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The narrowing sex gap in life expectancy:

Effects of sex differences in the age pattern of mortality

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Abstract

Using data from the Human Mortality Database for 29 high-income national populations (1751-2004), we review trends in the sex gap in *e*(0). Widening of the sex gap during most of the 1900s was largely due to slower mortality decline for males than females, which previous studies attributed to behavioural factors (e.g., smoking). More recently, the gap began to narrow in most countries, which researchers tried to explain with these same factors. However, our decomposition analysis reveals that for the majority of countries recent narrowing is primarily due to sex differences in the age pattern of mortality rather than declining sex ratios in mortality: the same rate of mortality decline produces smaller gains in *e*(0) for women than men because female deaths are less dispersed across age (i.e., survivorship is more rectangular). This study demonstrates that sex differences in the age pattern of mortality affect trends in the *e*(0) sex gap.

Keywords: Life expectancy, Mortality, Sex gap, Sex difference, Sex ratio, Mortality Decline

Introduction

Over the last 200 years, women in Europe and North America have enjoyed higher life expectancy at birth (*e*(0)) than men (Tabutin and Willems 1998). Indeed, sex differences in mortality are often wider than those between other subgroups (e.g., race/ethnicity). Yet, available data suggest that the sex gap in *e*(0) changed considerably in the past (Stolnitz 1956; United Nations 1988; Tabutin and Willems 1998). It was relatively small in the late 1800s but grew rapidly through most of the twentieth century.

Widening of the sex gap was accompanied by substantial rises in the sex ratio of age-specific death rates. Those rises are usually attributed to sex differentials in trends of behavioural and social risk factors such as cigarette smoking, heavy drinking, violence, and occupational hazards (Preston 1970; Lopez 1983; Vallin 1993; Waldron 1985, 1986).

However, around the 1980s, the sex difference in *e*(0) began to narrow in many industrialized countries in Northern Europe, Northern America, and Oceania (Meslé 2004). Waldron (1993) notes that sex ratios in mortality also stopped increasing in the 1980s. Trovato and Lalu (1996) demonstrate that between the early 1970s and the late 1980s the sex gap in *e*(0) narrowed substantially (-1.85 in Hong Kong to -0.26 years in USSR) in nine countries, including most English-speaking countries (i.e., USA, Canada, U.K., Australia) as well as several other countries (i.e., Hong Kong, Iceland, Austria, Finland, USSR). Yet, during this period, the sex gap continued to widen among Eastern European countries and, to a somewhat lesser extent, Southern European countries and Ireland. More recently, declines in the sex gap have become apparent among other countries in Western Europe, and the sex differential appears to be levelling off among several countries in Southern Europe, although the sex gap continues to increase in Japan (Meslé 2004).

Attempts to explain recent narrowing have focused on causes of death that contributed to the trend (Trovato and Lalu 1998; Pampel 2003; Elo and Drevenstedt 2005; Trovato and Heyen 2005) as well as behavioural and medical factors that would reduce sex ratios in mortality rates and thereby narrow the sex gap in $e(0)$ (Waldron 2005; Nault 1997; Valkonen and Poppel 1997; Pampel, 2002; Conti et al. 2003; Gjonca et al 2005; Preston and Wang, 2006). Those factors include increased smoking among females while prevalence declined among males (Waldron 2005), and advances in medical treatments for cardiovascular disease that may have benefited men more than women (Waldron 1995).

Factors that have different trends by sex (or have differential effects by sex) could reduce sex ratios in mortality rates and thereby, narrow the sex gap in $e(0)$. The increase in sex mortality ratios slowed down, ceased, or reversed in many countries during the last quarter of the twentieth century, which probably explains the cessation of the widening trend in the sex gap in $e(0)$. Sex mortality ratios declined recently in some countries, which could account, at least in part, for the narrowing sex gap in those populations.

However, changes in life expectancy depend on age patterns of mortality as well. Previous studies have indicated that the gain in life expectancy produced by declines in age-specific death rates tends to be smaller if deaths are more concentrated in a narrow age range—that is, the survival curve is more rectangular (Keyfitz 1985, chapter 3; Vaupel 1986; Vaupel and Canudas-Romo 2000). Because the age distribution of deaths is usually less dispersed for women than men (at least in recent decades), the sex difference in $e(0)$ may narrow as mortality declines, even if age-specific rates of mortality decline are the same for both sexes. Thus, the recent narrowing of sex gap may have resulted primarily from sex differences in the age pattern of mortality, rather than slower mortality decline for females than males. Little attention has been given to this demographic mechanism of sex gap narrowing. In this study, we explore this hypothesis using data from the Human Mortality Database (HMD, www.mortality.org) for 29 national populations with high quality mortality statistics. We begin by reviewing trends in the sex gap in $e(0)$ across time and place. The long time series (starting as early as 1751), the wide range of countries, and the application of uniform methods across countries allow us to evaluate the universality and generalizability of these trends.

We investigate the following research questions regarding the recent narrowing of the sex differential in life expectancy. First, when did a sustained narrowing of the sex gap in $e(0)$ begin, and how did the onset vary across countries? Second, to what extent and in what manner does the dispersion of deaths across age differ between males and females? Third, were the age groups that were the biggest contributors to widening the sex gap in $e(0)$ similar to those that contributed to the recent narrowing in the sex gap? Finally, to what extent did changes in the sex gap result from a sex difference in the age-specific rates of mortality decline versus sex differentials in the age pattern of mortality?

'Differential dispersion' hypothesis of sex gap narrowing: Relationship between gain in life expectancy and age pattern of mortality

Previous studies have shown that the gain in life expectancy depends on the age pattern of mortality. In particular, for a given rate of mortality decline, the gain in $e(0)$ is larger if the age pattern of mortality is more dispersed. Such a result occurs because a population that suffers more premature deaths (i.e., the survival curve is less rectangular) has more to gain by reducing mortality at younger and middle ages. Thus, if the level of mortality dispersion differs between males and females, the same age-specific rates of mortality decline for both sexes (i.e., sex mortality ratios remain unchanged) can change the sex gap in $e(0)$.

Keyfitz (1985) has shown that if all age-specific death rates decrease at the same rate *ρ(t)*, the resulting rate of absolute change in life expectancy at time *t* can be expressed by

$$\frac{de(0)}{dt} = G(t)\rho(t) \ , \tag{1}$$

where $G(t) = -\int_0^{\infty} l(x,t)\ln(l(x,t))dx > 0$, $\rho(t) = -(\partial\mu(x,t)/\partial t)/\mu(x,t)$ for any age *x*, $\mu(x,t)$ is the force of mortality at age *x* and time *t,* and *l(x)* is the proportion of those who survive from birth to age *x* in the life table at time *t*.

The corresponding relative change in life expectancy is given by $(de(0)/dt)/e(0) = H(t)\rho(t)$, where $H(t) = G(t)/e(0)$, which is called the life table entropy. *H* and *G* are used as measures of mortality variability because they tend to be smaller if deaths are concentrated in a very narrow age range (Nusselder and Mackenbach 1996; Wilmoth and Horiuchi 1999)—for example, if almost everyone survives to $e(0)$ and then dies soon thereafter. Thus, Equation (1) indicates that the gain in $e(0)$ tends to be smaller if the age distribution of deaths is less dispersed.

Pollard (1982) has demonstrated a paradox whereby the differential in $e(0)$ between two populations could *widen* even when the differences in age-specific death rates *decrease*. Keyfitz' formula suggests that this paradox is possible if the level of mortality dispersion differs between the two populations, because the relation between $e(0)$ gain ($de(0)/dt$) and mortality change (*ρ*) depends on mortality dispersion (*G*).

Using Keyfitz' equation and the Gompertz model, Vaupel and Canudas-Romo (2000) have revealed that the increase in life expectancy resulting from mortality decline tends to be smaller if adult mortality rises more steeply with age. Suppose that the risk of mortality above age *z* follows the Gompertz equation, $\mu(x,t) = a(t)e^{bx}$, where *a(t)* represents the overall level of mortality at time t, and b is the rate of relative change in mortality with age, which is assumed to remain constant over time. If mortality risk declines over time at the same rate $\rho(t)>0$ for all ages greater than z:

$$\frac{\partial \ln \mu(x,t)}{\partial t} = \frac{\partial \ln a(t)}{\partial t} = -\rho(t)\,, \tag{2}$$

the result is a vertical downward shift of logarithmic mortality curve (i.e., $\ln \mu(x,t) - \rho(t)\Delta t$ for age $x \geq z$), as illustrated in Figure 1. We could also interpret it as a horizontal shift to the right (i.e., the population survives to an older age before reaching a given level of mortality). Let $y(c,t)$ represent the age at which the risk of mortality is equal to a given level c, so that $c = a(t)e^{b \cdot y(c,t)}$, which can be rewritten as $y(c,t) = \dfrac{\ln c - \ln a(t)}{b}$; $y(c,t)$ is the inverse function of $\mu(x,t)$ where y corresponds to x and c corresponds to μ. Therefore, the rate of this horizontal shift is:

$$\frac{\partial y(c,t)}{\partial t} = -\frac{\partial \ln a(t)/\partial t}{b} = \frac{\rho(t)}{b}, \text{ for any } c \geq \mu(z,0)\,. \tag{3}$$

[INSERT FIGURE 1 ABOUT HERE]

Thus, a vertical shift of $-\rho(t)\Delta t$ in the logarithmic curve of Gompertzian mortality can also be interpreted as a horizontal shift of $\rho(t)\Delta t / b$. This implies that for a given rate of mortality decline $\rho(t)$, a steeper mortality curve (i.e., greater b) leads to a smaller horizontal shift in the mortality curve.

