

# The narrowing sex differential in life expectancy in high-income populations: Effects of differences in the age pattern of mortality

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*Using data from the Human Mortality Database for 29 high-income national populations (1751–2004), we review trends in the sex differential in  $e(0)$ . The widening of this gap during most of the 1900s was due largely to a 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, and researchers tried to explain this reversal with the same factors. However, our decomposition analysis reveals that, for the majority of countries, the recent narrowing is due primarily 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 for men because women's deaths are less dispersed across age (i.e., survivorship is more rectangular).*

**Keywords:** life expectancy; mortality; sex difference; sex ratio; mortality decline

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## 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 Willem 1998). Indeed, sex differences in mortality are often wider than those between other subgroups (e.g., race/ethnicity). Yet, available data suggest that the gap in  $e(0)$  between the sexes changed considerably in the past (Stolnitz 1956; United Nations Secretariat 1988; Tabutin and Willem 1998). It was relatively small in the late 1800s but grew rapidly through most of the twentieth century.

The widening of the 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; Waldron 1985, 1986; Vallin 1993). But around the 1980s, the 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 gap in  $e(0)$  narrowed substantially (–1.85 years in Hong Kong to –0.26 years in USSR) in nine countries, including most English-speaking countries (i.e., USA, Canada, UK, Australia) as well as several other countries (i.e., Hong Kong, Iceland, Austria, Finland, USSR). Yet, during this period, the gap continued to widen among Eastern European countries and, to a somewhat lesser extent, among Southern European countries and Ireland. More recently, the sex differential has begun to decline among other countries in Western Europe and appears to be levelling off among several countries in Southern Europe, although it continues to increase in Japan (Meslé 2004).

Attempts to explain the 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 2006) as well as behavioural and medical factors that would reduce sex ratios in mortality rates and thereby narrow the sex differential in  $e(0)$  (Nault 1997; Valkonen and Van Poppel 1997; Pampel 2002; Conti et al. 2003; Gjonca et al. 2005; Waldron 2005;

Preston and Wang 2006). Those factors include increased smoking among women while its prevalence declined among men (Waldron 2005), and advances in medical treatments for cardiovascular disease that may have benefited men more than women (Waldron 1995).

Those factors that have different trends by sex (or have differential effects by sex) could reduce sex ratios in mortality rates and thereby narrow the gap in  $e(0)$ . The increase in these 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 differential in  $e(0)$ . The ratios declined recently in some countries, which could account, at least in part, for the narrowing gap in  $e(0)$  between the sexes in those populations.

However, changes in life expectancy also depend on age patterns of mortality. 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, chap. 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 the sex differential may have resulted primarily from differences in the age pattern of mortality, rather than from slower mortality decline for women than men. Little attention has been given to this demographic mechanism for narrowing the gap between the sexes. In this study, we explore this hypothesis using data from the Human Mortality Database (HMD, [www.mortality.org](http://www.mortality.org)) for 29 national populations with high-quality mortality statistics. We begin by reviewing trends in the sex difference 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 about the recent narrowing of the sex differential in life expectancy. First, when did a sustained narrowing of the differential in  $e(0)$  begin, and how did its 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 the widening of the differential in  $e(0)$  similar

to those that contributed to its recent narrowing? Finally, to what extent did changes in the differential result from differences between the sexes in: (i) the age-specific rates of mortality decline, and (ii) the age pattern of mortality?

### The ‘differential dispersion’ hypothesis: 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., the mortality ratios of the sexes remain unchanged) can change the sex differential in  $e(0)$ .

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

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

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  $l(x, t)$  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. Both  $H$  and  $G$  are used as measures of variability in mortality 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 the paradox that the differential in  $e(0)$  between two populations could *widen* even when the differences in age-specific death rates *decrease*. Keyfitz’s 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 ( $\rho$ ) depends on mortality dispersion ( $G$ ).

Using Keyfitz's 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  $z$ :

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

the result is a vertical downward shift of the 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  $y(c, t) = (\ln c - \ln a(t))/b$ ;  $y(c, t)$  is the inverse function of  $\mu(x, t)$  where  $y$  corresponds to  $x$  and  $c$  corresponds to  $\mu$ . Therefore, the rate of this horizontal shift is

$$\frac{\partial y(c, t)}{\partial t} = -\frac{\partial \ln a(t)/\partial t}{b} = \frac{\rho(t)}{b},$$

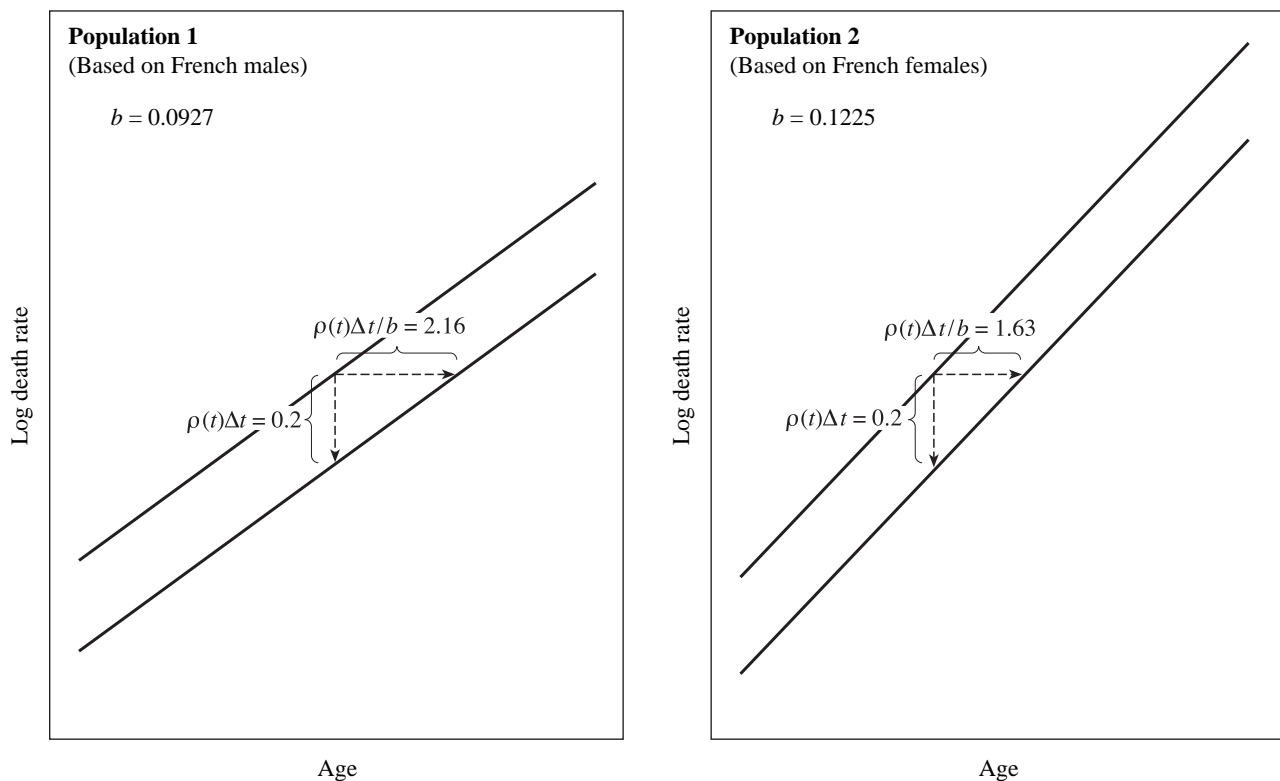
(3)

for any  $c \geq \mu(z, 0)$ .

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



**Figure 1** Effect of steepness of mortality curve on horizontal mortality shift

*Source:* The graphs 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

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 and 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's equation) but also intuitively. A steep rise of mortality with age from a low to a 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 sex ratios (male/female) of mortality tend to decline at older ages across various populations (Lopez 1983; United Nations Secretariat 1988), suggesting that adult mortality curves tend to be steeper for women than for men. Thus, we would expect similar rates of mortality decline for men and women to result in greater gains in  $e(0)$  for the former than the latter, thereby reducing the gap between the sexes.

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 difference of 8.1 years (HMD 2006). Hypothetically, if these mortality rates declined by 35 per cent at all ages for both males and females, the resulting life expectancy would be 75.0 for males and 82.2 for females, a difference of 7.2 years. Thus, the difference between the sexes would narrow by 0.9 years even if sex ratios in mortality rates remained *unchanged*.

## Data and methods

The data were obtained from the Human Mortality Database (HMD 2006) and represent 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 have data for more than 100 years (see Table 1). Because war can have a big impact on the sex ratio in mortality, we exclude the data for the periods

during the First World War (1914–19) and the Second World War (1939–45) from all of the analyses.

