Early-life conditions, lifetime income and mortality risk in Italy

by Michele Belloni [a,e,g], Adriaan Kalwij [c,e,f] and Rob Alessie [b,e]

[a] Ca'Foscari University of Venice michele.belloni@unive.it
[b] University of Groningen
[c] Utrecht University School of Economics
[e] Network for Studies on Pensions, Aging and Retirement (Netspar)
[f] Tilburg University
[g] CeRP - Collegio Carlo Alberto

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Abstract

In this study, we obtain insights into the impact of early-life conditions on old-age mortality in Italy. We capture cohorts' life conditions by means of mortality rates at different early-life stages and exploit exogenous variation provided by a series of abrupt mortality events which severely affected specific cohorts. We also test whether lifetime income – approximated by the amount of the pension benefit - is health protective against bad circumstances experienced in early life.
Early-life conditions have a long-lasting effect on males' mortality. Results suggest the existence of a considerable "scarring" effect: the death probability of a male born in 1932 alive at age 65 is 15% lower than that of a male born at the beginning of the XX-th century due to improved early-life conditions. For females, we do not find a significant impact of early-life conditions. We do not find evidence that income is health protective. On the contrary, we find that especially mortality of richer individuals is affected by circumstances in early life.

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1. Introduction

This paper analyzes the impact of individuals' early-life conditions on mortality at older ages in Italy. A wide empirical literature worldwide exists on this topic. In this literature, it is outlined that the risks of death can be linked through either "scarring" or "immunity" (Preston et al., 1998). Scarring determines a positive link across death risks at different ages. Certain diseases or conditions acquired in childhood and at birth may permanently weaken the survivor and increase her probability to contract a disease and die at subsequent ages. Many possible causal pathways connecting early life experiences and later mortality have been discussed (for a short review, see, e.g., Elo and Preston, 1992). Barker (1994, 1995) suggests that preconditions for coronary heart disease, hypertension, stroke, diabetes, and chronic thyroiditis are initiated in utero and become manifest much later in life (so called "fetal origin hypothesis"). Fridlizius (1989) points out that exposure to certain infectious diseases (such as smallpox, and measles, see also Lunn, 1991) early in childhood may determine a permanent dysfunction of the immune system and increase the risk of contracting other infectious diseases in later life stages. Fogel (1994) proposes scarring mechanisms based on malnutrition in utero and during early life and development of chronic diseases. Immunity leads to an inverse association across death risks over the life cycle. Individuals exposed to certain diseases such as influenza acquire immunization and might me expected to have a lower probability to contract similar diseases at subsequent ages.[1]

Various studies use mortality rates (in utero, at birth and during childhood) as early-life conditions indicators. The evidence is inconclusive. Two pioneering studies are Kermack et al. (1934) and Pearson (1912). The former study finds that mortality up to age 15 correlates positively with subsequent cohort mortality in Great Britain and Sweden. The latter discovered instead an inverse association between infant mortality and death rates from age 1 to 5 for England and Wales. More recently, Bengtsson and Lindstrom (2003) use historical data on parish registers in Sweden and find a

[1] At the population level, there can also be a selection effect, i.e. as a consequence of a bad event, weaker members of a cohort die, while survivors are stronger and live longer. The effect of immunity and selection go in the same direction.

strong positive correlation between local infant mortality rate and mortality at ages 55-80 while do not find any effect of the disease load on mothers during pregnancy. Catalano and Bruckner (2006) find a positive relation between mortality at ages up to 5 and life expectancy at age 5 for Sweden, Denmark, England and Wales. Bruckner and Catalano (2009) report that the association between infant mortality and later mortality decreases with age and disappears in adulthood. Myrskylä (2010) uses cohort-level data for six European countries and finds that early-life conditions have a transitory effect and potentially only little influence old-age mortality. For Italy, Caselli and Capocaccia (1989) use life table data for the cohorts 1882-1953 to study the relation between mortality risk at birth and during childhood and mortality risk at ages 25-79. They find an age-varying effect: mortality risks are positively correlated up to age 45 and negatively correlated (but to a small extent, and only for males) at old ages.[2]

Full life course epidemiology (see, e.g., Hall et al., 2002; Ben-Shlomo and Kuh, 2003; Kuh and Ben-Shlomo, 1997) expands the links between early life and later mortality further to take into account timing and order of different events throughout life. Early-life conditions may have a direct permanent effect on adult health and mortality, or an indirect effect, via its influence on obtained socioeconomic status (Preston et al., 1998; Bengtsson and Broström, 2009).[3] Moreover, the effect of early-life conditions on later mortality may depend on the socioeconomic status: wealthier or richer individuals may ("buy" health, in economic jargon) and recover from bad circumstances experienced in early life.

