419 | 4.169 | 4.029 | 3.866 | 3.793 |
| 58 | 4.860 | 4.538 | 4.276 | 4.128 | 3.956 | 3.879 |
| 59 | 5.006 | 4.664 | 4.390 | 4.234 | 4.052 | 3.971 |
| 60 | 5.163 | 4.798 | 4.511 | 4.345 | 4.153 | 4.067 |
| 61 | 5.334 | 4.940 | 4.640 | 4.464 | 4.261 | 4.170 |
| 62 | 5.514 | 5.093 | 4.777 | 4.590 | 4.375 | 4.279 |
| 63 | 5.706 | 5.257 | 4.922 | 4.724 | 4.496 | 4.394 |
| 64 | 5.911 | 5.432 | 5.077 | 4.867 | 4.625 | 4.517 |
| 65 | 6.136 | 5.620 | 5.244 | 5.020 | 4.762 | 4.648 |
| | percentage deviation with respect to 1995-09 | | | | | |
| 57 | - | -6.38 | -11.67 | -14.64 | -18.10 | -20.38 |
| 58 | - | -6.63 | -12.01 | -15.06 | -18.60 | -20.93 |
| 59 | - | -6.83 | -12.30 | -15.43 | -19.06 | -21.45 |
| 60 | - | -7.07 | -12.62 | -15.83 | -19.56 | -22.01 |
| 61 | - | -7.39 | -13.01 | -16.31 | -20.12 | -22.63 |
| 62 | - | -7.64 | -13.37 | -16.75 | -20.66 | -23.23 |
| 63 | - | -7.87 | -13.74 | -17.20 | -21.21 | -23.84 |
| 64 | - | -8.10 | -14.10 | -17.66 | -21.76 | -24.45 |
| 65 | - | -8.41 | -14.54 | -18.19 | -22.39 | -25.14 |

Note: percentage points; the first two columns report coefficients determined by laws n.335/95 and n.247/07; the following columns report forecasted coefficients (scenario B07) for selected triennial retirement periods. Coefficients are computed using forecasted cross-sectional life tables for the year indicated in italics at the top of each column.

and 2007 laws and valid respectively for the period 1995-2009 and 2010-12. In the following columns, the table shows forecasted coefficients (scenario B07) for selected retirement periods (2019-21, 2028-30, 2040-42, 2049-51). Conversion coefficients are expected to decrease considerably, as a consequence of the increased longevity. For example, the coefficients which will be applied in 2019-21 are expected to be 12-15 % lower than those legislated in 1995 (see lower part of the table). These findings are in line with those found by Caselli, Peracchi, Balbi, and Lipsi (2003) in their low-mortality scenario.¹⁷ The bottom part of table 3 shows that the relative reduction in the coefficients is higher for older retirement ages and concentrated in the first years of the simulation.

Table 4 reports the forecasted conversion coefficients according to the AB scenario for

¹⁷ Caselli, Peracchi, Balbi, and Lipsi (2003) predict conversion coefficients for 2020 using projected ISTAT life tables (base 2002, central and low-mortality scenarios, unpublished forecasts).

Table 4. Forecasted conversion coefficients by age, cohort and gender: AB scenario

| Age | Selected cohorts | | | | | |
|-----------|------------------|-------|-------|-------|-------|-------|
| | 1950 | 1960 | 1970 | 1980 | 1990 | 2000 |
| | Males | | | | | |
| 57 | 4.044 | 3.896 | 3.769 | 3.656 | 3.557 | 3.466 |
| 58 | 4.140 | 3.985 | 3.851 | 3.734 | 3.629 | 3.534 |
| 59 | 4.240 | 4.079 | 3.939 | 3.816 | 3.706 | 3.606 |
| 60 | 4.347 | 4.178 | 4.031 | 3.902 | 3.787 | 3.682 |
| 61 | 4.459 | 4.282 | 4.128 | 3.993 | 3.872 | 3.762 |
| 62 | 4.578 | 4.392 | 4.231 | 4.089 | 3.961 | 3.846 |
| 63 | 4.704 | 4.509 | 4.339 | 4.190 | 4.056 | 3.934 |
| 64 | 4.837 | 4.632 | 4.454 | 4.297 | 4.156 | 4.028 |
| 65 | 4.978 | 4.763 | 4.575 | 4.410 | 4.262 | 4.127 |
| | Females | | | | | |
| 57 | 3.925 | 3.767 | 3.627 | 3.498 | 3.375 | 3.255 |
| 58 | 4.014 | 3.848 | 3.701 | 3.566 | 3.437 | 3.312 |
| 59 | 4.108 | 3.933 | 3.779 | 3.637 | 3.503 | 3.371 |
| 60 | 4.208 | 4.024 | 3.862 | 3.713 | 3.571 | 3.433 |
| 61 | 4.313 | 4.119 | 3.949 | 3.792 | 3.643 | 3.499 |
| 62 | 4.425 | 4.220 | 4.041 | 3.875 | 3.719 | 3.568 |
| 63 | 4.543 | 4.327 | 4.137 | 3.964 | 3.800 | 3.640 |
| 64 | 4.668 | 4.440 | 4.240 | 4.057 | 3.884 | 3.716 |
| 65 | 4.802 | 4.560 | 4.348 | 4.155 | 3.973 | 3.797 |

selected cohorts. It highlights how heterogeneous the conversion coefficients would be, were they computed consistently with the cohort and gender differentials in life expectancy. For example, the coefficient applied to a male born in 1960 and retiring at age 60 ($\delta_{60,m}^{1960} = 4.178$) would be 3.9 % lower than that applied to a male retiring at the same age but born 10 years earlier ($\delta_{60,m}^{1950} = 4.347$). The coefficient applied to a male born in 2000 would be 15.3 % lower than that applied to an employee born in 1950 (cf. $\delta_{60,m}^{2000}$ with $\delta_{60,m}^{1950}$). Lower conversion coefficients would be applied to females, since they live longer than males; the provision of a joint annuity (see equation B 2, the indirect component) would only partially compensate the difference in life expectancy between genders in the main (direct) pension beneficiary.

Finally, table 5 shows - for each gender, retirement age and selected retirement year - the percentage deviation between coefficients computed according to the current legislation (scenario B07, table 3) and those computed according to the actuarial benchmark (AB scenario, table 4). For instance, the coefficient applied by law to a male employee retiring in 2010 at age 60 (4.798, see table 3) is 10.4 % higher than the coefficient computed according to the cohort- and gender-specific mortality rates ($\delta_{60,m}^{1950} = 4.347$, see table 4). The figures in the table are positive, meaning that the rules adopted by the 1995 reform

