

Inequality in Life Expectancies across Europe and the US*

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Abstract

We use harmonized household panel data from Europe and the US to provide comparable measurements of education and gender inequalities in total, healthy, and unhealthy life expectancies. Common across countries, the education advantage in total life expectancy is larger for males but the education advantage in (fewer) unhealthy years is larger for females. These results arise because the education advantage in survival is relatively more important for males, while the education advantage in health transitions is relatively more important for females. Across countries, the US stands out with the largest education gradient in healthy life expectancy.

JEL classification: I14, I24, J14, J16

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1 Introduction

The study of economic inequality has attracted a great deal of attention in the last decades. New data and methods have been developed, providing a fairly good picture of the differences in income and wealth inequality across countries, see the recent survey by Alvaredo et al. (2018). Unfortunately, less is known about the range of inequalities in health outcomes and mortality across countries. Health inequality might be more important in terms of welfare—see De Nardi et al. (2017)—, it has first order implications for public policy, and it is likely to become more relevant in the coming years with the aging of population. Furthermore, efforts to understand the determinants of health inequality can benefit from both the differences and the similarities to be found across countries.

This paper puts together harmonized panel data to compare, for the first time, the inequality in total life expectancy (LE), healthy life expectancy (HLE), and unhealthy life expectancy (ULE) between education and gender groups across Continental Europe, England, and the United States. In particular, we use data from the Survey of Health, Ageing and Retirement in Europe (SHARE), the English Longitudinal Study of Ageing (ELSA), and the Health and Retirement Study (HRS) for the time period 2002-2015. For every country-education-gender group, we estimate a three-state (good health, bad health, death) continuous time duration model tailored to match micro data obtained in discrete time at irregular intervals. The model estimates are then used to build country-specific three-state life tables by gender and education starting at age 50, which in turn are used to compute life expectancies for the different demographic groups.

We start by documenting the pervasiveness of several patterns across countries. First, as is well known, individuals who are more educated experience higher LE. Second, the education gradient in LE (the difference between college and non-college individuals) is larger for males than for females: the average across countries is 3.4 years for males and 2.2 years for females. Third, the education gradient is larger in HLE than in total LE, which means that college educated individuals spend less time in bad health despite living longer. And fourth, this education advantage in (fewer) unhealthy years is stronger among females: college educated females (males) spend 1.7 (0.6) fewer years in bad health. Finally, note that these last two results imply that there is a "compression of morbidity" across education groups: the more educated not only live longer but also spend less time in bad health. This is especially pronounced among females: every extra year of life for more educated females (males) is associated with almost 6 (almost 2) fewer months in bad health.

Next, the estimated three-state survival model allows us to provide a decomposition of mortality and morbidity dynamics that helps identify the drivers of the aforementioned interactions between education and gender. For instance, college educated individuals may live longer because they have better health at age 50, because they transit less frequently to states with higher mortality, because they recover more frequently from bad health, or because they survive longer conditional on their health. We find that education confers an advantage in both survival conditional on health and health transitions after age 50, but that their relative strength varies across genders: the education advantage in survival is relatively larger for men and the education advantage in health transitions is relatively larger for women.

Our data also display noteworthy heterogeneity across countries. Overall, taking into account both gender and education dimensions, inequality is largest in Eastern Europe and smallest in Scandinavia. The US stands out as the country with the largest inequality across education groups in HLE, which arises because it has very large education advantage in transition rates between health states. Different from other regions, Mediterranean females with college education have a very small advantage in total LE, but they have the largest advantage in (fewer) unhealthy years.

Finally, it is natural to ask whether spending in the national health system may help mitigate health inequalities. The countries in our sample differ substantially in this dimension: in 2010 public spending in health ranged from 4.6% of GDP in Poland (the lowest in our sample) to 8.8% in Denmark (the highest). We relate this variation with the cross-country heterogeneity in the educational gradients discussed above. We find that higher public spending in health correlates with lower inequality in LE across education groups. Inspecting the mechanisms through our counterfactual life expectancies, we conclude that public spending in health reduces the education advantage in survival rates but that it does not provide significant reductions in the education advantage in health transitions. This has the consequence of increasing life duration for the less educated mainly in the bad health state, which leaves the education gradient in HLE unchanged and increases the education advantage in (fewer) unhealthy years.

We acknowledge potential problems in using longitudinal survey data for survival analysis. While the SHARE, ELSA, and HRS surveys have carefully designed sampling frames and provide exit interviews to separate attrition from death events, there is always the possibility of biases in sample design, in response rates at baseline, or in sample retention. For this reason, we validate our survey data by, first, comparing our estimated life tables by gender to the ones from population data and, second, by running an attrition

analysis based on covariates that predict mortality. The Online Appendix provides all the details on this validation exercise, see Section E. In the paper, we only include results for the US, England, and ten SHARE countries whose life tables match the population tables best with no predictable attrition bias.

