From Death to Birth

Mortality Decline and Reproductive Change

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Infant Mortality and the Fertility Transition: Macro Evidence from Europe and New Findings from Prussia

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INTRODUCTION

Most attempts to understand secular fertility decline include some allusion to the European experience. It is generally thought that little or no relationship existed between fertility decline and infant mortality decline in Europe, or that the findings from relevant studies are inconsistent. We believe that these common perceptions are mistaken. When more attention is given to the varying methods of analyses, a more consistent picture emerges. We argue that it is particularly important to keep in mind whether studies are bivariate or multivariate; whether studies estimate cross-sectional relations between levels of fertility and of infant mortality, or instead focus on the relation of changes in these variables; and whether studies take into account the possibility that causality flows in both directions—from fertility to mortality as well as from mortality to fertility.

We estimate both the impact of infant mortality on fertility and the impact of fertility on infant mortality, using aggregate data from Prussia from 1875 to 1910 and fixed effects models with instrumental variables. This is followed by an extensive review of previous research on fertility and infant mortality within the historical European context.[1] By comparing our findings for Prussia with earlier research looking at both level and change effects, we find considerable evidence

[1]Our review of earlier research is restricted to those studies using aggregate data. There is a body of literature on fertility and infant mortality that uses micro-level data, see for example Knodel (1988). However, it is beyond the scope of this chapter to survey such studies.

for a positive association between the fertility level and the infant mortality level, as well as a positive association between fertility change and infant mortality change.

The Long Term

It is clear that in the very long run, in closed populations, fertility and mortality are linked because of the finiteness of the resource base, which implies that the average rate of natural increase n must not exceed zero more than slightly. This is an abstract argument. The historical reality has been that rapid natural increase sustained over long periods (say an average rate of natural increase greater than 0.02 over a period of more than two centuries) has not been observed except in frontier regions[2] such as North America. Much more typically, large populations appear to have had rates of natural increase of less than 1 percent per year until the Industrial Revolution, and usually with a strong positive statistical association between fertility and mortality in the cross section. Such a long-term positive association of fertility and mortality, and a limit to average rates of natural increase, can be explained in at least two ways. First, positive growth rates mean increasing population size and density, which under preindustrial conditions typically meant declining living standards. These in turn caused mortality to rise, or fertility to fall, and therefore growth rates to return toward zero. This is the Malthusian theory of population equilibration through negative feedback (Lee, 1987). Of course, emigration was another possible outcome, and technological progress or international trade might intervene between population growth and declining living standards. Second, it is sometimes argued that the sociocultural institutions governing fertility evolved in the context of some average mortality regime so as roughly to balance fertility and mortality on average, leaving rates of natural increase close to zero. In this version there is no feedback from population size to the vital rates; rather, growth rates themselves tend to have average levels not far above zero. Again, migration, technological progress, and international trade might play a role (Smith, 1977; Yule, 1906).

The constraint on average growth rates, and hence on fertility and mortality, implied by these theories and the positive association of fertility and mortality observed in the historical record, may also help to explain the long-run shape of the demographic transition (Lee and Bulatao, 1983, for example). Demeny (1968:502) gave a classic description of the transition: "In traditional societies, fertility and mortality are high. In modern societies, fertility and mortality are low. In between, there is demographic transition." In fact, in those national

[2]By frontier regions we mean from the point of view of agricultural populations, that is, not from the point of view of hunter and gatherer populations who may have occupied the area at relatively low density before the arrival of agriculturists.

populations that have "completed" the transition, fertility has dropped so low that growth rates may turn very substantially negative. Nonetheless, it is difficult to escape the conclusion that, in some vague and unspecified way, and despite all the accompanying structural changes in the economies and societies, the very long-term decline in fertility is ultimately due to a very long-term decline in mortality, or the two are interlinked. In fact, some theories link both declines to the same set of parental decisions concerning investment in children.