In the first part of the analysis, we examine the time trends in the sex difference in  $e(0)$  and determine the first year in which each country demonstrated a sustained decline in this difference. We calculate this difference ( $e_f(0) - e_m(0)$ ) 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 difference, we smooth the data using a 5-year moving average, where the difference for year  $t$  is calculated based on the mean for years  $t-2, t-1, \dots, t+2$ . This moving average gives zero weight to periods during the First and Second World War (i.e., value for 1946 is based on the average of 1946, 1947, and 1948 only). We begin by graphing the sex difference in  $e(0)$  across time and by country in order to ascertain which countries tend to have larger (or smaller) absolute differences between the sexes and how the trajectory changed over time. For each country, we define the onset of narrowing in the difference as the first year in which: (i) both the observed and smoothed differences were less than in the previous year; (ii) the smoothed difference fell to a level that has not since been exceeded; and (iii) the decline was not subsequently interrupted by an increase in the (smoothed) difference of more than 0.25 years (relative to the previous year) or a sustained increase for 3 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 narrowing of the gap between the sexes is due primarily 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 differential 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 gap between the sexes across periods in order to ascertain: (i) which age groups are the biggest contributors to the change in the gap, and (ii) to what extent the change in the gap results from sex differences in mortality change, and to what extent from sex differences in the age pattern of mortality. The first decomposition is straightforward, and is 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 the mortality rates for males and females:

**Table 1** Range of years covered, sex differential in  $e(0)$  during selected 5-year periods, and onset of narrowing of the sex differential in the populations of each of 29 high-income countries, 1900–2004

Population	Range of years	Sex differential in $e(0)$					Onset of sustained decline in sex differential in $e(0)$
		1900–1904	1925–29	1950–54	1975–79	2000–2004	
1. Australia	1921–2003	n/a	3.9	5.5	7.1	5.1	1981
2. Austria	1947–2002	n/a	n/a	5.2	7.1	6.0	1983
3. Belgium	1931–2002	n/a	n/a	5.0	6.6	6.2	1996
4. Bulgaria	1947–2003	n/a	n/a	3.4	5.1	6.9	1996
5. Canada	1921–2002	n/a	2.5	4.9	7.4	5.1	1979
6. Czech Republic	1950–2004	n/a	n/a	4.9	7.1	6.6	1991
7. Denmark	1835–2004	3.3	1.6	2.6	5.9	4.6	1982
8. England & Wales	1841–2003	3.8	4.1	5.2	6.1	4.5	1972
9. Finland	1878–2004	2.7	5.0	6.6	8.8	6.8	1978
10. France	1899–2003	3.5	4.4	5.9	8.1	7.3	1993
11. Germany, East (former)	1956–2002	n/a	n/a	n/a	5.8	6.7	1995
12. Germany, West (former)	1956–2002	n/a	n/a	n/a	6.7	5.7	1981
13. Hungary	1950–2001	n/a	n/a	4.3	6.4	8.5	1995
14. Iceland	1838–2004	4.4	2.8	4.4	6.0	3.8	1984
15. Italy	1872–2003	0.4	1.8	3.7	6.6	5.9	1992
16. Japan	1950–2004	n/a	n/a	3.5	5.2	7.1	<sup>1</sup>
17. Latvia	1959–2003	n/a	n/a	n/a	10.2	11.0	1995
18. Lithuania	1959–2003	n/a	n/a	n/a	9.6	11.2	<sup>1</sup>
19. Netherlands	1850–2004	2.9	1.5	2.5	6.4	4.8	1983
20. New Zealand	1948–2003	n/a	n/a	4.4	6.6	4.7	1980
21. Norway	1846–2004	3.1	2.8	3.5	6.4	5.1	1987
21. Portugal	1956–2005	n/a	n/a	n/a	7.1	6.7	1997
22. Russia	1959–2005	n/a	n/a	n/a	11.2	13.3	<sup>1</sup>
23. Slovakia	1950–2005	n/a	n/a	3.8	7.3	8.0	2001
24. Spain	1908–2004	n/a	3.1	4.6	6.0	6.8	1997
25. Sweden	1751–2004	2.6	2.1	2.8	6.1	4.5	1980
26. Switzerland	1876–2004	2.7	3.3	4.7	6.6	5.4	1992
27. Ukraine	1959–2005	n/a	n/a	n/a	9.0	11.4	<sup>1</sup>
28. USA	1959–2002	n/a	n/a	n/a	7.7	5.3	1976

<sup>1</sup>A sustained decline in the sex differential in life expectancy has not yet been observed.

Source: HMD 2006.

$$\begin{aligned}\mu_f(x) &= \beta(x)/\sqrt{\alpha(x)} \text{ and} \\ \mu_m(x) &= \beta(x) \cdot \sqrt{\alpha(x)}\end{aligned}\quad (4)$$

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 sex differential in  $e(0)$  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 gap in  $e(0)$  that would occur if the death rates of both females and

males 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 the life tables of females and males are identical. The mathematical formulation and statistical method for this decomposition analysis are described in more detail in the Appendix.

For the decomposition, we start with period death counts and estimates of exposure 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 5-year time intervals and 5-year age groups except for the first and last age intervals (0, 1–4, 5–9, ..., 85–89, 90+). Based on these  $5 \times 5$  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 gap between the sexes in  $e(0)$  for each pair of successive 5-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 5-year time intervals into the period during which the gap between the sexes was widening in all countries (1950–54 to 1975–79) and the period during which it began to narrow (1975–79 to 2000–2004).

## Results

### *Sex differences 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 for a period before the 1830s—the sex difference in  $e(0)$  was around 3 years during the late 1700s, but began to increase in the early 1800s reaching more than 4.5 years by 1830.

Starting around the mid-1800s, data became available for England and Wales, Denmark, Iceland, the Netherlands, and Norway. At this time, the sex differential in  $e(0)$  was around 2–3 years with the exceptions of Iceland and Sweden, where the advantage of females exceeded 4 years.

Data series for Italy, Finland, and Switzerland begin in the 1870s. The gap between the sexes remained between 2 and 4 years for seven of the nine countries observed. The two exceptions were: Iceland, with a gap greater than 4 years, and Italy, with a much smaller gap ( $\approx 0.5$  years in the 1870s). Throughout the second half of the nineteenth century the sex difference grew in England and Wales (from  $\approx 2.0$  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  $\approx 2.5$  years in 1900).

By 1900–1904, the 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 gap continued to grow in England and 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 differential declined between 1900

and 1920 in Denmark and the Netherlands. During 1925–29, the sex difference in  $e(0)$  ranged from 1.5 years in the Netherlands to 5.0 years in Finland among 13 countries.