The contribution of this paper to the literature is twofold. First, we obtain insights into the impact of early-life conditions on old-age mortality in Italy. No evidence based on micro data exists so far for this country. Second, we test whether lifetime income is health protective against bad circumstances experienced in early life.

[2] A recent series of studies show that the state of the business cycle at birth (Van Den Berg, Lindeboom and Portrait, 2006) and in the first years of life (Van Den Berg, Doblammer and Christensen, 2009) are important determinants of later mortality.

[3] Many studies find a negative association or causal relation between socioeconomic status and mortality; for a comparison across European countries, see, e.g., Mackenbach et al. (1997, 2008). Social inequalities in mortality are also found for European elderly; in comparison with younger age groups, elderly males typically show lower (higher) relative (absolute) differences in mortality rates by socioeconomic status (Huisman et al., 2004).

We capture early-life conditions by means of mortality rates during childhood. We allow for different postnatal human development stages to have a different impact on old-age mortality. The cohorts we analyze (1901-1936) experienced an overall declining trend in mortality. We identify scarring or immunization effects using further variation provided by a series of abrupt mortality events - such as the earthquakes of 1908 and 1915, the Word War I and the Spanish' flu of 1918 - which severely hit their childhood and adolescence lives.

We approximate individuals' socioeconomic status with their lifetime income. Individual information on mortality and lifetime income is obtained by a new available pension file drawn from an administrative archive held by the main Italian Social Security Institution. We have a precise proxy variable of the individual's lifetime income at our disposal, namely the amount of the pension benefit.

The paper proceeds as follows. Section 2 presents the data and sample selection. Section 3 describes the mortality risk model. Section 4 shows the main results and section 5 concludes.

2. Data and sample selection

Micro data

Information on individuals' mortality and income are obtained from 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. 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 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.[4]

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. 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. Individuals may retire earlier than 65 but this may be related to health (hence to mortality risk); we therefore did not take the time before age 65 into account. 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.

We proxy individual lifetime income by the amount of pension benefit received (Gaudecker and Sholz, 2007). 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. 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 below the threshold and exclude possible outliers with very high pension income by trimming the income distribution at the top per mille.

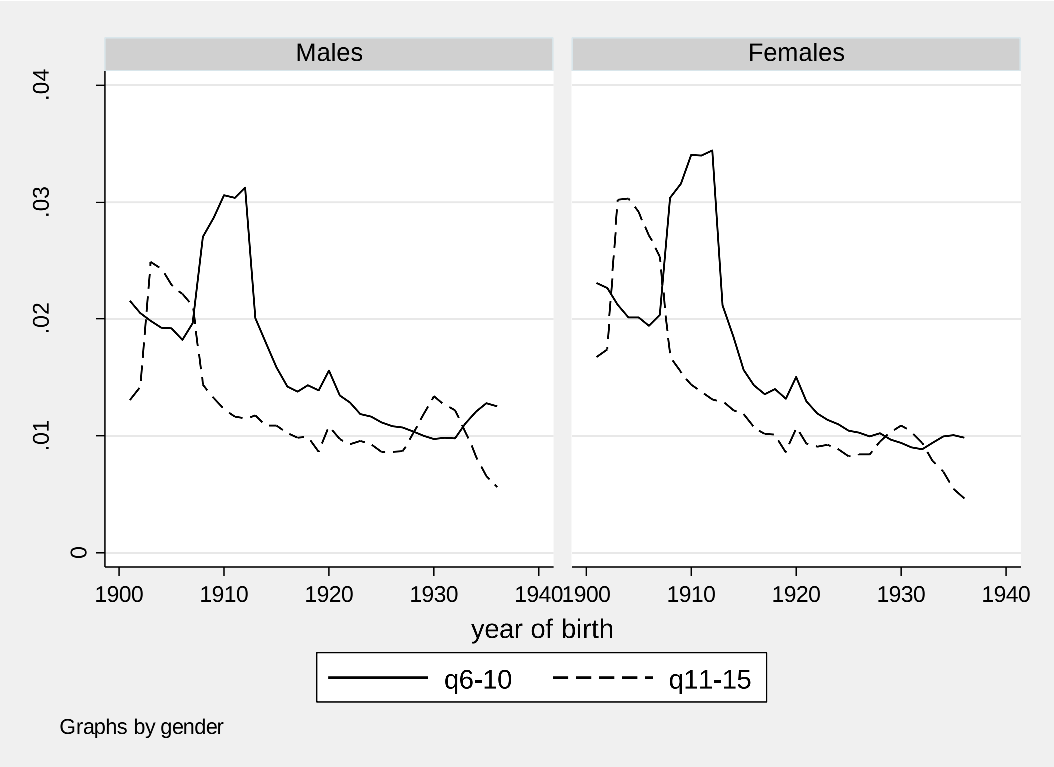
[4] When an individual obtains more than one pension during his or her life we examine the most recent ending date. In this way, it is possible to adequately deal with any inaccuracy resulting from the existence of other possible reasons for stopping a specific pension payment, such as conversion of 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).

