Demography, Vol. 27, No. 2, May 1990

Mortality Differentials by Marital Status: An International Comparison

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Although the greater longevity of married people as compared with unmarried persons has been demonstrated repeatedly, there have been very few studies of a comparative nature. We use log-linear rate models to analyze marital-status–specific death rates for a large number of developed countries. The results indicate that divorced persons, especially divorced men, have the highest death rates among the unmarried groups of the respective genders; the excess mortality of unmarried persons relative to the married has been generally increasing over the past two to three decades; and divorced and widowed persons in their twenties and thirties have particularly high risks of dying, relative to married persons of the same age. In addition, the analysis suggests that a selection process is operating with regard to single and divorced persons: the smaller the proportion of persons who never marry or who are divorced, the higher the resulting death rates.

The greater longevity of married people as compared with unmarried persons has been repeatedly demonstrated throughout the 20th century in a large number of countries. Although some analysts have argued that errors in reports of age and marital status can give rise to these observed differentials, the evidence strongly indicates that these differences remain sizable after correction for possible misreporting. In fact, these differences persist even when the effects of socioeconomic status and other observable factors are controlled.

Several hypotheses have been put forth to explain the more favorable mortality experiences of the married. One category of hypotheses posits a beneficial protective effect of marriage stemming from a variety of environmental, social, and psychological factors that make the married state a healthier one than the unmarried one. Departures from the married state may be particularly unhealthy: for example, some analysts argue that the widowed state is characterized by very high death rates in the early durations because of bereavement effects. A second category is based on the premise that marriage is selective: it selects healthier persons and leaves among the single a higher proportion of persons with serious health problems. Similarly, remarriage selects the healthiest individuals from among widowed and divorced persons. For example, a very recent study of a British cohort born in 1946 demonstrated that a significant fraction of single persons were handicapped (Kiernan 1988). Most researchers would maintain that a *combination* of selection and behavioral/environmental factors have, in fact, produced the observed differences.

In spite of the extensive research in this area, most of the studies have been confined to a particular country and even to a given year (or short time frame) within the country. As a result, there is little information of a comparative nature. For example, which unmarried state is generally characterized by the highest death rate? Has the excess mortality of unmarried persons compared with married persons generally increased or has it decreased over the past few decades? In which age groups do single and formerly married persons suffer the highest risks of dying relative to married people? One of the primary objectives of this analysis is to examine death rates by marital status across a large number of countries and for various time periods within each country in an attempt to answer such questions. Although several recent studies have been comparative in nature (Hu 1987; Kisker & Goldman 1987; Livi-Bacci 1985), they have either not specifically addressed these questions or failed to obtain satisfactory answers to them.

A second goal of this analysis is to attempt to assess the importance of selection factors in explaining the higher mortality rates of single and divorced persons. We do this by examining the relation between the relative size of these marital groups (i.e., the percentage of people in an age group who are single or divorced) and the magnitude of the mortality rate for these groups. The underlying hypothesis is that populations in which the vast majority of persons marry should be characterized by greater selectivity effects among those who remain single than populations in which substantial proportions never marry. Similarly, in populations in which divorce is rare or in which the majority of divorced persons remarry, those who are divorced or who remain divorced should be more selected with regard to their underlying health or mortality risks. Two studies (Kisker & Goldman 1987; Livi-Bacci 1985) examined such a relation between relative size and excess mortality among single persons and found a fairly strong correlation. Since neither study was multivariate in nature, however, there were no controls for other important factors, such as temporal trends in these variables.

Keep in mind that unfortunately, no analysis based on cross-sectional data can properly address this issue of selection and mortality. We can only stress the importance of large-scale longitudinal studies of health and mortality. We also recognize that an experimental design based on random assignment of persons to the various marital states, although justifiable on scientific grounds, will remain socially unacceptable.

Data

Since we make no allowance for possible errors in reports of numbers of deaths or numbers of persons, we confine our analysis to countries that have reasonably high-quality mortality and population data. The multivariate analysis is based on reported numbers of deaths and person-years of exposure (i.e., approximate midyear population), classified by age group, marital status, and gender for 16 developed countries. The data were obtained from published vital statistics and censuses for particular countries as well as from the *United Nations Demographic Yearbooks*.[1] Since one purpose of this study is to examine changes in mortality differentials over time, only countries that had the necessary data for a period spanning at least a decade (and preferably for at least two decades) were included in the analysis. The resulting countries (listed in Table 1) include two in North America, two in Asia, and the remainder in Europe. Although we would have preferred a wider geographic distribution of countries, this group of countries allows us to explore whether there are any regional deviations with regard to the two Far Eastern and two North American countries and whether there are any distinct geographic or cultural patterns within Europe.

For most of the countries included in this analysis, death rates by marital status were reported only for selected years (e.g., every fifth year), which are listed in Table 1.[2] In all but one country, the relevant data are categorized into five adult age groups: 20–24, 25–34, 35–44, 45–54, and 55–64.[3] In the majority of countries, death rates can be calculated for each of the four marital statuses (single, married, widowed, and divorced) for each of the selected years, but in several, the categories of widowed and divorced have been combined for some of the years.[4]

Log-Linear Models of Mortality

Our primary objective is to assess the effects of marital status on mortality. Since it is clear that these effects vary by age and time period, it is important to consider the effects of

Table 1. Countries (and Years) Included in the Analysis

Country	Years	No. of Cells Male	No. of Cells Female
Austria	1966, 1973, 1978	57	60
Canada	1966, 1971, 1976	57	60
Denmark	1959, 1965, 1972, 1979, 1984	93	94
England and Wales	1965, 1971, 1979, 1984	69	70
Finland	1950, 1959, 1965, 1972, 1979, 1984	105	106
France	1956, 1961, 1966, 1971, 1976, 1981	120	120
Hungary	1960, 1966, 1973, 1979, 1984	91	95
Japan	1955, 1960, 1965, 1970, 1975, 1980	120	120
Netherlands	1966, 1973, 1979, 1984	77	76
Norway	1950, 1960, 1965, 1970, 1974, 1984	66	64
Portugal	1950, 1960, 1966, 1972	72	74
Scotland	1960, 1965, 1973, 1979, 1984	75	77
Sweden	1955, 1960, 1965, 1970, 1975, 1980	114	119
Taiwan	1975, 1977, 1979, 1981, 1983, 1985	144	144
U.S.A	1940, 1950, 1960, 1980	80	80
West Germany	1950, 1965, 1972, 1979, 1984	90	96

each covariate, as well as potential interactions among them, on mortality. On the basis of previous studies, it is also evident that the effects of marital status on mortality vary substantially by country and gender. Rather than include country and gender as additional covariates, however, we have fit separate models for males and females in each of the 16 countries.

