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**Maurizio Franzini**

# Increasing Inequalities in Longevity among Italian Workers


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# Increasing Inequalities in Longevity among Italian Workers♣

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## Abstract

This article investigates the evolution of inequalities in life expectancy at 65 years old and all-cause mortality by socioeconomic position (SEP) among Italian workers.

Period life tables and negative binomial regression models are employed to estimate longevity inequalities. The empirical assessment is done on two administrative datasets, one covering the entire private sector employees' population in the years 1990–2019, the other covering all private and public workers of Turin for the years 1981–2019.

We find that social inequalities in life expectancy and all-cause mortality have widened among Italian workers during the last decades. Longevity improved for mid- to high SEP workers whereas it hardly changed for workers in the lowest SEP. We show that the estimated life expectancy gaps translate into a loss of pension wealth for the most disadvantaged workers, which is also increasing with time.


**Keywords:** Mortality; Life expectancy; Inequality; Socioeconomic Position; Income; Occupational Class; Population Health.

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**Sommario**

L’articolo analizza l’evoluzione delle disuguaglianze nella speranza di vita a 65 anni e nella mortalità per posizione socioeconomica basata su reddito e classe sociale occupazionale tra i lavoratori italiani. L'analisi utilizza dati amministrativi relativi all’intera popolazione dei dipendenti del settore privato per gli anni 1990-2019 e alla popolazione dei lavoratori del settore privato e pubblico dal censimento della città di Torino (Italia) per gli anni 1981-2019.

L’aspettativa di vita a 65 anni per diversi quantili di reddito e classi sociali occupazionali è stimata tramite la tecnica delle tavole di mortalità. Modelli di regressione binomiale negativa sono usati per stimare come cambia nel tempo l'effetto della posizione socioeconomica sulla mortalità controllando per numerosi fattori demografici e socioeconomici individuali e contestuali.

Le analisi mostrano un aumento delle disuguaglianze sociali nell'aspettativa di vita e nella mortalità per tutte le cause negli ultimi decenni. La longevità è migliorata per i lavoratori con un livello di posizione socioeconomica medio-alto, mentre è rimasta pressoché invariata per i lavoratori con reddito e classe occupazionale bassi. I divari di aspettativa di vita stimati si traducono inoltre in una significativa perdita di ricchezza pensionistica a sfavore dei lavoratori più svantaggiati, anch’essa in aumento con il tempo.

**Parole chiave:**

Mortalità; aspettativa di vita; disuguaglianza; posizione socioeconomica; reddito; classe professionale; salute della popolazione.

## 1. Introduction

Life expectancy around the world is increasing steadily, but the evidence on how uniform this improvement is for different segments of the workforce is still scant. Cross sectional evidence shows that there are relevant longevity inequalities along the socioeconomic position (SEP) dimension. Individuals with lower SEP are more likely to suffer from physical and mental illness, disability and to die at younger ages, resulting in a gap in average life expectancy between different socio-economic groups (WHO 2008).

An ongoing debate is discussing the causality of these associations. Health inequalities can lead to differences in SEP (e.g., Adams et al. 2003; Case and Paxson 2011; Chandra and Vogl 2010), especially in later working ages, when the health status translates into SEP via the ability to work and extend working life (Smith 1999, 2004). On the other way round, several empirical studies using quasi-experimental designs have also identified a causal effect of socioeconomic status on health or mortality (e.g., Lindahl 2005; Lleras-Muney 2005; Lundborg et al. 2016; Spasojevic 2010). There are in fact several ways in which SEP may directly and indirectly influence health and mortality via specific causal mechanisms (Geyer et al. 2006), ranging from material deprivation (Schrijvers et al. 1999), exposure to low quality jobs and harmful working conditions (Ardito et al. 2020), financial strain and emotional distress (Lachowska 2017), social isolation and detrimental health behaviors (e.g. Schrijvers et al. 1999; Dabergott 2021).

Inequalities in health, emerging or being amplified due to socio-economic disadvantages, are not only undesirable for ethical reasons, but they also imply a huge economic and welfare loss for governments, which will spend more on social welfare and health care, receiving lower revenues due to the reduced taxpaying capacity. It has been calculated that if the European population had the health of the 50% most educated, there would be 700,000 fewer deaths per year and 33 million fewer cases of ill health across the European Union (Mackenbach et al., 2011). Longevity inequalities have important consequences also in terms of redistribution and equity of pension systems. About two thirds of OECD countries employ life expectancy (at birth or at retirement) in several pension automatic adjustment mechanisms (De Tavernier and Boulhol, 2021). These include for example linking benefits to life expectancy at retirement through longevity factors in notional defined contribution schemes and linking the statutory retirement age to life expectancy changes registered in the population. By linking pension eligibility or formula to a unique life expectancy, these automatisms introduce a regressive redistribution of pension resources as more advantaged socio-economic and occupational

groups, who in general have a higher life expectancy than the average, receive more than actuarially fair, whereas pensioners who die earlier than the mean will receive less (De Tavernier and Boulhol, 2021; OECD 2018; Ardito & d'Errico 2018; Mazzaferro et al. 2012).

Are current improvements in life expectancies reducing these inequalities, or is the socioeconomic gradient widening? Recent years have witnessed an upsurge in research on this question. Many studies have shown that socioeconomic inequalities in life expectancies among the elders are increasing over time. However, the available evidence comes mainly from North America (see e.g. Currie and Schwandt 2016 and Bosworth 2018 for the US; Baker et al. 2019 for Canada) and from Northern European countries (Mortensen et al. 2015; Tarkiainen et al. 2011; Hann et al. 2020; Zarulli et al. 2012; Bär et al. 2020; Brønnum-Hansen-Baadsgaard, 2012). Only recently evidence have been published on the evolution of mortality inequalities in previously unexplored countries, e.g., Spain (González and Rodriguez-González, 2021), Portugal (Costa and Santana, 2021) and Czech Republic (Bertoli and Grembi, 2021)[6]. These latter studies, although having contributed substantially to the understanding on the evolution of mortality inequalities also in southern and eastern Europe, adopted aggregated area-level SEP or deprivation indicators, possibly underestimating the social gradients because of higher non-differential misclassification. Moreover, mortality rates were adjusted only for age and gender, with the risk of a residual confounding along other dimensions. In the case of Italy, although the socioeconomic mortality differentials have been extensively mapped (Leombruni et al. 2015; d'Errico et al. 2017; Lallo & Raitano, 2018; Petrelli et al. 2018; Alicandro et al. 2018; Petrelli et al. 2019), evidence on the evolution of the gradient in Italy is almost absent, as to the best our knowledge, only this study and that of Ghislandi and Scotti (2022) have explicitly analyzed it.

This study aims at contributing to the existing literature along the limits above mentioned. First, we present novel evidence on the evolution over the last 30 years of socio-economic differentials in mortality inequality and in life expectancy at 65 in Italy by analyzing life expectancy and mortality differentials in different periods. Second, thanks to the availability of very rich administrative data on individual work histories, we can examine very accurate individual-level measures of SEP rather than aggregated indicators at the area level. Finally, we can adjust our estimates with several individual- and job characteristics besides age and gender, to account for compositional differences across time. All these elements allow us to

[6] The works are part of a special issue on the evolution of mortality inequality in 12 OECD countries (Banks et al. 2021).

provide solid evidence on longevity and life expectancy inequality and their evolution, which can contribute to inform the debate around the revision of pension eligibility conditions for more disadvantaged categories.

Results indicate that in Italy the gap in life expectancy at 65 by income and occupational groups has increased over time, sustained by a different rate of growth in life expectancy, which has increased systematically for the advantaged categories, while it has barely improved for the disadvantaged groups. As a result, gaps in residual life expectancy at 65 by SEP have widened by 0.7 years among men and 0.2 among women over the last 30 years. Back of the envelope calculations show how these life expectancy gaps translate into pension entitlements loss for those having lower life expectancy. Our results imply that this penalty rose by about 60% over the analyzed period, resulting in a pension wealth loss among men of about 18,000 euro borne by the poorest 25%. Negative binomial regression analyses and the different sensitivity checks fully confirm the trend, showing that mortality inequalities between income groups are significantly widening, even when we control for several individual demographic and work-related characteristics.

