An extended decomposition of change in the sex gap in life expectancy in developed countries: period and cohort perspectives

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Abstract

To achieve a deeper understanding of the sex difference in mortality, this paper extends the existing decomposition methods on the changing sex gap in life expectancy. The new formula separates the change in sex gap in life expectancy into three components that capture: the effect from the sex difference in mortality improvement (ρ-effect), the effect from the sex difference in the life table distribution of deaths (f-effect), and the effect from the sex difference in remaining life expectancy (e-effect). These three components correspond to a comparison in speed of mortality reduction, dispersion of death distribution and survival advantage, respectively. Both period and cohort perspectives have been analysed in our study. To avoid the issues limited mortality data series causes, the truncated cohort life expectancy at birth is used for the cohort comparison. The results indicate that period and cohort analyses uncover different stories on the change in sex gap in life expectancy phenomenon in terms of these three effects.

Supplementary material for this article is available at: XXXXXXXXXXXXX

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Introduction

It is widely accepted that females have a higher life expectancy at birth than males (Barford et al. 2006). The sex difference in life expectancy has been variously studied in demography, actuarial studies, epidemiology and biology, following the pioneering work on sex differences in ill-health and mortality published by John Graunt in 1662. Indicative of wider interest, the topic attracted public attention in the twentieth century (Nathanson, 1984). Research into the underlying causes of sex-differential mortality indicates that social, environmental, biological, genetic and behavioural factors and their interactions all play a role (Lindahl-Jacobsen et al. 2016; Lopez, 1995; Luy, 2003; Madigan, 1957; Preston and Wang, 2006; Waldron, 1983, 1986, 1995; Wingard, 1982, 1984; Zarulli et al. 2017).

Many industrialised countries have experienced a similar trend in the sex gap in life expectancy at birth: from the 1950s to the 1970s the gap widened, but from the 1980s it has narrowed. This pattern has been used to inform analyses in relation to the epidemiologic transition, cardiovascular revolution, and behavioural changes (Booth et al. 2016; Bourbeau and Ouellette, 2016; Glei and Horiuchi, 2007; Pampel, 2002; Trovato and Lalu, 1997). To identify the age and cause of death contributions to the sex gap in life expectancy, many studies have used decomposition methods (Booth, 2003; Booth et al. 2016; Canudas-Romo et al. 2015; Tickle, 2016; Trovato and Lalu, 2007; Vallin and Meslé, 2001), like Andreev (1982) stepwise replacement decomposition, Arriaga (1984) discrete decomposition and Pollard (1982, 1988) continous decomposition. Through these decomposition results, researchers could describe the change of sex gap in life expectancy, but how to formulate and decompose the change in sex difference in life expectancy has not been fully explored. Relatively few studies have addressed this issue and measure the contributors to the time change in the sex gap in life expectancy (Canudas-Romo, 2003; Glei and Horiuchi, 2007; Zhang and Vaupel, 2008), and these all use methods developed by, or derived from, Vaupel and Canudas-Romo (2003). This paper further develops this approach and applies it to a wide range of populations with a view to enhancing understanding of the sex gap mortality transition.

Besides the demands on the sharper method, the phenomenon of the changing sex gap in life expectancy has yet to be fully investigated from the period and cohort perspectives. Existing research on the changing sex difference in mortality either focuses on only one perspective or uses age specific indicators (Beltrán-Sánchez et al. 2015; Glei and Horiuchi, 2007; Trovato and Lalu, 1996; Lindahl-Jacobsen et al. 2013). In fact, many studies have

found that cohort trends in demographic measures do not fully match period trends (Bongaarts and Feeney, 2002; Goldstein and Wachter, 2006; Kjærgaard and Canudas-Romo, 2017). Therefore, to get a better understanding of the change in sex gap in life expectancy phenomenon, we present both period and cohort perspectives.

In the following section, we introduce relevant existing decomposition methods for the life expectancy at birth and the derivation of the extended decomposition method, including that for the truncated cohort life expectancy. We then introduce the data used in the analysis. The results consist of period and cohort decompositions of temporal change in the sex gap. Finally, we discuss the findings and their implications.

Development of the extended decomposition method

There are three existing methods for decomposing change in the sex gap in life expectancy at birth: Canudas-Romo (2003), Glei and Horiuchi (2007), and Zhang and Vaupel (2008). Though these use different terms, all three methods are in fact derived from the Vaupel and Canudas-Romo (2003) decomposition method. The three methods are introduced, and their mathematical similarities and differences discussed, in the supplementary material. Among the three methods, the Glei and Horiuchi method is the most often cited and most influential. In this method, the mortality age pattern is determined by the slope of the Gompertz model. Glei and Horiuchi (2007) argue that since female mortality has a steeper Gompertz slope than male mortality, female adult mortality will increase more rapidly than male, leading to males having a larger gain in life expectancy at birth than females for the same rate of mortality reduction.

There are several limitations of the Glei and Horiuchi approach. First, the Gompertz slope is derived from adult mortality and is inappropriate for infant, child and old age mortality (Makeham, 1860; Siler, 1979) and for the 'accident hump' at young adult ages (Mazzuco et al. 2018). Second, a fixed Gompertz slope cannot capture fluctuations in empirical mortality change over age, i.e., in the life table ageing rate (LAR) (Horiuchi and Coale, 1990). Third, we note that the Gompertz slope is the inverse of life span disparity, denoted $e^{\dagger}(0,t)$, which is the age-aggregate of the product of remaining life expectancy, *e(a, t)*, and the distribution of life table deaths, *f(a, t)*, (Vaupel and Canudas-Romo, 2003). While $e^{\dagger}(0,t)$ is widely used (Nusselder and Mackenbach, 1996; Robine, 2001; Shkolnikov et al. 2011; Vaupel et al. 2011; Wilmoth and Horiuchi, 1999), it is affected by change in both

of its components, *e(a, t)* and *f(a, t)*. By using the Gompertz slope, the different effects of these components are not made explicit. Therefore, we prefer to further decompose the original Vaupel and Canudas-Romo (2003) decomposition.

Vaupel and Canudas-Romo decomposition

Vaupel and Canudas-Romo (2003) developed a method for the decomposition of change over time in life expectancy. Their method involves three components: the rate of mortality improvement, $\rho(a,t)$, the distribution of life table deaths, $f(a,t)$, and the remaining life expectancy, $e(a,t)$, where a denotes age and t denotes time. The change over time, or the time-derivative, is denoted by a superscript dot and the subscript t corresponding to time, the Vaupel and Canudas-Romo decomposition of the change in life expectancy is,

$$\dot{e}_t(0,t) = \int_0^{\omega} \rho(a,t)\, f(a,t)\, e(a,t) da, \qquad (1)$$

where ω is the highest age at death. The rate of mortality improvement is the time change in the force of mortality, $\mu(a,t)$, or its relative derivative with respect to time: $\rho(a,t) = -\frac{\partial \mu(a,t)}{\mu(a,t)\partial t}$. The remaining two components, the distribution of deaths and remaining life expectancy, are functions of the life table which also depend on the force of mortality (Preston et al. 2001). However, only the rate of mortality improvement represents change over time. Despite this, equation (1) shows that the final effect on change in life expectancy is also determined by the shape of the distribution of deaths, and subsequent survival or remaining life expectancy. Thus, these three components of the change in life expectancy are crucial for a full understanding the dynamics of mortality.

