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Research Article

A transition-based approach to measuring inequality

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A transition-based approach to measuring inequality

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Abstract

The measurement of inequality is often made using observed population-based distributions, such as the distribution of income or the distribution of members of different groups across neighborhoods. Unfortunately, such population distributions confound past and present behavior. Here, we advocate measuring inequality using behavior as reflected by transitions between categories of interest (e.g. income categories) over a specified time period, and show how such measures may be obtained from frequently available data. An illustrative example is provided for the case of income inequality.

The approach is then applied to analyze trends in inequality between men and women in the distribution of ages at death. Observed death distributions indicate that, since 1970, mortality in 4 Western countries experienced increases in inequality that recently leveled off. In contrast, life table death distributions, which solely reflect the implications of a given year's mortality rates, reveal a peak in inequality followed (in 3 of the 4 countries) by appreciable declines. The results are insensitive to whether inequality is measured by entropy, the Gini Index, or the Index of Dissimilarity. However, the type of population distribution analyzed—whether observed or inferred from the transitions—does make a significant difference in the results obtained. Because distributions derived from transitions reflect the inequality implications of behavior over a specified time interval, they are recommended for greater use in analyses of inequality.

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1. Introduction

Social inequalities are important indicators of social stratification, social cohesion, and individual well-being. As a result, ways to measure the extent of inequality in a population have received a great deal of methodological attention in the social sciences. Introductions to the large literature on the topic can be found in Cowell (1995), Jenkins (1991), and Sen (1997). A number of different measures, each with its own set of strengths and weaknesses, have been used in a variety of analyses.

For the most part, inequality measures have been based on population distributions observed at a given time with respect to a particular characteristic of interest. For example, measures of income inequality examine how members of different groups are distributed with regard to their level of income, while measures of residential segregation reflect how members of different groups are distributed with regard to place of residence. That approach is straightforward and reasonable, but such population-based measures for a given year (or period) do not reflect behavior at or near that time. Instead, they reflect the complex interplay of current and past behavior that produced the given year's population distribution. In short, conventional measures do not reflect the inequality implications of behavior at any readily identifiable time period. If the objective of an analysis is to do so, conventional approaches are inappropriate and can produce misleading measures of inequality levels and trends.

This paper provides a way to examine inequality based on transitions during a given period, i.e. on how persons move from one category of the characteristic of interest to another, rather than on the population distributions produced by past and present patterns of movement. Such transition rates or proportions reflect behavior during the given period, independent of previous behavior. Those transition values can be used to measure inequality because they imply a unique, long term population distribution. Thus the behavior of a given period can yield a theoretical population distribution that can be used with standard measures of inequality to calculate the inequality implied by behavior during that period.

The following sections describe how to calculate such transition-based measures of inequality. The long term, theoretical population distribution is first specified, as are the means for computing it. Procedures for determining the appropriate transition measures are then described. To illustrate the process, we first consider the case of income inequality. We then analyze time trends in the sex differential in mortality. For three different measures of inequality and data for four Western countries, we find that different trends in sex differentials emerge depending on whether observed or implicit (theoretical) distributions of deaths are used.

2. Transition measures and their associated population distributions

We want to consider behavior over a specific time period, say 1 to 5 years. Let the population distribution of a given group at the beginning of that interval be described by n-category column vector x_{t-1} whose jth element, $x_{j,t-1}$ gives the number of persons in category j at time $t-1$. For example, $x_{3,t-1}$ could denote the number of persons in income category 3 at time $t-1$. Analogously, the n-category vector x_t, with jth element x_{jt}, represents the population at time t. The behaviors of interest can, in general, be reflected by the elements of $n \times n$ population projection matrix A_t, which takes the time $t-1$ population to time t, and hence satisfies the projection relationship

$$x_t = A_t x_{t-1} \tag{1}$$

Let a_{ijt} denote the element in the ith row and jth column of A_t. Element a_{ijt} is essentially a proportion that represents the fraction of persons in the jth population category at time $t-1$ who are in the ith population category at time t. The elements of A_t thus reflect the results of transitions between categories over time, and describe population behavior over the $t-1$ to t interval.