The paper proceeds as follows. The next section describes the measures and statistical methods adopted. Section 3 presents the relevant results and section 4 includes robustness checks. In section 5 results are discussed and in section 6 conclusions are drawn.

## 2. Materials & Methods

### 2.1. Data and Samples

Our primary data come from the administrative archives of the Italian National Institute of Social Security (INPS). INPS data represents the most complete and up-to-date statistical source of information to study mortality differentials between socioeconomic and occupational groups in Italy. It offers detailed information on job spells, welfare benefits and demographic characteristics of all insured workers at individual level. We analyse data of the population of private sector employees using the INPS archives for the years 1990-2019. Furthermore, the analysis is replicated on an independent information source, the Turin Longitudinal Study (TLS), a longitudinal study based on the historical population registers and censuses, built to monitor metropolitan health variations linking social and health careers of individuals and families. Turin is one of the four largest cities for population in Italy. For the two data sources, it is possible to conduct mortality follow-up by linking administrative records with administrative mortality records up to very recent years, i.e., up to 2019. Advantages of using

the TLS are that it allows to test the validity of the results on the whole population of workers, rather than limited to private sector employees only, and to examine changes in differential life expectancy during a longer time span (as the first census dates to 1981).

For the analysis on private workers, the study period was broken down into three 5-year periods: 1990-1994; 1995-1999 and 2000-2004. For each 5-year period, we selected only workers with at least one-month job spell, born in Italy and aged 15-95 at the start of the 5-year period. Individuals were followed until the date of death (recorded at the year-month level) or until the end of the follow-up (equal to 20 years). For example, mortality follow up for workers of the 1990-1994 period went from 1 January 1990 to 31 December 2009 (Table 1). Henceforth the three samples will be referred to as the "1990", "1995" and "2000".

***Table 1 Sampling and mortality follow-up (FU) used to construct the analysis sample.***

| Period | Start FU<br>Start Sampling | End Sampling | End FU | #Persons (millions) | #Person years (millions) |
|---|---|---|---|---|---|
| «1990» | 1st Jan 1990 | 31st Dec 1994 | 31st Dec 2009 | 12.50 | 245.06 |
| «1995» | 1st Jan 1995 | 31st Dec 1999 | 31st Dec 2014 | 12.98 | 255.11 |
| «2000» | 1st Jan 2000 | 31st Dec 2004 | 31st Dec 2019 | 14.22 | 279.65 |

*Source: Authors' elaboration on Italian Social Security data, INPS, years 1990-2019*

For each person in the three periods, all job spells observed during the relative five-year window were piled. Individual career variables such as average weekly wage, prevalent geographical area of work, main sector of activity, main occupational class, were constructed as averages or modes, weighted by the length of the work episode. An indicator of work intensity was defined as the proportion of weeks worked over the period of observation and categorized in low (<20%), mid (20-80%) and high (>80%). To define weekly wage, we took the sum of reported employment inflation-adjusted earnings divided by the total number of weeks worked and constructed an average weekly wage over the 5-year window for each period. The final analysis dataset is described in Table 2.

As can be seen from Table 2, the composition of the three samples has changed over time, reflecting some of the macro-trends characterizing the Italian labour market during these years. We observe an increase in the average age of the samples over time, reflecting the ageing of the working population driven by overall ageing and recent pension reforms that have promoted the tightening of retirement eligibility conditions. We observe a change in the composition of

sectors, with an increase of the services sector at the expense of the manufacturing sector, reflecting a general process of tertiarization of the economy. The most striking compositional change evident in Table 2 regards the female sample, that in the 2000 group increased by 25% with respect to the 1990 group, against a most modest increase experienced by males, coherently with the positive trend of female participation to the labour market. The results on the female subsamples will be interpreted in the light of this compositional change.

***Table 2 Distribution of socioeconomic variables by period and gender, avg. or % of total persons.***

| | **Men** | | | **Women** | | |
|---|---|---|---|---|---|---|
| | **1990** | **1995** | **2000** | **1990** | **1995** | **2000** |
| Age | 34.73 | 36.44 | 37.76 | 30.29 | 31.04 | 32.46 |
| Geographic area of work: | | | | | | |
| North | 0.56 | 0.56 | 0.54 | 0.62 | 0.62 | 0.60 |
| Centre | 0.19 | 0.20 | 0.20 | 0.21 | 0.21 | 0.20 |
| South | 0.25 | 0.24 | 0.26 | 0.18 | 0.18 | 0.20 |
| Weekly Wage (real) | 525.46 | 554.35 | 556.00 | 400.25 | 419.90 | 425.45 |
| Occupational class: | | | | | | |
| Blue-Collar | 0.71 | 0.70 | 0.69 | 0.54 | 0.53 | 0.52 |
| White-Collar | 0.27 | 0.28 | 0.29 | 0.45 | 0.46 | 0.48 |
| Executives | 0.02 | 0.02 | 0.02 | 0.004 | 0.006 | 0.004 |
| Economic sector: | | | | | | |
| Primary & Construction | 0.17 | 0.18 | 0.18 | 0.02 | 0.03 | 0.03 |
| Manufacturing | 0.44 | 0.41 | 0.37 | 0.38 | 0.33 | 0.28 |
| Services | 0.39 | 0.42 | 0.45 | 0.60 | 0.64 | 0.69 |
| Work intensity: | | | | | | |
| Low (<20%) | 0.16 | 0.17 | 0.15 | 0.22 | 0.23 | 0.22 |
| High (>80%) | 0.51 | 0.49 | 0.53 | 0.38 | 0.35 | 0.35 |
| *#Persons* | 7,974,849 | 8,100,832 | 8,589,521 | 4,520,664 | 4,874,688 | 5,634,472 |
| *% Change w.r.t. 1990* | | +2% | +8% | | +8% | +25% |

*Notes: Authors' elaboration on Italian Social Security data, INPS, 1990-2019*

### 2.2. SEP Measures

The two SEP indicators used in this paper are based on income and occupational social class.

As an indicator of income, we used the average weekly wage. Individuals were ranked into quantiles of average weekly wage, which were calculated with cut-off points derived from the final dataset of analysis separately for men and women, and for each time-period.

Using income quantiles as SEP measures offers several advantages as they group individuals into equal shares of the population, helping to address the problem of measurement error in mortality rates for small groups when analysing age classes with low deaths counts, and limiting problems related to changing composition or measures definition over time, by looking at

category bins of similar size from one year to another. The main analysis is based on income quartiles to guarantee higher bin size.

Occupational social class is categorized according to the European Socio-economic Classification (ESeC) into three categories (executives, white-collar and blue-collar), as the original variable in the INPS data does not allows further disaggregation. This indicator was used in previous studies on life expectancy inequality based on INPS and TLS data (Leombruni et al. 2015; d'Errico et al. 2017), hence it allows an easier comparison and replicability of previous results, although the categorization based on occupational class suffers from changing composition and size over time.

### 2.3. Data analysis

To calculate life expectancy, we use life table methodology, as it is a standard tool used in demography and actuarial science to analyze death rates and calculate life expectancies at various age (Smith and Bradshaw, 2006; Ayuso et al. 2021). A life table shows, for each age and year, what the probability is that a person will die before his or her next birthday. There are two different types of life table: cohort and period. A cohort life table is used to represent the overall mortality rates of a certain cohort (year of birth) over the course of their lifetime. The period life table represents mortality rates during a specific period for a certain population. Despite cohort life tables produce more accurate measures of life expectancies, this approach is less frequently used by national governments since it is more data demanding, and subject on many assumptions and uncertainly, since it incorporates the gender specific expected mortality future improvement specific to each birth cohort. In contrast, almost all countries generally adopt period life tables to generate life expectancies and mortality rates that link longevity to pension benefits and eligibility changes (Ayuso et al. 2021). In Italy as well, the life expectancy measure adopted by Ministry of Labor to feed the complex set of pension rules is derived by period life tables rather than cohort life tables (Law 122/2010, art. 12)[7]. This is why we opt to adopt this approach to enhance the policy relevance of our contribution.

