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ACTUARIAL NEUTRALITY WHEN LONGEVITY INCREASES: AN APPLICATION TO THE ITALIAN PENSION SYSTEM

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Abstract

As a possible solution to the severe crises of the PAYGO pension systems induced by population aging and by low economic growth, many countries are recently switching from a defined benefit (DB) to a notional defined contribution (NDC) scheme. Particularly important for the NDC systems are the rules which establish how to incorporate into the pension *formulae* life expectancies and their changes.

This work investigates these issues focusing on the Italian NDC system, introduced by the 1995 reform and gradually taking the place of the previous DB one. The methodology follows a representative agents approach, in which each agent describes a cohort, either involved in the transition to the new rules or of steady-state. Cohort-and-gender-specific mortality projections, developed *ad hoc* for this work, are exploited to establish a benchmark against which actuarial fairness and neutrality can be assessed.

The NDC Italian pension system is almost actuarially fair and neutral. However, some noticeable distortions from our benchmark exist. A remarkable one is the use in the pension *formulae* of mortality tables which do not incorporate longitudinal trends in mortality rates.

Keywords: social security, notional defined contribution, actuarial neutrality, actuarial fairness, money's worth measures, demographic trends, cohort-specific mortality projections.

JEL codes: J11, J14, H55, J16, C23, D31.

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1 Introduction

As a possible solution to the severe crises of the PAYGO pension systems induced by population aging and by low economic growth, many countries are recently switching from a defined benefit (DB) to a notional defined contribution (NDC) scheme. Between the European countries, for example, there are Sweden, Italy, Poland and Latvia. The scheme guarantees a return equal to the economy growth rate and a pension compatible with the population-based life expectancy, shifting both the macroeconomic and the mortality risk from the public budget to the worker. If the economy is expected to grow less in the future years, a lower pension is paid, and the same occurs if life expectancy has recently increased. Financial sustainability of the system is therefore expected to improve (see e.g. Palmer 1999 on the Swedish reform).

The (N)DC pension formula is based on the principles of actuarial fairness and actuarial neutrality. According to the first principle, the expected present value of pension benefits is equal to the present value of contributions paid during the working career. This formula therefore strengthens the insurance role and reduces the redistributive impact of the pension system with respect to the DB one. According to the second principle, an early retiree receives a lower pension for an expected longer period, and the expected present value of net resources she benefits is the same (for the same amount of contributions paid) she would have benefited in case of postponed retirement. The (N)CD formula, therefore, potentially leaves financial considerations at margin out of the determinants of the retirement age.

Whether these principles guarantee desirable features of a public pension system is much discussed in the literature. In the debate the characteristics of DB and (N)DC schemes are compared, as well as those of different types of pension financing (PAYGO, funded or mixed systems).[1] While the theoretical analysis of different pension systems goes beyond the scope of this paper, it is important to note how these studies (e.g. Disney 1999) often highlight how strongly the comparison depends on the practical implementation of the principles, i.e. on what is established by the normative details.

Particularly important for the NDC systems are the rules which establish how to incorporate into the pension *formulae* life expectancies and their changes. Consider, for example, the following cases and their possible consequences. If the population-based value of mortality is taken into account but its heterogeneity within the population is not (i.e. the insurance concept of risk-pooling), there is a redistribution from shorter to longer living individuals. Although in some cases, it can be considered desirable in a "solidaristic" view (or, as between genders, in a "family" view) of public pension programmes, in others it cannot. This occurs, for example, whenever it is statistically documented that the richest individuals are also the longer living. Moreover, if the pensions are adapted with a delay to the changed mortality rates, redistribution between cohorts and financial difficulties can easily arise. Finally, the same effects can be produced

[1] Selected examples are Diamond (2004), Lindbeck and Persson (2003) and Gomes and Michaelides (2003).

by an imperfect measure of mortality rates. In particular, they can occur if mortality tables are cross-sectional, i.e. do not disentangle cohort from time effects in the evolution of mortality. Due to the fact that the introduction of NDC systems is recent, these issues have not been sufficiently analyzed so far.

Under the pressure of increasing deficits and an expected older and older population, in 1995 a reform introduced in Italy a NDC pension system which is gradually replacing the previous DB one. In this paper we exploit the Italian reform to investigate in details some of the issues introduced above. We provide an empirical evaluation of the actuarial fairness and actuarial neutrality of the new Italian pension system, in light of the expected changes in mortality. We look at both the long transition from the DB toward the NDC rules, and at the future steady state.

We develop the methodology of Ferraresi and Fornero (2000) - who compute money's worth measures (mwm henceforth) for the Italian pension system in a scenario characterized by constant mortality - by having endogenous (i.e.mortality-related) transformation coefficients (a determinant of the pension benefit, see section 3.1). We follow a representative agents approach, in which each agent describes a cohort, either involved in the transition or of steady-state. We also include some intragenerational heterogeneity, distinguishing the agents born in the same year according to gender, occupation, sector and geographical area. One feature of the model is that it incorporates a rather precise (and complex) normative framework, corresponding to the rules applying in the transition and in the new steady state. Each representative agent is defined by a stylized, extremely simplified, working career and by an estimated life-cycle average wage profile. The cohort dimension of the model is especially developed: both the estimation of the average wage profile and the projections of mortality rates disentangle the cohort from the age and the time effects in their evolutionary process. A weakness of the model is that it incorporates only two sources of uncertainty, on wages and on the lifetime. Nevertheless, given that the interest lies on long-run comparisons and on differentials in mortality more than on heterogeneous working careers, we believe that the model represents a suitable tool for this kind of analysis.

Two different mortality projections, developed *ad hoc* for this paper, are exploited. The first one, more standard, is *cross-sectional*, while the second, which disentangle cohort and time effects in the evolution of mortality, is *longitudinal*. The first one is used in a scenario which reproduces the present legislation, while the second is used to simulate a more actuarially neutral system, which we take as a benchmark of actuarial fairness and neutrality. For each agent we compute money's worth measures in both of the scenarios, and we quantify how much the system departs from the benchmark by comparing the two sets of results.

The paper proceeds as follows. Section 2 describes life expectancy in Italy, first focusing on its historical trends and then proposing *ad hoc* cohort-specific mortality projections. Section 3 describes the DC pension formula introduced by the 1995 reform, the model and the *formulae* used to compute mwm. It also proposes a definition of actuarial fairness and neutrality based on these *formulae*. Section 4 shows the results, focusing first on the forecasts of the transformation coefficients (4.1) and then on the mwm (4.2 and 4.3). Section 5 draws some conclusions. The paper is completed by four appendixes, describing the methodology used to forecast mortality (A), the *formulae* used to predict transformation coefficients (B), the model details (C) and the estimation of wage profiles (D).

2 Life expectancy variations in Italy

2.1 Historical trends and differentials

Longevity, reflected in the ever higher proportion of the elderly in the total population, is one of the most striking phenomena of post-modern societies. In recent times,the main factor in determining this situation has been the rapid decline in mortality rates at old ages, dictated by the continuous decline in overall mortality, which began long ago with improvements in public and private hygiene, the progress of medicine and also the continuous improvement in general economic conditions, levels of education and life styles.

At the individual level, a higher income and a higher standard of education have positive repercussions on the risk of death, as they facilitate consumption habits and standards of living that guarantee greater protection of health and living conditions. At aggregate level, the increase in income is accompanied by higher investments in public health, with improvements in the availability of services and access to medical care, as well as in the promotion of medical research.

In Italy, this process developed during the 20th century. It started in the first half of the century, with a general improvement in hygiene - e.g., the major land reclamation projects and the fight against malaria - and more effective treatments of infectious and respiratory diseases which dominated the epidemiological framework of the time. Following a major turnaround in the second half of the century, the levels of life expectancy improved, reaching the highest-level countries towards the end of the 90s. Infectious diseases gradually disappeared and infant mortality continued to decline. After a further improvement in the 60s and 70s, however, the increase in life expectancy at birth tended to slow down again, grinding to a standstill at old and oldest-old ages (table 1). Indeed, there had been a change in the epidemiological traits: the main causes of death were diseases of the circulatory system and neoplasms, both considerably affected by life styles inherited unconsciously from the past (in particular smoking and diet) or which were the direct effect of new dangers, such as road accidents.

Therefore, medicine had to measure up to new challenges (today, more than 70 percent of deaths are to be ascribed to diseases of the circulatory system and neoplasms) and develop a fine-tuned and efficient campaign of prevention and cure before it met with the first signs of success, which started to appear in the second half of the 80s, marked by an increasingly evident decline in diseases of the circulatory system. On the other hand, deaths caused by neoplasms

	Males			Females		
Period	e_0	e_{60}	e_{80}	e_0	e_{60}	e_{80}
1950-1953	63.7	16.0	5.0	67.3	17.5	5.5
1960-1962	67.2	16.7	5.7	72.3	19.3	6.4
1970-1972	69.0	16.7	5.8	74.9	20.2	6.7
1979-1983	71.0	16.9	5.8	77.7	21.3	7.2
1992	73.8	18.7	6.7	80.4	23.2	8.1
2000	76.5	20.4	7.3	82.5	24.9	9.2

Table 1: Comparison of e_0, e_{60} and e_{80} by gender from the 50s to the year 2000

Note: Italy, general mortality.
Source: ISTAT, life tables, various years.

continued to increase until the end of the 90s with a recent slight reversal of trend that, taking into account progress in the field of biomedical research, seems to be continuing.

It is this change in mortality causes that marks the second phase of the evolution of mortality in the second half of the 80s.[2] Following the decline in infantile mortality to almost "minimum terms", the barrier to achieving old age finally crumbled. With the mortality conditions prevailing at the start of 1900, only 37-38 percent of persons born at that time had any hope of reaching 65 years of age whereas, in the present conditions, about 90 percent of new-born boys and almost 95 percent of new-born girls can achieve this target (Maccheroni 1999). These results have had beneficial effects (although not to the same extent) on all age groups and also on both genders although, as is known and highlighted in table 1, females remain ahead. Recently, however, male life expectancy is tending to increase slightly more quickly than that of the female population, so that the gap, after reaching its peak in the 80s, is now slightly narrowing.

Although the gender differentials in mortality are the most widely known (also because they were the first to be studied in the past), those depending on social class and place of residence are equally important.

As is known, mortality in the most underprivileged social groups is higher than that in the rest of the population. Analyses in this field draw their results from a complex statistical base that generally links current recordings of mortality with those returned by censuses. In Italy, the statistics concerned refer to the 1981 and 1991 censuses. In this framework, one of the most frequently used characteristics - as in our case - is education level, because it remains relatively stable throughout a person's life and also because, in addition to elements that identify social-economic status, it also reflects differences as regards to access to information and the ability to benefit from cultural resources, which are consid-

[2] A recent further positive note has been the decline in deaths of young people due to AIDS, as a result of more wide-scale prevention and the introduction of more effective therapies, interrupting the sudden upswing in mortality for this cause that was triggered in the 90s. On the other hand, pathologies such as mental disorders and diseases of the nervous system at old ages have continued to increase from these years.

ered important factors in determining life styles and therefore affect the quality and duration of life.

Area	Educational qualification				
	1	**2**	**3**	**4**	**5**
			Males		
			Age 18–59		
North	188.59	140.73	103.01	69.79	47.11
Center	169.30	134.29	102.96	78.78	57.89
South	170.84	114.70	100.28	74.40	62.92
ITALY	172.57	128.51	102.32	72.29	52.75
			Age 60–74		
North	107.73	104.01	98.61	82.48	72.30
Center	111.90	101.36	100.21	91.97	70.47
South	104.87	96.76	107.70	94.96	86.53
ITALY	105.60	102.22	101.15	87.07	75.06
			Females		
			Age 18–59		
North	212.78	109.50	106.69	88.42	75.42
Center	188.09	104.26	106.02	84.73	87.23
South	149.54	97.59	96.93	79.79	91.21
ITALY	168.10	100.52	104.30	85.89	82.97
			Age 60–74		
North	103.27	99.88	103.88	87.81	85.55
Center	108.92	97.78	97.24	90.03	94.02
South	106.97	97.59	89.94	76.12	76.59
ITALY	109.34	98.07	98.96	84.93	84.86

Table 2: Index numbers of age-standardized mortality rates* by gender, age, educational qualification and geographical area of residence: years 1991-1992

*See note 3.
Note: educational qualification: 1=illiterate and without qualification, 2=junior school leaving certificate, 3=lower middle school leaving certificate, 4=upper middle school leaving certificate, 5=degree.
Source: ISTAT (2001).

The relationships that can be immediately established between education level, gender, age, place of residence and the corresponding mortality levels expressed in age-standardized death rate index numbers[3] (table 2) reveal a multifaceted framework. in fact, the table shows the pronounced reverse gradient of mortality - the higher one rises in the social ladder, the lower the risk of death - and the fact that this gradient varies by geographical area. The differential mortality associated with education level is particularly accentuated in the active age group (18-59 years) especially for males and in particular in the

[3]ISTAT doesn't supply life tables by social status; so we use standardized mortality rates which permit synthetic and correct comparisons of mortality levels.