Related Literature. The size of the education gradients of LE and HLE are consequential for many economic questions. First, because gradients are big, forecasts of future gains in LE and HLE need to keep track of the changes in education attainment of the underlying population. For instance, according to Case and Deaton (2017), the growth in college attainment explains almost one half of the reduction in age-adjusted mortality in the US between 1910 and 2000. Second, the redistributive power of retirement pensions may be partly eroded by the longer life expectancies of richer individuals (Brown (2002) or Fuster et al. (2003)). Third, the HLE gradient is crucial for effective outcomes in postponing retirement age for a sustainable pay-as-you-go retirement system (see Blundell et al. (2016)). And fourth, the ULE gradient is critical for planning long-term care as well as for predicting related expenses (De Nardi et al. (2016), Bueren (2017), or Braun et al. (2019)).

There are reliable measures of the education gradient in LE in the US using both data from death registers (Meara et al. (2008)) and from household surveys (Pijoan-Mas and Ríos-Rull (2014)). However, this is less so in Europe for two reasons. First, death registers in Europe do not record information on education. Despite this, some recent papers (Avendano et al. (2011) or Mackenbach et al. (2008)) have made progress by linking the death registers with census data to obtain education and gender-specific death rates. However, the advantage of this census-linked mortality data set is somewhat reduced because the lack of older population (above 80 years of age) and imperfect data harmonization across countries (with often non-representative country samples). And second, regarding survey data, only the European Community Household Panel has been used (ECHP, covering years 1994 and 2000 in Majer et al. (2010)). A limitation of the ECHP is its small number of old individuals, the lack of exit interviews for distinguishing attrition from death events, and the absence of data for East European countries. Besides the ECHP, HLE gradients in Europe are documented only by Maki et al. (2013) applying the Sullivan method to census-linked mortality data.

A recent literature has brought important findings about the widening gap of education gradients of LE in the US (Meara et al. (2008) and Montez et al. (2011)), the increase in mortality rates of low-educated white males (Case and Deaton (2017)), or the larger

decline in mortality in richer counties for the 50+ population (Currie and Schwandt (2016)). The increase in the education gradient of mortality has also been documented in several European countries (Mackenbach et al. (2015b) and de Gelder et al. (2017)). Our work does not explore cohort phenomena as the time span of our underlying data is relatively short. However, our methods can be used to analyze time changes when these longitudinal surveys span longer time periods (and can be already applied to the HRS panel).

2 Methodology

In longitudinal surveys, the time span between two consecutive observations that form an individual transition are often not perfectly regular across countries, waves, and individuals. This Section describes the multistate duration model we use to estimate health transitions from microdata obtained at irregular intervals, the estimation method, and the use of the estimated model to compute LE, HLE, and ULE education gradients and decompositions.

Statistical Model. We define 3 states: 0 (dead), 1 (alive-unhealthy), 2 (alive-healthy). The death state is absorbing, while individuals are allowed to transition between healthy and unhealthy as they age.¹ A typical measurement at wave w is (a_w, h_w, x_w) where a_w is age, h_w is health state, and x_w are socio-economic variables. Every individual in our sample is observed in at least two (not necessarily consecutive) waves, so our empirical model is based on the transition probabilities $P(h_{w+1}|a_{w+1}, a_w, h_w, x_w)$, where $w + 1$ is the next wave of observation for each individual, which is at an arbitrary distance $a_{w+1} - a_w$ from wave w .

We interpret the transitions between states as the outcome of independent competing risks in continuous time, but we assume that the underlying hazard rates are constant between birthdays and that one transition at most occurs between any two birthdays and between an observation (wave) and the nearest birthday. To obtain the likelihood contribution for $P(h_{w+1}|a_{w+1}, a_w, h_w, x_w)$, we need to combine probability contributions for: (i) complete 1-year intervals between birthdays, (ii) incomplete intervals between $[a_w, \text{int}(a_w) + 1]$ and $[\text{int}(a_{w+1}), a_{w+1}]$, where $\text{int}(a)$ is a function that returns the integer part of any age a . For instance, an individual who is observed in two waves separated by two and a half years will provide one or two contributions of type (i) and two con-

¹We stress that our model differs from the often used illness-to-death models with irreversible transitions to bad health.

tributions of type (ii). Furthermore, we need to integrate over all possible trajectories between (a_w, h_w) and (a_{w+1}, h_{w+1}) because health is allowed to change every year but it is unobserved between interviews.