Extended decomposition of change in the sex gap

The new decomposition of the change over time in the sex gap in life expectancy is based on the three components of equation (1). The new decomposition method first applies equation (1) to the time change in life expectancy for each sex *s*, $\dot{e}_t(0,t,s)$, obtaining three components $\rho(a,t,s)$, $e(a,t,s)$ and $f(a,t,s)$ for each sex. These time changes in male and female life expectancy are then compared by taking the partial derivative with respect to *s*:

$$\dot{e}_{t,s}(0,t) = \frac{\partial \dot{e}_{t,s}(0,t,s)}{\partial s} = \frac{\partial}{\partial s}\int_0^{\omega} \rho(a,t,s)\, f(a,t,s)\, e(a,t,s)da$$

$$= \int_0^{\omega} [\dot{\rho}_s(a,t,s)\, f(a,t,s)\, e(a,t,s) + \rho(a,t,s)\, \dot{f}_s(a,t,s)\, e(a,t,s)$$

$$+\rho(a,t,s)\, f(a,t,s)\, \dot{e}_s(a,t,s)]da. \qquad (2)$$

The right-hand side of equation (2) has three components, each including a partial derivative with respect to sex: the first term captures the effect of the sex difference in the rate of mortality improvement, *ρ(a,t,s)*, and is called the "*ρ-effect*"; the second term captures the effect of the sex difference in the life table death distribution, *f(a,t,s)*, and is called the "*f-effect*"; and the third term captures the effect of the sex difference in remaining life expectancy, *e(a,t,s)*, and is called the "*e-effect*". The total effect, the sum of the three terms, captures the slope of the change in the sex gap in life expectancy.

This extended method is used for decomposing the time change of the sex gap (as illustrated by the green line) in Figure 1. When the sex gap widens (the green line shows an increasing tendency), the total effect in equation (2) will be positive. When the sex gap narrows (the green line shows a decreasing trend) the total effect will be negative. At the turning point (the green dot in Figure 1) the total effect equals to zero and the components on the right of equation (2) offset each other.

[*Figure 1 about here*]

The *ρ-effect* captures the differential speed of mortality reductions between females and males. A positive *ρ-effect* indicates females have a faster mortality decline than males; and a negative *ρ-effect* indicates the reverse. The *f-effect* quantifies to what extent the change in the sex gap in life expectancy can be explained by the sex difference in the life table deaths distribution. It is an indicator of the disparity in age at death or variability at death, between the two sexes. A positive *f-effect* indicates females have more dispersed death distribution, while a negative means male death distribution is more dispersed than female. The *e-effect* is the survival advantage indicator. A positive *e-effect* indicates females have more cumulative advantages than males in survivorship. It should be noted that when both sexes experience mortality increases, or one sex has an overwhelmingly mortality increase, leading to negative *ρ* values, the direction of the *f-effect* and *e-effect* are inverted, which indicates that when the

mortality increases (negative ρ), the more dispersed death distribution population (male) and the more survival advantage population (female) will play the widening and narrowing sex gap role, respectively (for detailed age contribution see the supplementary material). Nevertheless, the interpretation of the latter effects remains as above. These situations can easily be identified by looking at the direction of *e-effect*, which in general should be positive, since females have more cumulative survival advantage than males.

Using continuous equations and notations for the decomposition analysis is more succinct than the discrete expression. Similar mathematical logic could be found in other articles with decomposition analysis (Bergeron-Boucher et al. 2015; Mogi and Canudas-Romo, 2018). Following this notion, we assume all the measures in equation (2) to be functions of continuous variables. The derivative of these measures with respect to its continuous variable, be this time or sex, is comparable to the stepwise replacement decomposition suggested by Andreev and his colleagues (Andreev, 1982; Andreev et al. 2002): estimating one differential component and keeping the rest fixed in a stepwise replacement procedure. Thus, we can consider the derivative of a measure as the transition from one population to another: $\dot{e}_t(0,t,s)$ as the transition for a population at two given times, or $\dot{e}_s(0,t,s)$ the transition between females and males. Other methods using derivatives and assuming continuous expressions of variables have shown small discrepancies in their estimations (Canudas-Romo and Guillot, 2015; Horiuchi et al. 2008). The methods for estimating equation (2) for data available at discrete times are discussed in the supplementary material.

The extended method in equation (2) can also be applied to cohort data. Furthermore, it can be applied to any age range. We use the truncated cohort life expectancy at birth, which is defined as,

$$ {}_Te_0^c = \int_0^T l(x,c)dx = \int_0^T e^{-\int_0^x \mu(a,c+a)da}\, dx, \qquad (3) $$

where ${}_Te_0^c$ is the truncated cohort life expectancy at birth. The subscript T is the age at truncation, c denotes birth cohort, and $\mu(a,c+a)$ is the force of mortality at age a and time $a+c$ for birth cohort c. Note that $l(0,c)=1$ is the radix of the life table. Analogous to steps in the period analysis, first the change over cohort in truncated life expectancies for both of the sexes is calculated, followed by the comparison between sexes. For example, the changing sex gap in truncated cohort life expectancies at age 50, or ${}_{50}e_0^c$, for those born in

1950 versus those in 1960 are decomposed. The supplementary material includes the detailed procedures for the discrete approach.

Data

The data for the analyses were obtained from the Human Mortality Database (HMD, 2018). The HMD contains high quality historical data combining vital statistics and census counts or official population estimations in both period and cohort perspectives. 13 countries have data for more than 100 years and 10 countries have data back to at least 1950. Standard methods applied to all populations over time allow comparisons over these two dimensions (Wilmoth et al. 2017). These high quality data, long time series and standard methods of HMD enable us to do the international comparison and the construction of truncated cohort life expectancies.

For the period data, we use single-year period life table data which starts from 1950 to 2010. The reasons for this data selection are: First, prior to 1950, the sex gap in life expectancy was dramatically enlarged by the two World Wars (1914-1918 and 1939-1945) and the Spanish Flu pandemic (1918-1920), as well as some period events such as the Spanish Civil War (1936-1939). Our decomposition method could capture these changes, showing the considerabe peak around the year that such events happened, however, our aim is at overall time trends. Second, restriction to the post-WWII era increases the number of populations meeting data requirements, with a missmatch of available information before that. Countries with less than 40 years of data were excluded to ensure a time series of adequate length, and countries with a total population of less than one million in 2010 were omitted so as to avoid substantial random fluctuations. This produced a total of 33 populations for inclusion in our analysis (see Table 1).