For purposes of statistical estimation, we assume that the numbers of deaths (D) in the various categories of age group, year, and marital state are independent Poisson random variables with means equal to the product of the number of person-years of exposure (N) in that category and an underlying death rate (μ). We can express the expected number of deaths as

$$E(D_{ijk}) = N_{ijk} \times \mu_{ijk},$$

where i, j, and k denote the age group, year, and marital group, respectively. This expression leads naturally to a log-linear model for the expected number of deaths:

$$\log E(D_{ijk}) = \log N_{ijk} + \log \mu_{ijk}.$$

The right side of the equation consists of the following terms: the (natural) logarithm of the exposure in each cell (the "offset") and the logarithm of the underlying hazard or mortality risk, which is the target of our model building. In this analysis, we use four separate linear models for $\log(\mu)$, ranging in complexity from a simple additive model to one that incorporates several interaction terms. The use of several models rather than a single "best" model allows us to accomplish our two-fold objective: (1) to use the models in an *exploratory* fashion, that is, to summarize and describe the patterns of mortality across the 16 countries (e.g., the average age-specific patterns of mortality for each marital state), and (2) to find a model that fits all of the observed data sets well so that we can assess whether the size of the marital group is an important variable in explaining the observed differentials in mortality.

For example, we use the following additive model to determine the relative ranking of the four marital states with regard to the level of mortality, controlling for age group and period:

$$\log E(D_{ijk}) = \log(N_{ijk}) + \eta + \alpha_i + \beta_j + \gamma_k,$$

where α_i denotes the effect of the ith age group, β_j the effect of the jth year, and γ_k the effect of the kth marital group. In this additive model, each of the three covariates is treated as categorical (i.e., as a series of dummy variables). When we introduce interaction terms involving time in subsequent models, however, such as the interaction between year and marital group, we treat year as a continuous variable (Y) and consider both linear and quadratic terms. This decision was reached after we tested a considerable number of different models. The resulting statistics indicate a substantially better fit when the main effect of period is modeled as a factor, rather than as a continuous variable. There is only a modest reduction in fit, however, and often a substantial gain in degrees of freedom, when interactions involving year are modeled with linear and quadratic terms as opposed to a full set of dummy variables.

One additional covariate is considered in the modeling process: the size of the marital state. Since size is measured as the percentage of all persons in that age group who are in the particular marital status, this variable is most naturally treated as continuous. As described later, we explored various transformations of size as well as interactions between size and the other covariates. In our final model (model 4) we use the logarithm of size (R), and in order to explore the different effect of this variable across marital states, we include it in the form of an interaction term with marital status.

The four models used in this analysis are presented in Table 2.[5] The first three models are used to describe and summarize the patterns of mortality across countries and genders. The first model, which is the only additive or proportional-hazard model, is used to explore the effects of being in a particular marital state, controlling for age group and period, on the death rate. The second model, which includes an additional term to capture the interaction between age and marital status, is used to describe the average age-specific mortality patterns in the various marital states. The third model, which includes additional interaction terms for the relationship between year and marital state, allows us to explore changes in mortality differentials by marital status over time. Finally, the fourth model is used to determine the importance of the size of the marital group in explaining these mortality differentials, once we have taken into account the most important interactions among the other covariates.

The actual number of observations (deaths and exposure counts) for each data set (i.e., country and gender) was determined by the number of categories of age group, year, and marital status for which deaths and persons were reported. This number of cells varies across the data sets for two reasons: (1) There is variation across countries in the number of periods and age groups for which data are published. (2) We eliminated all cells in which the number of persons was less than 50 (small cells occurred frequently for widows in the first age group). The number of cells included in the analysis, which ranges from about 60 in Austria and Canada to 144 in Taiwan, is shown on the right side of Table 1. The models were fit to the observed numbers of deaths in these cells with the statistical package GLIM.[6]

Table 2. Models Used in Analysis

Model	$\log(\mu_{ijk})$
1. $A + P + M$	$\eta + \alpha_i + \beta_j + \gamma_k$
2. $(A \times M) + P$	$\eta + \alpha_i + \beta_j + \gamma_k + (\alpha\gamma)_{ik}$
3. $(A \times M) + P + [M \times (Y + Y^2)]$	$\eta + \alpha_i + \beta_j + \gamma_k + (\alpha\gamma)_{ik} + a_kY + b_kY^2$
4. $(A \times M) + P + [(A + M) \times (Y + Y^2)] + (M \times R)$	$\eta + \alpha_i + \beta_j + \gamma_k + (\alpha\gamma)_{ik} + a_kY + b_kY^2 + c_iY + d_iY^2 + e_kR$

Note: A denotes age group (categorical variable); P denotes year (categorical variable); M denotes marital status (categorical variable); Y denotes year (continuous variable); R denotes the logarithm of the percentage in the marital state (continuous variable).

One problem that arose during the model-fitting process was the determination of goodness of fit. Since for each data set, the total exposure consists of either the male or female adult population (ages 20–64) for several years in the specified country, the sample sizes are enormous. Thus any relatively parsimonious model is almost certain to be rejected on the basis of the resulting chi-squared statistic and degrees of freedom, even though the model visually appears to fit quite well (Rodríguez & Cleland 1988). In fact, in all of the countries, the first three models considered here would be rejected on the basis of the resulting chi-squared statistic[7] (i.e., the p value that results from a comparison of the given model with the saturated model is always less than .01). For the majority of countries, the fourth model would also be rejected on this criterion. (Statistical comparisons of the models, however, based on differences in the chi-squared statistic, indicate that for almost all countries, the fourth model is significantly better than the others.)

As a consequence of these large exposures, we calculated two additional measures of fit: the percentage reduction of the deviance of the null model, a measure that corresponds to R^2 in classic linear regression, and the chi-squared goodness-of-fit statistic divided by the total exposure in a particular data set. This latter measure has been proposed by several statisticians (e.g., Bishop, Fienberg, & Holland 1975; Rodríguez & Cleland 1988) as a way of converting the chi-squared test of significance into a measure of *absolute* lack of fit. By either of these two measures, any of the four proposed models would be judged to fit almost all of the data sets exceedingly well. For example, the reduction in deviance from the null model exceeds 98% in the majority of countries for even the simple additive model; for the fourth model, it is close to 100% in almost all cases. The standardized chi-squared statistic is no larger than 10^{-6} in all cases; Rodríguez and Cleland consider a value of .03 to represent a close fit. The chi-squared statistic, the corresponding degrees of freedom, and the percentage reduction in deviance are shown in Table 3 for each of the four models.

Results

Relative Mortality Ratios

Our first objective is to determine the relative ranking of marital states with regard to the levels of mortality. Since the divorced and widowed states clearly consist of persons far older than those in the single state, one needs to control for the underlying ages. We use the simple additive model, based on the factors of age, year, and marital state, to estimate the effects of marital status across countries and genders:

$$\log E(D_{ijk}) = \log(N_{ijk}) + \eta + \alpha_i + \beta_j + \gamma_k,$$

or

$$\log \mu_{ijk} = \eta + \alpha_i + \beta_j + \gamma_k. \tag{1}$$

In this representation, γ_i, $i = 1, \ldots, 4$, denotes the effects of being married, single, widowed, and divorced, respectively.