These very long-run relations, both theoretical and empirical, are based on some sort of slow-acting feedback operating through the macro-economic or macro-societal level. The posited mechanisms might be expected to operate over the course of a century or more, but not over the course of decades. For this reason, they are of little relevance for questions about the policy-relevant time frame of adjustment over the medium range of, say, 5-30 years.

The Short Term

Although the long-run historical relation of fertility and mortality is doubtless positive, it is equally true that the empirical relationship over short-run fluctuations has been consistently negative in historical populations. This has been established by a large number of studies of time series of births and deaths, once the long-term trends in the data have been statistically removed. It is easy to think of reasons to expect either a positive or a negative association of the two vital rates. For a positive relationship, note that high mortality will break many marriages, particularly those of older couples, and that the subsequent remarriages of widows and widowers might result in higher fertility than if the marriages had been unbroken. Furthermore, higher mortality would free land holdings and create other economic opportunities permitting new marriages that would have high fertility. In existing unions, high infant and child mortality would interrupt breastfeeding, eliminating its contraceptive effects, and therefore lead to earlier conceptions and a temporary increase in fertility. Reconstitution studies have often demonstrated this lactation interruption effect. On the behavioral side, if we are not dealing with a natural fertility regime, we might expect couples who have experienced the loss of a child to attempt to replace it with another birth sooner than they normally would have, or by having one more birth than originally intended. However, many historical demographers dispute that this actually occurs to any appreciable degree, except in special subpopulations (Knodel, 1978).

For a negative relationship, note that many factors that tended to raise mortality would also tend to reduce fertility. For example, low real incomes apparently had this effect, as did unusually hot summer months or unusually cold winter months (see Lee, 1981; Galloway, 1986, 1988, 1994). The variation of such factors in the short term would have led to a negative bivariate association of fertility and mortality, but if observable, they can be netted out in multivariate studies. Perhaps more important were unobservable influences, of which ill health dominated. Fluctuations in morbidity both raised mortality and reduced fertility, leading to a strong negative association of short-term fluctuations in fertility and mortality, even after controlling for observed fluctuations in real incomes or grain prices and temperature. The estimated, strongly negative association of short-run variations in fertility and mortality is not very informative about structural or causal influences of mortality on fertility or the reverse. The many short-run studies of the relation of fertility to mortality in historical populations will therefore not answer the question before us.

The Medium Term

For policy makers, the most relevant time frame for fertility and infant mortality interactions is probably the medium term, say 5-30 years. Assuming a couple has some notion of a desired number of surviving offspring, infant and child mortality should be positively associated with the number of births. A couple can assume that some unknown number of offspring will die, and then stop when they think they have enough children (often called hoarding behavior). Or the couple can wait to see if the last child born survives past a certain age. If the child dies, the couple can then engage in replacement reproductive behavior. Both strategies are types of "inventory control" (Preston, 1978:10) leading to some desired number of surviving offspring. Within either strategy, the number of births should decline as infant mortality declines.

There are also ways in which declining mortality might alter the desired number of surviving children. It reduces the costs of achieving a given target number, and therefore might raise the target by increasing discretionary income. Alternatively, declining mortality might raise the rate of return on investments in children, which could lead to a substitution of quality for quantity, and a reduced target. Here, however, we concentrate on the fixed target scenario.

Once infant mortality begins to decline, it might take some time for couples to perceive the effect of infant mortality on child survivorship, which would ultimately lead to changes in fertility. It is also possible, however, that the effect could be almost immediate, as couples hear about and read about mortality decline.[3]

[3]It would be difficult to test for very short lags because censuses (from which we derive most of our independent variables) are nearly always at least 5 years apart, and because it is very likely that the level of infant mortality rates at year t will be highly correlated with the level of infant mortality rates at years $t-2$, $t-3$, or $t-4$. Using Prussian data and the model shown in Appendix Table 6A-2, we added the variable infant mortality lagged 5 years and found that, in the fixed effects model, which estimates changes, the regression estimate on infant mortality with no lag was 0.267 whereas the regression estimate on infant mortality lagged 5 years was -0.062, suggesting that the lagged variable was relatively unimportant.

