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*Descriptive Findings*

# Divergence without decoupling: Male and female life expectancy usually co-move

**Andrew Noymer**

**Viola Van**

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*Demographic Research*: Volume 31, Article 51
*Descriptive Findings*

# Divergence without decoupling: Male and female life expectancy usually co-move

**Andrew Noymer** [1]

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## Abstract

**BACKGROUND**
Divergence of male and female life expectancy is a well-documented phenomenon. Co-movement is a heretofore-neglected aspect of changes in male and female mortality.

**OBJECTIVE**
We develop a new framework for life expectancy sex differentials in time series, using co-movement/anti-movement and convergence/divergence.

**METHODS**
We apply this framework to the Human Mortality Database (HMD), assessing co-movement between male and female life expectancy with the nonparametric test of Goodman and Grunfeld (1961).

**RESULTS**
For every country in the HMD (except three with short spans of data), male and female mortality statistically co-move. This applies even in cases, including ones such as Russia that are well-discussed in the literature, that show extreme divergence between the sexes. The results are reasonably robust to subsetting with a 25-year time-window for all countries.

**CONCLUSIONS**
Male and female life expectancy co-move even when the life expectancy sex differential increases. The sex divergence in life expectancy needs to be (re-)considered in light of the fact that male and female life expectancy usually co-move, reflecting overall societal factors.

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# 1. Introduction

Women almost always have longer life expectancy than men. This fact is well-ingrained in mortality research, even if the root causes are still debated. While the contribution of behavior (e.g. alcohol or tobacco use) to life expectancy sex differentials is beyond doubt, the counterfactual condition of what mortality differences would look like in the absence of ostensibly changeable behaviors is less clear. A biological basis for some portion of observed mortality sex differentials is plausible (Waldron 1983; Garenne and Lafon 1998), although in many settings there is also a social component (see, e.g., Das Gupta 1987; Voland et al. 1997). Even for infant mortality, the role of biology in sex differences is debated (Pongou 2013). The long-held interest in mortality sex differentials has generated a large body of literature (e.g., Ciocco 1940a,b; Enterline 1961; Retherford 1975; Preston 1976, 1977; Lopez and Ruzicka 1983; Nathanson 1984; Ram 1993; Vallin 1995; Trovato and Lalu 1996; Kalben 2002; Pampel 2002; Meslé 2004; Preston and Wang 2006; Glei and Horiuchi 2007; Luy 2009; Rogers et al. 2010; Oksuzyan et al. 2010; Medalia and Chang 2011; Clark and Peck 2012; Kageyama 2012; Sawyer 2012; Seifarth, McGowan, and Milne 2012; Lindahl-Jacobsen et al. 2013; Thorslund et al. 2013; Oyen et al. 2013), a complete review of which is beyond the current scope.

Our question is whether the forces that shape time series of life expectancy do so in a similar way for males and for females. We do not attempt to explain the existence of the mortality sex differential per se, nor to decompose it among behavioral or biological components. Our interest is in changes of male and female life expectancy, paired at the country level, on a year-to-year basis. This analysis may be indirectly applicable to questions of behavior or biology, but is not designed to test crisp hypotheses about which matters more. Rather, we analyze divergence of male and female life expectancy trajectories using a new analytic framework, that of co-movement, described below.

The divergence of male and female life expectancy has been a topic of much interest, especially in the last 20 years or so (Vallin 1990, 1995; Shkolnikov, Meslé, and Vallin 1995; Meslé and Vallin 1998; Jasilionis et al. 2011). Adding another dimension, co-movement, uses the same data to reveal different aspects of human mortality, and how these reflect the conditions of the living. This paper examines male:female co-movement of life expectancy in 40 countries around the world. We find that – even when they diverge – males and females almost always statistically co-move.

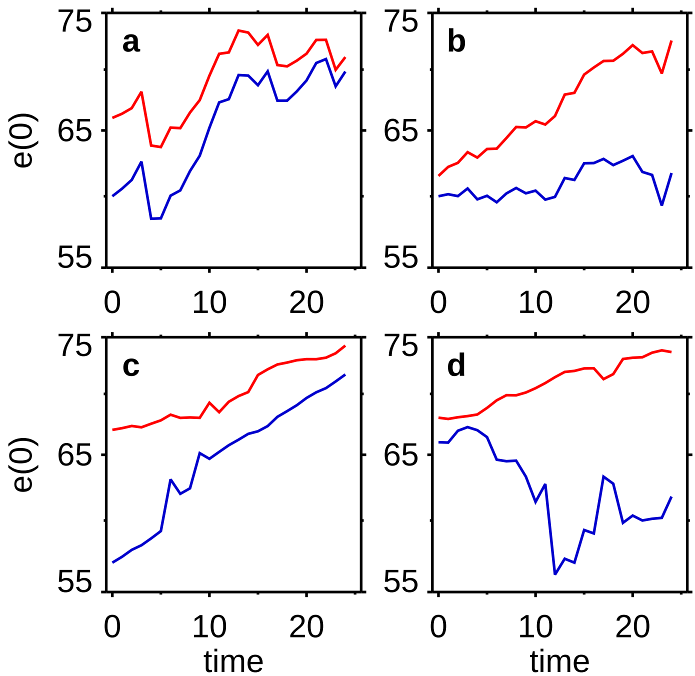
To set the stage further for the analysis that follows, consider two scenarios that could account for the divergence of male and female life expectancy. One case is that certain sex-specific factors are superposed on overall secular trends of mortality. Examples of these factors include but are not limited to behaviors such as tobacco and alcohol use. This superposition leads to widening sex differentials of mortality because it represents a consistent drag on one sex more than another without altering the common secular trend, thus implying co-movement. Another case would be that divergence occurs because male and female life expectancy have become decoupled (or never were coupled), moving their own independent ways with separate drift and noise parameters. This implies lack of co-movement: once serial correlation of each series has been dealt with, two independent random walks are not expected to co-move except by chance. Indeed, given the sometimes-enormous sex differences in life expectancy, what is to say – in theory at least – that life expectancy time series need be coupled in the first place? This paper brings an analytic framework and empirical data to bear on these contrasting scenarios.

## 2. Co-movement and convergence

We use the terms "co-movement" and "divergence/convergence" in a specific way. Hereinafter, if two time series are co-moving, then when one goes up, the other one does too, and if one declines, they both decline. Thus, co-movement between two time series means common signs of first differences that are statistically distinguishable from random movement. Statistical co-movement means that the preponderance of moves are in the same direction, not that the two time series need be in lock-step. Only the direction of the movements matters; the magnitudes need not be correlated. The opposite of co-movement is anti-movement. In other words, male and female life expectancies are usually both trending upward, so they have common movements just because of this; correcting for serial correlation takes this into account. Convergence has a more intuitive definition: in a specified time window, if two time series are closer together at the end than at the beginning, they are converging. The opposite of convergence is divergence.