Figure 1 compares the shifts in Gompertz mortality for two populations, both of which experience the same relative mortality decline (indicated by the vertical arrows). Yet, Population 2 (based on French females) has a smaller shift toward older ages (indicated by the horizontal arrows) because the mortality curve is steeper than for Population 1 (based on French males).

Vaupel and Canudas-Romo (2000) have shown that the absolute change in life expectancy at age z is also approximately equal to $\rho(t)/b$. The Gompertz model fits adult mortality reasonably well (except at very old ages). For industrialized countries during recent decades, the gain in $e(0)$ is mainly due to changes in adult mortality. Moreover, among those populations, the rate of mortality decline does not vary markedly across adult ages (Wilmoth & Horiuchi 1999); thus, adult mortality changes can be approximated well by parallel vertical shifts. Consequently, in those populations, we would expect the gain in $e(0)$ produced by a given extent of mortality decline to be smaller if mortality rises more steeply with age.

The formula by Keyfitz and that by Vaupel and Canudas-Romo are consistent with each other, although the former is based on the survival curve and the latter, the mortality curve. The consistency is clear not only mathematically (Vaupel and Canudas-Romo used Keyfitz' equation) but also intuitively. A steep rise of mortality with age from a low to high death rate should concentrate the majority of deaths into a relatively narrow age range, resulting in a sharp drop-off in the survival curve.

Empirical evidence demonstrates that male-to-female mortality ratios tend to decline at older ages across various populations (Lopez 1983; United Nations 1988), suggesting that adult mortality curves tend to be steeper for females than for males. Thus, we would expect similar rates of mortality decline for men and women to result in greater gains in $e(0)$ for the former compared with the latter, thereby reducing the sex gap.

A numerical simulation demonstrates this effect. Based on the observed mortality rates for France during 1975-79, life expectancy at birth was 69.5 for males and 77.6 for females, resulting in a sex gap of 8.1 years (HMD 2006). Hypothetically, if these mortality rates declined 35% at all ages for both males and females, the resulting life expectancy would be 75.0 for males and 82.2 for females, a sex gap of 7.2 years. Thus, the sex gap would narrow by 0.9 years even if sex ratios in mortality rates remained *unchanged*.

Data and methods

Data come from the Human Mortality Database (HMD 2006) and comprise 29 national territories (counting former East and West Germany separately). These data series cover a period as long as 254 years (1751-2004: Sweden). All but eight countries have data back to at least 1950, and ten countries have more than 100 years of data (see Table 1). Because war can have a big impact on the sex ratio in mortality, we exclude the data for the periods during World War I (1914-19) and World War II (1939-45) from all of the analyses.

[INSERT TABLE 1 ABOUT HERE]

In the first part of the analysis, we examine the time trends in the sex gap and determine the first year in which each country demonstrated a sustained decline in this gap. We calculate the sex gap in life expectancy ($e_f(0) - e_m(0)$) using the HMD estimates of period $e(0)$ by country, sex and calendar year. In order to better identify the general time trends in the sex gap, we smooth the data using a five-year moving average, where the sex gap for year t is calculated based on the mean for years t-2, t-1,…t+2. This moving average gives zero weight to periods during World War I and World War II (i.e., value for 1946 is based on the average of 1946, 1947 & 1948 only. We begin by graphing the sex gap in $e(0)$ across time and by country in order to ascertain which countries tend to have larger (or smaller) absolute sex differences and how the trajectory changed over time. For each country, we define the onset of narrowing in the sex gap as the first year in which: a) both the observed and smoothed sex gap declined since the previous year, b) the smoothed sex gap fell to a level that has not since been exceeded, and c) after which, the decline was not interrupted by an increase in the (smoothed) sex gap of more than 0.25 years (relative to the previous year) or a sustained increase for three or more consecutive years.

In the second part of the analysis, we compute dispersion indicators *H* and *G* and Gompertzian slope *b* by sex for 1975-79 in order to explore the possibility that the sex gap narrowing is primarily due to a sex differential in the age pattern of mortality. If *H* and *G* are larger for males than females, while the reverse is true for *b*, the sex gap in *e*(0) is likely to narrow even if the rates of mortality decline are the same for both sexes.

In the third part of the analysis, we decompose the change in the sex gap across periods in order to ascertain: a) which age groups are the biggest contributors to the change in the sex gap, and b) to what extent the change in the sex gap results from sex differences in mortality change versus sex differences in the age pattern of mortality. The first decomposition is straightforward, based on the calculation of the life expectancy as a function of age-specific death rates. The second decomposition makes use of the fact that sex-specific mortality can be expressed as a function of the sex ratio in mortality and the geometric mean of female and male mortality:

$$\mu_f(x) = \beta(x)/\sqrt{\alpha(x)} \text{ and } \mu_m(x) = \beta(x)\cdot\sqrt{\alpha(x)} \quad , \tag{4}$$

where $\mu_f(x)$ and $\mu_m(x)$ are forces of mortality at age *x* for females and males, respectively, $\alpha(x) = \mu_m(x)/\mu_f(x)$, and $\beta(x) = \sqrt{\mu_m(x)\cdot\mu_f(x)}$. All of these functions of age vary over time, but for simplicity the subscript for time is not shown.

The above pair of simple formulations makes it possible to convert the effects of sex-specific mortality ($\mu_f(x)$ and $\mu_m(x)$ effects) on the *e*(0) sex gap into a sex ratio effect ($\alpha(x)$ effect) and an average mortality effect ($\beta(x)$ effect). The sex ratio effect results from a sex difference in the rate of mortality decline, because the sex ratio changes if and only if the rate of mortality decline differs between females and males. The average mortality effect represents the change in the $e(0)$ sex gap that would occur if both female and male death rates followed the same rate of decline observed for the geometric mean of sex-specific death rates. The average mortality effect can be interpreted as an age pattern effect, because the effect is larger when there are greater sex differences in the life table pattern (more precisely, greater sex differences in the $d(x)e(x)$ function, where $d(x) = l(x)\mu(x)$); the effect is zero if female and male life tables are identical. Mathematical formulation and statistical methodology of this decomposition analysis are described in more detail in the Appendix.

For the decomposition, we start with period death counts and exposure estimates by country, calendar year, sex, and age. In order to reduce random fluctuations of death rate for each age-time combination, data for each population (separately by sex) are pooled into five-year time intervals and five-year age groups except for the first and last age intervals (0, 1-4, 5-9,…85-89, 90+). Based on these 5x5 data, we calculate period death rates and life expectancy at birth ($e(0)$) using standard methods (Wilmoth et al. 2005). We then decompose the change in the sex gap in $e(0)$ for each pair of successive five-year time intervals. For presentation purposes, we aggregated the effects for the 20 age groups into ages 0-19, 20-39, 40-59, 60-79, 80+. We further aggregated the effects for successive five-year time intervals into the period during which the sex gap was widening in all countries (1950-54 to 1975-79) and the period during which the sex gap began to narrow (1975-79 to 2000-04).

Results

Sex gap in life expectancy across time and place

Sex differences in $e(0)$ across time for each country are presented in Figure 2. Countries are grouped that share similar trajectories. In Sweden—the only country for which we have data prior to the 1830s—the sex difference in $e(0)$ was around three years during the late 1700s, but began to increase in the early 1800s reaching more than 4.5 years by 1830.

[INSERT FIGURE 2 ABOUT HERE]

Starting around the mid-1800s, data became available for England & Wales, Denmark, Iceland, the Netherlands, and Norway. At this time, the sex gap in $e(0)$ was around two to three years with the exceptions of Iceland and Sweden, where the female advantage topped four years.