Elevated infant mortality tends to shorten the birth interval because the death of an infant curtails lactation amenorrhea along with its contraceptive effects. An increase in infant mortality will cause an increase in fertility, *ceteris paribus*, although few children may ultimately survive, of course. This short-term phenomenon can persist over time, becoming an important factor over both the medium and the long term.

Fertility variations, whether deliberate or accidental, can also affect infant and child mortality as demonstrated by a host of contemporary studies. There is good reason then to expect that exogenous increases in infant and child mortality caused increases in marital fertility and that exogenous increases in fertility caused increases in infant and child mortality. (However, this micro-level reasoning about motives and relations does not translate exactly to the macro level because of expected nonlinearities in the relationships.)

Infant and Child Mortality

Matthiessen and McCann (1978) provide a useful overview of the findings of historical studies of macro-level data, with an emphasis on the early results of the European Fertility Project. They are particularly critical of the use of infant mortality as the explanatory mortality index, because they find that in practice other more appropriate measures, such as mortality of children age 0-15, began to decline earlier than did infant mortality, so that the European Fertility Project's studies of timing, for example, are of little value. When they reexamine the timing of the fertility transition in relation to ${}_{15}q_0$, they find that mortality decline almost always preceded fertility decline. We believe that it is very difficult to estimate the onset of secular ${}_{15}q_0$ decline. In general, we suggest that there is often no clear point at which one can categorically state that mortality or fertility has begun to decline, a suggestion with which Matthiessen and McCann (1978:52) clearly agree. Concerning the onset of infant mortality decline, van de Walle is appropriately cautious, noting that "in most instances we are left ignorant of past trends: the data do not allow us to go back in time and the existence of an earlier decline cannot be ascertained" (1986:213).

It might be useful to address the issue of infant (under age 1) versus child (age 1-9) mortality in multivariate analyses of secular fertility decline. When a husband or wife thinks about procreation in terms of offspring survivorship, he or she considers both infant and child mortality (Matthiesson and McCaan, 1978:52). Although infant mortality rates can generally be found in most historical registration material, the more detailed measures of child mortality are often unavailable. However, in a high infant mortality regime, the bulk of infant and child deaths will be infant deaths, and infant mortality should be very highly correlated with infant and child mortality combined. Using 1890-1891 male mortality data for the 36 provinces (Regierungsbezirke) of Prussia, we find that the correlation r between ${}_1q_0$ and ${}_5q_0$ is 0.96, between ${}_1q_0$ and ${}_{10}q_0$ is 0.96, and between ${}_1q_0$ and

${}_{15}q_0$ is 0.95 (Königliches Statistisches Bureau, 1904:135-147). The range of ${}_1q_0$ in the provinces is 109-273. A similar analysis of the 15 largest cities in Prussia from the same source reveals respective r's of 0.98, 0.97, and 0.91 with a range of ${}_1q_0$ of 170 to 326. Plots of each of the six graphs reveal essentially a straight line with no outliers. It seems likely that the infant mortality rate is an adequate proxy for infant and child mortality when using aggregate data in high infant mortality populations.

Although it is difficult to say much about the timing or onset of secular infant and child mortality decline, we can examine their relative speed. It is clear from Figure 6-1 that ${}_1q_0$ and ${}_{15}q_0$ generally declined at about the same rate over the decades from the 1870s to around 1925. The quality of German infant and child mortality data before 1875 is questionable. Such consistency lends further support to the notion that infant mortality is an adequate proxy for infant and child mortality, at least in Germany.