North, where mortality levels at the bottom end of the social ladder are four times those of the upper end. The inverse correlation is less marked in the case of females; this can be explained by the fact that the incidence of breast cancer is directly correlated with social status (a point against females with a higher education grade) and the differences in mortality are less accentuated in the Center and South compared with the North.

Although - as mentioned earlier - over the years far-reaching changes have occurred in the epidemiological traits and there has been a reduction in mortality in the various social classes, it is generally considered that in developed countries this social-economic gradient of mortality has continued notwithstanding in-depth changes in health and social policy. The first disconcerting results in this field were formulated for Great Britain in the Black Report (1980)[4] that caused a great stir by casting doubt on the assumption that a national health service with universal, egalitarian access to the services provided (such as that existing at that time in Great Britain and which would be subsequently imported to many other countries, including Italy) might, on its own, represent the most suitable instrument for improving the state of health and also for reducing social disparity in this area. Theoretical parity of access to health services does not however imply effective equality in using these as, once again, the "messages" from the health system are interpreted differently according to education level and social position.

Another element that reflects the complexity of the process of evolution of mortality and the considerable scope of this differential emerges when the role of the territorial factor is considered. Differential mortality also changes according to the geography of mortality by cause. In this case also, although the continuous progress achieved as regards survival has reduced the variability of regional differences, particularly in the case of males, these continue to be clearly marked, maintaining the positions of the various regions in a framework that has remained more or less unchanged since the 60s. However, current trends suggest that major changes are in the offing also from a geographical point of view, as is also reflected in table 3.

As confirmed by recent trends, the expectations of life are highest in Central Italy. The situation of the other two distributions is however more articulated. Distinguishing by gender, the expectations of life are generally lowest in the North (Lombardia, Piemonte - Valle d'Aosta, Friuli Venezia-Giulia and Liguria) for males, due to the higher incidence of neoplasms; the same situation is found in the South, especially in Campania. As far as females are concerned, mortality tends to be higher in the South due to the higher incidence of diseases of the circulatory system; this is true not only for Campania but also for Sicily where diabetes is also a major factor of mortality risk.

The current epidemiological traits are, however, destined to change, as suggested by the therapeutic potential of biomedical research; this change will occur depending on the time required to test the new results achieved. However, gen-

[4]This is not obviously an English anomaly; similar results were recorded, in the same period in Holland and Sweden. In France social disparity in relation to death does not seem to have faded (Desplanques 1993; Leclerc, Fassin, Granjean, Kaminski, and Lang 2000).

eral optimism in this field, reflected in international literature, suggests that, at least in the short term, further positive surprises are in the pieline.

Regions	Males				Females			
	e_0		e_{60}		e_0		e_{60}	
	60-62	88-92	60-62	88-92	60-62	88-92	60-62	88-92
PiemonteV.A.	66.6	73.3	*16.2*	*18.6*	72.5	80.1	*19.1*	*23.2*
Lombardia	65.4	72.6	*15.1*	*17.9*	72.0	80.1	*18.4*	*23.2*
Trentino A.A.	65.2	73.1	*15.4*	*18.6*	72.0	80.9	*18.3*	*23.9*
Veneto	66.8	73.4	*15.9*	*18.4*	72.9	80.8	*18.8*	*24.0*
Friuli-V.G.	66.9	72.8	*15.6*	*17.9*	73.2	80.3	*19.1*	*23.3*
Liguria	69.0	73.5	*17.1*	*18.7*	74.3	80.4	*20.2*	*23.5*
Emilia-Rom.	68.1	74.2	*16.5*	*19.3*	74.2	80.9	*19.8*	*24.0*
Toscana	69.1	74.7	*16.9*	*19.3*	74.0	81.1	*19.7*	*23.8*
Umbria	69.4	75.0	*16.9*	*19.5*	73.5	80.9	*19.0*	*23.7*
Marche	69.2	75.4	*16.8*	*19.9*	74.2	81.3	*19.1*	*24.2*
Lazio	68.6	74.0	*17.0*	*18.8*	73.5	80.1	*19.9*	*23.4*
Abruzzo-Mol.	68.9	74.9	*17.3*	*19.7*	72.3	80.7	*19.0*	*23.5*
Campania	66.2	72.8	*16.6*	*18.1*	70.2	78.6	*18.8*	*21.9*
Puglia	67.1	74.6	*17.5*	*19.5*	70.6	80.1	*19.3*	*23.1*
Basilicata	67.5	74.8	*18.0*	*19.7*	70.3	79.9	*18.7*	*22.9*
Calabria	69.0	74.3	*18.2*	*19.5*	71.6	80.0	*19.7*	*23.0*
Sicilia	68.5	73.9	*17.9*	*19.0*	71.3	79.0	*19.2*	*22.1*
Sardegna	69.4	73.9	*19.3*	*19.4*	73.4	80.5	*20.8*	*23.4*

Table 3: Comparison of e_0 and e_{60} by region and gender from the 60s to the 90s

Note: Italy, general mortality.
Source: ISTAT, life tables, various years.

2.2 Cohort-specific mortality forecasts

As already mentioned, the silent revolution that has completely remodeled the previous mortality patterns offers everyone or almost everyone the possibility of a longer life - therefore deaths are concentrated in old or oldest-old groups - so that future evolution will be conditioned by tendencies that emerge in the last phase of life. However, it is precisely on this point that the theories diverge.

According to the theory of the *compression of mortality*, the first to be formulated (Fries 1980, Fries 1983 and Fries 1989), the average life span cannot exceed the limit of 85 years and, as many more people gain years of life, there will be a process of concentration of their deaths in a narrow interval astride this threshold. However, this process is not clearly defined due to the presence of individuals with exceptional characteristics that die at a very old age. According to the theory of the *expansion of mortality* (Myers and Manton 1984; Olshanski, Carnes, and Cassel 1993), our life span will continue to increase, as confirmed by statistics regarding the maximum life span gradually reached by persons who have died in the last forty years in developed countries and also the decline in the mortality of the oldest old (80 years and more). Therefore, according to the current declining trend of mortality, future birth cohorts will live even longer in an evolution that does not exclude the previous process. In this context, the "cohorts" effect, already recognized as a further factor of differential mortality, would continue to play its positive role (Caselli 1990).

Recent experience has lead to the conclusion that the limit assumed by the first of the two theories can be considered "narrow". Therefore, although the hypothesis of future evolution of survival levels according to the theory of expansion of mortality continues to be the more credible, uncertainty remains as regards the presumable levels that may be reached by life expectancy. In this case also, the only solution is to refer (as we will see below) to the indications of the experts, bearing in mind that both theories implicitly assume that the current process of development and progress will continue into the future in the wake of the progress recorded recently in all fields and in particular in the sector of biomedical research.[5] Studies of mortality can boast a long sequence of methodological approaches that have delineated future prospects (Tabeau, Berg, Heathcote, and Heathcote 2001). Although in the past, in some cases, these prospects seemed to belong to "futurology" (Maccheroni 1999), they then proved to be understated as regards the real evolution of the phenomenon.

Therefore, aware of the margins of uncertainty of a long-term forecast, we use a "limit" scenario that forms the underpinnings of the methodological system constructed to simulate the process of mortality of Italian cohorts born between 1940 and 2000, in order to formulate a forecast whose time horizon becomes gradually broader moving from the oldest cohorts to the most recent ones. In particular, the scenario used also considered the results of various recent seminars that outlined some of the most important traits for human survival.[6] In fact, the twenty-first century is considered 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 aging"; in this context of protection of health, the endogenous mortality process could generate a life expectancy of close on 110 years and a maximum life span of just over 120 years.

To formulate our forecast (the methodology used is illustrated in appendix A), we first of all insert the scenario concerned in a timeline - presumably about 2150 - assuming a passage from the current mortality to the limit mortality pattern. Initially, this approach produced a forecast of mortality for fictitious cohorts, i.e. year by year, as in the case of most forecasts, so that the results

[5]The prospects assumed by the so-called pessimistic scenarios differ considerably as these consider that life expectancy may decline due to the combined effect of epidemics even more lethal than AIDS, of environmental disasters due to climate change, a continuing devastating economic depression, etc.; however, experts consider that the probability of the occurrence of a scenario of this type is low (Vaupel 2003).

[6]Reference is made in particular to "Human longevity, individual life duration and growth of oldest-old population", International Union Scientific Study of Population (IUSSP), Montpellier, October 26 2000, and "Health, ageing and work. Strategies for the new welfare society in the larger Europe", 2nd Geneva Association Health and Aging Conference, Trieste, October 21-22 2004.

obtained in this phase can be compared with those returned by recent experience.[7]

Of all the forecasts produced in the last few years, in table 4 we have indicated the most recent forecast formulated by both ISTAT, which stretches to 2030 and RGS, which runs to 2050.[8] Although the second forecast is less recent (1996), it has been used to construct a demographic base for life annuity insurance and therefore for problems similar to those dealt with in this study. As can be noted, the expected increases in longevity at 2030 expressed in terms of life expectancy are very similar to each other as regards males and slightly less optimistic in our forecast as regards females. The differences at 2050 are to be ascribed to the fact that the RGS forecast envisages a future evolution in the context of a process of compression of mortality (hence the substantial stabilization of life expectancy already in 2030) while our forecast reflects the framework of the expansion of mortality theory. Fig. 1 shows the characteristics for females, for whom this process is even more marked than for males.

Forecast	**Males**		**Females**	
	2030	**2050**	**2030**	**2050**
Our computations	81.8	83.2	86.1	88.7
ISTAT*	81.4	n.a.	88.2	n.a.
RGS**	81.4	81.7	87.3	87.6

Table 4: Comparison between predicted period life expectancies at birth

Note: n.a. = not available, *central mortality scenario, **low mortality scenario.
Source: RGS (General Accounting Office) and ISTAT (2002).

The passage from the forecast of mortality by fictitious cohorts to that by birth cohorts is straightforward. As the results of the former cover a much longer period of time than the effective life span of the most recent cohort considered, all the data necessary to describe the future process of mortality of the cohorts born between 1940 and 2000 is available in the age/time matrix of the results of the forecast by fictitious cohorts, so that the corresponding life tables can be obtained immediately.

Therefore, the results of a forecast embrace gradually increasing periods of time as we move from one cohort to the next. In fact, for the first cohorts, the forecast refers only to mortality in the last age interval, while for the most recent ones it covers their entire life. This specific aspect of the scope of the forecast is clearly visible in figure 2 which shows, for each cohort, the future female life expectancy trend at each age expressed as an index number whose basis is the corresponding life expectancy of the ISTAT life table by fictitious

[7] However, we must take into account the fact that this type of very long-term forecast is based on a methodological approach which is different from that of "current" demographic forecasts, which usually cover a period of thirty/fifty years.

[8] ISTAT (2002) projects Italian population up to 2051. However, mortality is forecasted only up to 2030 and assumed to be constant in the following period.

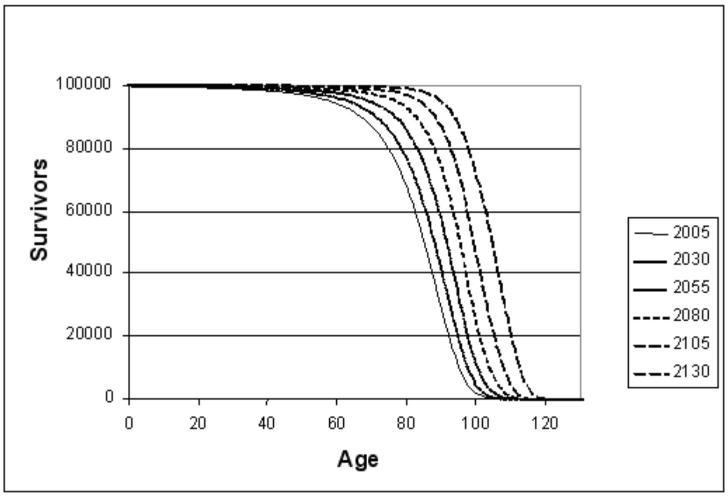

Figure 1: Forecasted period survival curves: females

Source: our computations.

cohort of 1990.[9]

As we can see in figure 2, the index numbers highlight the cumulative effects deriving both from the reduction in mortality that had already occurred between 1990 and 2000 - the "step" that is the starting point of the surface representing the trend of the index numbers - and from the expected future evolution of each cohort. From the 1990 ISTAT life table we can see that the expectation of life in old ages tends to decrease much more slowly than before and, as a consequence, the surface in figure 2 presents a "hunch" at about the age of 80.

