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LIFETIME INCOME AND OLD AGE MORTALITY RISK IN ITALY OVER TWO DECADES

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Abstract

BACKGROUND

European studies highlight a widening of relative inequalities in general mortality by socioeconomic status from the 1970s to the 1990s. Few studies are available for Southern European countries; they show that these countries represent an exception to these trends. Available evidence on this for Italy is for a specific city only and nationwide evidence does not exist yet.

OBJECTIVE

This paper examines the association between lifetime income and old age mortality risk, referred to as the income–mortality gradient, in Italy during the 1980s and 1990s. We investigate the shape of the gradient. Most importantly, we analyze the evolution of the gradient between these two decades.

METHODS

We use data drawn from an administrative pension archive held by the main Italian social security institution and proxy individual lifetime income with pension income. We use non-standard Cox proportional hazard models, where the positions and number of the knots in the spline function for income are determined by the data.

RESULTS

The shape of the income–mortality gradient shows two discontinuities for males and one for females; these kink points are situated almost at the same percentiles of the income distribution during the 1980s and the 1990s. The estimated associations are negative and stronger at higher income levels. The income–mortality gradient widens over time for males and remains unchanged for females. Accounting for regional effects explains most of the widening in the gradient over time for males.

CONCLUSIONS

Our findings show for both males and females that mortality risk decreases with income. Once controlled for regional differences, and in contrast with the trends observed in many other European countries, the relative difference in mortality risk between high and low-income individuals is rather stable over time in Italy.

1 - Introduction

Since the seminal works of Antonovsky (1967) and Kitagawa and Hauser (1973), many studies have quantified the differences in mortality risk across socioeconomic groups in various countries.[1] Individuals' socioeconomic status (SES) is commonly defined using income, wealth, or educational attainment (Galobardes et al., 2006). A significant negative correlation between SES and mortality is nearly always found; for a comparison across European countries, see, e.g., Mackenbach et al. (1997, 2008) and Huisman et al. (2004); for recent single-country studies see, e.g., Hupfeld (2011) and Gaudecker and Scholz (2007) for Germany), Leombruni et al. (2010) for Italy, Kalwij et al. (2013) for the Netherlands, Cristia (2009) for the US and Attanasio and Emmerson (2003) for the UK.

There is an ethical and ideological consensus—often echoed by the agendas of supranational institutions and official government statements (CDC, 2012; CSDH, 2008; COM, 2009; COM, 2007)—that differentials in health and life expectancy by social position must be tackled. Besides stressing that converging life expectancies across social categories can contribute to the general social and economic development of societies and to improving international rankings of life expectancy (Wilmoth and Dennis, 2007), the consensus also emphasizes the need for improving information systems for monitoring inequalities and evaluating policies and interventions (Kunst et al., 2004).

European studies highlight a widening of relative inequalities in mortality by SES from the 1970s to the first half of the 1990s and attribute this to changes over time in the social distribution of cardiovascular disease mortality, which in most countries declined relatively more in the upper socioeconomic groups, and to health-related behaviors such as smoking (Mackenbach et al., 2003). Findings for Southern European countries (Dalmau-Bueno et al., 2010; Borrell et al., 2008), as well as for Italy specifically (Mackenbach et al., 2003), represent an exception to these trends. The change in the SES-mortality gradient over time has been smaller and the change in cardiovascular disease mortality is more homogenous across socioeconomic groups in Southern than in other European countries. The comparative research by Mackenbach et al. (2003)

[1] Grossman (Grossman, 1972) developed a theoretical economic model that explains health differences across socioeconomic groups.

cited above considers two broad socioeconomic groups: high versus low educational level, and manual versus non-manual occupations. Italy contributed to this research, as well as to other international studies such as Mackenbach (2008), with the Turin Longitudinal Study (TLS), a system aimed at monitoring mortality and morbidity variations since 1971, by combining census data with health information databases (Costa et al., 1988). Empirical evidence on a changing income-mortality gradient over time on a national scale is however still lacking for Italy as the TLS covers only residents of the city of Turin, who may experience relatively high health inequality[2], and has no individual income information.

Although most data published on inequality refer to people in the working-age group, social inequalities in mortality are also found for European elderly males. In comparison with the younger group, elderly males typically show lower (higher) relative (absolute) differences in mortality rates by SES (Huisman et al., 2004). Results for females are more heterogeneous; in specific countries, including Italy (Turin), a slighter decrease of relative inequality with increasing age is found (see Huisman et al., 2004). Moreover, the age pattern of inequality is shown to differ with regard to different indicators of social position, such as home ownership, which can be considered a marker of material circumstances (Galobardes et al., 2006), and income. Pension benefits are often closely related to lifetime earnings and have therefore previously been used to measure elderly's SES (see, e.g., Gaudecker and Sholz, 2007).