Figure 1 illustrates a $2 \times 2$ lay-out of co-movement and convergence in male and female life expectancy for four simulated populations. The top row shows co-movement, while the bottom row shows no co-movement. The left column shows convergence while the right column shows divergence. There is no reason that male and female life expectancy in a given country cannot be like any of the quadrants of Figure 1. Possibilities not shown include parallel trajectories (neither convergence nor divergence), and anti-movement (which, *a priori*, seems unlikely in real life expectancy data). Figure 1 is a schematic of a novel way of conceptualizing divergence/convergence of demographic timeseries, because it introduces a second dimension, that of co-movement or lack thereof.

**Figure 1: Examples of convergence/divergence (left column is convergence; right column is divergence) and co-movement (top row, co-movement; bottom row, no co-movement). Simulated life expectancy trajectories for men (blue) and women (red). (a) co-movement with convergence; (b) co-movement with divergence; (c) no co-movement, with convergence; (d) no co-movement, with divergence**




## 3. Materials and methods

We use all available period life expectancy data from the Human Mortality Database. For France, the series has been supplemented at the beginning of the 19th century by 10 data points from Vallin and Meslé (2001). Table 1 provides basic descriptive statistics for the 40 countries (alphabetically by HMD country abbreviations). Minimum life expectancy usually occurs at or close to the start of the data set, and maximal values occur most recently, almost always at or near the end of the series. A prominent counterexample are Russian males, whose life expectancy peaked in 1964, as is well documented (Shkolnikov, Meslé, and Vallin 1995; Cockerham 1999; Gavrilova et al. 2008; Billingsley 2011).

**Table 1: Descriptive statistics**

| | | | Males | | | | Females | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Start | End | minimum | | maximum | | minimum | | maximum | |
| Country | date | date | Value | Year | Value | Year | Value | Year | Value | Year |
| Australia | 1921 | 2009 | 59.12 | 1921 | 79.70 | 2009 | 63.24 | 1921 | 84.16 | 2009 |
| Austria | 1947 | 2010 | 58.69 | 1947 | 77.70 | 2010 | 63.80 | 1947 | 83.15 | 2010 |
| Belarus | 1959 | 2012 | 62.18 | 2002 | 69.35 | 1964 | 71.98 | 1959 | 77.67 | 2012 |
| Belgium | 1841 | 2012 | 31.66 | 1866 | 77.77 | 2011 | 33.27 | 1866 | 82.86 | 2011 |
| Bulgaria | 1947 | 2010 | 52.54 | 1947 | 70.31 | 2010 | 55.70 | 1947 | 77.26 | 2009 |
| Canada | 1921 | 2009 | 55.85 | 1923 | 79.00 | 2009 | 58.05 | 1923 | 83.39 | 2009 |
| Chile | 1992 | 2008 | 71.42 | 1993 | 75.33 | 2008 | 77.47 | 1992 | 81.13 | 2008 |
| Czech Republic | 1950 | 2011 | 61.97 | 1950 | 74.71 | 2011 | 66.85 | 1950 | 80.86 | 2011 |
| Denmark | 1835 | 2011 | 36.65 | 1835 | 77.70 | 2011 | 39.83 | 1853 | 81.83 | 2011 |
| Estonia | 1959 | 2011 | 60.64 | 1994 | 71.09 | 2011 | 71.93 | 1959 | 80.99 | 2011 |
| Finland | 1878 | 2009 | 26.32 | 1918 | 76.51 | 2009 | 38.94 | 1881 | 83.14 | 2009 |
| France | 1806 | 2012 | 23.77 | 1813 | 78.51 | 2012 | 32.42 | 1871 | 85.02 | 2011 |
| GDR | 1956 | 2011 | 65.84 | 1957 | 77.05 | 2011 | 70.52 | 1957 | 82.85 | 2011 |
| FRG | 1956 | 2011 | 65.82 | 1957 | 78.25 | 2011 | 70.89 | 1956 | 82.94 | 2011 |
| Hungary | 1950 | 2009 | 59.85 | 1950 | 70.21 | 2009 | 64.25 | 1950 | 78.23 | 2009 |
| Ireland | 1950 | 2009 | 63.51 | 1951 | 77.41 | 2008 | 66.11 | 1951 | 82.23 | 2009 |
| Iceland | 1838 | 2010 | 16.76 | 1882 | 79.78 | 2008 | 18.82 | 1846 | 83.84 | 2010 |
| Israel | 1983 | 2009 | 73.10 | 1983 | 79.63 | 2009 | 76.53 | 1983 | 83.32 | 2009 |
| Italy | 1872 | 2009 | 23.50 | 1918 | 79.22 | 2009 | 28.33 | 1918 | 84.24 | 2009 |
| Japan | 1947 | 2009 | 49.78 | 1947 | 79.61 | 2009 | 53.65 | 1947 | 86.42 | 2009 |
| Latvia | 1959 | 2011 | 58.71 | 1994 | 68.53 | 2011 | 72.23 | 1994 | 78.52 | 2011 |
| Lithuania | 1959 | 2011 | 62.52 | 1994 | 68.30 | 1964 | 70.55 | 1959 | 79.06 | 2011 |
| Luxemburg | 1960 | 2009 | 65.42 | 1964 | 78.02 | 2008 | 71.82 | 1962 | 82.90 | 2009 |
| Netherlands | 1850 | 2009 | 29.88 | 1859 | 78.53 | 2009 | 31.86 | 1859 | 82.64 | 2009 |
| New Zealand | 1948 | 2008 | 66.97 | 1948 | 78.37 | 2008 | 70.74 | 1949 | 82.34 | 2008 |
| Norway | 1846 | 2009 | 43.34 | 1848 | 78.62 | 2009 | 45.78 | 1862 | 83.08 | 2009 |
| Poland | 1958 | 2009 | 62.62 | 1959 | 71.48 | 2009 | 68.34 | 1959 | 79.92 | 2009 |
| Portugal | 1940 | 2012 | 45.86 | 1941 | 77.25 | 2012 | 49.89 | 1941 | 83.50 | 2011 |
| Russia | 1959 | 2010 | 57.38 | 1994 | 64.89 | 1964 | 71.07 | 1994 | 74.79 | 2010 |
| Slovakia | 1950 | 2009 | 59.14 | 1950 | 71.36 | 2009 | 62.55 | 1950 | 78.95 | 2009 |
| Slovenia | 1983 | 2009 | 66.77 | 1983 | 75.77 | 2009 | 74.88 | 1983 | 82.30 | 2009 |
| Spain | 1908 | 2009 | 29.82 | 1918 | 78.48 | 2009 | 30.69 | 1918 | 84.55 | 2009 |
| Sweden | 1751 | 2011 | 17.15 | 1773 | 79.79 | 2011 | 18.79 | 1773 | 83.67 | 2011 |
| Switzerland | 1876 | 2011 | 38.38 | 1876 | 80.28 | 2011 | 41.52 | 1877 | 84.68 | 2011 |
| Taiwan | 1970 | 2010 | 66.32 | 1970 | 76.24 | 2010 | 71.42 | 1970 | 82.37 | 2010 |
| England and Wales | 1841 | 2011 | 33.38 | 1918 | 79.05 | 2011 | 38.14 | 1849 | 82.94 | 2011 |
| Scotland | 1855 | 2011 | 38.77 | 1864 | 76.40 | 2011 | 41.48 | 1864 | 80.85 | 2011 |
| Northern Ireland | 1922 | 2011 | 53.70 | 1924 | 77.82 | 2011 | 54.75 | 1925 | 82.39 | 2011 |
| Ukraine | 1959 | 2009 | 61.21 | 1995 | 68.45 | 1964 | 72.20 | 1959 | 75.19 | 1989 |
| United States | 1933 | 2010 | 58.34 | 1934 | 76.37 | 2010 | 62.34 | 1934 | 81.21 | 2010 |

