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Lifetime Income and Old Age Mortality Risk in Italy Over Two Decades

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Lifetime income and old age mortality risk in Italy over two decades[*]

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Abstract

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 extend the literature by estimating the income–mortality gradient using Cox proportional hazard models, where the positions of the knots in the spline function for income are determined by the data, and, most importantly, by providing empirical evidence for Italy on the evolution of the income–mortality gradient between the 1980s and the 1990s. In addition, we investigate the importance of controlling for regional differences. For this purpose, we use data drawn from an administrative pension archive held by the main Italian social security institution and proxy individual lifetime income with the amount of individual pension benefit.

We find that the shape of the income–mortality gradient is characterized by two discontinuities (knots) for males and one discontinuity for females. For both genders, the estimated associations between income and mortality risk are negative and stronger at higher income levels.

Results for the change in the income–mortality gradient between the 1980s and 1990s are sensitive to inclusion in the model of controls for regions. Without controlling for regional differences, the income-mortality gradient widens between the 1980s and the 1990s due to a rather surprising positive gradient for males whose income is below the 66th percentile in the 1980s. One possible explanation for this result could be the presence of manual workers with relatively high salary and high mortality during the 1980s in the industrialized Northwest. After controlling for regional differences, we find no empirical evidence of a change in the income–mortality gradient, including the position of the knots, over time.

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. Income, wealth, or educational attainment are used to define and classify individual socioeconomic status (SES) (Galobardes et al., 2006). A significant negative correlation between SES and overall mortality is nearly always found, with varying intensity; for comparison across European countries, see, e.g., Mackenbach et al. (2008), Mackenbach et al. (1999) and Mackenbach et al. (1997).

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.

European studies highlight a widening of relative inequalities in general 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), as well as for Italy specifically (Mackenbach et al., 2003), represent an exception to these trends. The time variation in the SES-mortality gradient has been smaller in Southern than in other European countries and the change in cardiovascular disease mortality is more homogenous across socioeconomic groups. The comparative research by Mackenbach et al. (2003) cited above considers two broad socioeconomic groups: high versus low educational level, and manual versus non-manual occupations. Italy contributed to this research 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). The TLS covers only residents of the city of Turin, so information on a national scale is still lacking.

Most data published on inequality refer to people in the working-age group; social inequalities in general 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 (see, e.g., Sullivan and von Wachter, 2009).

Finally, it has been observed that the decrease in mortality risk becomes less marked with increasing income, following a concave relationship, as a result of 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. We extend Leombruni et al. (2010), using a 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, and more importantly, we provide empirical evidence for Italy on the evolution of the income-(old age) mortality gradient between the 1980s and the 1990s. Such quantification is still lacking for Italy; available evidence (Mackenbach, 2003) considers other SES indicators and refers to the city of Turin.

This paper is organized as follows. The Materials and methods section presents the data, sample selection and the econometric approach used in the analyses. The Results section describes our main findings. Finally, the Discussion section discusses the findings and sets forth our conclusions.

2 - Materials and methods

Study population and income measure

We exploit a pension file 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, 1933 to 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. 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, early retirement, disability, survivors'). In addition, there is data on individual 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 last payment received.

Our study includes ex-private sector employees aged 65 and older. Age selection (65+) is dictated by the data, which is on pension beneficiaries, and by the retirement rules applied to employees. Until 1994, males (females) could claim an old age pension at age 60 (55). After a phase characterized by gradual increments, the minimum age for the old age pension was set at 65 (60) for males (females) in 2001. Empirically, at age 65 almost everyone is retired in the analyzed period (Belloni and Alessie, 2009). To facilitate comparisons, we apply the same age selection to both genders. Although we could have extended the analysis to younger individuals (e.g., to those aged 50+), this choice would have raised an issue of sample self-selection: those with worse health status may have retired earlier and be over-represented.

Individuals are assumed to be at risk of death starting from January 1979. From a first investigation, it appeared that the quality of the variable *date in which the pension flow ended* is rather poor before that date. Note that individuals retired before 1979 *are* included in the sample, provided that they are alive in January 1979. Even in the first years of the sample, therefore, there is information on mortality in our sample. We include cohorts born between 1901 and 1936. Cohorts born before 1901 are excluded, since coverage by the pension system for private sector employees was partial and participation was voluntary for them.