Finally, it has been observed that there is a concave relationship between income and mortality risk that may result from a diminishing return of health with increasing income (see, e.g., Mackenbach et al., 2005; Dowd et al. 2011; Hupfeld, 2011). However, the diminishing return hypothesis has been fertile ground for contrasting empirical evidence on mortality, as well as for various health measures (Rahkonen et al., 2000; Der et al., 1999; Marinacci et al., 2009). Leombruni et al. (2010) find that in Italy only the richest individuals benefit from significantly lower mortality risk.

The contribution of this paper is twofold. First, we investigate the shape of the association between income and old age mortality risk in Italy, and re-evaluate the diminishing return hypothesis for this country.

[2] Mackenbach et al. (2008) underline that urban areas – such as those included in their study (Turin, Barcelona, Madrid) - may be characterized by higher inequality in health than rural areas (see also Bos, Kunst and Mackenbach, 2002 and Hayward, Pienta and McLaughlin, 1997).

We extend Leombruni et al. (2010), using the method proposed by Dowd et al. (2011) and Molinari et al. (2001) of a non-standard free knots spline specification for the income–mortality gradient. Second, we extend Mackenbach et al. (2003) and provide empirical evidence for Italy on the evolution of the income-mortality gradient between the 1980s and the 1990s.

This paper is organised as follows. Section 2 presents the data and the statistical model used for the analysis. Section 3 reports our empirical findings and discusses the analysis and findings. Finally, section 4 concludes.

2 - Data and methods

2.1 - Dataset and income measure

We exploit a pension database drawn from an administrative archive held by the main Italian social security institution, Istituto Nazionale Previdenza Sociale (INPS). Our database reports pensions paid by INPS since its establishment in 1933 up to and including 2001. It covers approximately 1/90 of the ex-private sector workforce plus social assistance beneficiaries (in total approximately 289,000 individuals). Civil servants are therefore not included. While it is less evident that there are mortality differences between private and public sector workers,[3] for females the condition on having had employment may yield a non-random sample of females as their labor force participation is related to SES (Bratti, 2003). The data include all pension schemes managed by INPS. Major schemes cover private sector employees (Fondo Pensioni Lavoratori Dipendenti, FPLD fund) and the self-employed (artisans, traders, and farmers). Special schemes include, among others, miners, pilots, sailors, and clerical personnel. The following variables are available: month and year in which the pension was first paid to the individual, month and year in which the pension flow ended (if ended), pre-tax monthly pension amount, pension scheme, and benefit type (e.g., old age pension, early retirement, disability insurance and survivors benefits). In addition, there is data on individual

[3] The health status of civil servants may (indirectly) benefit of higher job security (Ferrie et al., 1998) and lower unemployment probabilities (for Italy, see Costa, Spadea and Cardano, 2004).

date and region of birth and gender. When an individual dies, INPS records the end of all pension payments the person had been receiving. We assume that the individual dies in the month of the last pension payment.

A first investigation showed that the quality of the variable *date in which the pension flow ended* is rather poor before January 1979. For this reason, we follow individuals aged 65 or over from January 1979 onwards.[4] The selection of 65+ is chosen since at age 65 most individuals are retired in the years covered by our data (Belloni and Alessie, 2009). Until 1994, males (females) could claim an old age pension at age 60 (55). After a period characterized by gradual increments, the minimum age for the old age pension was set at 65 (60) for males (females) in 2001. To facilitate comparisons, we apply the same age selection to both genders. Finally, we exclude individuals born before 1901 because coverage by the pension system for private sector employees from these cohorts was partial and participation was voluntary. Therefore, our selected data cover the cohorts born between 1901 and 1936.

An individual's lifetime income is considered a good measure for his or her SES (Sullivan von Wachter, 2009; Gaudecker and Sholz, 2007).[5] We proxy individual lifetime income by the amount of pension benefit received. This is a good proxy variable if we restrict our analysis to ex-private sector employees. In their case, the pension formula summarizes the salient characteristics of the working career: (last) average wages and seniority (years of contribution to the scheme). We exclude the self-employed since the benefits they receive are, given the pension rules, a bad proxy for their lifetime income.[6]

There is a minimum, but no maximum, pension benefit in Italy. If the accrued benefit is below the minimum pension and an earnings test is passed, the individual receives a social assistance benefit to make up for this difference. To reduce measurement error, we exclude individuals whose total pension income is

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[4] Individuals retired before 1979 *are* included in the sample if they are alive in January 1979. Therefore, also in the first years of the sample we have individuals of all ages.

[5] It is noteworthy that studies such as Martikainen (2009) and Kalwij (2013) have shown that —also for females— mortality is more strongly associated with individual income than with household income.