For the cohort analysis, we employed truncated life expectancies so as to avoid the problem of incomplete cohort data. Age-specific death rates from the single-year period life table were used to construct series of cohort mortality rates. Comparison with HMD cohort data for those countries with long enough series confirmed that using period data along the diagonal accurately matched cohort mortality (for the detailed see the supplementary material). To assure meaningful international comparison, we examined truncated cohort life expectancies from birth to ages 50, 70 and 90, denoted ${}_{50}e_0^c$, ${}_{70}e_0^c$, and ${}_{90}e_0^c$, for annual cohorts born in 1880 to 1960. Cohort data were constructed for 21 of the 33 countries (see Table 1):

12 were excluded because of data gaps or having less than 50 years of cohort data prior to 2000; these were Belgium, Belarus, East Germany, West Germany, Estonia, Ireland, Latvia, Lithuania, Taiwan, Poland, Russia, and Ukraine.

[Table 1 about here]

To reduce the fluctuation caused by no deaths at some young or old ages, the 5-year period life tables were used (HMD, 2018) in our period decomposition analysis. And, to generalise the changing sex gap pattern, in the cohort part, we combine the decomposition results by five birth cohorts, after decomposing the changes across each birth cohort.

Results

The period and cohort perspectives of changing sex-gap patterns are provided independently, followed in each case by their extended decomposition results. We grouped the analysed countries based on their turning point of the sex gap patterns. To simplify the results, we highlight four representative countries for the period perspective and two countries for the cohort perspective. The results for the rest of the countries are presented in the supplementary material.

Period perspective

The period sex gap in life expectancy is presented in Figure 2. Almost all countries experienced widening and narrowing patterns for the sex gap in life expectancy at birth from 1950-1954 to 2005-2009, except for Japan, Taiwan, and some former Soviet Union countries. The turning point in the sex gap, seen as the onset of sustained reduction in sex gap in life expectancy in Figure 1, was used to regroup countries in Figure 2. Then, based on the average of individual country's turning points, four changing sex-gap patterns were found: Classical, Delayed, Eastern Europe, and Eastern Asia. In the supplementary material, the four changing sex-gap patterns (Figure S-3) and the turning point for each country are provided (Table S-1). A similar pattern-grouping of countries was found by Liu et al. (2012) only varying slightly the turning points of the groups, probably caused by the different databases used.

The Classical pattern consists of 13 countries including the Scandinavian (Denmark, Finland, Norway, and Sweden) and Anglophone countries (Australia, Canada, New Zealand, England and Wales, Ireland and USA). In this pattern, the turning point of the sex gap in life expectancy at birth occurred around 1980. This group of countries has a converging trend, where the sex gaps had a wide distribution ranging from 2 to 7 years in 1950 narrowing down from 4 to 7 years in 2010. The Delayed pattern includes 12 countries, predominantly from Central and Southern Europe, e.g. Czech Republic, East Germany, Hungary, Italy, Portugal, Spain, and Switzerland. The turning point occurred around 1994, which is around a 14-year lag from the classical pattern. The distribution of the number of years in the sex gap remains at 4 years difference, ranging from 3 to 6 years in 1950 to 4 to 9 years in 2010. The Eastern Europe pattern is composed of 6 former Soviet Union countries, including Belarus, Estonia, Latvia, Lithuania, Russia, and Ukraine. The narrowing sex gap started very recently from 2000s and with a relatively higher sex gap compared with other groups. This cluster of countries has a great fluctuation from the late 1980s to the mid-1990s. Finally, the Eastern Asia pattern includes Japan and Taiwan, whose sex gaps have an increasing trend during the almost entire period of observation. After 2005, in both Japan and Taiwan, the sex gaps in life expectancy at birth have a slight decrease. For the spatial distribution of these group see Figure S-4 in the supplementary material.

[Figure 2 about here]

Based on the four changing sex-gap patterns, we selected England and Wales, France, Japan, and Ukraine to represent the Classical, Delayed, Eastern Asia and Eastern Europe patterns, respectively. Figure 3 presents the changing sex-gap decomposition of equation (2) for those four selected countries and the decompositions for the rest of the countries are provided in the supplementary material.

As seen in the line of the total change in Figure 3, the narrowing sex-gap phenomenon starts earlier in England and Wales followed by France. In Japan, this narrowing sex-gap phenomenon just occurred in the latest period and in Ukraine great fluctuations are observed from the mid-1980s to the mid-1990s. Figure 3 also includes the three components of our extended decomposition method in equation (2). In general, the mortality speed indicator (ρ-*effect*) shows a reverse direction pattern around the turning point, while the death variability indicator (*f-effect*) and cumulative survivorship indicator (*e-effect*) are relatively constant and remain negative and positive respectively during the entire period. In England and Wales,

France, and Japan, the *f-effect*, is negative, and has a narrowing sex gap role because male's life table deaths distribution is more dispersed than females; the *e-effect* is positive and widens the sex gap, because females have more survivorship advantages than males.

In England and Wales (Classical pattern), when the sex gap widens (1950-70), the positive *ρ-effect* is the dominant effect and contributes more than 70 per cent of widening sex gap trend. This indicates females have a faster mortality decline. However, when the sex gap narrows (1970-2010), the *ρ-effect* becomes negative, which suggested that males have faster mortality reduction than females, and the *f-effect* becomes the main contributor, indicating the more dispersed death distribution (male) could lead more potential to increase the life expectancy, narrowing the sex gap. It should be highlighted that the absolute *ρ-effect* has an increasing (1970-2005) (-0.01 to -0.35) and then a decreasing (2005-10) (-0.35 to -0.26) pattern. In France (Delayed pattern), similar to England and Wales, when the sex gap widens (1950-95) the positive *ρ-effect* is the major contributor, and when the sex gap narrows, the total trends are largely contributed to the negative *f-effect*. In contrast with the Classical pattern, in France (Delayed pattern) after 2005, the *ρ-effect* starts to increase from -0.17 to -0.19 years. The French trends of the three components follow closely, but with a time lag, the trends observed in England and Wales: i.e. the narrowing sex gap from 1995 to 2010 has similar pattern with the early narrowing period (1975-1990) in England and Wales.

Japan (Eastern Asia pattern) also have the same dominant effects as the Classical and delayed pattern in terms of changing sex gap. The *ρ-effect* is positive from 1950 to 2010, which indicates that females have more advantages in mortality reduction than males. However, the narrowing sex gap observed in 2005-2010, is driven by a large negative *f-effect*, which contributes -0.40 years. Finally, Ukraine (Eastern Europe pattern) has a different sex gap trend compared with the other countries. Almost all directions of *f-effect* and *e-effect* are inverted from 1960 to 2010, resultant of some years with mortality increase for both sexes and others of males have significant mortality increase compared to females. It should be emphasised that in the late 1980s, the *ρ-effect* is negative, opposing the neighbouring periods and contributing -0.50 years to the total trend (or 48 per cent of the total change).