For contributions of type (i) we specify two multinomial logits for the transition probabilities from each state $i = 1, 2$ at birthday a to state $j = 0, 1, 2$ at birthday $a + 1$. The covariates are age a itself and potentially variables for socio-economic status x . To ease notation, let's abstract from x . Define $a \in \{50, 51, \dots, \bar{a}\}$, where \bar{a} is the maximum age, and $f_{ij}(a) = \beta_{ij0} + \beta_{ij1}a$. The probability $p_{ij}(a)$ that an individual with health $i \in \{1, 2\}$ at birthday a transits into health $j \in \{0, 1, 2\}$ within a year is

$$p_{ii}(a) = \frac{1}{1 + e^{f_{ik}(a)} + e^{f_{i0}(a)}}, \quad p_{ik}(a) = \frac{e^{f_{ik}(a)}}{1 + e^{f_{ik}(a)} + e^{f_{i0}(a)}}, \quad \text{and } p_{i0}(a) = \frac{e^{f_{i0}(a)}}{1 + e^{f_{ik}(a)} + e^{f_{i0}(a)}},$$

where $k \neq i, 0$. For contributions of type (ii), define $\tilde{p}_{ij}(a, d)$ as the probability that an individual with health $i \in \{1, 2\}$ and age a transits into health $j \in \{0, 1, 2\}$ within a fraction d of a year (before reaching birthday $a + 1$). To compute $\tilde{p}_{ij}(a, d)$, first recover the hazard rates $\lambda_{ij}(a)$ from

$$\begin{aligned} 1 - p_{ik}(a) - p_{i0}(a) &= e^{-(\lambda_{ik}(a) + \lambda_{i0}(a))}, \\ p_{ik}(a) &= \frac{\lambda_{ik}(a)}{\lambda_{ik}(a) + \lambda_{i0}(a)} [1 - e^{-(\lambda_{ik}(a) + \lambda_{i0}(a))}], \end{aligned}$$

so that, for $k \neq i, 0$,

$$\begin{aligned} \tilde{p}_{ii}(a, d) &= e^{-(\lambda_{ik}(a) + \lambda_{i0}(a))d}, \\ \tilde{p}_{ik}(a, d) &= \frac{\lambda_{ik}(a)}{\lambda_{ik}(a) + \lambda_{i0}(a)} [1 - e^{-(\lambda_{ik}(a) + \lambda_{i0}(a))d}], \\ \tilde{p}_{i0}(a, d) &= \frac{\lambda_{i0}(a)}{\lambda_{ik}(a) + \lambda_{i0}(a)} [1 - e^{-(\lambda_{ik}(a) + \lambda_{i0}(a))d}]. \end{aligned}$$

Given the objects $p_{ij}(a)$ and $\tilde{p}_{ij}(a, d)$, we write the likelihood $P(h_{w+1}|a_{w+1}, a_w, h_w)$ of

any given individual transition as

$$P(h_{w+1}|a_{w+1}, a_w, h_w) = \begin{bmatrix} \mathbb{1}_{h_w=1} & \mathbb{1}_{h_w=2} \end{bmatrix} \begin{bmatrix} \tilde{p}_{11}(\text{int}(a_w), d_1) & \tilde{p}_{12}(\text{int}(a_w), d_1) \\ \tilde{p}_{21}(\text{int}(a_w), d_1) & \tilde{p}_{22}(\text{int}(a_w), d_1) \end{bmatrix} \\ \prod_{a=\text{int}(a_w)+1}^{\text{int}(a_{w+1})} \begin{bmatrix} p_{11}(a) & p_{12}(a) \\ p_{21}(a) & p_{22}(a) \end{bmatrix} \\ \begin{bmatrix} \tilde{p}_{11}(\text{int}(a_{w+1}), d_2) & \tilde{p}_{12}(\text{int}(a_{w+1}), d_2) & \tilde{p}_{10}(\text{int}(a_{w+1}), d_2) \\ \tilde{p}_{21}(\text{int}(a_{w+1}), d_2) & \tilde{p}_{22}(\text{int}(a_{w+1}), d_2) & \tilde{p}_{20}(\text{int}(a_{w+1}), d_2) \end{bmatrix} \begin{bmatrix} \mathbb{1}_{h_{w+1}=1} \\ \mathbb{1}_{h_{w+1}=2} \\ \mathbb{1}_{h_{w+1}=0} \end{bmatrix},$$

where $\mathbb{1}$ is an indicator function with $d_1 = \text{int}(a_w) + 1 - a_w$, and $d_2 = a_{w+1} - \text{int}(a_{w+1})$. Similar expressions for the likelihood contribution of a given individual transition can be derived when the information on health h_w or/and h_{w+1} is incomplete (the survival status is known but not whether the individual is healthy or unhealthy).

Finally, in the data we observe N of such individual transitions. Because we consider those N transitions independent, the full likelihood can be written as

$$p(H|\boldsymbol{\beta}) = \prod_{n=1}^N P(h_{w+1}^n | a_{w+1}^n, a_w^n, h_w^n), \quad (1)$$

where H represents all the health transitions in the sample and $\boldsymbol{\beta}$ is the vector of β_{ijl} parameters.

Estimation. In order to reduce the uncertainty of estimated parameters in samples with a small number of observed transitions, we rely on Bayesian techniques and constrain the space of possible $\boldsymbol{\beta}$ to satisfy a set of priors: conditional on surviving, the probability of remaining in good health and the probability of moving from bad to good health both decrease with age; conditional on being in good or bad health, the probability of surviving decreases with age; and the probability of dying is greater when in bad health than in good health. Then, the posterior distribution of $\boldsymbol{\beta}$ is

$$p(\boldsymbol{\beta}|H) \propto p(H|\boldsymbol{\beta}) \cdot p(\boldsymbol{\beta}), \quad (2)$$

In order to sample from the posterior distribution, we use Markov Chain Monte-Carlo (MCMC) methods with a random-walk Metropolis algorithm. See Appendix B for a detailed description of all procedures.