**Table 2: Descriptive statistics, female−male difference**

| | Female−Male Difference | | | | maximum 25-year divergence | |
|---|---|---|---|---|---|---|
| | Minimum | | Maximum | | | |
| | value | date | value | date | magnitude | start date |
| Australia | 3.39 | 1928 | 7.17 | 1980 | 3.25 | 1945 |
| Austria | 4.77 | 1949 | 7.21 | 1978 | 2.26 | 1949 |
| Belarus | 6.42 | 1962 | 12.24 | 2005 | 3.47 | 1987 |
| Belgium | −0.47 | 1845 | 8.96 | 1940 | 3.42 | 1920 |
| Bulgaria | 2.84 | 1948 | 7.58 | 1994 | 3.18 | 1970 |
| Canada | 2.14 | 1924 | 7.41 | 1978 | 3.75 | 1936 |
| Chile | 5.77 | 2006 | 6.71 | 1996 | 3.75 | 1992 |
| Czech Republic | 4.79 | 1953 | 7.87 | 1990 | 2.32 | 1952 |
| Denmark | 1.32 | 1869 | 6.21 | 1979 | 3.46 | 1950 |
| Estonia | 7.57 | 1959 | 12.92 | 1995 | 3.82 | 1971 |
| Finland | 1.87 | 1878 | 23.67 | 1941 | 20.00 | 1917 |
| France | 0.69 | 1858 | 25.31 | 1915 | 22.49 | 1891 |
| GDR | 4.47 | 1958 | 7.50 | 1994 | 2.30 | 1970 |
| FRG | 4.69 | 2011 | 6.75 | 1980 | 1.90 | 1987 |
| Hungary | 3.80 | 1954 | 9.46 | 1994 | 3.97 | 1968 |
| Ireland | 2.24 | 1950 | 5.90 | 1979 | 2.96 | 1950 |
| Iceland | −2.70 | 1891 | 10.70 | 1906 | 12.49 | 1867 |
| Israel | 3.43 | 1983 | 4.40 | 1998 | 0.40 | 1983 |
| Italy | −0.25 | 1885 | 16.46 | 1917 | 16.19 | 1893 |
| Japan | 3.31 | 1951 | 6.95 | 2003 | 1.81 | 1951 |
| Latvia | 7.10 | 1959 | 13.52 | 1994 | 4.93 | 1970 |
| Lithuania | 6.04 | 1960 | 12.68 | 2007 | 4.19 | 1960 |
| Luxemburg | 4.71 | 2008 | 7.85 | 1977 | 2.44 | 1977 |
| Netherlands | 1.28 | 1933 | 10.42 | 1945 | 8.77 | 1921 |
| New Zealand | 3.47 | 1949 | 6.70 | 1973 | 3.23 | 1949 |
| Norway | 2.19 | 1862 | 6.86 | 1986 | 3.26 | 1949 |
| Poland | 5.60 | 1958 | 9.22 | 1991 | 2.89 | 1967 |
| Portugal | 4.03 | 1941 | 7.50 | 1996 | 2.24 | 1942 |
| Russia | 8.30 | 1959 | 13.69 | 1994 | 3.32 | 1970 |
| Slovakia | 3.41 | 1950 | 8.87 | 1990 | 3.79 | 1954 |
| Slovenia | 6.53 | 2009 | 8.25 | 1985 | 1.72 | 1985 |
| Spain | 0.87 | 1918 | 9.35 | 1937 | 7.42 | 1913 |
| Sweden | 1.24 | 1779 | 6.65 | 1789 | 3.22 | 1954 |
| Switzerland | 2.06 | 1883 | 6.99 | 1991 | 2.75 | 1894 |
| Taiwan | 4.83 | 1985 | 6.41 | 2006 | 1.56 | 1982 |
| England and Wales | 1.64 | 1849 | 20.02 | 1917 | 16.80 | 1893 |
| Scotland | 2.01 | 1871 | 7.99 | 1944 | 4.81 | 1918 |
| Northern Ireland | 0.21 | 1927 | 7.14 | 1974 | 4.31 | 1937 |
| Ukraine | 6.41 | 1962 | 11.92 | 2007 | 3.35 | 1972 |
| United States | 3.61 | 1933 | 7.73 | 1975 | 2.77 | 1933 |

Table 2 provides descriptive statistics for life expectancy sex differentials. There is much international variation in the timing of maximum and minimum sex differences in life expectancy. The right two columns of Table 2 give the 25-year window for which the divergence between male and female life expectancy is greatest. For example, for Australia, the period 1945–69 has the greatest 25-year divergence between the sexes in life expectancy: in 1945 the sex difference was 3.68 years (females, 70.35; males, 66.67) and 25 years later the difference was 6.93 (females, 74.70; males, 67.77), for a divergence of 3.25 years. Note, in this example, the end of the 25-year window does not correspond to the single-year maximal sex differential (which was 7.17 years, in 1980), nor does the start correspond to the minimal sex differential (3.39, in 1928). However, there is no 25-year window in the Australian data in which the end-minus-start life expectancy sex differential exceeds 3.25 years.

