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Regional Mortality in the United States at Ages 80 and Older: An Analysis of Direct Estimates, 1959–2011

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Abstract

The almost-extinct cohort method was used to produce direct mortality estimates for states of the United States in the period 1959–2011 and at ages 80 and older. The estimates produced by this method were found unreliable for data for the 1960s, due to heavy age misreporting in the U.S. data on deaths. However, following dramatic improvements in the quality of U.S. data at older ages over the last four decades, mortality estimates for the period 2000–2011 were found to be reasonably good. In 2000–2011, levels of mortality in the United States were shown to be very similar to average levels of mortality in Japan (with the exception of Japanese females) and in 12 European countries with high longevity. Disparities in mortality among U.S. states were also comparable with disparities existing in the 13 high-longevity countries. Overall, mortality was lower in Western and Northeastern U.S. states and higher in Southern U.S. states. Hawaii stood out as a state with exceptionally high survival rates at advanced ages.

The views expressed in the paper do not imply the expression of any opinion on the part of the United Nations Secretariat or Duke University.

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Introduction

Despite the importance of mortality data for decision making and determining public policy, our knowledge about mortality trends and variations at advanced ages in the United States remains incomplete and dated. For example, the latest State Life Tables are available for 1999–2001 published as a part of the U.S. Decennial Life Tables for 1999–2001 (Arias et al., 2008), and the latest complete life tables by sex with an open-ended age category 100+ are available for the year 2012 (Arias, 2016). One of the reasons for the lack of up-to-date information on mortality trends and the within-country variability of mortality levels is the questionable quality of directly computed death rates at advanced ages, especially at ages 85 and older (Anderson, 1999; Coale and Kisker, 1990; Elo and Preston, 1994; Bennett and Olshansky, 1996). Therefore, the construction of U.S. life tables requires deriving mortality estimates at advanced ages by using *indirect* methods, rather than using directly computed death rates that may be smoothed by an appropriate graduation technique. The approach used by the National Center for Health Statistics (NCHS) for the adjustment of U.S. death rates at advanced ages involves the incorporation of information from Medicare records in either of two ways: (1) using observed rates of increasing mortality with age in the Medicare data set (or a subset of Medicare records deemed to be the most reliable) for extrapolation of U.S. death rates into higher ages (Anderson, 1999); or (2) substituting directly computed death rates for those from the Medicare data set (Arias et al., 2008). In both cases, directly computed death rates are assumed to be reliable until some age (e.g., 65), and then indirectly computed death rates (based on Medicare data) are used beyond this age to produce a complete set of death rates needed to compute a life table.

Anomalous age patterns of mortality at advanced ages in the United States are commonly attributed to age misreporting in censuses and in death registrations (Coale and Kisker, 1990; Elo and Preston, 1994). The problem of age misreporting is not limited to the United States (Bennett and Garson, 1983; Garson, 1991; Coale and Li, 1991; Bourbeau and Lebel, 2000) and is closely linked to the development of universal and complete birth registration systems. Persons with birth certificates are much more likely to report their ages accurately, leading to more accurate estimates of age patterns of mortality (Anderson, 1999; Rosenwaike and Hill, 1996; Kannisto, 1988). As the U.S. birth registration system became complete in 1933 (Hetzel, 1997), the quality of age reporting on death certificates is expected to improve over time, as more people who reach advanced ages will possess a birth certificate.

Data quality is likely to vary across states in the United States. States were admitted to the birth registration area in different years—when their birth registration reached 90 percent completeness. Northeastern states were generally admitted first, in 1915, and Southern states were generally admitted last. The process was completed in 1933 with the admission of Texas. In some states (e.g., Massachusetts), birth records are on file and available for an entire area since the middle of the 19th century. The high volume of migrants from countries with incomplete or nonexistent birth registration systems at the beginning of the 20th century also could be an important factor affecting the quality of age reporting in some states.

In this article, we estimated the levels and trends in death rates at ages 80 and older for each state in the United States and performed comparisons with 13 countries with reliable data (Thatcher et al., 1998). The objectives of this research were twofold: first, to evaluate quality of direct mortality estimates for individual U.S. states, and second, to examine whether data quality has improved over time in the United States.

Data

Annual life tables for ages 80 and over and for the period from 1959 to 2011 have been computed for each U.S. state, based on individual death certificate records disseminated by NCHS in the form of Mortality Detail Files available since 1959. Survivor estimates for non-extinct cohorts needed for application of the “almost extinct cohort” method or method of “extinct generations” (Thatcher et al., 2002) are based on intercensal estimates of the resident population of the United States produced by the U.S. Census Bureau (2011). Data for Alaska were excluded from the analysis, because the almost-extinct cohort method did not produce plausible estimates of mortality for this state—either due to small population sizes at advanced ages and/or due to severe data inaccuracies. The data for Alaska are included in the estimates for the United States as a whole.

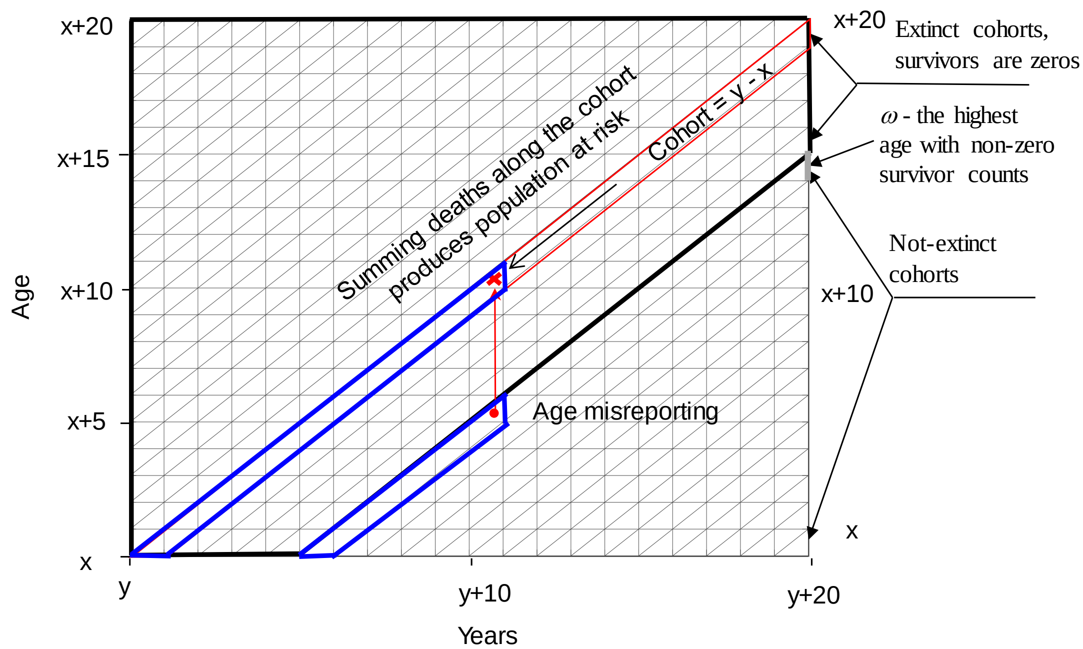
For comparison, we selected 13 countries with high longevity and whose mortality data are considered to be of the best quality available: Austria, Denmark, England and Wales, Finland, France, Germany (West), Iceland, Italy, Japan, the Netherlands, Norway, Sweden and Switzerland (Thatcher et al., 1998). Mortality estimates for the United States and Denmark were obtained from the World Mortality Trends website (see http://www.mortalitytrends.org [accessed May 2016]), and the estimates for the rest of the countries from the Human Mortality Database (see http://www.mortality.org [accessed May, 2016]).

Methods

The almost-extinct cohort method was used to produce mortality estimates at ages 80 and older and for years 1959–2011 for each U.S. state and for the comparison countries (Vincent, 1951; Kannisto, 1988; Thatcher et al., 2002, Andreev *et al.*, 2003, Wilmoth et al., 2007). The estimates were based on data on deaths from vital registration and data on population from censuses or population registers.

The method requires deaths tabulated by calendar year, age, and the Lexis triangle (so the deaths could be arranged into cohort mortality histories) and estimates of survivors for non-extinct cohorts. In this case, the population at risk could be obtained by cumulating deaths from the highest age attained by cohorts working backward along cohorts (Fig. 1). Summing deaths starts with 0 for extinct cohorts and with the remaining count of survivors for non-extinct cohorts. Mortality estimates produced by the almost-extinct cohort method are based almost entirely on the death data; it is only for the last few years that the population at risk is a mixture of survivor estimates and the population produced by summing deaths. Going further back in time reduces the share of survivors of non-extinct cohorts in the population at risk, and mortality estimates become increasingly based on cumulated deaths. For extinct cohorts, the population at risk is based entirely on death data, because no survivors are included in computation of the population at risk. This method has been used extensively over the last two decades for the estimation of mortality at advanced ages (Andreev et al., 2003; Wilmoth et al., 2007).

Fig. 1. Illustration of the Almost-Extinct Cohort Method and the Effect of Age Misreporting on Mortality Estimates



Effect of Age Misreporting on Mortality Estimates Produced by the Almost-Extinct Cohort Method

The almost-extinct cohort method is simply based on population accounting in a closed population. It produces correct results if migration is negligible, age reporting in deaths and population counts are accurate, death and population registrations are complete, and the coverage of death and population statistics is the same. Errors in both death and population data will bias estimates of death rates produced by this method. Errors in survivor estimates, however, will only bias mortality estimates in the lower-right triangle of the data (see Fig. 1). Typically, only death rates in the last few years are affected, because the deaths cumulated above a certain age in a cohort will eventually dominate over the survivor estimates for the population-at-risk estimates at this age.

Age misreporting in death data, however, affects mortality estimates in a more extensive way. Figure 1 illustrates the effect of age misreporting for a single death: for a death registered in the year $y + 10$ at age $x + 5$ (Fig. 1, red circle, cohort $y – x + 5$), it is assumed that age was reported as $x + 10$, and thus overstated by 5 years of age (red cross, Fig. 1). Excluding a single death from the $y – x + 5$ cohort not only reduces by 1 the number of deaths at age $x + 5$ in this cohort, but also reduces the population-at-risk for all ages below $x + 5$ in the same cohort, because the misreported death is excluded from the summing process in the almost-extinct cohort method. Similarly, the death with misreported age will be registered at age $x + 10$ in the cohort $y – x$, increasing not only the true number of deaths at age $x + 10$ but also the population-at-risk estimates for all ages below $x + 10$ in this cohort.