Therefore, the results of the forecast suggest greater increases in survival once again for the female component of each cohort and, more generally, proportionally increasing gains as each cohort ages over the years and more noteworthy gains for the cohorts that are youngest today.

3 Money's worth measures of social security

3.1 The defined contribution pension formula

The 1995 pension reform changed the old DB pension system into a (N)DC one. Different rules are however applied to different workers, according to the seniority accrued in 1995. Workers with at least 18 years of seniority are exempt

[9]This comparison is made because the 1990 life table refers to the current rules for calculating the transformation coefficients to calculate pensions with the (N)DC method (see section 3.1) and provides a preliminary view of how the scenario changes using the prospective life tables for cohorts.

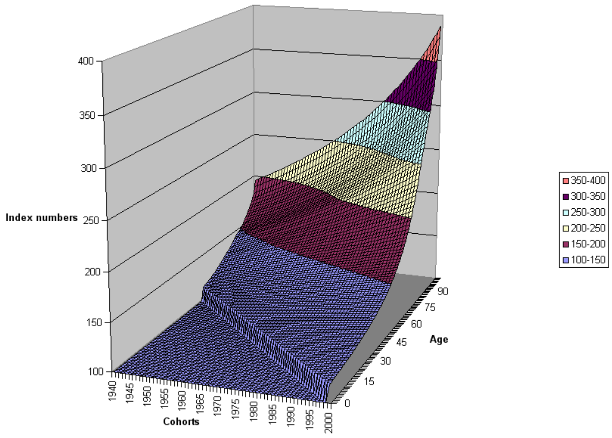

Figure 2: Forecasted life expectancy from 2000 onwards for the cohorts 1940-2000: females

Note: index numbers with base = 100 the life expectancy at the corresponding ages of the 1990 ISTAT life tables.
Source: our computations.

from the reform, younger workers are treated with a *pro rata* method, and workers enrolled from 1996 onward fall completely under the new provisions. This mechanism generated a very long and complicated transition to the new steady-state. The (N)DC pension formula, for retirement at age x, is given by:

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

where c_i is the contribution (in nominal prices) paid by the worker when seniority is i, a is seniority at retirement, $\bar{g}_j$ is the geometric mean of nominal GDP growth rate in the 5 years preceding the year in which seniority is j, and δ_x is the transformation coefficient for retirement at age x. The pension computation therefore requires two steps. In the first one, the individual accrued fund at retirement (the amount in squared brackets) is computed. It is given by the contribution paid in the last year of work, plus the contributions paid in the preceding working years, each of them capitalized at the various $\bar{g}_j$'s up to the year before retirement. In the second step the pension benefit is computed, multiplying the accrued fund at retirement times the transformation coefficient specific for the retirement age x.

The transformation coefficient δ_x is given by the inverse of the expected present value of an annuity of one euro revertible to the spouse, as in the following equation:[10]

[10]Equation (2) is a simplified version of the formula established by law. Its full version,

$$
\begin{aligned}
\delta_x &= \left(\frac{\sum_{s=m,f} dir_{x,s} + ind_{x,s}}{2} - k \right)^{-1} && \text{(2a)} \\
dir_{x,s} &= \sum_{t=0}^{\Omega-x} \frac{\ell_{x+t,s}}{\ell_{x,s}} (1+g_f)^{-t} && \text{(2b)} \\
ind_{x,s} &= \theta \sum_{t=0}^{\Omega-x} \frac{\ell_{x+t,s}}{\ell_{x,s}} \left(1 - \frac{\ell_{x+t+1,s}}{\ell_{x+t,s}}\right)(1+g_f)^{-(t+1)} a^W_{x+t+1} && \text{(2c)} \\
x &\in [57, 65] && \text{(2d)}
\end{aligned}
$$

where $\frac{\ell_{x+t,s}}{\ell_{x,s}}$ is the gender-s-specific survival probability at age $x+t$, conditional on being alive at age x, g_f is the (long-run) expected GDP growth rate, a^W_{x+t+1} 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 to take into account of different frequencies in pension payments (e.g. k is different if pensions are paid bimonthly instead of annually).

The normative provisions on the transformation coefficients more important for our analysis are summarized in the following points:

- life expectancies are assumed to be homogeneous within the population. Differences between genders are averaged-out in equation (2a);
- $\ell_{x,s}$ (and a^W_{x+t+1}) are based on cross-sectional mortality tables provided by ISTAT;
- the only dependant who is considered in the computation is the widow(er) (although the pension can be paid also to other survivors). Many assumptions, which we provide in appendix B, define $ind_{x,s}$. Between them, we underline the age-difference between the spouses (the wife is 3 years younger than her husband), and the quota of the pension revertible to the widow(er) ($\theta = 0.6$);
- there is a specific transformation coefficient for each retirement age between 57 and 65. Although this age bracket represents the retirement window established by the law for workers who fall under the new provisions, early workers - with at least 40 years of seniority - are allowed to retire before and exploit the age-57 coefficient. There is no mandatory retirement age, but retiring older than 65 implies the application of the age-65 coefficient;[11]

which we use in the computations, is shown in appendix B.

[11] Retirement before age 65 in the new system is allowed only if the following requirements are met: *a)* the accrued pension benefit is above a given threshold (1.2 times the social allowance), and *b)* 5 years of seniority are accrued. Some of these rules have been changed by the 2004 reform. See e.g. Marano and Sestito (2005) for details.

- to compute the pension of the workers entitled to retire in the period 1996-2005 according to the new rules, the reform established $g_f = 0.015$ and $\ell_{x,s}$ based on ISTAT 1990 mortality tables.[12] To account for their possible changes, a revision of the coefficients is scheduled every ten years. It is not an automatic application of the new mortality and macroeconomic data recorded by ISTAT in the previous decade, but an outcome of a political agreement between the Ministries, the Parliamentary Commissions and the Trade Unions. Importantly, the update only affects the flow of new pensions and does not modify the past stock.

3.2 The representative agents approach

As done by Ferraresi and Fornero (2000), we compute mwm (see e.g. Geanakoplos, Mitchell, and Zeldes 2000, Coile and Gruber 2000b and Coile and Gruber 2000a) for the Italian pension system following a representative agents approach.[13] We consider agents who represent the cohorts involved in the transitional phase toward the new rules and the cohorts of steady state. Although the focus of the analysis is more on differentials among cohorts, we include some intragenerational heterogeneity. In particular, we distinguish the agents born in the same year according to gender, occupation, sector, and geographical area.[14] We concentrate on private sector employees, because they represent the most numerous group of workers in Italy, and we consider them as representative of the whole population of workers.

We assume that each worker enters into the labor market at age 22 and does not interrupt the career until retirement. The latter assumption overestimates seniority, but is in part offset by a later entry into the labor market with respect to what descriptive statistics show.[15] Moreover, we took into account that many discontinuities in the working career are covered by notional contributions (e.g. motherhood, unemployment spells, sickness and disability, period of military service, and post-school education) and therefore these periods are accounted for to compute both minimum requirements and accrued benefits. Given our assumptions, and what is established by the law about minimum requirements,

[12] Due to the gradual implementation of the reform, almost nobody entitled to retire in the decade 1996-2005 fall automatically under the application of the NDC system. Although a voluntary switch from the DB to the (N)DC rules is allowed, to date only few people applied. Consequently, macroeconomic and demographic assumptions to compute transformation coefficients established for the decade 1995-2005 have been used seldom.

[13] A more technical description of the simulation model is provided in appendix C.

[14] In details, we distinguish males from females, blue from white collars, workers in the North-West, North-East, Center or South and Islands, workers in the industry, construction, services, or other sectors. Considering the number of cohorts involved in the computation (53 from the cohort 1947 to the cohort 2000) we end up with a total of 3392 representative agents. A full intragenerational analysis, performed on a real population of workers with heterogeneous careers, is performed by means of a microsimulation model in Borella and Coda (2005) and in Vagliasindi, Bianchi, and Romanelli (2004).

[15] Statistics we computed from the "Working History Italian Panel" - confirmed by The "Bank of Italy's Survey on Households, Income and Wealth" - show average entries into the labor market ranging - according to gender, cohort and occupation - between age 17 and 21.

workers are allowed to retire from age 57, regardless of the pension formula they are subject to.[16] These stylized life cycles are illustrated in table 5 for the different cohorts considered in this analysis.

cohort	**career start**	**seniority in 1995**	**pension formula**	**retirement period**
1947	1969	26	DB	2004-2009
...	...	...	...	...
1955	1977	18	DB	2012-2017
1956	1978	17	*pro rata*	2013-2018
...	...	...	...	...
1972	1994	1	*pro rata*	2029-2034
1973	1995	0	DC	2030-2035
...	...	...	...	...
2000	2022	0	DC	2057-2063

Table 5: Cohort differences in the pension regime

Notes: careers start at age 22; seniority in 1995 and the first retirement year are determined assuming continuous careers; the pension formula is determined by the seniority in 1995 as explained in section 3.1; the retirement period is defined considering a period of 6 years starting at the minimum requirements.
Source: our computations.

The average lifetime wage profile of each agent is estimated by means of a *Mundlak-type random effects model*. The logarithm of wage is regressed on an age spline, on a set of dummies for time, sector and geographical area, and on an individual effect capturing the cohort effect and unobserved heterogeneity.[17] The panel data model allows us to disentangle age, time, and cohort effects. Given the long-run horizon of the analysis, there may exist relevant lifetime financial wealth differences between the oldest and the youngest cohorts. Differences in lifetime financial wealth can offset differences in the social security wealth, induced by pension reforms. Therefore, measuring cohort effects is very important. The long run horizon requires to make predictions for young (out-of-sample) cohorts. Traditional econometric models, which use dummies to capture cohort effects, require additional assumptions for out-of-sample predictions. We therefore adopt an alternative modeling procedure, which follows the theory of Heckman and Robb (1985), and the application of Kaptein, Alessie, and Lusardi (2003) to The Netherlands. We assume that wages differ across cohorts only

[16]For workers under the DC regime age 57 is one of the requirements to retire (see section 3.1). Given the model assumptions, actually it is the only one binding for our DC agents at age 57 (the conditions on the accrued pension and on the seniority are satisfied at that age). For workers under the DB regime, instead, 35 years of seniority are required to access a seniority pension, regardless of their age. DB agents, thus, reach minimum requirements at age 57 because of the assumptions on their working career.

[17]In order to avoid the perfect multicollinearity trap between age, time and cohort, we restrict the time dummies to add up to zero and to be orthogonal to a linear trend, as suggested in Deaton and Paxson (1994).

due to the macroeconomic conditions when the individuals enter into the labor market. These conditions are summarized by productivity growth and are then approximated by GDP per capita. Dataset, estimates and statistical tests are described in appendix D.

In the left panel of figure 3 we plot the estimated profiles by age, and in the right panel the estimated cohort effects, i.e. the relation between GDP per capita and year of birth.[18] From the left panel we see that males have higher and more dynamic wages than females, and that the same is true for white collar *versus* blue collar. From the right panel we instead see that male white collar is the socioeconomic group with the strongest cohort effects, and that female blue collar is the group with the lowest.

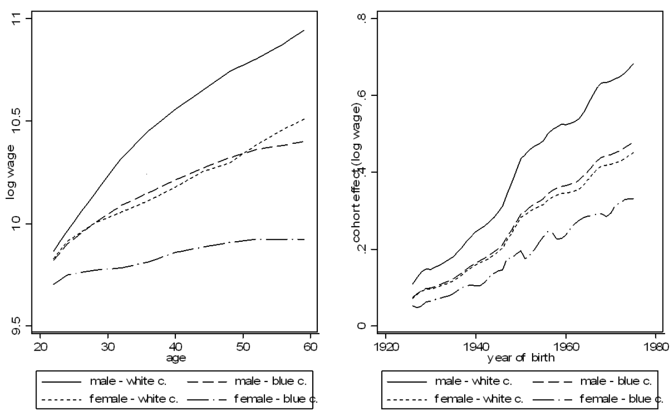

Figure 3: Wage estimates: age (left) and cohort effects (right)

Notes: selected representative agents. Log-wage is in thousands lira, prices 1995. Reference individual: North-West, other sectors, cohort 1949.
Source: our computations.

Computing pensions and mwm requires to set up a macroeconomic scenario covering a long period. While up to year 2004 we use historical series (e.g. about GDP growth rate, inflation, real wages growth rate, average return of public bonds, see table 11 for details), for the future we need to make long-run assumptions for some important variables. In particular, we assume a GDP growth rate equal to 1.5 percent and an interest rate equal to 2 percent. The assumptions we use are in line both with what was established by the 1995 reform, and with the assumptions of official simulations of the Italian pension expenditure (MLSP 2002a, MLSP 2002b MLSP 2002b). The spread of 0.5 points respects dynamic efficiency, but do not exaggerate the dominance of the

[18]While the increasing trends shown in the right panel reflect the economic growth over time, the short-time variations capture the recessions and allow for the identification of the cohort effects.

market with respect to the PAYGO return, given the deterministic scenario.