There are two ways to find the long term (or stable) population distribution implied by A_t. One is simply to project the initial population far into the future by repeated applications of A_t. Ultimately, the relative number of persons in the different population categories will become constant, and those ratios provide the desired population distribution (Keyfitz 1977). Alternatively, we can examine the latent (or eigen) structure of A_t. [To simplify the mathematical presentation, only the basic eigenstructure equations are given here; for a more complete discussion, see Caswell (2001) or Schoen (2006).] Matrix A_t is associated with n roots (or eigenvalues), denoted by λ, which satisfy the matrix equation

$$|A_t - \lambda I| = 0 \tag{2}$$

where the vertical bars indicate a determinant and I is the $n \times n$ identity matrix. With nonnegative elements a_{ijt}, A_t has a unique, real, and largest (or dominant) root whenever $A_t^{n(n-2)+2}$ has all positive elements (which is usually the case for the matrices discussed here). Associated with dominant eigenvalue λ_t is dominant right eigenvector u_t, defined by

$$\lambda_t u_t = A_t u_t \tag{3}$$

The long term distribution of the population implied by A_t is provided by n-element right eigenvector u_t. It is the only right eigenvector of A_t whose elements all have real, nonnegative values. The elements of u_t are only specified up to a scaling factor, but their relative sizes are unique.

Computationally, finding the eigenstructure of a matrix can be quite tedious. In practice, however, mathematical software (e.g. Maple, Mathematica, S+) are readily available. They make finding the roots and eigenvectors of a matrix, and hence the long term population distribution implied by a given population projection matrix, a rather simple task.

When inequalities in mortality are of interest, as in a later application here, a different way of finding the theoretical population distribution of interest may be preferable. Observed age-group-specific death rates (i.e. numbers of deaths for a given group in a given age interval divided by the number of persons in that group in that age interval) may well be available and can be used to calculate a life table (Preston et al. 2001). The life table shows the number of persons in a hypothetical cohort who survive from birth to every age under the given death rates and, via subtraction, gives the number of persons in the cohort dying in every age interval. Those life table cohort deaths can provide an appropriate theoretical distribution.

3. Specifying the projection matrix

Several frequently available sources of data make it possible to create the requisite population projection matrix for many potential areas of application. The essential requirement is that the data provide each person's category at the beginning and end of the projection interval. Prospective or panel data can provide contemporaneous information on each individual's beginning and ending status. Retrospective data can provide information on the previous status of those surviving. [On the assumption of no (or known) differential mortality between categories, survivorship can also be taken into account (Schoen 1988:76-79).] The element in the ith row and jth column of the population projection matrix is then given by the ratio of (i) the number of persons in category i at the end of the interval who were in category j at the beginning of the interval divided by (ii) the total number of persons in category j at the beginning of the interval.

The logic of the population projection matrix relates persons in each category of the ending population to their category at the beginning of the interval. Persons who are born or who migrate into the population of interest during the projection interval must be related to an appropriate initial category, but there is generally a reasonable way to do so. Births can usually be given the category of their parents. Persons initially residing outside the area of interest can be related to their end of interval residential category. Persons who initially had no income can be related to a "no income" category. Such behaviorally plausible assignments make it possible to fully specify the values used in the calculation of the elements of the population projection matrix.

Several assumptions are implicit in the structure of the projection matrices. One is that the observed population is homogeneous with respect to the risk of transfer. Another is the Markovian assumption that only the current state affects the risk of transfer, with past history irrelevant. Both assumptions are strong and typically violated. Past experiences often affect transfer risks, as do individual characteristics such as length of time in a particular category. Events in a specified time interval may be influenced by features of preceding intervals, which are not considered. Even in short time intervals, the risks of transition may change substantially over time. Investigators need to satisfy themselves that the model is sufficiently realistic for the purposes intended. The choice of categories also embodies judgments about the nature of the behavior being studied. For example, a study of income inequality over time should be based on categories that reflect the income distribution over the entire time span considered.

4. An illustration involving income inequality

The unequal distribution of income across groups is a longstanding focus of inequality research. To provide an illustration of how the transition-based approach can be applied, we examine the inequality between two time periods, 1963-66 and 1966-70. That is directly analogous to comparing the income inequality between men and women, between single persons and married persons, or between any specified pair of groups. We use data examined by Shorrocks (1976), which are shown in the Appendix Table. Those data give, for five income categories, (i) the income distribution of Shorrocks' 1963 sample (total number of persons = 800), (ii) a transition (projection) matrix showing the distribution of income in 1966 of persons in each income category in 1963, and (iii) a transition matrix showing the distribution of income in 1970 of persons in each income category in 1966. The Appendix Table also shows the estimated sample population by income category in 1966, obtained by projecting the 1963 population to 1966 using the 1963-66 projection matrix, and the estimated 1970 sample population obtained by projecting the 1966 population using the 1966-70 projection matrix. For the conventional approach, we consider the 1966 income distribution as characterizing "Group A", the 1970 income distribution as characterizing "Group B", and assess the inequality between those groups.