Table 5. Forecasted conversion coefficients: percentage deviation between current legislation (Table 3) and AB scenario (Table 4)

| Age | Selected retirement years | | | | |
|-----------|---------------------------|------|------|------|------|
| | 2010 | 2020 | 2030 | 2040 | 2050 |
| | <i>Males</i> | | | | |
| 57 | 10.5 | 8.1 | 7.9 | 6.6 | 6.5 |
| 58 | 10.5 | 8.1 | 7.9 | 6.6 | 6.5 |
| 59 | 10.4 | 8.0 | 7.8 | 6.5 | 6.7 |
| 60 | 10.4 | 8.0 | 7.8 | 6.4 | 6.3 |
| 61 | 10.3 | 7.9 | 7.8 | 6.4 | 6.3 |
| 62 | 10.3 | 7.9 | 7.7 | 6.3 | 6.2 |
| 63 | 10.2 | 7.8 | 7.7 | 6.2 | 6.1 |
| 64 | 10.2 | 7.8 | 7.6 | 6.1 | 6.1 |
| 65 | 10.2 | 7.7 | 7.6 | 6.1 | 6.0 |
| | <i>Females</i> | | | | |
| 57 | 14.1 | 12.0 | 12.3 | 11.7 | 12.5 |
| 58 | 14.1 | 12.0 | 12.4 | 11.8 | 12.6 |
| 59 | 14.1 | 12.1 | 12.5 | 11.8 | 12.7 |
| 60 | 14.0 | 12.1 | 12.5 | 11.9 | 12.8 |
| 61 | 14.1 | 12.1 | 12.6 | 11.9 | 12.8 |
| 62 | 14.1 | 12.2 | 12.6 | 12.0 | 12.9 |
| 63 | 14.1 | 12.2 | 12.7 | 12.0 | 12.9 |
| 64 | 14.1 | 12.2 | 12.7 | 12.0 | 13.0 |
| 65 | 14.1 | 12.2 | 12.8 | 12.0 | 13.0 |

to compute conversion coefficients - which rely on cross-sectional life tables and disregard cohort improvements in longevity - provide pension benefits which are more generous than what the benchmark suggests. This “premium” is lower for males than for females, since it is partly compensated (see equation 2a) by the application of unisex conversion coefficients averaging-out differences in mortality rates between genders.

6.2 Actuarial fairness

The PVR by cohort for a typical male white-collar worker retiring at age 60 is reported in figure 4. As specified in section 4, we consider three alternative scenarios (B07, B95 and AB), we assume a long-run GDP growth rate equal to 1.5 % and an risk-less interest rate equal to 2 %. Results for other agents, retirement ages and macroeconomic assumptions are reported in section 7.

Figure 4 highlights that the pre-reform system is extremely generous, providing much more than actuarially fair benefits. Cohorts belonging to the DB scheme (1945-1953), in fact, receive in terms of social security benefits up to 150 % of what they paid during

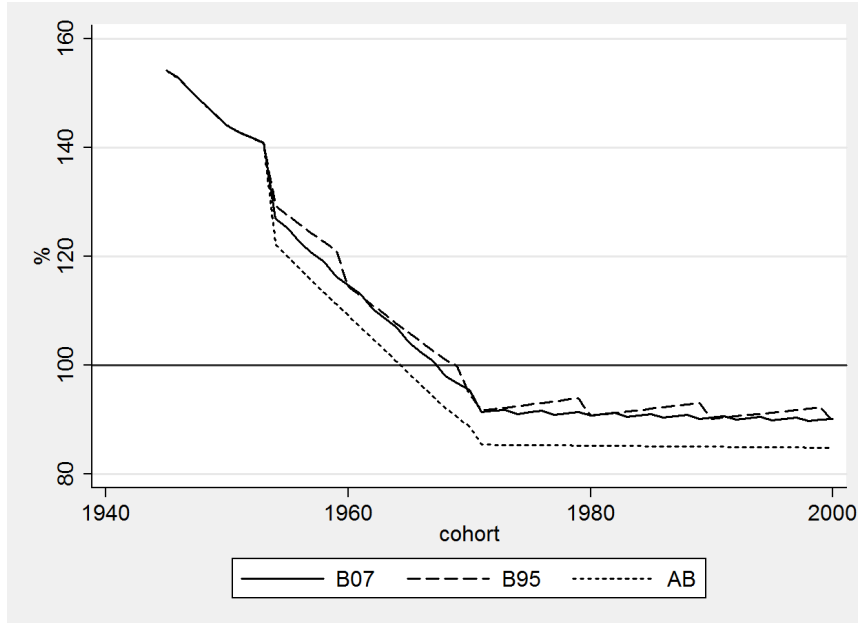
their working career. The PVR sharply falls (-14 percentage points, from 141 to 127 %) between the 1953 and 1954 cohorts, as a consequence of the classification of workers according to either the DB (cohort 1953) or the PR (cohort 1954) rules (see Ferraresi and Fornero 2000, Fornero and Castellino 2001). Throughout the transitional phase toward the DC rules (cohorts 1954-1970) the PVR exhibits a decreasing trend due to the increasing weight of the DC quota in the computation of pension benefits of younger cohorts. For cohorts younger than 1967, the system is less than actuarially fair. For cohorts born in 1971 and later, the indicator ranges between 89.7 and 91 %, indicating that the steady state DC regime is less than actuarially fair.

Results for the steady state can be attributed to dynamic efficiency (see Fornero and Castellino 2001): in an economy where the risk-less rate of return from financial markets dominates its rate of growth, only funded DC schemes can guarantee “pure” actuarial fairness. Regardless of macroeconomic conditions, a comparison across scenarios shows that the Italian pension system does not meet the weaker requirement of *quasi*-actuarial fairness on average (Palmer 2006). The PVR in the B07 and B95 scenarios is higher (by 5-6.5 percentage points) than in the AB scenario, meaning that the steady state DC rules are more than *quasi*-actuarially fair. This result depends on the rules adopted for the computation of conversion coefficients: since cohort effects in mortality are disregarded, retired cohorts will likely live longer than what accounted for in the computation of their pension benefit.

Finally, figure 4 shows that the PVR is characterized by an irregular course in the B07 and B95 scenarios. Periodical adjustments of the conversion coefficients, implemented to *ex-post* counteract increased longevity, generate discontinuities in the pension treatment of adjacent cohorts. The 2007 reform, by reducing from 10 to 3 years the temporal validity of the conversion coefficients, was quite effective in reducing redistribution across adjacent cohorts (cf. discontinuities in the B07 and B95 scenarios).¹⁸ The PVR has a smooth course in the AB scenario, where cohort- and gender-specific longevity increases are fully offset by corresponding reductions in conversion coefficients. For this reason, this simulation is similar to a kind of “static mortality case”. Accordingly, our results for the AB scenario are very similar to those reported in previous studies (Ferraresi and Fornero 2000) which assume static mortality.

¹⁸ Consider, for example, cohorts 1971 to 1974. A male white-collar worker born in 1971 is aged 60 in 2031. Under the current rules (B07), his PVR is equal to 91.4 %. A worker born in 1972 (1973) is expected to live slightly longer than one born in 1971, while his pension benefits will be computed using the same conversion coefficient of the previous cohort. As a consequence, he will enjoy a slightly higher PVR=91.6 (91.9). A worker born in 1974 and retiring at age 60 will be hit by the revision of the coefficients fixed by law in 2034 (NPVR=91.1). However, the reduction in the PVR suffered by the 1974 cohort with respect to the 1973 cohort is small, i.e. -0.8 percentage points. According to the 1995 legislation (coefficients revised in 2040), the effect of longevity improvements for 9 years combined with less frequent revisions of the coefficients would generate a much more pronounced cycle in the indicator: the PVR would increase by 2.1 percentage points (from 91.8 to 93.9 %) between the 1971 and 1979 cohorts and decrease by 3.2 points between the 1979 and 1980 cohorts due to the coefficients revision.