In order to create life tables for each of the three periods under analysis, we calculated period specific mortality rates for five-year age classes, sex, income and occupational class, as the ratio between the number of individuals who died in the age interval by the total population-years at risk in that age interval. Subsequently, to reduce fluctuation, we constructed abridged period life tables using 5-year age intervals with a final age interval of 85+ to estimate life expectancy

[7] Available at the following link: https://www.rgs.mef.gov.it/_Documenti/VERSIONE-I/Selezione_normativa/L-/L30-07-2010_122.pdf (last access: 06/06/2023)

and confidence intervals (CIs) using the method described by Chiang (1968), with adjustments to the standard errors as proposed by Silcocks et al (2001), as described by Eayres and Williams (2004). We examined life expectancy at the age of 65 years. The baseline for age covers the age range from 15 to 95, whereas the mortality follow-up is 20 years; consequently, we follow all workers for the same length but no individual in the sample is really observed ageing from age 15 to ages 95..

Using 5-year age intervals (rather than individuals' years of life) is standard in adult mortality analysis as it helps to minimize noise due to measurement error and to avoid zero deaths and person-years <5000 in some of the cells (Toson and Baker, 2003). Our measure of life expectancy inequality at 65 is the difference between life expectancy at 65 between the top and the bottom SEP category. We estimate its standard error and 95% CIs assuming that the difference between two life expectancies is normally distributed, as shown by Eayres and Williams (2004).

Since estimation of the trend in life expectancy differentials might be biased by omitted variables or by labour market compositional changes across years, we complement the life expectancy analysis by running a negative binomial regression model, to estimate mortality rate ratio (MRR) for the lowest-SEP compared to the highest-SEP group, controlling for age (15 categories), time period (3 categories), region of birth (20 categories), macro-region of work (5 categories), economic activity (11 ISIC rev. 4 categories), work intensity (3 categories: low, mid, high), firm size (3 categories: low, mid, high), occupational class (3 categories), income quantiles (4 categories). To do so, we build individual sociodemographic strata defined by the above-mentioned characteristics and count the number of deaths and person-years at risk within each stratum. All the analyses were stratified by sex. Given that the dependent variable (total number of deaths) was over dispersed, i.e. it has variance larger than the mean, negative binomial (NB) is to be preferred to Poisson model, as NB incorporates an additional parameter that accounts for overdispersion making estimates more efficient and less prone to downward bias (Cameron and Trivedi, 1986). NB models the expected number of events (deaths, Y) as function of known covariates $X$ and the time at risk measured in person years ($py$). The estimated regression coefficient indicates by how much the log of the expected count of deaths (technically, the death rate) varies due to a unitary change of $X$, holding constant the remaining vector of sociodemographic characteristics shared by individuals within each cell.

We first test the presence of social inequality in mortality by income quartiles ($IncomeQ_k$) by estimating the following equation separately by sex:

Eq. (1) $\log(Y) = b_0 + \sum_{k=1}^{4} IncomeQ_k + X + log\ (py)$

Where the vector of controls X includes occupational class (3 categories), age (15 categories), period (3 categories), region of birth (20 categories), macro- region of work (5 categories), economic activity (11 ISIC categories), work intensity (3 categories), firm size (3 categories).

To test the presence of a trend in social inequality in mortality, we introduce interaction terms between our main SEP measure ($IncomeQ_k$) and period, to compare the MRR for low-income versus high-income across the 1990, the 1995 and the 2000 samples:

Eq. (2) $\log(Y) = b_0 + \sum_{k=1}^{3} IncomeQ_k * \sum_{i=1990}^{2000} group_i + X + log\ (py)$

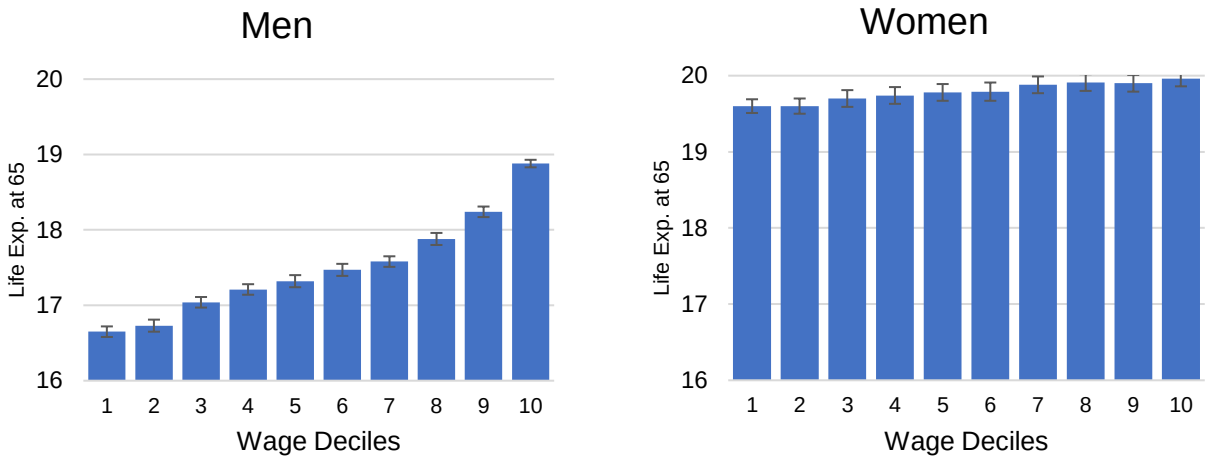
Sensitivity analyses were conducted using a different administrative data source covering the entire working population (private and public sectors), comparing different measures of SEP, extending the periods covered by the analysis and using different sample specifications (see Section: "Robustness").

## 3. Results

### 3.1. Life expectancy inequality

Figure 1 plots life expectancy (and 95% CIs) at age 65 in Italy by income decile, calculated separately for males and females, for the most recent group among the three taken into consideration. Several elements confirm the results on social mortality differentials highlighted by previous literature. A direct socio-economic gradient in life expectancy is present in both sexes, steeper for men than for women. Among male private sector employees belonging to the 2000 group, the gap between the highest and lowest income decile is 2.2 years (95% CI 2.14, 2.32). Among women, the same gap is of 0.36 years (95% CI 0.23, 0.50), roughly corresponding to four months.

