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The Cross-sectional Average Inequality in Lifespan (CAL†): A Lifespan Variation Measure That Reflects the Mortality Histories of Cohorts

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ABSTRACT Lifespan variation is a key metric of mortality that describes both individual uncertainty about the length of life and heterogeneity in population health. We propose a novel and timely lifespan variation measure, which we call the cross-sectional average inequality in lifespan, or CAL†. This new index provides an alternative perspective on the analysis of lifespan inequality by combining the mortality histories of all cohorts present in a cross-sectional approach. We demonstrate how differences in the CAL† measure can be decomposed between populations by age and cohort to explore the compression or expansion of mortality in a cohort perspective. We apply these new methods using data from 10 low-mortality countries or regions from 1879 to 2013. CAL† reveals greater uncertainty in the timing of death than the period life table–based indices of variation indicate. Also, country rankings of lifespan inequality vary considerably between period and cross-sectional measures. These differences raise intriguing questions as to which temporal dimension is the most relevant to individuals when considering the uncertainty in the timing of death in planning their life courses.

KEYWORDS Lifespan variation • Cohort mortality • CAL • Decomposition • Formal demography

Introduction

Population health has traditionally been judged by mean levels, including life expectancies and age-standardized death rates. Increasingly, however, more attention is being paid to the variability in the age at death, also known as lifespan variation. At the population level, lifespan variation quantifies the heterogeneity of survival in the population; at the individual level, it is a metric of lifetime uncertainty. If monitoring life expectancy indicates the average progress in increasing longevity, lifespan variation highlights the equality of mortality decline across ages (van Raalte et al. 2018).

To date, most empirical studies of long-term trends in lifespan variation have been based on period life tables (Alvarez et al. 2019; Colchero et al. 2016; Németh 2017; Permanyer and Scholl 2019; Shkolnikov et al. 2003; Smits and Monden 2009; van

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Raalte and Caswell 2013; van Raalte et al. 2018; Vaupel and Canudas-Romo 2003; Vaupel et al. 2011). By contrast, trends in cohort lifespan variation have rarely been examined, largely because the most recent completed cohorts were born at a time when infant mortality was still high. Thus, lifespan variation for these cohorts continues to far exceed the most recently observed period levels (Engelman et al. 2010; Wilmoth and Horiuchi 1999). Moreover, monitoring the lifespan variation of extinct cohorts is unlikely to provide useful insights for making macro-level policy decisions or micro-level life course decisions.

A key unanswered question is the extent to which individuals internalize their surrounding mortality conditions, and how this information is used in decision-making processes. Imagine two populations with identical period age–specific mortality: one who experienced a rapid decline in mortality in the past 50 years, and another who experienced only a moderate decline in mortality over a longer time span. Individuals from the two populations would have been exposed to different levels of mortality over their lifetimes. This raises intriguing questions about whether people base their subjective survival expectations (i.e., the most likely age at death) and the level of survival uncertainty (i.e., the uncertainty surrounding that age) on current mortality rates, in line with changing period-based indicators, or whether they are more influenced by the health and survival trajectories of their contemporaries from similar birth cohorts, such as school friends, spouses, and siblings. Alternatively, it is possible that people adjust their survival expectations on the basis of the survival histories of a broader mixture of family members, colleagues, and neighbors who may have died some years in the past, and who were from mixed birth cohorts. If the last scenario is closest to reality, an approach to monitoring mortality that mixes periods and cohorts could prove useful.

For this reason, we introduce to the demographic literature a new lifespan variation measure based on the historical mortality experiences of all birth cohorts present at a given time. The measure builds on the cross-sectional average length of life (CAL) approach (Brouard 1986). Because CAL averages over the past survival of all cohorts present at a given time (Guillot 2003; Riffe and Brouard 2018), this measure can be especially informative when mortality conditions are changing. The same logic that led to the development of CAL as an indicator of the average length of life of cohorts can be used to assess the variability in length of life. To this end, we have developed CAL†, a new lifespan variation measure that has mathematical properties similar to those of the life disparity ($e^{\dagger}$) measure (Vaupel and Canudas-Romo 2003). CAL† can be interpreted as representing the variation in mortality over ages of an average cohort present at a given time by combining the cohort's mortality experiences in a cross-sectional approach.

In this study, we present empirical data used in the illustration of the CAL† measure and its formulation. We examine the evolution of CAL† in populations with long mortality series and compare these trends with the developments of period and cohort $e^{\dagger}$. We compare the development of the negative correlation of CAL† and CAL with that of $e^{\dagger}$ and life expectancy at birth in period and cohort perspectives. Finally, we calculate the gap in CAL† between populations and demonstrate how it can be decomposed by age and cohort.

Data and Methods

Our data set covers all countries and regions from the Human Mortality Database (HMD) (2019) with complete mortality data from 1879 to 2013, including Denmark, England and Wales, Finland, France, Italy, the Netherlands, Norway, Scotland, Sweden, and Switzerland. We calculate from these data the cohort death probabilities for all of the age and cohort combinations using deaths by Lexis triangles and population counts, and follow the HMD protocol to construct cohort life tables (Wilmoth et al. 2017).

Because CAL† and life disparity ($e^\dagger$) have similar mathematical properties, we start by describing $e^\dagger$. Later, we introduce CAL† and the age and cohort decomposition of the difference between two CAL†.

Life disparity is a lifespan variation measure derived from a life table. At time t, $e^\dagger(t)$ is the average of the remaining life expectancy at age x given by $e(x)$, weighted by the age-at-death distribution, $d(x)$ (Vaupel and Canudas-Romo 2003). However, $e^\dagger(t)$ not only reflects the life table variation in the age at death (Nusselder and Mackenbach 1996; Vaupel et al. 2011), it also quantifies the average number of years lost at the time of death.

Life disparity is closely related to a life table quantity that is now usually referred to as the life table entropy (Keyfitz 1977). This quantity was first derived by Leser (1955) to calculate the effect of a proportional change in age-specific mortality on life expectancy, which he called the elasticity of life expectancy. Keyfitz and Golini (1975) noted the relationship of this quantity to classical entropy, as well as to changes in the rectangularization of the survival curve. Mitra (1978) pointed out that the life table entropy could also be interpreted as a weighted average of the product of life expectancy and the force of mortality. Almost a decade later, Vaupel (1986) and Goldman and Lord (1986) demonstrated that life table entropy could be expressed as the ratio given by life disparity and life expectancy at birth: $e^\dagger/e_0$. Following this relation and defining the radix of the life table as equal to 1 ($\ell(0,t)=1$), life disparity can be written as

$$e_i^\dagger(t) = -\int_0^\omega \ell_i(x,t)\ln\left[\ell_i(x,t)\right]dx, \tag{1}$$

where $\ell_i(x,t)$ is the survival function at age x and time t; i is defined as p or c if the survival function comes from a period or a cohort life table, respectively; and ω is the oldest age attained in the population or by the cohort born at time t. More details of this equation can be found in the online supplemental material (see OSM-1).