Computing life expectancies. The posterior distribution of parameter estimates delivers a distribution of transition probabilities or multi-state life tables $p_{ij}(a)$ for each country-gender-education sample. Given a multi-state life table we apply standard formulas described in Appendix C to compute total life expectancy (LE) at age 50 and its two components HLE and ULE, i.e. the expected number of years spent in the good and bad health state, where by construction $LE = HLE + ULE$. We report means and standard errors of these measures implied by the posterior distribution.

Decomposing differences in life expectancies. The fact that two individuals from different populations differ in their expected LE is the result of asymmetries in transition probabilities across states or in differences in initial conditions. For example, college educated people could live longer because they are healthy initially, because they tend to transit to states with higher mortality less frequently, or because they survive longer conditional on health. Of course, each of these forces could be at play simultaneously moving the difference in LE in opposite directions.

In order to disentangle the sign and the strength of these forces, we follow Pijoan-Mas and Ríos-Rull (2014) and compute *counterfactual education gradients* in LE, HLE, and ULE as if each gender-education subgroup differed only in its initial health *distribution* at age 50, or in health *transitions* conditional on being alive, or in probabilities of *survival* conditional on health.² That is, for each gender, we first compute the share of individuals in good and bad health state for the whole population and for each education level. Then we estimate the transition probabilities for the whole population and for each education subgroup. Finally, we compute counterfactual life expectancies for each education subgroup if it had: a) the transition probabilities of the whole population but its true education-specific initial condition; b) the survival probabilities conditional on health and the initial condition of the whole population but its true transition probabilities conditional on being alive; c) the transition probabilities conditional on being alive and the initial condition of the whole population but its true survival probabilities conditional on health.

²See Solé-Auró et al. (2015) for a similar decomposition of differences in disease-specific mortality between Europe and the US into contributions from larger prevalence at age 50, larger incidence after age 50, or lower survival.

3 Data

We merge all available waves of the SHARE data plus waves 6 to 11 of HRS and waves 1 to 7 of ELSA in order to use data in a comparable time frame. Every individual-wave *observation* refers to a transition between the given wave and the next available (not necessarily consecutive) one, keeping track of age, the date of interview, health status or the date of death in the case the individual did not survive, as well as gender and education. We exclude from the sample individuals without any transition (with only one observation or in their last wave) and those individuals with below 50 and above 90 years of age. Individuals with missing information on health but known survival status are kept as they also provide valuable likelihood contributions. We focus on education as a measure of socio-economic status because it is a good approximation to lifetime income and on gender because it is an important dimension of inequality across individuals. *College education* is defined as a completed degree at a tertiary educational institution (college or university) with ISCED 1997 codes 5-6, whereas *non-college education* corresponds to all the remaining ISCED codes 0-4.³

A person in a *healthy state* is defined as having no limitations with the following six activities of daily living (ADL): dressing (including putting on shoes and socks), walking across a room, bathing or showering, eating, getting in and out of bed, and using the toilet. Limitations in ADL are widely used in the economic literature (Dwyera and Mitchell (1999) for labor supply, Ameriks et al. (2019) for savings, or Braun et al. (2019) for long-term care insurance).⁴

In Table D.1 in Appendix D for every country we report the number of waves for which the survey was run, the interval of years for which the survey was conducted, and the number of respondents. The HRS and ELSA samples, with around 30,000 and 15,000 respondents, respectively, are much larger than any country sample in SHARE which range between around 2,100 and 6,900 individuals. In order to obtain more precise estimates from larger samples and to better organize our results, we group SHARE countries into four geopolitical regions and estimate the model with the pooled data separately for Western Europe (Austria and France), Eastern Europe (Czechia, Estonia, Poland and Slovenia), Mediterranean (Italy and Spain), and Scandinavia (Denmark and Sweden). Results for individual countries are in presented in Appendix D.

³See Table D.2 in Appendix D for a comparison of the education distribution in our country samples with the one from the population from Eurostat and OECD for the relevant age groups.

⁴The ADL scale was first proposed by Katz et al. (1963). For international comparability, Chan et al. (2012) find good equivalence for the ADL items between the HRS and SHARE, but less so with ELSA. For an educational gradient in the incidence of difficulties in ADLs see Cutler and Lleras-Muney (2010).

4 Results

This Section presents results on life expectancy (LE), healthy life expectancy (HLE), and unhealthy life expectancy (ULE) of four demographic groups: males and females with and without college education. Define the *education gradient* as the difference between a longevity outcome of the college educated minus the same outcome of the non-college educated individuals of the same gender in the same country. We report gradients separately for the four European regions, England, and the US.⁵ We first document patterns which are pervasive across countries, and then we report a few salient differences.