Early-life conditions data

We merge the micro data on pension benefits described above with cohort-level (*yob*) death probabilities at ages 0-15 from Human Mortality Database (2012) by gender (q_j^{yob} where *j=0,…,15*). These data are presented in figure 1 separately for the age groups 0 and 1-5 (q_0^{yob}, q_{1-5}^{yob}, top panel), 6-10 and 11-15 (q_{6-10}^{yob} and q_{11-15}^{yob}, bottom panel). The figure makes clear that the cohorts we analyze experienced an overall decreasing trend in mortality at all childhood ages. It also reveals a series of mortality peaks, corresponding to some known historical events, which halted this increase: the devastating earthquakes of 1908 and 1915, Word War I (1915-1918), the 1918 Spanish flu and the 1919-20 smallpox outbreaks.

Mortality was very heterogeneous across Italian regions, especially in correspondence of the above outlined abrupt events. The 1908 earthquake hit Calabria and Sicily while in 1915 earth tremors shocked the Marsican area and the Abruzzi. The Spanish' flu epidemic affected mainly Central and Northern Italy, although some regions such as Liguria and Veneto were left relatively untouched Pinnelli and Mancini (1999).

Figure 1 - death probabilities by year of birth, age group and gender



3. Method

The statistical analysis is performed on monthly data, from January 1979 to December 2001 (follow-up 276 months), separately by gender. A preliminary analysis of the association between lifetime income and survival was obtained from Kaplan-Meier survival estimates by gender and income quintiles.

We estimate a Cox model (Cox, 1972) by gender. We assume that the individual mortality risk $\theta(t, X_i, q^{yob(i)}, g(Y)_i)$ depends on survival time (age) t, a set of time-invariant characteristics X for individual i (X_i), mortality probabilities during childhood experienced by the cohort to which the individual i belongs ($q^{yob(i)}$), and a function of lifetime (pension) income $g(Y)_i$. Ln(θ) is linear in X_i, $q^{yob(i)}$ and $g(Y)_i$.

In the empirical specification we allow for different postnatal human development stages to have a different impact on old-age mortality. Specifically, we consider the following age groups separately:

infancy (first year of life), preschooler (ages 1-5), primary school age (6-10), preteen-adolescence (11-15). Formally, $q^{yob(i)} = \{q_0^{yob(i)}, q_{1-5}^{yob(i)}, q_{6-10}^{yob(i)}$ and $q_{11-15}^{yob(i)}\}$. X_i includes year of birth, a full set of region of birth dummy variables, retirement age and a dummy variable indicating whether the individual has received a disability pension.

A linear cohort trend (year of birth) is included in the model to capture long-term improvements of circumstances in early life. The estimation of the effect of $q^{yob(i)}$ variables on mortality exploits therefore (exogenous) cyclical deviations around this trend.[5] Figure 1 shows that mortality rates at various child developments stages are not perfectly collinear which allows us for a joint estimation of the effect of all q^{yob_i}. As outlined in the introduction, early-life conditions may have an impact on later mortality in two ways: through ``scarring'' and/or ``immunity''. The two effects counteract, and empirically only the net effect can be identified. If the scarring effect prevails, parameter estimates of $q^{yob(i)}$ variables have a positive sign; viceversa they have a positive sign.