Data series for Italy, Finland, and Switzerland begin in the 1870s. The sex gap remained between two and four years for seven of the nine countries observed. The two exceptions were: Iceland, with a gap greater than four years, and Italy, with a much smaller sex gap (≈0.5 years in the 1870s). Throughout the second half of the nineteenth century the sex difference grew in England & Wales (from ≈2.0 years in the 1850s to nearly 4.0 years by 1900), whereas it tended to decline in Sweden (from more than 4.0 years in the 1850s to ≈2.5 years in 1900).

By 1900-04, the sex gap remained by far the lowest in Italy (0.4 years) and the highest in Iceland (4.4 years), but ranged from 2.6 to 3.8 years among the other eight countries with available data (Table 1). After 1910, the sex gap continued to grow in England & Wales, Finland, and France, and in the 1920s, began to increase in Italy and Spain as well. On the other hand, after increasing somewhat in the late 1800s, the sex gap declined through the 1920s in Denmark and the Netherlands. During 1925-29, the sex difference in $e(0)$ was ranged from 1.5 years in the Netherlands to 5.0 years in Finland among 13 countries.

Before and after World War II, there were several notable changes in the sex gap. For Spain, the sharp pre-war rise in the sex gap reflects excess mortality among males due to the Spanish Civil War (1936-39); after the war, $e(0)$ for males returned to the historical trend with a corresponding decline in the sex gap from the peak levels attained at the end of the Spanish Civil

War. As for France, the sex gap increased during the pre-war period because males did not keep pace with females in terms of $e(0)$ gains, but males "caught up" in the post-war period (so that the gap returned to the historical trend line). In contrast, the sex gap in Finland increased markedly between 1938 and 1946 because females made greater gains in $e(0)$ than males—perhaps due to the aftermath of World War II and the Wars with the Soviet Union (1939-40, 1941-44); the gap narrowed again during the post-war period as males began to catch up with females.

By the 1950s, the sex gap in $e(0)$ was growing rapidly in all 21 countries observed and continued to increase through the 1970s. During 1950-54, the sex gap remained the lowest in the Netherlands (2.5 years) and highest in Finland (6.6 years). By 1975-79, countries of the former USSR (Latvia, Lithuania, Russia, and the Ukraine) exhibited the largest sex differences (9.0 to 11.2 years) among the 29 countries observed, while others ranged from 5.1 (Bulgaria) to 8.8 (Finland) years.

In the last couple decades of the twentieth century, the sex gap in $e(0)$ began a steady decline in most countries, although the timing of this reversal varied across countries. By 2000-04, six countries exhibited a sex gap of less than five years.

Onset of sustained decline in the sex gap in life expectancy

The first year of a sustained decline in the sex gap in $e(0)$ for each country is presented in the last column of Table 1. The narrowing trend started in England & Wales in 1972, followed in the later 1970s and early 1980s by the other English-speaking countries (United States, Canada, Australia, and New Zealand), Finland, former West Germany, and Sweden. The rest of the Scandinavian countries (Denmark, Iceland, and Norway) as well as Austria and the Netherlands followed suit in the remainder of the 1980s. Other Western European countries (Italy, Switzerland, France, Belgium, Portugal, and Spain) did not demonstrate a substantial narrowing of the sex gap until the 1990s.

The Eastern European countries (Czech Republic, former East Germany, Hungary, Latvia, Bulgaria. and Slovakia) were also late to exhibit a substantial decline in the sex gap, and in fact, Russia, Lithuania, and the Ukraine have yet to evidence such a decline. With the exception of East Germany, all of these Eastern bloc countries experienced sustained periods of stagnation or even declines in life expectancy after 1960; in 2000, these eight countries had the lowest levels of $e(0)$ among the 29 countries in this analysis. Males were particularly hard hit by the increases in mortality, which probably accounts for the continued widening of the sex gap. Among East Germans, $e(0)$ declined briefly after 1989 for males, but otherwise followed an upward trend like all other countries in this analysis. Japan is unique in the respect that despite impressive gains in $e(0)$—in fact, Japanese females enjoyed the highest and males the second highest $e(0)$ among these 29 countries—the sex gap also continues to widen with no evidence of narrowing.

Sex differences in the age pattern of mortality: life table entropy and Gompertzian slope

Table 2 shows sex-specific estimates of life table entropy (Keyfitz, 1985) and Gompertz slope (for ages 40-89) in 1975-79. As expected, the results reveal that for all 29 countries during 1975-79 the life table entropy parameters (H and G) are larger for males than females, suggesting that the age distribution of deaths is more dispersed for the former compared with the latter. Conversely, the Gompertzian slope is larger for females than males in every country, indicating mortality rises more steeply across adult ages for females, which is consistent with deaths being concentrated in a narrower age range. Both of these results support the argument that given the same rate of mortality decline for both sexes, males should enjoy greater gains in $e(0)$ than females—thereby narrowing the sex gap.

[INSERT TABLE 2 ABOUT HERE]

Contributions to widening versus narrowing of the sex gap by age group

In this section of the analysis, we decompose by age group the contribution to changes sex gap in *e*(0) during the period when the gap was widening (1950-54 to 1975-79) and the period when the gap began to narrow (1975-79 to 2000-04). From 1950-54 to 1975-79, the sex gap in *e*(0) widened in all 21 countries observed, with an increase ranging from +0.9 years (England & Wales) to +3.9 years (the Netherlands). As shown in Figure 3, ages 40 and older were the biggest contributors to this increase, especially ages 60-79. During the same period, the youngest age group (0-19) actually had a narrowing effect on the sex gap for all but three countries (Italy, Japan, and New Zealand).

[INSERT FIGURE 3 ABOUT HERE]

During 1975-79 to 2000-04, the sex gap in *e*(0) narrowed for 19 of 29 countries, with a decline ranging from -0.3 years (Belgium) to -2.4 years (United States). Many of the same ages that contributed to widening the sex gap in the earlier period also contributed to the narrowing in this recent period: the biggest contributors to reducing the sex gap were ages 40-79 (Figure 4). In contrast, ages 80 and older continued to have a widening effect on the sex gap for all but one country (United States). One of the notable differences between countries with the largest reduction in the sex gap during this period (United States, Canada, Finland) and those with the smallest decrease in the sex gap (Italy, Czech Republic, Belgium) is the contribution of those aged 80 and older: among the latter countries, the oldest ages had a greater offsetting effect (i.e., widening the gap by almost as much as other age groups narrowed the gap) than in the former.

[INSERT FIGURE 4 ABOUT HERE]

Changes in the sex mortality gap: effects of sex ratios versus age pattern of mortality

Changes in sex ratios in mortality (or equivalently, sex differences in rates of mortality change) certainly affect the absolute sex gap in *e*(0), yet this gap is also affected by changes in the overall level of mortality if males and females have different age patterns of mortality. Here, we determine how much of the widening (or narrowing) in the sex gap in *e*(0) was due to increasing (or decreasing) sex ratios in mortality rates versus sex differences in the age pattern of mortality.

Figure 5 reveals that among all 21 countries observed, the widening sex gap between 1950-54 and 1975-79 was entirely due to increasing sex ratios in mortality rates. In contrast, sex differences in the age pattern of mortality had a dampening effect on the sex gap. For example, in Finland, increasing sex ratios contributed 3.8 years to widening the sex gap, whereas sex differences in the age pattern of mortality reduced the sex gap by 1.6 years, resulting in a net increase of 2.2 years.

[INSERT FIGURE 5 ABOUT HERE]

Between 1975-79 and 2000-04, the sex gap increased further in most Eastern European countries as well as Spain and Japan. (Among those 10 countries, the gap actually started to narrow in six countries, but the onset is too recent for the narrowing to offset the widening in the 1980s and 1990s.) In all of these countries, increasing sex ratios in mortality contributed to the widening sex gap (Figure 6).

[INSERT FIGURE 6 ABOUT HERE]

Yet, Russia and the Ukraine stand out because most of this increase resulted from sex differences in the age pattern of mortality. These two countries are unique in that *e*(0) actually *declined* over this 25-year period (Russia: -3.2 years for males, -1.1 years for females; Ukraine: -3.1 years for males; -0.6 years for females; data not shown), whereas *e*(0) increased for all other countries in this analysis. Because males were hit harder than females by increasing mortality, growing sex ratios in mortality accounted for part of the widening sex gap in *e*(0). In addition, sex differences in the age pattern of mortality also played a role because the steeper rise in mortality with age among females meant that they experienced smaller losses in *e*(0) than males. These results imply that the sex gap still would have widened even if the rate of mortality increase had been the *same* for males and females.