Together with life expectancy, indices of lifespan variation such as $e^\dagger$ are measures calculated from one life table for one period or for one cohort. To also include the mortality histories of all cohorts present at a given time, Brouard (1986) introduced the cross-sectional average length of life (CAL). This measure was further developed by Guillot (2003). CAL(t) is a summary mortality measure that takes into account the mortality conditions of all cohorts present at a given time t, and it is calculated as

$$\mathrm{CAL}(t) = \int_0^\omega \ell_c(x,t-x)dx, \tag{2}$$

where t is the time period at which the measure is calculated, and $\ell_c(x,t-x)$ is the cohort life table survival function at age x for the cohort born in year $t-x$. CAL is comparable to life expectancy, as it describes the length of life of a population as an average of all cohort-specific survivals. A comparable measure of lifespan variation that includes the historical mortality information of all cohorts present at a given time remains to be defined. To fill this gap, we propose a measure analogous to life disparity in Eq. (1), defined as

$$\mathrm{CAL}^{\dagger}(t) = -\int_0^{\omega} \ell_c(x,t-x) \ln\left[\ell_c(x,t-x)\right] dx. \tag{3}$$

To further justify the use of CAL† as a measure of variation in age at death, Table A1 in the online supplemental material (see OSM-2) shows the mathematical similarities between life expectancy and life disparity, and between the two measures of CAL and CAL†, including cohort information. Mathematically, CAL† cannot be expressed as the average life years lost in the way that is possible for $e^{\dagger}$, because the remaining life years of some cohorts are unknown. The online supplemental material (see OSM-3) presents CAL†'s mathematical impediments to interpret this novel measure as the average numbers of years lost due to death. Nevertheless, by averaging cohort survival data, it can be thought of as broadly representing the average exposure to lifespan variation across cohorts in a cross-sectional approach. This interpretation is further supported by a strong correlation between CAL† and the standard deviation in lifespan calculated over the same data set (see Figure A1 in OSM-4). Finally, the online supplemental material (OSM-4) shows in detail the calculations that underlie CAL†.

Decomposition of the Difference Between Two CAL†

There is strong interest, especially among demographers, in disentangling changes and differences in aggregate measures when comparing two populations or a single population over time. We introduce the age and cohort decomposition of the difference between two CAL†, which allows us to make such comparisons. Decompositions can be performed by any underlying parameter such as age, cause of death, and cohort. Here, our aim is to decompose a gap between two CAL† into age and cohort contributions. Since we are dealing with cross-sectional measures, traditional age decomposition methods for life expectancy (Andreev et al. 2002; Arriaga 1984; Vaupel and Canudas-Romo 2003) and variation in the age at death (Aburto et al. 2019; Gillespie et al. 2014; Shkolnikov et al. 2003; Shkolnikov et al. 2011; Zhang and Vaupel 2009) have to be extended to include age and cohort components (Canudas-Romo and Guillot 2015; Nepomuceno and Canudas-Romo 2019).

To assess the differences between two CAL†, we examine the derivatives of CAL† with respect to the variable of interest (e.g., a comparison between two time periods or between two populations) as

$$\dot{\mathrm{CAL}}^{\dagger}(t) = -\int_0^{\omega} \ell_c(x,t-x)\left(\ln\left[\ell_c(x,t-x)\right]+1\right)\sum_{i=0}^{x-1} \frac{{}_1\dot{p}_i(t-x)}{{}_1p_i(t-x)}\, dx, \tag{4}$$

where the notation of a dot on top of a variable denotes the derivative of a function with respect to the variable of interest; this notation has been shown to be useful in developing succinct equations (Vaupel and Canudas-Romo 2003). We denote as ${}_1p_i(t-x)$ the probability of surviving from age i to $i+1$ for the cohort reaching age x at time t. By expressing the cohort survival function as the product of single age probabilities of surviving from age zero to age x—that is, $\ell_c(x,t-x) = {}_1p_0(t-x)\ {}_1p_1(t-x)\ldots\ {}_1p_{x-1}(t-x)$—we can separate the derivative of cohort survivals into age contributions (Canudas-Romo and Guillot 2015). The relative derivatives with respect to the age-specific cohort survival probabilities—denoted as $\frac{{}_1\dot{p}_i(t-x)}{{}_1p_i(t-x)}$ in Eq. (4)—correspond to the contribution of age i, for the cohort reaching age x at time t to the overall change in $\mathrm{CAL}^{\dagger}(t)$. More details of how Eq. (4) was derived are presented in the online supplemental material (see OSM-5).

Equation (4) includes a component of change—the relative derivatives of a single age cohort probability of surviving—and two weighting functions—the component of CAL and the component of $\mathrm{CAL}^{\dagger}$ at time t, or the cohort survival function (see Eq. (2)) and the cohort disparity component (see Eq. (3)), respectively (see Table A1 in OSM-2 for more details). Thus, although taking the age-specific survival component of change into account is vital to understand the dynamic of this measure of variability, the achieved levels of survival and variability must also be included in the comparison.

Since we used empirical data by single calendar year and single age, we discretized Eq. (3) and Eq. (4) in our illustration of $\mathrm{CAL}^{\dagger}$. The age and the cohort decomposition of the $\mathrm{CAL}^{\dagger}$ differences were performed at time t, between each country/region and the average population representing the average mortality of the 10 selected countries/regions, with each population being given the same weight in the average. Detailed results for all populations are available in an interactive web app at https://caldagger.shinyapps.io/CALdagger/. R code to reproduce the analysis can be obtained from https://osf.io/n3945/.

Results

Figure 1 presents the trends of period ($e_p^{\dagger}$) and cohort ($e_c^{\dagger}$) life disparity, and of $\mathrm{CAL}^{\dagger}$ by sex. The six panels show declining trends, albeit with different levels. Variation in the age at death transitions from levels as high as 25 years for period and cohort $e^{\dagger}$ in the late nineteenth century to levels as low as 9–12 years for $e_p^{\dagger}$ and $\mathrm{CAL}^{\dagger}$ in the twenty-first century, with $\mathrm{CAL}^{\dagger}$ being somewhat higher than $e_p^{\dagger}$. The trend of $\mathrm{CAL}^{\dagger}$ is smoother than the trends of $e_p^{\dagger}$ and $e_c^{\dagger}$ because it is less affected by period fluctuations. For instance, Figure 1 shows two peaks in the period life disparity: one around 1918 due to World War I and the Spanish Flu and another around 1945 as a result of World War II. These peaks reflect the immediate impact of period mortality shocks on the $e_p^{\dagger}$.