Including (lifetime) income allows us to disentangle direct effects of early-life conditions on later mortality from their indirect effects which can occur through socioeconomic status (see introduction). We specify $g(Y)_I$ as a set of dummy variables for each income quartile. This chosen specification relies on a preliminary descriptive data analysis (see next section); it also draws from previous literature for Italy on the income-mortality association (Belloni et al., 2012; Leombruni et al., 2010) which reports evidence that mortality differentials are concentrated in the upper part of the income distribution. In the most flexible proposed specification, we allow for the effects of early-life conditions on mortality to depend on income. In this way, we investigate whether income is health protective against early life circumstances. To this purpose, we include in the model a full set of interaction terms between income quartiles and $q^{yob(i)}$ variables (a total of 12 interaction terms).

[5] We tried alternative specifications based on the cycle/trend decomposition of mortality probabilities using the Hodrick-Prescott filter (with filter values 100 and 500). This is a well-known approach in the literature of early life conditions (Bengtsson and Lindstrom, 2003; Van den Berg et al., 2006; Van den Berg et al., 2009). We also experimented with including in these models time effects, which we modeled with calendar year dummy variables (from 1979 to 2001). Results were very similar.

We also tested the *foetal* origins hypothesis (Barker, 1994) by including in the model female mortality rates at fertile ages (between age 20 and 40) in the conception year (i.e. $q_{20-40}^{yob(i)-1}$ in our notation). This additional variable should capture mothers' conditions during pregnancy. We find no evidence that intrauterine conditions affect mortality at old age. We also investigated the effect of socioeconomic conditions at birth on mortality (see introduction). Unlike Van den Berg, Lindeboom and Portrait (2006), we do not find a significant effect of the state of the business cycle at birth. It should be recognized that most of the studies which find an effect of GDP at birth-type of variables on later mortality pertain to periods dating back to the XIX century or earlier, whereas we analyze mortality in more recent times.

4. Results

Table 1 reports descriptive statistics for the estimation samples. Altogether, there are 33,542 failures (deaths). Male (female) median survival time is equal to 191 (257) months (15.9 and 21.4 years respectively). Average monthly pension income is higher for males than for females. Figures 2 and 3 show Kaplan-Meier survival estimates by quartile of lifetime (pension) income for males and females respectively. Male median survival times are equal to 177, 188, 188 and 215 months for the four income groups; the corresponding figures for females are 249, 257, 262 and 266. The absolute difference between poorest and richest group of males (females) is thus equal to 38 (17) months and the relative difference is 21 (7)%. These figures show that differential mortality is much more sizable for males than for females; they also show that while for males it is especially the upper income quartile which benefits of lower mortality than the rest of the sample, for females the poorest income group is characterized by a sensibly higher mortality than the other groups.

Table 1 – Descriptive statistics

		Males	***Females***
Subjects		49045	34123
Failures (%)		22907	10635
		(.47)	(.31)
Survival time, median*		191	257
Monthly pension income (€, 2009 prices):			
Mean		1102.17	581.67
Standard deviation		1791.25	872.27
Percentiles:	*25st*	437.81	340.54
	50th	735.58	444.57
	75th	1268.88	520.76
Retirement age, mean**		58.43	55.70
Disability pension, %		.22	.25