The projection matrices in Panels B and C of the Appendix Table have theoretical population distributions by income category (i.e. dominant eigenvectors) associated with them. Those theoretical income distributions, scaled to yield a total of 800 persons, are shown in the last two columns of Panel A of the Appendix Table. To apply the transition-based approach, we consider the 1963-66 experience as characterizing Group A, the 1966-70 experience as characterizing Group B, and measure the inequality between the theoretical distributions implied by those two transition matrices.

The results, using two standard measures of inequality, are the following:

Measure	Inequality Based on	
	Data Populations	Theoretical Populations
Gini Index	0.1002	0.2282
Index of Dissimilarity	0.0894	0.2147

For both measures, the inequality based on the theoretical populations is more than twice as large as that based on the sample (data) populations. The result is not surprising, because the composition of data populations generally changes gradually, as it represents a complicated weighted average of past behavior. In contrast, theoretical populations are based solely on the behavior of a specified period. As a result, the transition-based approach reflects the full impact of the differences between the 1963-66 and 1966-70 patterns of transitions.

This illustrative application indicates some noteworthy features of the transition-based approach. First, it readily adapts to any number of categories. If there are n categories of interest, the projection matrix has n rows and n columns. Second, the method imposes no restrictions on movements between states. Persons can move freely, and repeatedly, between the states in the model as they do in the actual population. In addition, as long as the underlying rates do not change, the duration of the period of observation does not influence the theoretical population distribution. A longer duration affects the roots of the projection matrix, but not the eigenvectors that determine the theoretical population of interest.

5. Mortality trends and differentials

To investigate whether the use of observed versus theoretical distributions can make a significant difference in substantive analyses of inequality, we examine the case of sex differentials in mortality. The life table approach, as described above, is used to generate the theoretical distribution of deaths based on the experience of a hypothetical cohort, and three different inequality measures are used so that the effect of using observed versus theoretical death distributions can be clearly distinguished from inequality measure effects.

In general, research has found an inverse relationship between inequality in the timing of death and the average length of life. As life expectancy has increased, inequality in age at death has typically declined (Shkolnikov et al. 2003). Edwards and Tuljapurkar (2005) examined crossnational differences in the age pattern of mortality using the standard deviation of life table ages at death above age 10. They found striking differentials since

1960 which could not be explained by aggregate socioeconomic inequalities. As they noted, that has important policy implications, as it points to the need to reduce inequality in longevity as well as to increase life expectancy.

Goesling and Firebaugh (2004) used data for 169 countries to examine trends in between-country life expectancy between 1980 and 2000. Continuing a trend that began in the first half of the 20th century, inequality in the distribution of life expectancy across countries declined in the 1980s. During the 1990s, however, between-country inequality increased. Changes in mortality differentials by sex were explored by Glei and Horiuchi (2007). Using long term data on life expectancy at birth $[e(0)]$ from 29 high-income countries, they found that the difference between a country's male and female $e(0)$ values widened during most of the 1900s. Recently, however, the gender gap in longevity narrowed in most countries considered. Here, we seek to further examine trends in mortality sex differentials to see whether they differ when observed deaths are used instead of life table deaths. Rather than using the difference between male and female longevity, we employ three standard measures of inequality. In every population considered, we calculate our measures separately for life table and observed distributions of deaths, to see whether the transition-based life table distributions and the observed population distributions depict different trends in inequality. To the best of our knowledge, such a comparison has never previously been made.

6. Mortality data and methods

We use data for four countries: England and Wales 1840-2002, Italy, 1880-2002, Sweden, 1840-2002, and the United States, 1900-2002. For the first three countries, all data are from the Human Mortality Database. For the United States, life table values for 1900-1998 are from Haines (2006:Tables Ab988-1047), and data on observed deaths are from (i) the U.S. National Center for Health Statistics (NCHS) (2007) for the years 1900 through 1946 and (ii) the Human Mortality Data Base for the years 1947 through 2002.

In all cases, sex-specific death rates and numbers of deaths were determined, by single years of age, for each data year examined. Single years of age were available in the Human Mortality Database. For the United States, linear interpolation was used to obtain single year values from grouped data. The highest age recognized was 104.