Figure 4. PVR by cohort at age 60: male white-collar workers, alternative scenarios



Note: percentage points; B07: current legislation; B95: conversion coefficients revised every ten years; AB: *quasi*-actuarial benchmark; 36 years of seniority; $g_f = 0.015$, $r = 0.02$

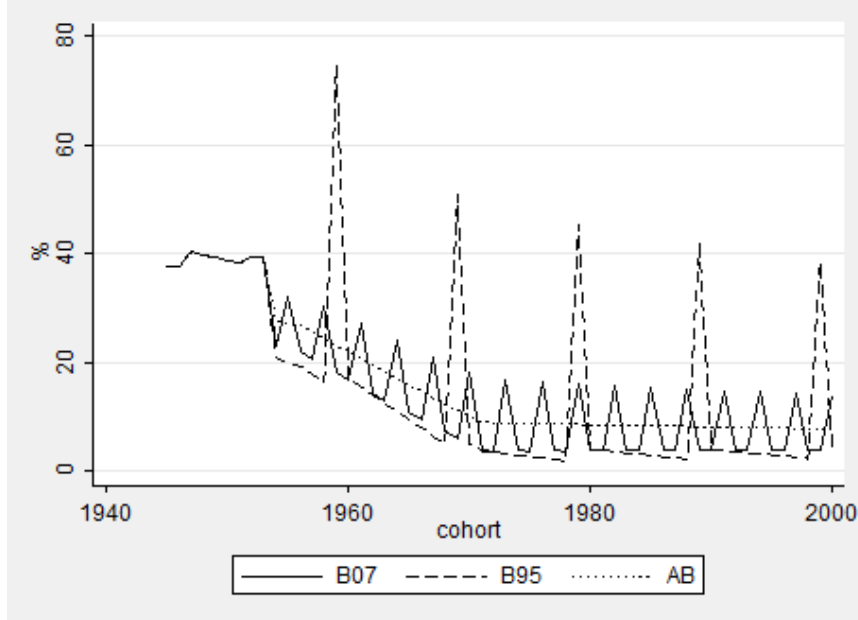
6.3 Actuarial fairness at margin

Figure 5 shows the TAX by cohort for male white-collar workers retiring at age 60. In the DB scheme the indicator reaches a value of 40 %, which represents a strong financial incentive to claim pension benefits at minimum requirements (see e.g. Belloni and Alessie 2009). The application of PR rules and the full application of the DC formula drastically reduce the implicit taxation of continuing working. However, the taxation is not eliminated in the steady state, which is less than actuarially fair at margin for most of the cohorts: excluding retirement years in which conversion coefficients are updated (see next paragraph), the indicator ranges between 3.8 and 4.1 %. Due to dynamic efficiency, even in the DC scheme, it is optimal to retire at minimum requirements and invest the accumulated wealth at the market return $r > g_f$.

Figure 5 (B07) points out the existence of big spikes of TAX (equal to about 11-12 percentage points) for the years in which conversion coefficients are updated. An individual considering staying at work in the year before the revision of the conversion coefficients would face a very strong constraint: his social security wealth would be considerably reduced if he kept working.¹⁹ The 2007 reform, by restricting the time validity of the

¹⁹ Compare, for example, workers born in 1972 and 1973. The former are aged 60 in 2032. They face a TAX equal to 3.8 % ($\delta_{60} = 4.295$ in 2032, they expect $\delta_{61} = 4.41$ in 2033). Workers born in 1973 face the revision of the coefficients when they reach age 60 and are taxed at a rate equal to 16.8 % if they postpone retirement by one year (they expect δ_{61} to fall from 4.41 to 4.36 due to the revision in 2034).

Figure 5. TAX by cohort at age 60: males white-collar workers, alternative scenarios



Note: percentage points; B07: current legislation; B95: conversion coefficients revised every ten years; AB: *quasi*-actuarial benchmark; 36 years of seniority; $g_f = 0.015$, $r = 0.02$

conversion coefficients to three years, reduced the implicit taxation in a sizable way (the B95 scenario shows spikes equal to 38-45 percentage points, cf. spikes in the B07 scenario).

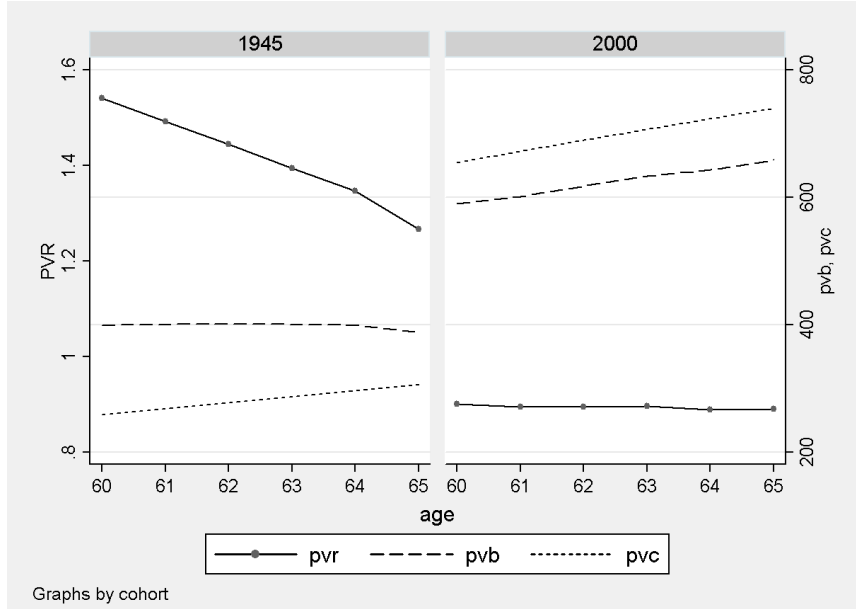
A comparison across scenarios shows that the Italian NDC system is more than *quasi*-actuarially fair at margin, meaning that it generates an implicit tax which is lower than that determined by the spread $r - g_f$. Indeed, in the B07 scenario, the TAX is 4-5 percentage points lower than in the AB scenario (excluding the spikes). This “extra-return” embedded in the current rules is due to the upward-biased difference between the conversion coefficients applied to adjacent retirement ages, which reflects the application of cross-sectional mortality rates in their construction. In the AB scenario, the TAX has a smooth trend since, in this simulation, longevity differences by cohort are compensated by appropriate differences in conversion coefficients applied to different cohorts.

7 Sensitivity

In this section, we provide a sensitivity analysis of the main results illustrated in the previous section comparing findings for alternative retirement ages (60-65), different agents (male blue-collar workers and females) and different macroeconomic settings. The full set of results for the base macroeconomic case is reported in Appendix C.

In the DB scheme, results show a negative relationship between the PVR and the retirement age whereas in the DC scheme the indicator is almost invariant with respect to retirement age. To better understand these findings, in Figure 6 we report the PVR as well as its components - present value of benefits (PVB) and present value of contributions

Figure 6. PVR, PVB and PVC for alternative retirement ages, cohorts 1945 (DB) and 2000 (DC)



Note: PVR left axis; pvb and pvc=1,000 euros, prices 2009, right axis; male white-collar workers; current legislation (B07); $g_f = 0.015$, $r = 0.02$

(PVC), see equation 4 - for two selected cohorts (1945, DB and 2000, DC) retiring at ages 60-65. For the DB cohort, the negative impact on the PVB of the shorter retirement period associated with later retirement dominates over the positive effect related to the modest increase in the pension benefit (2 % for each additional working year up to 40 years of seniority, plus the increase in pensionable earnings). Therefore, both the PVB and the PVR are lower for later retirement ages. For the DC cohort, the (actuarially-related) increase in the pension benefit associated with later retirement is more pronounced than in the DB case, and the consequent increase in the PVB almost perfectly compensates the increase in the PVC.²⁰ The TAX rate (see tables in Appendix C), on the contrary, tends to increase with the age of retirement in the DB scheme, thus providing increasingly stronger incentives to retire. The indicator becomes especially high when the individual accrues 40 years of seniority. In this case, in fact, there is a no annual return for the additional years of work. The TAX is instead almost constant with respect to the retirement age in the DC scheme.