ANALYSIS OF PRUSSIAN DATA

From a theoretical perspective it seems likely that higher infant mortality should ultimately be associated with higher fertility, and vice versa. We attempt to evaluate both the effect of infant mortality on fertility and the effect of fertility on infant mortality using fixed-effects models and instrumental variables estimation applied to data from 407 Prussian Kreise (administrative districts) and 54 cities in Prussia from 1875 to 1910.[4]

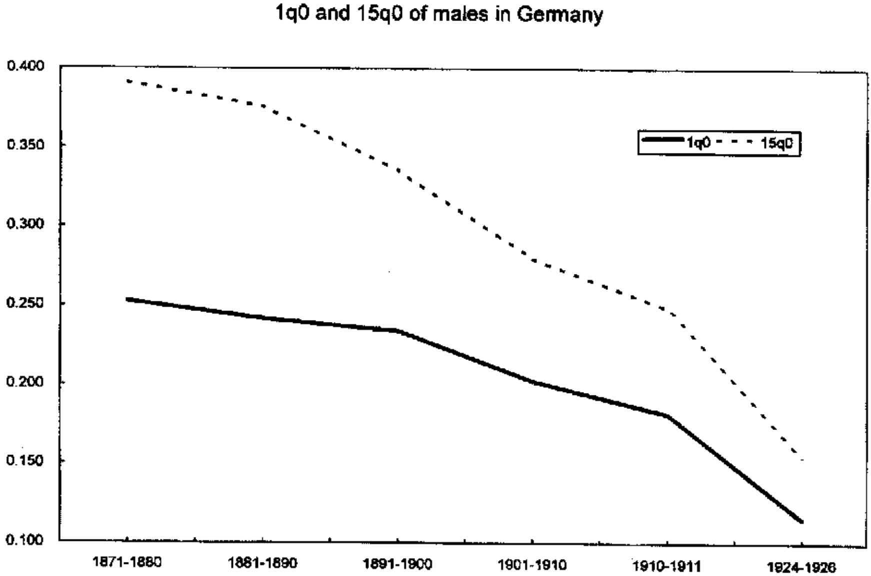
The following equation system describes the structural relationships between fertility and mortality:

$$F_{i,t} = \alpha_1 Y_1^F + \alpha_2 X_{i,t} + \alpha_3 Z_{i,t}^F + \alpha_4 M_{i,t} + \tilde{\alpha}_{5,i} + \tilde{\delta}_t + \tilde{\varepsilon}_{i,t},$$

$$M_{i,t} = \beta_1 Y_1^M + \beta_2 X_{i,t} + \beta_3 Z_{i,t}^M + \beta_4 F_{i,t} + \tilde{\beta}_{5,i} + \tilde{\theta}_t + \tilde{\upsilon}_{i,t}.$$

Here F and M refer to appropriate measures of fertility and infant (and/or child) mortality in the subpopulation of region i at time t. Y is a matrix of unchanging characteristics of the regions that influence fertility or mortality (indicated by superscripts). X is a matrix of changing influences on both fertility and mortality in the regions. Z refers to changing variables in the regions that influence just fertility or just mortality, respectively. $\tilde{\alpha}_{5,i}$ and $\tilde{\beta}_{5,i}$ are disturbances or fixed effects in the two equations that do not change over time, but are specific to the regions. $\tilde{\delta}_t$ and $\tilde{\theta}_t$ are disturbances to the two equations that are the same across all regions, but that vary over time. Finally, $\tilde{\varepsilon}_{i,t}$ and $\tilde{\upsilon}_{i,t}$ are disturbances to the

[4]See Galloway et al. (1994, 1995) for details regarding these two data sets.