Three measures of inequality are applied to both the behaviorally-based and observed distributions: an entropy index (H), the Gini Index (G), and the Index of Dissimilarity (D). All are leading measures of inequality, possessing most, if not all, of the properties desired in such a measure. In the main, past work has found them to be quite closely related, though each is sensitive to different distributional features and thus yields slightly different results (White 1986). Entropy is a measure of information content (or randomness), and has often been used in mortality analysis (Edwards and Tuljapurkar 2005; Vaupel 1986). Here it is applied following the manner described in White (1986), i.e.

$$H = (H^* - \hat{H})/H^* \tag{4}$$

The overall life course component is

$$H^* = -\sum_k p_k \; lnp_k \tag{5}$$

with the sum over k spanning both gender groups, p_k denoting the proportion of the total number of deaths that are male (or female), and ln indicating the natural logarithm. The minus sign makes the measure a positive quantity, as the logarithm of a proportion is a negative number. The age-group specific component is

$$\hat{H} = -\sum_i \left[d(i)/\ell(0)\right] \sum_k \left[d_k(i)/d(i)\right] ln \left[d_k(i)/d(i)\right] \tag{6}$$

where the sum over i ranges over all ages, the sum over k spans both gender groups, $d(i)$ indicates the total number of life table [or observed] deaths at age i, and $d_k(i)$ indicates the life table [or observed] number of group k deaths at age i. The symbol $\ell(0)$ denotes the total number of observed deaths or the initial (age 0) size of the life table cohort, including both males and females. In the life table calculations, we assumed the conventional sex ratio at birth (srb) of 105 males per every 100 females (though only H is sensitive to the srb used).

The Gini Index is one of the most frequently used measures of inequality, and is a standard measure in studies of income inequality (White 1986). For either gender group k, it can be written as

$$G = \frac{\sum_i \sum_j d(i)\, d(j)\, |\, [d_k(i)/d(i)] - [d_k(j)/d(j)]\, |}{2\, \ell(0)^2 [d_k/\ell(0)]\, [1 - d_k/\ell(0)]} \tag{7}$$

where the sums over i and j range over all ages, the vertical bars indicate an absolute value, and d_k is the total number of deaths in reference gender group k. The Index of Dissimilarity is another frequently used measure, and is a standard in studies of residential segregation (White 1986). For gender groups m and f, it can be written

$$D = (1/2) \sum_i |d_m(i)/d_m - d_f(i)/d_f| \tag{8}$$

where the sum over i ranges across all age groups.

7. Results of the mortality analyses

Figures 1, 2, and 3 show inequality measures H, G, and D, respectively, for England and Wales, Italy, Sweden, and the United States for each available year from 1900 to 2002. Similarly, Tables 1, 2, and 3 present those inequality measures for available years ending in the digit "0", beginning in 1840 and adding the most recent year, 2002. In every case, separate calculations were made based on observed deaths and life table deaths. The particularly pronounced spikes in inequality in England and Wales are associated with the 1918 Pandemic and World War II. The latter may also reflect administrative procedures followed in recording wartime deaths (personal communication from Seamus Spark, January 3, 2007). To a lesser extent, a similar war-related phenomenon may have occurred in Italy during 1944 and 1945.

The results indicate that, over most of the 20th century, gender inequality in mortality was increasing in all four countries considered. That is consistent with previous work, and is reflected by all three measures and by both observed and life table death distributions. Since about 1970, however, the trends depicted by observed and life table distributions have diverged.

In England and Wales, for all 3 measures, the observed death distributions indicate that inequality levels have been roughly constant since at least 1976. In contrast, the life table death distributions, for all three measures, peak in 1972 and have been declining since. Between 1975 and 2000, the life table based H declined by 27%, the G by 23%, and the D by 25%. The pattern in the United States is quite similar. According to all three measures, there is little recent change when observed death distributions are used. In contrast, life table distributions indicate that gender inequality peaks in the United States in 1975, and declines substantially afterwards. The pattern in Sweden is much the same. When observed deaths are used, there is little change in the three inequality measures since the early 1980s. However, when life table deaths are employed, all three measures show peaks in the mid-1980s and substantial declines thereafter. Italy deviates from the other three countries, but only in degree. When observed death distributions are used, inequality in Italy increases until the late 1980s, and then changes little through the 1990s. Life table death distributions show a small peak in inequality in 1991, followed by modest (5-9%) declines through 2002.

Table 1: Entropy measure (H) values of male/female inequality in mortality in four western countries, 1840-2002

A. Based on observed deaths

Year	England and Wales	Sweden	Italy	United States
1840	NA	0.00932	NA	NA
1850	0.00417	0.01039	NA	NA
1860	0.00446	0.00756	NA	NA
1870	0.00310	0.00810	NA	NA
1880	0.00302	0.00741	0.00160	NA
1890	0.00369	0.00722	0.00229	NA
1900	0.00434	0.00634	0.00219	0.00277
1910	0.00579	0.00620	0.00244	0.00369
1920	0.00887	0.00746	0.00301	0.00324
1930	0.00963	0.00611	0.00192	0.00332*
1940	0.01854	0.00766	0.00248	0.00544
1950	0.01663	0.00769	0.00488	0.01031
1960	0.03008	0.01059	0.01327	0.01787
1970	0.04463	0.01860	0.02537	0.02528
1980	0.05098	0.03134	0.04238	0.03652
1990	0.05071	0.04008	0.05191	0.04529
2000	0.04903	0.03959	0.05289	0.04377
2002	0.04853	0.04134	0.05324	0.04320