Among the remaining eight countries shown in the top half of Figure 6, increasing sex ratios accounted for the vast majority, if not all, of the widening sex gap. Nevertheless, we should note that mortality trends varied across these countries. Whereas in Japan, Spain, and East Germany *e*(0) gained more than five years over this 25-year period for both males and females, the other countries saw much smaller gains in *e*(0), reflecting more modest declines in mortality for both sexes. Because of these limited mortality reductions, sex differences in the age pattern of mortality did less to offset the widening effect produced by growing sex ratios in mortality for the latter countries than for the former. Nonetheless, for Japan, Spain, and East Germany, the narrowing effect on the sex gap resulting indirectly from large declines in mortality were not sufficient to offset the widening effect of substantial increases in sex ratios in mortality.

Among the 19 countries where the sex gap narrowed since the mid-1970s, declining sex ratios in mortality account for only part of the reduction in the sex gap (bottom half of Figure 6). Sex differences in the age pattern of mortality also had a substantial narrowing effect (from -0.4 years in Denmark to -1.8 in Portugal). In fact, for seven of these countries (Austria, Belgium, Czech Republic, France, Italy, Portugal, and West Germany), the decreased sex gap was due entirely to this effect; the change in sex ratios—which increased for many age groups, particularly among young adults and at older ages—actually had a widening effect. In

Switzerland, the decreased sex gap was also mostly (84%) due to sex differences in the age pattern of mortality. Among the eleven countries with the biggest declines in the sex gap, both factors made substantial contributions; declining sex ratios accounted for 30% (Finland) to 73% (United States) of the reduction in the sex gap. For 12 of these 19 countries, sex differences in the age pattern of mortality contributed substantially more to the narrowing of sex gap than changes in sex mortality ratios did, thereby lending support to the hypothesis that the recent narrowing may be primarily due to sex differences in the age pattern of mortality.

Discussion

The historical trends in the sex gap in $e(0)$ reviewed in this paper suggest there were period effects experienced by virtually all countries, although they affected some countries earlier than others. After historic widening of the sex gap among all countries observed here, most of these countries have demonstrated substantial narrowing of the sex gap in recent years. Thus, attempts to explain sex differentials in $e(0)$ should look longitudinally at factors that have changed over time. Cross-sectional studies of variation across country may not reveal the causal factors involved.

To understand this recent narrowing, it seems logical to focus on behavioural and medical factors that have different trends by sex (or have a differential effect by sex) and thereby, produce different rates of decline in age-specific mortality and change mortality sex ratios. Nonetheless, it may be misleading to look only for such factors. The sex gap in $e(0)$ varies not only with sex ratios in mortality rates, but also the level and age pattern of mortality. Our analyses show that although widening of the sex gap was almost entirely due to increasing sex ratios, the widespread narrowing of the sex gap in recent years is explained only in part by declining sex ratios. For the majority of countries, the reduction in the $e(0)$ sex gap resulted in large part from sex differences in the age pattern of mortality. Nonetheless, we must keep in mind that these countries vary widely in terms of population size (from 288,472 in Iceland to 289 million in the USA as of January 1, 2003; HMD, 2006). Among the 19 countries where the sex gap narrowed since the mid-1970s, the U.S. (where declining sex ratios accounted for 73% of narrowing) accounts for more than 40% of the total population. Altogether, the countries where the majority of narrowing of the sex gap resulted from declining sex ratios represent about 60% of the total population.

Reasons for the steeper age-related rise in mortality for females than males are not fully known. There seem to be at least two possible explanations for this sex difference. First, unhealthy life styles such as smoking, excessive drinking, and occupational hazard among middle-aged men may raise their death rates (Waldron 2005). The prevalence rates of those factors and their mortality effects may decline with age, partly because of age-related behavioural changes and partly because of selective survival, thereby making the male mortality slope less steep. Second, women at reproductive ages benefit greatly from their sex hormones, particularly oestrogen. Postmenopausal changes in the sex hormone status may make the age-related increase in mortality faster for females than males (Horiuchi 1997). However, assessment of effects of these and other factors on sex differences in the age pattern of mortality are beyond the scope of this paper and awaits further investigation.

A few cautionary remarks seem important. First, our results do not indicate that changes in sex mortality ratios played only a minor role in the *reversal* of the sex gap trend. Logically, the reversal occurs in two steps: the gap ceases to widen and then narrows. Thus, the major reason for the *cessation of widening* and that for the recent *narrowing* can be different. The sharp contrast between Figures 5 and 6 suggests that the considerable increase of sex mortality ratios in various countries slowed down, ceased, or reversed, thereby stopping the widening of the *e*(0) sex gap. Preston and Wang (2006) demonstrate that sex differences in smoking patterns by cohort could explain this reversal in the U.S. However, Figure 6 indicates that for the majority of countries, trends in sex mortality ratios played a lesser role in narrowing the sex gap than in the cessation of widening.

Second, our results do not necessarily suggest that behavioural and medical factors are unimportant for the recent narrowing of the *e*(0) sex gap. Those factors not only directly contribute to the narrowing by lowering sex mortality ratios in some countries, but also indirectly contribute to the narrowing phenomenon to the extent that they reduce overall mortality for both sexes and to the extent that they concentrate deaths into a narrower age range for females than males.

Third, the slower gain in life expectancy among women than men in recent years should not be simply attributed to higher female than male life expectancy. If the increase in *e*(0) tends to slow down at low mortality levels (Olshansky et al. 1990), the recent narrowing may not be surprising: women typically have higher *e*(0) than men, thus the *e*(0) gain for females is expected to be smaller than that for males, if other things are equal for both sexes. However, life expectancy trends in industrialized countries during recent decades are generally linear and do not indicate a tendency for the *e*(0) gain to be smaller at lower mortality levels (White 2002).

In conclusion, the results of this research seem to illustrate the risk of interpreting observed demographic trends without considering mathematical relationships at the aggregate level. In this case, the assumption that recent narrowing of the sex gap results from sex differences in the rate of mortality decline leads one to focus on sex differences in socio-behavioural and biomedical factors. Yet, the reality is that a substantial proportion of the reduction in the sex gap stems from sex differences in the age pattern of mortality rather than sex differences in the rate of mortality decline. Equation (1), which relates life expectancy gain with mortality decline and mortality variability (Keyfitz 1985), turned out to be the key to understanding the recent narrowing of sex gap in $e(0)$. Thus, this study clearly demonstrates that trends in the sex gap in life expectancy can be significantly affected (not only by sex differentials in the rate of mortality decline but also) by sex differences in the age pattern of mortality.

Appendix: Decomposition of changes in the sex gap in life expectancy

Mathematical Relationships

In this appendix, we follow the conventional life table notation and denote the force of mortality, death density, proportion surviving, and life expectancy at age x ($x \in [0,\omega]$) and time t by $\mu(x,t)$, $d(x,t)$, $l(x,t)$ and $e(x,t)$, respectively. Subscripts m and f indicate males and females, and ω is the highest age attained. The life expectancy at birth in the life table at time t is given by

$$e(0,t) = \int_0^{\omega} l(x,t)\,dx = \int_0^{\omega} e^{-\int_0^x \mu(y,t)dy}\,dx\,. \tag{A.1}$$

Differentiation of the above equation with respect to $\ln \mu(x,t)$ leads to

$$\frac{\partial e(0,t)}{\partial \ln \mu(x,t)} = -d(x,t)e(x,t) \tag{A.2}$$

The sex gap in *e(0)* can be expressed as:

$$\begin{aligned} e_f(0,t) - e_m(0,t) &= \int_0^{\omega} e^{-\int_0^x \mu_f(y,t)dy}\,dx - \int_0^{\omega} e^{-\int_0^x \mu_m(y,t)dy}\,dx \\ &= \int_0^{\omega} e^{-\int_0^x \alpha(y,t)^{-1/2}\beta(y,t)dy}\,dx - \int_0^{\omega} e^{-\int_0^x \alpha(y,t)^{1/2}\beta(y,t)dy}\,dx \end{aligned} \tag{A.3}$$

where $\alpha(x,t) = \mu_m(x,t) / \mu_f(x,t)$ and $\beta(x,t) = \sqrt{\mu_m(x,t) \cdot \mu_f(x,t)}$. It follows from these definitions of $\alpha(x,t)$ and $\beta(x,t)$ that

$$\frac{\partial \ln \mu_m(x,t)}{\partial t} = \frac{\partial \alpha(x,t)/\partial t}{2\alpha(x,t)} + \frac{\partial \beta(x,t)/\partial t}{\beta(x,t)} \tag{A.4}$$

and

$$\frac{\partial \ln \mu_f(x,t)}{\partial t} = -\frac{\partial \alpha(x,t)/\partial t}{2\alpha(x,t)} + \frac{\partial \beta(x,t)/\partial t}{\beta(x,t)} \quad . \tag{A.5}$$

By differentiating equation (A.3) with respect to age and making use of equations (A.2), (A.4), and (A.5), the rate of absolute change in the sex gap in life expectancy can be expressed as

$$\begin{aligned} \frac{\partial}{\partial t}\left\{e_f(0,t) - e_m(0,t)\right\} = &\int_0^\omega \left\{\frac{d_m(x,t)e_m(x,t) + d_f(x,t)e_f(x,t)}{2}\right\} \frac{\partial \alpha(x,t)/\partial t}{\alpha(x,t)}\, dx \\ &+ \int_0^\omega \left\{d_m(x,t)e_m(x,t) - d_f(x,t)e_f(x,t)\right\} \frac{\partial \beta(x,t)/\partial t}{\beta(x,t)}\, dx \end{aligned} \quad . \tag{A.6}$$

This equation is the mathematical basis of the decomposition analysis in this study. The right hand side of the equation has two terms. The first term indicates effects of changes in mortality sex ratios (i.e., effects of sex differences in the rate of mortality change). The second term indicates effects of changes in the geometric means of male and female mortality, which depends on sex differences in the age pattern of mortality: if there were no sex differences in the $d(x,t)e(x,t)$ function, then changes in mortality level would have no effect on the sex gap because this term would cancel out.