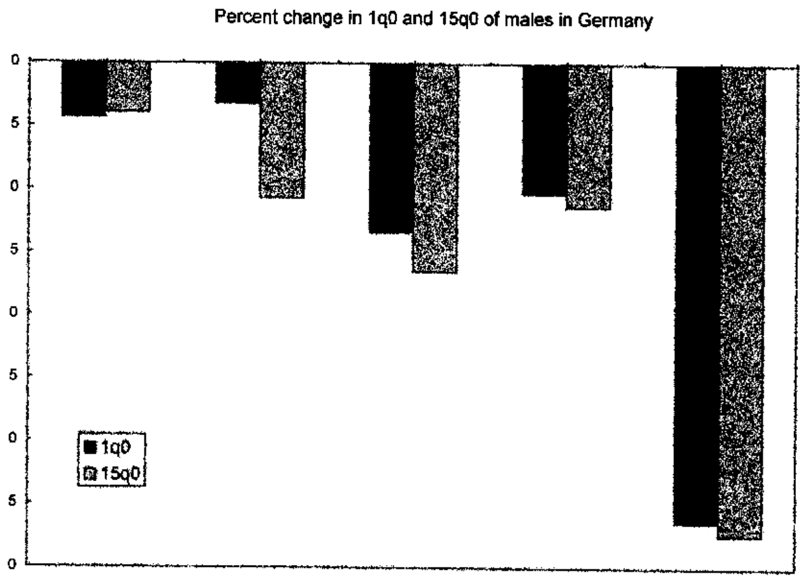


FIGURE 6-1 Level and change of male infant and child mortality rates in Germany, 1871-1926. SOURCE: Statistischen Reichsamt (1930:168).

equations that are specific to time period and to region. Because our data represent non-overlapping 5-year averages, the short-term relations will be largely masked, so long-term and medium-term relations will dominate the estimates.

We use this pair of equations to attempt to approximate a far more complicated dynamic pair of equations that would explicitly include the long-run adjustment processes that bring fertility and mortality to similar levels. In the given equations, the fixed-effect terms are used to represent the outcome of these long-term adjustment processes. The coefficients α_4 and β_4, which represent the influence of mortality on fertility and fertility on mortality, therefore abstract from these long-run adjustment processes and represent only the medium-term influences of one on the other. For example, they would not reflect the possible development of social institutions to motivate high fertility in the face of high mortality. When we estimate this model, which includes fixed effects, therefore, the estimated coefficients should reflect only the medium-term adjustment processes that we believe to be particularly informative for policy considerations. We often refer to such fixed-effect estimates as change estimates, because they are shaped entirely by the relation of changes over time within each Kreis, and not at all by differences in fertility and mortality between Kreise.

By contrast, estimates of these equations based on a single cross section, as are common in the literature, mainly reflect the outcome of the long-term adjustment processes and of correlations of right-hand variables with persistent features of the different geographic units, such as agrarian system, local culture, political orientation, or breastfeeding practices, yielding coefficient estimates that are inconsistent or irrelevant for policy. When suitable instruments are available, two-stage least squares can be used to avoid the biases arising from correlations of right-hand variables with persistent influences; however, suitable instruments can rarely be found in this context. Furthermore, instruments will not solve the problem that long-term adjustment processes probably dominate cross-sectional estimates.

We expect that over the long run, the levels of fertility and mortality will be positively correlated. These long-run correlations would show up in correlations of the fixed effect disturbances $\tilde{\alpha}_{5,i}$ and $\tilde{\beta}_{5,i}$. These correlations would bias any estimate of the structural coefficients relating fertility and mortality directly. Furthermore, the system described by these equations is simultaneous, with both fertility and mortality endogenous. Attempts to estimate the equations by ordinary least squares would yield biased estimates.