Table 1: (Continued)

B. Based on life table deaths

Year	England and Wales	Sweden	Italy	United States
1840	NA	0.03185	NA	NA
1850	0.02657	0.03434	NA	NA
1860	0.02681	0.03048	NA	NA
1870	0.02736	0.03061	NA	NA
1880	0.02862	0.02953	0.02550	NA
1890	0.02987	0.02820	0.02588	NA
1900	0.02990	0.02927	0.02535	0.02685
1910	0.03182	0.02828	0.02566	0.02854
1920	0.03456	0.02800	0.02534	0.02642
1930	0.03683	0.02660	0.02740	0.02952
1940	0.04871	0.02825	0.02905	0.03525
1950	0.04958	0.02938	0.03248	0.04537
1960	0.06443	0.03743	0.04600	0.05789
1970	0.07250	0.05062	0.05681	0.06949
1980	0.07198	0.06834	0.07442	0.07069
1990	0.06553	0.06666	0.07388	0.06327
2000	0.05337	0.05788	0.06918	0.04953**
2002	0.05002	0.05440	0.06849	NA

* US NCHS data for 1931; ** Haines data for 1998; NA: Not available

Table 2: Gini index values of male/female inequality in mortality in four western countries, 1840-2002

A. Based on observed deaths

Year	England and Wales	Sweden	Italy	United States
1840	NA	0.11840	NA	NA
1850	0.08505	0.12344	NA	NA
1860	0.08681	0.10373	NA	NA
1870	0.07150	0.10769	NA	NA
1880	0.07108	0.10641	0.05202	NA
1890	0.07782	0.10622	0.06111	NA
1900	0.08348	0.10149	0.06022	0.06544
1910	0.09501	0.10346	0.06379	0.07899
1920	0.11803	0.11223	0.06967	0.07281
1930	0.11847	0.09936	0.05733	0.07179*
1940	0.17072	0.11633	0.06401	0.09508
1950	0.16551	0.11267	0.09204	0.13223
1960	0.22969	0.13495	0.15234	0.17443
1970	0.27791	0.18096	0.21279	0.20795
1980	0.28914	0.23591	0.27487	0.24796
1990	0.28624	0.26244	0.30273	0.27465
2000	0.28545	0.26001	0.30469	0.27046
2002	0.28500	0.26592	0.30508	0.26898

Table 2: (Continued)

B. Based on life table deaths

Year	England and Wales	Sweden	Italy	United States
1840	NA	0.11599	NA	NA
1850	0.06805	0.13449	NA	NA
1860	0.07094	0.10424	NA	NA
1870	0.07671	0.10536	NA	NA
1880	0.08883	0.09615	0.05547	NA
1890	0.10060	0.08542	0.05903	NA
1900	0.09971	0.09670	0.05215	0.07318
1910	0.11303	0.08894	0.05511	0.09214
1920	0.13175	0.08266	0.05057	0.06712
1930	0.14487	0.06932	0.07710	0.10030
1940	0.20544	0.08878	0.09409	0.14228
1950	0.21276	0.09700	0.12216	0.19869
1960	0.26949	0.15656	0.20098	0.24995
1970	0.29217	0.22062	0.24680	0.28564
1980	0.28535	0.28221	0.30462	0.28761
1990	0.26279	0.27626	0.30279	0.26384
2000	0.22569	0.24257	0.28686	0.21637**
2002	0.21470	0.23043	0.28515	NA

* US NCHS data for 1931; ** Haines data for 1998; NA: Not available

Table 3: Dissimilarity index values of male/female inequality in mortality in four western countries, 1840-2002

A. Based on observed deaths

Year	England and Wales	Sweden	Italy	United States
1840	NA	0.09064	NA	NA
1850	0.06063	0.09469	NA	NA
1860	0.06316	0.07815	NA	NA
1870	0.05088	0.07892	NA	NA
1880	0.05154	0.07655	0.03926	NA
1890	0.05724	0.07406	0.04670	NA
1900	0.06143	0.07163	0.04438	0.04812
1910	0.06858	0.07405	0.04571	0.05506
1920	0.08578	0.08099	0.05052	0.05585
1930	0.08578	0.06874	0.04254	0.05247*
1940	0.12187	0.08598	0.04712	0.06720
1950	0.12244	0.07995	0.06924	0.09652
1960	0.17236	0.10054	0.11566	0.12824
1970	0.21394	0.14090	0.17022	0.15544
1980	0.22475	0.17836	0.21009	0.18684
1990	0.21679	0.19831	0.22593	0.20054
2000	0.21558	0.19325	0.23155	0.19740
2002	0.21435	0.19902	0.23186	0.19613