***Figure 1 Life expectancy at 65 and 95% CIs by income deciles and sex, 2000 sample.***

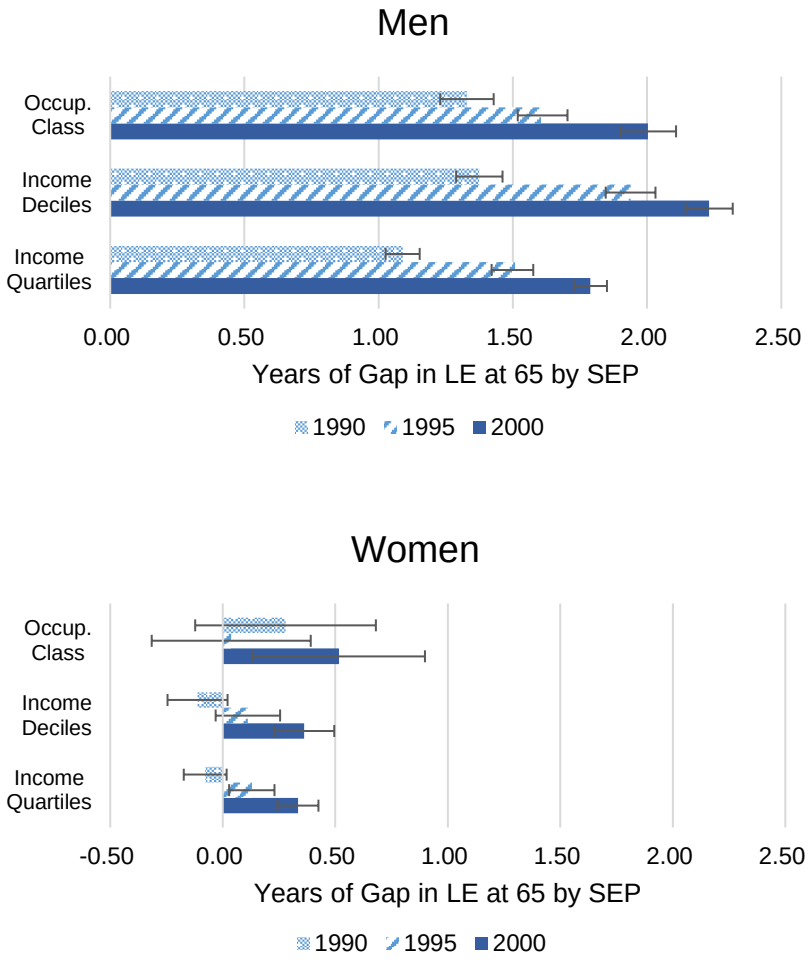



*Source: Authors' elaboration on Italian Social Security data, INPS, 1990-2019. Mortality follow-up is 20 years.*

A similar pattern emerges when using as SEP indicator the occupational social class (Appendix Table 4, panel A). Although occupational social class ranks individuals only in three broader categories (executives, white-collar, blue-collar), the highest SEP group presents a life expectancy at 65 years 2 years higher than blue-collar workers among men (95% CI 1.90, 2.11) and 0.52 among women (95% CI 0.13, 0.90).

In further analyses we restricted the comparison to the change in life expectancy at 65 between the highest and the lowest income quantiles and occupational classes, in the 1990, 1995 and 2000 samples. The results displayed in Figure 2 show that social inequality in life expectancy has increased with time, regardless the indicator adopted and in both sexes. Starting from income, we find that among men the life expectancy disadvantage experienced by people in the least 25% of the income distribution increased over the last 30 years, raising from 1.09 years in 1990 (95% CI 1.03, 1.15) to 1.79 years in 2000 (95% CI 1.73, 1.85). For women, we find insignificant difference in life expectancy at 65 between the richest and the poorest quartile in 1990 (95% CI: -0.17, 0.02) and a small but significant advantage of about 2 and 4 months of life expectancy for the richest quartile in the 1995 and in 2000 group, respectively; these results confirm the presence of a trend of increasing life expectancy inequality also for female private sector employees, although of a lower size.

***Figure 2 Trend of difference in life expectancy at 65 for highest compared to lowest SEP, by sex and different SEP measures.***




*Source: Authors' elaboration on Italian Social Security data, INPS, 1990-2019. Mortality follow-up is 20 years. LE, life expectancy. SEP, socioeconomic position. Income deciles (q10), income deciles. Income (q4), income quartiles.*

We observe very similar results using the prevailing occupational class as indicator of SEP, to those obtained by classifying workers according to income. As shown in Figure 2, among men aged 65 years old, executives could expect to live 1.33 years more than blue collars in 1990 (95% CI 1.23, 1.43) and 2.0 years more in 2000 (95% CI 1.90, 2.11). Although very imprecisely estimated, also among women it is possible to detect a significant difference in life expectancy of about 0.5 years to the advantage of executives in the 2000 group.

In Table 4 in the appendix, we present in detail also the absolute changes in life expectancy for all SEPs, to assess whether – although with a different pace – there have been improvements in life expectancy across all population segments. For both men and women and for all the SEP measures used, a gain in life expectancy from the 1990 to the 2000 group is present only among

highest- a mid-SEP, while a positive trend in life expectancy is almost absent for lowest-SEP workers.

### 3.2. Regression analysis

Table 3 presents mortality rate ratios (MRRs) and 95% CIs from negative binomial regression models for all-cause mortality in the 1990, 1995 and 2000 groups. Results from Equation 1 demonstrate that in all the samples there were significantly more deaths in the lowest-income groups (Q1) and in the lowest occupational class (Blue-collar) controlling for period, age, region of birth and several work characteristics, and mutually adjusting for the two dimensions of SEP. Among men, lowest-income and blue-collar workers exhibit mortality rate 1.57 and 1.47 times higher compared to the highest-income group and to executives, respectively. For women, the mortality rate ratio among the poorest income group individuals was 1.10, compared to the high-SEP group. Hence, both income and occupational class appear to be significant independent predictors of mortality, although the MRR of one SEP dimension decreases when the other is included[8].

[8] In model 1, the inclusion of occupational class lowers the MRR associated to low-income (Q1) by about 12% for men and 3% for women.

***Table 3 Mortality Rate Ratios from Negative Binomial Regression Model, by sex***

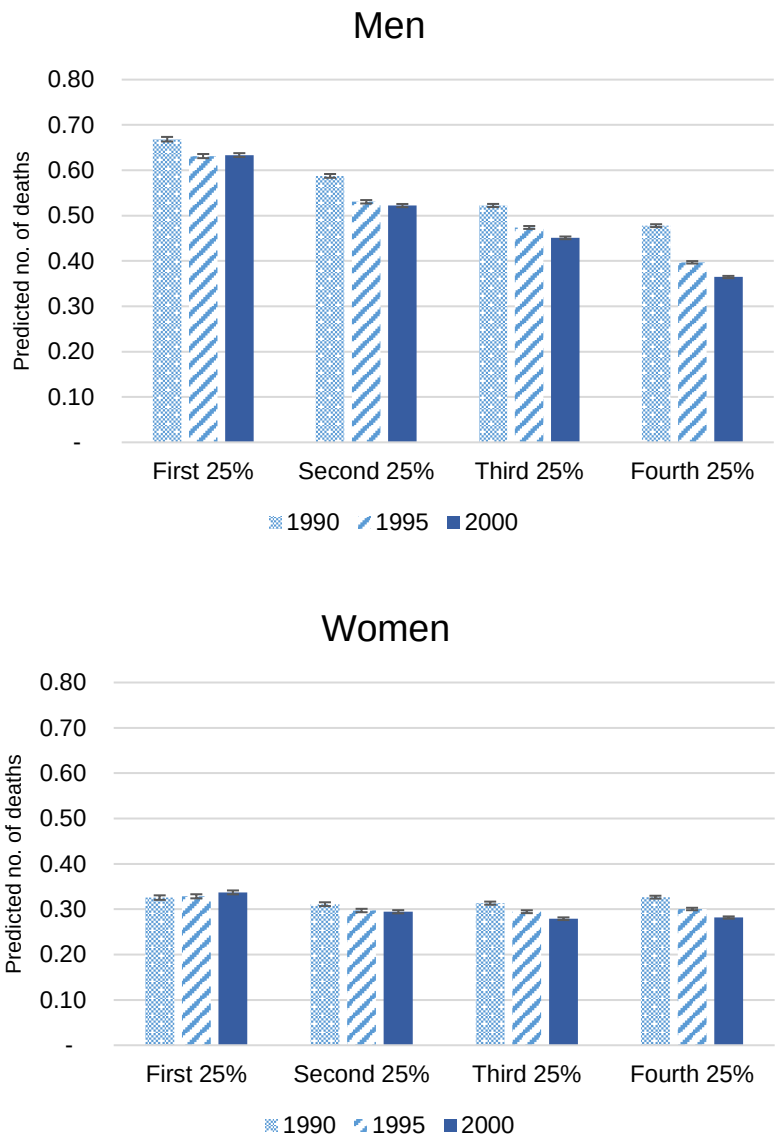
| | Men | | Women | |
|---|---|---|---|---|
| | **Equation 1 MRR [95% CI]** | **Equation 2 MRR [95% CI]** | **Equation 1 MRR [95% CI]** | **Equation 2 MRR [95% CI]** |
| Income quartile | | | | |
| $Q_1$ (first 25%) | 1.573*** | | 1.102*** | |
| | [1.562,1.584] | | [1.088,1.115] | |
| $Q_2$ (second 25%) | 1.330*** | | 0.998 | |
| | [1.321,1.339] | | [0.986,1.010] | |
| $Q_3$ (third 25%) | 1.173*** | | 0.976*** | |
| | [1.166,1.179] | | [0.966,0.987] | |
| $Q_4$ (fourth 25%) | 1 | | 1 | |
| | [1.000,1.000] | | [1.000,1.000] | |
| Interaction Year X $Q_k$ | | | | |
| year1990 X $Q_1$ | | 1.399*** | | 0.997 |
| | | [1.385,1.414] | | [0.977,1.017] |
| year1990 X $Q_2$ | | 1.230*** | | 0.953*** |
| | | [1.217,1.242] | | [0.933,0.972] |
| year1990 X $Q_3$ | | 1.093*** | | 0.960*** |
| | | [1.083,1.103] | | [0.941,0.978] |
| year1995 X $Q_1$ | | 1.590*** | | 1.093*** |
| | | [1.574,1.607] | | [1.072,1.114] |
| year1995 X $Q_2$ | | 1.336*** | | 0.989 |
| | | [1.323,1.350] | | [0.970,1.008] |
| year1995 X $Q_3$ | | 1.194*** | | 0.980** |
| | | [1.182,1.205] | | [0.962,0.998] |
| year2000 X $Q_1$ | | 1.737*** | | 1.197*** |
| | | [1.720,1.755] | | [1.176,1.218] |
| year2000 X $Q_2$ | | 1.432*** | | 1.045*** |
| | | [1.418,1.446] | | [1.027,1.064] |
| year2000 X $Q_3$ | | 1.237*** | | 0.991 |
| | | [1.225,1.248] | | [0.974,1.007] |
| Year | | | | |
| 1990 | 1 | 1 | 1 | 1 |
| | [1.000,1.000] | [1.000,1.000] | [1.000,1.000] | [1.000,1.000] |
| 1995 | 0.892*** | 0.831*** | 0.954*** | 0.920*** |
| | [0.888,0.896] | [0.823,0.838] | [0.945,0.963] | [0.905,0.936] |
| 2000 | 0.859*** | 0.763*** | 0.928*** | 0.863*** |
| | [0.855,0.863] | [0.756,0.770] | [0.920,0.937] | [0.849,0.877] |
| Occupational Class | | | | |
| Blue-collar | 1.471*** | 1.477*** | 1.110*** | 1.110*** |
| | [1.451,1.490] | [1.458,1.495] | [1.062,1.160] | [1.062,1.160] |
| White-collar | 1.200*** | 1.208*** | 1.058*** | 1.061*** |
| | [1.192,1.210] | [1.198,1.215] | [1.021,1.095] | [1.024,1.099] |
| Executives | 1 | 1 | 1 | 1 |
| | [1.000,1.000] | [1.000,1.000] | [1.000,1.000] | [1.000,1.000] |
| Obs. | 3,051,727 | 3,051,727 | 1,084,263 | 1,084,263 |