²⁰ A lower PVR associated with older retirement ages does not allow us to infer that there is an incentive to retire (and viceversa). For this kind of considerations, we have to refer to the TAX. As already explained in section 5 (see equation 5) both the SSW associated with immediate retirement and the expected SSW associated with retirement in one year are evaluated at the same point in time in the TAX. When we compare PVR associated with different retirement ages, we instead refer to computations made at different points in time and based on currently accrued pension benefits.

As regards the comparison across genders and occupations, it turns out that the 1995 reform changed the implicit redistribution of the Italian pension system from “wage-based” to “mortality-based”. In the DB scheme, the actuarial indicators are particularly affected by the shape of the lifetime wage profile. Figure 7 reports the PVR by cohort for retirement at age 60, for four types of agents: male blue-collar, male white-collar, female blue-collar and female white-collar workers. In the DB scheme, the flattest age-wage profiles are typically associated with lower PVR. If, for instance, we consider the 1945 cohort, the indicator is equal to 154 % for male white-collar workers (steepest profile) and to 126.5 % for female blue-collar workers (flattest profile). The flattest lifetime wage profiles also typically correspond to a higher implicit taxation: the corresponding TAX for the above reported examples are equal to 37.7 (male white-collar) and 64.3 % (female blue-collar). These results are explained by the DB pension formula, which accounts for the wages in the last part of the working career only. By postponing retirement, the implicit taxation is then particularly high for a flat age-wage profile, since in this case pensionable earnings do not increase. In the (N)DC scheme, the indicators are affected by group-specific (in our exercise, gender-specific) mortality rates. Figure 7 illustrates that the PVR is approximately 90 % for males and around 95 % for females. Females are favoured, since their longer life expectancy is not accounted for in the computation of conversion coefficients.²¹

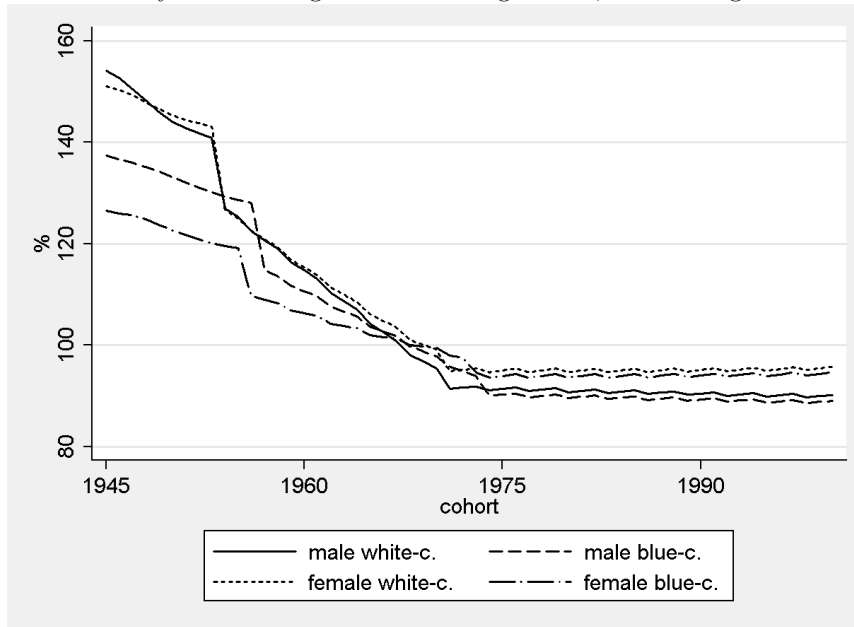
Finally, we report the results of a sensitivity analysis with respect to the macroeconomic scenario. Note that the sign of $\Delta\text{PVR}/\Delta i$ and $\Delta\text{TAX}/\Delta i$, $i = \{r, g_f\}$, in both the DB and the DC scheme is known *a priori* (see section 5) and is reported in the upper part of table 6 (“Qualitative impact”). While r affects the results in both the DB and the DC scheme, g_f affects them only in the DC scheme. To provide a quantification of these effects, we simulate the PVR and the TAX under a number of different values for g_f and r . We restrict our analysis to the sets $r \in [0.01, 0.03]$ and $g_f \in [0.005, 0.025]$, s.t. $r > g_f$. We consider two “extreme” cohorts: 1945 (DB) and 2000 (DC).

Simulated results for male white-collar workers retiring at age 60 are illustrated in detail in figure 8 and summarized in the bottom part of table 6. Each panel in figure 8 shows the value assumed by a given indicator and for a given cohort, as a function of the two macro variables. Each (iso-)line displays the set $\{g_f, r\}$ such that the indicator has the same value; each line refers to a decile of the simulated PVR or TAX distribution: lighter lines are associated with a higher value of the indicator. In the bottom part of table 6 (“Simulation results”), we report the corresponding 25th, 50th and 75th percentiles of the simulated PVR and TAX distributions.

Results in figure 8 are fully consistent with the expected effects described above. We refer to the upper part of table 6 for their interpretation. The simulated PVR (TAX) tends

²¹ In the DC scheme, the TAX is slightly negative for females (see appendix C), while it is slightly positive for males. As for males, two effects counteract: dynamic efficiency, which tends to increase the tax, and the upward-biased difference in the conversion coefficients applied to adjacent retirement ages, which leads to a reduction in the TAX (the “extra return” described in 6.3). As compared to males, females benefit from favourable gender-neutral conversion coefficients. Therefore, the latter effect is stronger for females, which explains the difference in the sign of the TAX between the two genders.

Figure 7. PVR by cohort at age 60: current legislation, different agents



Note: scenario B07; $g_f = 0.015$, $r = 0.02$

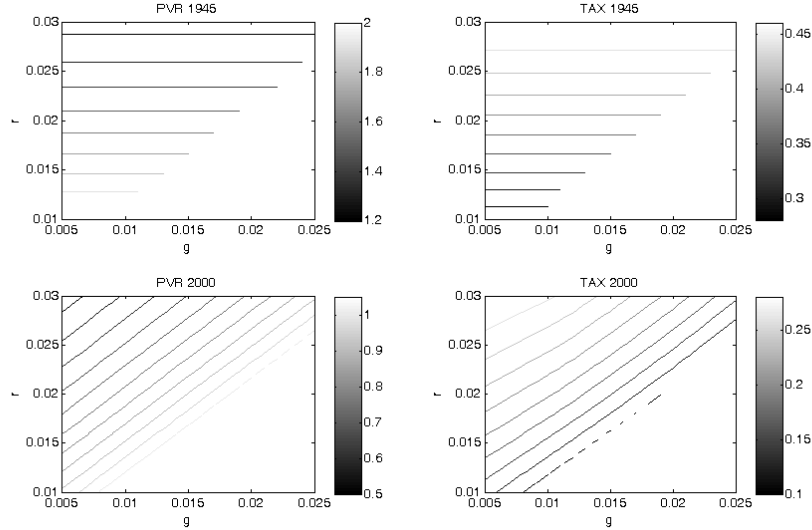
to be lower (higher) than in the base case, since in most cases, the spread $r - g_f$ is greater than 0.5 percentage points.²² Empirical evidence shows that the interquartile range turns out to be quite restricted, especially for the TAX (e.g. $43.9 - 35.5 = 8.4$ percentage points for the 1945 cohort). The reference thresholds for the two indicators, i.e. 1 for the PVR and 0 for the TAX, are almost never “crossed” in either the DB or the DC scheme. Overall, the main conclusions derived for the base case do not seem to change dramatically for a wide range of macroeconomic conditions.