With respect to the way transformation coefficients are computed and updated to longevity changes, we consider the following two scenarios:

- *scenario 1* represents the *present legislation.* We use cross-sectional mortality tables to predict how the transformation coefficients of the 1995 reform will change in the next decades. To this aim we apply formula (10) in appendix B and we follow the rules described in section 3.1;
- *scenario 2* represents our *benchmark* against which to assess actuarial fairness and neutrality. We use longitudinal mortality tables to predict theoretical cohort-and-gender specific transformation coefficients for the next decades. To this aim we apply formula (11) described in appendix B.

The first scenario captures the effects of the normative changes in the pension formula on mwm during the transition to the steady state and the effects of the expected increase in longevity that young generations will benefit. The second scenario captures only the first effect, because the second is neutralized by a perfect (with respect to cohort and gender) adjustment in the coefficients. By a comparison between the two scenarios we can thus isolate the impact on money's worth of the second effect.

3.3 Actuarial principles and mwm *formulae*

We compute three mwm: social security wealth (SSW), net present value ratio ($NPVR$) and tax/subsidy rate (TAX). The first and the second are used to define and evaluate actuarial fairness, and the third to define and evaluate actuarial neutrality.

The *social security wealth*, computed when t_1 years of seniority are accrued and related to a "retirement plan" with a years of seniority, is given by the present value of pension benefits from retirement up to death (including a survivor's component) net of the present value of contributions to pay from t_1 up to the last year of work. Under the assumptions of the model, for $t_1 \in [1, a+1]$ and evaluated in t_2, it is given by:[19]

$$SSW^a_{t_1,t_2} = -\sum_{j=t_1}^{a} c^*_j (1+r)^{t_2-j} + \left[P(e+a) \frac{1}{\delta^{co}_{e+a,s}} \right] (1+r)^{t_2-(a+1)} \qquad (3)$$

[19]Financial flows are anticipated. When $t_1 = a + 1$, all c^*_j were already paid, and do not appear in the formula. Notice that, given that we are only interested in computing the indicator before (or exactly when) the flow of pensions starts, we provide the formula for $t_1 \leq a + 1$. A more general formula would be otherwise needed. In agreement with most of the literature (see e.g. Coile and Gruber 2000b), we assume that the individual is alive at retirement. Other studies, like Wilke (2005), take into account of mortality rates before retirement.

where c_j^* is the annual contribution (at constant prices) paid at the beginning of the year during which the worker accrues a seniority equal to j, P is the pension (paid at the beginning of each year, starting from retirement), e is the age of entry into the labor market, and r is the time-constant financial discount rate. $\delta_{e+a,s}^{co}$ is the transformation coefficient specific for age $e+a$, cohort co and gender s, computed according to formula (11) and restricting $g_f = r$. Notice how in this formula, and in the following ones in this section, transformation coefficients *formulae* are exploited twice. First, they are used to compute $P(e+a)$, as in (1). To this aim we use either formula (10) - if we compute the pension in scenario 1 - or formula (11) - if we compute it in scenario 2. Second - and regardless of the scenario - the inverse of formula (11) is used to compute the present value of pension benefits.

The *net present value ratio* is given by the ratio between the present value of pensions received from retirement up to death (including a survivor's component) and the present value of contributions paid throughout the whole working career. It has the intuitive interpretation of how much the system returns back to the worker for each euro paid. This indicator, as well as the previous one, measures therefore the generosity of the pension system, i.e. its *global* incentives. $NPVR$ facilitates intergenerational comparisons because - on the opposite to SSW - is not affected by the economic growth. Under the assumptions of the model, and when seniority at retirement is equal to a, it is given by:

$$NPVR^a = \frac{P(e+a)\dfrac{1}{\delta_{e+a,s}^{co}}}{\sum\limits_{j=1}^{a} c_j^*(1+r)^{a+1-j}} \tag{4}$$

If $NPVR = 1$ we define the pension system actuarially fair.[20]

To introduce the third indicator, which measures *marginal* incentives to retire, we need first to define the marginal variation in SSW, called *accrual*. If the "retirement plan" is postponed by one year, from $a = a'$ years of seniority to $a = a' + 1$, it is given by:

$$Accr_{t_2}^{a'} = SSW_{t_1,t_2}^{a'+1} - SSW_{t_1,t_2}^{a'} \tag{5}$$

Defining $SSW_{t_1,t_2}^{a'+1}$ and $SSW_{t_1,t_2}^{a'}$ according to (3), and substituting them into equation (5), we obtain:

[20] Disney (2004) asserts that an actuarially fair pension system accumulates the individual's contributions up to the retirement date, and the present value of these contributions is equal to the present value of the pension benefits paid both to the contributor in retirement and to any dependant who is eligible for a survivor's pension. In order to fulfill these conditions the pension system should *i)* provide a pension accrual which is formula-based such as to be proportional to contributions, with the accrual rate constant across earnings, individuals and cohorts, and *ii)* reevaluate the accumulated contributions, and the pension in payment, with an interest rate that matches the rate of return that the participant could have obtained by investing in a risk-free asset in the market. Both Disney (2004) and Brown (2000) stress how it also should take into account of any *ex-ante* differentials in mortality across socioeconomic groups in the population.

$$
\begin{aligned}
Accr_{t_2}^{a'} &= -c^*_{a'+1}(1+r)^{t_2-(a'+1)} + \\
&\left[P(e+a'+1)\frac{1}{\delta^{co}_{e+a'+1,s}} - P(e+a')\frac{1}{\delta^{co}_{e+a',s}}(1+r)\right] \\
&(1+r)^{t_2-(a'+2)} \qquad (6)
\end{aligned}
$$

From formula (6) it can be seen how, if retirement is postponed by one year, SSW varies for two reasons. First, it decreases because of the additional contribution to pay. Second, it may increase due to the difference between the present values of pensions determined by the alternative retirement options (indicated between squared brackets). The sign of the difference in squared brackets is however undefined, because two effects counteract there: a (generally) higher pension and a shorter expected retirement period. In order to have an accrual equal to zero, the second effect must be positive and equal to the first negative effect.[21] To make the indicator independent on the level of wages, the accrual is divided by the expected wage for the additional year of work. If the "retirement plan" is postponed by one year, the TAX computed in t_1 and evaluated in t_2 is thus given by:

$$
TAX^{a'}_{t_1,t_2} = \frac{-Accr^{a'}_{t_2}}{E_{t_1}\left[w_{a'+1}\right](1+r)^{t_2-(a'+1)}} \qquad (7)
$$

If $TAX = 0$ we define the pension system actuarially neutral. If TAX is positive (i.e. *Accr* is negative), the pension system favors early retirement by imposing an implicit taxation on the continuation of the working activity.

4 Computing mwm with dynamic mortality tables

4.1 Forecasts of the transformation coefficients

Table 6 shows the forecasted transformation coefficients for the next decades, according to the rules established by the 1995 reform (scenario 1). It highlights how coefficients will be substantially reduced, if the expected increases in longevity actually occur. With respect to the coefficient of a worker retiring in the first decade of application of the reform, for example, the coefficient of a worker retiring in the decade 2006-2015 will be 6-8 percent lower, and that of a worker retiring in the decade 2046-2055 will be even 20-24 percentage points lower. Our findings are very similar to those in Caselli, Peracchi, Balbi, and

[21] Basically, the individual who continues to work for an additional year on one side "looses" because of 1) the additional contribution to pay, 2) the pension not received in the first year, and 3) the expected retirement period is then shorter. On the other side, she "gains" because of the (likely) higher pension received throughout the retirement period.

Lipsi (2003).[22] Notice, however, that a reduction in the forecasted coefficients does not imply a reduction in the expected pension. Due to the gradual implementation of the reform, in fact, most of the workers retiring in the next (say two) decades will be paid a DB or a *pro rata* pension. According to the assumptions of our model, in particular, DC pensions will be paid to workers retiring from 2030 (see table 5). For workers subject to the *pro rata* rules the expected reduction in the pension is lower than the reduction in the forecasted coefficients, and for workers subject to the DB rules there is no expected reduction at all.

			Retirement year			
Age	**L. 335/95**	**2006-15**	**2016-25**	**2026-35**	**2036-45**	**2046-55**
57	4.72	4.422	4.231	4.067	3.925	3.799
58	4.86	4.543	4.342	4.169	4.019	3.886
59	5.006	4.673	4.459	4.277	4.118	3.978
60	5.163	4.811	4.584	4.391	4.223	4.075
61	5.334	4.957	4.717	4.512	4.335	4.178
62	5.514	5.114	4.859	4.641	4.453	4.287
63	5.706	5.281	5.01	4.779	4.579	4.403
64	5.911	5.46	5.17	4.925	4.712	4.526
65	6.136	5.651	5.343	5.081	4.855	4.658

Table 6: Forecasted transformation coefficients by age and retirement year: present legislation

Notes: percentage points; scenario 1 (projected cross-sectional mortality tables, formula 10, assumptions in appendix B); $g_f = 0.015$.
Source: our computations.

Table 7 shows the forecasted transformation coefficients according to our benchmark (scenario 2). It highlights how heterogeneous the transformation coefficients would be, were they computed consistently with the cohort and gender differentials in life expectancy. With respect to the coefficient of a worker born in 1948, for example, the coefficient of a worker born ten years later would be 5 percent lower, while that of a worker born in 1998 even 20 percent lower. Males, who are expected to live shorter than females, would have a higher coefficient. Percentage differences in the coefficients, from 3 to 7 points according to the cohort and to the retirement age, do not however completely reflect the differences in life expectancy. The provision of a joint annuity, in fact, dramatically reduces the gender gap in the coefficients which would exist if only the life expectancies of the pensioner - and not of the widow(er) - were taken into account.[23]

Table 8 finally shows, for each gender and for each age, the percentage

[22] The authors forecast the coefficients up to 2020 using both ISTAT 1990 and ISTAT 1997 mortality tables. They find that, if the more recent tables were used in place of the older ones, coefficients would have to be revised by 3 percent.

[23] Given that in formula (10) $\Psi = 0.6$, the differences in life expectancy between genders is not completely offset by the opposite differences in the survivor's component of the annuity.

	Cohort					
Age	**1948**	**1958**	**1968**	**1978**	**1988**	**1998**
			Males			
57	4.071	3.918	3.787	3.673	3.571	3.478
58	4.167	4.008	3.871	3.751	3.644	3.547
59	4.269	4.103	3.959	3.834	3.722	3.619
60	4.376	4.203	4.052	3.921	3.803	3.696
61	4.49	4.308	4.151	4.012	3.889	3.776
62	4.61	4.42	4.254	4.109	3.98	3.861
63	4.737	4.538	4.364	4.211	4.075	3.95
64	4.872	4.662	4.48	4.319	4.176	4.045
65	5.016	4.795	4.603	4.434	4.283	4.145
			Females			
57	3.949	3.791	3.649	3.518	3.395	3.275
58	4.039	3.873	3.724	3.587	3.458	3.333
59	4.135	3.96	3.804	3.66	3.524	3.393
60	4.235	4.052	3.887	3.736	3.594	3.456
61	4.342	4.149	3.976	3.817	3.668	3.523
62	4.455	4.251	4.069	3.902	3.745	3.593
63	4.575	4.36	4.167	3.991	3.826	3.667
64	4.702	4.475	4.271	4.086	3.912	3.744
65	4.837	4.596	4.382	4.186	4.003	3.827

Table 7: Forecasted transformation coefficients by age, cohort and gender: benchmark

Notes: percentage points; selected cohorts; scenario 2 (cohort-and-gender-specific mortality tables, formula 11, assumptions in appendix B); $g_f = 0.015$.
Source: our computations.

differences between the transformation coefficients computed according to the present legislation (table 6) and those computed according to the benchmark (table 7). To define the retirement years of each cohort we assume continuous careers. Given that the figures in the table are positive for both genders and for every cohort, it turns out that the Italian pension system provides an excess return with respect to what the benchmark would grant. This result can be explained by the use of cross-sectional, instead of longitudinal, mortality tables to measure life expectancies. Given that the number of years spent on average by each individual into retirement is long and given that the pension benefit is not revised (according to mortality changes) afterward, the system provides more resources than what the benchmark indicates. For females a further element applies: they have the same coefficients of males but higher life expectancies.