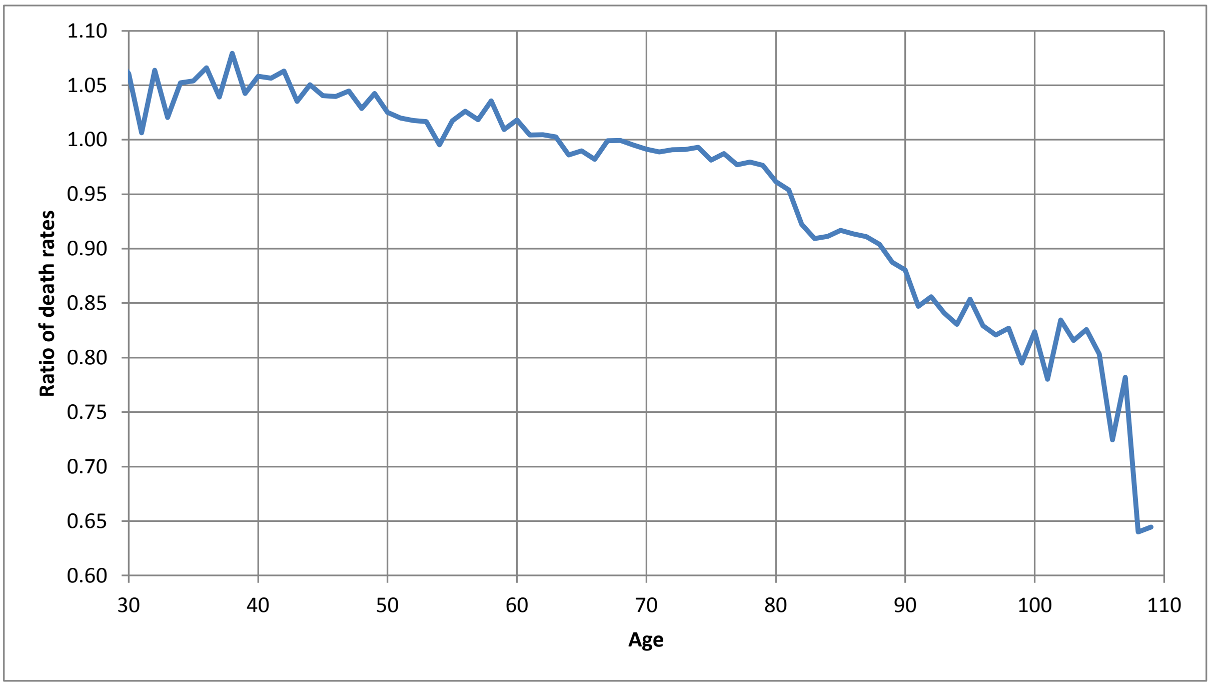
This simple example clearly shows that misreporting of a single death affects mortality estimates for two cohorts for all years, and for all ages before the year of registration. For a younger cohort ($y – x + 5$), all death rates below $x + 5$ are affected, and the misreported death rates are higher than the true death rates because misreporting removes deaths from the population-at-risk. If mortality is constant over time and the true death rates in both cohorts are the same, this type of age misreporting will produce an artificial increase in death rates over time, due to the effect on mortality in two cohorts—i.e., the reduction of death rates in the older cohort and the increase of death rates in the younger cohort. If mortality rates are declining (as is typical of contemporary mortality trends in high-longevity countries), this type of age misreporting will dampen rates of mortality improvement.

To gain further insights into the effects of age misreporting on mortality estimates by the almost-extinct cohort method, we conducted a simple simulation study. We assumed that the population is stationary with mortality represented by the 1960 U.S. male decennial life table (NCHS, 1964) with life expectancy at birth and at ages 80 and 95 being 66.8, 6.0, and 2.4 years, respectively. We further assumed that age misreporting is symmetrical and independent of age, with no net overstatement or understatement of age at death, and with a 95% chance that the misreported age will be less than five years below/above the true age at death (see Note 1 in Appendix A). We applied this model to each death in the life table and recomputed a life table based on misreported deaths.[1]

Figure 2 shows the ratio of the misreported death rates to the true death rates. Up to age 65, the misreported death rates are higher than the true death rates; however, the excess mortality is generally small (less than 5%). At ages 65 to 75, the true and misreported death rates are similar, and age misreporting has no effect on the death rates. At ages above 75, the misreported death rates become progressively lower than the true death rates—approximately 5% lower at age 80, 15% lower at age 90, and 20% lower at age 100. The effect of age misreporting on remaining life expectancy is more uniform. Estimated life expectancy in the life table based on misreported deaths is higher than in the original life table at all ages above age 25. Especially pronounced are the differences in estimated life expectancy at the highest ages: 6.4 vs. 6.0 years at age 80, and 2.9 vs. 2.4 years at age 95. Even if age misreporting produces a crossover between true and misreported death rates, the upward bias in the death rates below age 65 is not enough to offset the downward bias in death rates above age 75 all the way back to age 25. Only at age 25 does life expectancy in both life tables become close (45.2 years).

[1] We use 1 million for the life table radix instead of 100,000 to get more stable simulation results.

Fig. 2. Ratio of Death Rates in the Population with Simulated Age Misreporting to True Death Rates



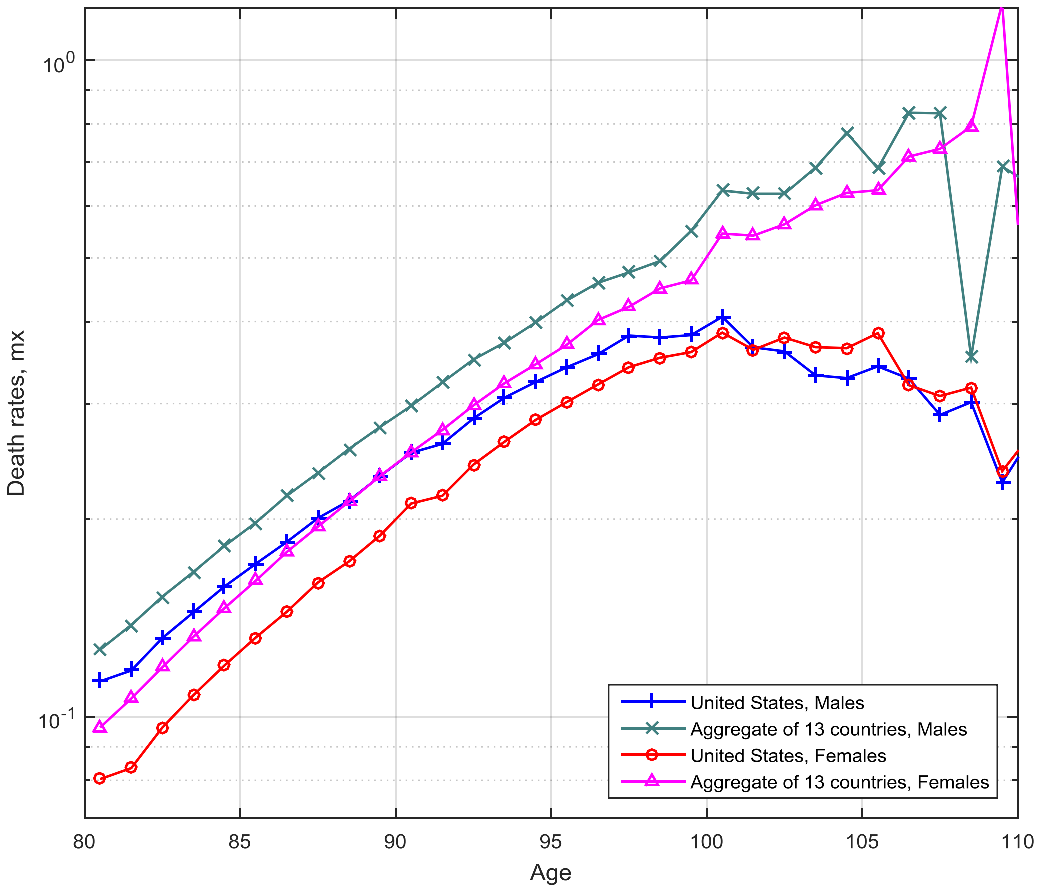
The simulation study of symmetric age misreporting produced results similar to that of Preston *et al.* (1999), who reported that under common scenarios of age misreporting (understatement, overstatement and symmetric), life expectancy beginning at age 80 will be overestimated in the data affected by age misreporting. In our simulation, this result holds true as well: misreported life expectancy is higher for all ages beginning at about age 25. The effect of age misreporting on death rates, however, is not as uniform on remaining life expectancy. Symmetric age misreporting does not bias death rates downward uniformly, but instead produces a mortality crossover: the misreported death rates are *lower* than true death rates at ages 75 and older, but *higher* below age 65. Age misreporting has little effect on death rates only in the age group 65–75 (the misreported and true death rates are approximately the same). A mechanism to explain why even symmetric age misreporting could have a dampening effect on death rates at the oldest ages is further outlined in Note 2 (Appendix A).

A common manifestation of age misreporting is not only implausibly low death rates at the highest ages, but also implausibly low rates of increasing mortality with age. The rate of mortality increase with age in the life table with misreported deaths will be lower than the rate of increase in the original life table, as a consequence of progressive downward bias in death rates introduced by age misreporting. In our simulation exercise, death rates in the original life table were increasing at rates 0.074, 0.083, 0.065 and 0.066 over the age intervals 60–79, 80–89, 90–99 and 100+, respectively. In the life table based on misreported deaths, the rates of increase over the same age intervals were only 0.073, 0.077, 0.059 and 0.054—i.e., lower by approximately 2%, 8%, 10% and 22%, respectively.
Another common manifestation of age misreporting is the spurious temporal increases in death rates. That is, if the quality of data is improving over time and age reporting is getting more accurate, observed death rates will be increasing over time simply because the biased-downward misreported death rates will be approaching true levels of mortality due to the vanishing effect of age misreporting on death rates. If the quality of data is improving and mortality is declining at the same time, observed mortality dynamics will be complex, as these two factors are acting in opposite directions. One may observe, for example, an initial increase in death rates due to improvements in data quality, followed by a decline in death rates due to reductions in mortality.