²² In the base case ($g_f = 0.015$ and $r = 0.02$), PVR is equal to 154.1 for the 1945 cohort and to 90.2 for the 2000 cohort, while TAX is equal to 37.5 and 14.1 respectively (see Appendix C).

Table 6. Sensitivity analysis with respect to g_f and r : qualitative impact and simulation results

| | | Cohort (scheme) | | | |
|--------------|--|----------------------------|--------------|--------------|--------------|
| | | 1945 (DB) | | 2000 (DC) | |
| | | <i>Qualitative impact:</i> | | | |
| | | Δ PVR | Δ TAX | Δ PVR | Δ TAX |
| Δr | | - | + | - | + |
| Δg_f | | 0 | 0 | + | - |
| | | <i>Simulation results:</i> | | | |
| percentiles: | | PVR | TAX | PVR | TAX |
| 25 | | 126.2 | 35.5 | 69.7 | 13.3 |
| 50 | | 141.4 | 40.3 | 80.4 | 17.3 |
| 75 | | 163.2 | 43.9 | 92.3 | 20.9 |

Note: percentage points; male white-collar workers retiring at age 60 in the B07 scenario; $r \in [0.01, 0.03]$, $g_f \in [0.005, 0.025]$, $r > g_f$, 21×21 simulations

Figure 8. Sensitivity analysis with respect to g_f and r : isolines of PVR and TAX for the 1945 and 2000 cohorts

Note: male white-collar workers retiring at age 60 in the B07 scenario; $r \in [0.01, 0.03]$, $g_f \in [0.005, 0.025]$, $r > g_f$, 21×21 values; each (iso-)line display the set $\{g_f, r\}$ such that the indicator gets the same value; each line refers to a decile of the simulated PVR or TAX distribution; lighter lines are associated with an higher value of the indicator.

8 Conclusions

In this study, we analysed the actuarial features of the Italian pension system after the 1995 reform. We considered both the pre-reform DB, the *pro-rata* and the new (N)DC rules applied to INPS-FPLD (i.e. private sector) employees. Although our analysis focuses on actuarial differences across cohorts, we also provide some evidence of within cohorts redistribution by distinguishing workers by gender and occupation. We computed two actuarial indicators, namely the present value ratio (PVR) and the implicit tax rate (TAX). Improving on the existing literature (Ferraresi and Fornero 2000, Fornero and Castellino 2001), we allowed for dynamic mortality in their computation. We particularly investigated the actuarial properties of the rules designed by the 1995 reform (as modified by the 2007 law) to handle increasing longevity. To this aim, we built projected cohort- and gender-specific mortality tables based on a limit demographic scenario recently depicted by demographic experts. Results for the current legislation were compared with those stemming from a scenario where conversion coefficients, used to transform the accumulated sum at retirement into the annuity, are based on cohort- and gender-specific mortality rates.

The old DB pension rules provide extremely generous, much more than actuarially fair, pension benefits. The steepest age-wage profiles are typically associated with higher PVR. If, for instance, we consider the 1945 cohort, the indicator is equal to 154 % for male white-collar workers (steepest profile) and to 126.5 % for female blue-collar workers (flattest profile). The DB scheme also provides strong incentives to early retirement (TAX up to 70 % for female blue-collar workers). Throughout the transitional phase toward the DC rules, the generosity and the implicit taxation of the system progressively decrease.

Due to dynamic efficiency, the steady-state (N)DC regime is less than actuarially fair (the PVR is about 90 % for males and 95 % for females). For the same reason, the new regime is less than actuarially fair at margin for most of the retiring cohorts and agents (excluding the years in which conversion coefficients are updated, the TAX is about 4 %). The implicit redistribution of the Italian pension system changed from “wage-based” (DB) to “mortality-based” (NDC).

As a consequence of the rules adopted to compute conversion coefficients and to update them in response to increasing longevity, the future DC steady state is more than *quasi*-actuarially fair on average (at margin). Periodical *ex-post* adjustments in the conversion coefficients generate discontinuities in the pension treatment of adjacent cohorts as well as strong incentives to retire prior to each adjustment. The 2007 reform, reducing from 10 to 3 years the temporal validity of the conversion coefficients, was quite effective in improving actuarial fairness. Nevertheless, actuarial unfairness is still embedded in the Italian DC system, a problem that mostly depends on the use of historical cross-sectional mortality rates in the computation of conversion coefficients. Retired cohorts will likely live longer than what accounted for in the computation of their pension benefits. This discrepancy occurs because cohort effects in mortality are disregarded by the Italian law.

Finally, our simulations highlight that adopting cohort-specific conversion coefficients based on projected mortality tables would drastically improve actuarial fairness across cohorts and retirement ages. Since in that case pension computations would be based

on mortality projections, the implementation of such a policy would mean dealing with both possible discretionary policy interventions and uncertainty. For this reason, operations and proceedings by the demographic experts in charge of producing mortality tables would need to be as transparent as possible. Revisions would need to be performed by a strictly technical team. Uncertainty related to the projections can be indeed high for young cohorts, but is more limited for those approaching retirement (unfortunately, we cannot quantify projection uncertainty in our exercise). As Palmer (2006) pointed out, *quasi-actuarially* fair NDC pension systems can opt for either uncertainty *at* retirement - i.e. relying on projected cohort mortality when annuitizing, but keeping constant benefits for retirees - or uncertainty *after* retirement - i.e. regularly adjusting the pensions of retirees on the basis of updated historical statistics. Current Italian social security rules are not a feasible *quasi-actuarially* fair alternative.

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Appendices

A Mortality forecasts: methodology

The construction of our life tables by cohort involves several steps. The first step consists in converting the indication of the experts on the human being limit characteristics (Robine, Crimmins, Horiuchi, and Zeng 2006) into an appropriate survival function. Following previous studies (see e.g. Duchêne and Wunsch 1993, Maccheroni 1998), we model the trend in endogenous mortality $q_{end}(x)$ by means of a Weibull model. It generates a process of mortality of an initial closed contingent of individuals of the same age according

to the following survival function:

$$\ell(x) = \exp \left[- \left(\frac{x-a}{m} \right)^b \right] \quad x \geq a; a \geq 0; b, m > 0 \quad (\text{A } 1)$$

The derivative of (A 1) provides the corresponding function of deaths which corresponds to the Weibull statistical distribution:

$$f(x) = \begin{cases} \frac{b}{m} \left(\frac{x-a}{m} \right)^{b-1} \exp \left[- \left(\frac{x-a}{m} \right)^b \right] & x \geq a; a \geq 0; b, m > 0 \\ 0 & \text{otherwise} \end{cases}$$

where we interpret $f(x)dx$ as the fraction of components of an initial contingent that are eliminated at age $[x, x + dx)$. The Weibull model, in particular the three-parameters type, permits fairly effective control of important features of the survival reference scenario, such as life expectancy at birth, the Lexis point and also the threshold below which mortality is considered avoidable. They are all functions of the parameters in equation (A 1) (Maccheroni 1998).

We then build the limit probabilities of dying due to exogenous or accidental causes $q_{eso}(x)$, i.e. accidents, traumas, etc. $q_{eso}(x)$ is valorised analyzing the trend of this set of causes of death for a group of developed countries,²³ thus constructing an initial table of minimum mortality by age. A fifth degree polynomial, which provides the analytical form of $q_{eso}(x)$, is then adapted to the empirical function obtained in this way.

The limit probabilities of dying $q_{lim}(x)$ ($x = 0, 1, \dots$) is obtained as the suitably perequated sum of (the discrete version of) the original components $q_{end}(x)$ and $q_{eso}(x)$. The abridged limit survival function is shown in table A 1. The synthetic characteristics of this mortality model are $e_0 = 109.4$ years and an extreme age of life $\omega = 125$.