The countries in the HMD differ in the length of their series. Therefore, we perform a separate analysis, using only the data from the 25-year window for each country. Looking at 25-year subsets is one way to put all the countries on an equal footing, ensuring that differences in statistical significance are not a function of differences in sample size. We choose the 25-year subset in which there is the most divergence to give a conservative estimate of co-movement, since periods of divergence theoretically ought to be when there is the weakest mortality coupling between the sexes. Chile (1992–2008, 17 years) is the only country in the HMD with fewer than 25 years of life expectancy data.

Using the Goodman-Grunfeld (1961) nonparametric test of co-movement between two time series, we analyze signs of first-differences. This approach examines if male and female life expectancy both increase, or both decrease, or move in either permutation of opposite directions, on a year-to-year basis. The Goodman-Grunfeld test is based on the $\chi^2$ statistic, but introduces a correction for serial correlation; important, since life expectancy is not a stationary process. The test is also well-suited to small sample sizes (Goodman and Grunfeld 1961).

Figure 2 presents all the data underlying this study. For each country, two panels are shown. The bottom panel is a time series plot of male and female life expectancy for all the available years. The 25-year window of maximum divergence of male and female life expectancy is indicated by gray shading. The top panel for each country shows the corresponding female minus male life expectancy differential, which is always positive, except for Iceland in 1891; Italy in 1883, 1885–87, 1889; and Belgium in 1842–45, 1848, 1850–52.

**Figure 2: Male and female life expectancy for 6 countries, as labeled. Within each country, the top panel shows the female minus male life expectancy sex differential, and the bottom panel shows male and female life expectancy, with the span of the 25-year window of maxiumum divergence indicated by gray shading**

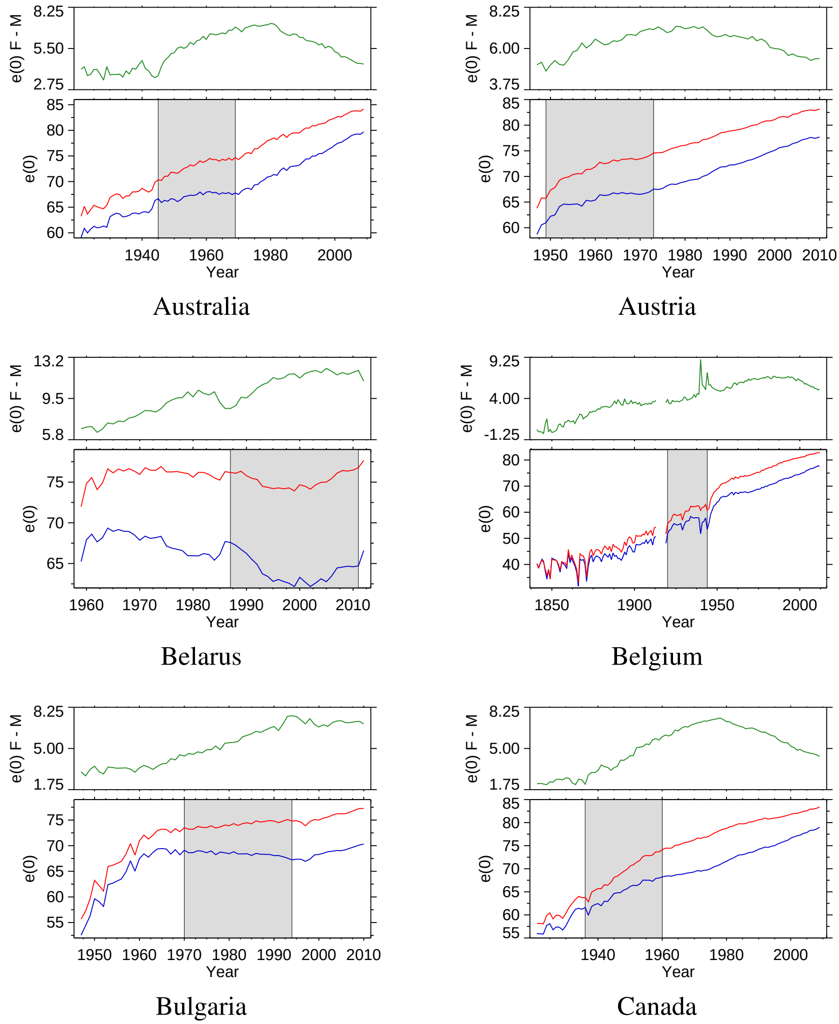
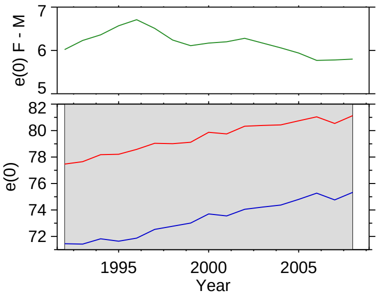
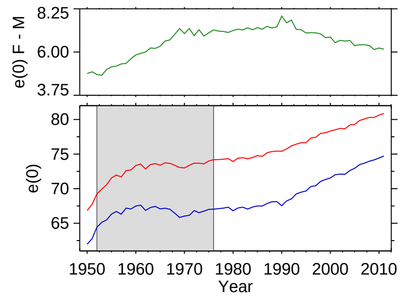