[Figure 3 about here]

Cohort perspective

Figure 4 shows the changing sex-gap pattern of the truncated life expectancies between birth and ages 50, 70 and 90, across 21 countries. The observed patterns could be roughly divided

into two groups, based on the patterns that the nine long mortality series countries presented. One group (England and Wales, France, and Italy) has two evident peaks for the sex gap in ${}_{50}e_0^c$ and ${}_{70}e_0^c$. The two peaks are concentrated around the 1895 and 1925 birth cohorts. For the sex gap in ${}_{90}e_0^c$, the peak for the 1895 birth cohort is visible but we do not have full data series for the 1920s birth cohorts and cannot see the second peak. England and Wales was selected as the representative country for this two-peak group. The other group (Denmark, Finland, Netherlands, Norway, Switzerland and Sweden) has one small peak located around the 1920s birth cohort for the changing sex gap in ${}_{50}e_0^c$ and ${}_{70}e_0^c$; but the sex gap in ${}_{90}e_0^c$ shows a monotonous increasing trend from 1880 to 1920 birth cohort. We selected Sweden as the representative country for this one-peak group. Besides the difference in the peak between the two groups of countries, the magnitude of sex gap in the two-peak group is larger than in the one-peak group for all truncated life expectancies. Finally, in England and Wales and Sweden the changes of sex difference in ${}_{50}e_0^c$ and ${}_{70}e_0^c$ are relatively constant for birth cohorts from 1930 to 1960, and their level of sex gap are similar.

[Figure 4 about here]

Figure 5 shows the decomposition of the time change in the sex gap in cohort life expectancies for England and Wales and Sweden. The extended decomposition results for the rest of the countries are provided in the supplementary material. In Figure 5, the black line captures slopes of the changing sex gap in ${}_{50}e_0^c$, ${}_{70}e_0^c$ and ${}_{90}e_0^c$ across cohorts. Positive total effect indicates the sex gap in cohort truncated life expectancy widens, while the negative indicates the reverse. The patterns shown in Figure 5 confirm the fluctuations observed in Figure 4. In England and Wales (two-peak group), the positive total effect for the birth cohorts 1880 to 1895 was followed by a negative total effect for the birth cohorts 1895 to 1905, followed again by a positive total effect from 1900 to 1925 birth cohorts and a negative total effect from 1925 to the last available birth cohort. For the last truncated life expectancy to age 90, ${}_{90}e_0^c$, only the first three fluctuations (or, positive-negative-positive total effect) can be observed. For Sweden (one-peak group), the changing sex-gap pattern is relatively constant in terms of total effect for ${}_{50}e_0^c$ and ${}_{70}e_0^c$, compared with the patterns in England and Wales. The changing sex-gap pattern in ${}_{90}e_0^c$ shows a steady increasing tendency (positive total effect) from 1880 to 1895 and to the 1920s birth cohort.

Similar with period decomposition results, the *f-effect* and *e-effect* are relatively steady, while the *ρ-effect* is more varied across birth cohorts in both countries and for each truncated

life expectancy, except for the cohorts born in 1880 to 1900 in England and Wales. Without the exception happened between 1880 to 1900 birth cohorts in England and Wales, in general, when the sex gap widens, the main effect is *ρ-effect*, which accounts for more than 50 per cent of the total changes, indicating females have a faster mortality decline across cohorts than males. When the sex gap narrows, the *ρ-effect* becomes negative, but the dominant effect is *f-effect*, which contributes more than 60 per cent of the narrowing trend, suggesting that male death distribution is still more dispersed than female across cohorts. Finally, the *e-effect* is relatively steady and plays a minor widening sex gap role, indicating females keep their advantages in survivorship across cohorts.

For England and Wales birth cohort born in 1885 to 1895, the *ρ-effect* shows an accelerated decreasing trend, the *f-effect* is positive and *e-effect* is negative in ${}_{50}e_0^c$, ${}_{70}e_0^c$ and ${}_{90}e_0^c$. This extreme episode is influenced by the dramatic increase in male mortality during the first quarter of the twentieth century, leading to the widening sex gap trend. Males born in the late 1890s are less likely to experience the event, which leads male has a faster mortality decline than female across cohorts.

Surprisingly, in Sweden, the decomposition of changing sex gap in ${}_{90}e_0^c$ shows that females have a faster mortality reduction than males in cohorts from 1880 to 1920. The positive *ρ-effect* could be observed, indicating females have a faster mortality decline than males across birth cohorts.

[Figure 5 about here]

Discussion

As noted by Horiuchi et al. (2008), comparing two populations is one of the most desirous tasks among demographers – either two populations at the same time or the same population at two time points. Our extended decomposition could take both dimensions (differences in two populations at two time points) at the same time and we applied the extended method to changing sex gap in life expectancy at birth. Our extension decomposes the changing sex gap in life expectancy into the effect of the sex difference in mortality improvement (*ρ-effect*), the sex difference in the death distribution (*f-effect*), and the sex difference in remaining life expectancy (*e-effect*). These three effects are used to describe the comparison in the speed of mortality reduction, death dispersion and survival advantages, respectively.

To provide a better understanding of the changing sex-gap phenomenon in human mortality history, both period and cohort perspectives were examined. For the period analysis, four changing sex-gap patterns were identified: Classical, Delayed, Eastern Europe, and Eastern Asia. Based on the cohort analysis, we grouped the changing sex-gap patterns into one-peak and two-peak patterns. In terms of the *ρ-effect*, *f-effect* and *e-effect*, the different patterns have different contributions in period and cohort perspectives. Our results show that the *ρ-effect* is highly correlated with the changing sex gap in life expectancy, changing from positive *ρ-effect* and widening sex gap, to negative values and narrowing sex gap, while the other components are relatively constant. In both the period and cohort analyses, we found that the *f-effect* almost always produces a narrowing of the sex gap, because the male death distribution is more dispersed than the female. The *e-effect* is almost always positive, because females have lower mortality than males at all ages and hence a cumulative survivorship advantage in remaining life expectancy.

Previous research has identified three broad factors and their interaction as determinants of the sex gap in life expectancy: 1) behavioural factors, 2) political events, public health intervention and some periodical events, 3) some cultural and socio-environmental factors, and 4) biological and genetic factors (Luy, 2003; Madigan, 1957; Preston and Wang, 2006; Waldron, 1983, 1995; Wingard, 1982, 1984; Zarulli et al. 2017). In the following paragraphs, we link these explanatory factors with the results of our decomposition.