Note: * Kaplan-Meier survival estimates, ** years

Figure 2 – Kaplan-Meier survival estimates by quartile of pension income: males

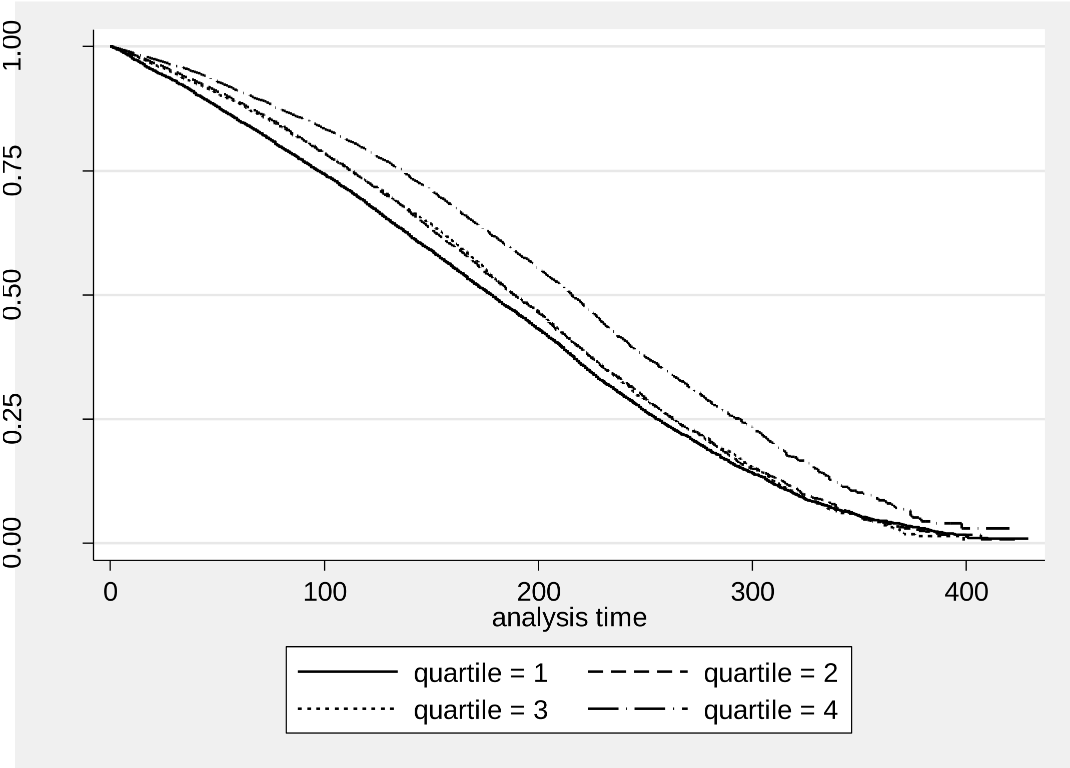


Figure 3 – Kaplan-Meier survival estimates by quartile of pension income: females