Mortality Levels in the Period 1959–1969

In the first decade (1959–1969) of the analyzed period, U.S. mortality estimates produced by the almost-extinct cohort method are significantly lower than mortality in the pooled data for the 13 countries (Fig. 3). The U.S. male life expectancy at age 80 is estimated to be 5.9 years, whereas it is only 5.3 years (more than a half year lower) in the 13 aggregated countries. For females, the difference is even larger, close to one year (7.2 vs. 6.2 years, respectively). The rates of mortality increase with age in the United States are lower than in the 13-country aggregate, and the difference is exacerbated with advancing age. Moreover, the rates of mortality increase in the United States appear anomalous at advanced ages in males in females. By about age 97, the death rates stop increasing, and after age 100, they even start to decline. The sex ratio of mortality also appears abnormally low after age 100, with the female survival advantage largely disappearing after age 100 and with the direct estimate of male e_{100} (2.85 years) even higher than that of females (2.79 years). In the pooled data, conversely, the female mortality advantage persisted for the entire age range.

Fig. 3. The Force of Mortality by Age in Years 1959–1969 for the United States and the Aggregated Data for 13 Countries

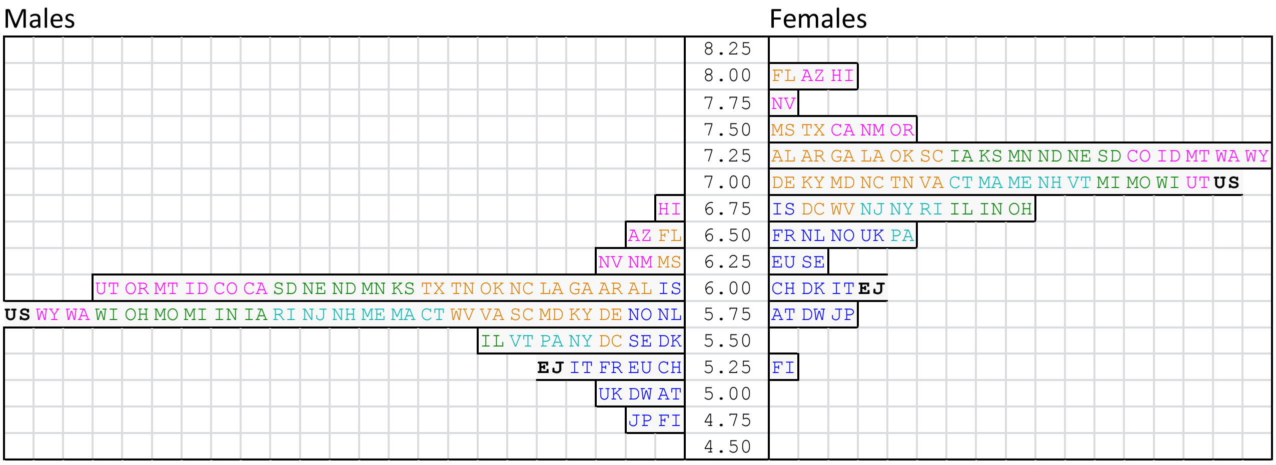


Note: See Appendix B for the complete list of countries included in the aggregated data.

Finally, there are clearly visible signs of age heaping in the U.S. data. The death rates at ages 80, 90 and 100 in males and females are elevated, and the death rates at ages 81, 91 and 101 are lowered. This pattern indicates that too many deaths were reported at ages 80, 90 and 100 at the expense of underreporting deaths at adjacent ages. Age heaping does not exist in the pooled data for the 13 comparison countries. Instead, death rates increase gradually until very high ages, when estimates start to become affected by random noise. All empirical evidence in Figure 3 suggests that direct estimates of U.S. mortality in the 1960s are severely affected by age misreporting and cannot be accepted at face value. Coale and Kisker (1990), who arrived at the same conclusion, observed similar age patterns of mortality—i.e., leveling off and declines—in directly computed death rates for whites and nonwhites at the time of the 1980 census.

Figure 4 shows that distributions of e_{80} by U.S. states and by the 13 comparison countries barely overlap. In general, life expectancy at age 80 is higher in all U.S. states than for the comparison countries, indicating that no U.S. state had good data for this period.

Fig. 4. Distribution of Life Expectancy at Age 80 in the Period 1959–1969

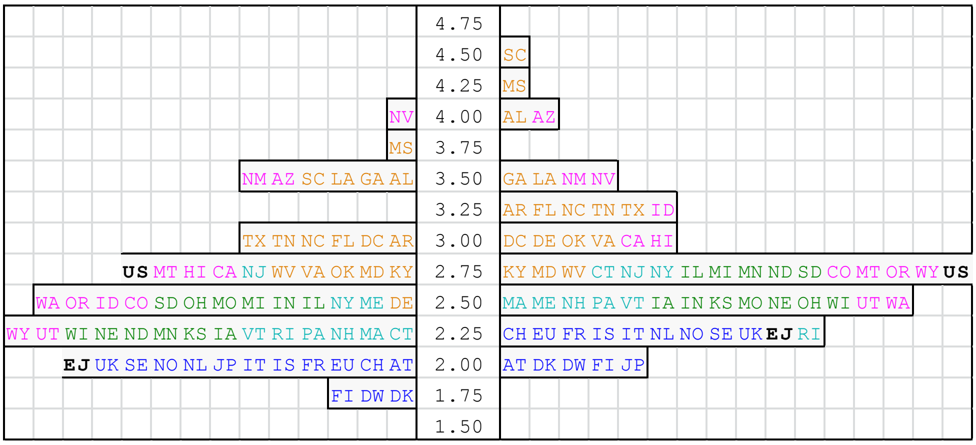


Note: Northeastern U.S. states are shown in cyan, Midwestern U.S. states in green, Southern U.S. states in brown, Western U.S. states in magenta, and the 13 high-longevity countries selected for comparison in blue. The United States and the 13-country-aggregate are shown in bold font. The list of abbreviations is given in Appendix B.

None of the U.S. states had estimates of e_{80} that were close to that of the 13-country aggregate (see U.S. state markers and EJ marker in Fig. 4). In only a few U.S. states (Illinois, Vermont, Pennsylvania, and New York), e_{80} for males was estimated to be close to that of Sweden. We also examined male death rates in Pennsylvania more closely, because for females in Pennsylvania, e_{80} was also close to that in other European countries (e.g., France and United Kingdom), yet higher than in Sweden. The examination reveals a crossover between death rates in Pennsylvania and Sweden: Death rates in Pennsylvania were higher than in Sweden until approximately age 85, and then drop below Swedish levels after this age. Estimated life expectancy at age 80 was lower in Pennsylvania than in Sweden (5.6 vs. 5.7 years), whereas e_{95} was estimated to be significantly higher (2.5 vs. 2.0 years). This pattern indicates that old-age mortality in Pennsylvania is affected by age misreporting in a similar way to the rest of the U.S. states.

At age 95, life expectancy in any U.S. state is consistently higher than in any of the 13 high-longevity countries (Fig. 5). The small overlap between distributions of e_{80} observed in Figure 4 virtually disappears in Figure 5. An exception is females of Rhode Island, where e_{95} was comparable with levels observed in the European countries. Implausibly high levels of e_{95} (i.e., more than 4 years, or approximately double the level in high-longevity countries) were observed in the female populations of Arizona, Alabama, Missouri and South Carolina, and among males of Nevada. For comparison, the highest life expectancy at age 95 in the world, in Japanese females in 2000–2011, was only 3.6 years. For comparison, the highest life expectancy at age 95 in the world—in Japanese females in 2000–2011—was only 3.6 years. Figure 5 lends further evidence that all U.S. states were severely affected by age misreporting in the 1960s.

Fig. 5. Distribution of Life Expectancy at Age 95 in the Period 1959–1969



Note: Northeastern U.S. states are shown in cyan, Midwestern U.S. states in green, Southern U.S. states in brown, Western U.S. states in magenta, and the 13 high-longevity countries selected for comparison in blue. The United States and the 13-country-aggregate are shown in bold font. The list of abbreviations is given in Appendix B.

In sum, unless there are strong reasons to believe that U.S. mortality at advanced ages in the 1960s was truly significantly lower (see, e.g., Himes, 1994; Manton and Vaupel, 1995) than that observed in other low-mortality populations, direct mortality estimates for U.S. states and the United States as a whole suggests that the almost-extinct cohort method did not produce useful mortality estimates for this period. Even if the method relies exclusively on death data, the age misreporting that appears to exist in death certificate data rendered the U.S. estimates useless. A decline in death rates after approximately age 97 (illustrated in Figure 3) is perhaps the clearest depiction of age misreporting. The amount of age misreporting in the U.S. data during this period must be more severe than the amount of age misreporting used in our simulation study, as the simulated death rates were still increasing at all ages, although biased downward and at slower rates. The quality of data in Southern U.S. states was found to be worse than in Northeastern and Midwestern U.S. states—a pattern that is roughly with phasing in of the birth registration system in the United States. According to Hetzel (1997), U.S. states in the Northeast were generally admitted first to the birth registration area, around 1915, and states in the South were generally admitted last, about 10 to 15 years later, in the late 1920s.