The relationship between sex differences in behavioural factors (e.g. smoking, alcohol, eating diet, risky driving, etc.) with sex-differential mortality has been emphasised by many scholars from both cohort and period perspectives (Beltrán-Sánchez et al., 2015; Pampel, 2001, 2002; Preston, Glei, and Wilmoth, 2011; Trovato and Heyen, 2006; Trovato and Lalu, 2007; Yu and Booth, 2014). Among these, the sex difference in the timing and duration of smoking has been shown to be related to the changing sex gap (Pampel, 2002, 2003). As Preston et al. (2011) found, in the 1950s the attributable risk from smoking for females was negligible in most developed countries, but it increased rapidly in Australia, Canada, Denmark, Netherlands, New Zealand, UK and USA (Group 1), though it remained low in France, Italy, Portugal and Spain (Group 2). These two groups of countries coincide respectively with the countries in the Classical and Delayed patterns of sex-gap change. In addition, Group 2 countries might have a Mediterranean Diet (Preston et al. 2011), which has been shown to be related to better health and could provide favourable mortality conditions

(especially, lower rate of cardiovascular diseases and cancer) for females (Knoops et al. 2004). The turning point in the Delayed pattern has around 15 years lag from the Classical pattern, and the three effects also capture the same lagging composition pattern. This lag might imply that the gender equality process plays a role in the changing sex gap dynamic through the diffusion of healthy/risky behaviour. Finally, in cohort results, the changing sex gap in the shorter truncated life expectancies (${}_{50}e_0^c$ and ${}_{70}e_0^c$) show an increasing and decreasing pattern in the last two cohorts' comparisons in all analysed countries (Figure 4). This could be partially explained by smoking. Similar results have been observed by Beltrán-Sánchez et al. (2015) for cohort age-specific death rates. Excess male mortality at adult ages due to smoking increased across cohorts born in 1880-1899 to 1900-1919 and then decreased across cohorts born in 1900-1919 to 1920-1935.

The effects of period events on the changing sex gap have been identified in both period and cohort results. In the Eastern Europe pattern, a zig-zag phenomenon from the late 1980s to the early 1990s was observed. The former narrowing sex gap trend is due to Gorbachev's anti-alcohol campaign (1985–1988) (Bhattacharya et al. 2013; Grigoriev et al. 2014; Shkolnikov and Nemstsov, 1997), with males receiving more benefits from this public health intervention than females, leading to a significant negative *ρ-effect*. However, the later widening sex gap is due to the social crisis caused by the dissolution of the Soviet Union (Shkolnikov et al. 1998; Vallin et al. 2018). This caused massive psychological stress, the spread of unhealthy life styles (high alcohol content consumption, and smoking) and painful social and economic reform, leading to a health crisis, especially for males (Brainerd and Cutler, 2005; Cockerham et al. 2006; Shkolnikov et al. 2001). For the cohort perspective, the two-peak pattern (including England and Wales, France, and Italy) is heavily distorted by WWI (1914-1918), the Spanish flu (1918-1920) and WWII (1939-1945), because adult males participation in the military events and males have a higher probability of death during epidemics than females (Zarulli et al. 2017). The first cohort peak in Figure 4 includes the former two of these events, resulting in a first peak higher than the second. When these wars and the epidemic occurred, positive *ρ-effect* and *f-effect* are detected for birth cohorts in 1880 to 1885. Studies have observed that young male adults had high mortality during the WWI and during the flu pandemic (Viboud et al. 2012; Gagnon et al. 2013). Later born cohorts, especially young male adults, are gradually less exposed to these harsh period events, showing the increasing negative *ρ-effect* gradient, or catching up of males, leading to the narrowing sex-gap trend. The one-peak pattern (including Denmark, Finland, Netherlands,

Norway, Switzerland and Sweden) is mainly influenced by WWII, but the magnitude varies across these countries. Like Switzerland and Sweden, whose governments kept neutral during WWII, Finland was dramatically influenced by the wars with the Soviet Union (Andreeva et al. 2017).

The cultural impacts on the changing sex gap are more substantial in the Eastern Asia pattern, where the sex gap in life expectancy is increasing during the period. The reasons for this increasing sex-gap pattern are various. Besides the sex difference in smoking and alcohol consumption (Liu et al. 2013; Trovato and Heyen, 2006), the traditional cultures, social norms and gender roles might be the underlying mechanism. In Japan, for example, society has a long history of division of labour – female's place is at home, while male's place is at work (Sugihara and Katsurada, 2000) which could help females avoid a hazardous environment and working stress. This might be one reason for Japanese females always having a faster mortality reduction than males from 1950 to 2000, leading to the positive *ρ-effect*.

Finally, our results also could be linked with the biological and genetic. As shown in the period results, the total effect in the most advanced countries in terms of the sex gap transition (like Australia, Canada, Demark, England and Wales, Netherlands, Sweden, and USA) has a declining tendency, especially seen in the *ρ-effect*. In the Classical pattern, a convergence in the sex gap across these countries is observed. These might suggest the tendency of changing sex gap is approaching a ceiling and the recent years' male advantages in mortality improvement is disappearing in most advanced countries. Given that the three effects are reducing over time (or the total effect is close zero), there might be a biological difference between females and males under the current epidemiological environment. Similar phenomenon also could be observed in the cohort analysis. The sex gap in ${}_{50}e_0^c$ and ${}_{70}e_0^c$ changes in a slow pace for cohorts born after the 1930s, which might suggest that the room for further decline of sex difference in premature death is almost zero.

The assessments of these explanatory factors of the sex gap are out of the scope of the present study. Our contribution can be included among the research focusing on the first stage of demographic research, the discovery stage (Billari, 2015). Work on the second stage of demographic research, namely explanatory research, will benefit from using the methodology here presented. Given the heterogeneity of cultural, institutional, and socio-economic contexts across the developed countries included in our analysis, further research is

required to investigate the different explanatory mechanisms that trigger the changing sex gap.