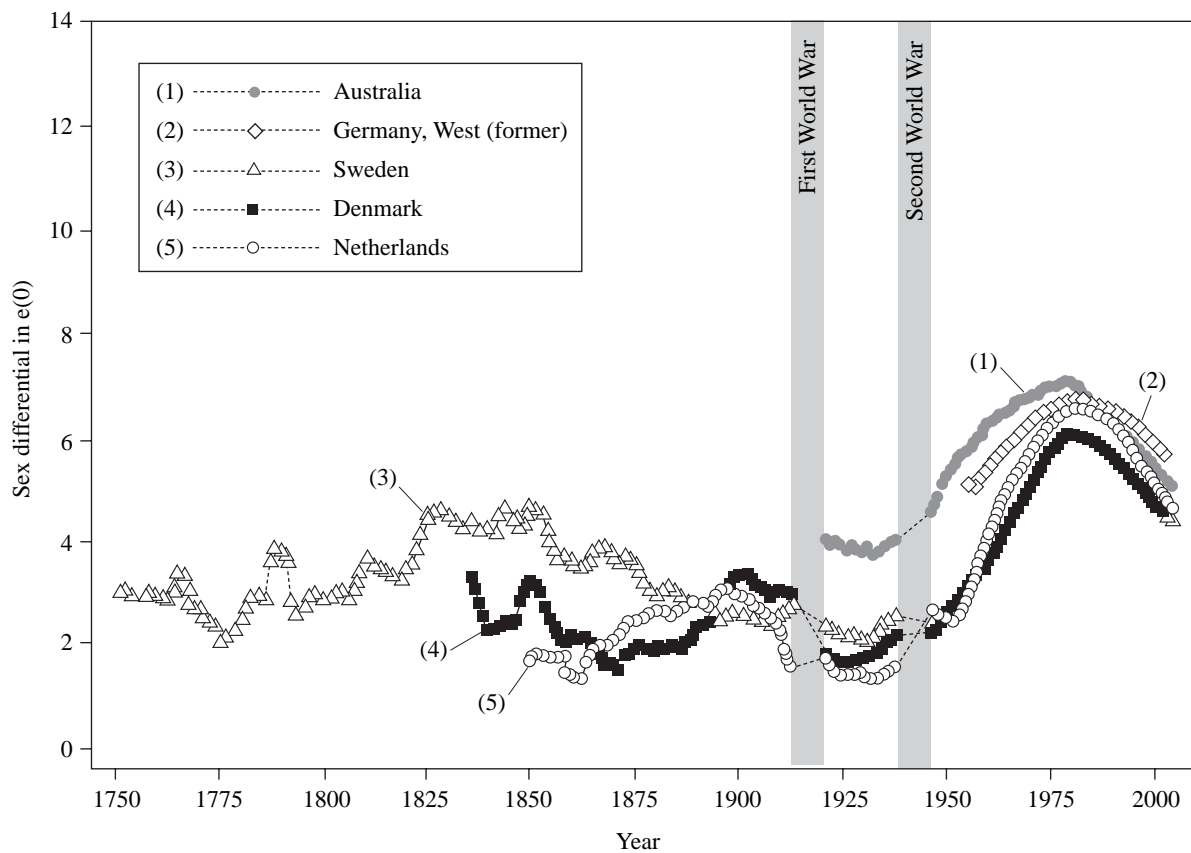
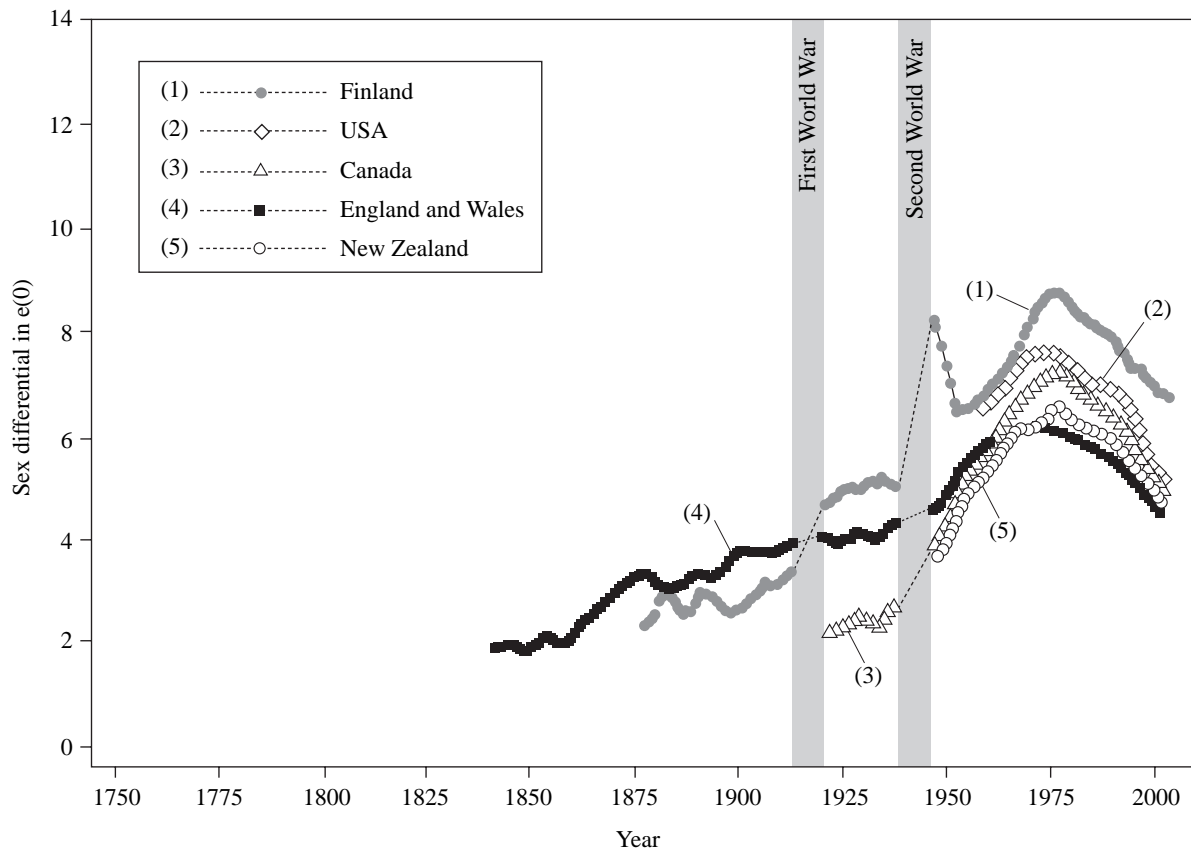
Before and after the Second World War there were several notable changes in the sex differences in  $e(0)$ . For Spain, the sharp pre-war rise in the difference reflects excess mortality among men owing to the Spanish Civil War (1936–39); after the war,  $e(0)$  for males returned to the historical trend with a corresponding decline in the difference between the sexes from the peak levels attained at the end of the Spanish Civil War. As for France, the gap increased during the pre-war period because gains in  $e(0)$  for males did not keep pace with those for females, but males ‘caught up’ in the post-war period (so that the gap returned to the historical trend line). In contrast, the gap in Finland increased markedly between 1938 and 1946 because females made greater gains in  $e(0)$  than males—perhaps owing to the aftermath of the Second World War 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 gap in  $e(0)$  was growing rapidly in all 21 countries observed and continued to increase through the 1970s. During the period 1950–54, it remained lowest in the Netherlands (2.5 years) and highest in Finland (6.6 years). By 1975–79, the countries of the former USSR (Latvia, Lithuania, Russia, and Ukraine) exhibited the largest sex differences (9.0–11.2 years) among the 29 countries observed, while others ranged from 5.1 (Bulgaria) to 8.8 years (Finland).

In the last two decades of the twentieth century, the gap in  $e(0)$  began a steady decline in most countries, although the timing of this reversal varied across countries. By 2000–2004, six countries exhibited a gap of less than 5 years.

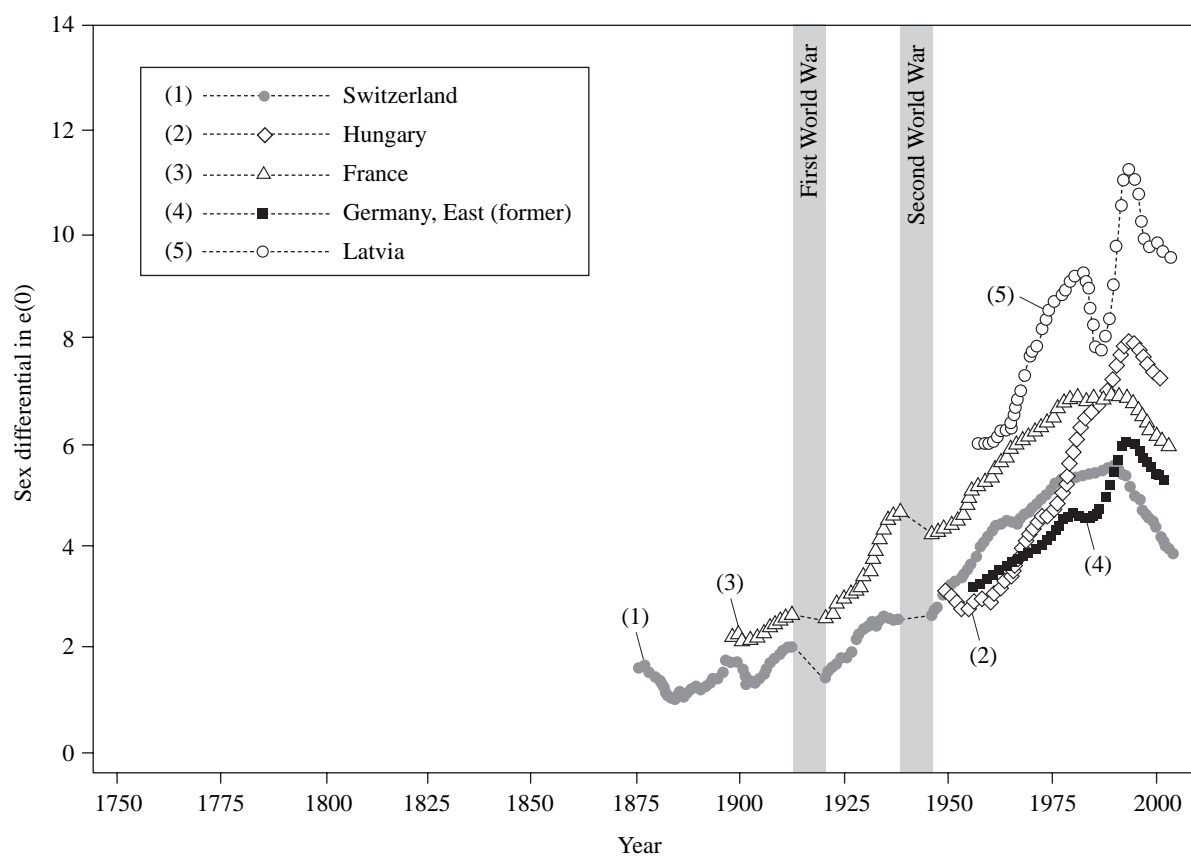
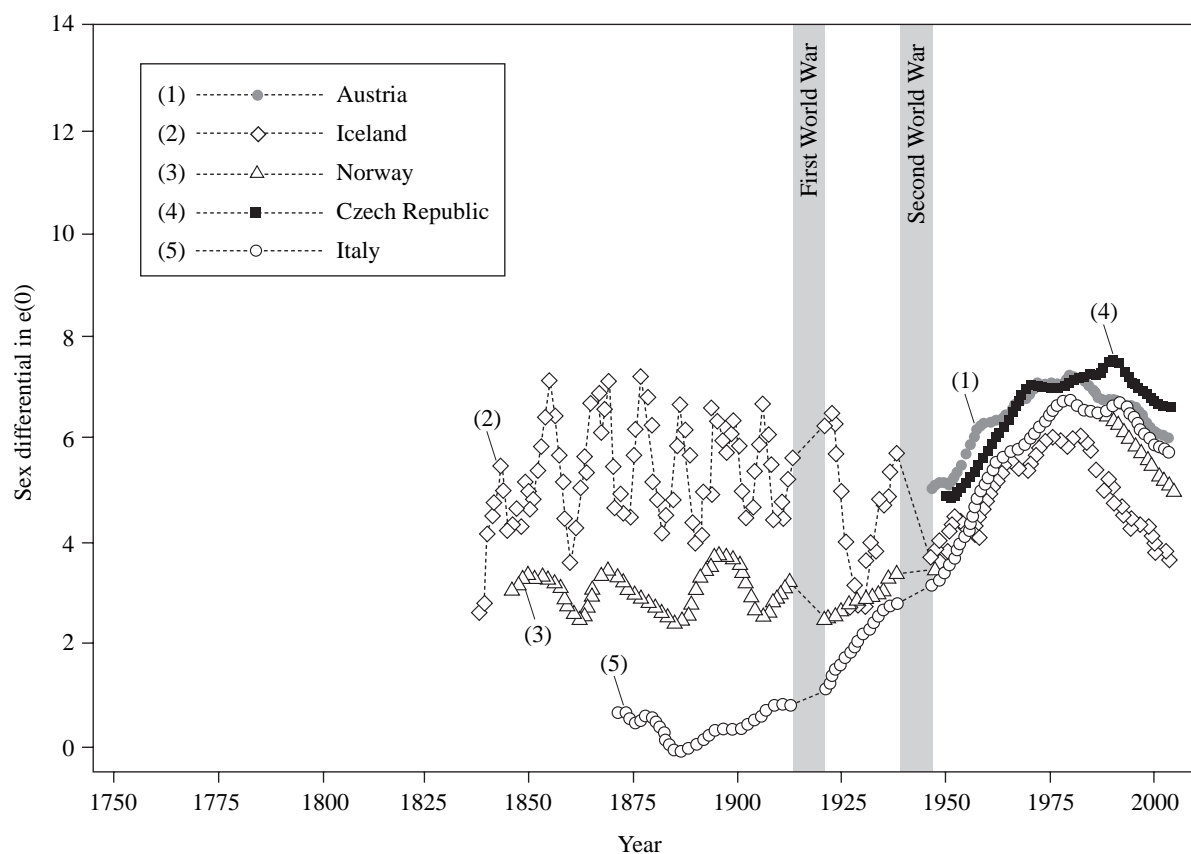
### *Onset of sustained decline in the sex differential in life expectancy*