Equation (A.6) is closely related to Keyfitz' equation about the change in life expectancy. Keyfitz (1985) has shown that if the forces of mortality decline at the same rate at all ages (i.e., if $\partial \ln \mu(x,t)/\partial t = -\rho(t)$ for any $x \in [0,\omega]$), then

$$\frac{de(0,t)}{dt} = G(t)\rho(t) \text{ where } G(t) = -\int_0^\omega l(x,t) \ln l(x,t) dx \, . \tag{A.7}$$

Goldman and Lord (1986) and Vaupel (1986) have indicated that *G(t)* can also be expressed as:

$G(t) = \int_0^{\omega} d(x,t)e(x,t)dx$.

If $\rho(t)$ is the same for males and females, then $\partial\alpha(x,t)/\partial t = 0$ and $\partial\beta(x,t)/\partial t = -\beta(x,t)\rho(t)$ for any *x*. Substituting these into equation (A.6), we get

$$\begin{aligned} \frac{\partial}{\partial t}\left\{e_f(0,t) - e_m(0,t)\right\} &= \int_0^{\omega} \left\{d_f(x,t)e_f(x,t) - d_m(x,t)e_m(x,t)\right\} \rho(t)\, dx \\ &= \left\{G_f(t) - G_m(t)\right\}\rho(t) \ , \end{aligned} \qquad \text{(A.8)}$$

where $G_m(t)$ and $G_f(t)$ are male and female versions of *G(t)*. Obviously, equation (A.8) can be considered a sex-difference variant of Keyfitz' equation.

Statistical Procedures

Equation A.6 decomposes the change in the sex gap in life expectancy during an infinitesimal time period, based on the continuous-time and continuous-age framework. However, we actually need to decompose changes in the sex gap during relatively long periods, using data by discrete age groups for discrete time periods. Let $\theta(t)$ represent the sex gap in $e(0)$ at time *t*: $\theta(t) = e_f(0,t) - e_m(0,t)$. The change in the sex gap between t_1 and t_2 can be expressed as: $\lambda(t_1, t_2) = \theta(t_2) - \theta(t_1)$. Changing from continuous to discrete formulation, let $M_{m,i}(t)$ and $M_{f,i}(t)$ be the age-specific death rates among males and females, respectively, for the i^{th} age group in period *t*. Thus, $\alpha_i(t) = M_{m,i}(t)/M_{f,i}(t)$ and $\beta_i(t) = \sqrt{M_{m,i}(t) \cdot M_{f,i}(t)}$.

Referring back to equation (A.3), the sex gap can be expressed as a function of α's and β's:

$$\theta(t) = f[\alpha_1(t), \alpha_2(t), \ldots \alpha_n(t), \beta_1(t), \beta_2(t), \ldots \beta_n(t)]. \qquad \text{(A.9)}$$

The contribution of i^{th} age group to the change in the sex gap during the infinitesimal time interval between *t* and $t+\Delta t$ is:

$$\left(\frac{\partial\theta(t)}{\partial\alpha_i(t)}\frac{d\alpha_i(t)}{dt}+\frac{\partial\theta(t)}{\partial\beta_i(t)}\frac{d\beta_i(t)}{dt}\right)\Delta t\ , \tag{A.10}$$

which is equivalent to $\left(\frac{\partial\theta(t)}{\partial M_{m,i}(t)}\frac{dM_{m,i}(t)}{dt}+\frac{\partial\theta(t)}{\partial M_{f,i}(t)}\frac{dM_{f,i}(t)}{dt}\right)\Delta t$.

Using the line-integral method of decomposition, which was applied in several previous studies (Horiuchi et al., 1999; Wilmoth and Horiuchi 1999; Pletcher et al. 2000; Wilmoth et al. 2000), the change in the sex gap between t_1 and t_2 can be decomposed:

$$\lambda(t_1,t_2)=\sum_i\int_{t_1}^{t_2}\frac{\partial\theta(t)}{\partial\alpha_i(t)}\frac{d\alpha_i(t)}{dt}dt+\sum_i\int_{t_1}^{t_2}\frac{\partial\theta(t)}{\partial\beta_i(t)}\frac{d\beta_i(t)}{dt}dt\ . \tag{A.11}$$

The first term on the right hand side is the mortality sex ratio effect, and the second term is the overall mortality level effect (which interacts with sex differences in the age pattern of mortality). The effect of the i^{th} age group is:

$$\int_{t_1}^{t_2}\frac{\partial\theta(t)}{\partial\alpha_i(t)}\frac{d\alpha_i(t)}{dt}dt+\int_{t_1}^{t_2}\frac{\partial\theta(t)}{\partial\beta_i(t)}\frac{d\beta_i(t)}{dt}dt\ . \tag{A.12}$$

We use this method to decompose the change in the sex gap in *e*(0) from each five-year time period to the next into the effects due to changes in death rates for each of 20 age groups (0, 1-4, 5-9,…85-89, 90+), and within each age group, the effects due to changes in sex ratios versus sex differences in the age pattern of mortality.

Between consecutive five-year time periods, we assume that all changes in the logarithms of age-specific death rates are proportional to each other (i.e., there is a continuous function $\phi(t)$ such that

$$\frac{\ln M_{k,i}(t)-\ln M_{k,i}(t_1)}{\ln M_{k,i}(t_2)-\ln M_{k,i}(t_1)}=\phi(t) \tag{A.13}$$

for sex *k* (m or f), any age group *i,* and any *t* between t_1 and t_2, and where $\phi(t)$ ranges from 0 to 1). For example, if $\phi(t) = 0.6$, then every variable $\ln M_{k,i}(t)$ has completed 60% of its "trip" from t_1 to t_2. This proportionality assumption is equivalent to the assumption of constant age pattern of mortality change in the Lee-Carter model. It also means that the line integral of equation (A.9) is calculated along the straight line from

$[\ln M_{m,1}(t_1), \ln M_{m,2}(t_1), \ldots \ln M_{m,n}(t_1), \ln M_{f,1}(t_1), \ln M_{f,2}(t_1), \ldots \ln M_{f,n}(t_1)]$ to

$[\ln M_{m,1}(t_2), \ln M_{m,2}(t_2), \ldots \ln M_{m,n}(t_2), \ln M_{f,1}(t_2), \ln M_{f,2}(t_2), \ldots \ln M_{f,n}(t_2)]$ in the *2n*-dimensional space. Because the proportionality is assumed for changes in a short period (i.e., between two consecutive five-year time intervals), but not for changes in a long period such as several decades, deviations from this assumption are unlikely to be substantial.

Starting with 1950-54, the decomposition method was applied to pairs of successive five-year time periods for each of 29 populations, which resulted in a total of 286 pairs (because not all countries have data back to 1950). The number of intervals used for numerical integration was set at five. Accordingly, the period from 1950 to 2004 is divided into 50 intervals (5 x 10 pairs of successive five-year time periods). The proportional error of numerical integration was less than one percent for all but two pairs, both of which had virtually no change in the sex gap (< 0.1 years) between the two successive periods and an absolute error no greater than 0.001 years.