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). For the same reason, we were forced to exclude the self-employed. (However, we include ex-employees who also receive self-employment pensions, provided that the latter is a minor portion of their total pension income). Due to the pension rules, benefits they received for most of the period covered by the data were not informative with respect to lifetime labor income.

There is a minimum, but no maximum, pension benefit in Italy: if the accrued benefit is below the minimum pension, the difference is subsidized, provided that an earnings test is passed. To reduce measurement error, we exclude individuals whose total pension income is below the threshold. We exclude possible outliers with very high pension income by trimming the income distribution at the top per mille.

Statistical analysis

The analysis is performed on monthly data, from January 1979 to December 2001, and separately by gender. 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 thus end up with a total of 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 separate Cox models (Cox, 1972) by decade: individuals alive in December 1990 are treated as right censored in the first period and as left truncated in the second period. Mortality risk is assumed to depend on survival time 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 to compare 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 the 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.

Free knots models outperformed 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 testing 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 the testing results, which are reported in the appendix, 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.

To formally check whether the slope parameters jointly changed over time, we implement a Hausman test, which takes into account dependence across samples (right-censored individuals in the 1980s sample also belong to the 1990s sample); the variance of the difference between the two income parameters vectors is obtained by paired bootstrapping (Cameron and Trivedi, 2005).

3 - Results

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. 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. Throughout this period, the income distribution moved to the right and became somewhat more compressed. There was a moderate reduction in pension income inequality (see Gini’s indices).

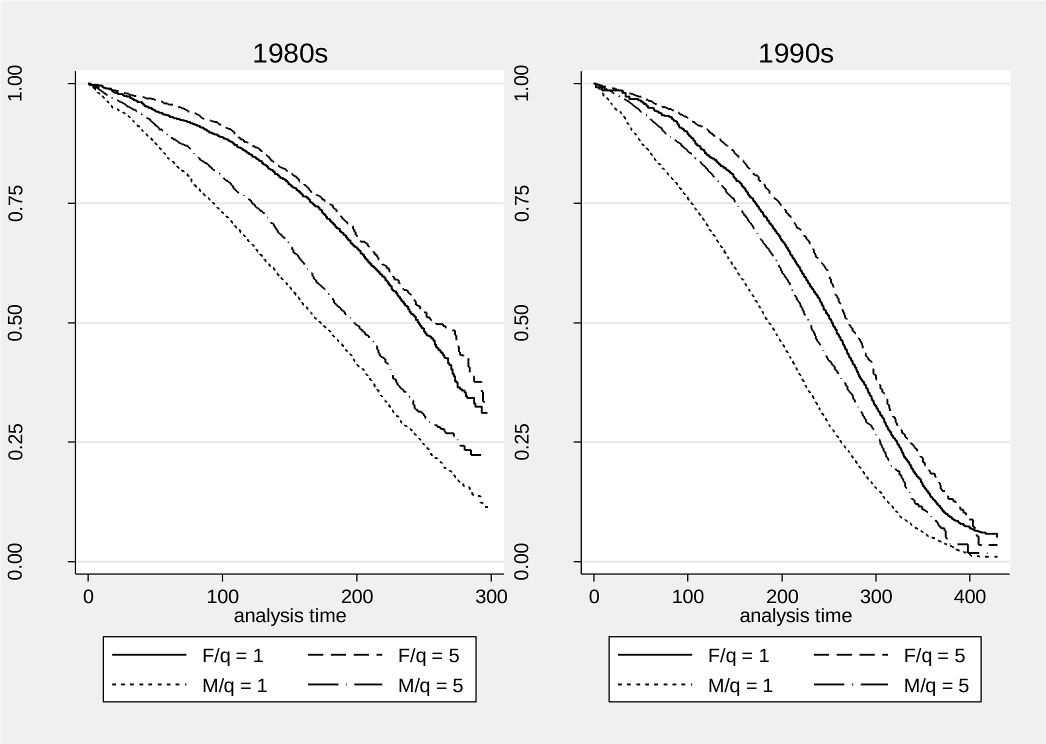
Table 1 – Descriptive statistics

		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:	*1th*	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).