The first year of a sustained decline in the difference between the sexes in  $e(0)$  for each country is presented in the last column of Table 1. The narrowing trend started in England and Wales in 1972, and was followed in the later 1970s and early 1980s by the other English-speaking countries (USA, Canada, Australia, and New Zealand), Finland, former West Germany, and Sweden. The rest



**Figure 2 (a,b)** Sex differential (female — male) in  $e(0)$  across time and country, based on a 5-year moving average, 1751–2004

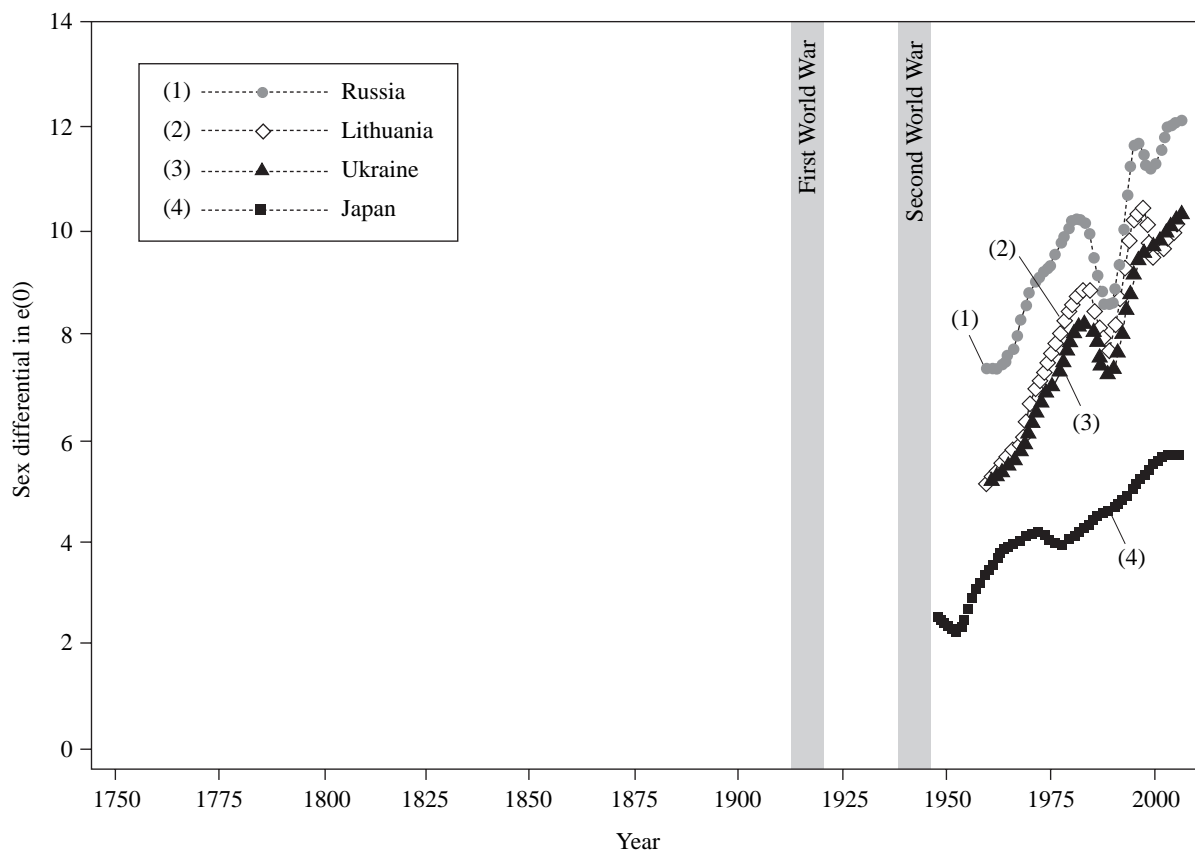
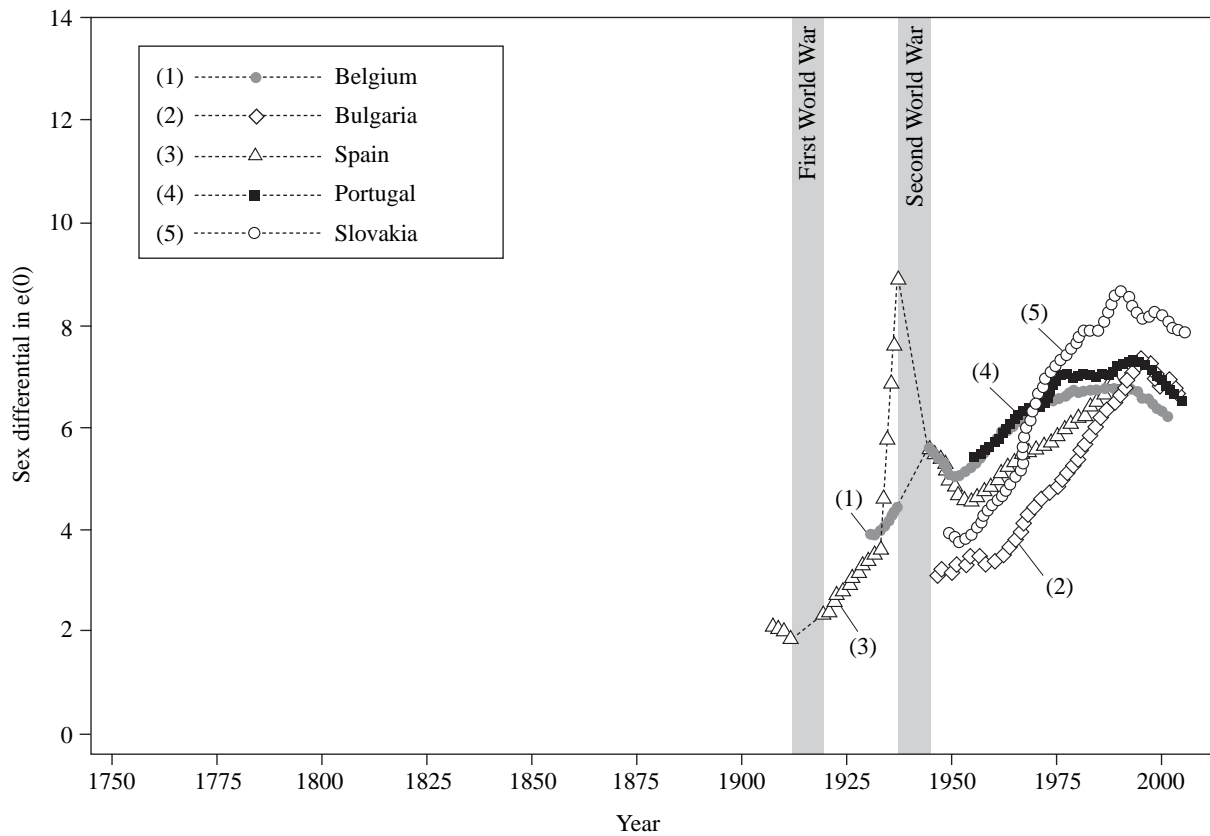
Source: As for Table 1



**Figure 2 (c,d)** Sex differential (female — male) in  $e(0)$  across time and country, based on a 5-year moving average, 1751–2004

Source: As for Table 1





**Figure 2 (e,f)** Sex differential (female – male) in  $e(0)$  across time and country, based on a 5-year moving average, 1751–2004

Source: As for Table 1

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 gap until the 1990s.