	Retirement year					
Age	**2005**	**2015**	**2025**	**2035**	**2045**	**2055**
			Males			
57	16	12.8	11.7	10.8	9.9	9.2
58	16.1	12.9	11.8	10.8	10	9.3
59	16.3	13	11.9	10.9	10	9.3
60	16.5	13.1	11.9	10.9	10.1	9.3
61	16.7	13.2	12	11	10.1	9.3
62	16.9	13.4	12.1	11.1	10.1	9.4
63	17.1	13.5	12.2	11.1	10.2	9.4
64	17.3	13.6	12.3	11.2	10.2	9.4
65	17.6	13.7	12.4	11.2	10.3	9.5
			Females			
57	19.5	16.6	15.9	15.6	15.6	16
58	19.8	16.8	16.1	15.8	15.8	16.2
59	20	17	16.3	16	16	16.3
60	20.2	17.2	16.5	16.2	16.1	16.5
61	20.5	17.4	16.7	16.3	16.3	16.7
62	20.7	17.6	16.9	16.5	16.5	16.9
63	20.9	17.7	17	16.7	16.7	17
64	21.1	17.9	17.2	16.9	16.9	17.2
65	21.4	17.9	17.4	17	17	17.4

Table 8: Transformation coefficients: percentage differences between present legislation (table 6) and benchmark (table 7)

Notes: selected retirement years; retirement year for each cohort and retirement age is defined assuming continuous careers.
Source: our computations.

4.2 An assessment of the actuarial fairness

The net present value ratio by cohort in the two scenarios is shown in figure 4. There we focus on one type of representative agent for each cohort (male, blue collar, industry, North-West), and we assume she retires at minimum requirements (at age 57, with 35 years of accrued seniority). Old cohorts receive back from the DB system up to 1.5 times what they paid. The $NPVR$ exhibits a decreasing trend along the transitional phase, showing that the return is reduced by the gradual implementation of the new rules. The system in steady state is almost actuarially fair. Precisely, the figure shows a NPVR slightly less than one. This last result follows from the macroeconomic assumptions on the economic growth and on the interest rate. The NDC system gives an excess return compared to the benchmark (table 8), but it brings it back in terms of the financial opportunity-cost of being in the system, even in our "conservative" macroeconomic scenario.

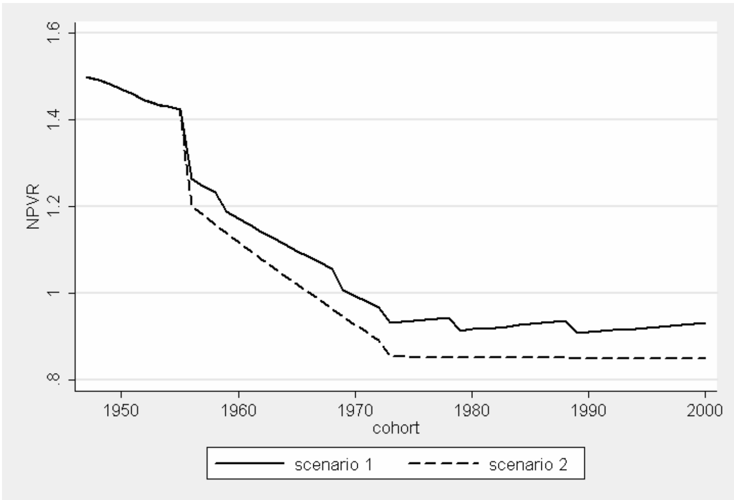

Figure 4: Net present value ratio by cohort: alternative scenarios

Notes: scenario 1: present legislation, scenario 2: actuarial benchmark; representative agent: male, blue collar, industry, North-West; 35 years of seniority; $g_f = 0.015, r = 0.02$.
Source: our computations.

The $NPVR$ in scenario 1 shows two kinds of discontinuities, determined by the reform provisions. The first one is due to the rigid classification of workers, based on the seniority in 1995 (see Ferraresi and Fornero 2000 and Fornero and Castellino 2001), while the second is due to the revision of the transformation coefficients and is more interesting for this study. The $NPVR$ in scenario 2, instead, shows results very similar to Ferraresi and Fornero (2000): by offsetting cohort-and-gender longevity changes by means of changes in the transformation coefficients, we end up in a static mortality scenario similar to their one.

Focusing on the second kind of discontinuity, the figure shows how the combined effects of continuous mortality changes and discrete adjustments in the coefficients generate a kind of cycle. Its values and fluctuations can be measured by means of the difference between the $NPVR$ in scenarios 1 and 2. Two points are worth noting. First, there is a difference, equal to 6 percentage points of $NPVR$, existing in every year of the simulation (excluded the first years, still under the DB rules). It is caused by the use of cross-sectional instead of longitudinal mortality tables in the computation of the coefficients, and confirms the results of table 8. Second, there is a ten-years fluctuation, caused by the delay with which the coefficients are updated. The cycle reaches its maximum in the year before the revision of the coefficients ($3 + 6 = 9$ points of $NPVR$), and is at its minimum immediately afterward (6 points). Based on this empirical evidence, we can thus rank the two mentioned causes of distortion from the benchmark. Not incorporating cohort effects when measuring longevity changes counts for two thirds (i.e. 6 points out of 9) of the total distortion, while the delay in the revision of the coefficients counts for the residual one third.[24]

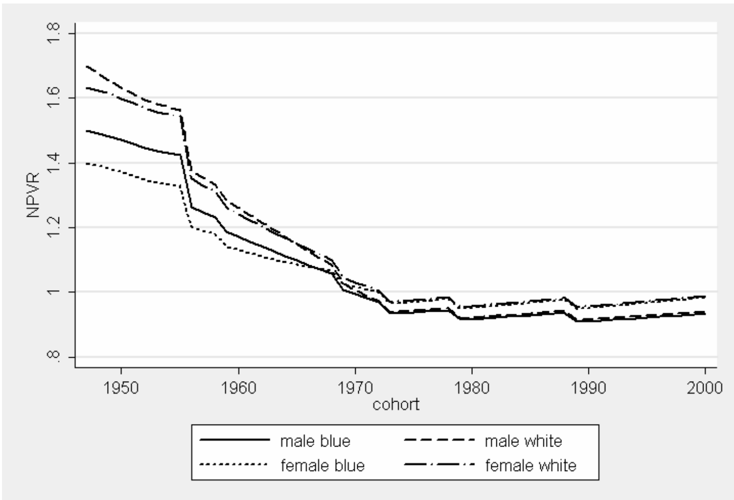

Figure 5: Net present value ratio by cohort, gender and occupation: present legislation

Notes: scenario 1; representative agent: industry, North-West; 35 years of seniority; $g_f = 0.015, r = 0.02$.
Source: our computations.

Exploiting the intragenerational heterogeneity of the model, we can shed some light on actuarial fairness within cohorts. Figure 5 shows the $NPVR$ in scenario 1 for four representative agents: male blue collar, male white collar,

[24]This result finds a confirmation at macroeconomic level in MLSP (2002a). The study highlights how pension expenditure would not substantially change, if coefficients were revised yearly instead of every ten years.

female blue collar and female white collar. It highlights how the reform, though globally reduced intragenerational redistribution, changed its nature from *wage-based* to *mortality-based*. In particular white collars, characterized by a steep wage profile, are favored by the weak correlation between contributions and pensions typical of the DB formula in the old system, but are treated as well as blue collar in the (N)DC one. Females, penalized with respect to males by their flatter wage profiles in the DB system, are instead favored in the DC one.

Finally, figure 6 shows social security wealth by cohort, gender and occupation. It underlines how the economic growth benefited by the younger generations can offset the restrictive effects of the 1995 reform. In particular, the SSW of male white collars - the socioeconomic group characterized by the strongest cohort effects - will be even higher for younger than for older workers.

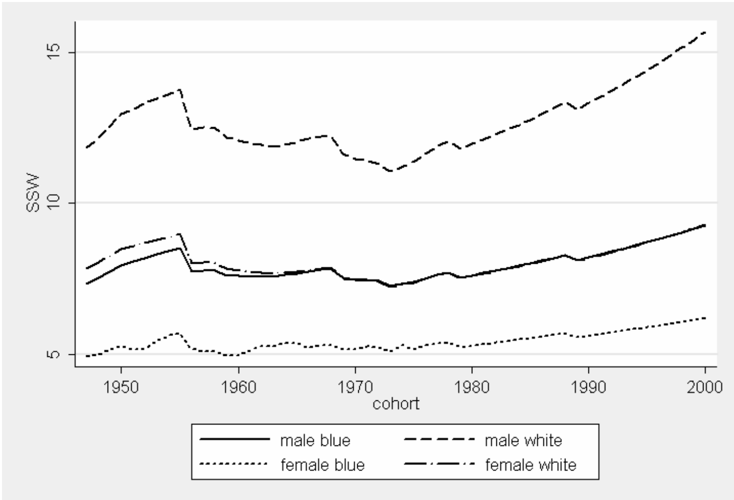

Figure 6: Social security wealth by cohort, gender and occupation: present legislation

Notes: thousand millions lira, prices 1995; scenario 1; representative agent: industry, North-West; 35 years of seniority; $g_f = 0.015, r = 0.02$.
Source: our computations.

4.3 An assessment of the actuarial neutrality

Figure 7 shows the tax rate by cohort in different scenarios. Implicit taxation is extremely high in the DB system, reaching a value of 50 percent. Workers subject to the old rules are therefore strongly encouraged to leave the labor force as soon as possible. The *pro rata* mechanism before, and the complete application of the new rules then, drastically reduce the implicit taxation to a maximum value of 4 percent. Retirement choices in the NDC system are therefore almost actuarially neutral.

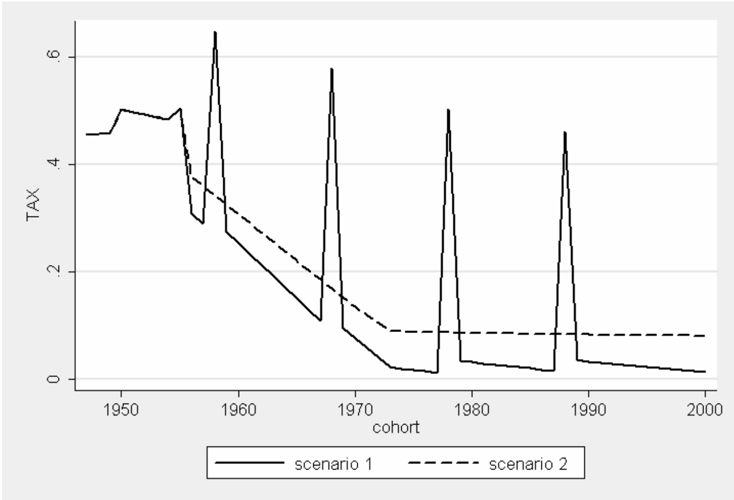

Figure 7: Tax rate by cohort: alternative scenarios

Notes: scenario 1: present legislation, scenario 2: actuarial benchmark; representative agent: male, blue collar, industry, North-West; 35 years of seniority; $g_f = 0.015, r = 0.02$.
Source: our computations.

There are, however, some important exceptions to this general result. First, a worker who has to decide whether to retire or to continue to work in the year before the revision of the transformation coefficients faces a very strong constraint. Her pension and her social security wealth, in fact, would be considerably cut if she continued to work. Scenario 1 is, in fact, dominated by spikes of taxation of 30-40 percentage points, even in steady state. Second, as far as dynamic efficiency holds (and there are no transaction costs or market failures), it is efficient to retire at minimum requirements and to invest the accumulated wealth at the market return. Our macroeconomic assumptions explain thus the residual 4 percent implicit taxation described above. Third, as already mentioned, retirement is allowed also outside the age bracket 57-65. Even if the assumptions of the model exclude these cases from the analysis and thus we do not provide any empirical evidence, it is worth mentioning that retirement choices are, for them, not actuarially neutral. Early workers (who have accrued 40 years of seniority) can exploit the age-57 coefficient and are thus favoured, while those who retire older than age 65 are penalized by the application of the age-65 coefficient.[25]

[25] A minor distortion is highlighted by comparing the alternative scenarios. Apart in the years characterized by the spikes, taxation is lower in scenario 1 than in scenario 2. The present legislation gives an extra-return - with respect to what is actuarially neutral according to our benchmark - to each additional year of work. In other words, the difference between coefficients at subsequent ages is upward biased, because the "correct" conditional mortality probability at each age includes longitudinal differences in mortality.