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Table 1. Range of years covered, sex gap in *e*(0) during selected five-year periods, and onset of narrowing for each population

Population	Code	Range of years	Sex gap in *e*(0)					Onset of sustained decline in sex gap in *e*(0)
			1900-04	1925-29	1950-54	1975-79	2000-04	
1. Australia	AUS	1921-2003	na	3.9	5.5	7.1	5.1	1981
2. Austria	AUT	1947-2002	na	na	5.2	7.1	6.0	1983
3. Belgium	BEL	1931-2002	na	na	5.0	6.6	6.2	1996
4. Bulgaria	BGR	1947-2003	na	na	3.4	5.1	6.9	1996
5. Canada	CAN	1921-2002	na	2.5	4.9	7.4	5.1	1979
6. Czech Republic	CZE	1950-2004	na	na	4.9	7.1	6.6	1991
7. Denmark	DNK	1835-2004	3.3	1.6	2.6	5.9	4.6	1982
8. England & Wales	ENW	1841-2003	3.8	4.1	5.2	6.1	4.5	1972
9. Finland	FIN	1878-2004	2.7	5.0	6.6	8.8	6.8	1978
10. France	FRA	1899-2003	3.5	4.4	5.9	8.1	7.3	1993
11. Germany, East (former)	GDR	1956-2002	na	na	na	5.8	6.7	1995
12. Germany, West (former)	FRG	1956-2002	na	na	na	6.7	5.7	1981
13. Hungary	HUN	1950-2001	na	na	4.3	6.4	8.5	1995
14. Iceland	ISL	1838-2004	4.4	2.8	4.4	6.0	3.8	1984
15. Italy	ITA	1872-2003	0.4	1.8	3.7	6.6	5.9	1992
16. Japan	JPN	1950-2004	na	na	3.5	5.2	7.1	[1]
17. Latvia	LVA	1959-2003	na	na	na	10.2	11.0	1995
18. Lithuania	LTU	1959-2003	na	na	na	9.6	11.2	[1]
19. Netherlands	NLD	1850-2004	2.9	1.5	2.5	6.4	4.8	1983
20. New Zealand	NZL	1948-2003	na	na	4.4	6.6	4.7	1980
21. Norway	NOR	1846-2004	3.1	2.8	3.5	6.4	5.1	1987
22. Portugal	PRT	1956-2005	na	na	na	7.1	6.7	1997
23. Russia	RUS	1959-2005	na	na	na	11.2	13.3	[1]
24. Slovakia	SVK	1950-2005	na	na	3.8	7.3	8.0	2001
25. Spain	ESP	1908-2004	na	3.1	4.6	6.0	6.8	1997
26. Sweden	SWE	1751-2004	2.6	2.1	2.8	6.1	4.5	1980
27. Switzerland	CHE	1876-2004	2.7	3.3	4.7	6.6	5.4	1992
28. Ukraine	UKR	1959-2005	na	na	na	9.0	11.4	[1]
29. United States	USA	1959-2002	na	na	na	7.7	5.3	1976

Source: HMD (2006).

[1] A sustained decline in the sex gap in life expectancy has not yet been observed.

Table 2. Life table entropy (*H* and *G*) and Gompertzian slope (*b*) by sex in 1975-79 for each population

Population	*H*		*G*		*b*	
	Females	Males	Females	Males	Females	Males
Australia	0.15	0.18	11.3	12.6	0.098	0.090
Austria	0.14	0.18	10.8	12.7	0.109	0.092
Belgium	0.14	0.18	10.8	12.2	0.106	0.092
Bulgaria	0.15	0.19	11.2	13.0	0.109	0.093
Canada	0.15	0.18	11.6	13.1	0.095	0.086
Czech Republic	0.14	0.18	10.6	12.3	0.109	0.091
Denmark	0.14	0.17	11.0	12.1	0.097	0.090
England & Wales	0.14	0.17	11.1	11.7	0.100	0.094
Finland	0.13	0.18	10.2	12.5	0.109	0.084
France	0.14	0.18	10.8	12.8	0.105	0.085
Germany, East (former)	0.14	0.17	10.4	12.1	0.110	0.095
Germany, West (former)	0.14	0.18	10.9	12.5	0.107	0.093
Hungary	0.16	0.20	11.7	13.4	0.102	0.088
Iceland	0.14	0.17	11.1	12.7	0.098	0.087
Italy	0.14	0.18	10.6	12.4	0.110	0.092
Japan	0.13	0.16	10.0	11.4	0.110	0.097
Latvia	0.16	0.23	11.9	15.1	0.099	0.074
Lithuania	0.16	0.24	12.3	15.7	0.097	0.072
Netherlands	0.13	0.16	10.4	11.7	0.106	0.093
New Zealand	0.15	0.18	11.8	12.7	0.094	0.090
Norway	0.13	0.16	10.0	11.8	0.111	0.093
Portugal	0.16	0.21	12.1	14.2	0.107	0.090
Russia	0.17	0.25	12.6	15.6	0.096	0.071
Slovakia	0.15	0.20	11.1	13.5	0.106	0.085
Spain	0.14	0.17	10.7	12.4	0.110	0.094
Sweden	0.13	0.16	10.2	11.6	0.109	0.096
Switzerland	0.13	0.17	10.2	12.0	0.110	0.095
Ukraine	0.16	0.22	12.0	14.6	0.099	0.076
United States	0.16	0.19	12.2	13.6	0.090	0.081

Source: As for Table 1.

H = Life table entropy (Keyfitz, 1985), calculated using HMD estimates of *l(x)* for ages 0,1,….110+.

G = Numerator of Keyfitz's *H*.

b = Slope parameter from a Gompertz model fitted to death rates for ages 40-44, 45-49,…85-89 using weighted least squares.

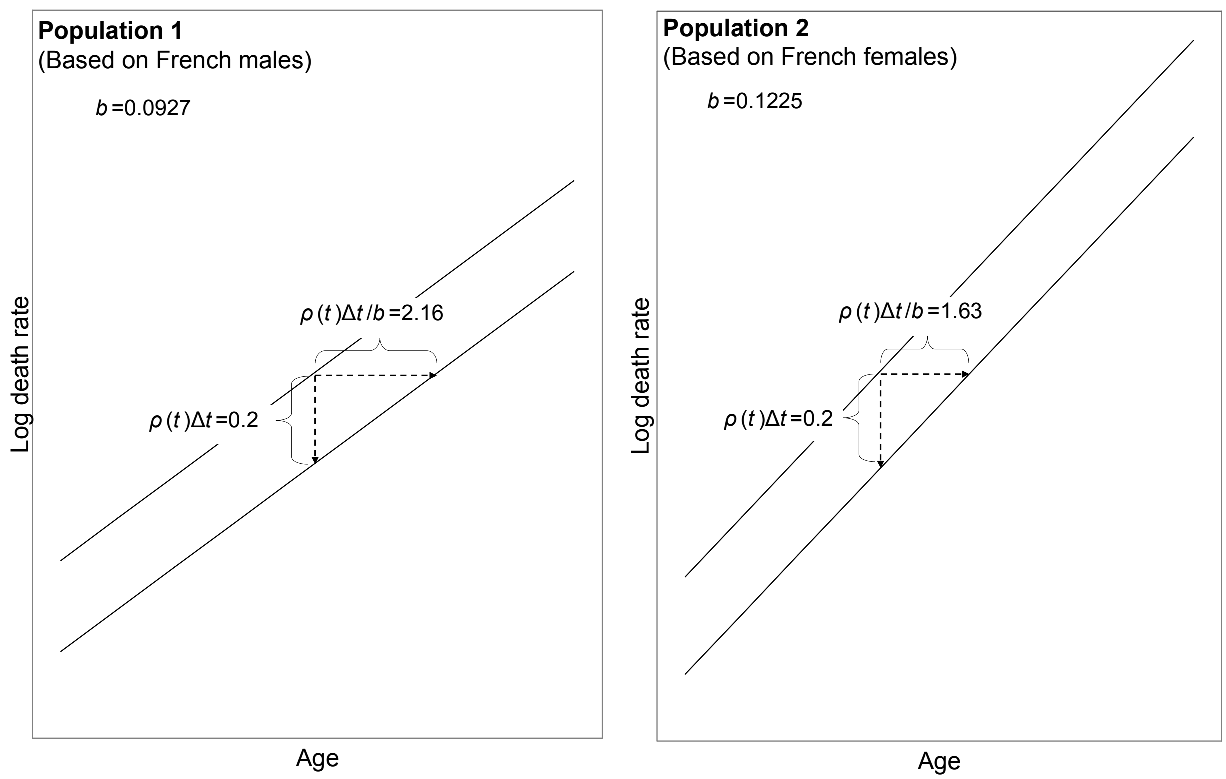

Figure 1. Effect of steepness of mortality curve on horizontal mortality shift
Source: The graphs shown above assume Gompertz mortality. Values of b come from Horiuchi et al. (2003) based on fitting a Gompertz model to mortality rates at ages 65-89 among French males and females between 1979 and 1994.