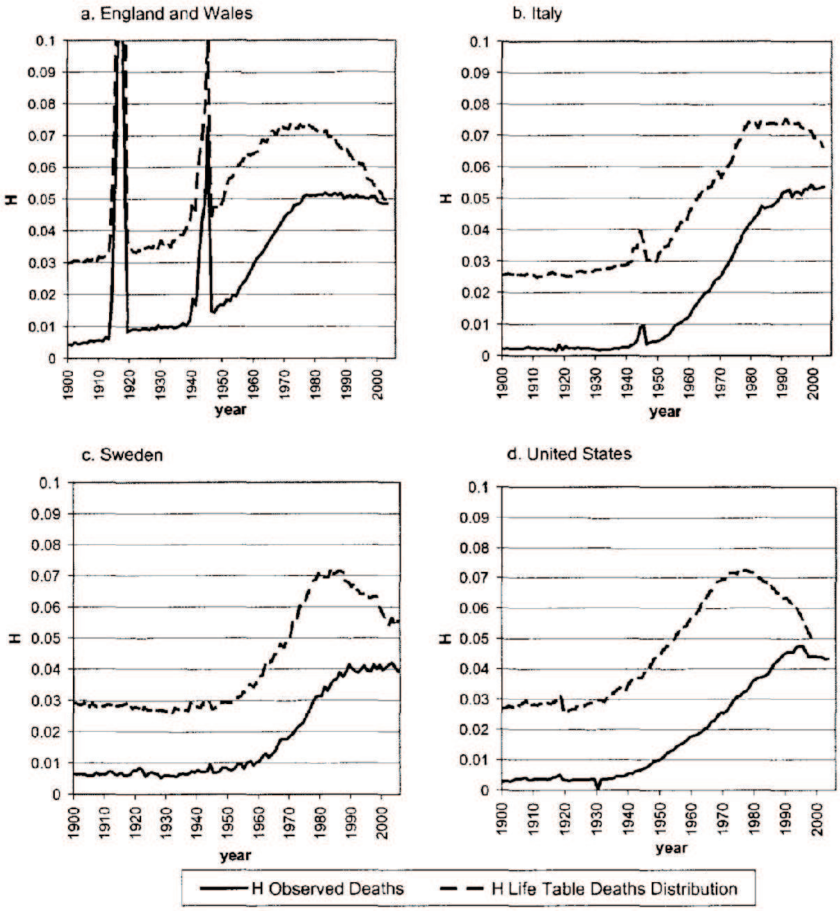
Table 3: (Continued)

B. Based on life table deaths

Year	England and Wales	Sweden	Italy	United States
1840	NA	0.08784	NA	NA
1850	0.04756	0.10344	NA	NA
1860	0.04976	0.07472	NA	NA
1870	0.05352	0.07778	NA	NA
1880	0.06271	0.06826	0.04176	NA
1890	0.07335	0.05892	0.04432	NA
1900	0.07592	0.06933	0.03816	0.05052
1910	0.08425	0.06254	0.04044	0.07006
1920	0.09912	0.05716	0.03550	0.04603
1930	0.10856	0.04859	0.05519	0.07540
1940	0.14894	0.06288	0.06784	0.11012
1950	0.15856	0.06976	0.08765	0.15240
1960	0.20351	0.11533	0.14843	0.19254
1970	0.22222	0.16743	0.18790	0.21894
1980	0.21860	0.21184	0.22791	0.21986
1990	0.20132	0.20883	0.22499	0.19967
2000	0.16910	0.17959	0.21373	0.16387**
2002	0.16122	0.17171	0.21462	NA