Education gradients in life expectancy. Table 1 shows the average education gradients in LE for men (Panel A, column 1) and women (Panel B, column 1) in each region. As is well-known, the LE education gradient is always positive for both males and females. More importantly, the gradient tends to be larger among males than among females (on average across all countries 3.4 vs. 2.2 years).

Education gradients in healthy and unhealthy life expectancy. Tables 2 and 3 display the education gradients in HLE and ULE. Education gradients in HLE are always positive and larger than the gradients in LE, and more so among females: on average across countries, the education gradients in HLE and LE are 3.9 and 2.2 years for females and 4.0 and 3.4 years for males.⁶ This means that the gradient in ULE is negative (more educated individuals spend fewer years in bad health) and much larger for females (-1.7 years vs. -0.6 years for males).⁷

Decomposition of education gradients. Tables 1-3 also report the decomposition for the LE, HLE, and ULE gradients when the education types differ only in their health distribution at age 50 (columns denoted with a subscript D), only in health transitions conditional on being alive (T), or only in survival conditional on health (S).

⁵The actual levels of the life expectancy for each of the four demographic groups in each country are reported in Table D.3 and D.4 of Appendix D, while the levels pooling across education groups are reported in Table D.5. The education gradients country by country are reported in Table D.6. The corresponding life tables for each group and country are available online at the authors' web pages.

⁶For several definitions of health and socio-economic status, Crimmins and Cambois (2003) also document a greater gradient for HLE than LE. Maki et al. (2013) apply the Sullivan method to census-linked mortality data and calculate educational gradients in healthy life expectancy between the ages of 30 and 79 years with similar results in all countries. Using data from the ECHP, Majer et al. (2010) find larger HLE gradients and a big female advantage in LE (however, they do not have data on Eastern Europe, England and the US).

⁷Note that the identity $LE = HLE + ULE$ carries over to the education gradients.

First, differences in health across education groups at age 50 are inconsequential as most individuals in their fifties are healthy (all estimates in D-columns are very small).

Second, the decomposition of LE gradients in Table 1 shows that college educated individuals tend to live longer because of advantages in both health transitions and survival. Among males, the advantage in survival is very significant and explains around 3/4 of the observed gradients in LE (2.7 out of 3.4 years). In contrast, differences in survival rates between education groups are relatively less important for females as they account for about 3/5 of the LE gradient (1.3 out of 2.2 years). Note that gender differences in the education gradient in LE are entirely driven by the survival gradient (1.3 years females, 2.7 years males) as the education gradient driven by health transitions is very similar between men and women.⁸

Third, the decomposition of the HLE gradients in Table 2 shows that educational advantage in health transitions is as large as the education advantage in survival for males (1.9 and 2.0 years, respectively). Instead, for females the gradient in transition is much larger (2.8 years) while that of survival is small (1.0 year).

And finally, more educated individuals spend fewer years in bad health because of their better health transitions and despite their better survival functions. The total negative education gradient in ULE in Table 3 is composed from a positive gradient due to mortality differences (0.7 and 0.3 for each gender, respectively) and from a large, negative gradient due to health transitions (-1.1 and -1.9 years, respectively). In other words, absent the differences in health transitions, the college-educated individuals would display larger (not smaller) ULE due to their better survival rates in bad health. Finally, note that the total education advantage in unhealthy years is much larger among females (-1.7 vs. -0.6 years) and that is mostly due to their greater advantage in health transitions.

Cross-country differences. Tables 1-3 also show heterogeneity across countries. The most significant findings are as follows. Overall, taking into account both gender and education dimensions, inequality in LE, HLE and ULE is largest in Eastern Europe and smallest in Scandinavia. The US stands out as the country with the largest education gradient in HLE. Our decompositions show that this happens because in the US the advantage conferred by education through transition rates between able and disabled states is very large.⁹ Finally, different from other regions and countries, Mediterranean

⁸With the exception of the Mediterranean countries, where the education advantage in health transitions turns out to be substantially larger among females.

⁹Differences in health transitions in the US also make a large contribution to the male gradient in LE. This result is related to Pijoan-Mas and Ríos-Rull (2014), who find that the gradient in LE in the US is almost exclusively explained by the different health transitions across education groups. Our results

females with college education have a negligible advantage in total LE compared to non-college females. However, they have a strikingly large advantage in (fewer) unhealthy years as they spend 3.5 fewer years in the disabled state (1.6 vs 5.1).

5 Discussion

Compression of morbidity across education groups. The “compression of morbidity” hypothesis states that improvements in longevity over time are associated with decreases in morbidity.¹⁰ But the compression question is also of interest in cross-sectional comparisons of populations that differ by socio-economic status or gender. Our analysis of education gradients and their decompositions shows there is a “compression of morbidity” across education groups (education groups that live longer spend fewer years in bad health), and that this compression occurs because the education advantage in health transitions is sufficiently important relative to the educational advantage in conditional survival rates.