Figure 8 shows intragenerational differences in taxation. Results on actuarial neutrality parallel those on actuarial fairness under this respect. Both in the old and in the new system, workers with a more dynamic career (typically males and white collars) as well as those who live longer (typically females) have a lower taxation. Nevertheless, the NDC system gives a much higher weight to the *mortality-based* differences than what the DB one does.[26]

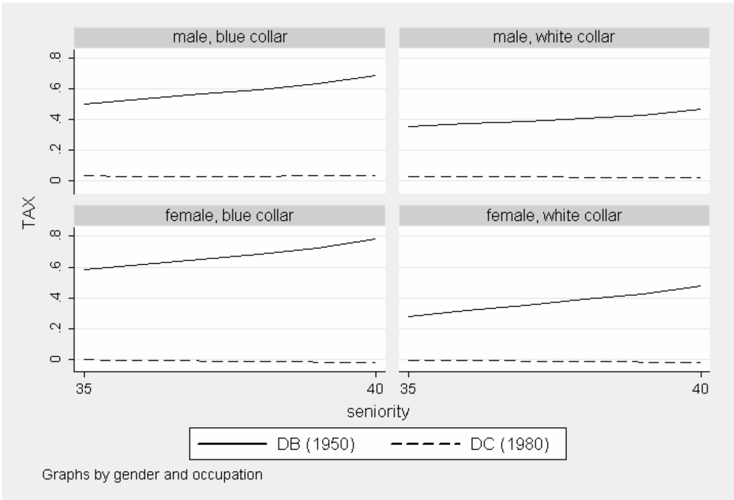

Figure 8: Tax rate by seniority, gender and occupation: selected cohorts

Notes: scenario 1; 35 years of seniority; $g_f = 0.015, r = 0.02$.
Source: our computations.

Finally, we briefly consider - although in an informal way - the interrelations between the two actuarial characteristics we examine in this study, i.e. fairness and neutrality. In general, the two principles are quite different, and a pension system can be close to one of them but far from the other.[27] However, from our results emerges how for the Italian pension system this is not the case. While the old DB scheme is far from both of the principles, the *pro rata* method before and the full application of the new (N)DC rules afterward made it pretty close to both of them.

It is also worth underlying the implications of actuarial neutrality for the actuarial fairness of the system. For this purposes, instead of focusing on inter-generation comparisons between representative agents retiring at the same age,

[26] Figure 8 also shows, for the cohorts 1950 (DB) and 1980 (DC), how taxation varies as the representative worker accrues seniority. It highlights how, while in the old system the implicit taxation increases as the worker continues to work, in the new one it keeps almost constant.

[27] For example, Disney and Johnson (2001) show how the UK public pension system is almost actuarially neutral but far from being actuarially fair. On the contrary, Börsch-Supan and Schnabel (1999) show that the German pension system before the 1992 reform was almost actuarially fair but not neutral.

we evaluate how the degree of actuarial fairness for each representative agent varies according to her retirement age. If the system is almost actuarially neutral - i.e. the social security wealth computed for a given agent for alternative “retirement plans” remains constant - as occurs with the new DC rules, there is almost no change in the degree of actuarial fairness of the agent when the retirement date is postponed. If instead the system heavily penalizes the continuation of the working activity, as in the old DB scheme and in the new one when transformation coefficients are revised, the degree of actuarial fairness of the agent remarkably reduces if retirement is postponed. For example, the 1950 cohort, characterized by a tax rate between 55 and 82 percent (depending on the age of retirement, gender and occupation), faces a decline in $NPVR$ from 1.5 (for retirement at age 57) to 1.2 (for retirement 6 years later).

5 Conclusions

Particularly important for the NDC systems, in order to insure actuarial fairness and neutrality, are the rules which establish how to incorporate into the pension *formulae* life expectancies and their changes. In this study we investigate these issues focusing on the Italian NDC scheme, introduced by the 1995 reform and gradually taking the place of the previous DB one. Our model exploits cohort-and-gender-specific mortality projections to establish a benchmark against which actuarial fairness and neutrality can be assessed.

The NDC system is almost actuarially fair. However, some noticeable distortions from the benchmark exist. A first one, concerning the methodology used to compute the pension, comes from the use of mortality tables which do not incorporate longitudinal trends in mortality. Another one comes from the delay with which pensions are revised in order to adapt to the increased longevity. We find evidence that the first distortion is more important than the second.

Looking at the reform from the point of view of the allocation of risk, two contrasting effects emerge. The reform, linking the benefit to the contributions, reduces the often perverse wage-based redistribution of the DB scheme. Introducing risk-pooling, at the same time, it brings about a potential mortality-based redistribution. Whenever the rich individuals are also the longer living, it redistributes in a wrong direction. Our mortality tables do not distinguish mortality between different levels of income (or wealth), and thus we cannot analyze these aspects in details. Within the simplified intragenerational framework of our model, we can however provide some evidence on the redistribution between genders. Females take advantage (but the provision of a joint annuity mitigates the gain) from having, *ceteris paribus,* the same pension of males in spite of a higher life expectancy. A justification of this favorable treatment has to be found within a public pension system which focuses on the family, instead on the single individual.

The new pension formula is almost actuarially neutral. However, the revision process introduces discontinuities of treatment. Due to the delay, in fact, a relevant expected cut in the pension level - as well as in social security wealth

- strongly constraints the retirement choice of workers in the previous year.

Interpretation of the results should take into account the limits of the analysis, suggesting further refinements and research. First, the model allows only for a simplified intragenerational analysis. Studying both early workers and workers with discontinuous careers, for example, can be more consistently done within a microsimulation model. Moreover, it is very important to provide a sensitivity analysis to the main macroeconomic and demographic assumptions - proposing alternative high and low mortality scenarios. Finally, the evidence we find crucially depends on the choice of the benchmark against which actuarial fairness and neutrality are compared. Working on its definition, possibly in a interdisciplinary perspective, is surely worth of further research.

Appendixes

A Mortality forecasts: methodology

As already mentioned in section 2.2, our mortality forecast considers a scenario based on the assumption that one of the limits of the human species is a life expectancy at birth of 110 years and a maximum life span of 120 years. One of the first problems to be addressed is that of converting these indications into a survival function that makes it possible to achieve the level of analytical detail by age required to construct a forecast.

The limit characteristics of mortality indicated above are therefore considered in order to describe endogenous mortality whose trend, as occurred previously in other experiments (Duchêne and Wunsch 1993, Maccheroni 1998), can be represented using 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:

$$\ell(x) = exp\left[-\left(\frac{x-a}{m}\right)^b\right] \qquad x \geq a; a \geq 0; b, m > 0 \tag{8}$$

The derivative of (8) 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 parameter type, permits fairly effective control 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 of (8) (Maccheroni 1998).

Having established the limit probabilities of dying due to endogenous causes $q_{end}(x)$ - i.e. linked to deterioration of the body's power of resistance, drawn from (8) - we construct the limit probabilities of dying for 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,[28] 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 on the empirical function obtained in this way.

The limit probabilities of dying $q_{lim}(x)(x = 0, 1, ...)$, can be obtained as the suitably perequated sum of the discretized version of the original components $q_{end}(x)$ and $q_{eso}(x)$, thus obtaining the complete table of which only the

[28] Norway, Holland, Belgium, Spain, France, Finland, Italy, Great Britain, Sweden, Austria, Canada, USA, Australia, Japan, New Zealand.

abridged survival function is shown in table 9. The synthetic characteristics of this mortality model are $e_0 = 109.4$ and an extreme age of life ω of just over 125 years.

$\mathbf{Age_x}$	$\ell_\mathbf{x}$	$\mathbf{Age_x}$	$\ell_\mathbf{x}$
0	100000	*65*	99642
1	99997	*70*	99566
5	99992	*75*	99454
10	99983	*80*	99280
15	99968	*85*	98985
20	99948	*90*	98242
25	99923	*95*	95017
30	99894	*100*	85236
35	99863	*105*	67726
40	99831	*110*	42180
45	99799	*115*	15306
50	99768	*120*	1980
55	99735	*125*	40
60	99695		

Table 9: Limit life table: survivors from 100000 live births

Source: our computations.

The following two steps are then necessary in order to produce the forecast:

1. establishing a time frame for the scenario provided by the $q_{lim}(x)$ in relation to the recorded Italian mortality trends. With regard to the above, a procedure based on the logit model (Brass 1971, Zaba 1979) is used. That is to say, the limit survival function $\ell_{lim}(x)$ obtained from the $q_{lim}(x)$ is taken as standard and, using the observed survival functions (Maccheroni and Locatelli 1999) $\ell_{x,t}(t = 1940, 1941, , 1998)$, the historical series of parameters a_t and b_t of the following relationship are studied:

$$Y_{x,t} = a_t + b_t Y_x^{lim} \tag{9}$$

   where Y_x^{lim} is the logit of the limit life table and $Y_{x,t}$ that 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 of parameters a_t^* and b_t^* (in order to establish if and when $a_t^* \to 0$ and $b_t^* \to 1$ when t diverges). In our case, these results are obtained when $t = 2143$ for females and when $t = 2170$ for males.

   Once again on the basis of (9), it is then possible to obtain the succession of the projected life tables that reflect the characteristics of the limit situation in the period of time that stretches from 1999 to the two previous extremities of time.

2. linking the mortality models, obtained as explained above (characterised by evidently innovative structural aspects compared with the present situation), with the on-going process of evolution of mortality because the limit scenario used (table 9) is more or less "disconnected" from this. According to the main hypotheses outlined in section 2.2, the gap between current mortality and the limit situation are bridged by a type of evolution that reflects the theory of expansion of mortality (Myers and Manton 1984).

   This step is carried out in two phases. The first phase defines the evolution of mortality resulting from the more 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 obtained by envisaging first a stationary situation, then a tendentially decreasing evolution. The second phase synthesizes for each year the projections of the two mortality models - obtained both from the logit relationship (9) and from the above-mentioned model of the projection - using an average constructed by generally assigning gradually increasing linear weights to the life tables projected obtained from (9) and decreasing linear weights to those obtained with the exponential model.

   Therefore, the projected life tables resulting from this synthesis (which were further perequated) produce a new succession of life tables that, initially, take into account recent mortality trends and gradually - approaching the two end dates of the forecast for males and females - assume the characteristics of the limit life table.

B Transformation coefficients *formulae*

Transformation coefficients in scenario 1 (present legislation) are computed according to the following formula:

$$\begin{aligned}
\delta_{e+a} &= \left(\frac{\sum\limits_{s=m,f} dir_{e+a,s} + ind_{e+a,s}}{2} - k \right)^{-1} \\
dir_{e+a,s} &= \sum_{t=0}^{\Omega-e+a} \frac{\ell_{e+a+t,s}}{\ell_{e+a,s}} (1+g_f)^{-t} \\
ind_{e+a,s} &= \Psi\Phi_s \sum_{t=0}^{\Omega-e+a} \frac{\ell_{e+a+t,s}}{\ell_{e+a,s}} (1 - \frac{\ell_{e+a+t+1,s}}{\ell_{e+a+t,s}})(1+g_f)^{-t}\Theta_{e+a+t,s} \\
& \sum_{\tau=1}^{\Omega-e+a-t+\varepsilon_s} \frac{\ell_{e+a+t+\tau-\varepsilon_s,-s}}{\ell_{e+a+t+1-\varepsilon_s,-s}} (1 - \ell^{ved}_{e+a+t+\tau-\varepsilon_s,-s})(1+g_f)^{-\tau}
\end{aligned} \tag{10}$$

while in scenario 2 (actuarial benchmark) they are computed according to the following one:

$$
\begin{aligned}
\delta^{co}_{e+a,s} &= (dir^{co}_{e+a,s} + ind^{co}_{e+a,s})^{-1} \\
dir^{co}_{e+a,s} &= \sum_{t=0}^{\Omega^{co}-e+a} \frac{\ell^{co}_{e+a+t,s}}{\ell^{co}_{e+a,s}}(1+g_f)^{-t} \\
ind^{co}_{e+a,s} &= \Psi\Phi_s \sum_{t=0}^{\Omega^{co}-e+a} \frac{\ell^{co}_{e+a+t,s}}{\ell^{co}_{e+a,s}}(1-\frac{\ell^{co}_{e+a+t+1,s}}{\ell^{co}_{e+a+t,s}})(1+g_f)^{-t}\Theta_{e+a+t,s} \\
&\quad \sum_{\tau=1}^{\Omega^{co}-e+a-t+\varepsilon_s} \frac{\ell^{co+\varepsilon_s}_{e+a+t+\tau-\varepsilon_s,-s}}{\ell^{co+\varepsilon_s}_{e+a+t+1-\varepsilon_s,-s}}(1-\ell^{ved}_{e+a+t+\tau-\varepsilon_s,-s})(1+g_f)^{-\tau}
\end{aligned}
\tag{11}
$$

Symbols in *formulae* (10) and (11) - together with those used in appendix C - are described in table 10.

Some of the parameters value have been fixed by law. In particular:

$$
\varepsilon_s = \begin{cases} +3 & \text{if} \quad s = m \\ -3 & \text{if} \quad s = f \end{cases} \qquad \Phi_s = \begin{cases} 0.9 & \text{if} \quad s = m \\ 0.7 & \text{if} \quad s = f \end{cases}
$$

$$
\Psi = 0.6 \qquad g_f = 0.015 \qquad k = 0.42 \ \text{(bimonthly anticipated)}
$$

$$
\ell_{e+a,s} = \text{ISTAT 90 mortality tables (if} co + e + a \leq 2005)
$$

Additionally we assume: 1) $\ell_{e+a,s}$ (if $co + e + a > 2005$) are obtained by projected cross-sectional mortality tables, and 2) $\ell^{ved}_{e+a,s}$ and Φ_s are constant over time.