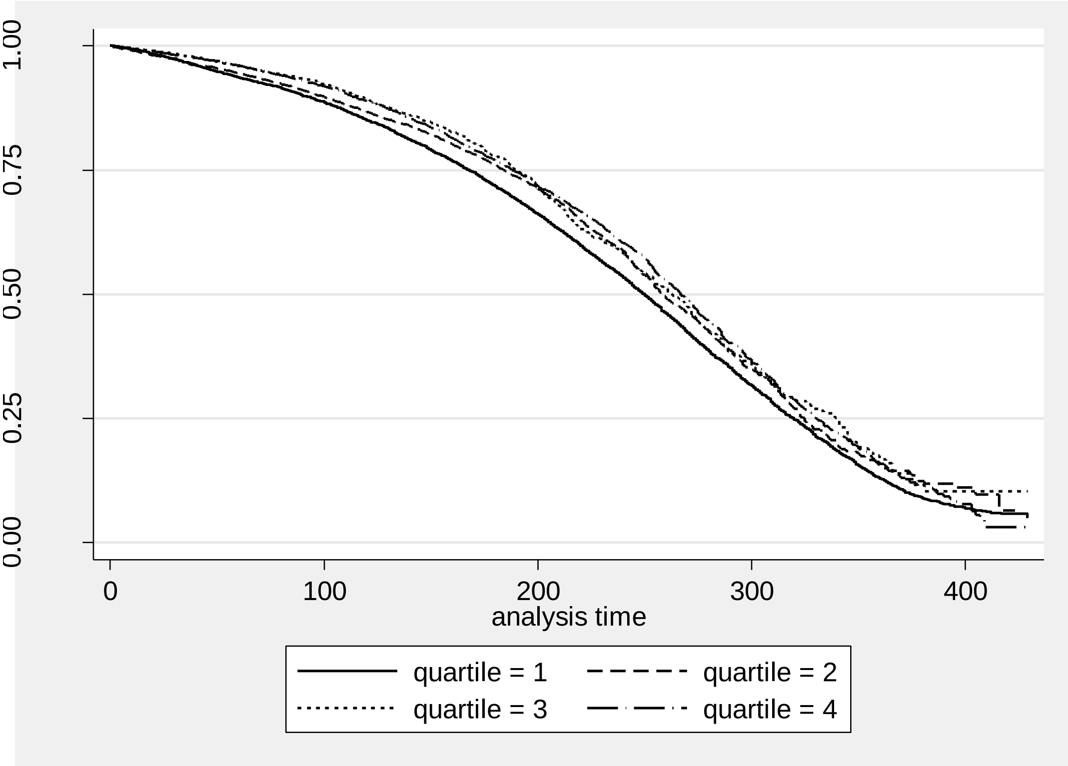


Table 2 reports parameters estimates for the Cox models (coefficients). For both genders, it shows results for three alternative specifications (spec. M-I,M-II,M-III; F-I,F-II,F-III). In spec. (I) the effect of lifetime income on mortality is restricted to be zero; spec. (II) relaxes the assumption of spec. (I) and includes quartiles of total pension income as additional explanatory variables (see section 3); spec. III) adds to spec. II) a full set of interaction terms between income quartiles and $q^{yob(i)}$ variables (a total of 12 interaction terms). In all shown specs. we control for region of birth (dummy variables), retirement age and whether the individual has received a disability pension (dummy variable).[6]

Early-life conditions have a long-lasting effect on male mortality, see model (M-I). Between the child development stages included in the model, preschooler (ages from 1 to 5) seems to matter the most to determine old-age mortality: the estimate of the $q_{1-5}^{yob(i)}$ variable is highly significant (p-value=0.007). Also circumstances during primary school age (from 6 to 10) are found important (p-value=0.099; the null hypothesis of equality between the effect of $q_{1-5}^{yob(i)}$ and $q_{6-10}^{yob(i)}$ is not rejected by

[6] Results are found robust to different sets of control variables. We also included the individual' month of birth as additional control, with no effect on the results.

a Wald-test). Positive signs for these two variables suggest the existence of a “scarring effect”: individuals grown in worse times have higher death probabilities at old ages than those grown in better times. We do not find a significant effect on male mortality in old age of life conditions during infancy, as well as during preteen-adolescence (ages 11-15). For females, we do not find a significant impact of $q^{yob(i)}$ variables on old age mortality (see model F-I).[7]

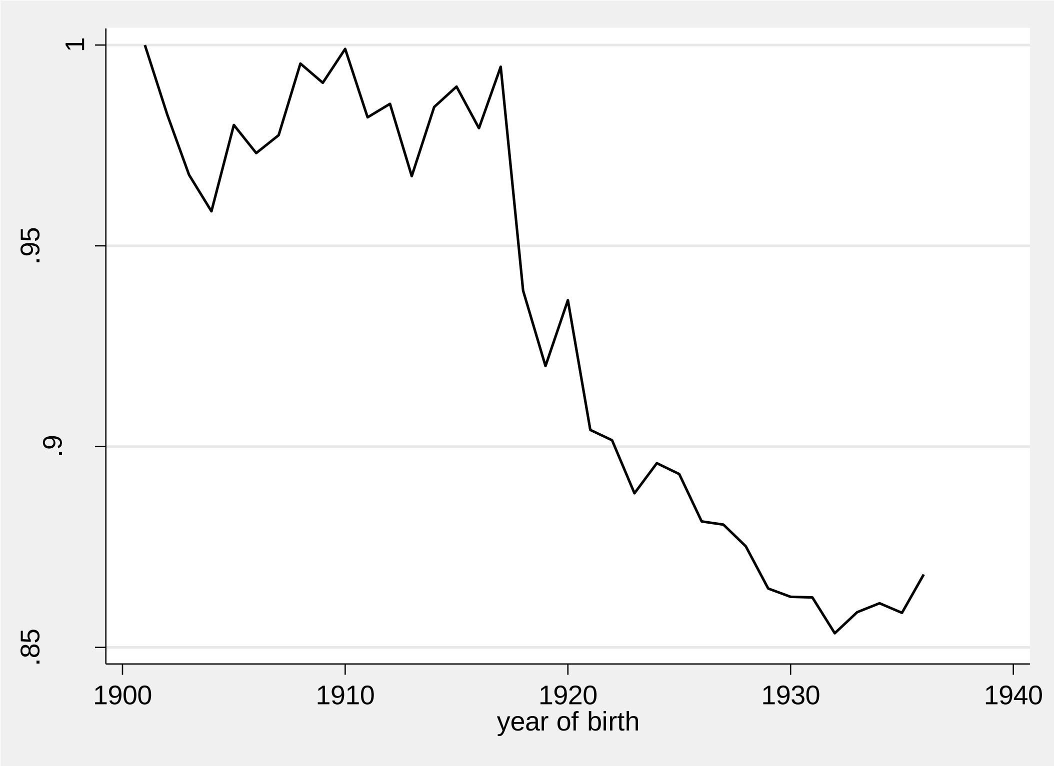
Higher lifetime income is significantly associated to lower mortality risk (spec. M-II and F-II). The risk of death of a male in the top income quartile is 23% lower than that of a male in the bottom one (M-II; hazard ratio: exp(-.26)=.77). For females, we find a much lower income-mortality gradient (F-II; hazard ratio: exp(-.09)=.91), with differential mortality more homogenously distributed throughout the whole income distribution. These results qualitatively confirm previous findings for Italy on the association between income and mortality (Belloni et al., 2012; Leombruni et al., 2010). Importantly, specs. (II) show that estimated parameters for the $q^{yob(i)}$ variables do not vary meaningfully if lifetime income is included in the model (cf. model M-I versus M-II, and model F-I versus F-II).