Mortality Changes Between 1959–1969 and 2000–2011

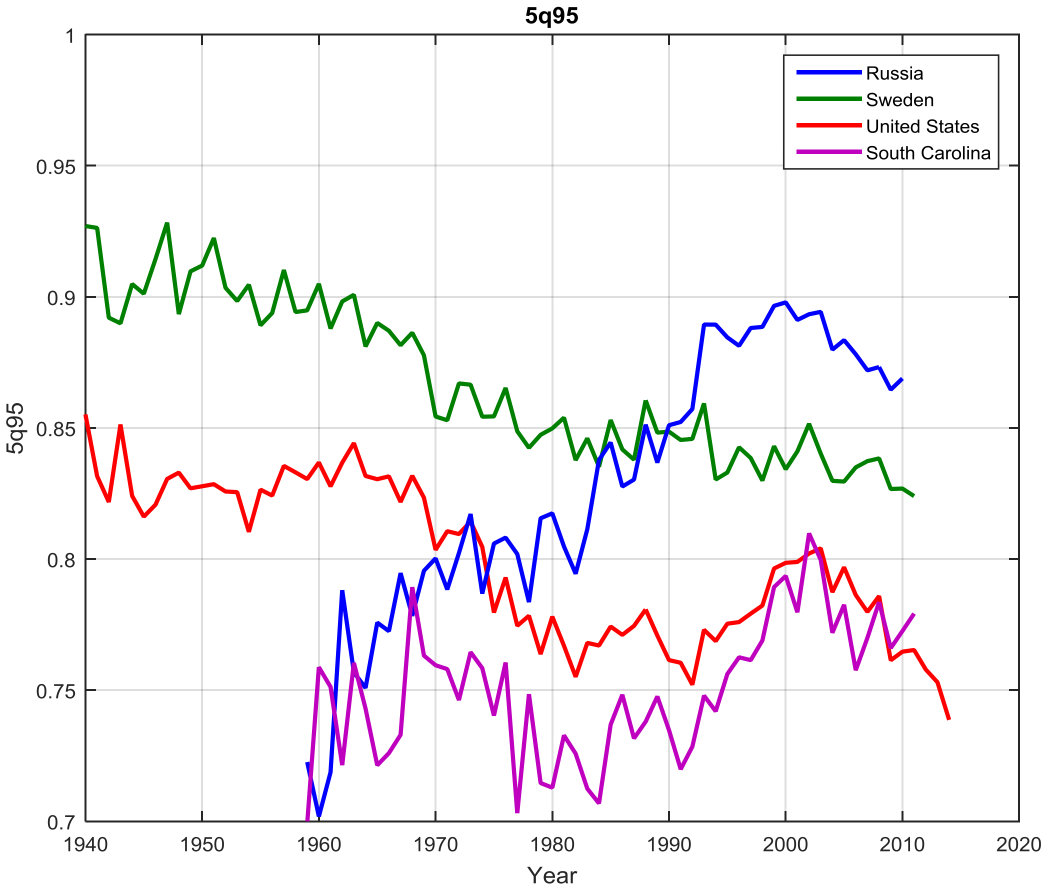
As the cohorts born in years when birth registration was considered complete are now entering old ages, the quality of U.S. data at advanced ages must be improving. People born in 1915, the year when the first 11 U.S. states (mostly from the Northeast) were admitted to the birth registration area (Hetzel, 1997), have reached age 80 in 1995 and age 100 in 2015. People born in 1933, the year when the last U.S. state (Texas) was admitted to the birth registration area, have reached age 80 in 2013. Although estimates for 1959–1969 are likely to be severely biased downward by age misreporting, examining changes in death rates between 1959–1969 and 2000–2011 is informative. If death rates are not changing over time, improvements in data quality will lead to temporal increases in death rates. The downward bias in mortality estimates due to age misreporting will disappear over time, and death rates will recover to their true levels. Visually, one will observe an increasing trend in death rates. For example, if age misreporting decreases a death rate at some age by 15% in 1959–1969, and if the data become free of age misreporting by 2000–2011, one would observe a 15% increase in the death rate at this age over time simply due to improvements in the quality of age reporting. The observed increase in mortality, of course, is spurious and does not reflect deterioration in health conditions.

Anticipated increases in death rates due to improvements in data quality are confounded, however, by reductions in death rates at advanced ages in the last decades in the high-longevity countries (Kannisto et al., 1994). If a death rate declines exactly by 15%, the 15% increase due to improvements in data quality will be offset by this decline, and visually, no changes in death rates will be observed over time. As reductions in death rates were higher at lower ages (Kannisto et al., 1994), and as age misreporting manifests itself more at higher rather than lower ages, a sensible approach would be to examine mortality trends at higher ages to detect any changes in the quality of data.

Figure 6 shows trends in female ${}_5q_{95}$ in the United States with illustrative examples from Russia, Sweden and South Carolina. The ${}_5q_{95}$ in Sweden was stable for a long time at a very high level (approximately 0.9) and only started to decline in the 1960s. Over time, the decline was persistent and gradual. This trend in Sweden is typical of a country with good data quality for the entire period and with declining mortality since the 1960s. Russia illustrates a case typical of a country where ${}_5q_{95}$ was initially increasing due to improvements in data quality. The first census in the Russian Empire took place in 1897; the people who were born shortly after the census (at the beginning of the 20th century) reached age 95 by only the late 1990s. The increasing trend in Russian ${}_5q_{95}$ appears consistent with the development of vital statistics and civil registration in the country. Reductions in death rates since 2000 may be attributed to the first improvements in mortality in this age group. It is worth noting that the level of ${}_5q_{95}$ in Russia at the peak in 2000 is equal to approximately 0.9. This level is very consistent with the long-term level of Swedish mortality; in Sweden, ${}_5q_{95}$ was mostly stable from 1860 to 1950 (at the level of 0.917) and only dropped to the level of 0.902 in the 1950s.

Observed patterns of change in ${}_5q_{95}$ over time for U.S. states could be generally classified as belonging to either of two categories: the trends similar to the observed trend in South Carolina and the trends similar to that of the United States as a whole (Fig. 6). Changes in ${}_5q_{95}$ over time in South Carolina resemble the Russian trend and could be interpreted as driven mainly by improvements in data quality until 2000 and by mortality reductions after that year. The trend observed in the United States as a whole, however, is somewhat puzzling, particularly the decline in death rates from the mid-1960s to the early 1990s. The cohorts reaching age 95 in this period included a high proportion of migrants who immigrated to the United States in the beginning of the 20th century. If age reporting was worse among migrants than among the native-born population, the observed decline in death rates could be artificially induced by the deterioration of data quality. An alternative explanation is that the observed decline is genuine, either due to the rapid progress made against mortality over this period or due to migration. Migrants could be of better health than their native-born counterparts, and the decline could be caused by a large share of robust individuals in cohorts with high proportions of migrants. The trend in ${}_5q_{95}$ for the United States appears to be inconsistent with the hypothesis of persistent and gradual improvements in data quality. Trends in ${}_5q_{95}$ similar to South Carolina are also observed in Alabama, Arizona, Louisiana, Mississippi and New Mexico for both sexes, and in Nevada and Georgia for males only. For the rest of the U.S. states, trends in ${}_5q_{95}$ are similar to that of the United States as a whole. To summarize, the observed trends in ${}_5q_{95}$ in the U.S. states only partially support the argument that data quality at ages 80 and older was improving over time.

Fig. 6. Illustrative Examples of Trends in Probability of Dying, $_5q_{95}$, from 1959 to 2011 in Russia, Sweden, United States and South Carolina, Females



Further insights into data quality can be provided by examining the rates of mortality increase with age. The rates of mortality increase with age are not commonly analyzed in assessments of data quality; however, such assessments may be both illustrative and suggestive in this regard. Age misreporting introduces downward bias into death rates in such a way that the discrepancy with true death rates widens progressively with age (Fig. 2). Depending on the amount, age misreporting may result in decelerating rates of mortality increases with age, artificial mortality plateaus, or even declines in death rates with age. Generally, at older ages, the slopes of age-specific death rates are steeper for lower-mortality levels because the magnitude of the reduction in death rates over time was higher at lower ages (Kannisto et al., 1994; Andreev, 2004). If death rates are estimated to be low with slow rates of increase with age, this may be a manifestation of age-misreporting problems.

To explore these issues, we examined mortality levels and rates of mortality increase with age separately in three age intervals: 80–89, 90–99 and 100+. We used three age intervals because the Gompertz model is not applicable for the entire age range (see Perks, 1932; Thatcher et al., 1998); however, for shorter age intervals, one could assume that death rates are increasing at a constant rate. For the present purpose, this assumption should be reasonable. Mortality was measured by age-specific probabilities of dying ($_{10}q_{80}$, $_{10}q_{90}$) and life expectancy at age 100 (e_{100}), and rates of mortality increase with age were estimated by Poisson regression (Note 3, Appendix A). In the period 1959–1969, mortality in all U.S. states was significantly lower for all three indicators in males and females than in the 13 high-longevity countries, with differences that exacerbated with advancing age (Fig. 7). Unlike mortality levels, the rates of mortality increase with age in the lowest age interval (80–89) were similar between the U.S. states and the 13 comparison countries. The average slope of death rates for the United States as a whole was an estimated 0.083 for males and 0.099 for females. For the 13-country aggregate, the estimates were only moderately higher for males (0.088) and nearly identical for females (0.099). For several U.S. states, mostly in the Deep South, the slopes were found to be implausibly low. For example, males in Mississippi, Louisiana, Delaware, South Carolina and the District of Columbia had estimated slopes less than 0.075, and females in Mississippi, Louisiana, the District of Columbia and New York had estimated slopes less than 0.090.

Slopes of death rates in the U.S. data become increasingly implausible with age. For males, the average rate of mortality increase with age in the 13-country aggregate was an estimated 0.070 at ages 90–99, significantly higher than in the United States (0.057). In 41 of the U.S. states, the rates were lower than the 13-country average. Extremely low slopes of death rates (less than 0.03) were found in Alabama, Mississippi, Georgia, New Mexico and South Carolina. After age 100, estimates of the slopes for death rates become completely implausible—i.e., death rates were *declining* in 22 out of 27 U.S. states with sufficient data after age 100. For the United States as a whole, death rates were also declining at an average rate of –0.042. In contrast, death rates in the 13 aggregated countries increased for the entire age range at an average rate of 0.022 (Fig. 7).

For females at ages 90 and older, differences in mortality levels and rates of mortality increase with age between U.S. states and the 13 high-longevity countries are similar to those of males. The slope of death rates at ages 90–99 in the United States was an estimated 0.069, slightly lower than estimates for the 13-country aggregate (0.074). Over age 100, similar to males, death rates for U.S. females were found to be declining at a rate of –0.027, whereas death rates continued to increase at a rate of 0.033 for the 13 high-longevity countries. For individual U.S. states, the declines in female death rates were observed in 20 out of 36 states with sufficient data. The observed declines in female death rates after age 100 indicated that the estimates for U.S. females are as implausible as those for U.S. males. The lower rate of decline for females (–0.027) than for males (–0.042) suggests that data for U.S. females are less affected by age misreporting than the data for males. The levels of mortality and rates of mortality increase with age presented in Figure 7 further supports the finding that U.S. data at older ages are severely affected by age misreporting in the 1960s and that age misreporting is observed from data in all U.S. states (with some variations).