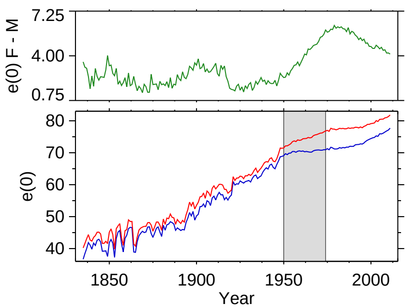
**Figure 2: (Continued)**



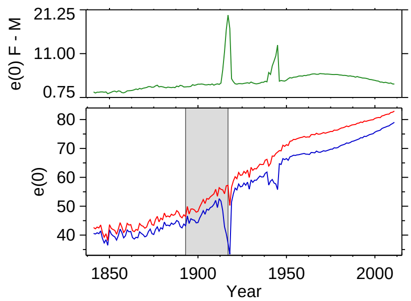

Chile



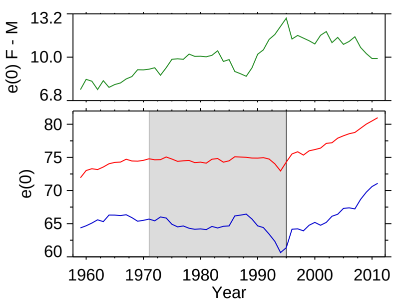

Czech Republic



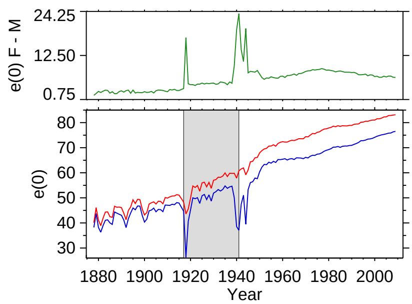

Denmark




England & Wales

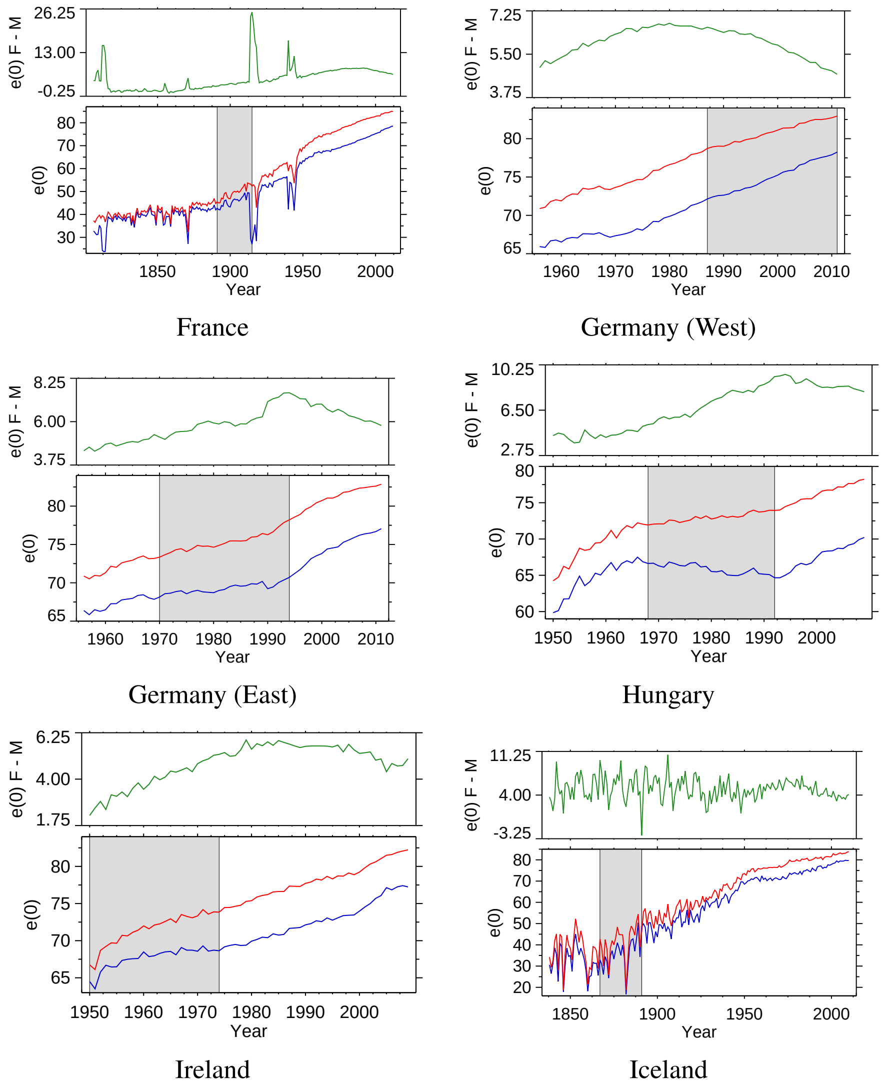



Estonia




Finland

**Figure 2: (Continued)**



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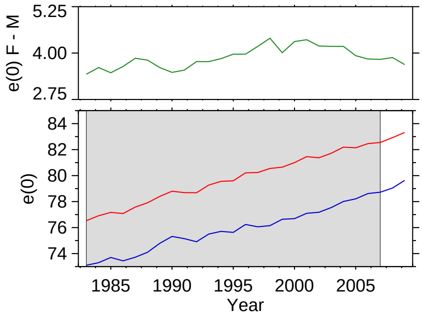
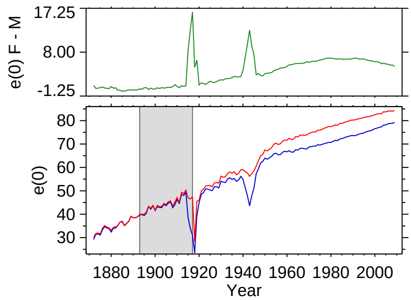
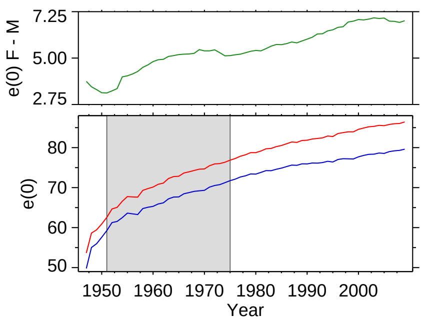