While they differed in magnitude, all three lifespan variation measures ($e_p^{\dagger}$, $e_c^{\dagger}$, and $\mathrm{CAL}^{\dagger}$) declined as longevity increased ($e_{0,p}$, $e_{0,c}$, and CAL), as presented in Figure 2. This negative correlation suggests that lower levels of lifespan variation are consistent with higher levels of longevity. Unlike in Figure 1, the associations of the trends

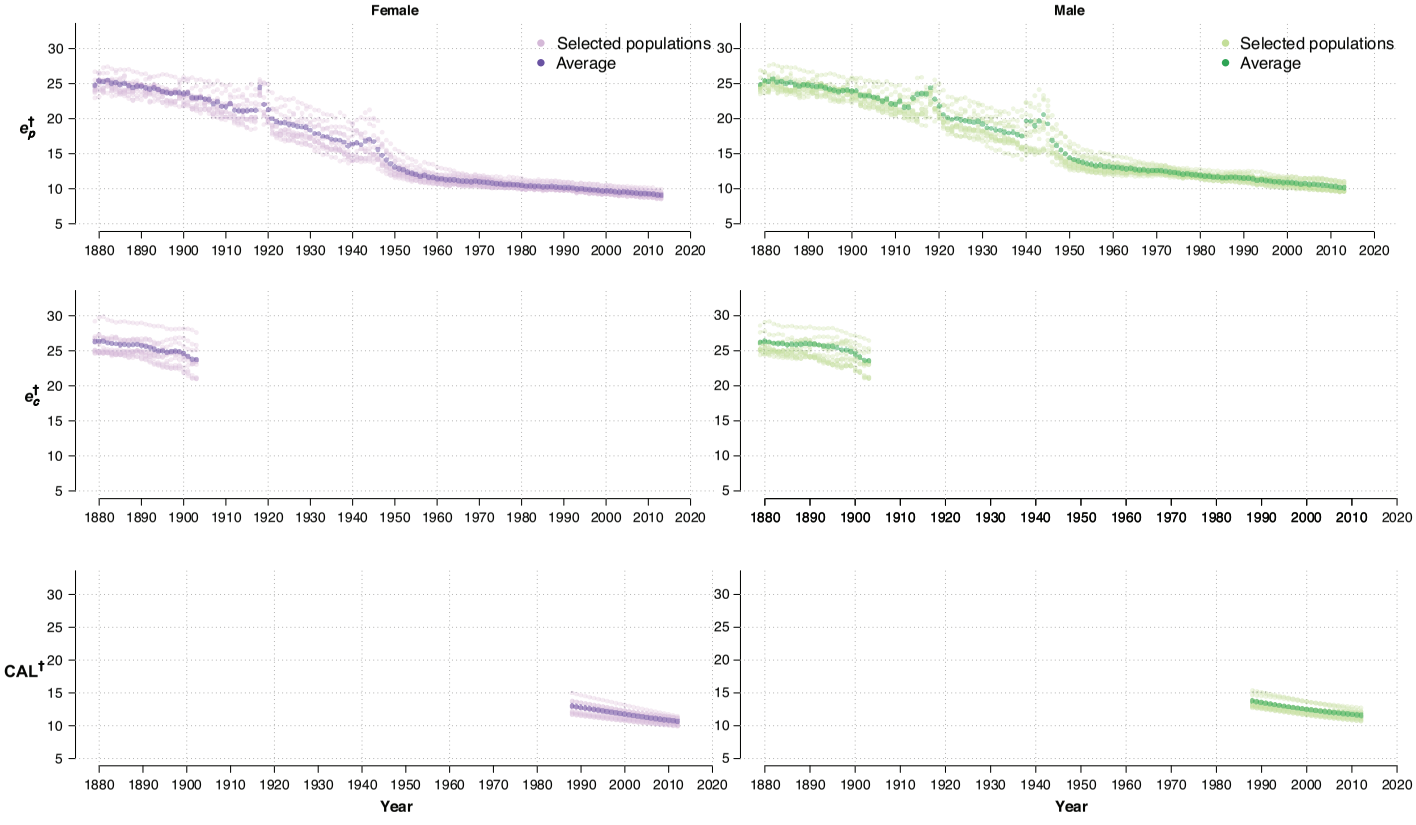

Fig. 1 Trends of period ($e_p^\dagger$), cohort ($e_c^\dagger$), and cross-sectional ($CAL^\dagger$) lifespan variation measures for Denmark, England and Wales, Finland, France, Italy, the Netherlands, Norway, Scotland, Sweden, and Switzerland, for females and males, 1879–2013. *Source:* Authors' calculations based on the Human Mortality Database (2019).