From the mid-1960s to the mid-2000s, levels of mortality and rates of mortality increase with age have changed dramatically (Fig. 7). Probabilities of dying, ${}_{10}q_{80}$, declined for all countries and for all U.S. states. For males in the 13-country aggregate, ${}_{10}q_{80}$ is very similar to the level of mortality in U.S. males (0.660 vs. 0.656). Female mortality in the 13 aggregated countries also is similar to levels of mortality in U.S. females (0.514 vs. 0.548). The slopes of death rates at ages 80–89 have increased in virtually all U.S. states and in all countries included in the comparison. The average slopes for the 13-country aggregate are now 0.108 for males and 0.132 for females. For the United States, these estimates are 0.105 and 0.116, respectively. For the individual U.S. states, all estimated slopes are higher than 0.090 for males and higher than 0.109 for females.

Observed changes in mortality at ages 90-99, ${}_{10}q_{90}$, between 1959–1969 and 2000–2011 are principally different from those observed at ages 80–89. The probability of dying at ages 90–100 has actually *increased* in eight U.S. states for females and in 16 U.S. states for males.[2] For the remaining U.S. states and for all 12 European countries and Japan, ${}_{10}q_{90}$ has declined similarly to ${}_{10}q_{80}$. The slope of death rates in this age group has increased in all U.S. states and in all 13 countries. The greatest increases took place in Southern U.S. states. For example, the rate of mortality increase with age is now close to 0.097 for females in South Carolina, whereas in the mid-1960s, it was only 0.011 (nearly a 10-fold increase). Figure 7 shows that the levels and slopes of death rates are now approximately comparable between the U.S. states and the comparison countries.

[2] Females: Alabama, Arizona, Georgia, Louisiana, Mississippi, Nevada, Oklahoma, South Carolina. Males: the same states as for females plus Arkansas, Florida, Kentucky, Montana, New Mexico, North Carolina, Tennessee, Texas.

Fig. 7. Changes in $_{10}q_{80}$, $_{10}q_{90}$ and e_{100+} and Corresponding Rates of Mortality Increase with Age Between 1959–1969 and 2000–2011

Males

1959–1969 **2000–2011**

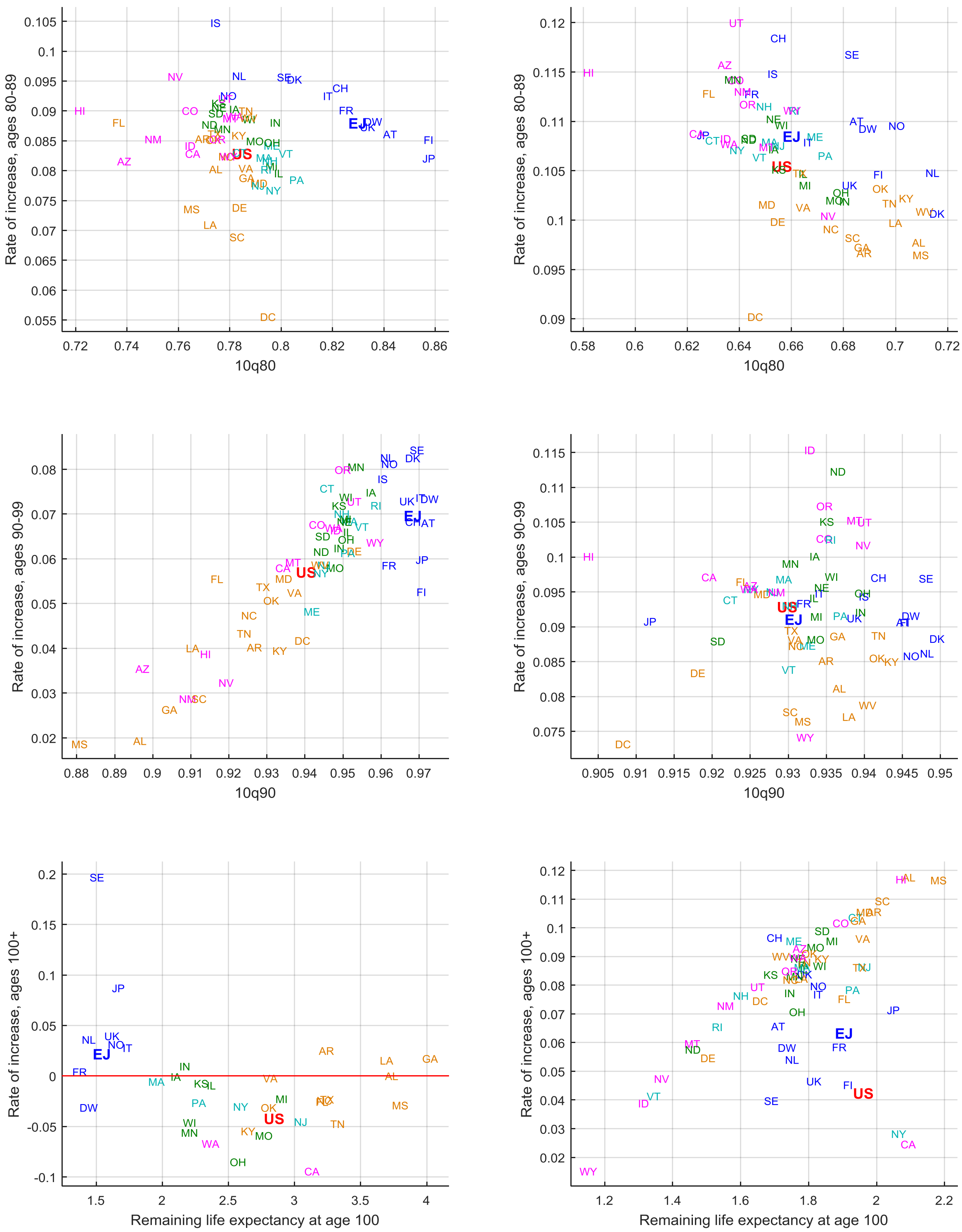
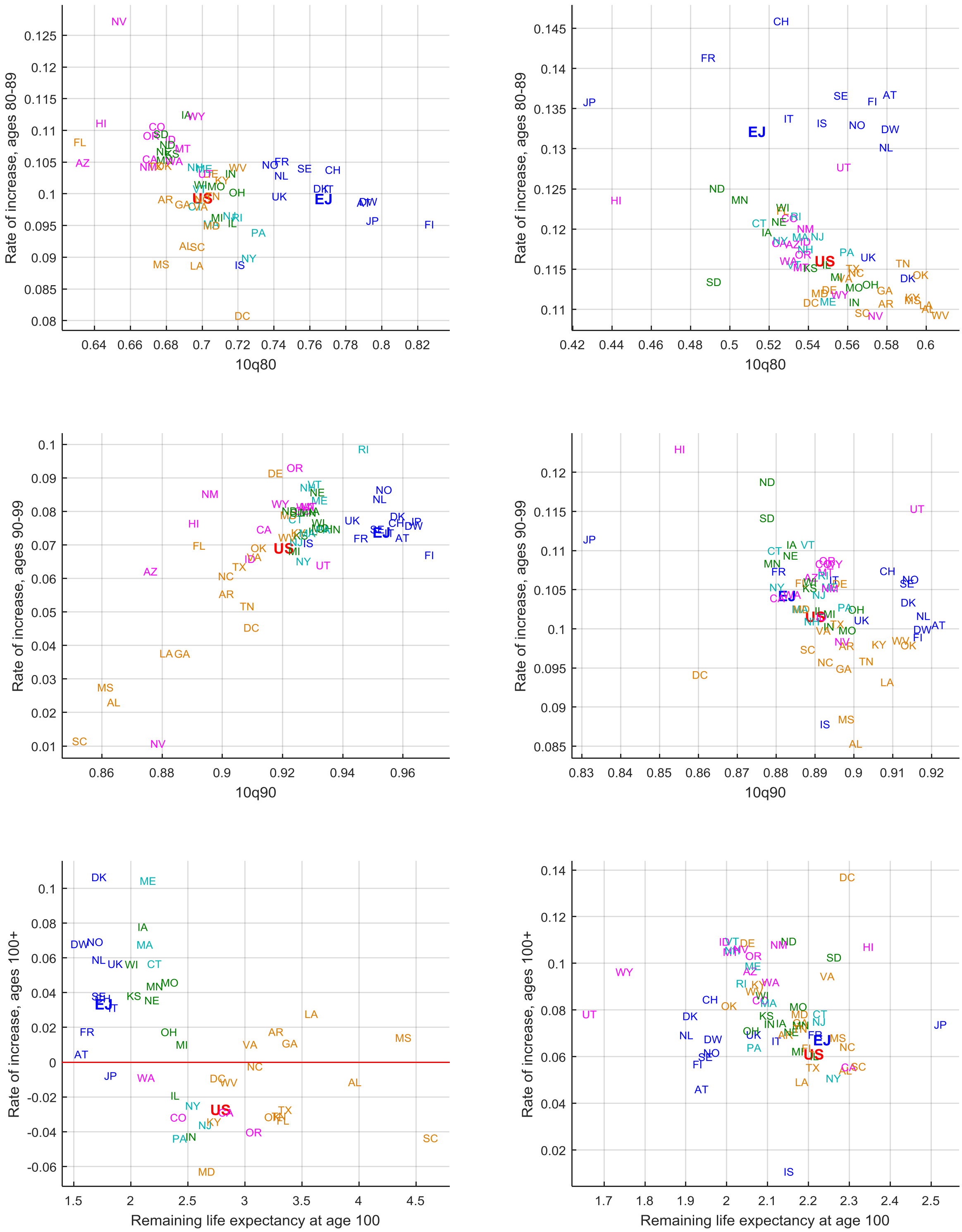


Fig. 7. (cont.)

Females

1959–1969 **2000–2011**



Note: Only the estimates based on 100 or more deaths are included. See Appendix B for the complete list of abbreviations.