*Notes: The table displays MRR adjusted for the displayed covariates plus: age (15 categories), region of birth (20 categories), macro- region of work (5 categories), economic activity (11 ISIC categories), work Intensity (3 categories: low, mid, high), firm size (3 categories: low, mid, high). Source: Authors' elaboration on Italian Social Security data, INPS, 1990-2019. Mortality follow-up is 20 years. *p < .05; **p < .01; ***p < .001*

In Equation 2, we investigated the presence of a trend in social inequality in mortality by substituting the income SEP variables with the interaction terms Year*SEP to Eq. 1. Examination of these interactions allowed understanding to what extent the SEP-mortality association differed significantly across the three periods. Results from Equation 2 show a clear social gradient for men and indicate that a low-income man from the 1990 had 40% higher death rate than highest-income group compared with 60% and 74% higher death rate in the 1995 and 2000 periods, respectively. The MRRs for the three interaction terms are significantly above 1.00 and confidence intervals do not overlap, indicating that the SEP-mortality association became stronger in the most recent cohort. For women, there is some evidence that mortality inequality increased with time, too. Whereas in the 1990 sample there was no disadvantage for low-income persons, we find 9% and 20% greater death rate among low-income compared to high-income women in the 1995 and 2000 periods, respectively.

To further explore our key findings, we employed results from the fully adjusted and interacted model to predict the standardized number of death events for each income quartile and group, separately for men and women and we plot them in Figure 3.

***Figure 3 Trend of Standardized Predicted Number of Deaths from Negative Binomial Regression Model (Eq. 2), by sex and SEP measure based on income quartile.***




*Notes: Predicted number of events from Negative Binomial Regression (Equation 2). 95% confidence intervals based on standard errors estimated using delta method. Source: Authors' elaboration on Italian Social Security data, INPS, 1990-2019. Mortality follow-up is 20 years.*

In Figure 3, we can interpret the difference between the 1990 and 2000 bars within each income group as an indication of by how much the predicted death rates reduced with time for that specific group. We can notice that the difference between the 1990 and 2000 bars is the narrowest for the lowest income quartile and that it widens with income, suggesting that high-income groups experienced a more marked reduction in death rate with time. This is striking for women, for whom we find no reduction at all among the poorest 25% while a small but significant drop in the death rate occurred for the top income only. These results confirm what

also the life expectancy analysis was indicating, i.e., that longevity is not improving for all in the same way and for some groups, it hardly improved.

## 4. Robustness

We have performed several checks to test the robustness of our results.

First, we have chosen to replicate the life expectancy analysis on the Turin Longitudinal Study (TLS), a different data source covering the entire population of workers of the city of Turin. TLS is a prospective study of mortality including persons residents in Turin (the fourth largest city of Italy) censused at the 1981, 1991 and 2001 national population census. The study offers large numerousness (on average one million of individuals over the three censuses) and it allows to observe the phenomenon on all workers strata and not only on those working in the private sector. Moreover, TLS has more retrospective depth than the INPS source, since it can observe temporal variations in longevity inequality starting from the 80s, i.e., one decade ahead the time periods covered by the main analysis.

The TLS design consists of a mortality follow-up of 19 years, one year less than in the main analysis. The SEP variables used is occupational social class, based on self-reported data of position in the profession, reclassified into four groups: 1) bourgeoisie (entrepreneurs, executives, highly skilled professionals), 2) middle class (clerical workers), 3) petite bourgeoisie (autonomous artisans and traders, with and without employees), and 4) working class (skilled and unskilled workers). Analysis of mortality by income groups is not possible as the variable is not available in the data. The analyses of TLS data show significant social inequalities in life expectancy at 65 years that magnify with time among men (Figure 4 in the Appendix). The disadvantage in years between the lowest and the highest occupational class was 0.64 in 1981 (95% CI 0.31, 0.96), 1.22 in 1991 (95% CI 1.06, 1.38) and 1.50 in 2001 (95% CI 1.18, 1.80), thus confirming the main findings of the paper based on INPS data. For women, no significant differences in life expectancy emerge in TLS data in all the years, hence undermining the hypothesis that longevity inequality is an emerging phenomenon for female workers, at least when the reference population also encompasses the public sector (Figure 4 in the Appendix).

Secondly, a possible concern is that the SEP measures used in the main analysis, defined as the prevalent SEP looking at 5 years of potential job spells, do not accurately summarize workers' social conditions, which might have changed over time. However, if we construct the SEP

variables using information from 15-years of working career, selecting individuals who were present in all the three samples, the estimated gap in life expectancy at 65 years between lowest-SEP and highest-SEP is quantitatively the same as in the main analysis for both men and women. Results are shown in table 5 in the Appendix.

Third, we have extended the life expectancy analysis to more recent cohorts of workers to increase the longitudinal perspective of the analysis to four time points (rather than three), at the expense of shortening the mortality follow up to 15 years (rather than 20). The results displayed in Figure 6 in the Appendix confirm the overall results, showing that the disadvantage in life expectancy at 65 years for the lowest income groups is raising with time.