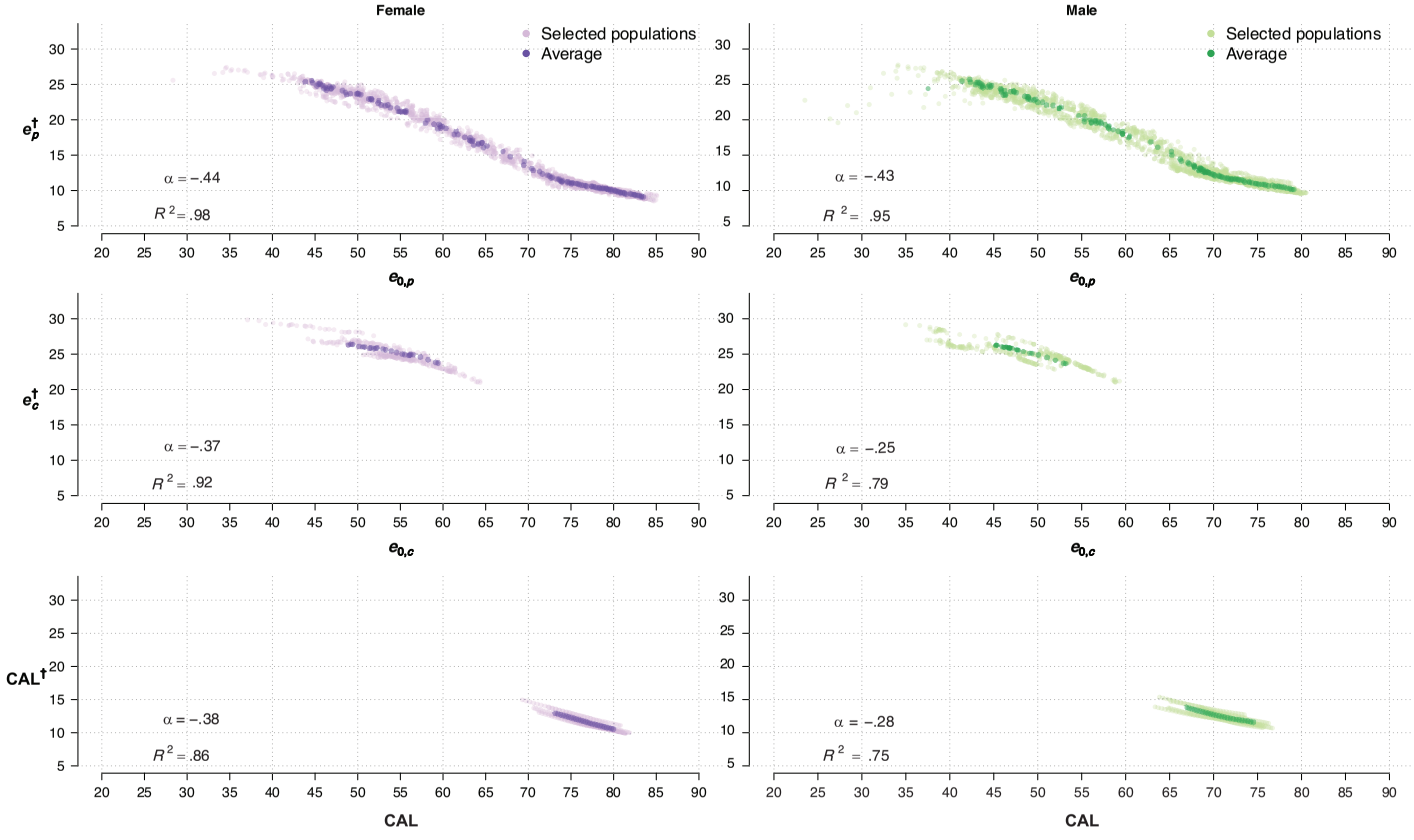

Fig. 2 Pair comparisons between lifespan variation and longevity measures for Denmark, England and Wales, Finland, France, Italy, the Netherlands, Norway, Scotland, Sweden, and Switzerland, for females and males, 1879–2013. α and R^2 correspond, respectively, to the slope and the coefficient of determination derived from linear regressions; ($e_{0,p}$) and ($e_{0,c}$) correspond, respectively, to the period and cohort life expectancy at birth. *Source:* Authors' calculations based on the Human Mortality Database (2019).

Table 1 Period and cross-sectional longevity and lifespan variation measures, and the gap in lifespan variation between each population and the average population, for females and males, 2013

Population	$e_{0,p}$	$e^{\dagger}_{p}$	Gap in $e^{\dagger}_{p}$	CAL	CAL†	Gap in CAL†	$\frac{\text{CAL}^{\dagger}}{e^{\dagger}_{p}}$
Female							
Scotland	81.06	9.96	0.80	77.39	11.12	0.46	1.12
Denmark	82.31	9.50	0.34	78.60	10.99	0.32	1.16
England and Wales	82.96	9.43	0.28	79.64	10.62	–0.05	1.13
France	85.05	9.24	0.09	80.80	11.14	0.47	1.20
Netherlands	83.04	9.18	0.02	80.43	10.35	–0.32	1.13
Norway	83.60	9.01	–0.15	81.10	10.15	–0.52	1.13
Finland	83.83	8.86	–0.30	79.90	10.36	–0.31	1.17
Sweden	83.71	8.85	–0.31	81.39	9.91	–0.76	1.12
Italy	84.95	8.67	–0.48	79.03	11.45	0.79	1.32
Switzerland	84.75	8.53	–0.63	81.86	10.06	–0.61	1.18
Average	83.47	9.16	—	79.95	10.67	—	1.16
Male							
France	78.77	11.07	0.85	73.50	12.73	1.13	1.15
Scotland	77.03	11.03	0.80	72.56	11.94	0.34	1.08
Finland	77.88	10.75	0.53	72.79	11.92	0.33	1.11
England and Wales	79.23	10.32	0.10	75.18	11.21	–0.38	1.09
Denmark	78.26	10.21	–0.01	73.94	11.57	–0.03	1.13
Norway	79.65	9.82	–0.40	75.78	11.09	–0.51	1.13
Italy	80.25	9.75	–0.47	73.64	12.16	0.56	1.25
Switzerland	80.52	9.70	–0.52	76.25	11.43	–0.17	1.18
Netherlands	79.42	9.68	–0.54	75.57	10.78	–0.82	1.11
Sweden	80.10	9.61	–0.62	76.69	10.71	–0.89	1.11
Average	79.08	10.22	—	74.54	11.60	—	1.13

Note: The table is ordered by the gap in $e^{\dagger}_{p}$.

Source: Authors' calculations based on the Human Mortality Database (2019).

in Figure 2 are related to the pace of change in variation and longevity. The slope (α) presented in each panel of Figure 2 was derived from linear regressions, controlling for the number of observations in each pair comparison. While the levels of the CAL measure more closely resemble the end point of the period measures ($e^{\dagger}_{p}$ versus $e_{0,p}$), their pace of change, captured by α, is closer to that observed among the cohort pair measures ($e^{\dagger}_{c}$ versus $e_{0,c}$). The slope of the association between variation and longevity corresponds to a 0.38-year decline in variation for each additional year of longevity in the female cross-sectional (i.e., CAL versus CAL†) dimension, compared to 0.37 years in the cohort dimension and 0.44 years in the period dimension (the corresponding values for males are 0.28, 0.25, and 0.43 years).

Table 1 covers the period ($e_{0,p}$ and $e^{\dagger}_{p}$) and cross-sectional measures (CAL and CAL†) in 2013, and comparisons of indices of lifespan variation between each country and the average population, which is the non-population-weighted composite. The gaps in $e^{\dagger}_{p}$ and in CAL† reveal which populations have higher or lower levels of lifespan inequality than the average. Positive gaps correspond to greater lifespan variation in the index than in the average population. The gaps differ according to the lifespan variation measure used. For Swedish females, the gaps in both $e^{\dagger}_{p}$ and CAL† suggest

that their lifespan inequality was lower than that of the average population, but the gap in CAL† was double that of the gap in $e_p^{\dagger}$ in 2013. For some populations, such as those in Italy and in England and Wales, both the magnitude and the direction of the gap differed depending on the measure used. Italy had above-average variability when the cross-sectional measure was applied, but below-average variability when the period measure was applied; the opposite was the case for England and Wales.