[6] However, we include ex-employees who also receive self-employment pensions if the latter relates to only a minor portion of their total pension income.

below the threshold and exclude possible outliers with very high pension income by trimming the income distribution at the top per mille.

2.2 - Statistical analysis

The analysis is performed on monthly data, from January 1979 to December 2001. We split the whole period into two sub-periods of similar length—January 1979–December 1990 (the “1980s,” follow-up 144 months) and January 1991–December 2001 (the “1990s,” follow-up 132 months). We analyze males and females separately and thus end up with four distinct estimation samples.

A preliminary analysis of the association between lifetime income and survival was obtained from Kaplan-Meier survival estimates by gender and income quintiles for the two analyzed decades.

Further, as recently proposed by Dowd et al. (2011), we estimate a Cox model (Cox, 1972) by decade (and gender). Individuals that survived up to December 1990 are treated as right censored observations in the first period and as left truncated observations in the second period. Mortality risk is assumed to depend on survival time, (age) and a set of time-invariant individual characteristics, including lifetime (pension) income. We mainly consider models with income as the only explanatory variable since we are especially interested in the unconditional association between lifetime income and mortality risk.

Following Dowd et al. (2011) and Molinari et al. (2001), we implemented non-standard free knots spline specifications to model the income–mortality gradient, i.e. a Cox model where the position of the knot(s) in the spline function is determined by the data. Such a model is particularly suitable for analyzing a change in the shape of the income–mortality gradient over time. The change in an estimated income slope coefficient indicates either a deeper or a weaker association between lifetime income and risk of death, while a change in the position of a knot suggests that such association applies to either a wider or a narrower part of the population. As in the standard Cox model, parameter estimates—including the knot(s) parameter(s)—are obtained by maximum likelihood.

As it turned out, free knots models with only a few knots already outperform standard Cox model such as (log-) linear, quintile, (log-) linear spline with 5 and 10 knots located at corresponding quintiles of the (gender- and period-specific) income distribution. Log-transformed income variables turned out to better fit the data. Results of hypotheses tests were rather insensitive with respect to criteria (LR-test, AIC, BIC) and periods. When different criteria led to different results, to ease interpretation, we opted for models with a lower number of knots, and the same number of knots for the two periods. Based on these statistical tests we selected a free knot spline Cox model with two knots in the case of males and with one knot in the case of females. We refer to the appendix for further details.

To test the joint hypotheses of changes in the slope parameters over time, we implement a Hausman test. Hereby we take into account dependence across samples as right-censored individuals in the 1980s sample also belong to the 1990s sample. The variance of the difference between the two parameters vectors is obtained by paired bootstrapping (Cameron and Trivedi, 2005).

3 - Results

3.1 - Empirical findings

Table 1 reports descriptive statistics. In total, there are 33,542 failures (deaths). Male median survival time increased from 178 months (14.8 years) in the 1980s to 203 months (16.9 years) in the 1990s. Corresponding values for females are 246 and 263 months, respectively.[7] Average monthly pension income is higher for males than for females and higher in the 1990s than in the 1980s; this mainly reflects economic growth. In addition, there was a moderate reduction in pension income inequality between the 1980s and the 1990s.

Figure 1 reports Kaplan-Meier survival estimates by decade, gender and income quintile (first *versus* fifth). In the 1980s (left graph), male median survival is equal to 172 months for the poorest group, and 197 months for the richest group, revealing an absolute difference in life expectancy of 25 months and a relative difference of 15%. Corresponding values for females are 244 and 258 months, showing a smaller gradient

[7] Mean survival times (not reported) are somewhat higher than life expectancies at age 65 reported in HMD (2012): +0.2 years for males in the 1980s and +1 year for males in the 1990s; corresponding values for females are +0.5 and +1.6 years. It should however be realized that the two populations are different, since we consider ex-workers whereas HMD includes the whole population.

(absolute difference 14 months, relative difference 6%). In the 1990s (right graph), male median survival is found to equal 185 months for the first quintile and 228 months for the fifth quintile (absolute difference 43 months, relative difference 23%). Corresponding values for females are 252 and 270 months (absolute difference 18 months, relative difference 7%). From these figures, it turns out that differential mortality is sizable for males and further increases over the analyzed period. Females experience a smaller gradient, which only slightly increases from the 1980s to the 1990s. By comparing median survival time by income quintile with the unconditional values reported in Table 1, we may infer that differential mortality is concentrated in the upper part of the income distribution.