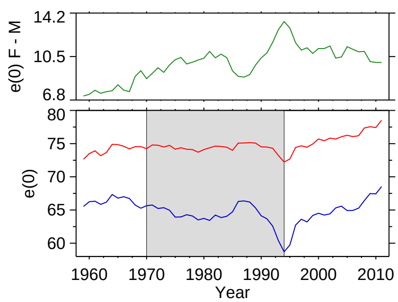

Israel



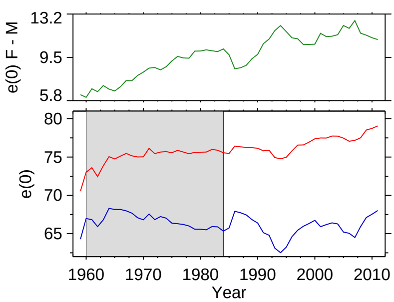

Italy



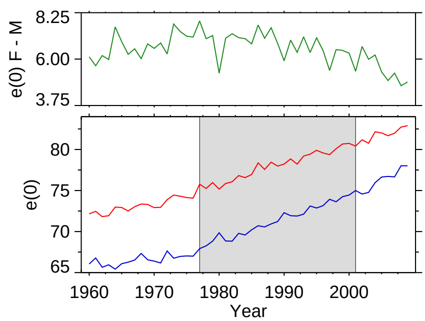

Japan




Latvia

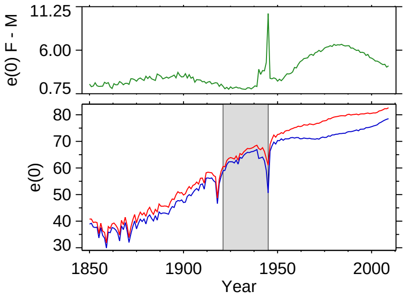



Lithuania

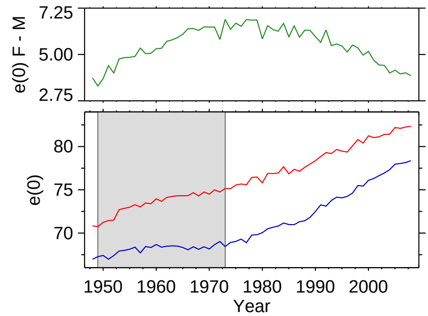



Luxembourg

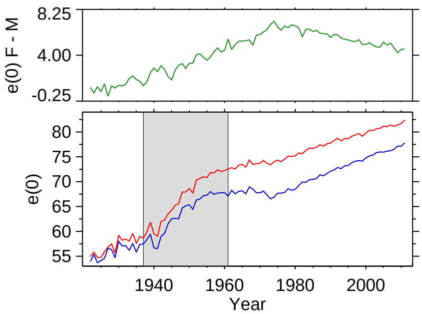
**Figure 2: (Continued)**



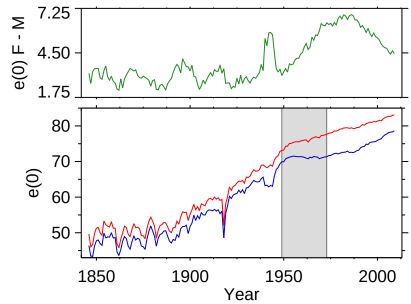

Netherlands



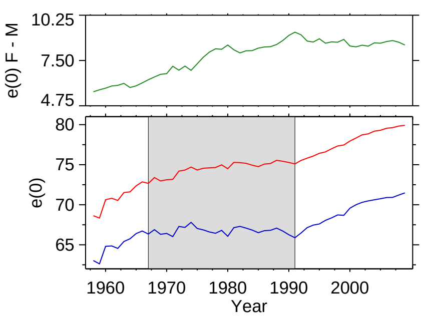

New Zealand



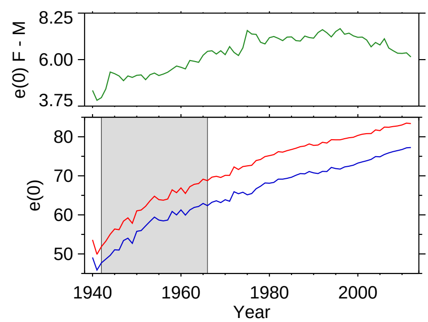

Northern Ireland




Norway




Poland




Portugal

**Figure 2: (Continued)**

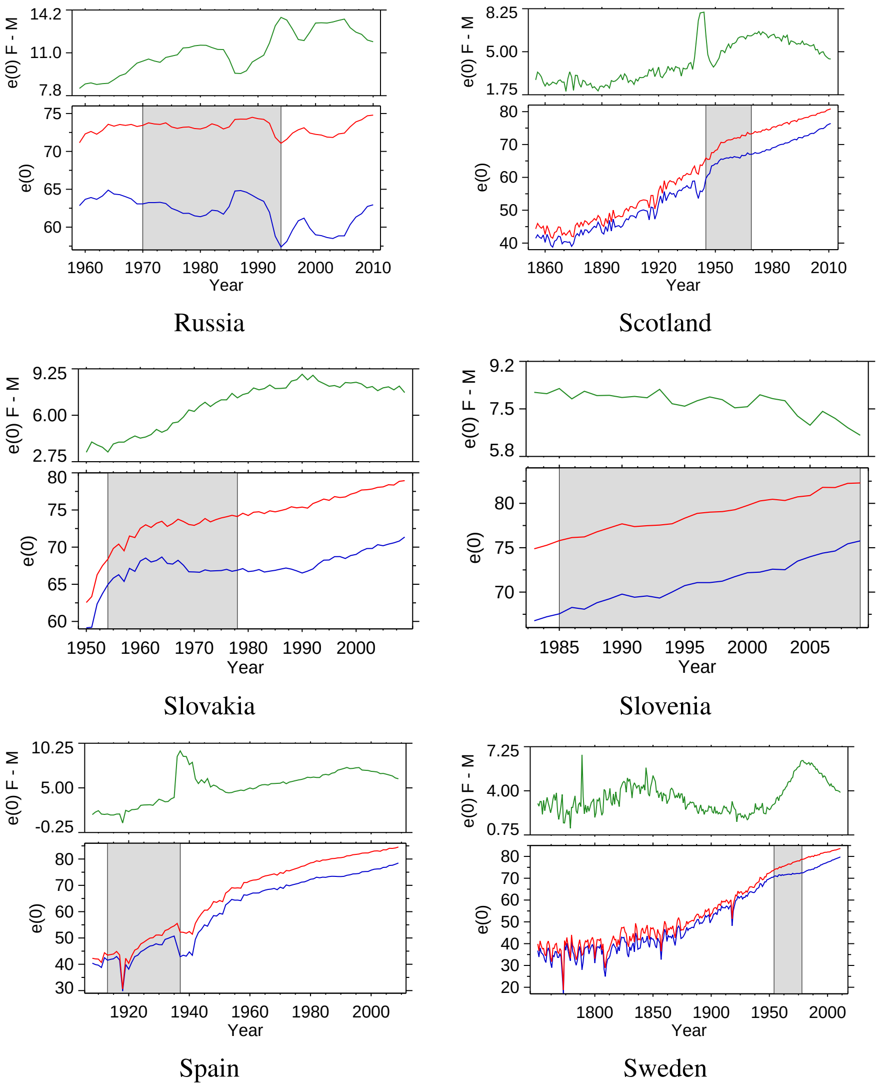
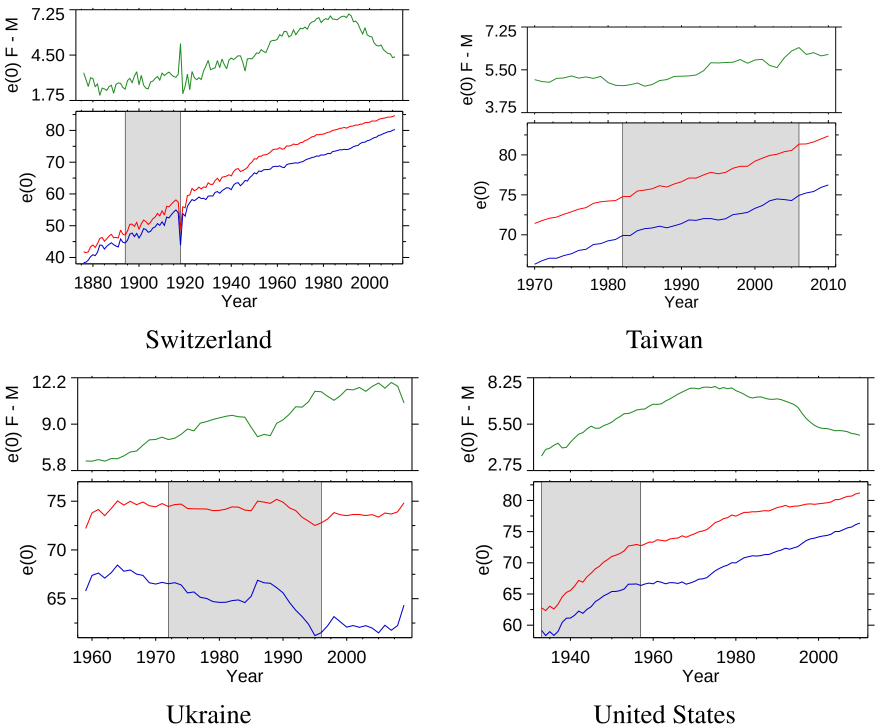