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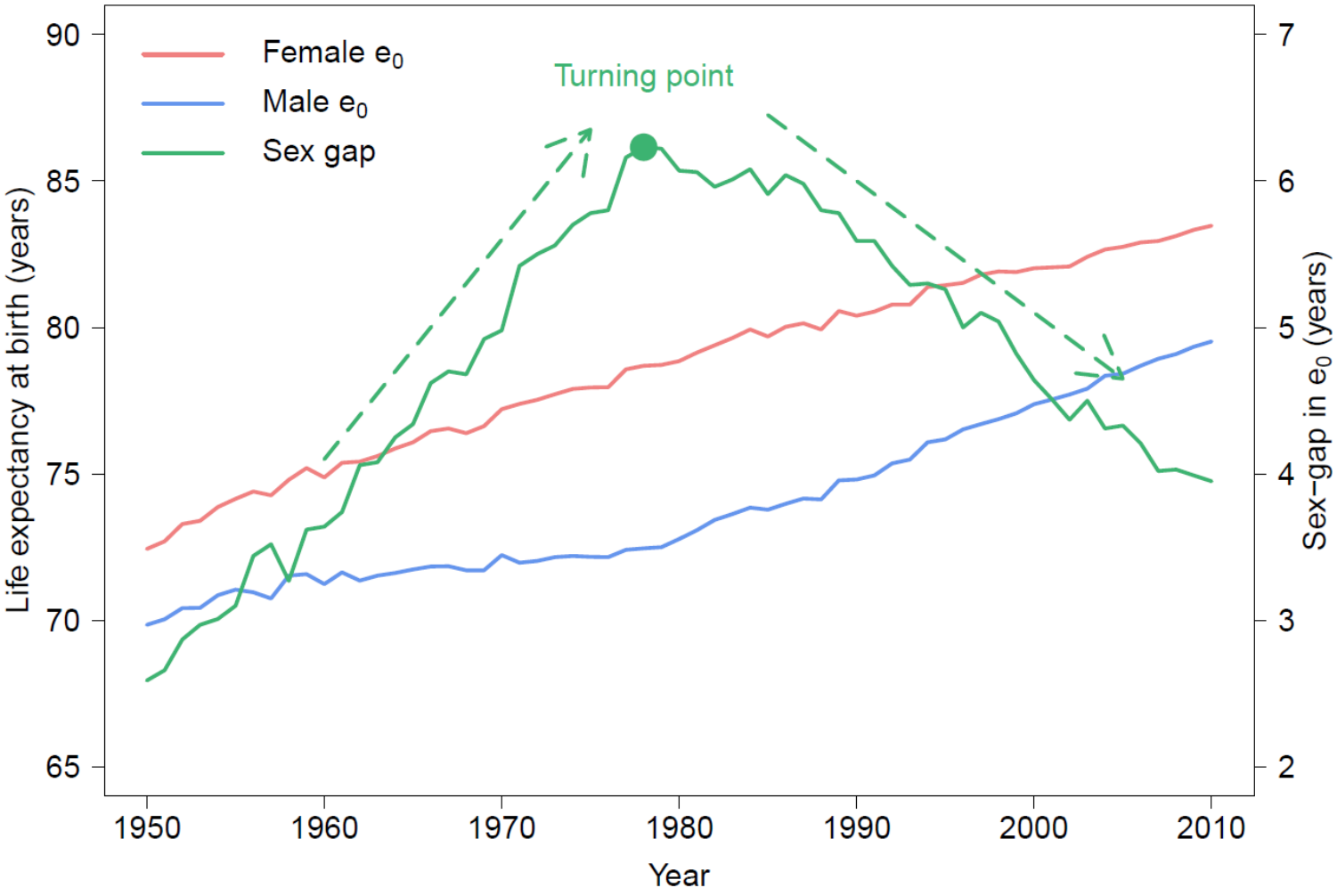
Figures & Tables

Table 1 Countries, periods and birth cohorts of data used in the analysis

Country	Period	Cohort	Country	Period	Cohort
Australia	1950-2010	1925-1960	Netherlands	1950-2010	1880-1960
Austria	1950-2010	1950-1960	Norway	1950-2010	1880-1960
Belgium	1950-2010	-	New Zealand	1950-2010	1950-1960
Bulgaria	1950-2010	1950-1960	Portugal	1950-2010	1940-1960
Canada	1950-2010	1925-1960	Slovakia	1950-2010	1950-1960
Switzerland	1950-2010	1880-1960	Sweden	1950-2010	1880-1960
Czech	1950-2010	1950-1960	USA	1950-2010	1935-1960
West Germany	1950-2010	-	Belarus	1960-2010	-
Denmark	1950-2010	1880-1960	East Germany	1960-2010	-
Spain	1950-2010	1910-1960	Estonia	1960-2010	-
Finland	1950-2010	1880-1960	Latvia	1960-2010	-
France	1950-2010	1880-1960	Lithuania	1960-2010	-
England & Wales	1950-2010	1880-1960	Poland	1960-2010	-
Hungary	1950-2010	1950-1960	Russia	1960-2010	-
Ireland	1950-2010	-	Ukraine	1960-2010	-
Italy	1950-2010	1880-1960	Taiwan	1970-2010	-
Japan	1950-2010	1950-1960			

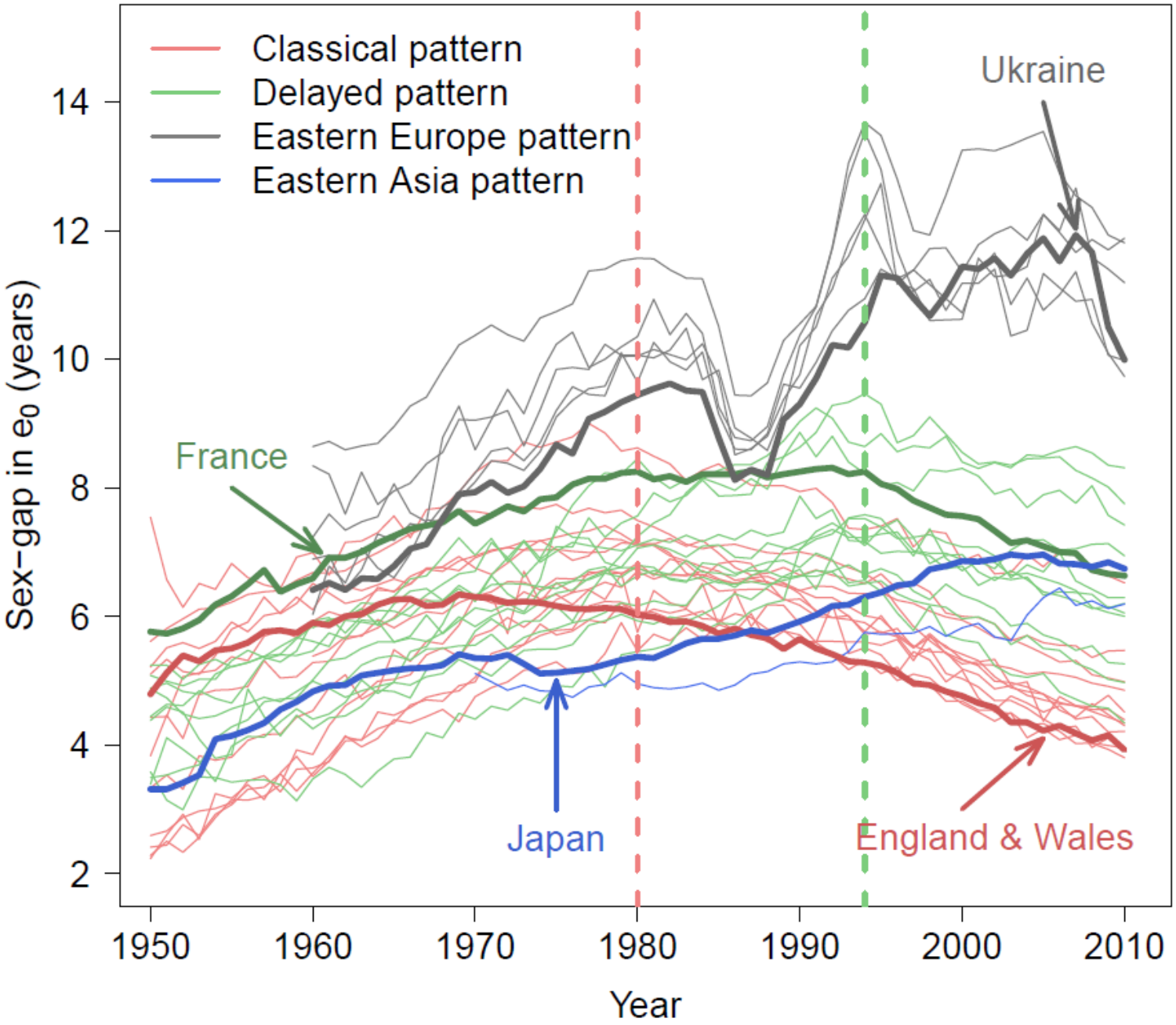
Source: HMD (2018)