In the calculation of parameter estimates, GLIM treats the first category of each factor as a reference cell, and as such it is assigned a parameter estimate of zero. Hence all remaining parameter estimates for a given factor can be interpreted as effects relative to that for the reference cell. Since married is taken to be the reference category for marital status, the parameter estimates $\hat{\gamma}_2$, $\hat{\gamma}_3$, and $\hat{\gamma}_4$ measure the effects of being in the single, widowed, and divorced states, respectively, *relative* to being married. Since we are modeling the *logarithm* of the force of mortality, however, we must exponentiate the parameter estimates to determine the effect of the various covariates on the death rate itself.

Table 3. Comparison of Model Fitting

Country	Total person-years (in 1000s)	Null Model		Model 1			Model 2			Model 3			Model 4		
		χ^2	df	χ^2	df	% χ^2 reduction [a]	χ^2	df	% χ^2 reduction [a]	χ^2	df	% χ^2 reduction [a]	χ^2 [b]	df	% χ^2 reduction [a]
Males															
Austria	5,784	35,227	56	433	47	98.77	157	36	99.56	116	30	99.67	18*	18	99.95
Canada	17,077	91,571	56	1,228	47	98.66	574	35	99.37	224	29	99.75	170	17	99.81
Denmark	6,918	39,592	92	545	81	98.62	180	70	99.54	128	64	99.68	84	52	99.79
England and Wales	54,921	385,996	68	1,782	58	99.54	549	46	99.86	321	40	99.92	44	28	99.99
Finland	7,584	54,624	104	1,197	92	97.81	701	80	98.72	611	74	98.88	136	62	99.75
France	82,680	492,761	119	6,997	107	98.58	3,094	95	99.37	1,601	89	99.68	329	77	99.93
Hungary	14,312	95,334	90	1,067	79	98.88	578	67	99.39	367	61	99.61	178	49	99.81
Japan	172,558	839,040	119	18,477	107	97.80	7,319	95	99.13	2,571	89	99.69	805	77	99.90
Netherlands	15,434	81,791	76	658	66	99.20	170	54	99.79	134	48	99.84	41*	36	99.95
Norway	5,205	23,801	65	407	54	98.29	215	43	99.10	180	39	99.24	48	27	99.80
Portugal	9,000	49,933	71	1,604	61	96.79	1,318	50	97.36	1,067	44	97.86	110	32	99.78
Scotland	7,047	59,203	74	311	63	99.47	169	52	99.71	80	46	99.87	45*	34	99.92
Sweden	13,755	63,277	113	1,031	101	98.37	430	90	99.32	195	84	99.69	109	72	99.83
Taiwan	22,914	21,811	143	1,274	130	94.16	489	115	97.76	455	109	97.92	248	95	98.86
U.S.A.	167,782	108,537	79	11,625	69	89.29	6,556	57	93.96	4,287	51	96.05	430	39	99.60
West Germany	80,017	463,823	89	5,966	78	98.71	3,201	66	99.31	2,495	60	99.46	797	48	99.83
Females															
Austria	6,438	20,740	56	182	47	99.12	57	35	99.73	33	29	99.84	20*	17	99.91
Canada	17,053	45,584	59	497	50	98.91	126	38	99.72	80	32	99.82	25*	20	99.95
Denmark	6,754	21,003	93	252	82	98.80	165	70	99.22	140	64	99.33	76	52	99.64
England and Wales	56,044	198,593	69	1,329	59	99.33	763	47	99.62	631	41	99.68	108	29	99.95
Finland	7,947	25,258	105	580	93	97.70	464	81	98.16	386	75	98.47	71*	63	99.72
France	82,603	216,197	119	1,979	107	99.08	635	95	99.71	381	89	99.82	187	77	99.91
Hungary	15,185	53,358	94	679	83	98.73	364	71	99.32	235	65	99.56	138	53	99.74
Japan	181,511	507,333	119	11,829	107	97.67	5,803	95	98.86	2,597	89	99.49	611	77	99.88
Netherlands	14,960	35,858	75	380	65	98.94	142	53	99.60	122	47	99.66	44*	35	99.88
Norway	5,035	12,497	63	258	52	97.93	173	42	98.61	163	38	98.69	45	26	99.64
Portugal	9,568	26,074	73	1,185	63	95.45	1,021	52	96.08	959	46	96.32	57	34	99.78
Scotland	7,144	29,937	76	252	65	99.16	135	53	99.55	120	47	99.60	43*	35	99.86
Sweden	13,597	36,913	118	466	106	98.74	185	94	99.50	120	88	99.68	84*	76	99.77
Taiwan	21,281	11,639	143	789	130	93.22	233	115	98.00	215	109	98.15	151	95	98.71
U.S.A.	172,540	610,050	79	8,384	69	98.63	6,654	57	98.91	4,374	51	99.28	519	39	99.91
West Germany	88,953	281,524	95	3,761	84	98.66	2,547	72	99.10	1,776	66	99.37	576	54	99.80

[a] Relative to the null model.

[b] Contrary to convention, * indicates that the model would not be rejected at the 5% level ($p > .05$).

For much of the analysis, we use these exponentiated parameter estimates as a measure of mortality differentials by marital status. Removing ourselves temporarily from the jargon of multivariate models, we note that these estimates are *relative mortality ratios* (RMRs) that adjust for specified covariates. For example, in the additive model, $\exp(\hat{\gamma}_2)$ is the (estimated) death rate of the single population relative to the (estimated) death rate of the married population, adjusted for period and age group.[8] RMRs, which always consider the death rate of the unmarried group relative to the married population, have commonly been used to compare differentials across ages, over time, or among countries (e.g., Gove 1972, 1973; Kobrin & Hendershot 1977; Livi-Bacci 1985). The RMRs used in most previous studies, however, are based directly on reported death rates and usually do not control for other correlates of mortality.[9]

Figure 1 shows the estimated RMRs for males and females in each of the three unmarried groups in each of the 16 countries. Note that because these estimates are derived from an additive model (model 1), the estimated RMRs are constant over age group and period. The graph reveals several interesting findings. First, the ratios are systematically larger for males than for females: the ratios for males lie between 1.6 and 3, with an average of 2.0; with only two exceptions, the ratios for females are below 2, and they average to about 1.5. These gender differences, which have been pointed out by many researchers (e.g., Gove 1972, 1973; Koskenvuo, Kaprio, Lonnqvist, & Sarna 1986), have led to considerable speculation about the ways in which the protective effects of marriage are of a greater health benefit to men than women.