Figure 1 reports Kaplan-Meier survival estimates by gender and income quintile (first *versus* fifth) for the two analyzed decades. In the 1980s (left panel), 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 (absolute difference 14 months, relative difference 6%). In the 1990s (right panel), 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 are characterized by 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.

Figure 1 – Kaplan-Meier survival estimates by gender (F for females and M for males) and income quintile (q = 1 and q = 5): 1980s and 1990s



Parameter estimates for the selected free knots spline models are reported in Table 2 and predictions based on these estimates are presented in Figures 2 and 3. 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%. This surprising result is shown in Figure 2 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), such association is found to be 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), such association becomes much steeper: a 1% increase in lifetime income reduces mortality risk by 3.6% (*income 3*). In Figure 2 this is shown 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 discussion), 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 basically 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 found to be 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 gradient.

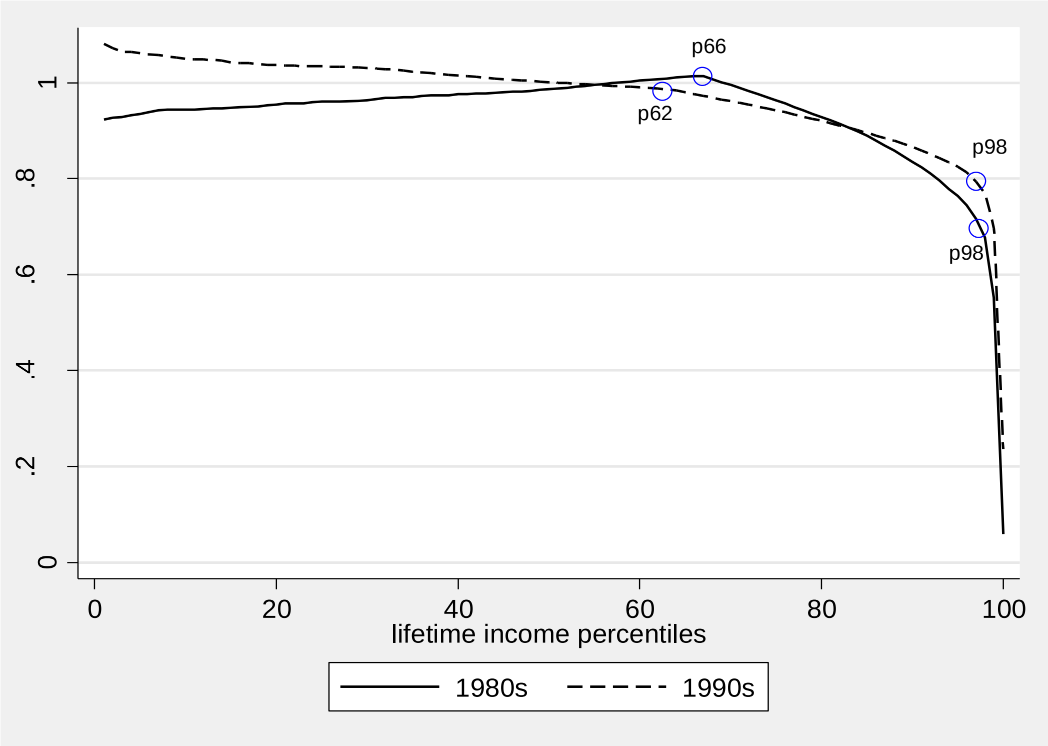
For females, the (single) knot is found positioned at the top of the income distribution (97th percentile in the 1980s, and 99th percentile in the 1990s). The sign of the estimated associations between income and mortality is found to be always negative for females. The size of such association is much smaller below the estimated knot—i.e., for almost the entire population—than above it. This result is similar to what is found for males, second knot. In the 1980s, the estimated association between income and mortality is not statistically different from zero below the knot. Also somewhat similarly to males, the association deepens below the knot (cf. *income 1* in the 1980s and 1990s) and weakens above it (cf. *income 2* parameters). Nevertheless, as the Hausman test shows (*p*-value 0.67), these changes are not statistically significant. This finding suggests an invariant income–mortality gradient for females over these two decades.