Furthermore, negative binomial regression results are also robust to different sample and model specifications. The exclusion from the samples of individuals with very low labour market attachment (i.e., who worked less than 20% of time at baseline); with age<45 at baseline; outliers in income (outside first and 99$^{th}$ percentile) and domestic migrants, leaves the results very similar to the estimates from main analysis, both qualitatively and quantitatively (Table 6 in the Appendix).

Finally, we investigate the presence of heterogeneity in longevity inequality by running a set of separate negative binomial regression fully adjusted models (Eq. 1) on different subpopulations. Figure 5 in the Appendix reports the MRR for lowest- compared to highest-income quartile and shows that mortality inequalities are larger in the North of Italy, in the primary and secondary sectors, and among workers employed in large firms. The pattern is extremely similar among men and women.

## 5. Discussion

This paper revealed an increasing gap in life expectancy at 65 between income quartiles and occupational classes. By following for 20 years three groups of workers employed in Italy during consecutive quinquennia, i.e., 1990-1994, 1995-1999 and 2000-2004, our analysis showed that the gap in life expectancy at 65 between the highest and the lowest income quartile widened with time. In the most recent group of workers, remaining life expectancy at age 65 for lowest-income quartile individuals was approximately 1.8 years less than that for highest income individuals among men, and the corresponding gap was 0.33 years for women. Back of the envelope calculations, using actual individual pension data observed in the 2000s, allow us to translate these life expectancy gaps into pension wealth loss associated to the increasing longevity inequality. Had the lowest-income quartile male workers had the same life

expectancy at 65 than top-quartile ones, they would have accumulated about 11,000 euro more using the longevity differentials estimated on the 1990 group, and 18,000 euro more on the 2000 group (Appendix Table 7). This implies an increase in the loss of pension wealth of about 63% for the poorest income group due to the widening of the social longevity gap. For female workers in the lowest-quartile group, the relative loss of pension wealth is much smaller due to the smaller gap in longevity, amounting to about 2,000 euro for the latest group.

Negative binomial regression analyses confirm the overall results obtained with the life expectancy analysis. Individual income and occupational class are about equally and independently strongly associated with mortality risk for men and women, conditional on age and sectorial composition, individual region of birth and work, firm size, and work intensity. In the fully adjusted model, we find that lowest income quartile males exhibit about 40% higher mortality rate than the top income in the 1990 group and about 74% in the 2000-2004 one. Among women, inequality in mortality is lower than men's but it has also increased with time, being null in the earliest years and significant in the most recent ones.

An important strength of our study was that both SEP and death rate were derived from individual-level administrative data covering the entire population of private sector workers. Thus, we were able to exploit a large sample size and to provide measures that are accurate and robust to reporting bias or loss to follow up. Moreover, by using individual-based SEP measures, we were able to avoid the problems of selective migration (Chetty et al. 2016), potential misclassification of individual exposure and underestimation of the social gradient in mortality generally found with SEP measures based on aggregated geographical areas. A further strength of our approach derives from the fact that ranking individuals along income quantiles substantially limits problems related to changing composition, definitions, shrinking, or growing groups over time, by always looking at category bins of similar size from one year to another. A limitation of our SEP indicators is that they are measured in a relatively short period, looking at only a 5-year window. However, in our sensitivity analysis we showed that extending the window to 15-years gives qualitatively and quantitatively the same results.

Similar to us, several other scholars reported an increasing gap in life expectancy between the highest and the lowest income among men (Blakely et al. 2005; Tarkiainen et al. 2011; Haan, et al 2020, Brønnum-Hansen and Baadsgaard 2012; Kalwij et al 2013; Chetty et al. 2016), a result which has been confirmed also using different SEP measures, such as area-level income (Currie and Schwandt 2016; Bär et al. 2020), occupational class (Martikainen et al. 2007; Bengtsson et al. 2020), and education level (Brønnum-Hansen and Baadsgaard 2012; Cutler, et

al. 2011). However, most of previous findings come from Nordic European countries, New Zealand, and the US, while none have focused on Italy or other South European countries using individual data. Furthermore, several previous studies assessed inequality in mortality by SEP groups adjusting only for age and sex, while we were able to provide estimates of the association of mortality with income quantiles controlling for several individual and contextual factors and, importantly, a second SEP dimension simultaneously, i.e. occupational social class.

Previous literature largely supports the results of women's smaller longevity inequalities than men (for example Mustard and Etches, 2003; Currie and Schwandt, 2016, Bengtsson et al. 2020). In contrast, evidence on the trends of female longevity inequality is less clear cut. Some studies show an increase in the female life expectancy gap between highest and lowest SEP over time, although of a smaller magnitude than what observed among males, coherently with our results (Brønnum-Hansen and Baadsgaard 2012; González & González 2021; Costa & Santana). In contrast, a slightly narrowing gap has been reported for women in New Zealand (Blakely et al 2005), while a Swedish study found that the social gradient emerged earlier for women than for men and then became equally stronger for both sexes (Bengtsson et al. 2020). In our study, the mortality gradient for women employed in the private sector was absent in the nineties and become significant only in the 2000s, years characterized by much higher female employment rates - as reflected in Table 2. This implies that results for women might be driven by changes in the underlying characteristics of women employed. We expect those working in years when participation rates were lower to be more positively selected in terms of their underlying health, compared to those working in times of higher participation rates. This implies that we may expect further increase of longevity inequality in the following years among women at work due to increasing female labour force participation in Italy.

The analysis of variations over time of life expectancy at 65 revealed that for individuals in the bottom income quartile or in the lowest occupational class life expectancy hardly changed over time, while a quicker improvement in life expectancy was evident among top-SEP individuals. The negative binomial models also support this finding, showing that lowest-SEP individuals experienced only a negligible reduction in mortality rate over the observed period compared to the higher-SEP categories. Similarly, evidence of stagnation of life expectancy among the most disadvantaged categories was found in the Finnish general population (1988-2007, LE at 35 by income quintiles, Tarkiainen et al. 2011), in the US (Currie & Schwandt, 2016) and in Germany by Haan et al. (2020).

Unfortunately, we cannot check directly in our data if other dimensions of health changed that may have contributed to this increasing inequality in mortality. However, some studies point to increasing alcohol-related mortality in the working-age population as the main reason for the stagnation of life expectancy in the lowest income quintile (Tarkiainen e al 2011), as well as cardiovascular diseases (Zarulli et al. 2012, Bär et al. 2020, Tarkiainen e al 2011 for men only), cancers (Zarulli et al. 2012) and smoking-related diseases (Currie and Schwandt 2016; Fenelon and Preston, 2012). We can expect that also some of the changes occurred in the working environments over the observed years might have contributed too, such as widening wage inequality (Devicienti et al. 2019), intensification of work (Adăscălițeiz et al. 2022), worsening of working conditions (d'Errico et al. 2022) and a tendency toward late exit from the labour market prolonging exposures to unfavorable working conditions (Carrino et al. 2020)[9]. A better understanding of what is driving the observed raises in longevity inequality at older age remains an important issue for future research.

## 6. Conclusion

Despite the abundance of evidence on the improvements of longevity over the last decades, we still know little about the distribution of these gains over time and across socio-economic groups.

This study shows that in Italy, disadvantaged social status at work, as measured by lower income or lower occupational grade, place individuals at a greater risk of dying and that inequality in mortality has widened over time. Different pace of growth in life expectancy, which have changed in more favorable manner for the most advantaged categories and even stagnated for some subgroups, drive the results.

Since life expectancy at retirement on average translates into the number of years of pension receipt, longevity inequalities have important implications also for pension policy. Today, most of OECD countries, among which Italy, adopt automatic adjusting mechanisms linking pension rules to life expectancy in the whole population. This implies that adjustments apply paradoxically also to subgroups whose life expectancy is stagnating, casting doubts about the fairness and progressivity of such pension systems.

[9] Several studies, adopting counterfactual approaches, suggest that pension reforms delaying retirement may negatively impact the health of workers exposed, before retirement, to worst working conditions, routine manual work, strain jobs (see for example: Blake and Garrouste 2019; Eibich 2015; Ardito et al. 2020).