The Eastern European countries (Czech Republic, former East Germany, Hungary, Latvia, Bulgaria, and Slovakia) were also late in exhibiting a substantial decline in the difference between the sexes, and in fact the evidence of such a decline has yet to appear for Russia, Lithuania, and Ukraine. 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 gap between the sexes. 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 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, life-table entropy parameters ( $H$  and  $G$ ) are larger for males than females during the period 1975–79, suggesting that the age distribution of deaths is more dispersed for the former than the latter. Conversely, the Gompertzian slope is larger for females than males in every country, indicating that 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 gap between them.

#### *Contributions to the widening vs. narrowing of the sex differential in $e(0)$ by age group*

In this section of the analysis, we decompose by age group the contribution to changes in the gap between the sexes in  $e(0)$  during the period when it was widening (1950–54 to 1975–79) and the period when it began to narrow (1975–79 to 2000–2004). From 1950–54 to 1975–79, the sex differential in  $e(0)$  widened in all 21 countries observed, with an increase ranging from +0.9 years (England and Wales) to +3.9 years (the Netherlands). As shown in Figure 3, those aged 40 and older were the biggest contributors to this increase, especially those aged 60–79. During the same period, the youngest age group (0–19) actually had a narrowing effect on the gap for all but three countries (Italy, Japan, and New Zealand).

During the period 1975–79 to 2000–2004, the gap in  $e(0)$  between the sexes narrowed for 19 of 29 countries, with a decline ranging from –0.3 years (Belgium) to –2.4 years (USA). The biggest contributors to reducing the gap were those aged 40–79 (Figure 4). Other age groups that contributed to widening the gap in the earlier period also contributed to the narrowing in this recent period. In contrast, the mortality of those aged 80 and older continued to have a widening effect on the gap for all but one country (USA). One of the notable differences between countries with the largest reduction in the gap during this period (USA, Canada, Finland) and those with the smallest decrease in the gap (Italy, Czech Republic, Belgium) is the contribution of those aged 80 and older: among the latter countries, the mortality of those at 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.

#### *Changes in the sex differential in $e(0)$ : Effects of sex ratios vs. age pattern of mortality*

Changes in sex ratios in mortality (or equivalently, sex differences in rates of mortality change) certainly affect the absolute difference in  $e(0)$  between the sexes, 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 gap in  $e(0)$  was due to increasing (or decreasing) sex ratios in mortality rates vs. sex differences in the age pattern of mortality.

**Table 2** Life-table entropy ( $H$  and  $G$ ) and Gompertzian slope ( $b$ ) by sex in 1975–79 for the population in each of 29 high-income countries

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
USA	0.16	0.19	12.2	13.6	0.090	0.081

$H$  = 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.

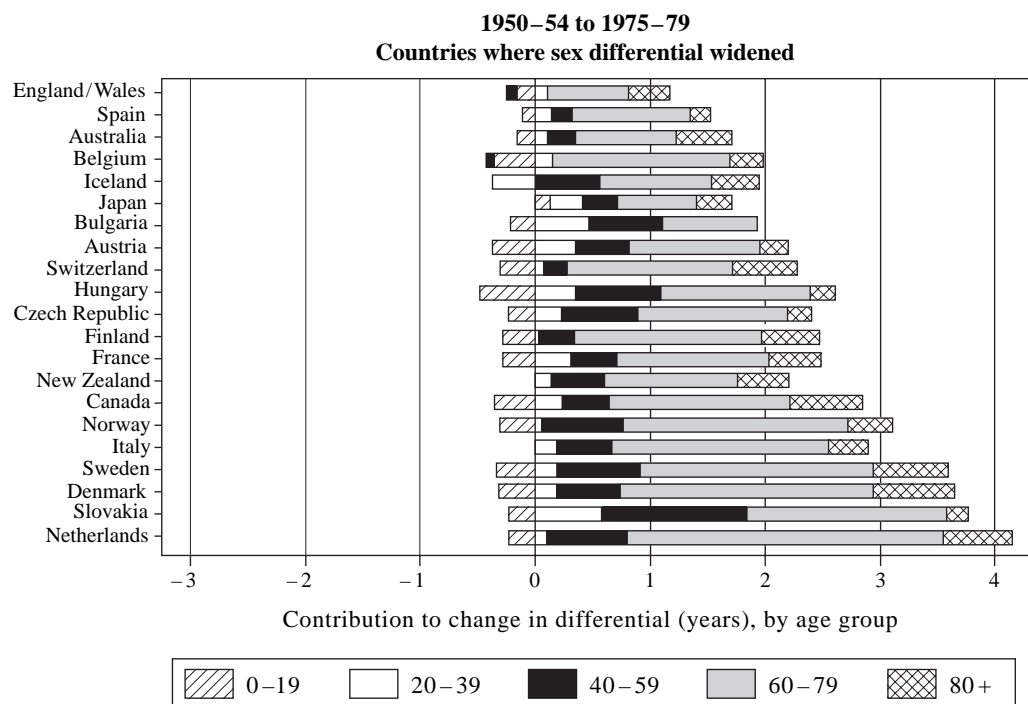
Source: As for Table 1.

Figure 5 reveals that among all 21 countries observed, the widening gap in  $e(0)$  between the sexes from the years 1950–54 to 1975–79 was due entirely to increasing sex ratios in mortality rates. In contrast, sex differences in the age pattern of mortality had a dampening effect on the gap. For example, in Finland, increasing sex ratios contributed 3.8 years to widening the gap, whereas sex differences in the age pattern of mortality reduced the gap by 1.6 years, resulting in a net increase of 2.2 years.

Between the years 1975–79 and 2000–2004, the gap increased further in most Eastern European countries as well as in Spain and Japan. (Among those ten 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 gap (Figure 6).

Yet, Russia and 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 affected more sharply than females by increasing mortality, growing sex ratios in mortality accounted for part of the widening 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 gap would

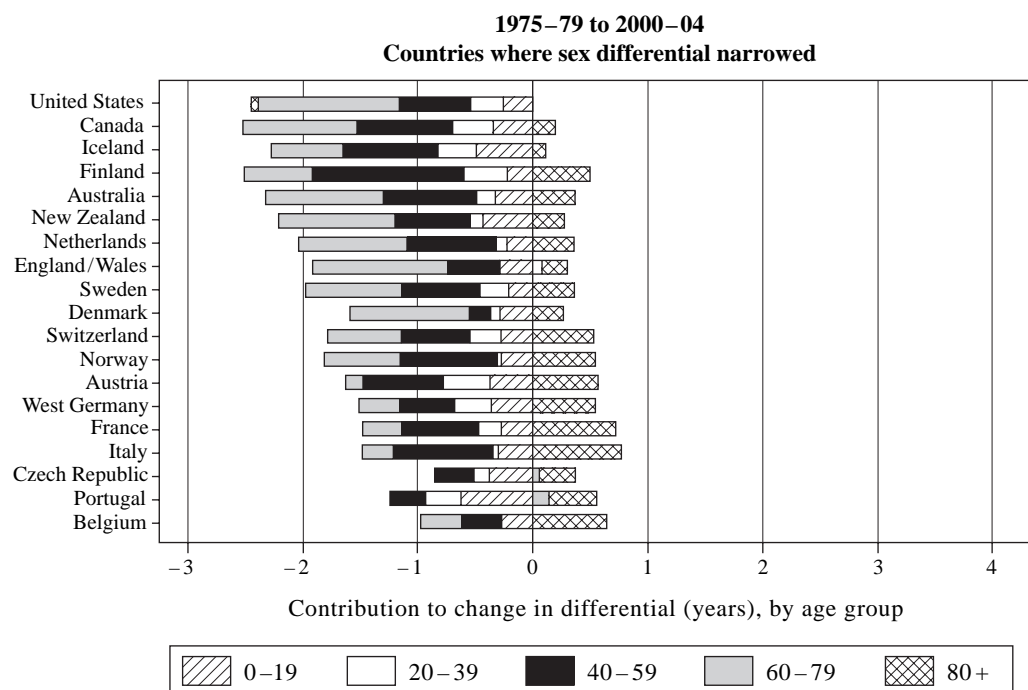


**Figure 3** Contributions to widening of the sex differential in  $e(0)$  by age group, 1950–54 to 1975–79  
Source: As for Table 1