Interactions between education and gender. At the same time, we find that the compression of morbidity across education groups is larger for females: the education advantage in LE is larger for males but the education advantage in (fewer) unhealthy years is larger for females. A careful look at our decompositions shows that we can trace interactions between gender and education to a simple regularity: across countries, the education advantage in conditional survival is relatively more important for males and the education advantage in health transitions is relatively more important for females.

Gender gaps. The interaction between gender and socio-economic status sheds new light on a related question. It is well-known that females have longer LE than males and that the gender gap or female advantage in LE is associated with more not less morbidity, a pattern often described as “women get sicker but men die quicker”.¹¹ A double-difference identity implies that the difference in gender gaps in LE, HLE, or ULE between the non-

differ from theirs in that we still find a role for the education gradient in survival, while they do not. This is likely due to their use of self-rated health measure, a very good predictor of mortality.

¹⁰Fries (1980) was the first to note that the delay in mortality in the US may have been associated to an even larger delay in the onset of disease or disability, thereby reducing the average time spent in poor health. This was in contrast to Gruenberg (1977), who argued that delays in mortality are associated to smaller delays in the onset of disease and hence to increases in ULE. Recent results by Cutler et al. (2013) confirm the compression of morbidity in the US since the 90’s. See Fries et al. (2011) for a survey of this literature.

¹¹Note that this represents a “failure of compression” of sorts, see Lorber and Moore (2002).

college and college populations must equal the difference in education gradients between men and women.¹² Our results in Tables 1-3, show that the education gradient in LE is 1.2 years larger for males than for females, while the (negative) education gradient in ULE is 1.1 years larger for females in absolute value. This means that the gender gap in LE is 1.2 years larger for the non-college population than for the college population while the (negative) gender gap in ULE is 1.1 years larger for the non-college. In other words, the “women get sicker but men die quicker” phenomenon is much more apparent when comparing women and men without college education. Indeed, we find that the “women get sicker but men die quicker” phenomenon is hardly present for men and women with college education.¹³

Cross-country differences in the literature. Our comparison of life expectancies between the US and Europe is novel. Interestingly, the US does not display particularly large inequality in LE across education groups as compared to Europe. This may be surprising given the prevalence of “deaths of despair” across less educated Americans, see Case and Deaton (2017). However, our data refers to the 50+ population, while the evidence on “deaths of despair” refers to middle aged (45-54) individuals. Instead, the US does show the largest levels of inequality in HLE for the 50+, which comes from the fact that the US displays the largest education gradient in health transitions but not in survival. Our results on differences in LE education gradients across European countries are qualitatively in line with other studies, although there are some important differences. In particular, using census-based mortality studies, Mackenbach et al. (2008), Mackenbach et al. (2015a) and Avendano et al. (2011) document that mortality differences are largest in Eastern Europe, intermediate in Nordic countries, and smallest in Mediterranean. Mackenbach (2017), refers to these results as the “Eastern Disaster”, the “Nordic Paradox”, and the “Southern Miracle”. The term “Nordic Paradox” highlights that one would expect the lowest inequality in mortality to arise in countries with low income inequality and strong welfare states, while the term “Southern Miracle” underscores the low inequality in mortality in countries where the welfare state is not so strong. Our

¹²In particular, letting the c and nc subscripts denote college and non-college and m and f denote male and female we can write:

$$(LE_{nc,f} - LE_{nc,m}) - (LE_{c,f} - LE_{c,m}) = (LE_{c,m} - LE_{nc,m}) - (LE_{c,f} - LE_{nc,f})$$

and the same applies for HLE and ULE.

¹³As shown in Table D.7 in Appendix D, the gender gap in ULE among the college educated is very small (0.3 years) and not statistically significant for the average of countries, and it is actually negative in the Mediterranean and Scandinavia.

results do confirm the “Eastern Disaster” but not the “Nordic Paradox” nor the “Southern Miracle”. There are several possible reasons for the discrepancy between our results and the ones obtained in the census-based mortality studies quoted above. First, the age range of the underlying populations are different: 50+ in SHARE, HRS, and ELSA, and 30-79 in the census-based mortality studies. Second, several of the census-based mortality studies are represented only by provinces or cities (Italy and Spain). And third, we have different countries representing Scandinavia (Denmark and Sweden compared to Finland and Norway).

Cross-country differences and public spending in health. The national health system is possibly the main policy tool available to governments to tackle health inequalities. The countries in our study differ substantially in their public spending in health. As a first step towards understanding how health policy affects different socio-economic groups, we explore the co-variation of the education gradients in life expectancies for our 12 countries with the fraction of GDP devoted to public spending in health. In particular, we regress the gradients of our three life expectancies in each country against the ratio of health spending over GDP in 2010. The top panel in Table 4 shows a negative and strong correlation between public spending on health and the gradient in LE for males. The estimated regression coefficient is -0.44, which indicates that the education gradient in LE for males is 1.8 years smaller when the health spending goes from 4.6% of GDP (the lowest in the sample, Poland) to 8.8% (the highest in the sample, Denmark). These 1.8 years represent 53% of the average LE gradient for males. We can further explore the relationship between health spending and the gradients of HLE and ULE separately.¹⁴ We find that all the effect goes to the ULE gradient: countries with more public spending in health tend to have smaller education gradients in LE, the same gradient in HLE, and larger (in absolute value) gradients in ULE.