Changes after age 100 have been even more dramatic. In all U.S. states, remaining life expectancy at age 100 has declined. Conversely, e_{100} has increased in all 13 comparison countries. The highest reductions, and similar to ages 90–99, took place in Southern U.S. states. For example, e_{100} declined from 4.63 to 2.32 over this period in South Carolina. The slopes of death rates at ages 100 and older have increased in all U.S. states and, for the most part, in the 13 comparison countries. In all U.S. states, and for both sexes, the rates of mortality increase with age are now positive; there are no observed declines in death rates after age 100 in the period 2000–2011 (unlike in the 1959–1969 period). Death rates after age 100 are now increasing in the United States at an average rate of 0.042 for males and 0.061 for females. Although somewhat higher, the rates are similar in the 13-country aggregate for males and females (0.063 and 0.067, respectively). At ages 90–99, the levels and rates of mortality increase with age are comparable between the U.S. states and the 13 comparison countries.

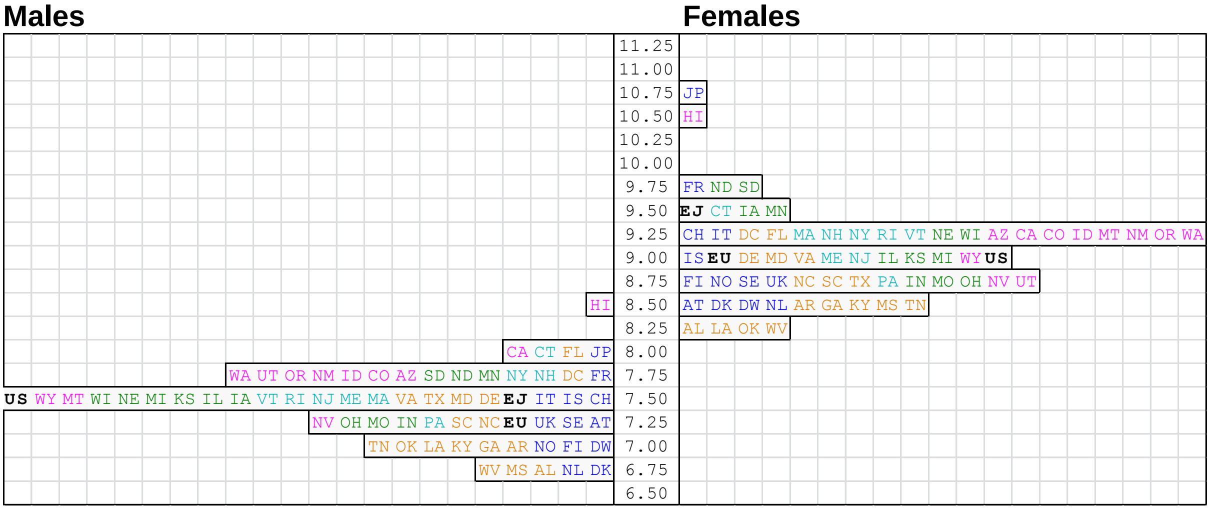
Our analysis of changes in mortality indicators over the period 1959–2011 provides indirect but convincing evidence that the quality of mortality data at older ages in the United States has increased dramatically. Even at very old ages (90 and older) in the latest years 2000–2011, the direct estimates of mortality produced by the almost-extinct cohort method did not exhibit idiosyncratic patterns commonly attributed to age misreporting. Our analysis could not identify a single U.S. state where data could be considered of good quality for the entire period of 1959–2011. By a varying degree, all U.S. states are affected by age misreporting in the 1960s—with Southern U.S. states and male populations having more data problems than others.

Mortality Levels in the Period 2000–2011

For the period 2000–2011, direct estimates of mortality and the mortality differences between the United States and the 13 comparison countries should be more accurate (i.e., reflective of reality), due to the observed improvements in quality of U.S. data. Distributions of life expectancy at age 80 for the U.S. states and the 13 comparison countries are provided in Figure 8. Because the survival of Japanese females is exceptionally high, we also added an aggregate (labeled EU) that includes only pooled data for the 12 European countries. For males, the U.S. e_{80} (7.66 years) is higher than the European average (7.41 years), and only marginally higher if Japan is included in the pooled data (7.62 years). Male mortality is found to be exceptionally low in Hawaii, with an estimated e_{80} equal to 8.70, surpassing the Japanese level (8.11 years) by nearly 0.6 years. States with the lowest mortality include California (8.11 years), Florida (8.04 years) and Connecticut (8.00 years), with e_{80} now over 8 years and as high as in Japan. No European country has reached that level—with France as the second-highest-longevity country (e_{80} = 7.80 years). Compared with the European average, 36 out of 50 U.S. states now have lower mortality. The lowest e_{80} values are observed in Denmark (6.84 years) and Netherlands (6.89 years), followed by West Virginia (6.93 years), Alabama (6.97 years) and Mississippi (6.97 years). Variations in e_{80} across U.S. states are only marginally different from variations in e_{80} across the 13 comparison countries.[3]

[3] The *p*-value of the two-sample Kolmogorov-Smirnov test is equal to 0.04, which is significant at the 0.05 level.

Fig. 8. Distribution of Life Expectancy at Age 80 in U.S. States and in the 13 High-Longevity Countries, 2000–2011



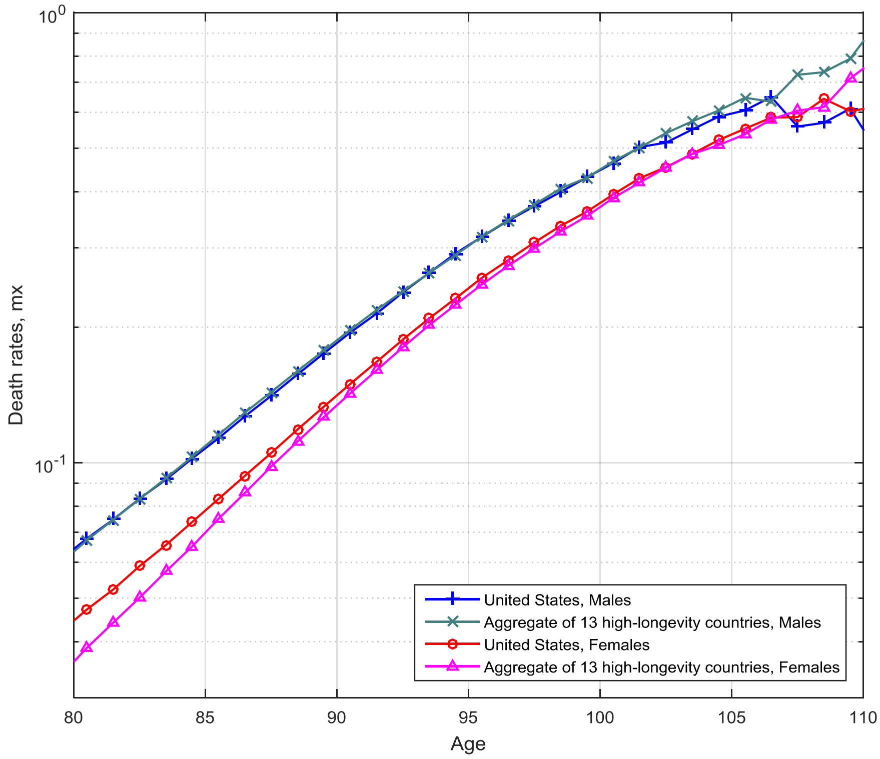
Note: Northeastern U.S. states are shown in cyan, Midwestern U.S. states in green, Southern U.S. states in brown, Western U.S. states in magenta, and the 13 high-longevity countries selected for comparison in blue. The United States and the 13-country-aggregate are shown in bold font. The list of abbreviations is given in Appendix B.

Survival beyond age 80 is significantly higher for females than for males in the U.S. states and for the 13 high-longevity countries. Distributions of e_{80} for males and females virtually do not overlap (Fig. 8), with the exception of Hawaiian males. On average, females can expect to live 1.5 years longer after reaching age 80 in the United States, 1.7 years longer in the 12 European countries, and 2.9 years longer in Japan. Japan and Hawaii enjoy outstandingly high survival at the oldest-old ages (10.97 and 10.65 years, respectively). In the rest of the female populations, e_{80} is below the Japanese level by more than one year. High life expectancies at age 80 ($e_{80} > 9.5$ years) were also found in France (9.96 years), North Dakota (9.93 years), South Dakota (9.88 years), Minnesota (9.74 years), Connecticut (9.63 years) and Iowa (9.55 years). With the exception of Connecticut, all U.S. states with high levels of female longevity are from the Midwest. Alternatively, relatively high rates of mortality ($e_{80} < 8.5$ years) prevail in Southern states, such as West Virginia (8.34 years), Louisiana (8.43 years), Alabama (8.45 years) and Oklahoma (8.45 years). The overall U.S. e_{80} (9.16 years) is slightly higher than the European average (9.12 years), and the death rates in 27 U.S. states (out of 50 states analyzed) are lower than the average European level. Applying the two-sample Kolmogorov-Smirnov test to distributions of e_{80} in U.S. states and the 13 high-longevity countries indicates no significant differences (p-value = 0.2).[4]

Figure 9 compares age schedules of death rates in the United States and in the aggregated data for the 13 high-longevity countries. For males, the two schedules of age-specific death rates are virtually indistinguishable at almost all ages. For example, the U.S. e_{95} was an estimated 2.64 years, and the 13-country aggregate was an estimated 2.62 years. For females, the approximately 0.5-year disadvantage in survival is largely attributed to the higher death rates of U.S. females at ages 80–84. In this age group, U.S. death rates are about 17% higher on average. However, the U.S. survival disadvantage disappears with increasing age as the two mortality schedules converge.

[4] Excluding Japan and Hawaii produces p-value = 0.13.