Differences in country rankings (Table 1) also reflect the strong negative correlation between lifespan variation and longevity. In Italy, the change of direction in gaps in $e_p^{\dagger}$ versus CAL† is consistent with the shift of direction in gaps in longevity ($e_{0,p}$ versus CAL). For both female and male Italians, the negative gap in $e_p^{\dagger}$ reflects the above-average longevity when $e_{0,p}$ is used; the positive gap in CAL† reflects the lower levels of CAL than the average. However, there are exceptions to this association, such as in England and Wales.

The last column of Table 1 shows the ratio between CAL† and $e_p^{\dagger}$ in 2013. Compared to $e_p^{\dagger}$, CAL† was about 12–32% higher for females and 8–25% higher for males. The most dramatic difference is found for Italian females, among whom CAL† was 32% higher than $e_p^{\dagger}$ in 2013. In contrast, the discrepancy between CAL† and $e_p^{\dagger}$ was smallest among Scottish males, at about 8%. In Sweden, the discrepancy was about 11–12% for both sexes. To contrast the dynamics of populations that experienced slow and rapid mortality transitions, we examine Sweden and Italy in more detail. The trends in the ratio between CAL† and $e_p^{\dagger}$ in Sweden and Italy from 1989 to 2013 can be found in the online supplemental material (see OSM-6).

To better assess the gap in CAL† between Sweden and Italy relative to that of the composite average population of 2013, Figure 3 presents the age-cohort contributions to the difference in CAL†, accumulated over cohort. In Sweden, CAL† was lower than the average (−0.76 for females and −0.89 for males) because of the below-average cumulative age-cohort contributions of all cohorts born in the early 1930s onward. In other words, all the cohorts born from the early 1930s contributed to compressing the Swedish 2013 mortality distribution compared to the average distribution. This dynamic was different in Italy, where CAL† was higher than average by 0.79 for females and 0.56 for males. In Italy, CAL† was higher because of the above-average cumulative age-cohort contributions to the gap of all cohorts present in 2013, except for those who were around age 90 and those who were very young in 2013. In both countries, cohorts that reached ages between 60 and 80 in 2013 contributed substantially to the gap in CAL†. The online supplemental material (see OSM-7) shows the Lexis surface for the cumulative age and cohort contributions to the difference in CAL† for Sweden and Italy.

The cumulative age-cohort contributions to the gap in CAL† (Figure 3) can be broken down by age- and cohort-specific contributions that led to their development. Figure 4 shows the age and cohort decomposition of the gap in CAL† (Sweden and Italy each compared to the average population) for males. The female age- and cohort-specific contribution to the difference in CAL† between Sweden and Italy each versus the average population is accessible in the app https://caldagger.shinyapps.io/CALdagger/.

For aid in interpreting Figure 4, red hues refer to age-cohort contributions that led to higher CAL† than the average, while blue hues refer to the opposite. At some ages, the contributions to the differences between Sweden and Italy versus the average

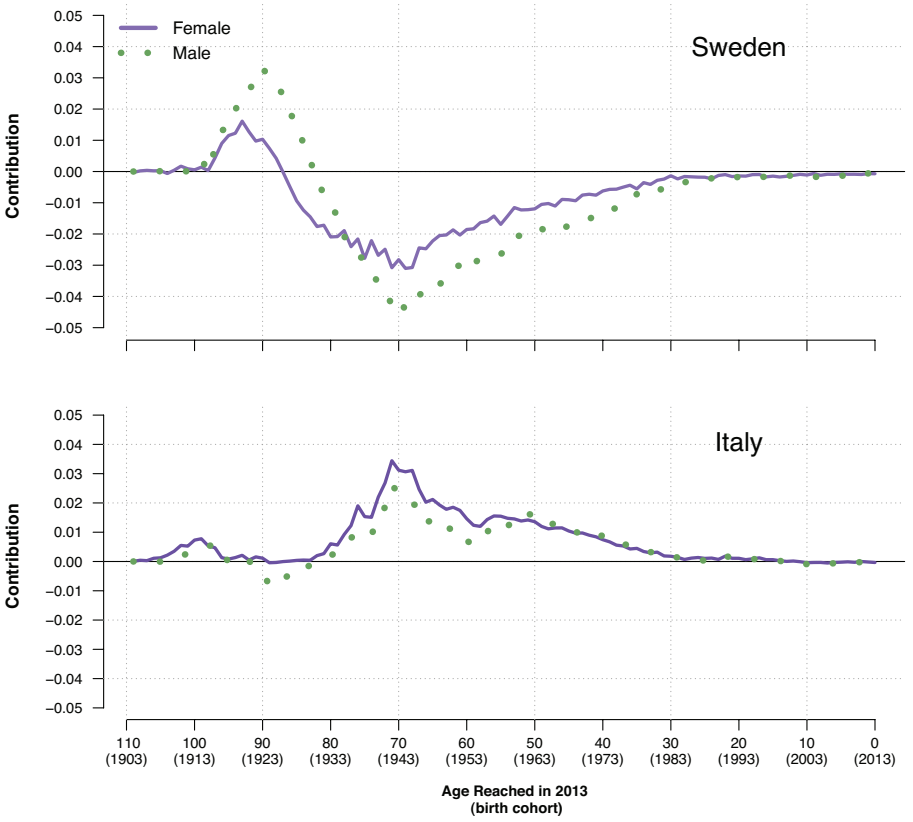

Fig. 3 Cumulative age-cohort contribution to the gap in CAL† in Sweden and Italy with respect to the average population, for females and males, 2013. *Source:* Authors' calculations based on the Human Mortality Database (2019).

population were zero (white). Specifically, we can see a white diagonal band (see the "crossover line") at which the contributions to the differences in CAL† of the male cohorts born around 1930 were zero from birth to the age reached in 2013. This crossover line is analogous to the threshold age of other lifespan variation measures, such as life disparity and the lifetable entropy. Progress in mortality needs to occur below a threshold age to reduce lifespan variation; mortality reductions above that threshold age increase lifespan variation (Aburto et al. 2019; Aburto et al. 2020; Vaupel et al. 2011; Zhang and Vaupel 2009). In our between-country comparison, this property can be translated to indicate that lower mortality than the average population below the crossover line contributes to lower CAL†, while lower mortality above such a diagonal line contributes to increasing CAL† compared to the average population (Table A2 in the OSM-8 helps to interpret these mortality dynamics). This also explains why there is a color switch in the Figure 4 panels above and below this line. Below the crossover line (i.e., for cohorts born after 1930), blue points would indicate mortality lower than the composite population because they are occurring for cohorts who in 2013 were at ages below the threshold age, which compress mortality.