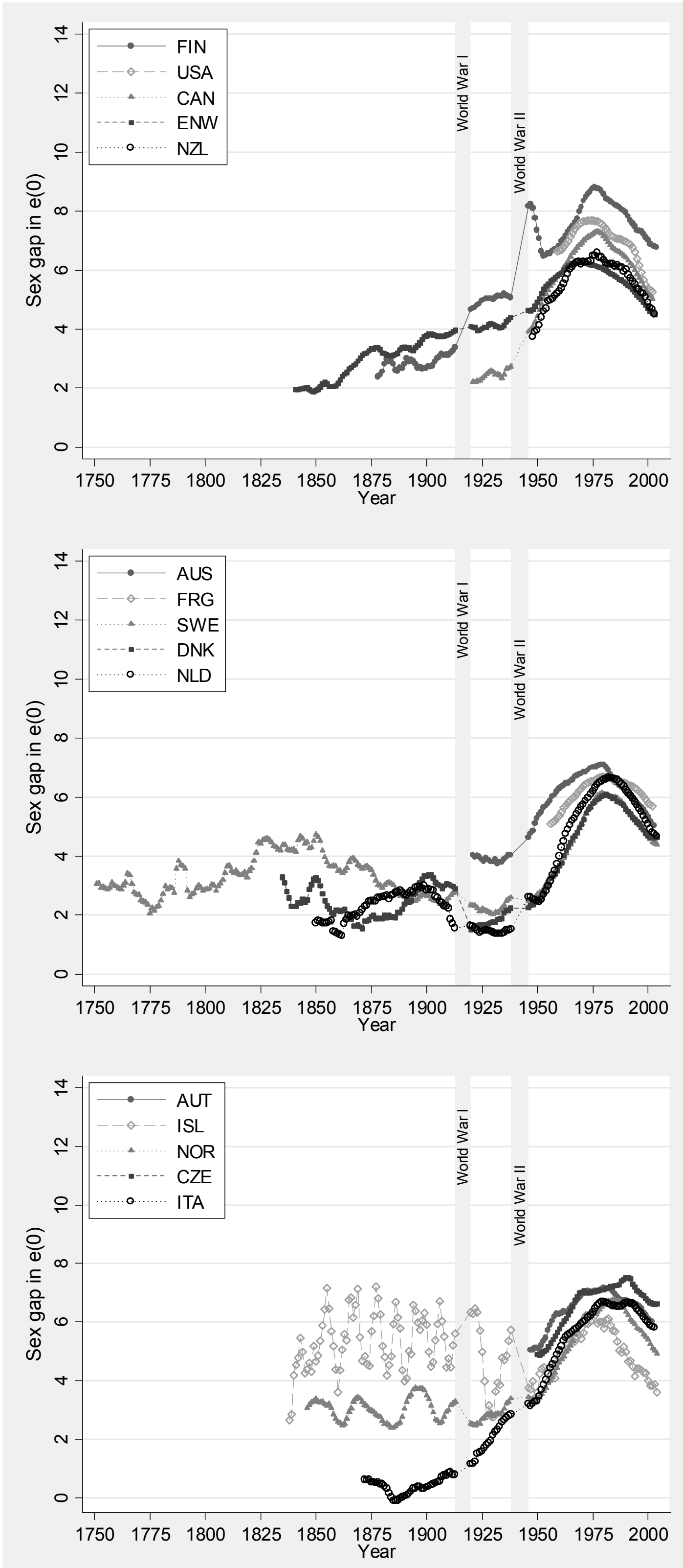

Figure 2 (continued on next page)

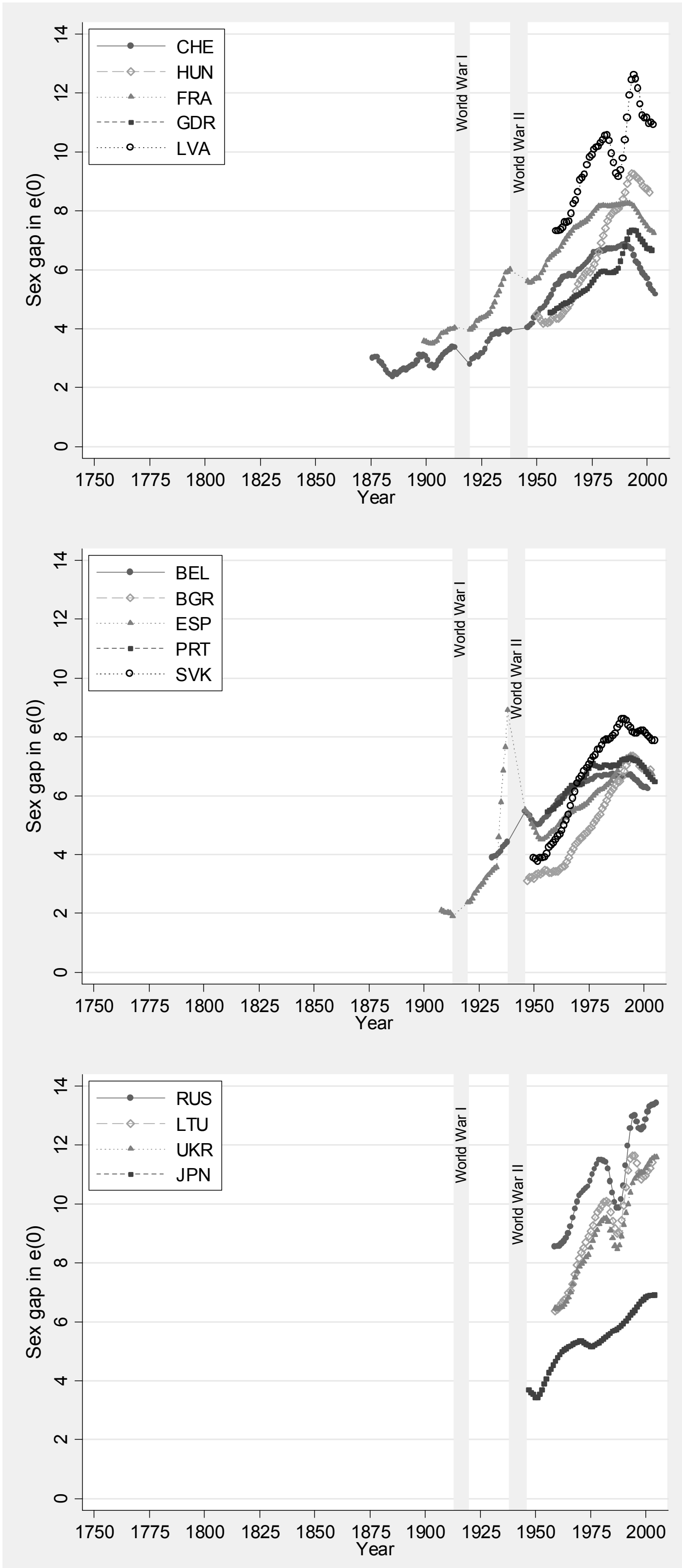

Figure 2. Sex gap (female – male) in $e(0)$ across time and country, based on a five-year moving average, 1751-2004. *Source*: As for Table 1. *Note:* See Table 1 for definition of country codes.

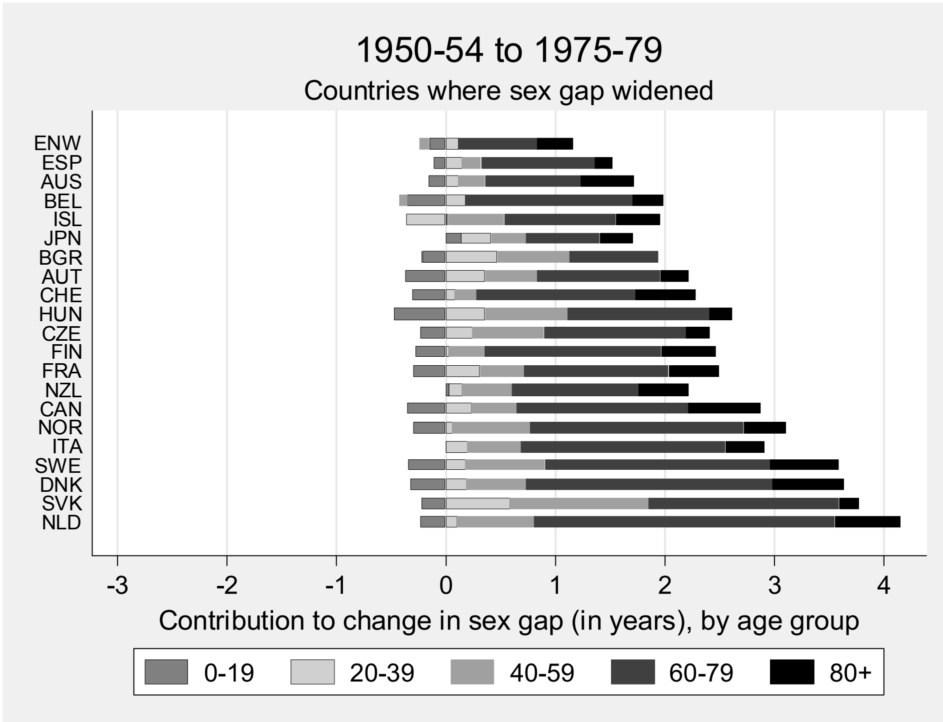

Figure 3. Contributions to widening of the sex gap in *e*(0) by age group, 1950-54 to 1975-79.
Source: As for Table 1.