The estimated mortality ratios also indicate that in all but three countries (Portugal, Taiwan, and France), divorced males have the highest ratios among unmarried men. Differences between single males and widowed males are not consistent and are generally not large. The findings for females are not as clearcut, although in more than half of the countries, divorced females also have the highest ratios. Although several studies based on data for a single country found that divorced persons suffer excess mortality relative to single

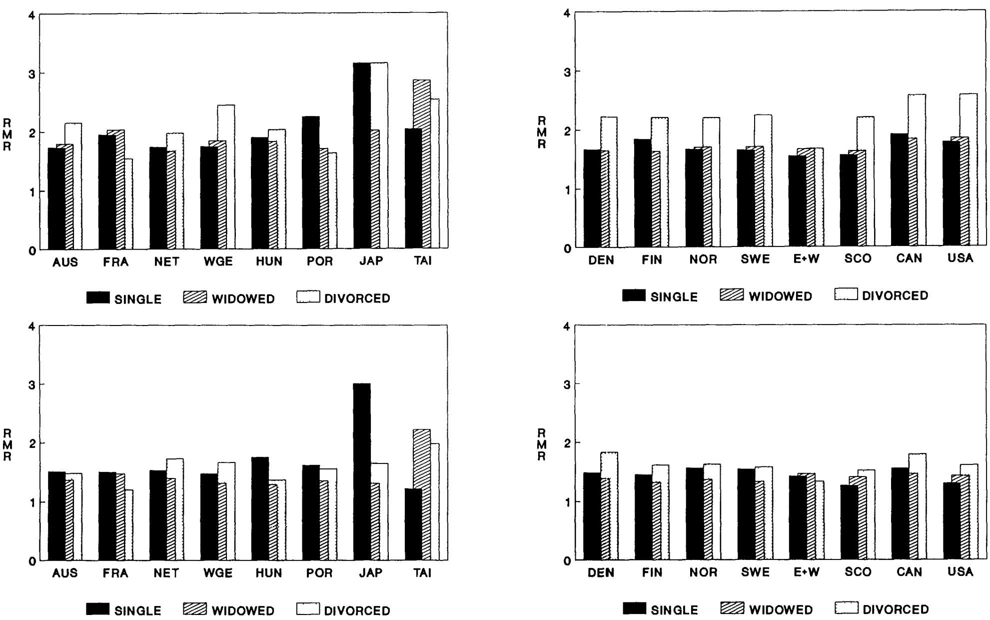

Figure 1. Relative Mortality Ratios (relative to married persons) by Country and Gender (top, males; bottom, females)

and widowed persons (e.g., Fox & Goldblatt 1982; Gove 1973; Klebba 1970; Koskenvuo, Kaprio, Kesaniemi, & Sarna 1980), a previous cross-country comparison (Kisker & Goldman 1987) did not uncover any systematic differences among the three unmarried groups.[10]

There are several outliers apparent from Figure 1. The most obvious one is the extremely high ratio (3.0) for single females in Japan, which is 70% higher than the next highest ratio for single women. Interestingly, both single and divorced men in Japan also have unusually high ratios. Kisker and Goldman (1987) discovered a similar phenomenon for single males and females in Japan. One question that we will address later in this article is whether these high mortality ratios are due to the very low proportions of people who remain single and who get divorced in Japan and hence to the possible strong selection factors that may be operating here. An alternative plausible explanation is that unmarried persons in this society suffer a greater amount of stress or are at a greater social and economic disadvantage relative to married persons than in Western societies.

The data in Figure 1 also suggest that divorced males in North America suffer greater excess mortality than their counterparts in European countries, although the differences are much smaller than for the Far Eastern countries. No clear regional patterns within the European countries emerge from these estimates.

Age-Specific Ratios

We use the second model, which includes an interaction term between age group and marital status, to describe the age patterns of the mortality ratios for the three unmarried states. This interaction term is by far the most important of any interaction between covariates. The resulting estimates of relative mortality ratios by age group[11] are shown graphically for each country in Figures 2 (males) and 3 (females) and indicate that the three unmarried states are characterized by quite different age patterns relative to married persons.

The figures reveal some interesting similarities across countries. Most striking are the extremely high ratios—typically greater than 5 and sometimes as high as 15—of young