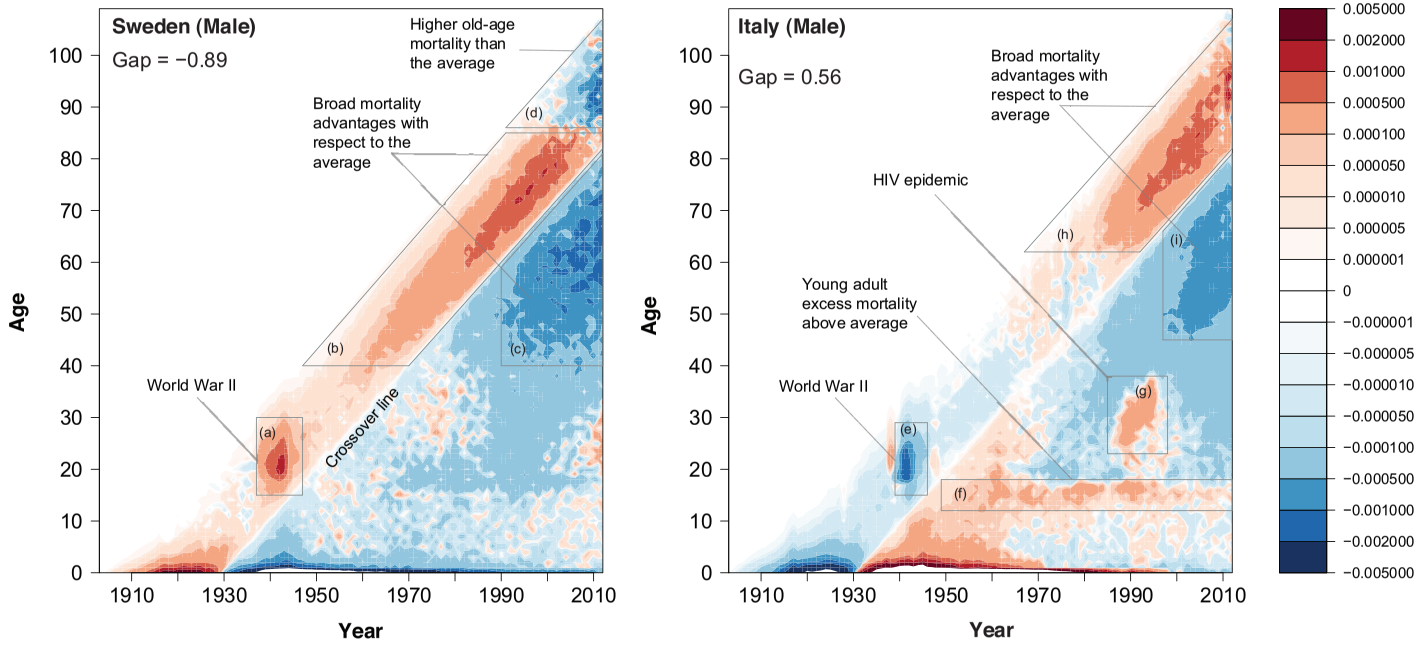

Fig. 4 Lexis surface for the age- and cohort-specific contribution to the difference in CAL† (2013) between Sweden and Italy, each compared to the average population, for males. Red hues indicate contributions to higher CAL† in Sweden and Italy each compared to the average population, and blue hues indicate contributions to lower CAL† in Sweden and Italy each compared to the average population. The sums of the age- and cohort-specific contribution to the difference in CAL† (2013) in the outlined areas are: (a)=0.043, (b)=0.237, (c)=−0.466, (d)=−0.055, (e)=−0.037, (f)=0.022, (g)=0.002, (h)=0.186, and (i)=−0.241. *Source:* Authors' calculations based on the Human Mortality Database (2019).

Above this line, blue points would indicate mortality that is higher than the average population, because these points come from cohorts who in 2013 were at ages above the threshold age.

In the comparison between Swedish males and the male average population (Figure 4), we can see that mostly lower mortality of cohorts born before 1930 contributed to increasing the Swedish variation in the age at death compared to the average, such as the lower mortality impact from World War II compared to the average population that increased the Swedish CAL† by 0.043 with respect to the average (area "a" in Figure 4). However, this trend reversed at the oldest ages, in which the Swedish higher old-age mortality decreased the Swedish CAL† by 0.055 compared to the average (area "d" in Figure 4). By contrast, most age-cohort combinations from cohorts born after the 1930s contributed to lower CAL† than the average population, which can also be seen by the cumulative age-cohort contribution of these cohorts in 2013, as presented in Figure 3. Ages below five, particularly for those born between 1930 and the early 1970s (aged 40–80 in 2013), were especially important. The broad survival advantages of these cohorts, as outlined by area "c" in Figure 4, led to a −0.466 change in the Swedish CAL† compared to the average. Also, Figure 4 complements Figure 3 by revealing that the substantial contribution to compressing the Swedish distribution of cohorts between ages 60 and 80 in 2013 started at ages below five.

In Italy, the oldest cohorts—that is, those born before 1930—contributed to reducing the Italian CAL† compared to the average over the first and young adult ages because of their higher mortality. The impact of World War II among Italian males, for instance, decreased the Italian CAL† by 0.037 with respect to the average (area "e" in Figure 4). However, at older ages, these cohorts experienced lower-than-average mortality, with the contributions from these ages to CAL† differences becoming positive, which summed to 0.186 (area "h" in Figure 4). When summed over cohorts (see Figure 3 and Figure A5 in OSM-7), the overall contribution from these cohorts led to higher-than-average CAL†, meaning that the lower mortality at older ages more than compensated for their higher mortality at younger ages. For the Italian cohorts born from the 1930s, Figure 4 shows higher mortality below age five and at young adult ages compared to the average. The young adult excess mortality between ages 10 and 20 as outlined by area "f" in Figure 4 was roughly constant over time for these Italian cohorts, which increased the Italian CAL† by 0.022 compared to the average.