C Description of the simulation model

The simulation model computes pensions and money's worth measures for a set of representative workers, using the rules of the present legislation. The model structure is also flexible enough to incorporate alternative pension regimes and therefore it allows to evaluate the effect of policy interventions on money's worth. It focuses on employees - both in the private (FPLD fund) and in the public sector (INPDAP fund) - while it does not incorporate self-employed. It provides a quantification of the impact on pension expenditure of various policy interventions which affect both the minimum requirements to retire and the benefit amount. The model is constituted by four modules: *data*, *control variables, computation,* and *aggregation*. Symbols are described in table 10.

symbol	**description**
$s, -s, co, e, a, a_{95}, \Omega/\Omega^{co}$	Gender (m=male, f=female), gender of the widow(er), cohort, age of entry into the labor market, seniority at retirement, seniority in 1995, life span according to cross-sectional/cohort-and-gender-specific mortality tables.
g_f	Expected long-run GDP growth rate.
$\ell_{x,s}/\ell_{x,s}^{co}$	Survivors of age x from 100.000 live births of gender s, according to cross-sectional/cohort-and-gender-specific mortality tables.
$\ell_{x,s}^{ved}$	Probability for the widow(er) of age x and gender s to marry again.
$\Theta_{s+t,s}$	Probability for the widow(er) of age $x+t$ and gender s to leave a family.
ϵ_s	Age-difference between the pensioner (of gender s) and the widow(er).
Ψ	Quota of the pension revertible to the widow(er).
Φ_s	Reduction in the widow(er)'s pension due to her(his) additional income (earning-test on widow(er)'s income).
k	Actuarial adjustment factor to take into account of different frequencies in pension payments.
$\overline{g}_x$	Geometric mean of nominal GDP growth rate in the 5 years preceding the year in which seniority is x.
r	Real interest rate.
w_x	Wage in constant prices when seniority is x.
w_x^*/w_x^{**}	Wage in nominal prices when seniority is x revalued according to the "quota A"/"quota B" law coefficients.
c_x/c_x^*	Contribution in nominal/constant prices when seniority is x.
δ_x	Transformation coefficient for retirement at age x.

Table 10: symbols

C.1 The *data* module

It is constituted by two sub-modules: *inputs* and *wage profiles*. The main macroeconomic and demographic inputs - which are passed to the computation module - are described in the text. The estimation of wage profiles is illustrated in appendix D. Data and sources are shown in table 11.

C.2 The *control variables* module

It assigns values to the following (global) variables:

- *gender*;
- *cohort*;
- *age of entry into the labor market*: constant within the population or conditional on cohort, gender and occupation;
- *seniority at retirement*: between 35 and 40 years;
- *kind of worker*: private or public employee;
- *occupation*: white or blue collar;
- *sector*: industry, manufacturing, services, others;
- *geographical area*: North-West, North-East, Centre, South and Islands;
- *institutional framework*: present legislation or alternative rules;
- *transformation coefficient*: values published in the law n. 335/95 or other values (see text and appendix B);
- *pension indexation*: price-linked or full wage-linked;
- *type of interest rate to apply to past years*: historical rates or projected (constant) rate;
- *various macroeconomic parameters*: e.g. type of working population (defined benefit, defined contribution and *pro rata*), last year in which there are new entrants in the working population and last year of the simulation.

C.3 The *computation* module

The module computes:

- *pension*, according to the values attributed to the following (global) variables: kind of worker, institutional framework, transformation coefficients.

Data	Sources
Nominal GDP growth rate (1992-2002)	OECD Statistical Compendium 2003-2. ISTAT (2003), *Annuario Statistico Italiano* (E12.3 – "Conti e aggregati economici delle Amministrazioni Pubbliche")
Inflation: consumer price index (1951-2003)	1951-1996: ISTAT (1997) *Annuario statistico italiano*; 1997-1998: ISTAT (1999 April), *Bollettino mensile di statistica*; other years: www.istat.it
Real wages growth rates (1951-1997)	OECD *Statistical Compendium*, (variable "Wage rate Business Sector" net of ISTAT inflation)
Revaluation coefficients for 2003 pensions: quota A and quota B of DB pensions (1920-2001)	G. Russo (2002), *Il calcolo delle pensioni, manuale operativo*, Il Sole 24 ore (ed.)
Contribution rates INPS-FPLD (1951-2000)	Castellino's computations (1995) on: INPS, (1970) *Raccolta di studi per i settant'anni dell'INPS e i cinquant'anni dell'assicurazione obbligatoria*; INPS (various years), *Notizie statistiche*; 1995-2000: G. Russo (2002), *Il calcolo delle pensioni, manuale operativo,* Il Sole 24 ore (ed.)
Contribution rates CPDEL-INPDAP (1951-2000)	INPDAP (1998), *Studio di fattibilità relativamente allo svolgimento di servizi amministrativi ai fondi pensione complementare*
Cohort-and-gender specific mortality tables	Our computations
Probability to marry again and probability to leave family	INPS (1989), *il modello INPS e le prime proiezioni al 2010*, Previdenza Sociale n.3/89
Average return of public bonds (1960-2000)	Bank of Italy (various years), *Assemblea generale ordinaria dei partecipanti*, appendix; Bank of Italy (2001), *Supplementi al Bollettino Statistico, Mercato Finanziario*, Anno XI, n.3. (1960-1965: medium-term risk-free interest rate returns; 1996-1997: gross average returns on BOT; 1997-2003: BOT rated on MOT)
INPS-FPLD and INPDAP contributors at the 31st of December 2000 by age, seniority and gender	INPS archives; R&P computations on: Ragioneria Generale dello Stato (1997), *Indagine sull'anzianità contributiva dei dipendenti statali.*

Table 11: Data and sources

Below we show pension *formulae* for private employees (FPLD), according to the present legislation.

Defined benefit pension$(a_{95} \geq 18)$:

$$P_{a_{95} \geq 18} = 0.02 \left[(a_{95} - 3) \frac{\sum_{i=0}^{4} w^{*}_{a-i}}{5} + (a - a_{95} + 3) \frac{\sum_{i=0}^{\beta_1(a_{95},a)-1} w^{**}_{a-i}}{\beta_1(a_{95}, a)} \right] \tag{12}$$

Pro rata and defined contribution pension$(a_{95} < 18)$:

$$P_{a_{95} < 18} = 0.02 \left[\beta_2(a_{95}) \frac{\sum_{i=0}^{\beta_3(a_{95},a)-1} w^{**}_{a-i}}{\beta_3(a_{95}, a)} + \left[a_{95} - \beta_2(a_{95})\right] \beta_4(a_{95}) \frac{\sum_{i=0}^{4} w^{*}_{a-i}}{5} \right] +$$

$$+ \left[c_a + \sum_{i=1-\beta_4(a_{95})+a_{95}\beta_4(a_{95})}^{a-1} c_i \prod_{j=i}^{a-1} (1 + \overline{g}_j) \right] \delta_{e+a} \tag{13}$$

where:

$$\beta_1(a_{95}, a) = \min \left\{ 10; Int[6,5 + \frac{2}{3}(a - a_{95})] \right\}$$

$$\beta_2(a_{95}) = \begin{cases} 0 \Leftrightarrow a_{95} \leq 0 \\ a_{95} \Leftrightarrow 0 < a_{95} < 3 \\ 3 \Leftrightarrow a_{95} \geq 3 \end{cases}$$

$$\beta_3(a_{95}, a) = \min\{a; 8 + a - a_{95}\}$$

$$\beta_4(a_{95}) = \begin{cases} 0 \Leftrightarrow a_{95} \leq 0 \\ 1 \Leftrightarrow a_{95} > 0 \end{cases}$$

- *Social Security Wealth*, *Net Present Value Ratio*, *Accrual*, *Tax/subsidy rate* which are described in the text.
- *internal rate of return,* the solution for x of the following equation:

$$\sum_{i=1}^{a} c^{*}_i (1 + x)^{1-i} = P(e + a) \frac{1}{\delta^{co}_{e+a,s}(x)} (1 + x)^{-a} \tag{14}$$

- *replacement rate*, the ratio between the firstly paid pension and the average wage of the last five years of work:

$$RR_a = \frac{P(e+a)}{\frac{\sum_{i=0}^{4} w_{a-i}}{5}} \tag{15}$$

C.4 The *aggregation* module

The module inherits the main assumptions of the others, i.e. the continuity of the working career and the macroeconomic scenario. It follows a "cell-based" approach: individuals are categorized according to their socioeconomic characteristics and according to their age and seniority. Each category represents a cell of a matrix; agents within each cell are treated as identical.

The main data sources of the aggregation module are the following populations:

1. contributors to the funds (FPLD and INPDAP) at the 31st of December 2000, by gender, age and seniority;[29]

2. new contributors to the funds from 1st of January 2001 onward, by gender and age.

The first population is exogenous, while the second is given by the difference between the working population - which diminishes over time due to deaths and to retirement exits – and a "target" population. The resulting total population is thus characterized by an exogenously fixed size and by an endogenous distribution by age, which becomes stationery after a very long simulation period (i.e. when the original working population is extinguished). The distribution by age of the population n. 2 mimic that of the population n. 1 at the 31st of December 2000, and is assumed to be stable along the simulation period.

The aggregation module first selects the correct population (1, 2 or both depending on the reform to simulate), then the workers involved in the reform and finally those who reach the minimum requirements. Retirement is assumed to occur as soon as minimum requirements are met. The impact on pension expenditure of the simulated reform is obtained by an output matrix, showing by row the annual flow of the effect on pension expenditure by year of retirement (workers retired in a given year will continue to generate flows of pensions as far as there is someone in the group still alive), and by column the effect on pension expenditure by year of simulation, generated by workers retired both in that year and in the previous ones.

[29]The population of contributors to FPLD and INPDAP differ in their composition by gender and by seniority. In FPLD there are more males (61 per cent) than females, while in INPDAP the composition by gender is more balanced. Around 53 per cent of both of the populations is subject to *pro rata* rules. In the public sector there are more workers subject to the DB scheme (35 per cent with respect to 21 percent of the private sector). The characteristics of the resulting flow of retirees - given the assumptions of continuous careers and retirement at the minimum requirements - closely reflect those of the original populations of contributors.

D The estimation of wage profiles

Figure 9 shows the average wage by age and by cohort for males. The vertical distance between the curves - which represents a first approximation of the cohort-and-time effect - is relevant and confirms the importance of disentangling cohort, age and time effects in the estimation of wage profiles. A similar figure can be drawn for females.

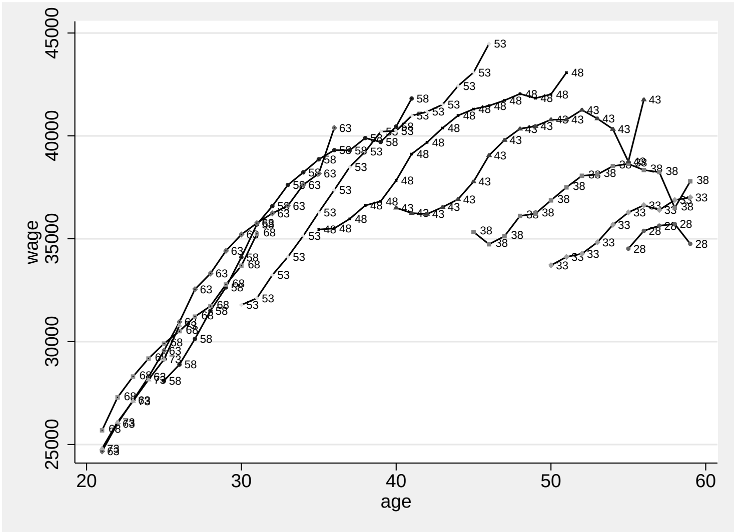

Figure 9: Average wage by age and cohort: males

Notes: annual wages in thousands lira prices 1995. Observations are grouped in 5-years-by-5-years cohort clusters (1926-1931,...,1971-1976). Each cluster is labeled as belonging to the year in the middle of the interval (1928,1933,...,1973).
Source: our computations.