Table 2 reports the parameter estimates for the models using a free knots spline for income, and Figures 2 and 3 present predictions based on these estimates. For males—1980s, a first knot (*knot 1*) is estimated at €805 monthly pension income (exp(6.691)); this corresponds to the 66th income percentile. In this period, for males with income below this threshold, the association between income and mortality risk is positive, although weak (*income 1*): a 1% increase in lifetime income increases the risk of death by 0.07%. Figure 2 shows this perhaps surprising result as the upward slope up to p66; we further investigate this issue below. For males with income above this level, we find a negative income–mortality association. In particular, for incomes between the 66th and 98th percentiles (*knot 2* is at approximately €7000 monthly pension income) this association is moderate: if income increases by 1%, the hazard of death is reduced by 0.25% (*income 2*). For top incomes (i.e., greater than the 98th percentile), the association becomes much steeper: a 1% increase in lifetime income reduces mortality risk by 3.6% (*income 3*). Figure 2 shows this result by the very steep downward slope after p98. It must be recognized that standard errors of the *income 3* slopes are large (especially in the 1980s) and, although we carefully checked influential observations (see section 3.2), results for top incomes should be considered with caution.

Results for males in the 1990s show that the positions of the two knots in the income–mortality gradient are by and large the same as for the previous decade. However, the *income 1* slope parameter has changed considerably over time: for males with income below the 62nd income percentile (p62), the estimated association is negative in the 1990s; a 1% increase in income reduces risk of death by 0.06%. The same unitary change reduces the risk of death of top earners in the 1990s by 1.4% (cf. with 3.6% in the 1980s; this difference is, however, not statistically different from zero). Little changes between the 1980s and the 1990s for males with lifetime income between the 62nd/66th and 98th percentiles (cf. *income 2* parameters). The result of the Hausman test (*p*-value 0.009) provides further evidence in favor of an overall changing income–mortality gradient over time.

For females, we find a (single) knot at the top of the income distribution (97th percentile in the 1980s, and 99th percentile in the 1990s). The sign of the estimated associations between income and mortality is negative for females over the whole income range. The size of such association is much smaller below the estimated knot—i.e., for almost the entire population—than above it. This result is similar to what we found for males (second knot). In the 1980s, the estimated association between income and mortality is not statistically different from zero below the knot. The association deepens below the knot (cf. *income 1* in the 1980s and 1990s) and weakens above it (cf. *income 2* parameters). However, individually, these time differences in the slope parameters are insignificantly different from zero and the Hausman test shows that these changes are jointly not statistically significant (*p*-value 0.67). All this suggests an invariant income–mortality gradient for females over these two decades.

Finally, we return to the perhaps surprising result in Figure 2 of the upward slope up to the 66th percentile. In Table 3 we present results for models that include area of birth as an additional covariate. We show results only for males as for females they do not vary substantially with respect to those presented in Table 2. The results show a negative estimate for the *income 1* coefficient that is not statistically different from zero for the 1980s. We also find a somewhat higher estimate for the *income 1* coefficient for the 1990s. Males living (and working) in the industrialized Northwest of the country have higher mortality, compared to those living in other parts of Italy (see ISTAT, 2012; Caselli et al., 2003). This is confirmed in our data by the area of birth dummies estimates. Meanwhile, males in the Northwest also have higher average income (€1230 versus €950 in the South & Islands). Conditional on living in a particular area, the effect of individual income on mortality is thus strengthened. We speculate that this effect occurs only for poorer males (*income 1*, i.e., up to the 62nd/66th lifetime income percentile) since it especially affects manual workers.

Results set forth in Table 3 also show that once controlled for geographic area, the estimate of the *income 1* parameter for males does not change over time in a significant way (cf. with Table 2). The

Hausman test suggests that—conditioning on living in a given area of the country—the association between income and mortality remained unchanged from the 1980s to the 1990s (*p*-value 0.07). Regional differences in mortality—higher mortality in the Northwest and particularly in the 1980s—and income thus explain a large part of the time evolution of the income–mortality association reported in Table 2.

3.2 - Robustness checks

We run an extensive analysis of the robustness of the results with respect to sample selection and model specification. We were particularly concerned about the sensitivity of the results on the shape of the income–mortality association to the inclusion/exclusion of earners at the two extremes of the income distribution. We thus experimented with including in the sample individuals whose pension income is below the minimum pension and trimming of top incomes to exclude possible outliers with different thresholds. We also compared the estimation results from standard Cox models with income quantiles or spline functions with exogenously determined knots as explanatory variables with the results from our models using the free knots spline Cox specifications. Our results turned out to be robust to all these modifications.