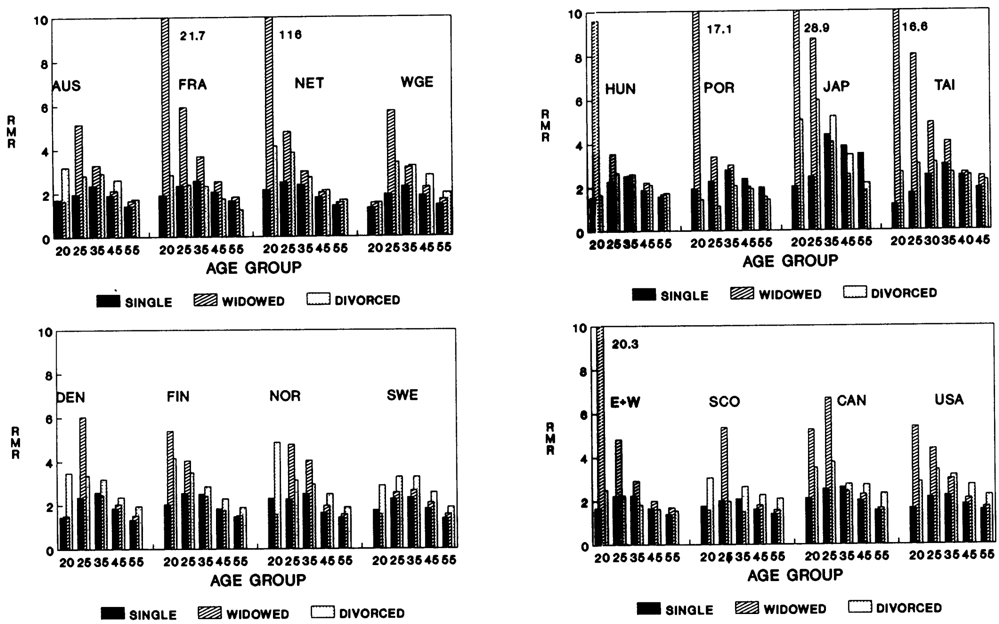

Figure 2. Age-Specific Relative Mortality Ratios by Country, Males

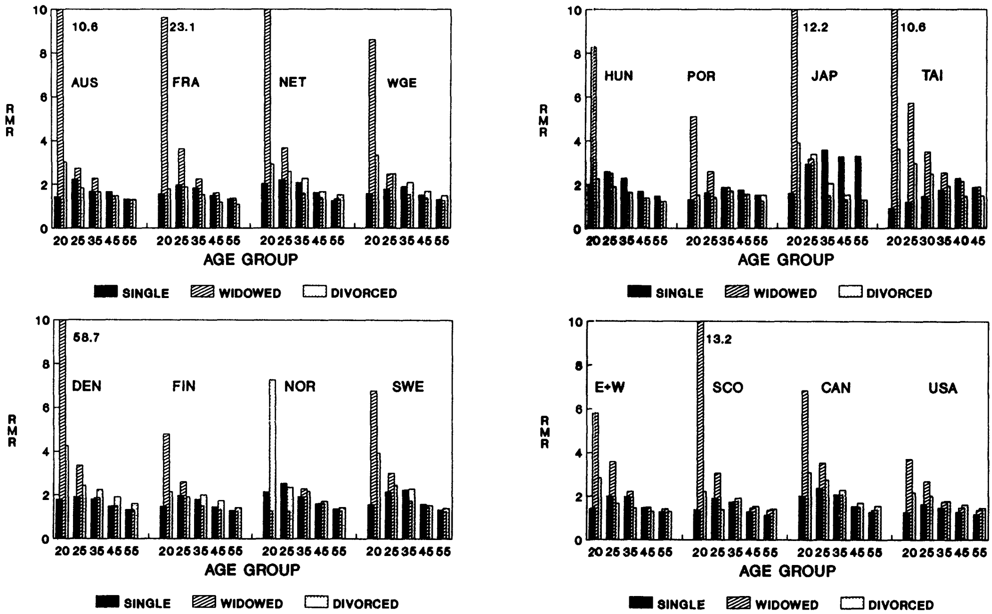

Figure 3. Age-Specific Relative Mortality Ratios by Country, Females

widows and widowers. Although the exposed population of young widows is very small in some countries, the similarity of this pattern across countries suggests a real phenomenon, rather than sampling error. Given the mixed evidence for bereavement effects (e.g., Cox & Ford 1964; Helsing & Szklo 1981; Kaprio, Koskenvuo, & Rita 1987; MacMahon & Pugh 1965), it is hard to imagine that bereavement could be an explanation for these age patterns. Rather, the results suggest that whatever caused the death of one spouse also placed the surviving partner at a higher risk of dying. In some cases, this could have been an accident that injured both persons but led one to die before the other. In other circumstances, it could have been an infectious disease (or another aspect of an unhealthy marital environment, such as drugs or alcohol) that debilitated the surviving spouse.

Although young divorcees are not at as high a risk of dying as young widows and widowers, they too are characterized by high mortality ratios relative to other age groups and to single persons. For example, in about half of the countries, the mortality ratios of divorced men and women aged 20–24 are greater than 3. For both the divorced and the widowed states, the mortality ratios generally decline with increasing age. That is, the differences in the risk of mortality (in relative terms) between formerly married persons and married persons is smallest for the oldest age groups, although even these ratios are above 1.5.

In contrast to formerly married persons, single persons are characterized by ratios that generally reach peak values in the age groups of 25–34 and 35–44; excess mortality of single persons is smallest in the youngest age group and the oldest. The former factor probably occurs because many persons—in some countries, the majority—are still single in their early twenties; selection factors, which draw healthier persons into the married state, would not be apparent by this young an age. By the age of 35, on the other hand, the single state is composed of individuals who, for a variety of reasons, are unlikely ever to marry. The declining ratios in the older age groups are more difficult to interpret. One could argue that this reflects an adjustment process whereby the frailest single persons die young and the healthier, or those most able to adjust to the single state, constitute an increasing proportion of the single group above the age of 40. One could also maintain that this pattern is a consequence of the differentially beneficial effects of the married state for persons of different ages, that is, that being married is especially beneficial in the thirties and forties.

Time Trends in Mortality Ratios

The two models considered so far make no allowance for possible interactions between time and marital status; that is, the resulting estimates are "average" effects over the years included in the analysis. Comparisons of model 2 with models that also include some form of interaction between year and marital status indicate a significant improvement of goodness of fit in all countries. As noted earlier, after considerable exploration of a variety of such models, we settled on model 3, in which year is treated as a factor for modeling the main effect and as a continuous variable, with both a linear and quadratic term, for modeling its interaction with marital state.

To compare the resulting trends in the excess mortality of the unmarried population across the 16 countries, each of which had different calendar years included in the analysis, we introduced several modifications to the calculation. Rather than examine the estimated RMR in each year, we consider the ratio of the estimated RMR in a given year to that in a standard year (usually 1955; see the Appendix for details). The resulting estimates of the relative RMRs, derived from model 3, are shown in Figures 4 and 5 for single men and single women. The corresponding estimates for widowed and divorced persons are described below, but the figures are not included here.

Several interesting trends emerge. First, single persons in most countries experienced a modest increase of about 10–20% in the RMR. Japan is clearly distinct in terms of its time trend in the mortality ratio: for both single males and females, the RMR declined by about 40% between 1955 and 1980. Although these data suggest that the extremely high ratios shown in Figure 1 for single persons in Japan might be primarily due to death rates

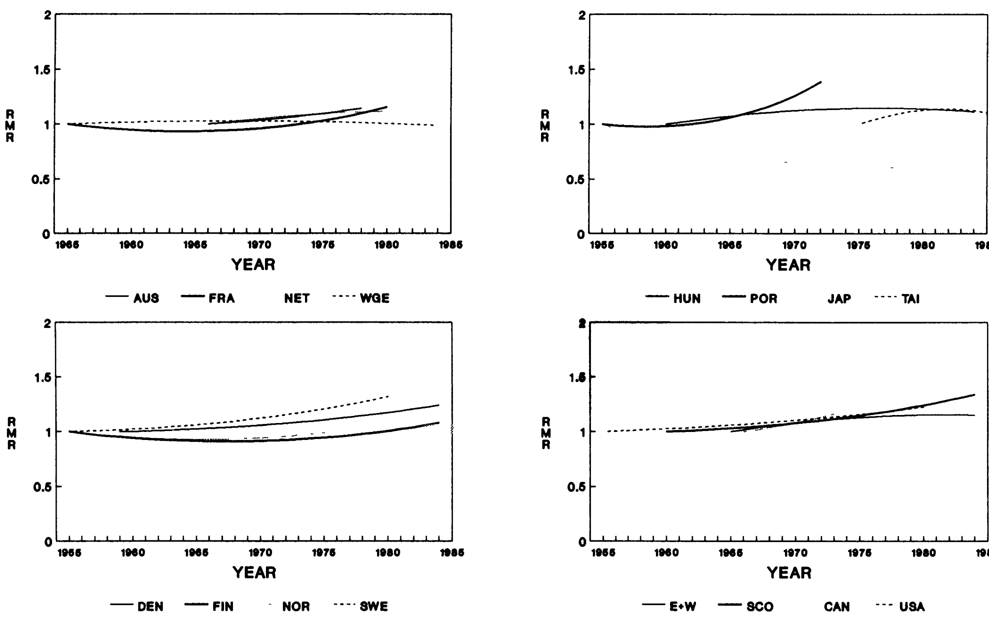

Figure 4. Trend in Relative Mortality Ratios (RMR in given year relative to RMR in starting year) by Country for Single Males