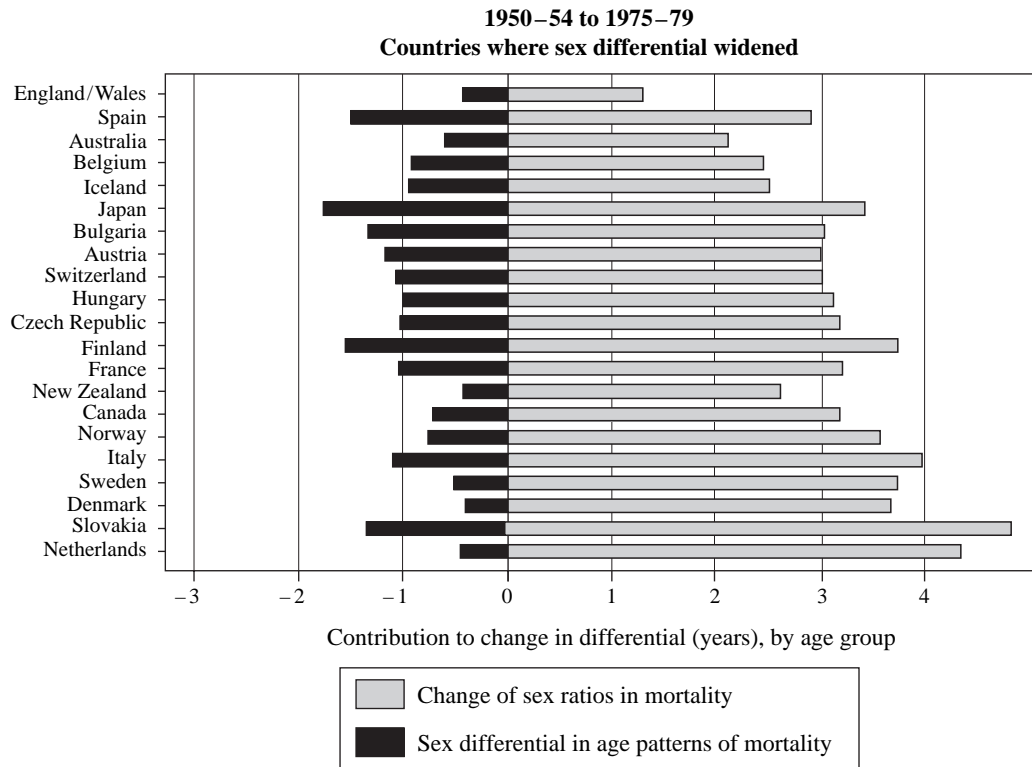
still have widened even if the rate of mortality increase had been the *same* for males and females.

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 gap between the sexes. Nevertheless, we should note that mortality trends varied across these

countries. Whereas in Japan, Spain, and East Germany  $e(0)$  gained more than 5 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



**Figure 4** Contributions to narrowing of the sex differential in  $e(0)$  by age group, 1975–79 to 2000–2004  
Source: As for Table 1



**Figure 5** Widening of sex differential in  $e(0)$  caused by changing sex ratios contrasted with widening caused by differential age patterns, 1950–54 to 1975–79

Source: As for Table 1

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 gap resulting indirectly from large declines in mortality was not sufficient to offset the widening effect of substantial increases in sex ratios in mortality.

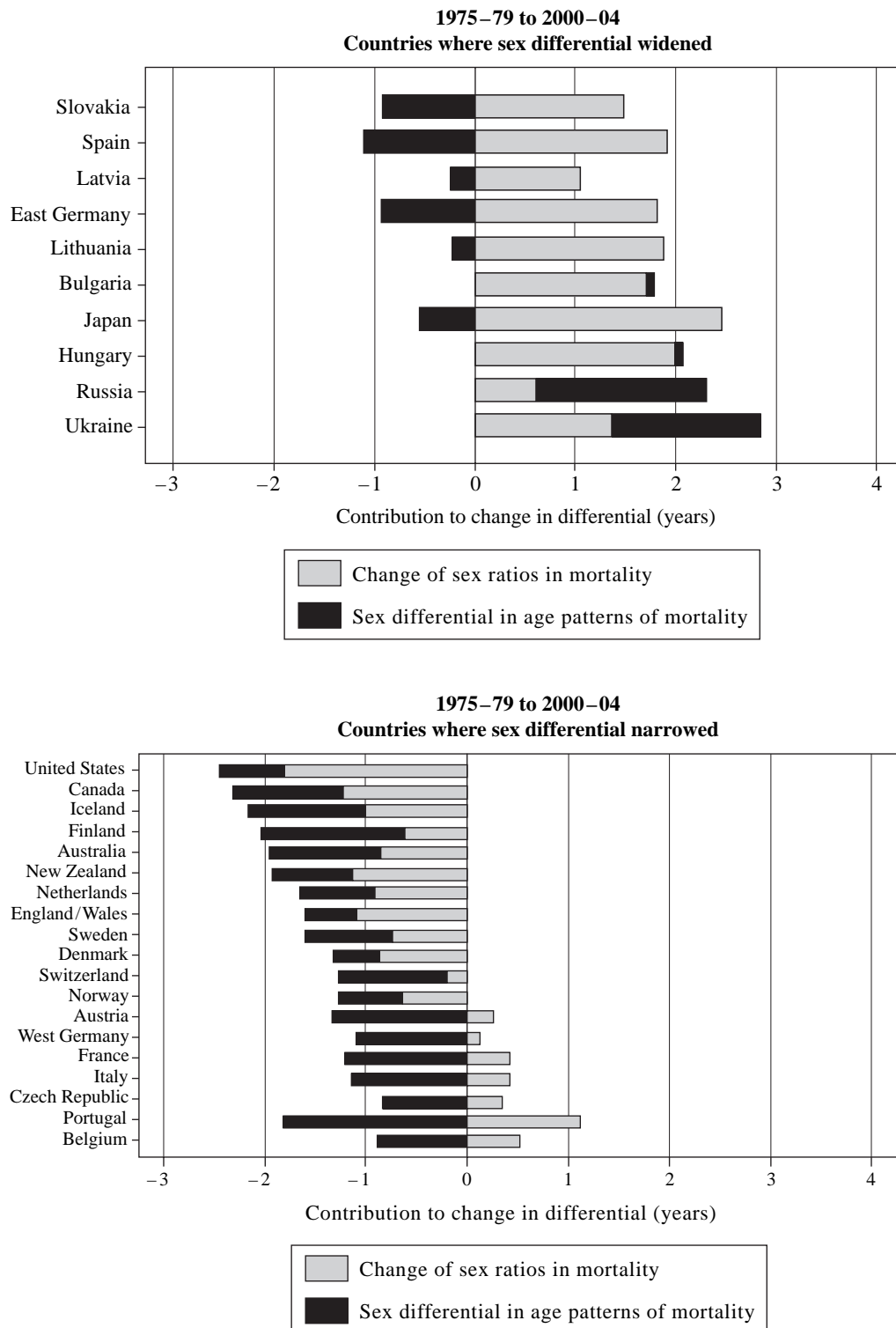
Among the 19 countries where the sex differential narrowed from the mid-1970s, declining sex ratios in mortality account for only part of the reduction in the gap (bottom half of Figure 6). Sex differences in the age pattern of mortality also had a substantial narrowing effect (from  $-0.4$  years in Denmark to  $-1.8$  years in Portugal). In fact, for seven of these countries (Austria, Belgium, Czech Republic, France, Italy, Portugal, and West Germany), the decreased 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 gap was also mostly (84 per cent) due to sex differences in the age pattern of mortality. Among the eleven countries with the biggest declines in the gap, both factors made substantial contributions; declining sex ratios accounted for between 30 per cent (Finland) and 73 per cent (USA) of the reduction in the gap. For

12 of these 19 countries, sex differences in the age pattern of mortality contributed substantially more to the narrowing of the gap than did changes in sex mortality ratios, thereby lending support to the hypothesis that the recent narrowing may be due primarily to sex differences in the age pattern of mortality.