We have gathered reassuring additional evidence that the potential measurement errors in mortality and income are likely to be of little influence on our findings. Recall that individuals' date of death is constructed using information on the date their pension flow ended. When an individual obtains more than one pension during his or her life we examine the most recent ending date. In this way, it is possible to adequately deal with any inaccuracy resulting from the existence of other possible reasons for stopping a specific pension payment, such as conversion of disability into old age pensions or temporary illness. Reconstructed death rates by age turned out to be similar to those reported in the Human Mortality Database (HMD, 2012). The potential issue of underestimation of individual total pension income may arise for private sector employees who contributed to other pension schemes managed by non-INPS institutions during their working career. For example, if an individual partly worked in the private and in the public sector for a sufficiently long period to accrue pension rights in both respective funds, we would underestimate individual's total pension income.[8] For the period 1995–2000, from a different pension file (drawn from the Casellario dei Pensionati archive), we obtain information on non-INPS pensions—i.e., public sector pensions and minor pension funds managed by large firms. We find that only a negligible percentage of individuals (1.7% of males and 0.3% of females) receive pensions from both INPS and non-INPS institutions.

In addition, the main empirical results for both genders are unaffected when controlling for other covariates such as year of birth, retirement age, or whether receiving a disability pension. These additional results are available from the authors upon request.

Finally, in our analysis individuals are at risk of death from age 65 onwards. They may, however, retire earlier but this may be related to health (hence to mortality risk); we therefore did not take the time before age 65 into account. Nevertheless, analyzing individuals at risk of death from the age when they first received a pension onwards (and this could be before age 65) did not affect our main findings in terms of both time evolution and shape.

3.3 – Discussion

The shape of the income–mortality association that we find is uncommon in the literature: whenever a curvilinear relationship is found, it is typically characterized by decreasing returns to income (Martikainen et al., 2009). This evidence, however, does not refer to Southern European countries. We already mentioned that the income–mortality association for these countries is found peculiar in terms of both strength and time changes. Similar to our findings, Leombruni et al. (2010) find that in Italy only retirees in the highest quintile of the income distribution benefit from lower death probability in comparison to the rest of the retired population. A possible explanation of these findings relates to confounding factors. We have already shown that the income–mortality association for poorer males in the 1980s can partly be explained by regional differences. Data on educational attainment and working career characteristics, which we do not have, could certainly yield further explanations. Findings for other countries show that, after adjusting for other SES

[8] To control for such issue, Gaudecker and Sholz (2007) use information on the number of years of pension-relevant insurance periods (and the type of health insurance coverage) and estimate mortality rates for different subsamples. We do not have this information in our data.

factors, the income–mortality association typically becomes less strong but remains characterized by decreasing returns to income (Martikainen et al., 2009).

Results for top earners, whose estimated income–mortality association is particularly steep, should be interpreted with caution. From the statistical point of view, we have already pointed out that estimates of the slope parameters above the highest positioned knot are relatively imprecise. Moreover, for top earners observed pension income can be a poor proxy of lifetime income for the following reasons. First, for these people, pension income could be a poor proxy of total earnings since they may have substantial income from financial assets and housing wealth. Alvaredo et al. (2012) report that in Italy, in 1990, 24% of incomes of the top-1% earners were either capital income or rents, while wages and pensions represented only 37%; corresponding figures for the top 10-to-5% were 7% and 83%, respectively. Top earners are typically executives and highly qualified workers who are often enrolled in special pension schemes. Their actual pension income can be higher than what is observed due to lump-sum compensation or additional annuities (unobserved in this data) that were often granted by employers to get rid of these high-cost workers or as a final premium for their working careers. Finally, the work-package agreement for top earners was extremely generous (often included private health insurance coverage and various in-kind benefits) and, therefore, their SES was higher than implied by their income.

Our results on the time evolution of the income-mortality association are in line with those in Mackenbach et al. (2003) for the city of Turin and with other SES indicators (broad educational and occupational categories). A widening of relative inequalities in general mortality by income during the 1980s and 1990s is a common finding for European countries. Such evidence, however, mostly does not include Southern European countries due to a lack of longitudinal data on mortality. As stressed by Mackenbach et al. (2003, 2008), studying Southern European countries is of major interest, since these countries are perhaps illustrative of the dynamics in the social patterning of behavioral risk factors, like smoking and diet, which perhaps differ from those of Nordic countries. In this study, we cannot analyze behavioral changes, nor do we have causes of death information to explore in depth the causes of any time change in the income–mortality gradient. Nevertheless, in this paper, we complete an important first step and show that for Italy, it is necessary to account for regional differences when examining changes in the income–mortality gradient, and that after doing so, that the income-mortality gradient turns out to be rather stable over time.

4 - Conclusions

This is the first Italian study on the time trend of the association between income and mortality risk. In addition, this study provides further empirical evidence on the shape of the income–mortality association for Italy. The analysis makes use of a sample of ex-private sector workers aged 65 and over from a newly available pension database drawn from an administrative archive held by the main Italian social security institution.

Our empirical results show that the association between lifetime income and old age mortality risk in Italy is negative and stronger for higher income levels (increasing returns to income). Once controlled for regional differences, and in contrast with the trends observed in many other European countries, the income-mortality gradient is rather stable over time for both males and females. This suggests that decreasing trends in cardiovascular disease mortality and in behavioural risk factors like smoking in Italy may be rather evenly distributed among people with different socioeconomic status.