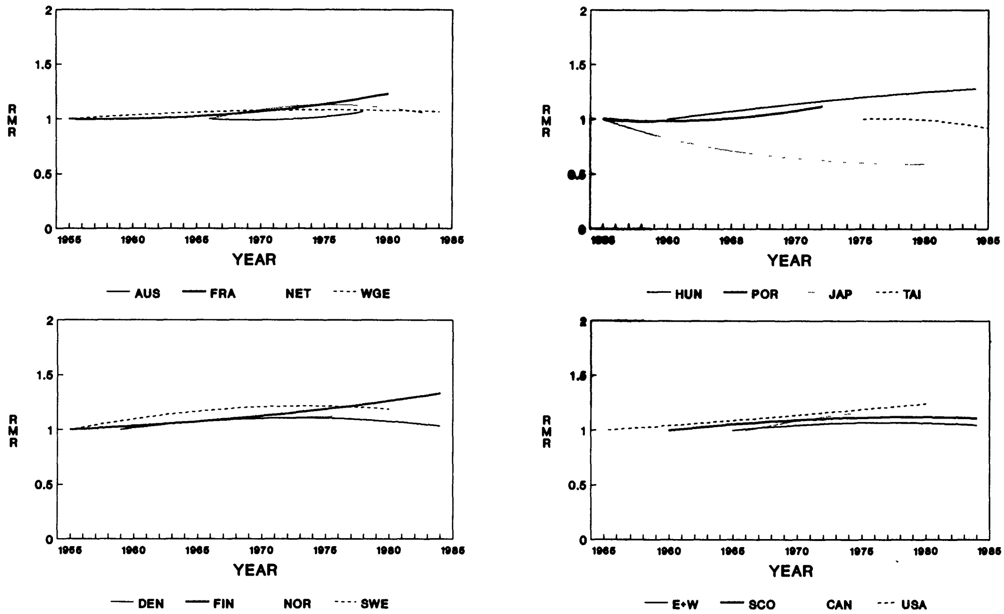

Figure 5. Trend in Relative Mortality Ratios (RMR in given year relative to RMR in starting year) by Country for Single Females

in the 1950s and 1960s, comparisons not shown here indicate that the RMR for single persons in Japan remains above those for other countries even in 1980.

The ratios for widowed persons also indicate modest increases in the RMR over the period under study, for most countries. Particularly large increases (e.g., above 30%) occurred among widowed males in Hungary, Portugal, Denmark, Finland, and Canada and among widowed females in West Germany and Hungary.

The trends in the RMR for divorced persons show the largest variation across countries. The majority of countries show some increase in the ratio: many of these increases are small in magnitude, but there are extremely large increases for males in several countries—85% in France, 55% in Japan, and 67% in Hungary. At the same time, there are some notable decreases in the ratios, such as those for Canadian and Portuguese males and females.

In summary, for the majority of countries included in this analysis, as well as for both genders, the excess mortality of each unmarried state (relative to married persons) has increased over the past two to three decades. This increase could be due to one of three factors: larger health benefits of marriage during a time of rapid technological development, a greater degree of selection in terms of the type of person who remains unmarried or whose marriage is terminated, or the inappropriateness of the RMR as a measure of differential mortality by marital status.

The first explanation derives from the notion that increases in stress that have accompanied rapid modernization have affected unmarried persons more than married persons (who at least have a partner with whom to share the stress). In addition, married persons may avail themselves more readily (because of financial or social reasons) of improvements in medical technology. The second interpretation is based on the fact that the proportion of persons who are single has generally decreased over the period under study; thus one could argue that the degree of selection has increased, with the single state containing higher proportions of unhealthy persons over time. This argument is not appropriate for divorced persons, since the relative size of the divorced group has generally increased in most countries.

The third explanation arises from the use of a ratio to measure mortality differentials: the RMR can increase while death rates for both the unmarried and married groups decline (a pattern that in fact occurs in most of the countries) as well as when the absolute difference between the two rates narrows.

Mortality and the Relative Size of the Marital Group

Two previous studies (Kisker & Goldman 1987; Livi-Bacci 1985) indicated that the excess mortality of single persons (relative to married persons) is related to the relative size of the single group. Specifically, these analysts demonstrated that in populations in which almost all persons marry, single persons are characterized by mortality rates substantially higher than married persons. By contrast, when a substantial fraction of the population is single, the excess mortality of single persons is considerably lower. The authors demonstrated this relation by calculating simple correlations between the percentage of persons in an age group who are single and the excess mortality of the single (i.e., the relative mortality ratio in that age group). Large negative correlations (usually below $-.5$ and often as low as $-.8$) resulted from a comparison across 26 developed countries around 1980 and from comparisons across time periods for males and females in France and Japan.[12] These correlations were substantial in virtually all age groups between 20–24 and 45–49 (Kisker & Goldman 1987).

An important limitation of these earlier analyses is the absence of any control for period effects. For example, it is possible that the relationship between the RMR and the percentage of single persons was driven by the temporal pattern of increasing ratios over the past several decades. Thus we use model 4, which incorporates all of the important interactions among age, year, and marital state in addition to a term involving the size of the marital group, to explore the importance of size in determining the mortality rates among unmarried persons. Since we speculate that size may be important for divorced persons as well as for single persons—that is, during periods in which very few people get divorced, those who do may be select with regard to health characteristics—we examine the resulting coefficients for both single and divorced persons of each gender.[13]

In model 4, the size of the marital group has been measured as the logarithm of the percentage of persons (of the relevant age group and gender) in that state, a variable denoted by R.[14] Since we are modeling the logarithm of the mortality rate, the resulting coefficients (e_k) can be interpreted as elasticities; that is, e_k measures the percentage change in the mortality rate for a 1% change in the relative size of the marital group.

The estimated coefficients for single and divorced persons in each country are shown in Table 4. In general, the results are consistent with the existence of a selection effect, particularly for single persons. In all but 4 of the 32 country–gender combinations, the resulting coefficient is negative, as expected: we speculated that, all other factors being held constant, the larger the relative size of the group, the lower would be the risk of dying. For the 28 negative cases, the average size of the coefficient is $-.34$ ($-.30$ among males and $-.37$ among females).[15] Thus a 10% increase in the percentage of persons who are single is associated with about a 3.4% decline (on average) in the death rate for single persons (in that age group).

There is also some support for the selection argument for divorced persons. Most of the coefficients are negative, and the average magnitude of the negative coefficients ($-.36$) is about the same as that for the single state. The coefficients are not, however, statistically significant in more than half of the cases.