Now suppose we difference all the variables, representing the differenced values by lower case letters. We will actually be using fixed-effects estimators, but looking at the effects of differenced variables is a simple way to examine the consequences of using the fixed-effect model. Differenced disturbance terms are represented by the same Greek characters, but without the tildes. In this case, the region-specific disturbances that do not change over time disappear, as do all other variables that do not change over time. We then have

$$f_{i,t} = \alpha_2 x_{i,t} + \alpha_3 z_{i,t}^F + \alpha_4 m_{i,t} + \delta_t + \varepsilon_{i,t},$$
$$m_{i,t} = \beta_2 x_{i,t} + \beta_3 z_{i,t}^M + \beta_4 f_{i,t} + \theta_t + \upsilon_{i,t}.$$

The problem of correlated fixed effects has been removed, but the simultaneity remains. However, we can use the variables z to identify instrumental or two-stage least-squares estimates of the coefficients. These should, in principle, be unbiased—except that there is an additional problem: We do not observe all the relevant x variables. For example, breastfeeding behavior, which affects both fertility and infant mortality, is unobserved.[5] But extended breastfeeding reduces both fertility and infant mortality, and failure to control for its influence will lead to a noncausal positive association of fertility and mortality. Put differently, unobserved and therefore omitted variables in the x matrix will induce a correlation of the error terms in the two equations ε and υ. To be concrete, the omission of breastfeeding variables from the estimated model will lead to a positive correlation in the disturbances, resulting in an upward bias in the estimated structural coefficients on fertility and mortality; the omission of health and illness variables may have the opposite effect. This could be a serious problem for which we have no remedy. Therefore, this potential source of bias must be kept in mind when interpreting the results.

Two-Stage Least-Squares Estimation

Our estimation model reflects the general theoretical perspective outlined above in the use of fixed effects and controls for the endogeneity of infant mortality. However, we have taken a rather eclectic approach to inclusion of socioeconomic influences on fertility and have not imposed any mathematical structure on the relations to be estimated beyond the usual assumption that our linear model approximates some true but unknown nonlinear specification.

One of the findings from our review of earlier research on fertility decline in Europe is that many studies that purport to say something about fertility decline (change) only examine fertility level. Preston (1978:1) states "a central problem in modern population studies . . . is . . . the degree to which changes in mortality can be expected to induce changes in fertility." This clearly poses the question in terms of changes and not of levels. We believe that this is indeed the appropriate question, and that it is not an issue that can be resolved by studying the relationship between levels of fertility and mortality.

[5]In Prussia, longitudinal data on breastfeeding are available only for Berlin. Breastfeeding in Berlin decreased significantly from 1885 to 1910 (Kintner, 1985:169-170) while both marital fertility and infant mortality were declining substantially. It is not known whether other areas in Prussia experienced similar trends.

In our analyses of pooled cross-sectional time series we first examine the "between" estimators, regressions on the means over time of each Kreise or city, that give us level effects. We also generate "within" or "fixed-effects" estimators, regressions that allow each Kreis or city to have its own intercept. This effectively measures how changes in our independent variables affect changes in our dependent variable and is the more appropriate approach for explaining fertility change.[6]

We have dealt at length with the theoretical expectations and empirical findings of fertility decline in 407 Prussian Kreise (Galloway et al., 1994) and 54 cities (Galloway et al., 1995) using quinquennial data from 1875 to 1910[7] and pooled cross-sectional time series ordinary least-squares methods.[8] A detailed analysis of infant mortality decline in the Kreise and cities of Prussia is in progress (Galloway et al., 1996).[9] We expect fertility to influence infant mortality and simultaneously we expect infant mortality to influence fertility. Instrumental variables estimation, two-stage least squares in this case, appears to be an appropriate method for estimating these effects. All the variables in all the models are defined in Appendix Table 6-A1. Our fertility model is shown in Appendix Table 6-A2 followed by a summary of regression results in Appendix Tables 6-

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[6]Brass and Barrett (1978:212) also favor pooled cross-sectional time series studies done of areas within a country.