Figure 4 reports the predicted change in the hazard, attributed to changes in early-life conditions across cohorts (hazard ratios, reference cohort 1901; spec. M-II).[8] In comparison with the oldest cohort in the data, younger cohorts live longer because of improved early-life circumstances. These improvements especially concern individuals born after the World War I, who experienced crucial improvements in living standards during their childhood with respect to the pre-war conditions. The figure highlights that the cohort of 1932 has the greatest benefit from improved early-life conditions: due to these changes, its risk of death would be almost 15% lower than that of the 1901 cohort.

[7] Results for females are in line with those in Caselli and Capocaccia (1989). For males these authors report evidence of a negative correlation across death risks over the life cycle for individuals aged 45 and older. However, the comparison with our results (age 65+) is difficult since they find an age-varying effect of early-life conditions on mortality. A tentative comparison can be made from figure 3 of Caselli and Capocaccia (1989). From this figure it looks that for ages 65+ the effect is very small; moreover, standard error of the estimates are not reported in the paper.

[8] In this simulation, we allow for the $q^{yob(i)}$ variables to vary by cohort, while keeping all the other variables constant. We compute: $\theta(t, X^{yob})/\theta(t, X^{1901}) = \exp((q^{yob(i)} - q^{1901(i)})\hat{\beta})$, where $\hat{\beta}$ is the vector of estimated parameters for the $q^{yob(i)}$ variables.

Figure 4 - Predicted change in the hazard, attributed to changes in early-life conditions across cohorts – hazard ratios, reference cohort 1901: males



Younger cohorts have a lower risk of death; we also find evidence of a wide heterogeneity in mortality risk across Italian regions (not reported in Table 2, region of birth dummy variables are jointly highly significant). It is also worth mentioning that earlier retirement age and having received a disability pension are found associated with higher mortality risk. These results are common to both genders.

Spec. (III) allows for the effect or early-life conditions on mortality risk to depend on income. We do not find evidence that income is health protective. On the contrary, we find a “scarring” effect of early-life conditions – more precisely of $q_{1-5}^{yob(i)}$ – on mortality risk only for individuals belonging to the top income quartile (the null that the total effect of $q_{1-5}^{yob(i)}$ for the top income quartile is equal to zero is rejected at 1% level; the same hypothesis for the other income groups, as well as for the other $q^{yob(i)}$ variables for all income groups are not rejected).