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## APPENDIX

### Appendix Tables

*Table 4 Trend in Life expectancy at 65 by SEP and sex*

**Panel A: LE at 65 by Occupational Social Class (1=blue-collar, 2=white-collar, 3=executives)**

| Group | SEP | LE at 65 | 95% LB | 95% UB | Abs. Change $LE_{2000}$-$LE_{1990}$ | % Change |
|---|---|---|---|---|---|---|
| | | | **WOMEN** | | | |
| 1990 | 1 | 19,55 | 19,51 | 19,60 | | |
| 1995 | 1 | 19,64 | 19,59 | 19,69 | | |
| 2000 | 1 | 19,74 | 19,70 | 19,78 | 0,19 | 1,0% |
| 1990 | 2 | 19,58 | 19,51 | 19,64 | | |
| 1995 | 2 | 19,76 | 19,69 | 19,83 | | |
| 2000 | 2 | 19,92 | 19,86 | 19,98 | 0,34 | 1,7% |
| 1990 | 3 | 19,83 | 19,43 | 20,23 | | |
| 1995 | 3 | 19,68 | 19,33 | 20,03 | | |
| 2000 | 3 | 20,25 | 19,87 | 20,63 | 0,42 | 2,1% |
| | | | **MEN** | | | |
| 1990 | 1 | 17,01 | 16,98 | 17,04 | | |
| 1995 | 1 | 17,22 | 17,19 | 17,25 | | |
| 2000 | 1 | 17,31 | 17,29 | 17,34 | 0,30 | 1,8% |
| 1990 | 2 | 17,51 | 17,46 | 17,55 | | |
| 1995 | 2 | 18,02 | 17,97 | 18,07 | | |
| 2000 | 2 | 18,40 | 18,36 | 18,45 | 0,89 | 5,1% |
| 1990 | 3 | 18,34 | 18,24 | 18,43 | | |
| 1995 | 3 | 18,82 | 18,73 | 18,91 | | |
| 2000 | 3 | 19,32 | 19,22 | 19,42 | 0,98 | 5,3% |

(Cont.)

(Cont. Table 4)

| Panel B: LE at 65 by Income Quartiles | | | | | | |
|---|---|---|---|---|---|---|
| **Group** | **SEP** | **LE at 65** | **95% LB** | **95% UB** | **Abs. Change $LE_{2000}$-$LE_{1990}$** | **% Change** |
| | | | **WOMEN** | | | |
| 1990 | 1 | 19,57 | 19,51 | 19,63 | | |
| 1995 | 1 | 19,61 | 19,54 | 19,67 | | |
| 2000 | 1 | 19,61 | 19,55 | 19,68 | 0,04 | 0,2% |
| 1990 | 2 | 19,63 | 19,54 | 19,72 | | |
| 1995 | 2 | 19,7 | 19,61 | 19,78 | | |
| 2000 | 2 | 19,76 | 19,69 | 19,82 | 0,13 | 0,7% |
| 1990 | 3 | 19,62 | 19,54 | 19,71 | | |
| 1995 | 3 | 19,69 | 19,6 | 19,77 | | |
| 2000 | 3 | 19,83 | 19,76 | 19,9 | 0,21 | 1,1% |
| 1990 | 4 | 19,49 | 19,42 | 19,56 | | |
| 1995 | 4 | 19,74 | 19,66 | 19,81 | | |
| 2000 | 4 | 19,95 | 19,88 | 20,01 | 0,46 | 2,4% |
| | | | **MEN** | | | |
| 1990 | 1 | 16,63 | 16,58 | 16,67 | | |
| 1995 | 1 | 16,72 | 16,67 | 16,77 | | |
| 2000 | 1 | 16,77 | 16,72 | 16,81 | 0,14 | 0,8% |
| 1990 | 2 | 16,84 | 16,79 | 16,9 | | |
| 1995 | 2 | 17,18 | 17,12 | 17,24 | | |
| 2000 | 2 | 17,22 | 17,17 | 17,27 | 0,38 | 2,3% |
| 1990 | 3 | 17,12 | 17,07 | 17,18 | | |
| 1995 | 3 | 17,41 | 17,36 | 17,47 | | |
| 2000 | 3 | 17,58 | 17,53 | 17,63 | 0,46 | 2,7% |
| 1990 | 4 | 17,72 | 17,67 | 17,76 | | |
| 1995 | 4 | 18,23 | 18,18 | 18,27 | | |
| 2000 | 4 | 18,55 | 18,51 | 18,6 | 0,83 | 4,7% |

*Source: Authors' elaboration on INPS Italian Social Security Data, 1990-2019*

***Table 5 Life expectancy at 65 and gap between low-SEP (bottom income quartile, Q1) and high-SEP (top income quartile, Q4), by sex***

| SEP | LE at 65 (95% CI) | Gap in LE at 65 (95% CI) |
|---|---|---|
| **Panel A: WOMEN** | | |
| Q1 only in 2000 | 19.40 (19.21, 19.59) | 0.54 (0.34, 0.74) |
| Q1 only in 1995, 2000 | 19.52 (19.34, 19.71) | 0.42 (0.23, 0.61) |
| Q1 in 1990, 1995, 2000 | 19.68 (19.58, 19.78) | 0.26 (0.14, 0.38) |
| Q4 in 1990, 1995, 2000 | 19.94 (19.87, 20.01) | Ref. cat. |
| **Panel B: MEN** | | |
| Q1 only in 2000 | 16.65 (16.54, 16.77) | 1.96 (1.84, 2.08) |
| Q1 only in 1995, 2000 | 16.66 (16.55, 16.77) | 1.95 (1.83, 2.07) |
| Q1 in 1990, 1995, 2000 | 16.78 (16.69, 16.86) | 1.83 (1.74, 1.92) |
| Q4 in 1990, 1995, 2000 | 18.61 (18.56, 18.66) | Ref. cat. |

*Source: Authors' elaboration on INPS Italian Social Security Data, 1990-2019. Balanced sample, composed of workers present in all the samples of 1990, 1995 and 2000. N= 4,734,427 individuals, corresponding respectively to the 55% and 46% of men and women in the unbalanced sample.*

***Table 6 Mortality Rate Ratios from Negative Binomial Regression Model, by sex and different subpopulations***

| | Age >45 MRR [95%CI] | At work >20% of time MRR [95%CI] | p1-p99 Income MRR [95%CI] | Domestic Migrants MRR [95%CI] | Domestic Migrants Excluded MRR [95%CI] |
|---|---|---|---|---|---|
| **Panel A: MEN** | | | | | |
| year1990xQ1 | 1.469*** | 1.465*** | 1.419*** | 1.502*** | 1.366*** |
| | [1.453,1.485] | [1.448,1.483] | [1.402,1.435] | [1.474,1.529] | [1.348,1.384] |
| year1995xQ1 | 1.695*** | 1.674*** | 1.589*** | 1.660*** | 1.551*** |
| | [1.677,1.713] | [1.654,1.694] | [1.571,1.607] | [1.631,1.689] | [1.531,1.571] |
| year2000xQ1 | 1.856*** | 1.825*** | 1.735*** | 1.785*** | 1.711*** |
| | [1.837,1.876] | [1.804,1.846] | [1.716,1.754] | [1.755,1.814] | [1.690,1.733] |
| N | 2,534,979 | 2,217,053 | 2,679,481 | 1,851,659 | 1,200,068 |
| **Panel B: WOMEN** | | | | | |
| year1990xQ1 | 1.018 | 1.014 | 1.031*** | 1.041** | 0.985 |
| | [0.996,1.040] | [0.991,1.038] | [1.008,1.055] | [1.003,1.081] | [0.962,1.009] |
| year1995xQ1 | 1.138*** | 1.128*** | 1.110*** | 1.161*** | 1.066*** |
| | [1.115,1.162] | [1.103,1.153] | [1.086,1.134] | [1.120,1.202] | [1.042,1.091] |
| year2000xQ1 | 1.246*** | 1.225*** | 1.202*** | 1.238*** | 1.183*** |
| | [1.223,1.269] | [1.200,1.250] | [1.179,1.225] | [1.199,1.278] | [1.159,1.208] |
| N | 896,169 | 738,968 | 964,319 | 710,869 | 373,394 |