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				

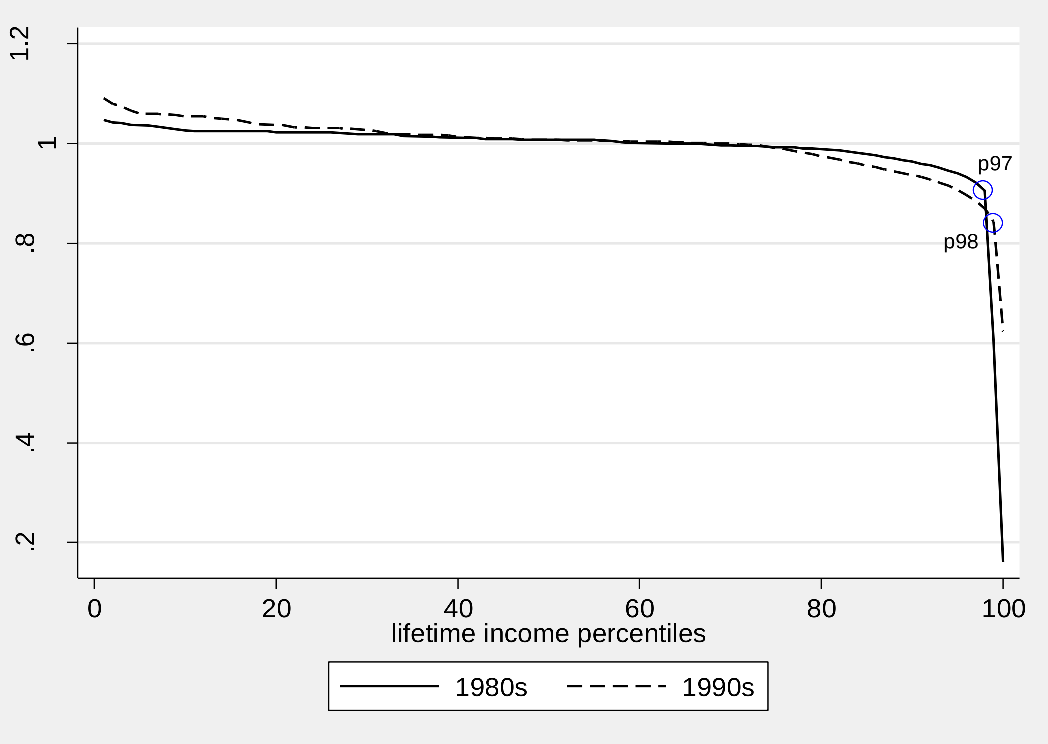
Notes: models by gender and decade; spline function with endogenously determined knots; *income x* is the percentage change in the hazard of death associated with a 1% increase in the lifetime (pension) income for levels of log-income in the interval [*knot x-1*, *knot x*]; *knot x* is the estimated 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.

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



Notes: Two free knots spline model; the circles indicate the percentiles in which the estimated knots are positioned.

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



Notes: One free knot spline model; the circles indicate the percentiles in which the estimated knots are positioned.

Finally, we return to the surprising result in Figure 2 of the upward slope up to p66. In Table 3 we present results for models that include area of birth as an additional covariate; we show results only for males. The results do not vary substantially with respect to those presented in Table 2 for females nor, for both genders, 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). Results show a negative estimate for the *income 1* coefficient, not statistically different from zero for the 1980s; also, a higher estimate for the *income 1* coefficient is found 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 the country (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 $62^{nd}/66^{th}$ lifetime income percentile), since it especially affects manual workers.

Results set forth in Table 3 also show that, controlling 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 (*p*-value 0.07) 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. Regional differences in mortality—higher mortality in the Northwest, particularly in the 1980s—and income thus explain a large part of the time evolution of the income–mortality association reported in Table 2.

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.