Are any regional patterns apparent from the coefficients presented in Table 4? The most obvious is the lack of support for the selection hypothesis in the Scandinavian countries. With only two exceptions, coefficients for the sizes of both the single and divorced groups (for men and women) in Denmark, Sweden, and Norway are not significant.[16] A similar finding resulted from the earlier analysis of Swedish data by Kisker and Goldman (1987),

Table 4. Coefficients of R for Single and Divorced Persons

	Single		Divorced	
Country	Coefficient	t	Coefficient	t
Males				
Austria	−.46	−1.66*	−.31	−1.23
Canada	−.34	−1.39	−.08	−.36
Denmark	−.34	−2.11*	−.29	−1.40
England and Wales	−.31	−2.87*	−.28	−2.02*
Finland	.19	1.21	−.26	−1.44
France	−.47	−7.10*	−.57	−15.95*
Hungary	.29	2.50*	.20	.67
Japan	−.13	−3.74*	−.22	−3.61*
Netherlands	−.47	−2.61*	−.22	−1.08
Norway	−.04	−.16	−.21	−.50
Portugal	−.40	−2.85*	−.50	−.97
Scotland	−.14	−.81	−1.22	−2.06*
Sweden	−.13	−.80	.28	1.36
Taiwan	−.50	−6.95*	−.53	−3.61*
U.S.A.	−.19	−7.55*	−.30	−15.69*
West Germany	.09	2.30*	−.46	−12.39*
Females				
Austria	−1.11	−2.18*	.23	.32
Canada	−.41	−1.05	.03	.06
Denmark	−.14	−1.13	−.45	−1.45
England and Wales	−.25	−2.57*	−.35	−2.18*
Finland	−.07	−.32	−.20	−.55
France	.26	2.06*	−.49	−7.92*
Hungary	−.22	−1.43	−.41	−.69
Japan	−.24	−9.47*	−.29	−7.79*
Netherlands	−.28	−1.72*	−.23	−1.12
Norway	−.01	−.10	−.14	−.25
Portugal	−.81	−3.22*	.71	1.41
Scotland	−.38	−2.64*	.55	1.13
Sweden	−.35	−3.53*	−.20	−.61
Taiwan	−.60	−3.72*	−.19	−.47
U.S.A.	−.39	−11.74*	−.52	−9.08*
West Germany	−.28	−6.30*	−.40	−5.53*

Note: Coefficients are from model 4. R denotes the logarithm of the percentage in the marital status.
* $p < .05$ (one-tailed test).

in which only a weak correlation existed between the RMR of single persons and the percentage single. The lack of importance of the size of the group in these countries is not altogether surprising given the high level of informal cohabitation. By contrast, the coefficients for the size of the divorced and single groups are always significant (and negative) for the United States, England and Wales, and Japan.

One anomaly apparent from Table 4 is the positive coefficient for single females in France. This is in striking contrast to the large negative correlations found by Kisker and Goldman (1987) for French females during the period 1886–1981. Although some of the differences may be due to the much longer period used in the earlier analysis, our findings suggest that part of the discrepancy is due to the inclusion of period interactions in our model. Over the time period included in our analysis, the RMR of single persons increased and the relative size of the single group declined (in most age groups). This apparent negative correlation is "captured" as a period effect in the model; if we omit the interaction terms between year and marital state, the coefficient e_k for single females would indeed be negative and significant. This example highlights the importance of controlling for temporal patterns in an examination of the effect of group size on mortality.

We speculated earlier that the high RMR for single persons in Japan and the notable decline in this RMR over the last few decades might be a consequence of the small relative size of the single group in Japan, compared with other countries, as well as of an increase in this relative size over the past two to three decades. Is this the case? By and large, the answer appears to be "no" in both instances. We reached this conclusion by using the estimated coefficients of e_k for Japan in model 4 to simulate the value of the RMR in Japan for various relative sizes of the single population. For example, for the cohort of Japanese females aged 35–44 in 1955, the reported percentage single equals 3.1, in contrast to the average value for the corresponding European cohorts of 13.5%. Increasing the size of the single group from 3.1% to 13.5% (with a corresponding decline in the size of the married population) would lower the estimated RMR for these women from 3.9 to 3.3. A similar calculation for the cohort of women aged 35–44 in 1980 would lower the estimated RMR from 3.4 to 2.9. Although the size of the single population has a clear effect on the relative mortality ratio, even the reduced values of the RMR are greater than those experienced in any other country in this analysis. Furthermore, simulated values based on several models indicate that if the relative size of the single group in Japan had been constant over the period 1955–1980, the decline in the RMR would have been smaller than that which actually occurred, but it still would have been substantial[17] (and also larger than that experienced in any other country).

Conclusions

Several consistent and striking patterns emerged from an exploration of marital status differentials in mortality across 16 developed countries. First, in all countries, the excess mortality of unmarried men (relative to married men) greatly exceeds that of unmarried women. Second, in most countries, divorced men have the highest death rates among the three unmarried groups. The corresponding finding for females is true in more than half of the countries. Third, age-specific effects indicate that widowed and divorced persons in their twenties and early thirties experience the highest mortality risks, risks that are sometimes 10 times as high as those for married persons of the same age. By contrast, among single persons, the highest risks are usually associated with men and women between the late twenties and early forties. Fourth, in the majority of countries, the relative mortality ratios of the three unmarried groups have increased over the past two to three decades. Japan emerges as the most apparent exception to these generalizations.

The results also indicate that the size of the marital group is related to mortality in most countries. These findings suggest that selection operates with regard to both single and divorced persons: The smaller the proportion of persons who never marry, the more likely that these people are characterized by high risks of mortality; similarly, the smaller the proportion of divorced persons, the more likely that the small group who obtained a divorce (and did not remarry) are select with regard to a variety of risk factors that increase their chances of dying. Even these findings, however, are not conclusive evidence for selection. It is possible that persons who belong to relatively small groups may experience greater stress because of their rarity and possible social isolation.

In general, the consistency of findings across a large number of countries suggests that similar processes are responsible for producing the higher death rates of unmarried persons in diverse social settings. Such results strengthen previous speculations about the importance of marriage in maintaining health and reducing the risk of dying and the increased stresses associated with both the single and the formerly married states. The next step is to obtain a better understanding of the ways in which marriage is beneficial to health and nonmarriage is hazardous. For example, why is it that divorced men have especially high risks of dying relative to married men? Why have these risks been increasing over time, even though divorce has become more common? Clearly, selection is not the entire answer. Demographers need to exploit both large-scale prospective studies, in which given individuals experience changes of marital state, and cause of death data, in which the excess mortality of the unmarried can be identified with particular disorders, to learn more about the causal mechanisms between marital status and mortality.

Notes

[1] Data sources: England and Wales (Registrar General 1945–1971; Office of Population Censuses and Surveys 1983). France (Levy 1974). Japan (Statistics and Information Dept. 1955–1980; Statistics Bureau 1955–1980). Sweden (Statistika Centralbyran 1955–1980). Taiwan (Ministry of the Interior 1976, 1986). United States (Grove & Hetzel 1968). Other countries (United Nations, *Demographic Yearbook*, selected years between 1958 and 1985).

[2] In the case of Taiwan, data were reported for every calendar year since 1975; we selected every third year to be included in the analysis.

[3] The exception is Taiwan, for which the data are classified into six age groups: 20–24, 25–29, 30–34, 35–39, 40–44, and 45–49. We do not consider age groups beyond 55–64 because in about half of the countries, the upper age category is 65 and above for at least some of the years included in this analysis.

[4] For these years, only data for the single and married populations were included in the analysis.

[5] All of these models are hierarchical so that inclusion of an interaction term such as $A \times M$ (age and marital status) implies that the main effects of A and M are also incorporated into the model. The expression $(A + M) \times (Y + Y^2)$ denotes an interaction between age and year (with linear and quadratic terms for year) and between marital state and year (also with both linear and quadratic terms for year). Although year is modeled as a series of dummy variables in terms of the main effect P, it is modeled as a continuous variable Y (which is measured as the difference between the given calendar year and the first year in the data set, for the particular country) in the interaction terms in the third and fourth models. In the fourth model, R denotes the logarithm of the percentage in the marital state.