*Notes: The table reports MRR for being in the lowest income quartile (Q1) versus top quartile, interacted for the year dummy (Eq. 2). Models also adjusted for the interactions Q2xYear (3 categories), Q3xYear (3 categories), Year (3 categories), age (15 categories), occupational class (3 categories), region of birth (20 categories), macro-region of work (5 categories), economic activity (11 ISIC categories), work Intensity (3 categories: low, mid, high), firm size (3 categories: low, mid, high). Source: Authors' elaboration on Italian Social Security data, INPS, 1990-2019. Mortality follow-up is 20 years. *p < .05; **p < .01; ***p < .001*

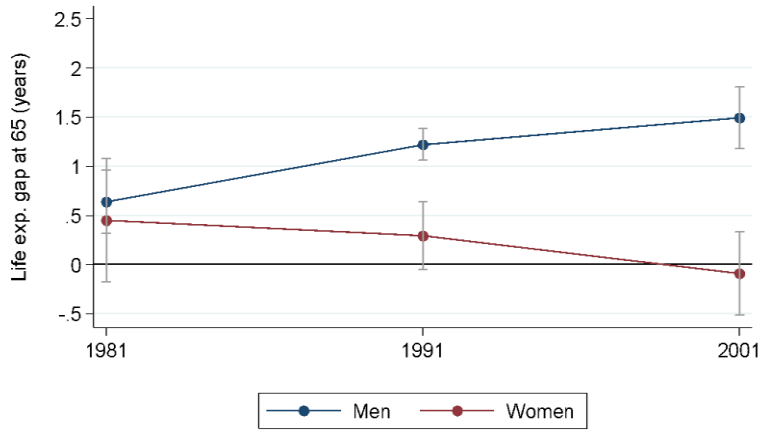
***Table 7 Back of the envelope calculation: loss of pension wealth for low-income workers and period***

| | | LE gap at 65 between Top-Bottom income Q4 | | | Pension loss associated to the LE gap (Annual pension*LE gap) | | |
|---|---|---|---|---|---|---|---|
| Gender | Annual pension in 2000 | 1990 | 1995 | 2000 | 1990 | 1995 | 2000 |
| Men | €10,387 | 1.09* | 1.51* | 1.78* | 11,322 | 15,684 | 18,489 |
| Women | €6,114 | -0.08 | 0.13* | 0.34* | n.s. | 795 | 2,079 |

*Notes: Pension loss for low-income group (lowest income quartile group) is computed as the amount of pension entitlements lost if they had not had lower life expectancy than highest income quartile group. Annual pension computed as the 25° percentile of old age pensions paid by INPS to private sector workers in 2000, expressed at 2021 price levels. Only old-age pensions paid to workers with at least 52 weeks of contributions were selected. LE difference and significance levels are those reported in Figure 2 (* p-value < .05). Source: Authors' elaboration on Italian Social Security data, INPS, 1990-2019.*

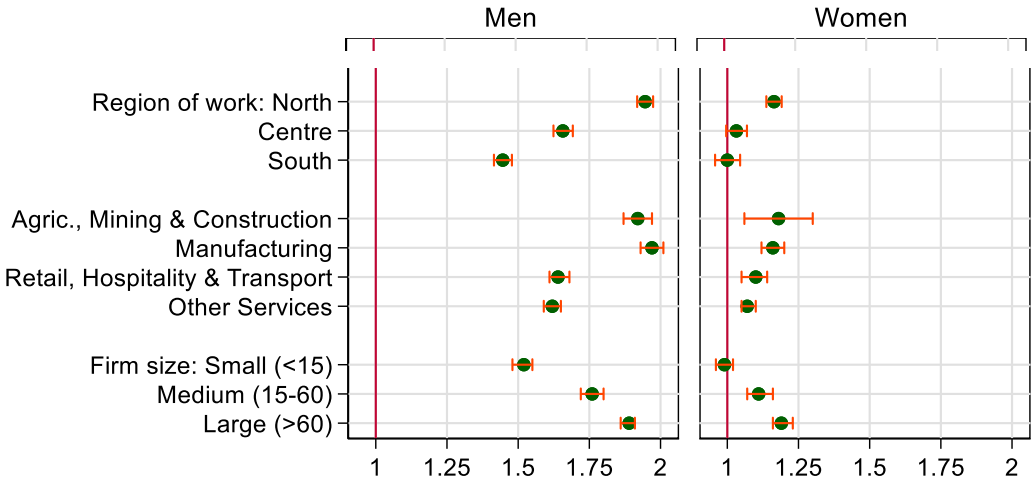
## Appendix Figures

***Figure 4 Trend in difference in life expectancy at 65 for last and first Occupational Class, by sex***



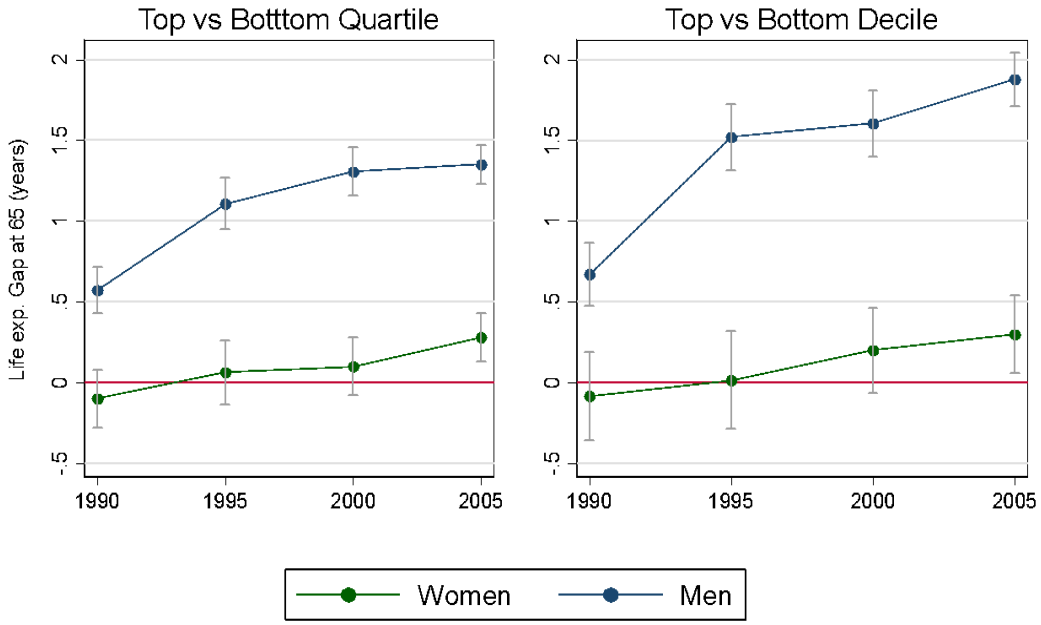

*Source: Authors' elaboration on Turin Census Data, TLS 1981-2019. Mortality follow-up is 19 years.*

***Figure 5 Mortality Rate Ratios from Negative Binomial Regression Model for lowest- compared to highest-income quartile on different subpopulations, by sex (Eq. 1)***




*Notes: The Figure displays MRR and 95% CI from Eq. 1 run separately on different subgroups and by sex. Models are also adjusted for age (15 categories), region of birth (20 categories), macro- region of work (5 categories), economic activity (11 ISIC categories), work Intensity (3 categories: low, mid, high), firm size (3 categories: low, mid, high). Source: Authors' elaboration on Italian Social Security data, INPS, 1990-2019. Mortality follow-up is 20 years.*

***Figure 6 Trend in difference in life expectancy at 65 for highest and lowest income quantile, by sex***




*Source: Authors' elaboration on INPS Italian Social Security Data, 1990-2019. Mortality follow-up is 15 years.*