Discussion

Summary of Results

We introduced CAL† as an indicator of lifespan variation that includes the mortality experience of all cohorts present at a given time. The CAL† trends were remarkably similar to those of life disparity from a period and a cohort perspective, albeit with different levels. Lower levels of CAL† were associated with higher levels of longevity, as measured by CAL. Variation in the age at death differed substantially between the cohort, period, and cross-sectional measures: CAL† was consistently higher than

period lifespan variation, although the extent of these differences varied across populations and over time. As a result, the cross-country rankings of lifespan variation changed greatly depending on the measure used. The age and cohort decomposition of the difference in CAL† revealed the large contributions of the mortality changes at young ages to lower or greater disparity—that is, which compressed or expanded the variability in the age at death—even for those cohorts who reached adult and old ages in 2013.

Interpretation

CAL† can be interpreted as the variation in mortality over age that all cohorts present in the cross section have been exposed to, on average, over their lifetime. For the populations and years that we examined, CAL† was always higher than period $e^{\dagger}$. This is not surprising, given that recent period mortality schedules may be expected to have much lower mortality over younger ages than the mortality levels experienced by a cross-sectional average of past cohorts over these ages.

Perhaps more interesting was our finding that the population-level rankings in lifespan variation varied considerably depending on whether the period or the cross-sectional average perspective was applied. The most dramatic example of the differences was found for Italian females: across the 10 countries investigated in 2013, CAL† was the highest, but period $e^{\dagger}$ was the second lowest, for this group. In contrast, Swedish females had a period $e^{\dagger}$ that was not much higher than that of their Italian counterparts, but a CAL† that was a full year and a half (around 15%) lower. The implication here is that Italian females were exposed to much greater variability in the age at death over their life course than Swedish females, even if the current period schedule of mortality suggests that lifespan variation has been similar for the two populations.

To highlight the substantive meaning behind the different CAL† and $e_p^{\dagger}$ levels, we can put these differences into historical perspective. For Italian females, for instance, the ratio between CAL† and $e_p^{\dagger}$ is 32% in 2013. This ratio corresponds, approximately, to 40 years of reductions in $e_p^{\dagger}$; more specifically, it corresponds to $e_p^{\dagger}$ reductions between 1970 and 2013 (36%) when $e_p^{\dagger}$ decreased from 11.78 to 8.67. Among Scottish males, the ratio of 8% between CAL† and $e_p^{\dagger}$ corresponds to reductions in $e_p^{\dagger}$ over more than three decades ($e_p^{\dagger}(1980) = 11.96$ and $e_p^{\dagger}(2013) = 11.02$); while for Swedish males, the ratio of 11% is similar to more than two decades of reduction in $e_p^{\dagger}$ ($e_p^{\dagger}\left(1990\right) = 10.87$ and $e_p^{\dagger}\left(2013\right) = 9.60$). These comparisons show that CAL† approaches $e_p^{\dagger}$ lagged by 20 to 40 years. In other words, the ratios between these two lifespan variation measures would be equivalent to 2–4 decades of improvements in $e_p^{\dagger}$.

Thus, if lifespan variation was a good indicator of individual uncertainty around the timing of death, it should accurately reflect how individuals assess their prospective chances of survival to future ages. Whether such individual assessments are best quantified by period, cohort, or the cross-sectional average mortality experience remains a key open question. As the Italian, Scottish, and Swedish examples demonstrate, the differences in the lifespan variation according to the temporal dimension examined are not trivial.

Uncertainty in the Timing of Death

Previous studies have shown that individuals internalize their surrounding mortality (Dormont et al. 2018; Hurd et al. 1998; Hurd et al. 2004; Perozek 2008). Precisely how observations of mortality progress influence people's subjective survival expectations and uncertainty about the timing of death is an intriguing question that warrants further investigation (Sasson 2016). Individuals might base their subjective expectation of survival on the mortality conditions experienced by their friends, spouses, parents, or other relatives. The death of a parent seems to affect both the subjective survival probabilities and survival uncertainty (Dormont et al. 2018; Hurd and McGarry 1995). Indeed, the individual's subjective lifetime uncertainty increases when one of her or his parents is deceased (Dormont et al. 2018).

By introducing $CAL^{\dagger}$, we proposed a measure of lifespan variation that reflects the tendency of individuals to base their lifetime uncertainty on the survival trajectories of family members, colleagues, and neighbors who died during their lifetime, but are of mixed birth cohorts. Our results indicated that uncertainty in the timing of death was greater when the mortality histories of several cohorts were taken into account. This finding is important because greater uncertainty is highly undesirable for most people.

People's expectations and uncertainty about their future survival are instrumental in their key life course decisions. There is evidence that subjective expectations affect the demand for insurance products (Brown 2001; Inkmann et al. 2011; O'Dea and Sturrock 2019). Individuals who are pessimistic about surviving to older ages are more likely to retire early, to save less for retirement, and to avoid purchasing annuities (Heimer et al. 2019; O'Dea and Sturrock 2019). It has been argued that whether people are optimistic about reaching an advanced age affects the likelihood of adopting a healthy lifestyle or of undergoing health screenings (Picone et al. 2004; Scott-Sheldon et al. 2010), and can even influence the timing of childbearing (Eisenberg and Schenker 1997; Geronimus et al. 1999; Nettle 2010; Rindfuss and Bumpass 1976).

Implicit within much of this literature linking subjective survival expectations with individual behaviors is that individuals base their life course decisions on achieving average levels of survival. But it is equally plausible that individuals also consider the degree of uncertainty in survival outcomes in planning their lives. Indeed, a French survey that measured the respondents' subjective survival probabilities of surviving to multiple ages showed that the level of subjective uncertainty about longevity (the standard deviation of their subjective expected age at death) closely matched life table standard deviations in the age at death, and had an association with risky behavior that was independent of their subjective life expectancy (Dormont et al. 2018).

Differences between individuals' and insurers' perceptions of lifetime uncertainty can have implications for the annuity market. Considering that annuities are primarily used as an income flow for retirees, if insurers underestimate individuals' perceptions of lifetime uncertainty, sales of annuities may be negatively affected. By showing that there are considerable differences between period, cohort, and cross-sectional lifespan variation measures that are in line with different subjective assessments of lifetime uncertainty, our results can shed light on the low demand for annuities (Agnew et al. 2015; Alexandrova and Gatzert 2019). Individuals who base their uncertainty about

their time of death on the survival trajectories of family and friends who have died in the past and are from different birth cohorts might see an annuity as a worse deal than people who base their survival uncertainty on current mortality rates, since they believe their chances of dying early are higher and their risk of outliving their savings is lower.