Forecasts are then obtained as follows:

1. by establishing a time frame for the scenario provided by $q_{lim}(x)$ in relation to the recorded Italian mortality trends. A procedure based on the logit model (see Brass 1971, Zaba 1979) is used: the limit survival function $\ell_{lim}(x)$ obtained from $q_{lim}(x)$ is taken as standard and, using the observed survival functions (Maccheroni and Locatelli 1999) $\ell_{x,t}$ ($t = 1945, 1946, \dots, 1998$), the historical series of parameters a_t and b_t of the following relationship are studied:

$$Y_{x,t} = a_t + b_t Y_x^{lim} \quad (\text{A } 2)$$

where Y_x^{lim} is the logit of the limit life table and $Y_{x,t}$ is the logit obtained from the observed $\ell_{x,t}$. Extrapolations are performed on the historical series of a_t and b_t in order to obtain new time sequences for parameters a_t^* and b_t^* (to establish if and when $a_t^* \rightarrow 0$ and $b_t^* \rightarrow 1$ when t diverges). In our case, these results are obtained when $t = 2143$ for females and when $t = 2170$ for males. On the basis of (A 2), it is then possible to obtain the sequence of projected life tables that reflect the characteristics of the limit situation in the period of time that stretches from 1999 to the two previously defined extremities of time;

²³ In details: Norway, Holland, Belgium, Spain, France, Finland, Italy, Great Britain, Sweden, Austria, Canada, USA, Australia, Japan, New Zealand.

2. by linking the mortality models obtained as explained above with the on-going process of evolution of mortality. The gap between current and limit mortality is bridged by assuming a type of evolution that reflects the theory of expansion of mortality (Myers and Manton 1984). This step is carried out in two phases. In the first one, we define the evolution of mortality resulting from the most recent trends. In the age groups where the mortality trend is decreasing, the evolution of mortality is obtained by extrapolating the recent observed historical series of $q_{x,t}$ with a conventional exponential model. In the groups where mortality is growing, the evolution is first obtained by envisaging a stationary situation and subsequently by considering a tendentially decreasing evolution. In the second phase, we synthesize the projections for each year by averaging the results of the two mortality models (the logit A 2 and the exponential model). The average assigns gradually increasing linear weights to the life tables projected obtained from the logit (A 2) and decreasing linear weights to those obtained with the exponential model. The resulting projected life tables, therefore, initially take into account recent mortality trends and gradually assume the characteristics of the limit life table.

Table A 1. Limit life table: survivors from 100,000 live births

| Age_x | ℓ_x | Age_x | ℓ_x | Age_x | ℓ_x |
|------------------------|----------------------------|------------------------|----------------------------|------------------------|----------------------------|
| 0 | 100,000 | 40 | 99,831 | 85 | 98,985 |
| 1 | 99,997 | 45 | 99,799 | 90 | 98,242 |
| 5 | 99,992 | 50 | 99,768 | 95 | 95,017 |
| 10 | 99,983 | 55 | 99,735 | 100 | 85,236 |
| 15 | 99,968 | 60 | 99,695 | 105 | 67,726 |
| 20 | 99,948 | 65 | 99,642 | 110 | 42,180 |
| 25 | 99,923 | 70 | 99,566 | 115 | 15,306 |
| 30 | 99,894 | 75 | 99,454 | 120 | 1,980 |
| 35 | 99,863 | 80 | 99,280 | 125 | 40 |

B conversion coefficients: formulas

Conversion coefficients in the B07 and B95 scenarios are given by equation (B 1) while in the AB scenario they are given by equation (B 2). Symbols are described in table B 1.

Table B 1. symbols

| Symbol | Description | Value |
|------------------------------|--|-------------------------------|
| x | retirement age | [57,65] |
| $s(-s)$ | gender (gender of the widow(er)) | m =male, f =female |
| co | cohort | [1945,2000] |
| g_f | Expected long-run real GDP growth rate | 0.015 |
| $\ell_{x,s}/\ell_{x,s}^{co}$ | Survivors of age x and gender s , according to cross-sectional/longitudinal mortality tables | our computations ^a |
| Ω/Ω^{co} | life span according to cross-sectional/cohort- and gender-specific mortality tables | our computations |
| $\ell_{x,s}^{ved}$ | Probability for the age- x gender- s widow(er) to re-marry | INPS 1989 |
| $\Theta_{s+t,s}$ | Probability for the age- $x+t$ gender- s widow(er) to leave the family | INPS 1989 |
| ϵ_s | Age difference between the (gender- s) pensioner and the widow(er) | +3(-3) if $s=m(f)$ |
| Ψ | Quota of the pension revertible to the widow(er) | 0.6 |
| Φ_s | Reduction in the survivor's benefit due to earning test | 0.9(0.7) if $s=m(f)$ |
| k | Actuarial adjustment factor | 0.4615 ^b |

^a δ_x for $co + x < 2010$ are computed according to ISTAT 1990 mortality tables; δ_x for $2010 \leq co + x \leq 2013$ are computed according to ISTAT 2002 mortality tables.

^b law 247/2007

$$\delta_x = \left(\frac{\sum_{s=m,f} dir_{x,s} + ind_{x,s}}{2} - k \right)^{-1} \quad (B 1)$$

$$dir_{x,s} = \sum_{t=0}^{\Omega-x} \frac{\ell_{x+t,s}}{\ell_{x,s}} (1+g_f)^{-t}$$

$$ind_{x,s} = \Psi \Phi_s \sum_{t=0}^{\Omega-x} \frac{\ell_{x+t,s}}{\ell_{x,s}} \left(1 - \frac{\ell_{x+t+1,s}}{\ell_{x+t,s}}\right) (1+g_f)^{-t} \Theta_{x+t,s}$$

$$\sum_{\tau=1}^{\Omega-x-t+\epsilon_s} \frac{\ell_{x+t+\tau-\epsilon_s,-s}}{\ell_{x+t+1-\epsilon_s,-s}} \left(1 - \ell_{x+t+\tau-\epsilon_s,-s}^{ved}\right) (1+g_f)^{-\tau}$$

$$\delta_{x,s}^{co} = (dir_{x,s}^{co} + ind_{x,s}^{co})^{-1} \quad (B 2)$$

$$dir_{x,s}^{co} = \sum_{t=0}^{\Omega^{co}-x} \frac{\ell_{x+t,s}^{co}}{\ell_{x,s}^{co}} (1+g_f)^{-t}$$

$$ind_{x,s}^{co} = \Psi \Phi_s \sum_{t=0}^{\Omega^{co}-x} \frac{\ell_{x+t,s}^{co}}{\ell_{x,s}^{co}} \left(1 - \frac{\ell_{x+t+1,s}^{co}}{\ell_{x+t,s}^{co}}\right) (1+g_f)^{-t} \Theta_{x+t,s}$$

$$\sum_{\tau=1}^{\Omega^{co}-x-t+\epsilon_s} \frac{\ell_{x+t+\tau-\epsilon_s,-s}^{co+\epsilon_s}}{\ell_{x+t+1-\epsilon_s,-s}^{co+\epsilon_s}} \left(1 - \ell_{x+t+\tau-\epsilon_s,-s}^{ved}\right) (1+g_f)^{-\tau}$$

C Detailed results

In this appendix we report detailed simulation results for the base macroeconomic case ($r = 0.02$, $g_f = 0.015$). Scenario B95 = law n. 335/1995; Scenario B07 = current legislation; Scenario AB = conversion coefficients are cohort- and gender-specific. DB = defined benefit; PR = *pro-rata*; DC = notional defined contribution. Indicators are in percentage points.