Figure 1 Sex gap in life expectancy at birth, Sweden, 1950-2010.



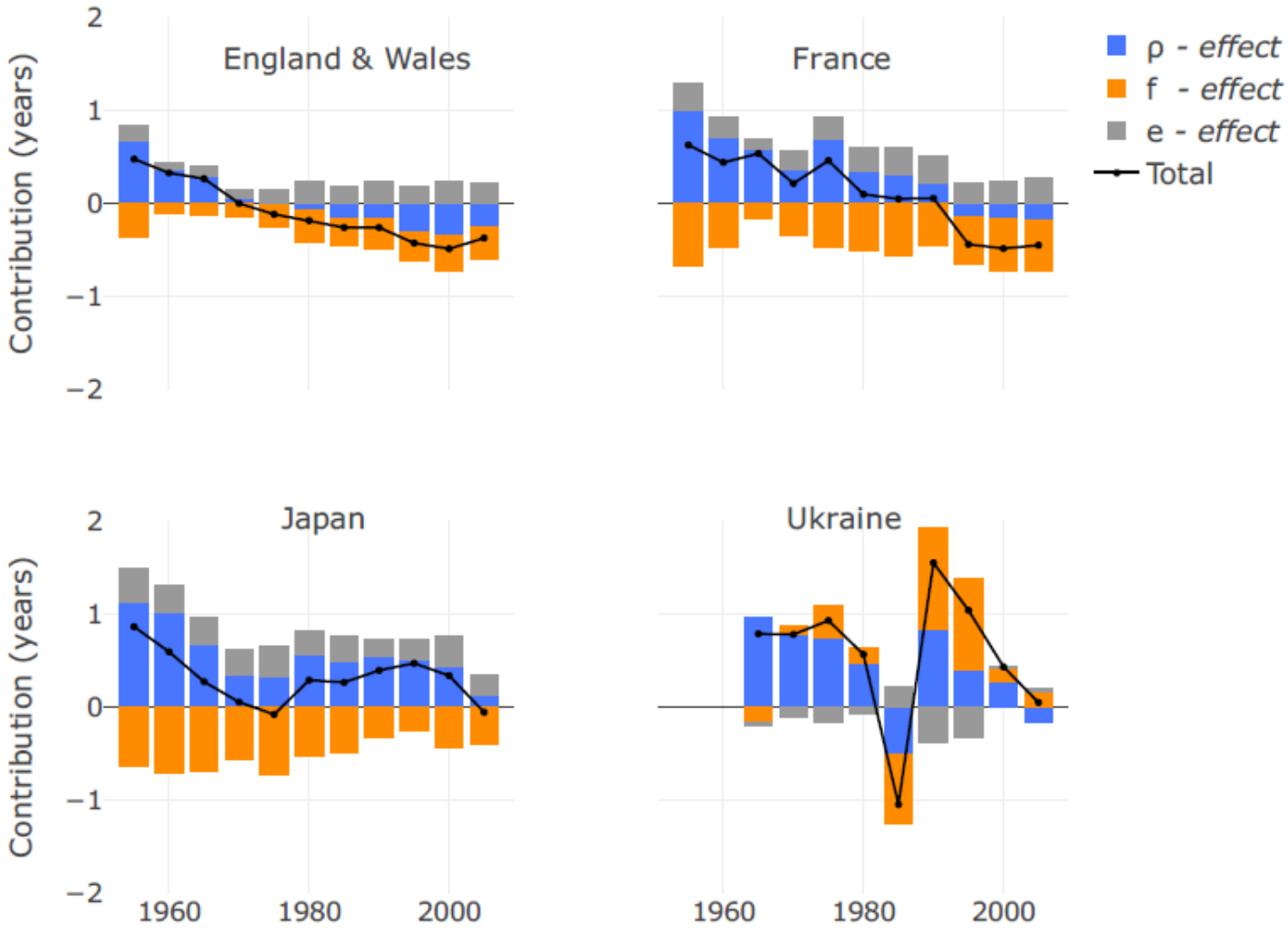
Source: Authors' calculations based on the single year life table (HMD, 2008)

Figure 2 Sex gap in life expectancy at birth by pattern, 1950-1954 to 2005-2009



Source: Authors' calculations based on the single-year life table (HMD, 2008)