## Discussion

The historical trends in the sex differential 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 the historical widening of the gap among all countries observed here, most of these countries have demonstrated a substantial narrowing of the 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



**Figure 6** Changes in sex differential in  $e(0)$  caused by changing sex ratios contrasted with changes caused by differential age patterns, 1975–79 to 2000–2004

Source: As for Table 1

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 gap in  $e(0)$  varies not only with sex ratios in mortality rates, but also with the level and age pattern of mortality. Our analyses show that although the

widening of the gap was due almost entirely to increasing sex ratios, its widespread narrowing in recent years is explained only in part by declining sex ratios. For the majority of countries, the reduction in the  $e(0)$  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 1 January 2003; HMD 2006). Among the 19 countries where the gap narrowed from the mid-1970s, the USA (where declining sex ratios accounted for 73 per cent of narrowing) accounts for more than 40 per cent of the total population. Altogether, the countries where the majority of the narrowing of the gap resulted from declining sex ratios represent about 60 per cent of the total population.

The reasons for the steeper age-related rise in mortality for females than for males are not fully known. There seem to be at least two plausible explanations for this sex difference. First, unhealthy lifestyles such as smoking, excessive drinking, and occupational hazards 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 mortality slope of males less steep. Second, the health status of women at reproductive ages benefits greatly from their sex hormones, particularly oestrogen. Postmenopausal changes in their hormone levels may make the age-related increase in mortality faster for females than males (Horiuchi 1997). However, assessments of the effects of these and other factors on sex differences in the age pattern of mortality are beyond the scope of this paper and await further investigation.

A few cautionary remarks seem warranted. First, our results do not indicate that changes in sex ratios in mortality played only a minor role in the *reversal* of the trend in the sex differential in  $e(0)$ . 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 in sex ratios in mortality for various countries slowed down, ceased, or reversed, thereby stopping the widening of the  $e(0)$  sex differential. Preston and Wang (2006) demonstrate that sex differences in smoking patterns by cohort could explain this reversal in the USA. However, Figure 6 indicates that for the majority of countries, trends in sex ratios in mortality played a lesser role in narrowing the 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 gap in  $e(0)$ . Those

factors not only directly contribute to the narrowing by lowering sex ratios in mortality for 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 attributed simply to longer life expectancy for females than males. 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, trends in life expectancy in industrialized countries during recent decades are generally linear and do not indicate a tendency for the gain in  $e(0)$  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 the recent narrowing of the gap between the sexes results from differences between them 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 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 gain in life expectancy with mortality decline and mortality variability (Keyfitz 1985), turned out to be the key to understanding the recent narrowing of the gap in  $e(0)$ . Thus, this study clearly demonstrates that trends in the 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 differential 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  $\omega$  is the highest age attained. The life expectancy at birth in the life table at time  $t$  is given by

$$e(0,t) = \int_0^{\omega} l(x,t) dx = \int_0^{\omega} e^{-\int_0^x \mu(y,t) dy} dx. \quad (A1)$$

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

$$\frac{\partial e(0,t)}{\partial \ln \mu(x,t)} = -d(x,t)e(x,t). \quad (A2)$$

The sex differential in  $e(0)$  can be expressed as  $e_f(0,t) - e_m(0,t)$

$$\begin{aligned} &= \int_0^{\omega} e^{-\int_0^x \mu_f(y,t) dy} dx - \int_0^{\omega} e^{-\int_0^x \mu_m(y,t) dy} dx \\ &= \int_0^{\omega} e^{-\int_0^x \alpha(y,t)^{-1/2} \beta(y,t) dy} dx \\ &\quad - \int_0^{\omega} e^{-\int_0^x \alpha(y,t)^{1/2} \beta(y,t) dy} dx \end{aligned} \quad (A3)$$

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

$$\frac{\partial \ln \mu_m(x,t)}{\partial t} = \frac{\partial \alpha(x,t)/\partial t}{2\alpha(x,t)} + \frac{\partial \beta(x,t)/\partial t}{\beta(x,t)} \quad (A4)$$

and

$$\frac{\partial \ln \mu_f(x,t)}{\partial t} = -\frac{\partial \alpha(x,t)/\partial t}{2\alpha(x,t)} + \frac{\partial \beta(x,t)/\partial t}{\beta(x,t)}. \quad (A5)$$

By differentiating equation (A3) with respect to age and making use of equations (A2), (A4), and (A5), the rate of absolute change in the gap in life expectancy can be expressed as

$$\begin{aligned} &\frac{\partial}{\partial t} \{e_f(0,t) - e_m(0,t)\} \\ &= \int_0^{\omega} \left\{ \frac{d_m(x,t)e_m(x,t) + d_f(x,t)e_f(x,t)}{2} \right. \\ &\quad \times \frac{\partial \alpha(x,t)/\partial t}{\alpha(x,t)} dx \\ &\quad + \int_0^{\omega} \{d_m(x,t)e_m(x,t) - d_f(x,t)e_f(x,t)\} \\ &\quad \times \frac{\partial \beta(x,t)/\partial t}{\beta(x,t)} dx. \end{aligned} \quad (A6)$$

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 sex ratios in mortality (i.e., effects of sex differences in the rate of mortality change). The second term indicates effects of changes in the geometric means of the mortality rates for males and females, 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 gap between the sexes because this term would cancel out.

Equation (A6) is closely related to Keyfitz's 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

$$\frac{de(0,t)}{dt} = G(t)\rho(t) \text{ where} \quad G(t) = -\int_0^{\omega} l(x,t)[\ln l(x,t)] dx. \quad (A7)$$

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 (A6), we get

$$\begin{aligned} &\frac{\partial}{\partial t} \{e_f(0,t) - e_m(0,t)\} \\ &= \int_0^{\omega} \{d_f(x,t)e_f(x,t) - d_m(x,t)e_m(x,t)\} \rho(t) dx \\ &= \{G_f(t) - G_m(t)\}\rho(t) \end{aligned} \quad (A8)$$

where  $G_m(t)$  and  $G_f(t)$  are male and female versions of  $G(t)$ . Obviously, equation (A8) can be considered a sex-difference variant of Keyfitz's equation.

### Statistical procedures

Equation (A6) decomposes the change in the 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 gap during relatively long periods, using data by discrete age groups for discrete time periods. Let  $\theta(t)$  represent the gap between the sexes in  $e(0)$  at time  $t$ :  $\theta(t) = e_f(0,t) - e_m(0,t)$ . The change in the 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 (A3), the gap between the sexes can be expressed as a function of  $\alpha$ 's and  $\beta$ 's:

$$\theta(t) = f[\alpha_1(t), \alpha_2(t), \dots, \alpha_n(t), \beta_1(t), \beta_2(t), \dots, \beta_n(t)]. \quad (\text{A9})$$

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

$$\left( \frac{\partial \theta(t)}{\partial \alpha_i(t)} \frac{d\alpha_i(t)}{dt} + \frac{\partial \theta(t)}{\partial \beta_i(t)} \frac{d\beta_i(t)}{dt} \right) \Delta t, \quad (\text{A10})$$

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 gap between  $t_1$  and  $t_2$  can be decomposed:

$$\begin{aligned} \lambda(t_1, t_2) = & \sum_i \int_{t_1}^{t_2} \frac{\partial \theta(t)}{\partial \alpha_i(t)} \frac{d\alpha_i(t)}{dt} dt \\ & + \sum_i \int_{t_1}^{t_2} \frac{\partial \theta(t)}{\partial \beta_i(t)} \frac{d\beta_i(t)}{dt} dt. \end{aligned} \quad (\text{A11})$$

The first term on the right-hand side is the effect of changes in the sex ratios in mortality, and the second term is the overall effect of mortality level (which interacts with sex differences in the age pattern of mortality). The effect of the  $i$ th age group is

$$\int_{t_1}^{t_2} \frac{\partial \theta(t)}{\partial \alpha_i(t)} \frac{d\alpha_i(t)}{dt} dt + \int_{t_1}^{t_2} \frac{\partial \theta(t)}{\partial \beta_i(t)} \frac{d\beta_i(t)}{dt} dt. \quad (\text{A12})$$

We use this method to decompose the change in the gap in  $e(0)$  from each 5-year time period to the next into the effects resulting from 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 resulting from changes in sex ratios and from sex differences in the age pattern of mortality.

Between consecutive 5-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

$$\frac{\ln M_{k,i}(t) - \ln M_{k,i}(t_1)}{\ln M_{k,i}(t_2) - \ln M_{k,i}(t_1)} = \phi(t) \quad (\text{A13})$$

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 per cent of its 'trip' from  $t_1$  to  $t_2$ . This assumption of proportionality is equivalent to the assumption of a constant age pattern of mortality change in the Lee–Carter model (Lee and Carter 1992). It also means that the line integral of equation (A9) is calculated along the straight line from  $[\ln M_{m,1}(t_1), \ln M_{m,2}(t_1), \dots, \ln M_{m,n}(t_1), \ln M_{f,1}(t_1), \ln M_{f,2}(t_1), \dots, \ln M_{f,n}(t_1)]$  to  $[\ln M_{m,1}(t_2), \ln M_{m,2}(t_2), \dots, \ln M_{m,n}(t_2), \ln M_{f,1}(t_2), \ln M_{f,2}(t_2), \dots, \ln M_{f,n}(t_2)]$  in the  $2n$ -dimensional space. Because the proportionality is assumed for changes in a short period (i.e., between two consecutive 5-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 the period 1950–54, the decomposition method was applied to pairs of successive 5-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 ( $5 \times 10$  pairs of successive 5-year time periods). The proportional error of numerical integration was less than 1 per cent for all but two pairs, both of which had virtually no change in the gap between the sexes ( $< 0.1$  years) between the two successive periods and an absolute error no greater than 0.001 years.