| cohort/scheme | Scenario B95 | | | | | Scenario B07 | | | | | Scenario AB | | | | | |
|---------------|--------------|------|------|------|------|--------------|------|------|------|------|-------------|------|------|------|------|------|
| | 60 | 61 | 62 | 63 | 64 | 60 | 61 | 62 | 63 | 64 | 60 | 61 | 62 | 63 | 64 | 65 |
| 1945 DB | 34.0 | 38.0 | 45.8 | 49.0 | 91.0 | 34.0 | 38.0 | 45.8 | 49.0 | 91.0 | 34.0 | 38.0 | 45.8 | 49.0 | 91.0 | 92.0 |
| 1946 DB | 34.3 | 42.2 | 45.3 | 48.5 | 90.8 | 34.3 | 42.2 | 45.3 | 48.5 | 90.8 | 34.3 | 42.2 | 45.3 | 48.5 | 90.8 | 91.8 |
| 1947 DB | 38.4 | 41.6 | 44.8 | 47.9 | 90.6 | 38.4 | 41.6 | 44.8 | 47.9 | 90.6 | 38.4 | 41.6 | 44.8 | 47.9 | 90.6 | 93.3 |
| 1948 DB | 37.8 | 41.1 | 44.2 | 47.4 | 92.3 | 37.8 | 41.1 | 44.2 | 47.4 | 92.3 | 37.8 | 41.1 | 44.2 | 47.4 | 92.3 | 93.2 |
| 1949 DB | 37.3 | 40.6 | 43.7 | 49.0 | 92.1 | 37.3 | 40.6 | 43.7 | 49.0 | 92.1 | 37.3 | 40.6 | 43.7 | 49.0 | 92.1 | 93.1 |
| 1950 DB | 36.8 | 40.1 | 45.3 | 48.7 | 92.0 | 36.8 | 40.1 | 45.3 | 48.7 | 92.0 | 36.8 | 40.1 | 45.3 | 48.7 | 92.0 | 93.0 |
| 1951 DB | 36.1 | 41.6 | 44.9 | 48.3 | 91.9 | 36.1 | 41.6 | 44.9 | 48.3 | 91.9 | 36.1 | 41.6 | 44.9 | 48.3 | 91.9 | 92.8 |
| 1952 DB | 37.8 | 41.3 | 44.6 | 47.9 | 91.7 | 37.8 | 41.3 | 44.6 | 47.9 | 91.7 | 37.8 | 41.3 | 44.6 | 47.9 | 91.7 | 92.7 |
| 1953 DB | 37.4 | 40.9 | 44.2 | 47.6 | 91.6 | 37.4 | 40.9 | 44.2 | 47.6 | 91.6 | 37.4 | 40.9 | 44.2 | 47.6 | 91.6 | 92.6 |
| 1954 PR | 20.9 | 20.8 | 20.8 | 20.6 | 98.4 | 22.9 | 34.4 | 24.1 | 24.3 | 38.3 | 25.8 | 29.9 | 30.5 | 31.1 | 31.9 | 32.6 |
| 1955 PR | 19.2 | 19.1 | 19.0 | 18.8 | 93.6 | 24.4 | 32.2 | 22.2 | 22.4 | 35.9 | 23.8 | 24.1 | 28.6 | 29.2 | 29.9 | 30.6 |
| 1956 PR | 17.6 | 17.4 | 17.3 | 89.0 | 22.4 | 22.6 | 20.5 | 20.6 | 33.6 | 21.8 | 22.1 | 37.6 | 27.4 | 28.0 | 28.6 | 30.0 |
| 1957 PR | 15.9 | 15.7 | 84.8 | 20.4 | 20.6 | 20.8 | 18.8 | 31.4 | 20.0 | 20.1 | 35.1 | 21.5 | 26.2 | 26.8 | 27.4 | 28.8 |
| 1958 PR | 14.2 | 80.8 | 18.6 | 18.7 | 18.9 | 19.0 | 29.2 | 18.2 | 18.3 | 32.8 | 19.5 | 19.7 | 25.0 | 25.5 | 26.1 | 26.8 |
| 1959 PR | 77.1 | 16.8 | 16.9 | 17.0 | 17.1 | 17.2 | 16.4 | 16.5 | 30.5 | 17.7 | 17.8 | 34.3 | 23.8 | 24.3 | 24.9 | 25.5 |
| 1960 PR | 15.1 | 15.2 | 15.2 | 15.3 | 15.4 | 14.8 | 28.3 | 15.8 | 16.0 | 31.9 | 17.2 | 22.6 | 23.1 | 23.6 | 24.3 | 24.9 |
| 1961 PR | 13.5 | 13.5 | 13.6 | 13.6 | 13.6 | 26.2 | 14.1 | 14.2 | 29.6 | 15.3 | 15.4 | 21.3 | 21.9 | 22.4 | 23.0 | 24.3 |
| 1962 PR | 11.9 | 11.9 | 11.9 | 11.8 | 11.8 | 12.5 | 12.5 | 27.3 | 13.5 | 13.5 | 31.0 | 20.1 | 20.6 | 21.2 | 21.8 | 23.0 |
| 1963 PR | 10.3 | 10.2 | 10.2 | 10.1 | 10.0 | 10.0 | 10.9 | 25.2 | 11.7 | 11.8 | 28.7 | 12.9 | 18.9 | 19.4 | 19.9 | 20.5 |
| 1964 PR | 8.7 | 8.6 | 8.5 | 8.4 | 8.3 | 71.6 | 23.1 | 10.1 | 10.1 | 26.4 | 11.1 | 11.1 | 17.8 | 18.2 | 18.7 | 19.3 |
| 1965 PR | 7.0 | 6.9 | 6.8 | 6.6 | 67.9 | 10.0 | 8.5 | 8.5 | 24.2 | 9.3 | 9.3 | 27.7 | 16.6 | 17.0 | 17.5 | 18.0 |
| 1966 PR | 5.4 | 5.2 | 5.1 | 64.2 | 8.3 | 8.2 | 6.9 | 22.1 | 7.7 | 7.7 | 25.4 | 8.6 | 15.4 | 15.8 | 16.3 | 16.8 |
| 1967 PR | 3.8 | 3.6 | 60.7 | 6.6 | 6.5 | 6.5 | 20.1 | 6.1 | 6.0 | 23.2 | 6.9 | 6.8 | 14.2 | 14.6 | 15.1 | 15.6 |
| 1968 PR | 2.2 | 57.4 | 5.0 | 4.9 | 4.4 | 4.7 | 4.6 | 4.5 | 21.0 | 5.2 | 5.1 | 24.4 | 13.1 | 13.4 | 13.9 | 14.4 |
| 1969 PR | 54.2 | 3.5 | 3.4 | 3.2 | 3.1 | 2.9 | 3.0 | 19.0 | 3.7 | 3.6 | 22.1 | 4.3 | 11.9 | 12.3 | 12.7 | 13.2 |
| 1970 PR | 2.0 | 1.8 | 1.7 | 1.6 | 1.4 | 1.2 | 17.0 | 2.2 | 2.0 | 20.0 | 2.7 | 2.6 | 10.7 | 11.1 | 11.5 | 12.0 |