Figure 4. Contributions to narrowing of the sex gap in *e*(0) by age group, 1975-79 to 2000-04.
Source: As for Table 1.

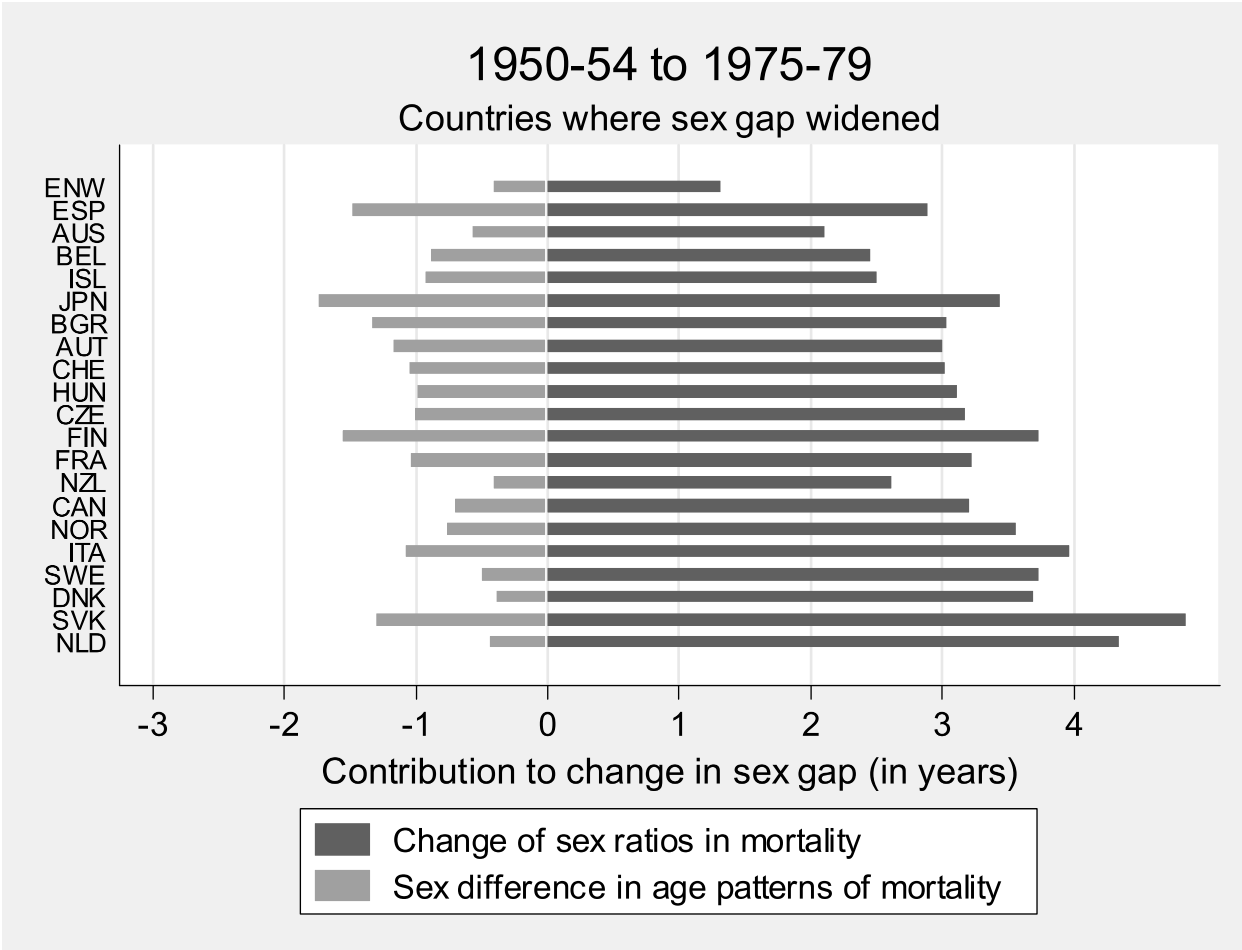

Figure 5. Widening of sex gap in *e*(0) due to changing sex ratios versus differential age patterns, 1950-54 to 1975-79.
Source: As for Table 1.

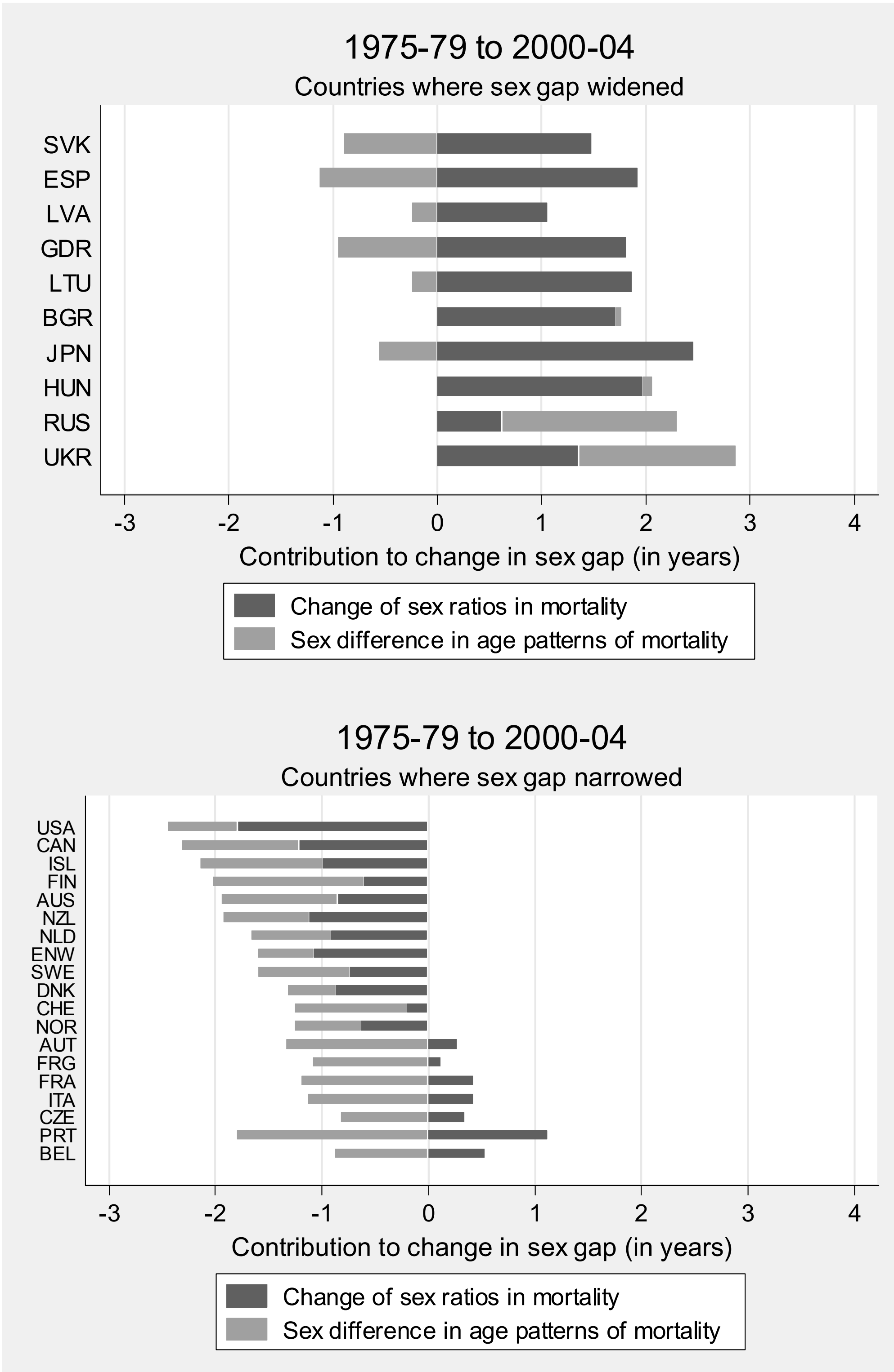

Figure 6. Changes in sex gap in *e*(0) due to changing sex ratios versus differential age patterns, 1975-79 to 2000-04.
Source: As for Table 1.