**Figure 2: (Continued)**



## 4. Results and discussion

Table 3 presents the results of the Goodman-Grunfeld tests. When using the entire time series for each country, only three countries do not have significant co-movement of life expectancy between the sexes. These are Chile, which has only 17 years of data, Luxembourg, which has ample data, and Slovenia (only 27 years of data). The negative test statistic for Luxembourg (Table 3) is nowhere near large enough to indicate statistically-significant anti-movement if the opposite one-sided $p$-value were to be calculated; it is simply non-significant. All the other G-G test statistics reject the null hypothesis of no co-movement at the 5% level. Clearly, the overall picture is that male and female life expectancies co-move.

Table 3 also gives the proportion of the co-movements which are congruent. This may be thought of like an effect size, as opposed to a test of statistical significance. For most of the populations, the proportion of congruent co-movements is high – for example of the 37 statistically-significant countries, 22 have an 85% or greater proportion of congruent co-movements. On the other hand, as the most extreme example of a modest effect size, consider New Zealand. With a test statistic of $2.439$, it is amply significant ($p = 0.0074$), while the co-movements are congruent $73.3\%$ of the time. This is clearly still a majority of congruent co-movements but shows that effect size does not always follow the $p$-value. Of the 37 statistically-significant countries, there are 10 with less than $80\%$ congruent co-movements ($80\%$ could nonetheless still be regarded as a somewhat strict criterion). Luxembourg shows only $51\%$ congruent co-movements, but of course it was far from significant in the G-G test.

The rightmost two columns of Table 3 are for the 25-year windows, and, as noted, serve two purposes. First, because these windows are all of the same width, it puts all the tests on an equal footing by using the same sample size. The choice of 25 years is arbitrary. With longer windows, more countries (e.g., Slovenia) would be using their entire series in both comparisons (with 25 year windows, only Chile has this problem). With shorter windows, the probability of a type II error increases. Second, it is a conservative test: it examines co-movement during the period of greatest divergence between male and female life expectancy, which is when one may theoretically expect the greatest propensity for lack of co-movement. By choosing windows of maximum divergence, we essentially are trying to categorize the countries as belonging to panel (b) or panel (d) of Figure 1.

**Table 3: Goodman-Grunfeld test statistics and $p$-values**

| | Entire Series | | | 25-year Window | | |
|---|---|---|---|---|---|---|
| | G-G test statistic | one-sided $p$-value | proportion congruent | G-G test statistic | one-sided $p$-value | proportion congruent |
| Australia | 5.610 | <.00005 | 0.852 | 3.468 | 0.0003 | 0.958 |
| Austria | 3.386 | 0.0004 | 0.873 | 2.239 | 0.0126 | 0.833 |
| Belarus | 3.979 | <.00005 | 0.792 | 2.514 | 0.0060 | 0.792 |
| Belgium | 9.895 | <.00005 | 0.892 | 2.820 | 0.0024 | 0.833 |
| Bulgaria | 5.627 | <.00005 | 0.873 | 2.097 | 0.0180 | 0.750 |
| Canada | 5.988 | <.00005 | 0.909 | 1.418 | 0.0781 | 0.792 |
| Chile | 1.047 | 0.1476 | 0.812 | 1.047 | 0.1476 | 0.812 |
| Czech Republic | 5.552 | <.00005 | 0.918 | 2.576 | 0.0050 | 0.833 |
| Denmark | 8.779 | <.00005 | 0.852 | 1.073 | 0.1416 | 0.708 |
| Estonia | 3.175 | 0.0007 | 0.750 | 1.593 | 0.0556 | 0.708 |
| Finland | 7.796 | <.00005 | 0.863 | 2.973 | 0.0015 | 0.875 |
| France | 11.574 | <.00005 | 0.917 | 4.089 | <.00005 | 0.958 |
| GDR | 4.926 | <.00005 | 0.909 | 3.229 | 0.0006 | 0.917 |
| FRG | 4.039 | <.00005 | 0.891 | — | — | 0.875 |
| Hungary | 3.901 | <.00005 | 0.780 | 2.465 | 0.0069 | 0.750 |
| Ireland | 2.711 | 0.0034 | 0.763 | 2.723 | 0.0032 | 0.833 |
| Iceland | 5.216 | <.00005 | 0.715 | 3.749 | 0.0001 | 0.917 |
| Israel | 1.930 | 0.0268 | 0.846 | 1.784 | 0.0372 | 0.833 |
| Italy | 9.243 | <.00005 | 0.920 | 3.172 | 0.0008 | 0.958 |
| Japan | 5.523 | <.00005 | 0.935 | 3.132 | 0.0009 | 1.0 |
| Latvia | 3.029 | 0.0012 | 0.731 | 1.496 | 0.0673 | 0.708 |
| Lithuania | 3.361 | 0.0004 | 0.731 | 1.614 | 0.0532 | 0.625 |
| Luxemburg | −0.088 | 0.5351 | 0.510 | −0.429 | 0.6660 | 0.542 |
| Netherlands | 8.682 | <.00005 | 0.862 | 3.702 | 0.0001 | 0.917 |
| New Zealand | 2.439 | 0.0074 | 0.733 | 2.026 | 0.0214 | 0.792 |
| Norway | 7.890 | <.00005 | 0.828 | 1.753 | 0.0398 | 0.667 |
| Poland | 4.548 | <.00005 | 0.863 | 2.328 | 0.0100 | 0.750 |
| Portugal | 5.984 | <.00005 | 0.903 | 4.349 | <.00005 | 1.0 |
| Russia | 4.708 | <.00005 | 0.843 | 2.781 | 0.0027 | 0.833 |
| Slovakia | 3.317 | 0.0005 | 0.780 | 2.709 | 0.0034 | 0.833 |
| Slovenia | 1.341 | 0.0899 | 0.846 | 1.224 | 0.1105 | 0.833 |
| Spain | 7.032 | <.00005 | 0.891 | 3.215 | 0.0007 | 0.917 |
| Sweden | 12.997 | <.00005 | 0.912 | 1.512 | 0.0653 | 0.708 |
| Switzerland | 7.100 | <.00005 | 0.859 | 4.322 | <.00005 | 1.0 |
| Taiwan | 2.854 | 0.0022 | 0.875 | 2.260 | 0.0119 | 0.833 |
| England and Wales | 11.073 | <.00005 | 0.941 | 3.400 | 0.0003 | 0.875 |
| Scotland | 8.090 | <.00005 | 0.846 | 3.373 | 0.0004 | 0.917 |
| Northern Ireland | 4.237 | <.00005 | 0.787 | 2.184 | 0.0145 | 0.833 |
| Ukraine | 4.163 | <.00005 | 0.820 | 1.671 | 0.0473 | 0.750 |
| United States | 6.760 | <.00005 | 0.948 | 3.790 | 0.0001 | 1.0 |