One way to interpret this result is that public spending in health allows less privileged individuals to live longer but in worse health, which would be consistent with public spending in health improving the survival of low educated individuals in bad health but not improving their health transitions. Consistent with this interpretation, the effect of public spending in health is apparent in the counterfactual gradients LE_s and ULE_s but not in LE_T and ULE_T , see the central and bottom panels in Table 4. Also consistent with this

¹⁴Because of the identity $LE = HLE + ULE$ and the linearity of the covariance operator, the regression coefficients for the gradients of HLE and ULE add up to the regression coefficient for the gradient of LE, which allows for a clean decomposition. In practice, however, this sum is not exact in Table 4 because we use the median and not the mean of the posterior distribution of the gradients in LE, HLE and ULE.

interpretation, recent quasi-experimental evidence from the 2008 Oregon Health Insurance Experiment finds no improvements in measured physical health outcomes among low income individuals in the US who have been randomly granted access to Medicaid, see Baicker et al. (2013).

Finally, for females, public spending on health is also negatively related to the education gradient in LE with a point estimate of -0.12, which is smaller and less precisely estimated than for males. This can be explained as follows: public spending in health seems to diminish the education gradients in LE by reducing the education advantage in survival and, as discussed in Section 4, survival differences across education groups matter less for females.

6 Conclusions

In this paper we study LE, HLE, and ULE across Continental Europe, England, and the United States, focusing on the interaction between education and gender. We develop statistical methods that are suitable for panel data obtained at irregular intervals. A compression of morbidity arises as a return to education. On the one hand, college-educated individuals experience a lower mortality in the bad health state, which prolongs its duration. But on the other hand, education has even greater health-protecting effect as the college-educated visit bad health less often and hence have lower average ULE. Our 3-state model allows for transitions from the unhealthy state back to the healthy state, which is a relevant aspect of the advantage conferred by higher education. A larger compression of morbidity occurs among females because their educational advantage in health transitions is relatively more important. Another implication of this is that college females live longer than college males with hardly any increase in unhealthy years, in partial rebuttal of the well-known “women are sicker but men die quicker” phenomenon. Beyond these common patterns there is heterogeneity across countries and further research may want to exploit these phenomena to better understand the underlying causes of health inequality. Cross-country correlations suggest that public spending in health diminishes the education gradients in LE by reducing the education advantage in survival rates when in bad health. Consequently, more public spending in health does not change the education advantage in HLE while it might increase the education advantage in ULE.

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Tables

TABLE 1: Decomposition of LE education gradients

	A. Males				B. Females			
	LE	LE _D	LE _T	LE _S	LE	LE _D	LE _T	LE _S
Western Europe	3.9 (0.7)	0.0 (0.0)	0.8 (0.2)	3.2 (0.7)	1.7 (0.6)	0.0 (0.0)	0.6 (0.2)	1.1 (0.7)
Eastern Europe	4.0 (0.8)	0.1 (0.0)	0.7 (0.2)	3.3 (0.8)	3.9 (0.6)	0.0 (0.0)	0.8 (0.2)	3.0 (0.6)
Mediterranean	3.0 (1.0)	0.0 (0.0)	0.7 (0.3)	2.6 (1.0)	0.7 (1.1)	0.0 (0.0)	1.4 (0.3)	-1.6 (1.2)
Scandinavia	2.1 (0.7)	0.0 (0.0)	0.7 (0.2)	1.6 (0.7)	2.3 (0.6)	0.0 (0.0)	0.8 (0.2)	1.5 (0.7)
England	3.4 (0.6)	0.1 (0.0)	1.1 (0.2)	2.5 (0.6)	1.2 (0.6)	0.0 (0.0)	0.5 (0.1)	0.6 (0.6)
US	3.6 (0.4)	0.1 (0.0)	1.4 (0.1)	1.9 (0.5)	3.2 (0.4)	0.1 (0.0)	1.3 (0.1)	2.0 (0.4)
Average	3.4 (0.4)	0.1 (0.0)	0.8 (0.1)	2.7 (0.4)	2.2 (0.4)	0.0 (0.0)	0.8 (0.1)	1.3 (0.4)

Notes: Median (standard deviation) of the life expectancy distribution from the posterior distribution of the estimated β parameters. LE: education gradient in life expectancy at age 50. Counterfactual education gradients when education types differ only in: health distribution at age 50 (LE_D), health transition conditional on being alive (LE_T), probability of survival (LE_S).