4 - Discussion

This is the first Italian study on the trend of association between income and mortality risk on a national scale; it also provides evidence on the shape of the income–mortality association for the country. The analysis is conducted in a sample of ex-private sector workers aged 65 and over; it is performed on a newly available pension file 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:

1. is negative, and stronger for higher income levels (i.e., increasing returns to income); is positive during the 1980s for males whose income is below the 66nd percentile, but this finding disappears once controlled for area of birth;
2. is represented by two discontinuities and at the 66nd and the 98th percentiles of the income distribution for males and one discontinuity at the 97th percentile for females; these kink points are almost at the same positions during the 1980s and the 1990s;
3. became steeper between the 1980s and 1990s for males, but regional differences in mortality and income explain a large part of this change; and
4. remained unchanged between the 1980s and 1990s for females.

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 was below the minimum pension and trimming of top incomes to exclude possible outliers with different thresholds. Standard Cox models with income quantiles or spline functions with exogenously determined knots as explanatory variables were estimated and compared with our non-standard Cox specifications. Results turned out to be robust to all these modifications.

This study's findings may be affected by several data limitations. Both the dependent and main independent variables of our model may potentially be measured with error: individuals' date of death is reconstructed by using information on the date their pension flow ended. Note that where an individual obtains more than one pension during his 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 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 that individual's total pension income. 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.

Our proxy for lifetime income is individual pre-tax income. Mortality has been shown to be more strongly associated with individual taxable income than with household taxable or disposable income (Martikainen et al., 2009). These associations can especially differ for females: in the literature, there is evidence that individual income, more than spousal income, is correlated with mortality. Kalwij et al. (2013) find that partner income is only weakly associated with mortality risk for females. Undoubtedly, results for females should be considered with caution; the general female population is particularly under-represented in the sample, since female participation in the labor market is lower than that of males (compared with population statistics from HMD, the percentage of females in the sample is about 8 percentage points lower) and is probably determined more by social and health conditions (see, e.g., Fokkema, 2002).

In our analysis, individuals are at risk of death starting from age 65. In this way, we avoid over-representation of individuals in poor health who tend to retire early from the labor market. Obviously, our sample is more selected because retirees who die before age 65 are not accounted for. Because of this, the estimated income–mortality associations may be lower than what we would have obtained by including all retirees. Our main findings in terms of both time evolution and shape remain unchanged when we include all retirees (age 50+), although the size of the estimated associations are indeed somewhat higher in this latter case.

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). Similar to our findings, Leombruni et al. (2010) find that only retirees in the top income quintile 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 results for poorer males in the 1980s, as well as the overall change in the income–mortality association, can be partly explained by regional differences. Data on educational attainment and working career characteristics could certainly yield clarification. Findings for other countries show that, after adjusting for other SES factors, the income–mortality association is typically flattened, although its shape is not evidently different (Martikainen et al., 2009).

Results for top earners, whose estimated shape of the income–mortality association is particularly steep, should be considered 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, this group’s observed pension income can be a poor proxy of their lifetime spending potential, for the following reasons. First, this group’s lifetime labor income (approximated in this exercise by pension income) can be a poor proxy of total earnings; earnings from financial and housing wealth represent a high proportion of total earnings. 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 often enrolled in special pension schemes; their actual pension income can be higher than what is observed: lump-sum compensation or additional annuities (unobserved in this data) were often granted by employers to get rid of these highly costly workers or as a final premium for their working careers. Finally, this group’s work-package agreement was extremely generous, often including private health insurance coverage and various in-kind benefits; their lifestyle was higher than what is expressed by income.

Our results on the time evolution of the income-mortality association confirm those in Mackenbach et al. (2003) for the city of Turin and for 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 mostly concerns the Nordic countries and the United Kingdom and not Southern EU countries, due to a lack of longitudinal data on mortality. As stressed by Mackenbach et al. (2003), 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. Obviously, in this study we cannot analyze behavioral changes, nor do we have cause-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, this gradient appears to be rather stable over time.

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 thus given by an "indirect" LR-test (Molinari, 2001) in which each model is compared against the log-linear model.

Table 4 – Model selection

Model:	**d.f.**	**-2* 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 against the log-linear model, *p*-value.