For the 25-year window, nine of the forty countries fail to achieve statistically-significant co-movement (Table 3). Three of these have already been discussed: in Chile, it is not a separate test, since the country only has 14 years of data; in Slovenia (27 years of data total), the 25-year window is nearly the same test as the overall series; and in Luxembourg, the 25-year window has a similar character as the overall 50-year data set. The six countries for which there is a lack of significant co-movement at the 5% level for the 25-year window, despite significant co-movement in the entire data series, are Canada, Denmark, Estonia, Latvia, Lithuania, and Sweden. During their periods of maximum divergence, these countries behave like panel (d) of Figure 1, with divergence accompanied by decoupling of male and female life expectancy (viz., lack of co-movement). However, among the six countries, only Denmark would fail to reject the null at the 10% significance level. There are 24 first-difference values in each of these windows – a sample size that is unusually low, even for country-level studies. For the 25-year window, no G-G test may be performed for West Germany (FRG in Table 3); this stems from the monotonic increase of male life expectancy in this time period.

This study has several limitations. The HMD is not a representative sample of the global population; it only includes countries whose data meet its quality control standards. Caution is therefore warranted when extrapolating from the HMD to the world as a whole. As noted, the HMD data series are not of equal length; this makes country-by-country comparisons of statistical significance problematic. The extreme case, Chile, has a data set with 16 first-differences. We have tried to address this in two ways. First, we use a Neyman-Pearson (not Fisherian) statistical framework (Lehmann 1993 discusses the distinction): results are considered as significant/not-significant without bias toward the large test statistics (i.e., small $p$-values) that come out of the longer time series. Second, we have done a set of comparisons using the same window size (25 years, maximum divergence) for all countries. The choice of maximum divergence is to make the test conservative: intuitively, one would expect divergence to favor non-co-movement, as in panel (d) of Figure 1. However, the choice of 25 years is arbitrary.

## 5. Conclusion

Preston and colleagues (1972) commented, “mortality conditions mirror those in the general society”. This makes sex differences all the more interesting, since mortality sex differentials change frequently. Extreme sex differences are associated with wartime. However, even in peacetime there are a number of large sex differentials, most notably Russia in 1994, when female life expectancy exceeded that of males by 13.7 years. All peacetime sex differences $\geq$11 years occur in former Soviet countries, in 1994 or later.

Given these extreme life expectancy sex differentials, and the divergences documented in Table 2, something is happening differently for males and females. In the HMD countries in the last half century or so, alcohol and tobacco use are heavily implicated in the divergence – and, in some cases, convergence, as women adopt previously "male" behaviors such as tobacco use (e.g. Preston and Wang 2006). This raises the question, are males and females *separate* barometers of conditions in the general society?

This study gives a nuanced answer to that question. When it comes to life expectancy, co-movement is the norm – all but three of the 40 HMD countries show significant co-movement when evaluated on their entire data series. Moreover, the preponderance (31 countries) still show significant co-movement when evaluated on the 25-year most-divergent subset. This is surprising given the extreme nature of these subsets: 25-years is a very short time window, stacking the deck in favor of type II errors; choosing the most-divergent period intuitively would seem to be a set-up for *lack* of co-movement.

We interpret these co-movement findings as follows. Severe or lenient mortality conditions act in the same direction on both sexes. Thus, changes in the overall mortality environment cause male and female life expectancy to co-move. Remarkably, even when male and female life expectancy trends are diverging, this usually applies. Life expectancy sex divergence has much to teach us, and reflects important aspects of society. However, it does not reveal much about "conditions ... in the general society", but, rather, about sex-specific risk behavior (alcohol abuse, tobacco use, violence) or labor force participation (Pampel and Zimmer 1989). On the other hand, the overall trend of mortality does indeed reflect conditions in the general society, but – as demonstrated by the analysis herein – these tend to act similarly on both sexes. As Vallin (1990) noted, "Time trends ... reflect general health progress, which results in a gain in life expectancy for both sexes, but more so for women, among whom adverse health behaviours have traditionally been less common."

Human mortality is a barometer of society, but one in which the mercury rises or falls for both sexes when it rises or falls for either (although not necessarily by the exact-same number of millimeters). Males and females are part of the same society, so divergence does not imply decoupling. Life expectancy time series by sex, without taking the female minus male difference, demonstrate society-wide factors, and these series co-move. This is not to diminish the importance of the divergence of male and female mortality; it is just to emphasize that these illuminate patterns and processes of sex-*specific* behavior such as smoking or drinking, not society-wide factors. Divergence does not imply decoupling: male and female life expectancy usually co-move, even when they diverge. This phenomenon is somewhat more pronounced when measured by statistical significance for co-movement in time series versus prevalence of co-movement on a year-to-year basis (where the highest proportion of congruent co-movements is 94.8%, for the United States). Nonetheless, a majority of countries saw congruent co-movements 85% of the time or more.

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