TABLE 2: Decomposition of HLE education gradients

	A. Males				B. Females			
	HLE	HLE _D	HLE _T	HLE _S	HLE	HLE _D	HLE _T	HLE _S
Western Europe	4.6 (0.7)	0.1 (0.0)	2.1 (0.5)	2.5 (0.6)	3.0 (0.7)	0.1 (0.0)	2.2 (0.5)	0.7 (0.5)
Eastern Europe	4.3 (0.7)	0.2 (0.0)	1.7 (0.4)	2.5 (0.6)	5.1 (0.6)	0.1 (0.0)	2.7 (0.4)	2.1 (0.4)
Mediterranean	3.4 (1.0)	0.0 (0.0)	1.6 (0.6)	1.8 (0.7)	4.1 (1.1)	0.0 (0.0)	4.6 (0.7)	-0.7 (0.7)
Scandinavia	2.9 (0.7)	0.1 (0.0)	1.8 (0.4)	1.1 (0.6)	3.7 (0.7)	0.1 (0.0)	2.4 (0.4)	1.1 (0.5)
England	4.7 (0.6)	0.3 (0.0)	2.9 (0.4)	1.6 (0.4)	3.0 (0.6)	0.2 (0.0)	2.4 (0.4)	0.3 (0.4)
US	5.3 (0.4)	0.3 (0.0)	3.2 (0.3)	1.5 (0.4)	5.1 (0.4)	0.4 (0.0)	3.5 (0.3)	1.3 (0.3)
Average	4.0 (0.3)	0.2 (0.0)	1.9 (0.2)	2.0 (0.3)	3.9 (0.4)	0.1 (0.0)	2.8 (0.3)	1.0 (0.2)

Notes: Median (standard deviation) of the life expectancy distribution from the posterior distribution of the estimated β parameters. HLE: education gradient in healthy life expectancy at age 50. Counterfactual education gradients when education types differ only in: health distribution at age 50 (HLE_D), health transition conditional on being alive (HLE_T), probability of survival (HLE_S).

TABLE 3: Decomposition of ULE education gradients

	A. Males				B. Females			
	ULE	ULE _D	ULE _T	ULE _S	ULE	ULE _D	ULE _T	ULE _S
Western Europe	-0.8 (0.4)	-0.1 (0.0)	-1.3 (0.3)	0.7 (0.2)	-1.3 (0.4)	-0.1 (0.0)	-1.6 (0.4)	0.4 (0.3)
Eastern Europe	-0.3 (0.3)	-0.1 (0.0)	-1.0 (0.2)	0.9 (0.2)	-1.2 (0.4)	-0.1 (0.0)	-1.8 (0.3)	0.9 (0.3)
Mediterranean	-0.4 (0.5)	-0.0 (0.0)	-0.9 (0.3)	0.7 (0.3)	-3.5 (0.4)	-0.0 (0.0)	-3.2 (0.5)	-0.9 (0.5)
Scandinavia	-0.8 (0.3)	-0.1 (0.0)	-1.1 (0.3)	0.5 (0.2)	-1.4 (0.3)	-0.1 (0.0)	-1.6 (0.3)	0.4 (0.2)
England	-1.3 (0.3)	-0.2 (0.0)	-1.8 (0.3)	0.8 (0.3)	-1.8 (0.4)	-0.2 (0.0)	-1.9 (0.3)	0.3 (0.3)
US	-1.7 (0.2)	-0.2 (0.0)	-1.8 (0.1)	0.4 (0.1)	-2.0 (0.2)	-0.3 (0.0)	-2.2 (0.2)	0.7 (0.2)
Average	-0.6 (0.2)	-0.1 (0.0)	-1.1 (0.1)	0.7 (0.1)	-1.7 (0.2)	-0.1 (0.0)	-1.9 (0.2)	0.3 (0.2)

Notes: Median (standard deviation) of the life expectancy distribution from the posterior distribution of the estimated β parameters. ULE: education gradient in unhealthy life expectancy at age 50. Counterfactual education gradients when education types differ only in: health distribution at age 50 (ULE_D), health transition conditional on being alive (ULE_T), probability of survival (ULE_S).

TABLE 4: Correlations between education gradients and public spending in health

A. Males			B. Females		
LE	HLE	ULE	LE	HLE	ULE
-0.44 (0.19)	-0.08 (0.23)	-0.35 (0.14)	-0.12 (0.35)	0.09 (0.34)	-0.18 (0.25)
LE _T	HLE _T	ULE _T	LE _T	HLE _T	ULE _T
0.13 (0.09)	0.34 (0.19)	-0.20 (0.11)	0.12 (0.10)	0.26 (0.28)	-0.12 (0.19)
LE _S	HLE _S	ULE _S	LE _S	HLE _S	ULE _S
-0.55 (0.21)	-0.41 (0.16)	-0.13 (0.06)	-0.23 (0.39)	-0.21 (0.21)	-0.01 (0.18)

Notes: Each entry reports the regression coefficient of the corresponding life expectancy gradient on health spending over GDP. The top block refers to the actual gradients, the middle block refers to the counterfactual gradients in which education groups only differ in transitions (subscript T), and the bottom block refers to the counterfactual gradients in which education groups only differ in conditional survival (subscript S). Standard errors in parenthesis.