5 – References

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Appendix: Testing

Table 4 shows likelihood-based goodness-of-fit measures: AIC, BIC, and an LR-test against the log-linear model. Notice that (i) free knots models with a different number of knots are not nested; (ii) all estimated models are nested in the log-linear model. An alternative test to AIC/BIC is an "indirect" LR-test (Molinari, 2001) in which each model is compared with the log-linear model.

Table 4 – Model selection

Model:	**d.f.**	**-2x log(Lik)**	**LR-test***	**AIC**	**BIC**
		Males - 1980s			
log-linear	1	180988	-	180990	180998
log-linear spline with 5 knots	5	180876	2,73E-23	180886	180928
log-linear spline with 10 knots	10	180846	3,99E-26	180866	180949
log-linear spline with 1 free knot	3	180852	2,98E-30	180858	180883
log-linear spline with 2 free knots	5	180821	4,67E-35	180831	180873
log-linear spline with 3 free knots	7	180819	7,15E-34	180833	180891
		Males - 1990s			
log-linear	1	229070	-	229072	229081
log-linear spline with 5 knots	5	228974	6,98E-20	228984	229027
log-linear spline with 10 knots	10	228940	1,19E-23	228960	229046
log-linear spline with 1 free knot	3	228935	5,97E-30	228941	228967
log-linear spline with 2 free knots	5	228928	9,54E-30	228938	228981
log-linear spline with 3 free knots	7	228928	3,44E-28	228942	229002
		Females - 1980s			
log-linear	1	62464	-	62466	62474
log-linear spline with 5 knots	5	62452	0,017351	62462	62502
log-linear spline with 10 knots	10	62442	0,008879	62462	62541
log-linear spline with 1 free knot	3	62437	0,000001	62443	62467
log-linear spline with 2 free knots	5	62437	0,000017	62447	62486
		Females - 1990s			
log-linear	1	118416	-	118418	118426
log-linear spline with 5 knots	5	118400	0,003019	118410	118452
log-linear spline with 10 knots	10	118396	0,017912	118416	118499
log-linear spline with 1 free knot	3	118405	0,003218	118411	118435
log-linear spline with 2 free knots	5	118398	0,001135	118408	118449

* LR-test: each model is compared with the log-linear model, *p*-value.

Tables:

Table 1 – Descriptive statistics

	Males		Females	
	1980s	***1990s***	***1980s***	***1990s***
Subjects	30,830	39,054	20,416	30,470
Failures (% of subjects)	9,991	12,916	3,653	6,982
	(0.32)	(0.33)	(0.18)	(0.23)
Survival time, median*	178	203	246	263
Monthly pension income (€, 2009 prices):				
Mean	1029	1198	541	604
Standard deviation	2173	1952	1073	913
Percentiles: 1st	207	227	207	218
10th	276	355	272	290
50th	533	828	360	452
90th	1821	2086	757	939
99th	9116	7386	3910	3185
*Gini's index of inequality***	0.53	0.45	0.39	0.35
Area of birth (%):				
Northwest	0.24	0.23	0.32	0.30
Northeast	0.23	0.22	0.26	0.26
Center	0.18	0.18	0.16	0.16
South & Islands	0.35	0.37	0.26	0.28

Note: * Kaplan-Meier survival estimates; ** ranges between 1 (= max inequality) and 0 (= no inequality).

Table 2 – Free knots spline Cox models: parameter estimates

Period/Parameter	income 1	knot 1	income 2	knot 2	income 3
			Males		
1980s	0.068	6.691	-0.248	8.857	-3.568
	(0.029)	(0.092)	(0.037)	(0.135)	(1.130)
		p66		*p98*	
1990s	-0.056	6.977	-0.201	8.689	-1.444
	(0.022)	(0.055)	(0.037)	(0.045)	(0.178)
		p62		*p98*	
Hausman test (*p*-value)*	0.009				
			Females		
1980s	-0.063	7.757	-1.867	-	-
	(0.042)	(0.220)	(0.754)		
		P97			
1990s	-0.102	8.238	-0.738	-	-
	(0.027)	(0.063)	(0.202)		
		p99			
Hausman test (*p*-value)*	0.673				

Notes: models by gender and decade; spline function with endogenously determined knots; *income x* is the percentage change in the hazard of death associated with a 1% increase in the lifetime (pension) income for levels of log-income in the interval [*knot x-1*, *knot x*]; *knot x* is the estimated knot of the log-income spline function; in italics we report the position of the estimated knot in the sample-specific log-income distribution; standard error in parenthesis; * χ^2-test on the equality of the joint income parameters in the 1980s versus the 1990s samples; the variance of the difference between the 1980s and the 1990s income parameters vectors is obtained by paired bootstrapping (100 samples); see Cameron and Trivedi (2005) equation (11.24) for details.