Fig. 9. The Force of Mortality by Age for the United States and the Aggregated Data for 13 Countries, 2000–2011

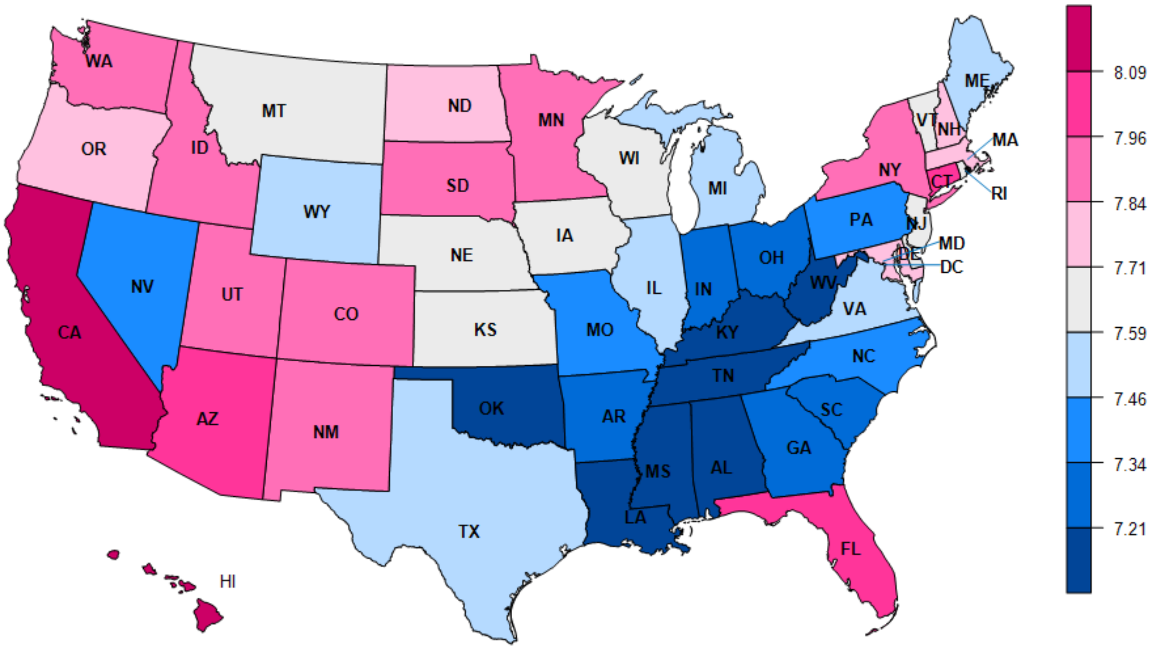


Note: See Appendix B for the complete list of countries included in the aggregated data

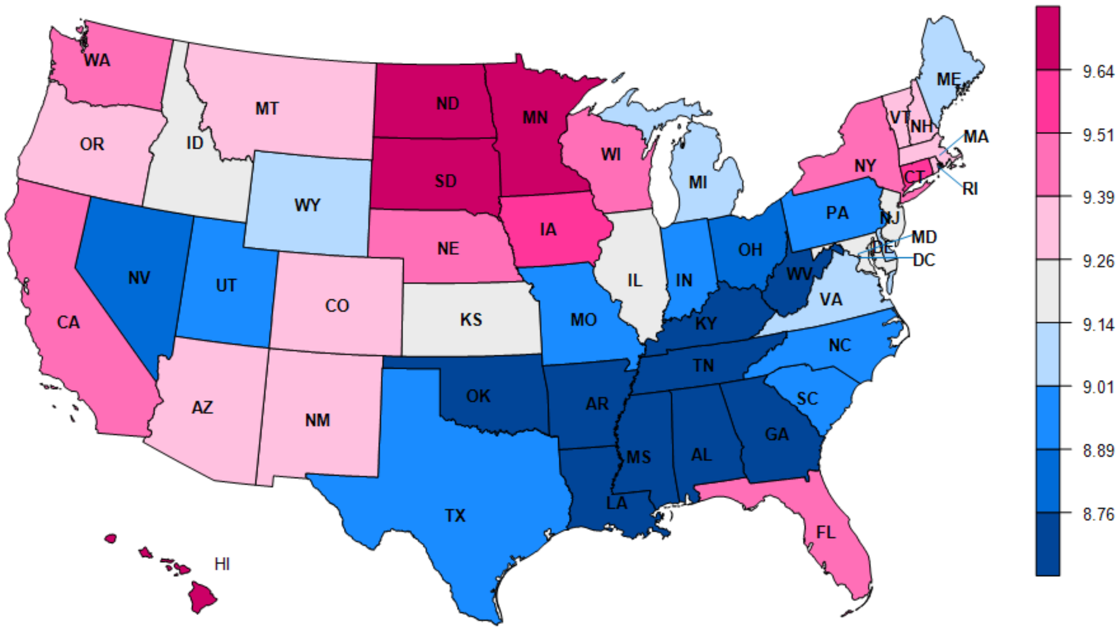
Figure 10 illustrates geographical variations in e_{80} across U.S. states. The scales for the levels presented in the maps were selected to be equidistant, with the initial levels and steps varying by sex. The scales were chosen so that the median of e_{80} for males (7.65) and females (9.20) falls within the middle (i.e., grey area). The blue hues depict U.S. states with e_{80} less than the median U.S. level. The lower e_{80} is as compared to the median, the darker is the blue hue for that state. Similarly, the magenta hues depict U.S. states with e_{80} higher than the median U.S. level; the higher e_{80} is to the median, the darker the magenta hue for that state. U.S. states with the highest mortality rates are depicted in the darkest-blue hues, and U.S. States with the lowest mortality rates are depicted in the darkest-magenta hues.

Fig. 10. Life Expectancy at Age 80, by U.S. State and by Sex, 2000–2011

Males



Females



The maps clearly illustrate a survival disadvantage in Southern U.S. states compared with other U.S. states. Particularly disadvantaged are the states located in the Deep South of the United States (Alabama, Louisiana and Mississippi), with areas of high mortality extending northeastward to Kentucky, Tennessee and West Virginia. Higher mortality rates are also seen in Arkansas and Georgia (for females) and in Oklahoma (for both sexes). Among Southern U.S. states, only Florida exhibits mortality rates that are lower than the median U.S. level. The U.S. state with the lowest level of mortality is Hawaii. Mortality is also generally lower in Western, Midwestern and Northeastern states (with the exception of Nevada and Maine). For males, California exhibits the highest level of longevity; for females, the highest levels of longevity are concentrated in Midwestern states, including Minnesota, North Dakota and South Dakota.

Discussion and Conclusions

Our analysis explored mortality levels and trends at older ages in the United States. We derived mortality estimates using the almost-extinct cohort method, an approach that has been used extensively over the past two decades for the estimation of death rates at advanced ages in high-longevity countries (Vincent, 1951; Kannisto, 1988; Thatcher et al., 2002; Andreev et al., 2003; Wilmoth et al., 2007). The reasons for selecting this method are twofold. First, it is widely believed that this method produces more reliable estimates of mortality at old ages because it uses only death registration data for both the numerator and denominator (Gallop, 2001; Rosenwaike, 1981).[5] Population counts at advanced ages that are enumerated from censuses are often highly inflated because of net age overstatements. Incentives to exaggerate one's age in self-reported population data largely stem from the prestige associated with old age in many cultures (Medvedev, 1974; Hendricks and Hendricks, 1977; Mazess and Forman, 1979; Kannisto, 1988). However, as noted by Palloni et al. (2016) based on their linkage of 2002 census data in Costa Rica with national voter registration data, the net result of age misreporting may be overstatements of age, although reported ages may be both under- and overstatements. Age at death, in contrast, is not self-reported, and there are fewer incentives to misreport age at death intentionally. The second reason for using the almost-extinct cohort method was the lack of detailed population data for all ages. National Statistical Offices, including the U.S. Census Bureau, do not publish detailed information about the composition of populations by age because of concerns about data quality and thus limit analyses of death rates to open-ended age groups. The almost-extinct cohort method is able to produce population estimates by single years of age to the highest age recorded in death registration data.

The almost-extinct cohort method relies on three main assumptions: that death registration is complete, migration can be ignored, and the reporting of age at death is accurate. For high-longevity countries with long-running civil registration systems, it is reasonable to assume that the first two assumptions are satisfied for ages 80 and older. The third is more questionable. Age at death in the United States is often reported by the next of kin, without formal verification with birth records (Bennett and Olshansky, 1996). Death certificate data published by the NCHS do not include information regarding data validation procedures. Therefore, we simply do not know whether age at death was provided by next of kin, was verified against a birth certificate, or was verified against the National Death Index database. Any misreporting of age at death will obviously affect mortality estimates produced by the almost-extinct cohort method. One would expect that death rates computed from death records with misreported ages will be biased downward (relative to true death rates), and the discrepancy with true data would widen progressively with age. In other words, age misreporting affects the slope of the force of mortality. Depending on the extent of age misreporting, one would expect a slowdown in the rate of mortality increase with age (a subtle and hard-to-detect problem), an observed plateau in mortality rates, or even declines in death rates with advancing age. This suggests that an examination of how rates of mortality increase with age may be useful for assessing the reliability of mortality estimates.

For example, if rates of mortality increase in a population are too low compared with rates observed in other countries/populations, it is likely that the mortality estimates are affected by age

[5] As discussed above only for non-extinct cohorts, the population at risk is a mixture of survivor estimates and the population produced by summing deaths

misreporting and, consequently, biased downward. The slope of the force of mortality in period life tables is not, however, a fixed quantity. An examination of the slopes of death rates should be done for comparable levels of overall mortality, because improvements in mortality over the last several decades have favored younger ages and tapered off with increasing age—leading to an increasing slope of death rates with diminishing levels of mortality at oldest-old ages (Kannisto et al., 1994; Andreev, 2004). The analyses conducted here provide further support to this observation. From 1959–1969 to 2000–2011, the age-specific probabilities of dying from ages 80 to 90 in the aggregated 13 countries declined from 0.768 to 0.514, while the average rate of mortality increase with age rose from 0.099 to 0.132. At ages 90–99 in the same period, ${}_{10}q_{90}$ declined from 0.953 to 0.883, and the slope of the death rates increased from 0.074 to 0.104. Similarly, at ages 100 and older, e_{100} increased from 1.77 to 2.2,3 and the slope increased from 0.033 to 0.067 (Fig. 7).