The decision to buy an annuity, as a way to create retirement income, varies across social groups. For instance, higher education has a positive effect on annuitization (Agnew et al. 2015). Thus, in a heterogeneous population, the subjective assessments of lifetime uncertainty may also differ across social groups. How these assessments are formed remains poorly understood. However, if socially disadvantaged groups, which are less likely to buy an annuity and have higher mortality levels, base their lifetime uncertainty on cross-sectional measures ($CAL^{\dagger}$ higher than $e_p^{\dagger}$), they may assess annuities as a worse deal than their actuarial period probabilities of dying. But if, on the other hand, socially advantaged groups base their lifetime uncertainty on projections of falling period mortality, they may perceive annuities as a better deal than when based on the $CAL^{\dagger}$ perspective. Thus, both social gradients in mortality and differences in perceptions of lifetime uncertainty across social groups may amplify inequalities in annuity demand.

Examining the differences in the levels of uncertainty about the timing of death across individuals and population groups might help us gain a better understanding of the motivations behind adverse health behavior. Individuals with high levels of lifetime uncertainty may prioritize goals with immediate outcomes, which could make them less likely to adhere to health promotion and disease prevention and to partake in risky behavior. Whether individuals follow the advice of public health programs to stop smoking, lose weight, exercise, and make other lifestyle changes greatly depends on their levels of engagement (Cameron and Best 1987; Middleton et al. 2013). Thus, individuals' perceptions of the timing of death should be included in health program discussions since the effectiveness of these programs depends on individual health behavior.

Measuring Lifespan Variation

Our results revealed that different levels of uncertainty in the timing of death arise from the mortality conditions taken into account by each measure. This finding was expected because the period, cohort, and cross-sectional lifespan variation measures consider, respectively, the current mortality rates, the past mortality conditions, and the mortality conditions previously experienced by the cohorts in a cross-sectional approach.

Consistent with previous studies (Engelman et al. 2010; Wilmoth and Horiuchi 1999), we found higher levels of cohort than period variation in the age at death. This result was largely attributable to the higher levels of infant mortality experienced by the extinct cohorts compared to those observed in the period basis. We also compared the period with the cross-sectional approach. Because it took high past levels of mortality into account, $CAL^{\dagger}$ revealed higher levels of variation in the age at death than the period life table–based index. The comparison between $CAL^{\dagger}$ and cohort life disparity is not straightforward. Cohort measures go from 1879 to 1903, with longevity values being lower than 65 and 60, respectively, for females and males, and the variation in the age at death between 21 and 30 for both sexes. In contrast to the cohort approach,

cross-sectional measures start from 1989 with longevity values all above 69 for females and 63 for males, and CAL† between 9 and 15. Thus, comparisons matching observed cohort measures with the cross-sectional ones (Guillot and Kim 2011; Guillot and Payne 2019), for both longevity and inequality in lifespan indices, are limited by our years of observation. Matching CAL† with the average number of years lost due to death for an actual cohort can strengthen the interpretation of CAL†. To do that, one should access cross-sectional indices before the year 1879, or forecast mortality into the future.

In interpreting the ratios between CAL† and the period life disparity, it is important to distinguish between current mortality rates and current mortality conditions. To clarify this point, we turn to the period life expectancies. Demographers usually describe life expectancy as the average lifespan of a hypothetical cohort of individuals who live their lives under current mortality conditions (Preston et al. 2001; Vaupel 2002; Vaupel 2008). However, the current mortality rates at each age result not only from the current epidemiological environment, but also from the selective historical mortality experiences of the particular cohort who has survived to each age (Vaupel 2002). This makes the interpretation of period expectancies (and variability) somewhat awkward (Guillot and Payne 2019). CAL† is an alternative measure of the variability in the age at death that unifies both past and present mortality in a cross-sectional approach.

CAL† tracks mortality changes by all the cohorts present at a given time, from birth to the present. This is an important advantage because over the last century most populations experienced substantial changes in mortality (Meslé and Vallin 2005; Meslé and Vallin 2011; Rau et al. 2017; World Health Organization 2000). In the past, survival improved faster at younger than at older ages as a result of reductions in infectious disease. Infant and child mortality decreased substantially over the last century. Older birth cohorts were at higher risk than younger cohorts of dying during the first years of life. More recently, mortality decline in modern developed countries has been achieved mainly by reductions in circulatory and other chronic diseases at adult and older ages (Bergeron-Boucher et al. 2020). As a result of these historical changes, the current population is a mix of several birth cohorts that faced a vastly different set of age-specific death rates through their life.

In recent decades, there is evidence that in some national populations (Aburto and Beltrán-Sánchez 2019; Aburto and van Raalte 2018; Edwards and Tuljapurkar 2005; García and Aburto 2019; Gillespie et al. 2014; Seaman et al. 2016) and socioeconomic groups (Brønnum-Hansen 2017; Permanyer et al. 2018; Sasson 2016; Seaman et al. 2019; van Raalte et al. 2014), lifespan variation has stalled or increased, often alongside increasing life expectancy. These findings refer mainly to populations in which midlife mortality has been stagnating or increasing, while mortality at very young and older ages has continued to improve. Because none of the populations we used in our illustration experienced sustained mortality increase over midlife, we were unable to examine how this dynamic has been evolving over single cohorts, or in a cross-sectional cohort perspective. Nevertheless, in addition to detecting a negative association between CAL and CAL†, our findings showed a slower slope in the association between CAL and CAL† than between the pairs of $e_{0,p}$, and $e_p^{\dagger}$. Thus, by taking into account past and present mortality, CAL and CAL† were found to move more gradually. Our results also showed that changes in CAL were associated with slower changes in CAL† when compared with changes in the period and cohort pairs of mortality measures.

One limitation of $CAL^{\dagger}$ is its interpretation as the average number of years lost due to death, as we mentioned in the Data and Methods section and mathematically showed in the online OSM-3. Another drawback of $CAL^{\dagger}$ is the requirement of historical mortality information for its calculation, which to date is available only for some developed countries. As long mortality series are becoming more available, $CAL^{\dagger}$ can be calculated for a broader number of populations. Completing the mortality histories of cohorts near extinction (for instance, those who have already reached the modal age at death) would also allow for a longer time series. However, efforts need to be made to measure the variation in the ages at death in a cross-sectional perspective in countries that lack the mortality history of cohorts. For those populations, a truncated version of $CAL^{\dagger}$, an approach similar to that presented by Canudas-Romo and Guillot (2015) for the CAL measure, may help the investigation of the lifespan variation in a cross-sectional perspective.

Conclusion

We introduced a novel lifespan variation measure to the demographic literature. This measure is a cross-sectional approach that reflects the experiences of individuals who base their level of lifetime uncertainty on both past and current mortality conditions. We highlighted the potential implications of higher levels of uncertainty about the timing of death on individual life course decisions. ■

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