D.1 Model specification

We consider the following specification:

$$\log y_{it} = \alpha + \sum_{j=1}^{j\max} \beta_j s_j(age_{it}) + \sum_{\tau=1986}^{1997} \gamma_\tau TD_{\tau it} + \\ + \sum_{j=1}^{3} \delta_j^1 \sec_{jit} + \sum_{j=1}^{3} \delta_j^2 area_{jit} + v_i + \varepsilon_{it} \quad (16)$$

where t defines the years (from 1985 to 1997), i defines the individuals, $\log y_{it}$ is the logarithm of wage, $s_j(age_{it})$ is the linear spline function of age (with j max knots to be determined by the data), TD_{jit} are time dummies, equal to 1 if $j = t$ and 0 otherwise, sec_{jit} are dummies for sector, equal to 1 if $sec = j$ and 0 otherwise, $area_{jit}$ are dummies for regions, equal to 1 if $area = j$ and 0 otherwise. v_i is an individual effect and ε_{it} is a idiosyncratic error, $i.i.d$ with $E(\varepsilon_{it}) = 0$ and $\sigma^2 = \sigma_\varepsilon^2$.

The individual effect v_i is furtherly specified to model cohort effects. The standard procedure to this aim is to include dummies for year of birth. A peculiar problem with this approach, however, is that it requires additional assumptions on the way the wage profile is vertically shifted, due to the cohort effect, in order to make out-of-sample predictions. The long run horizon of our analysis requires to predict wages for cohorts of workers born after 1976, the last one in the sample. We therefore adopt an alternative modeling procedure. Following Heckman and Robb (1985), we assume that cohort effects can be captured by the macroeconomic conditions when the individuals enter into the labor market, summarized by productivity growth and approximated by GDP per capita.

The inclusion of cohort effects in the model excludes the application of a fixed effects approach, since these are time-invariant covariates. The alternative random effects approach would require the assumption of orthogonality between all the explanatory variables and the individual unobserved component. To relax this assumption, at least as far as area and sector are concerned, we adopt a Mundlak approach. Accordingly to Mundlak (1978), a set of time-variant variables X_{it} are estimated in a random effects approach allowing for $E(v_i|X_{it}) \neq 0$, if their time-average is included in the model.[30]

The individual effect is thus modeled as:

$$v_i = \sum_{j=1}^{3} \delta_j^3 \overline{\text{sec}}_{ji} + \sum_{j=1}^{3} \delta_j^4 \overline{area}_{ji} + \sum_{j=0}^{n} \mu_j GDPPC(j+1)_i + \eta_i \tag{17}$$

where $\overline{\text{sec}}_{ji} = \dfrac{1}{T_i} \sum_{t=1}^{T_i} \text{sec}_{jit}, \overline{area}_{ji} = \dfrac{1}{T_i} \sum_{t=1}^{T_i} area_{jit}$ (T_i is the number of years for which individual i is observed) and η_i is an individual random effect capturing unobserved heterogeneity.

The last sum in (17) represents Almon's distributed lags (Almon 1989) of GDP per capita (of degree $j+1$, between age 16 and 25). The need for modeling GDP per capita in this way comes from the fact that the age of beginning of the working career is unknown and thus an average value of GDP over time is needed. With respect to other weighting methods, however, Almon's lags are more flexible to better fit the data. The term in equation (17) is constructed in the following way. Considering a polynomial of degree n as $\lambda_i = \mu_0 + i\mu_1 + i^2\mu_2 + ... + i^n\mu_n$ and plugging it into $\sum_{j=1}^{10} \lambda_j GDPPC_{26-j}$, (i.e. the linear combination of $GDPPC$ when the individual is aged $[16, 25]$), we obtain:

[30] We cannot include time-averages of age splines and time dummies because they would have been estimated only because the INPS panel is unbalanced. In a balanced panel, in fact, age splines and time dummies are perfectly collinear to the year of birth.

$$\mu_0 \sum_{j=1}^{10} GDPPC_{26-j} + \mu_1 \sum_{j=1}^{10} j \cdot GDPPC_{26-j} + ... + \mu_n \sum_{j=1}^{10} j^n \cdot GDPPC_{26-j} \quad (18)$$

Defining $\sum_{j=1}^{10} GDPPC_{26-j}$ as $GDPPC1 \ldots \sum_{j=1}^{10} j^n \cdot GDPPC_{26-j}$ as $GDPPC(n+1)$, we get the term indicated in equation (17), characterized by $n+1$ $GDPPC$ variables and unknown parameters.

A last point concerns time dummies in equation (16). As defined by equations (16) and (17), in fact, the model is not identified because of the perfect multicollinearity between age, time and cohort variables. Following Deaton and Paxson (1994) we therefore restrict the time dummies to add up to zero and to be orthogonal to a linear time trend.[31] Applying these restrictions we obtain:

$$TD'_{\tau t} = TD_{\tau t} - (\tau - 1)TD_{2t} + (\tau - 2)TD_{1t} \quad \tau = 3, ..., T \quad (19)$$

Equation (19), together with (16) and (17), defines the model.

D.2 Sample selection and specification tests

In the estimation we exploit the administrative archive "Estratti Conto INPS". The panel covers the period 1985-1997 and represents 1:365 of the total population of Italian private sector employees. We run separate regressions, according to gender and occupation (blue and white collars). We consider a sub-sample of full-time employees between age 21 and 59. The upper age bound is chosen in order to get partly rid of a strong selection process occurring in the labor market. Typically, in fact, individuals with lower wages tend to stay longer in the labor market (we observe a decline in the average wage by age above a given threshold). We assume that this selection occurs mainly starting from age 60, i.e. the age at which males were entitled for the old-age pension in the sampled period.

We perform a series of tests in order to specify both the number of knots of the spline for age and the degree of the Almon's polynomial. In addition, we test Deaton-Paxson restrictions on the time-dummies, the existence of residual (and of spurious) cohort effects not captured by our model, and the significance of the Mondlak's additional variables. We show the results of these tests at the bottom of tables 12 and 13, for males and females respectively.

The specification for age is determined by the data, progressively reducing the number of knots of the spline and comparing the resulting formulation with a model including a full set of age dummies. The χ^2-tests do not reject the

[31] Of course there are other ways to deal with the perfect multicollinearity between these variables. Our preference for applying Deaton-Paxson restrictions has two explanations. First, we better capture cohort effects, especially for females. We test it performing F-tests on the joint significance of the $GDPPC$ variables (not shown). Second, we do not need additional assumptions on the time effects when predicting wages for future years.

formulation at 10 equally-spaced knots. Concerning cohort effects, we choose an Almon's polynomial of second degree as a result of a testing-down procedure: we progressively reduce the degree of the polynomial until the χ^2-test rejects the hypothesis of equality with respect to a model with one degree less.[32]

To test the Deaton-Paxson restrictions, we perform an F-test on the restriction $TD_2 = -2TD_3 - 3TD_4 - ... - (T-1)TD_T$. We find that, for three of the four groups (female both blue and white collar and male white collar), it is not rejected by the data. Therefore, we adopt this specification to model time effects.

Residual cohort effects are tested adding to the model 46 year-of-birth dummies and looking at their joint significance.[33] The Wald-test shows different results accordingly to the gender and to the occupation. The year-of-birth dummies turn out to be insignificant for female white collar and only slightly significant for female blue collar meaning that, for this gender, the *GDPPC* variables completely capture the differences between cohorts. For males (both blue and white collar) the model instead captures cohort effects only partly. However, a meaningful result we obtain for males (as well as for females) is that the estimates for the age spline are robust with respect to the way cohort effects are modeled (*GDPPC* variables or cohort dummies). Therefore, we rely on this specification for both genders. Furthermore, we investigate if the *GDPPC* variables suffer from the "spurious regression problem", i.e. if they capture other cohort-specific unobserved characteristics which are positively correlated with GDP. To test this, we add to the model the year-of-birth-squared variable and we look at its significance. The variable is generally insignificant and therefore we conclude that our specification does not suffer from the spurious regression problem.

Finally, the t-test on the time-averages of regional and sector dummies - which provides qualitatively the same result of the Hausman's test - confirms that simple random effects estimates would be inconsistent.

D.3 Results

Estimates are presented in table 12 and 13, for males and females respectively. Profiles by age and cohort effects are shown in figure 3 and briefly commented in section 3.

[32] For female blue collar we choose a polynomial of fourth degree. Given that the testing-down procedure does not stop at 5 percent significance level, we need to fix a much higher significance level for this group.

[33] The sample includes 50 cohorts. We must subtract 4 (arbitrary chosen) dummies to avoid perfect multicollinearity: one between the cohort dummies, one to avoid multicollinearity between year of birth, year and age, and two because of $GDPPC1$ and $GDPPC2$.

variable		**blue collar**		**white collar**	
dep: log-wage		Coeff.	t	Coeff.	t
Spline for age:					
	24	0.039	24.5	0.049	14.18
	28	0.026	28.11	0.045	35.28
	32	0.021	24.83	0.043	45.35
	36	0.016	17.43	0.033	37.21
	40	0.016	17.06	0.027	29.79
	44	0.014	15.39	0.023	24.53
	48	0.013	13.65	0.023	23.17
	52	0.010	9.39	0.015	12.87
	56	0.005	4.14	0.018	11.66
	> 56	0.006	2.68	0.023	7.75
Cohort effect:					
	GDPPC1	7.783	6.1	13.806	6.11
	GDPPC2	-1.169	-4.57	-2.222	-4.89
Sector:					
	industry	0.069	6.92	0.042	4.11
	construction	0.049	4.36	0.050	3.44
	services	0.026	2.64	0.061	6.09
Area:					
	North-East	-0.030	-3.43	-0.010	-1.16
	Center	-0.044	-5.15	0.009	1.14
	South and Isl.	-0.035	-4.5	-0.034	-4.09
Obs.			103981		37510
Groups			15567		4784
Log L.			11679.9		17134.1
σ_u			0.267		0.306
σ_e			0.179		0.121
Specification tests:*					
LR test against a model with a full set of age dummies					
$\chi^2_{(28)}$		32.5	(0.255)	17.63	(0.935)
LR test against a model with					
a further reduction in the degree of Almon's polynomial					
$\chi^2_{(1)}$		20.9	(0)	23.91	(0)
Wald test for Deaton-Paxson restrictions					
$\chi^2_{(1)}$		26.16	(0)	4.98	(0.026)
Wald test for cohort residual effects					
$\chi^2_{(46)}$		88.59	(0)	134.45	(0)
t-test for year-of-birth squared					
t-value		3.68	(0)	1.15	(0.125)
Wald test for time-averages of sector and area					
$\chi^2_{(6)}$		456.05	(0)	225.19	(0)

Table 12: Random effects estimates of log-wage: males

* p-values are presented in parentheses.
Notes: log-wage is in thousands lira, prices 1995; constant, time dummies and time-averages of sector and area are omitted. Reference individual is: other sectors, North-West.
Source: our computations.

variable		**blue collar**		**white collar**	
dep: log-wage		Coeff.	t	Coeff.	t
Spline for age:					
	24	0.023	8.93	0.042	15.67
	28	0.006	3.47	0.022	14.91
	32	0.003	1.5	0.014	9.76
	36	0.007	3.62	0.014	9.42
	40	0.012	5.93	0.017	10.61
	44	0.006	3.25	0.019	10.44
	48	0.006	2.81	0.011	4.83
	52	0.004	1.79	0.022	7.89
	56	-0.001	-0.16	0.019	4.34
	> 56	-0.007	-1.07	0.016	1.8
Cohort effect:					
	GDPPC1	44.036	1.8	8.453	2.55
	GDPPC2	-33.512	-1.57	-1.331	-2.03
	GDPPC3	6.877	1.41	-	-
	GDPPC4	-4.128	-1.28	-	-
Sector:					
	industry	0.145	10.58	0.093	7.09
	construction	0.137	3.41	0.092	4.07
	services	0.049	4.28	0.058	4.62
Area:					
	North-East	0.033	1.58	0.032	1.6
	Center	-0.025	-1.11	0.029	1.58
	South and Isl.	0.018	0.73	-0.045	-1.63
Obs.			35232		28666
Groups			6443		4400
Log L.			-3532.4		2719.5
σ_u			0.371		0.358
σ_e			0.210		0.174
Specification tests:*					
LR test against a model with a full set of age dummies					
$\chi^2_{(28)}$		31.37	(0.301)	15	(0.979)
LR test against a model with a further reduction in the degree of Almon's polynomial					
$\chi^2_{(1)}$		1.63	(0.201)	4.1	(0.043)
Wald test for Deaton-Paxson restrictions					
$\chi^2_{(1)}$		1.9	(0.1678)	0.01	(0.926)
Wald test for cohort residual effects					
$\chi^2_{(46)}$		**60.93	(0.046)	41.54	(0.620)
t-test for year-of-birth squared					
t-value		0.99	(0.161)	-0.38	(0.648)
Wald test for time-averages of sector and area					
$\chi^2_{(6)}$		46.72	(0)	46.72	(0)

Table 13: Random effects estimates of log-wage: females

* p-values are presented in parentheses; ** 44 dof for this group.
Notes: log-wage is in thousands lira, prices 1995; constant, time dummies and time-averages of sector and area are omitted. Reference individual is: other sectors, North-West.
Source: our computations.

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