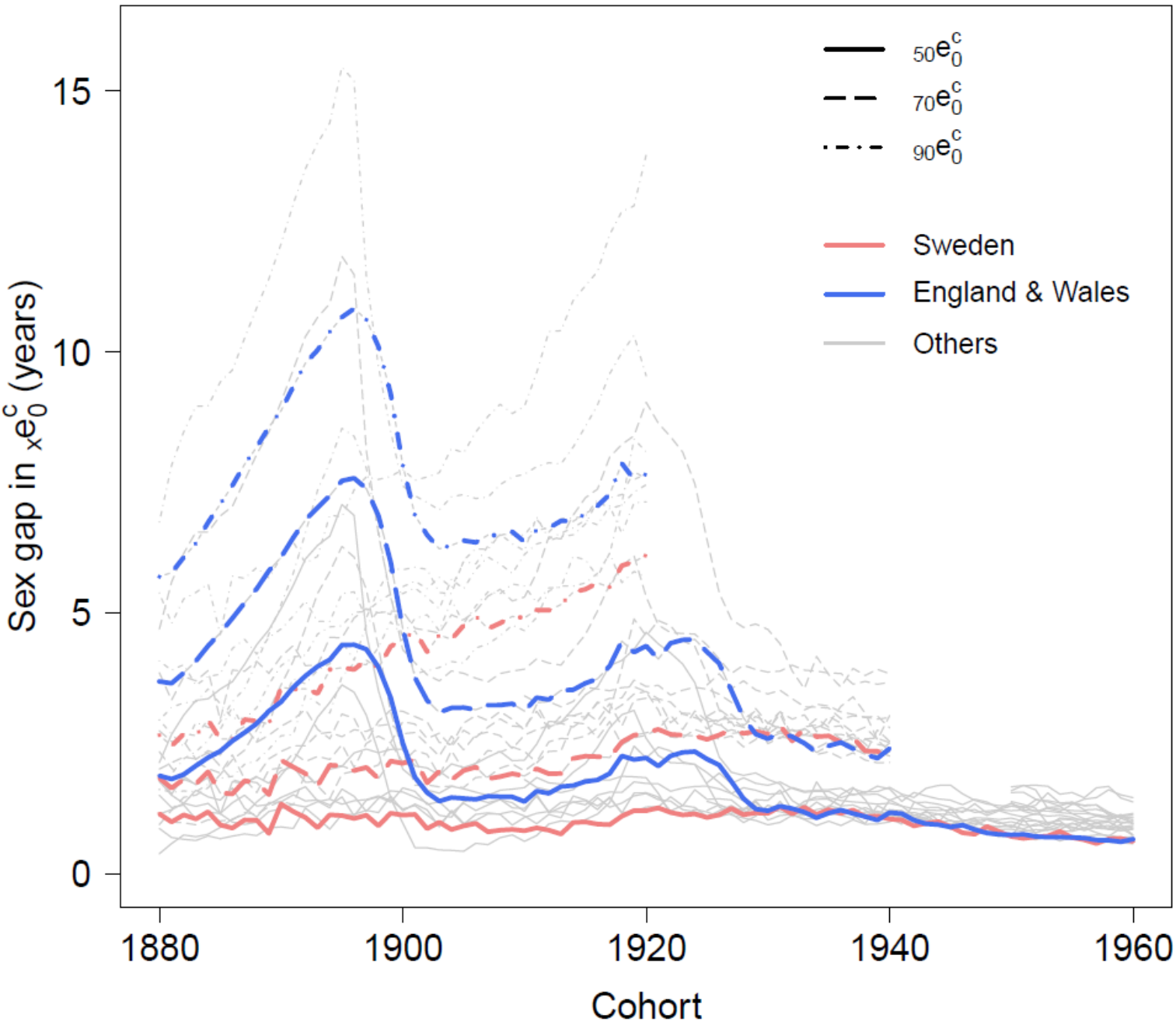
Figure 3 Decomposition of change in the sex-gap in life expectancy at birth, England and Wales, France, Japan, and Ukraine, 1950-2010



Source: Authors' calculations based on the 5-year life table (HMD, 2008)

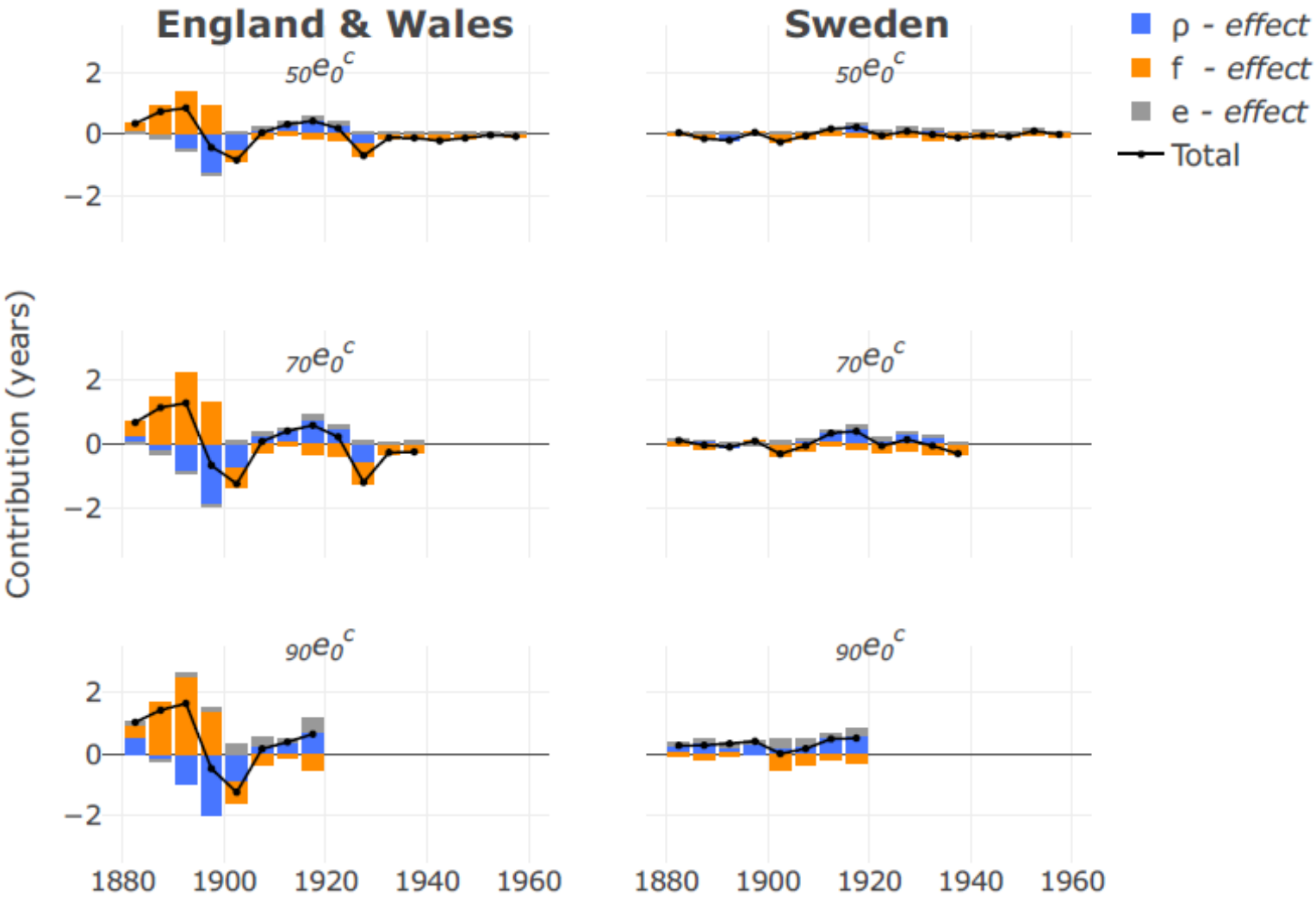
Note: 1965 means from 1960-1964 to 1965-1969.

Figure 4 Sex gap in cohort truncated life expectancies, Sweden, England & Wales, and others, cohorts born in 1880 to 1960



Source: Authors' calculations based on single year life tables (HMD, 2008)

Figure 5 Decomposition of change in the sex-gap in cohort truncated life expectancies at ages 50, 70 and 90 for England and Wales and Sweden



Source: Authors' calculations based on period single year life table data (HMD, 2008)