## Notes

- 1 Dana A. Gleis is at the University of California, Berkeley. E-mail: danaglei@berkeley.edu. Shiro Horiuchi is at Rockefeller University.
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## References

- Conti, S., G. Farchi, M. Masocco, G. Minelli, V. Toccaceli, and M. Vichi. 2003. Gender differentials in life ex-

- pectancy in Italy, *European Journal of Epidemiology* 18(2): 107–112.
- Elo, Irma T. and Greg L. Drevenstedt. 2005. Cause-specific contributions to sex differences in adult mortality among whites and African Americans between 1960 and 1995, *Demographic Research* 13(19): 485–520.
- Gjonca, Arjan, Cecilia Tomassini, Barbara Toson, and Steve Smallwood. 2005. Sex differences in mortality, a comparison of the United Kingdom and other developed countries, *Health Statistics Quarterly* 26(Summer): 6–16.
- Goldman, Noreen and Graham Lord. 1986. A new look at entropy and the life table, *Demography* 23(2): 275–282.
- HMD (Human Mortality Database). 2006. University of California, Berkeley (USA), and Max Planck Institute for Demographic Research (Germany). Available: [www.mortality.org](http://www.mortality.org) (accessed 10 October 2006).
- Horiuchi, Shiro. 1997. Postmenopausal acceleration of age-related mortality increase, *Journal of Gerontology, Series A, Biological Sciences and Medical Sciences* 52(1): B78–B92.
- Horiuchi, Shiro, John R. Wilmoth, and Scott D. Pletcher. 1999. A general method of decomposition analysis. Paper presented at the Annual Meeting of the Population Association of America, New York, 25–27 March.
- Horiuchi, Shiro, Caleb Finch, France Meslé, and Jacques Vallin. 2003. Differential patterns of age-related mortality increase in middle age and old age, *Journal of Gerontology, Series A, Biological Sciences and Medical Sciences* 58(6): 495–507.
- Keyfitz, Nathan. 1985. *Applied Mathematical Demography*, 2nd edn. New York: Springer-Verlag.
- Lee, Ronald D. and Lawrence Carter. 1992. Modeling and forecasting the time series of US mortality, *Journal of American Statistical Association* 87(419): 659–671.
- Lopez, Alan D. 1983. The sex mortality differentials in developed countries, in A. D. Lopez and L. T. Ruzicka (eds.), *Sex Differentials in Mortality: Trends, Determinants and Consequences*. Canberra: Australian National University, pp. 53–120.
- Meslé, France. 2004. Differences in life expectancy between the sexes: why is it reducing? Paper presented at Human Mortality over Age, Time, Sex, and Place: The 1st HMD Symposium, Rostock, Germany, 18–19 June. Co-sponsored by the Max Planck Institute for Demographic Research and the Center on the Economics and Demography of Aging, University of California, Berkeley. [Published in French in *Revue d'épidémiologie et de sante publique*.]
- Nault, François. 1997. Narrowing mortality gaps, 1978 to 1995, *Health Report* 9(1): 35–41.
- Nusselder, Wilma J. and Johan P. Mackenbach. 1996. Rectangularization of the survival curve in the Netherlands, 1950–1992, *The Gerontologist* 36(6): 773–782.
- Olshansky, S. Jay, B. A. Carnes, and C. Cassel. 1990. In search of Methuselah: estimating the upper limits of human longevity, *Science* 250(4981): 634–640.
- Pampel, Fred C. 2002. Cigarette use and the narrowing sex differential in mortality, *Population and Development Review* 28(1): 77–104.
- Pampel, Fred C. 2003. Declining sex differences in mortality from lung cancer in high-income nations, *Demography* 40(1): 45–65.
- Pletcher, Scott D., Aziz A. Khazaeli, and James W. Curtsinger. 2000. Why do lifespans differ? Partitioning mean longevity differences in terms of age-specific mortality parameters, *Journal of Gerontology, Series A, Biological Sciences and Medical Sciences* 55(8): B381–B389.
- Pollard, J. H. 1982. The expectation of life and its relationship to mortality, *Journal of the Institute of Actuaries* 109(1982): 225–240.
- Preston, Samuel H. 1970. *Older Male Mortality and Cigarette Smoking: A Demographic Analysis*. Berkeley: Institute of International Studies, University of California.
- Preston, Samuel H. and Haidong Wang. 2006. Sex mortality differentials in the United States: the role of cohort smoking patterns. Paper presented at the Annual Meeting of the Population Association of America, Los Angeles, 29 March–1 April.
- Stolnitz, George J. 1956. A century of international mortality trends: II, *Population Studies* 10(1): 17–42.
- Tabutin, Dominique and Michel Willems. 1998. Differential mortality by sex from birth to adolescence: the historical experience of the West (1750–1930), in *Too Young to Die: Genes or Gender?* New York: United Nations, pp. 17–52.
- Trovato, Frank and N. M. Lalu. 1996. Narrowing sex differentials in life expectancy in the industrialized world: early 1970's to early 1990's, *Social Biology* 43(1/2): 20–37.
- Trovato, Frank and N. M. Lalu. 1998. Contribution of cause-specific mortality to changing sex differences in life expectancy: seven nations case study, *Social Biology* 45(1/2): 1–20.
- Trovato, Frank and Nils B. Heyen. 2006. A varied pattern of change of the sex differential in survival in the G7 countries, *Journal of Biosocial Science* 38(3): 391–401.
- United Nations Secretariat. 1988. Sex differentials in life expectancy and mortality in developed countries: an analysis by age groups and causes of death from recent and historical data, *Population Bulletin*

- of the United Nations, No. 25-1988. New York: United Nations.
- Valkonen, Tapani and Frans Van Poppel. 1997. The contribution of smoking to sex differences in life expectancy: four Nordic countries and the Netherlands 1970–1989, *European Journal of Public Health* 7(3): 302–310.
- Vallin, Jacques. 1993. Social change and mortality decline: women's advantage regained or achieved?, in Nora Federici, Karen Oppenheim-Mason, and Solvi Sogner (eds.), *Women's Position and Demographic Change in the Course of Development*. Oxford: Clarendon Press, pp. 190–212.
- Vaupel, James W. 1986. How change in age-specific mortality affects life expectancy, *Population Studies* 40(1): 147–157.
- Vaupel, James W. and Vladimir Canudas-Romo. 2000. How mortality improvement increases population growth, in E. J. Dockner, R. F. Hartl, M. Luptacik, and G. Sorger (eds.), *Optimization, Dynamics and Economic Analysis: Essays in Honor of Gustav Feichtinger*. Vienna: Springer, pp. 350–357. Available: [www.demogr.mpg.de/Papers/Working/wp-1999-015.pdf](http://www.demogr.mpg.de/Papers/Working/wp-1999-015.pdf) (accessed 21 February 2006).
- Waldron, Ingrid. 1985. What do we know about causes of sex differences in mortality? A review of the literature, *Population Bulletin of the United Nations* (18): 59–76.
- Waldron, Ingrid. 1986. The contribution of smoking to sex differences in mortality, *Public Health Reports* 101(2): 163–173.
- Waldron, Ingrid. 1993. Recent trends in mortality ratios for adults in developed countries, *Social Science and Medicine* 36(4): 451–462.
- Waldron, Ingrid. 1995. Contributions of biological and behavioral factors to changing sex differences in ischaemic heart disease mortality, in A. D. Lopez, G. Caselli, and T. Valkonen (eds.), *Adult Mortality in Developed Countries: From Description to Explanation*. Oxford: Clarendon Press, pp. 161–178.
- Waldron, Ingrid. 2005. Gender differences in mortality: causes and variation in different societies, in P. Conrad (ed.), *The Sociology of Health and Illness-Critical Perspectives*. New York: Worth, pp. 38–55.
- White, Kevin M. 2002. Longevity advances in high-income countries, 1955–96, *Population and Development Review* 28(1): 59–76.
- Wilmoth, John R. and Shiro Horiuchi. 1999. Rectangularization revisited: variability of age at death within human populations, *Demography* 36(4): 475–495.
- Wilmoth, John R., Leo J. Deegan, Hans Lundström, and Shiro Horiuchi. 2000. Increase in maximum life span in Sweden, 1861–1999, *Science* 289(5488): 2366–2368.
- Wilmoth John R., Kirill Andreev, Dmitri Jdanov, and Dana A. Gleit (with Carl Boe, Michael Bubenheim, Dimiter Philipov, Vladimir Shkolnikov, and Pierre Vachon). 2005. Methods protocol for the Human Mortality Database. Available: [www.mortality.org/Docs/MethodsProtocol.pdf](http://www.mortality.org/Docs/MethodsProtocol.pdf) (accessed 6 December 2005).