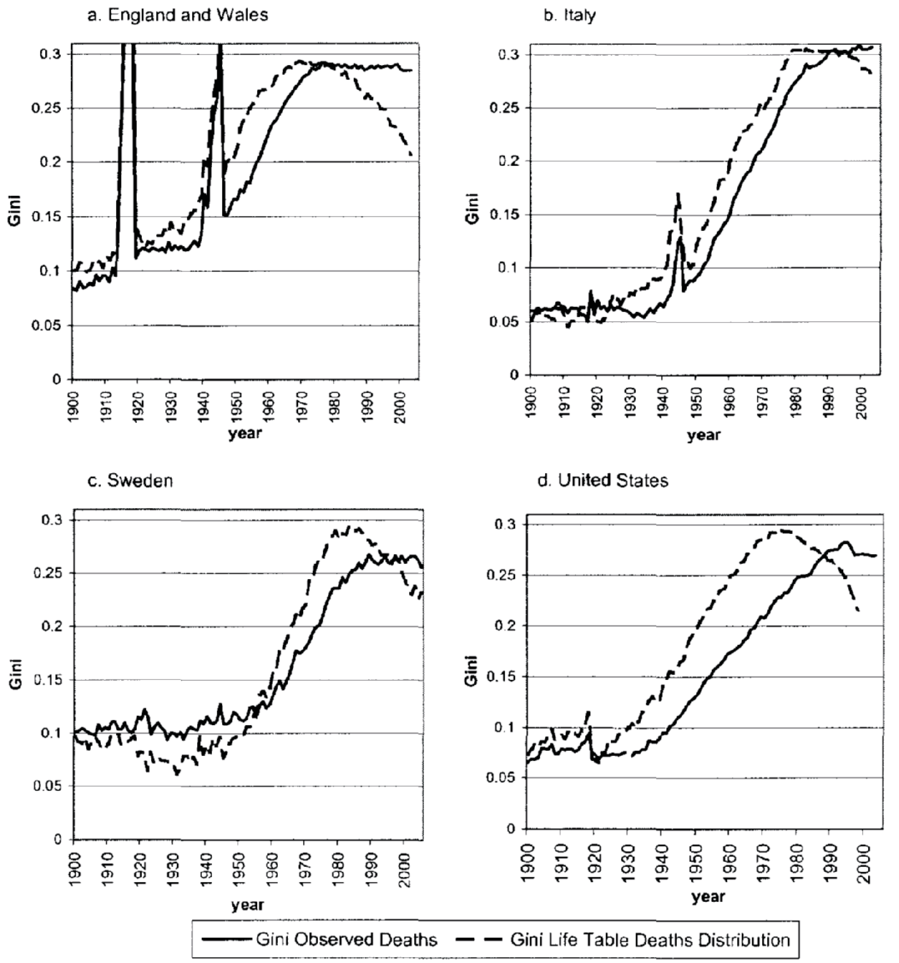
* US NCHS data for 1931; ** Haines data for 1998; NA: Not available

Figure 1: Entropy measure (H) of male/female inequalities in mortality over time in four western countries, 1900-2000



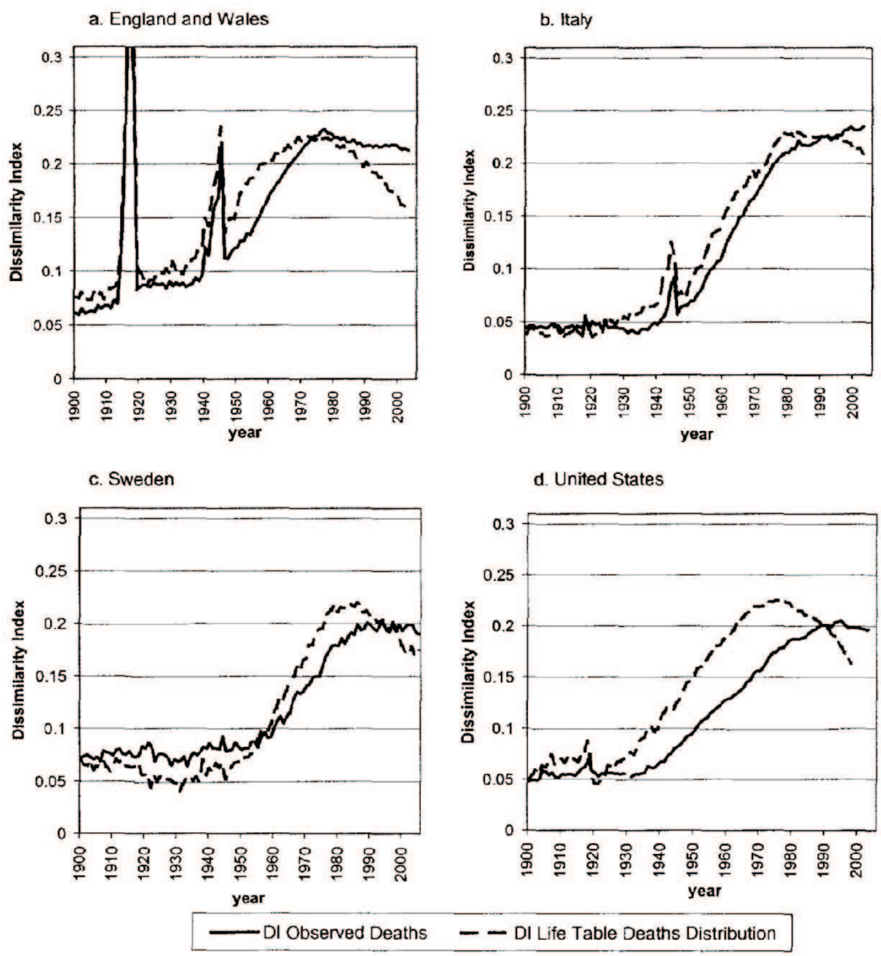
Note: Entropy values for England and Wales in several years are off scale. Those values are (i) for life table deaths: 1916 (0.14506), 1917 (0.19274), 1918 (0.14547) and 1945 (0.10075) and (ii) for observed deaths: 1917(0.12552)