| 1971 DC | 0.4 | 0.2 | 0.0 | -0.2 | -0.4 | -0.6 | 0.7 | 0.5 | 17.5 | 1.2 | 1.0 | 20.6 | 9.4 | 9.8 | 10.2 | 10.6 |
| 1972 DC | 0.0 | -0.2 | -0.4 | -0.6 | -0.8 | -1.1 | 0.3 | 16.3 | 0.9 | 0.7 | 19.2 | 1.4 | 9.4 | 9.7 | 10.1 | 10.5 |
| 1973 DC | -0.3 | -0.5 | -0.8 | -1.0 | -1.3 | -1.6 | 15.2 | 0.7 | 0.5 | 18.0 | 1.1 | 0.9 | 9.3 | 9.7 | 10.0 | 10.5 |
| 1974 DC | -0.6 | -0.9 | -1.2 | -1.4 | -1.7 | 65.1 | 0.5 | 0.3 | 16.8 | 0.9 | 0.7 | 19.7 | 9.3 | 9.6 | 10.0 | 10.4 |
| 1975 DC | -1.0 | -1.2 | -1.5 | -1.8 | 61.4 | 1.1 | 0.2 | 15.7 | 0.7 | 0.5 | 18.4 | 1.1 | 9.2 | 9.6 | 9.9 | 10.3 |
| 1976 DC | -1.3 | -1.6 | -1.9 | 57.8 | 0.8 | 0.6 | 14.7 | 0.5 | 0.3 | 17.3 | 0.8 | 0.6 | 9.2 | 9.5 | 9.9 | 10.3 |
| 1977 DC | -1.6 | -2.0 | 54.3 | 0.6 | 0.4 | 0.2 | 0.4 | 0.2 | 16.2 | 0.6 | 0.4 | 18.9 | 9.2 | 9.5 | 9.8 | 10.2 |
| 1978 DC | -2.0 | 51.1 | 0.5 | 0.3 | 0.0 | -0.3 | 0.1 | 15.1 | 0.5 | 0.3 | 17.7 | 0.7 | 9.1 | 9.4 | 9.8 | 10.1 |
| 1979 DC | 48.0 | 0.3 | 0.1 | -0.1 | -0.4 | -0.7 | 14.2 | 0.3 | 0.1 | 16.6 | 0.5 | 0.3 | 9.1 | 9.4 | 9.7 | 10.1 |
| 1980 DC | 0.2 | 0.0 | -0.3 | -0.5 | -0.8 | -1.2 | 0.2 | 0.2 | 0.0 | 15.5 | 0.4 | 0.1 | 18.2 | 9.1 | 9.4 | 9.7 |
| 1981 DC | -0.1 | -0.3 | -0.6 | -0.9 | -1.2 | -1.6 | -0.1 | 14.6 | 0.3 | 0.0 | 17.0 | 0.4 | 9.0 | 9.3 | 9.6 | 9.9 |
| 1982 DC | -0.4 | -0.7 | -1.0 | -1.3 | -1.7 | -2.1 | 13.7 | 0.2 | -0.1 | 16.0 | 0.3 | 0.0 | 9.0 | 9.0 | 9.3 | 9.6 |
| 1983 DC | -0.7 | -1.0 | -1.4 | -1.7 | -2.1 | -2.5 | 0.1 | -0.2 | 15.0 | 0.1 | -0.2 | 17.4 | 0.7 | 9.2 | 9.5 | 9.8 |
| 1984 DC | -1.0 | -1.4 | -1.7 | -2.1 | -2.5 | 58.9 | -0.2 | 14.0 | 0.0 | -0.2 | 16.4 | 0.1 | 8.9 | 9.2 | 9.5 | 9.8 |
| 1985 DC | -1.4 | -1.7 | -2.1 | -2.5 | 55.5 | -0.1 | 13.2 | 0.0 | -0.3 | 15.3 | 0.0 | -0.4 | 8.9 | 9.1 | 9.4 | 10.0 |
| 1986 DC | -1.7 | -2.1 | -2.5 | 52.2 | -0.1 | -0.5 | -0.1 | -0.4 | 14.4 | -0.1 | -0.5 | 16.7 | 8.9 | 9.1 | 9.4 | 10.0 |
| 1987 DC | -2.0 | -2.4 | 49.2 | -0.2 | -0.6 | -0.9 | -0.4 | 13.5 | -0.2 | -0.5 | 15.7 | -0.3 | 8.8 | 9.1 | 9.3 | 9.6 |
| 1988 DC | -2.3 | 46.3 | -0.3 | -0.6 | -1.0 | -1.4 | 12.7 | -0.4 | -0.6 | 14.7 | -0.4 | -0.7 | 8.8 | 9.0 | 9.3 | 9.6 |
| 1989 DC | 43.5 | -0.3 | -0.6 | -1.0 | -1.4 | -1.8 | -0.2 | -0.6 | 13.8 | -0.4 | -0.8 | 16.0 | 8.8 | 9.0 | 9.2 | 9.5 |
| 1990 DC | -0.3 | -0.7 | -1.0 | -1.4 | -1.8 | -2.2 | -0.6 | 13.0 | -0.5 | -0.8 | 15.0 | -0.6 | 8.7 | 9.0 | 9.2 | 9.5 |
| 1991 DC | -0.7 | -1.0 | -1.4 | -1.7 | -2.2 | -2.7 | 12.2 | -0.5 | -0.8 | 14.1 | -0.7 | -1.1 | 8.7 | 8.9 | 9.2 | 9.4 |
| 1992 DC | -1.0 | -1.3 | -1.7 | -2.1 | -2.6 | -3.1 | -0.5 | -0.8 | 13.3 | -0.7 | -1.1 | 15.4 | 8.7 | 8.9 | 9.1 | 9.4 |
| 1993 DC | -1.3 | -1.7 | -2.1 | -2.5 | -3.0 | -3.6 | -0.8 | 12.5 | -0.7 | -1.1 | 14.4 | -1.0 | 8.7 | 8.9 | 9.1 | 9.3 |
| 1994 DC | -1.6 | -2.0 | -2.4 | -2.9 | -3.4 | 53.6 | 11.7 | -0.7 | -1.1 | 13.6 | -1.0 | -1.4 | 8.6 | 8.8 | 9.0 | 9.3 |
| 1995 DC | -1.9 | -2.4 | -2.8 | -3.3 | 50.6 | -1.3 | -0.7 | -1.0 | 12.7 | -1.0 | -1.4 | 14.7 | 8.6 | 8.8 | 9.0 | 9.3 |
| 1996 DC | -2.2 | -2.7 | -3.2 | 47.6 | -1.2 | -1.7 | -1.0 | 12.0 | -1.0 | -1.4 | 13.9 | -1.4 | 8.6 | 8.8 | 9.0 | 9.2 |
| 1997 DC | -2.6 | -3.0 | 44.8 | -1.2 | -1.6 | -2.1 | 11.3 | -0.9 | -1.3 | 13.0 | -1.3 | -1.8 | 8.6 | 8.8 | 9.0 | 9.2 |
| 1998 DC | -2.9 | 42.2 | -1.2 | -1.6 | -2.0 | -2.5 | -0.9 | -1.3 | 12.3 | -1.3 | -1.7 | 14.1 | 8.5 | 8.7 | 8.9 | 9.2 |
| 1999 DC | -39.7 | -1.1 | -1.5 | -2.0 | -2.4 | -3.0 | -1.2 | 11.5 | -1.3 | -1.7 | 13.3 | -1.8 | 8.5 | 8.7 | 8.9 | 9.1 |
| 2000 DC | -1.1 | -1.5 | -1.9 | -2.3 | -2.9 | -3.4 | 10.8 | -1.2 | -1.6 | 12.5 | -1.7 | -2.2 | 8.5 | 8.7 | 8.9 | 9.1 |