For the period 1959–1969, the almost-extinct cohort method produced estimates for the U.S. states that were implausible in terms of both the levels of mortality and the rates of mortality increase with age (Fig. 7). Based on this, we conclude that the almost-extinct cohort method fails to produce useful mortality estimates at old ages for U.S. states in this period. The estimates we obtained were consistent with the estimates one would expect if the data on deaths were affected by serious age misreporting. This also is consistent with the earlier work of Coale and Kisker (1990) and Rosenwaike (1981). An analysis of estimates for individual U.S. states did not reveal a particular state with good-quality data; rather, the direct estimates of mortality do not appear credible in any U.S. state. Overall, Southern U.S. states appear to have more data problems than Northeastern and Midwestern U.S. states, a finding that is generally consistent with the development of the birth registration system in the United States (Hetzel, 1997). Moreover, the quality of data for males was inferior to that of females. Our conclusion is strongly supported by the fact that death rates for the United States as a whole appear to decline at about age 100 (Fig. 3)—a phenomenon nearly universally attributed to severe age misreporting. There is virtually no evidence in the literature on human mortality to suggest that death rates decline above a certain age. Indeed, studies by Bayo and Faber (1983) and Kestenbaum (1992) conclude that U.S. mortality increases continuously with age if the estimates are based on death records with extensive age validation.

The idiosyncratic patterns commonly attributed to age misreporting in U.S. death rates observed in the 1960s largely disappear in the last period analyzed (2000–2011). Instead, U.S. death rates continue to increase for all ages (Fig. 3), and the rates of mortality increase with age (Fig. 7) in U.S. states are comparable with the rates observed in the 13 high-longevity countries used for comparison. The only exceptions are the male populations of New York and California, where the rates of increase after age 100 are approximately 3% (somewhat lower than expected, given the overall level of mortality). We interpret these results as a reflection of dramatic improvements in the quality of U.S. data over time.

Based on this premise, a comparison of mortality estimates in 2000–2011 between the U.S. states and the 13 high-longevity countries allows us to draw the following conclusions. Our main finding is that levels of mortality at advanced ages in United States in 2000–2011 are very similar to the average levels of mortality in the high-longevity countries (with the exception of Japanese females). Analyzing data using the same extinct-cohort method for a somewhat earlier period, Manton and Vaupel (1995, p. 333) concluded, “For people 80 years old or older, life expectancy is greater in the United States than it is in Sweden, France, England, or Japan.” Our analysis does not support this conclusion. Instead, we find that past U.S. data appear to be affected by age misreporting and thus mortality estimates are biased downward. Therefore, the findings of Manton and Vaupel (1995) may be simply an artifact of age misreporting.

Examining geographic variations in life expectancy at age 80 in the United States reveals higher mortality in Southern U.S. states (Fig. 10) that extend northeastward into Appalachia—a pattern roughly consistent with that reported by Ezzati *et al.* (2008). With some exceptions, mortality is generally lower in Western and Northeastern U.S. states. We also found outstandingly high rates of survival in Hawaii—for males, it was the highest observed among all U.S. states and 13 comparison countries. For females, it was the second highest observed, second only to Japan. Especially low death rates for males also were observed in California, whereas for females, and somewhat unexpectedly, we observed especially high longevity in Midwestern U.S. states such as Minnesota, North Dakota and South Dakota. Among Southern U.S. states, only Florida had an e_{80} that was higher than the U.S.

median level. The disparities we found in old-age mortality among U.S. states were generally comparable with the disparities found in the 13 high-longevity countries. Variations in e_{80} among U.S. states were of the same magnitude as variations in e_{80} across the 13 high-longevity countries.

Several important issues could not be addressed in our analysis. Foremost, it was not possible to directly estimate the improvements in U.S. mortality that occurred at advanced ages over the last few decades. Unreliable estimates of death rates from the past that were produced by the almost-extinct cohort method did not provide a reliable starting level of mortality. Because past death rates are likely to be biased downward, directly computed rates of mortality improvement are biased downward as well. As a result, age misreporting conceals the true progress made in the reduction of death rates at advanced ages. The observed rates of mortality improvement may be interpreted as a lower bound—with the expectation that true improvements in mortality are presumably higher. Estimating adjusted rates of mortality improvement would require producing adjusted mortality estimates for the past. To date, there is no established methodology for performing this task. Further investigation into this problem would be an important research priority considering that none of the countries included in the major databases (HMD, KTDB) currently adjust for age misreporting or provide reliable indicators of data quality. For some countries (e.g., Sweden), quality data on old-age mortality have been available for more than a century, whereas for other countries (e.g., Russia), these data may not be usable until just recently (Fig. 6). The Kannisto mortality model employed by HMD methodology (Wilmoth et al., 2007) does not produce adjusted mortality estimates at advanced ages; but instead graduates and extrapolates death rates through age 110.

Another important issue that was unaddressed in our analysis is interstate and seasonal migration at older ages. The extinct cohort method assumes that migration is 0. It may be a valid assumption for the overall population; however, it may not be valid for all U.S. states. In particular, the low mortality estimates observed in Florida warrant further analysis. Florida is a popular destination for retirees and for vacationing. It would be worth exploring whether the results we obtained are sensitive to the tabulation of deaths by place of occurrence (our analysis uses deaths tabulated by the place of residence).

Better U.S. mortality estimates, both in terms of both their quality and timeliness, would be obtained by extending the death series from 2011 through 2014 (the latest currently available). Since 2004, however, information on place of residence has been excluded by NCHS from publicly available mortality files, and their access is prohibitively expensive for most of the research community. Establishing a sustainable database for monitoring disaggregated U.S. mortality trends by state and ethnicity would be another important area of priority. Such a database would be crucial for monitoring health disparities in the United States and for making forecasts of mortality.

Acknowledgments

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Appendix A. Notes

Note 1

If X is the true age at death, and X^* is random age misreporting, we assume that the misreported age at death is $\tilde{X} = X + X^*$. We further assume that $X^* \sim N(0,3.04)$ with 90% of probability mass lying between –5 and 5. This model assumes unintentional, random or symmetric age misreporting, with no net overstatement or understatement of age.

Note 2

Suppose that age at death could be misreported to be 1 year higher, 1 year lower or reported the same as the true age at death with equal probabilities. Further assume that the number of deaths is linearly declining over the age range from x – 1 to x + 1: $D_{x-1} = D_x + \Delta$ and $D_{x+1} = D_x - \Delta$, where D_x is number of deaths at age x and Δ is the difference of deaths at adjacent ages. The misreported number of deaths at age x will then be $\tilde{D}_x = \frac{1}{3}D_{x-1} + \frac{1}{3}D_x + \frac{1}{3}D_{x+1}$. The first and the last terms are reallocations of deaths from age groups x – 1 and x + 1, and the middle term is the remaining number of deaths in the age group x (not misreported). Obviously, the number of misreported deaths is equal to the number of true deaths, $\tilde{D}_x = D_x$, as $\tilde{D}_x$ is simply an average of deaths at ages x – 1, x, x + 1. In this particular case, age misreporting is not changing the number of deaths at age x. Similarly, due to age misreporting, the population at risk at age x, N_x, is inflated by reallocation of $\frac{1}{3}D_{x-1}$ from the age group x – 1 to the age group x and deflated by reallocation of $\frac{1}{3}D_x$ below age x. Age misreporting at ages x + 1 and over has no effect on the population at risk, because none of the deaths are reallocated below age x and will still be counted in the population at risk. This mechanism of age misreporting inflates the population at risk at age x by $\frac{\Delta}{3}$: $\tilde{N}_x = N_x + \frac{1}{3}D_{x-1} - \frac{1}{3}D_x = N_x + \frac{\Delta}{3}$. Consequently, the misreported probability of dying is lower than the true one, due to the inflation of the population at risk: $\tilde{q}_x = \frac{\tilde{D}_x}{\tilde{N}_x} = \frac{D_x}{N_x + \Delta/3}$. For age intervals with increasing deaths, for example, before the mode of deaths at adult ages, the effect of such age misreporting is positive, and the misreported death rates will be higher than observed.

Note 3

Estimates of the rates of mortality increase with age were computed by fitting Poisson regression assuming that observed deaths D_x are distributed according to the Poisson distribution $D_x \sim \text{Poisson}(m_x E_x)$ and with age x modeled as a continuous independent variable, $\ln m_x = \beta_0 + \beta_1 x$, where m_x is the death rate and E_x is exposure estimate at age x. If the regression is fitted, say, to the data at ages 80–89, $e^{\hat{\beta}_1}$ is an estimate of a rate of mortality increase with age or slope of death rates over this age range.

Appendix B. Abbreviations Used for U.S. States, Countries, Statistical Areas and Aggregates

Name	Ab.	Name	Ab.	Name	Ab.	Name	Ab.
Alabama	AL	Louisiana	LA	Ohio	OH	Austria	AT
Arizona	AZ	Maine	ME	Oklahoma	OK	Denmark	DK
Arkansas	AR	Maryland	MD	Oregon	OR	England and Wales	UK
California	CA	Massachusetts	MA	Pennsylvania	PA	Finland	FI
Colorado	CO	Michigan	MI	Rhode Island	RI	France	FR
Connecticut	CT	Minnesota	MN	South Carolina	SC	Germany (West)	DW
Delaware	DE	Mississippi	MS	South Dakota	SD	Iceland	IS
District of Columbia	DC	Missouri	MO	Tennessee	TN	Italy	IT
Florida	FL	Montana	MT	Texas	TX	Japan	JP
Georgia	GA	Nebraska	NE	Utah	UT	The Netherlands	NL
Hawaii	HI	Nevada	NV	Vermont	VT	Norway	NO
Idaho	ID	New Hampshire	NH	Virginia	VA	Sweden	SE
Illinois	IL	New Jersey	NJ	Washington	WA	Switzerland	CH
Indiana	IN	New Mexico	NM	West Virginia	WV	13-country aggregate[2]	EJ
Iowa	IA	New York	NY	Wisconsin	WI	12-country aggregate[3]	EU
Kansas	KS	North Carolina	NC	Wyoming	WY		
Kentucky	KY	North Dakota	ND	United States[1]	US		

[1]Data for the United States includes Alaska.
[2]The 13-country aggregate includes pooled data for Austria, Denmark, England and Wales, Finland, France, Germany (West), Iceland, Italy, Japan, the Netherlands, Norway, Sweden and Switzerland.
[3]The 12-country aggregate excludes Japan.