Table 3 – Free knots spline Cox models including area of birth as a covariate—males: parameter estimates

Period/Parameter	income 1	knot 1	income 2	knot 2	income 3	NE	C	SI
1980s	-0.012	6.714	-0.226	8.835	-3.449	-0.079	-0.129	-0.256
	(0.026)	(0.019)	(0.034)	(0.009)	(0.755)	(0.028)	(0.031)	(0.027)
		P67		*P98*				
1990s	-0.093	6.978	-0.188	8.688	-1.430	-0.054	-0.142	-0.135
	(0.022)	(0.038)	(0.036)	(0.021)	(0.177)	(0.026)	(0.028)	(0.024)
		P62		*P98*				
Hausman test (*p*-value)	0.072							

Notes: area of birth dummies: reference category is Northwest, NE=Northeast, C=Center, SI=South & Islands; see Table 2 for additional explanations.

Figures:

Figure 1 – Kaplan-Meier survival estimates by gender (F for females and M for males) and income quintile (q; 1 corresponds to the 1st quintile, 5 to the 5th quintile): 1980s and 1990s

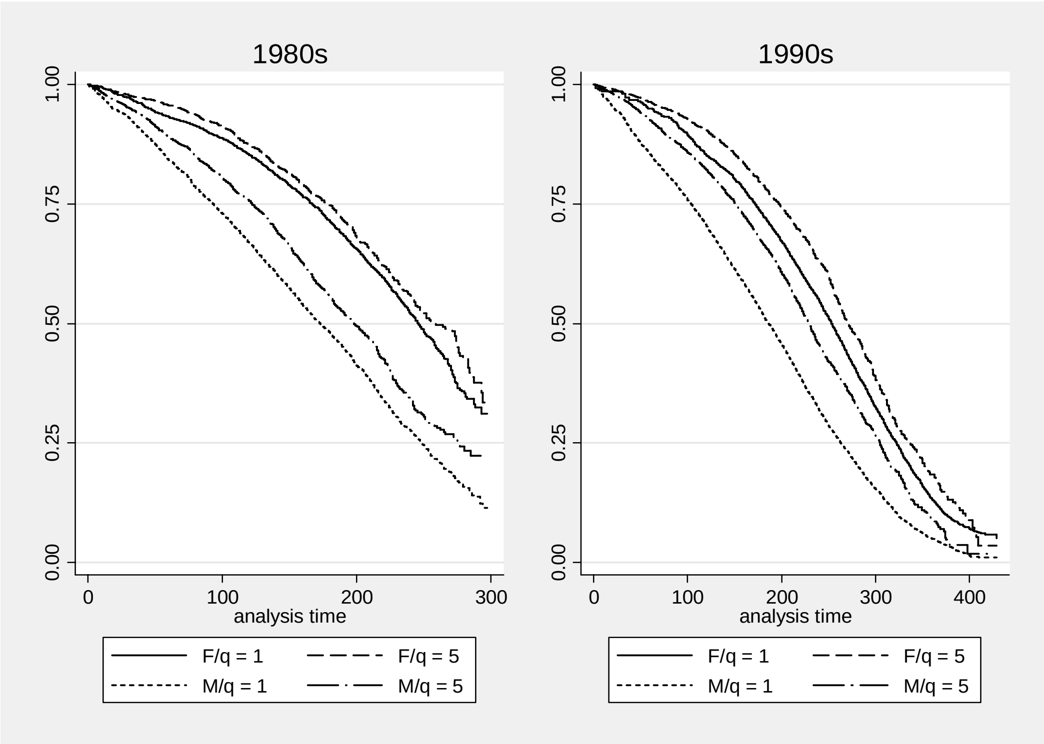
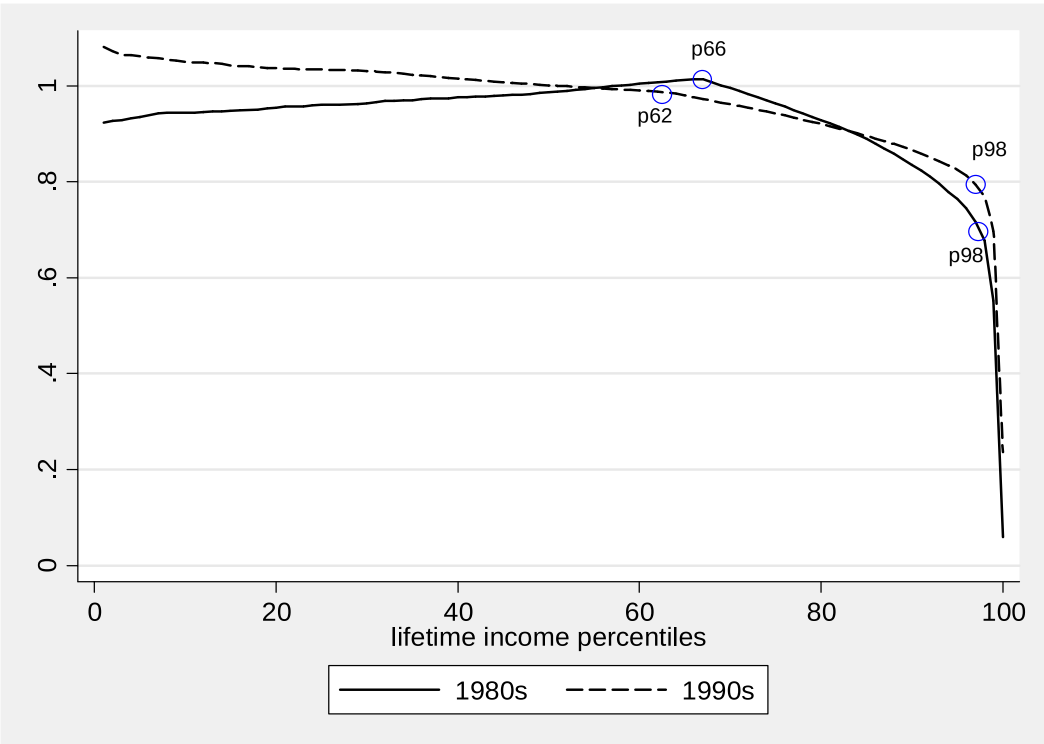
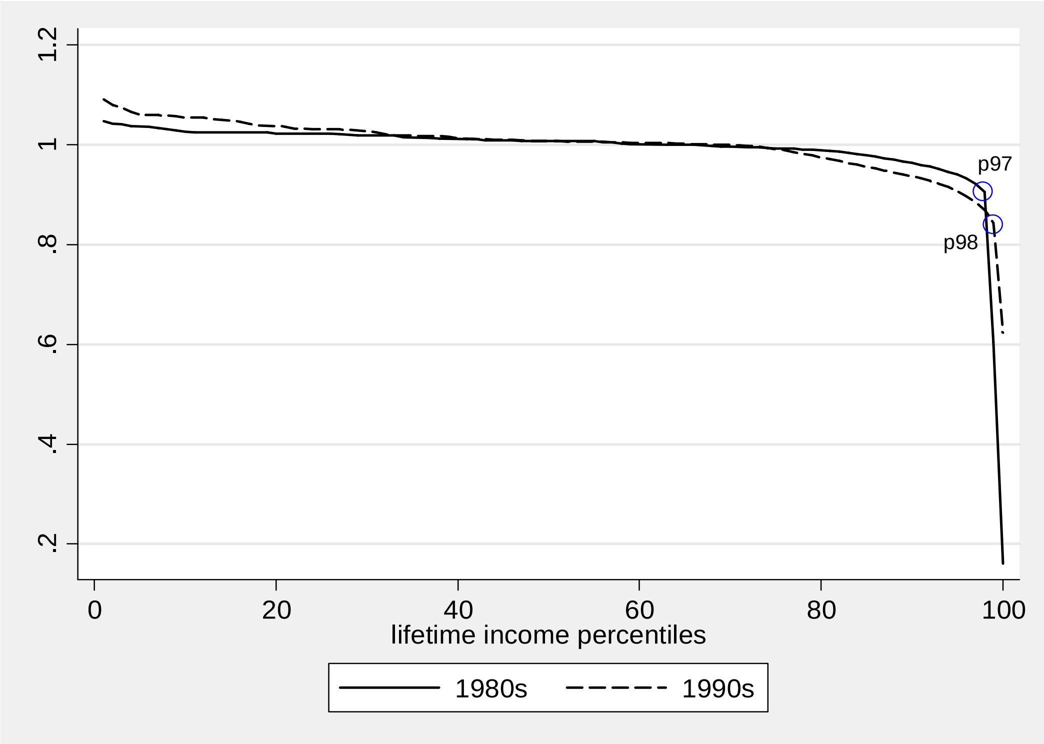


Figure 2 – Average predicted hazard ratios by income percentile and decade: males



Notes: Two free knots spline model, the circles indicate the positions of the estimated knots:.p62 is the 62nd percentile, p66 is the 66th percentile and p98 is the 98th percentile of the period-specific income distribution.

Figure 3 – Average predicted hazard ratios by income percentile and decade: females



Notes: One free knot spline model; the circles indicate the positions of the estimated knots. P97 is the 97th percentile and p98 is the 98th percentile of the period-specific income distribution.

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