Figure 2: Gini index of male/female inequalities in mortality over time in four western countries, 1900-2000



Note: Gini Index values for England and Wales in several years are off scale. Those values are (i) for life table deaths: 1916 (0.42598), 1917 (0.50929), 1918 (0.45388) and 1945 (0.33548) and (ii) for observed deaths: 1916(0.38156), 1917(0.43641) and 1918(0.34965)

Figure 3: Dissimilarity index of male/female inequalities in mortality over time in four western countries, 1900-2000



Note: Dissimilarity index values for England and Wales in several years are off scale. Those values are (i) for life table deaths: 1916 (0.31706), 1917 (0.3895) and 1918 (0.33861) and (ii) for observed deaths: 1916(0.30234) and 1917(0.36040)

8. Summary and conclusions

The trend in gender inequality in mortality since the 1970s is markedly different when inequality is based on observed death distributions rather than life table death distributions. The life table distributions, which show the implications of the death rates observed in a given year, show a peak in inequality followed by a notable decline. In England and Wales and the United States, a clear peak in the 1970s was followed by a marked decline. Italy showed a small peak in 1991 followed by a small decline. Sweden was intermediate, with a peak in the 1980s followed by a decline of over 20%. Those patterns do not depend on whether inequality is measured by entropy (H), the Gini Index, or the Index of Dissimilarity. They are also consistent with the pattern found by Glei and Horiuchi (2007) based on differences between male and female life expectancy at birth.

Recent trends in inequality in mortality are quite different when based on observed death distributions, which confound current mortality with past demographic behavior. Again the three inequality measures provide the same results—a longer continuation of the increase in inequality followed by a plateau extending to 2002. No decline in inequality was evident in any of the four countries considered. It is difficult to argue that including the effects of past fertility, mortality, and migration improves the measurement of gender inequality in mortality in any given year. The use of observed death distributions thus disguises recent trends in inequality.

Much of the literature on inequality, including most studies of income inequality and residential segregation, use data analogous to the observed death distributions. There is thus a real possibility that reported trends in inequality may be compromised by the use of data that confounds current and past behavior. The illustrative example on income inequality in Section 4 reinforces that view. The transition-based approach described here, which uses longitudinal or retrospective data to produce time-specific behavioral measures, is therefore recommended as a feasible method for measuring the inequality implications of behavior observed during a specified period of time.

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Appendix table: Illustrative data on income inequality

A. Population by income category

Category	1963 Sample	Projected 1966	Projected 1970	Theoretical, based on matrix for 1963-66	1966-70
1	76	85.1	124.5	95.7	188.1
2	212	204.5	180.3	199.4	144.1
3	256	245.3	210.3	236.1	160.3
4	164	175.8	163.5	179.5	139.0
5	92	89.3	121.4	89.3	168.5

B. Income category1963

		1	2	3	4	5
Income category In 1966	**1**	0.64	0.14	0.02	0.01	0.00
	2	0.29	0.56	0.22	0.04	0.01
	3	0.04	0.26	0.54	0.27	0.05
	4	0.03	0.03	0.21	0.54	0.27
	5	0.00	0.01	0.01	0.14	0.67

C. Income category1966

		1	2	3	4	5
Income category In 1970	**1**	0.78	0.22	0.05	0.00	0.01
	2	0.15	0.50	0.23	0.05	0.00
	3	0.07	0.24	0.45	0.23	0.05
	4	0.00	0.03	0.25	0.45	0.19
	5	0.00	0.01	0.02	0.27	0.75

Note: Income transition matrices and 1963 sample population from Shorrocks (1976).Projected and theoretical populations calculated as described in the text. In every case, the total number of persons is 800.