[6] GLIM stands for Generalized Linear Interactive Modeling, a program for fitting a family of models that includes the Poisson log-linear model as an example. The procedures for fitting these models are described in Baker and Nelder (1978) and Payne (1987).

[7] We use the likelihood ratio goodness-of-fit statistic, which can be expressed as $\chi^2 = 2 \Sigma D_{ijk} \times \log(D_{ijk}/\hat{D}_{ijk})$, where D_{ijk} is the observed death count and $\hat{D}_{ijk}$ is the fitted death count for age group i, year j, and marital status k. This statistic can be reexpressed as $\chi^2 = 2N \Sigma W_{ijk} \mu_{ijk} \log(\mu_{ijk}/\hat{\mu}_{ijk})$, where μ_{ijk} and $\hat{\mu}_{ijk}$ denote the observed and fitted death rates in cell i, j, k and W_{ijk} represents the fraction of the total exposure N in cell i, j, k. From this reexpression, one can directly see how the chi-squared statistic is a function of the total exposure N.

[8] From the equation for the additive model, we see that $\exp(\hat{\eta})$ and $\exp(\hat{\eta} + \hat{\gamma}_2)$ are the estimated death rates for married and single persons (in the first age group and first year), respectively. Thus it follows that $\exp(\hat{\gamma}_2)$ is the estimated RMR of single persons. Note that the estimated RMR would be the same for any age group and year.

[9] Undoubtedly, use of the RMR as a measure of mortality differentials has serious limitations because changes in the ratio are complicated to interpret: for example, changes over time in the relative mortality ratio of the single population could be due to temporal changes in either the death rates of the single or those of the married. Unfortunately, there is no ideal measure of mortality differentials (e.g., see Sheps 1961).

[10] Since Kisker and Goldman's (1987) study was not based on a multivariate model, they concluded only that the average excesses of mortality (across countries) were similar among single, widowed, and divorced persons.

[11] For example, for the first two age groups (20–24 and 25–34), the estimated RMRs of single persons would be equal to $\exp(\hat{\gamma}_2)$ and $\exp[\hat{\gamma}_2 + (\hat{\alpha}\gamma)_{22}]$, respectively.

[12] In the cross-cultural comparison the correlation coefficients were much larger for females than males, but within both France and Japan the coefficients were large for both sexes. Much smaller coefficients resulted from comparisons across periods within England and Wales and Sweden (Kisker & Goldman 1987).

[13] We do not examine the coefficients for widowed persons because the relationship between the magnitude of the death rate and the size of the widowed group is a very complex one, operating in both directions. In particular, the higher the death rates, the more likely persons are to become widowed and hence the larger the size of the widowed group.

[14] Overall, the logarithm of the percentage performed better, across countries, than the untransformed percentage in the group or than a threshold variable that was equal to the percentage in the state up to a certain value (e.g., 10%) and then simply leveled off at this value. The latter transformation was based on the assumption that as long as the relative size of the group was very small, size would be related to the mortality risk; however, once the relative size reached a large enough value, the potential selection effect would be undetectable.

[15] As shown in the table, the coefficients are statistically significant in about two-thirds of the cases.

[16] The only other countries for which this is generally true are Finland, Hungary, and Canada.

[17] For example, if the relative size of the single group is held constant at its value for 1955, the decline in the RMR for singles (for all age groups combined) would be 33% for males and 27% for females, in comparison with the estimated values of 39% for males and 41% for females.

Appendix

The time trends in the relative mortality ratios were derived on the basis of model 3:

$$\log \mu_{ijk} = \eta + \alpha_i + \beta_j + \gamma_k + (\alpha\gamma)_{ik} + a_k Y + b_k Y^2.$$

In this model, α_i denotes the effect of the ith age group, β_j the effect of the jth year (with year being treating as a categorical or dummy variable), γ_k the effect of the kth marital group (with $k = 1, \ldots 4$, denoting the effects of being married, single, widowed, and divorced, respectively), $(\alpha\gamma)_{ik}$ an interaction effect between age and marital status, Y the specified year in continuous notation (measured as the actual calendar year minus the starting year y_0—the first year for which data are included in this study), and a_k and b_k interaction effects between year and marital status k.

Consider, as an example, the estimation of the RMR for single persons (relative to married persons) in the third age group for year Y. On the basis of model 3, the estimated RMR is as follows:

$$\hat{\text{RMR}} = \exp[\gamma_2 + (\alpha\gamma)_{32} + a_2 Y + b_2 Y^2].$$

Note that this expression contains many fewer terms than in the original model because $\gamma_1 = (\alpha\gamma)_{12} = a_1 = b_1 = 0$ (i.e., these are the relevant reference cells, which are set equal to zero) and because the main effect for period (β_j) cancels out from the numerator and denominator of the RMR.

Now suppose that we consider the estimated RMR for calendar year $Y + y_0$ relative to the corresponding one (i.e., for single persons in the same age group) for the starting year y_0. Based on the preceding formula, we have

$$\frac{\hat{\text{RMR}}_{Y+y_0}}{\hat{\text{RMR}}_{y_0}} = \frac{\exp[\gamma_2 + (\alpha\gamma)_{32} + a_2 Y + b_2 Y^2]}{\exp[\gamma_2 + (\alpha\gamma)_{32}]} = \exp[a_2 Y + b_2 Y^2]. \qquad \text{(A.1)}$$

For example, $\hat{\text{RMR}}_{1960}/\hat{\text{RMR}}_{1955} = \exp[5a_2 + 25b_2]$. Equation (A.1), with modifications for different age groups and marital states, was used to generate the trends shown in Figures 4 and 5 for countries in which y_0 was 1955 or later. For countries in which y_0 was 1940 or 1950, the estimated RMR was calculated relative to 1955 so that results could be more easily compared with those from other countries. In these cases the values shown in the figures were determined from the following formula (still based on single persons):

$$\frac{\hat{\text{RMR}}_{Y+1955}}{\hat{\text{RMR}}_{1955}} = \exp\{a_2[Y - (1955 - y_0)] + b_2[Y^2 - (1955 - y_0)^2]\}, \qquad \text{(A.2)}$$

which reduces to formula (A.1) when $y_0 = 1955$. The estimated RMRs for divorced and widowed persons are analogous to those for single persons, with the coefficients a_2 and b_2 replaced by a_3 and b_3, and a_4 and b_4, respectively.

Acknowledgments

This research was funded by National Institute of Child Health and Human Development Grant R01 HD22414. We would like to thank Germán Rodríguez and Ellen Kisker for their assistance throughout the analysis and James Trussell and Lorenzo Moreno for their comments on an earlier version of this article.

[*Received April 1989. Revised August 1989.*]

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