Table 2 – Cox model for mortality risk: parameters estimates

VARIABLES	M-I	M-II	M-III	F-I	F-II	F-III
YQ2		-0.00485	-0.0509*		0.00280	-0.181***
		(0.0182)	(0.0298)		(0.0285)	(0.0552)
YQ3		-0.0180	-0.0786***		-0.0769*	-0.258***
		(0.0201)	(0.0305)		(0.0396)	(0.0601)
YQ4		-0.260***	-0.330***		-0.0864***	-0.280***
		(0.0216)	(0.0320)		(0.0294)	(0.0586)
q_0^{yob}	0.0209	-0.0697	-1.436	0.115	0.00826	-1.903
	(0.636)	(0.636)	(0.931)	(0.965)	(0.966)	(1.283)
$q_0^{yob} \times YQ2$			2.731*			1.038
			(1.458)			(2.154)
$q_0^{yob} \times YQ3$			3.096**			3.330
			(1.514)			(3.600)
$q_0^{yob} \times YQ4$			-0.136			4.957**
			(1.702)			(2.308)
q_{1-5}^{yob}	1.433***	1.460***	0.680	1.086	0.992	-2.290*
	(0.528)	(0.530)	(0.782)	(0.761)	(0.769)	(1.197)
$q_{1-5}^{yob} \times YQ2$			-0.173			4.655***
			(1.058)			(1.544)
$q_{1-5}^{yob} \times YQ3$			0.691			2.969
			(1.079)			(2.388)
$q_{1-5}^{yob} \times YQ4$			2.422**			2.709*
			(1.157)			(1.637)
q_{6-10}^{yob}	2.820*	3.077*	2.952	3.411*	3.392*	-1.067
	(1.708)	(1.714)	(2.152)	(2.037)	(2.059)	(2.591)
$q_{6-10}^{yob} \times YQ2$			-2.020			6.917*
			(2.901)			(4.036)
$q_{6-10}^{yob} \times YQ3$			-1.887			3.188
			(3.124)			(5.942)
$q_{6-10}^{yob} \times YQ4$			3.797			5.154
			(3.265)			(4.062)
q_{11-15}^{yob}	2.682	2.695	3.407	2.991	3.469	4.845
	(2.661)	(2.670)	(3.264)	(2.768)	(2.788)	(3.093)
$q_{11-15}^{yob} \times YQ2$			0.128			-8.109
			(4.394)			(5.416)
$q_{11-15}^{yob} \times YQ3$			-1.347			0.410
			(4.885)			(7.566)
$q_{11-15}^{yob} \times YQ4$			4.920			-7.354
			(5.502)			(4.781)
Year of birth	-0.0167***	-0.0147***	-0.0142***	-0.0135***	-0.0126***	-0.0141***
	(0.00331)	(0.00332)	(0.00334)	(0.00485)	(0.00487)	(0.00502)
Retirement age	-0.0012***	-0.0015***	-0.0015***	-0.0015***	-0.0015***	-0.0014***
	(0.000221)	(0.000224)	(0.000229)	(0.000288)	(0.000289)	(0.000297)
Disability p.	0.162***	0.119***	0.116***	0.168***	0.151***	0.141***
	(0.0171)	(0.0180)	(0.0181)	(0.0217)	(0.0225)	(0.0227)
Log-Lik	-219112	-219017	-219007	-96250	-96244	-96233

Notes: coefficients; standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1; region of birth dummy variables not reported (20 regions), YQ*x*=1 if the individual income belongs to quartile *x*, $q^{yob(i)}$ variables are mean-centered; see text for definition of $q^{yob(i)}$ variables, reference group: first income quartile, never received a disability pension

5. Conclusions

In this paper, we obtain insights into the impact of early-life conditions on old-age mortality in Italy. We capture cohorts' life conditions by means of mortality rates at different early-life stages and exploit exogenous variation provided by a series of abrupt mortality events which severely affected specific cohorts. We also test whether lifetime income – approximated by the amount of the pension benefit - is health protective against bad circumstances experienced in early life.

Early-life conditions have a long-lasting effect on males' mortality. Results suggest the existence of a "scarring" effect: males grown in worse times have higher death probabilities at old-ages than those grown in better times. For females, we do not find a significant impact of early-life conditions. The impact on old-age mortality of the historical improvement in early-life conditions experienced by the cohorts in our sample - especially those born after the World War I - is considerable. The death probability of a male born in 1932 alive at age 65 is 15% lower than that of a male born at the beginning of the XX-th century due to improved early-life conditions. Finally, we do not find evidence that income is health protective. On the contrary, we find that especially mortality of richer individuals is affected by circumstances in early life.

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