[7]Prussia became a state within Germany in 1871, but continued to maintain its own statistical bureau that published detailed demographic and economic data until the early 1930s. In 1910 the population of Prussia was just over 40 million, about 70 percent of Germany. If Prussia had been a country, it would have been the largest country in Europe excluding Russia. It covered most of modern-day Germany north of the Main River and most of the western half and northern quarter of modern-day Poland. Registration data are available annually and census data quinquennially from 1875 to 1910 for Kreise and major cities. Kreise are administrative units similar to U.S. census tracts though much larger (average population 60,000) and, like census tracts, tended to split over time. There were about 400 Kreise in Prussia in 1875 and about 600 by 1910, with the total area of Prussia being virtually constant. To maintain spatial consistency over time, we combined many of the split Kreise. This resulted in 407 Kreise, each with a constant area, from 1875 to 1910. Only seven of the 407 Kreise were 100 percent urban. To examine fertility decline within a strictly urban setting, we created another data set consisting of 54 cities, whose area we allowed to change over time, realizing that any area incorporated into a city was itself likely to be highly urban.

[8]We use general marital fertility rate (GMFR) as our measure of marital fertility in this and in all our previous analyses of Prussian Kreise. GMFR is defined as the number of legitimate births per 1,000 married women aged 15-49. A 5-year average centered on the census year is used. More detailed marital age structure data are not available. Coale and Treadway (1986:153) note that there would have been little difference in their findings if GMFR had been used instead of *Ig*. In fact we find that in 54 Prussian cities from 1875 to 1910 where data are available to calculate both *Ig* and the GMFR that *Ig* and GMFR are highly correlated ($r = 0.97$) and that the two measures are virtually interchangeable.

[9]The average infant mortality rate in Prussia in 1900 was 179, somewhat above that found in most less developed countries today. In 1992 Mozambique had an estimated infant mortality rate of 162, the highest of any county in the world (World Bank, 1994:214).

A3 and 6-A4. The infant mortality model can be found in Appendix Table 6-A5 with regression results summarized in Appendix Tables 6-A6 and 6-A7. We focus on the relationship between fertility and infant mortality. The findings for the other right-hand side variables have been discussed elsewhere (Galloway et al., 1994, 1995, 1996).

Estimation of the Fertility Equation

High fertility and the shorter birth intervals it involves might well cause higher infant and child mortality. To avoid the possible bias associated with ordinary least-squares estimation, we need instruments for infant mortality that are correlated with infant mortality but not correlated with the component of its variance that might be influenced by fertility. We believe that male mortality at older ages provides a nearly ideal instrument in this case. Fertility should affect only the mortality of children. We do not know the upper limit of the range of ages that might be affected by high fertility, so we avoided using the mortality even of teenagers. The mortality of women depends in part on maternal mortality, which would depend on fertility, so we avoided using female mortality. For these reasons, we decided to use the mortality of adult males. Because of the cost of data entry, we limited ourselves to male mortality in the 30-34 age group. This choice was based in part on an examination of a correlation matrix for mortality at different ages for both Prussian Regierungsbezirke and historical Swedish data, which showed that death rates at age 30-34 were relatively highly correlated with infant mortality. Other age groups had correlations that were nearly as high, so the exact choice makes little difference, and ideally we would have used mortality over a broader range.

The idea is that male mortality at age 30-34 is a useful index of the general level of mortality in the population, reflecting all local factors that influence mortality, such as standard of living, nutrition, general sanitary conditions, ecological and epidemiological conditions, and the quality of health care. At the same time, it does not reflect the particular influence of either breastfeeding conditions or of fertility and therefore should not correlate with marital fertility.

Unfortunately, age-specific death rates are available only for Regierungsbezirke (very large areas, similar to provinces) of which there were 36 in Prussia. Given 407 Kreise in Prussia, there were on average about 11 Kreise per Regierungsbezirk. We applied Regierungsbezirk male mortality at age 30-34 for each of the eight quinquennial periods from 1875 to 1910 to the Kreise or city within the Regierungsbezirk. This instrument captures only broad regional variations in mortality, but not local differences, reducing its usefulness.

Appendix Table 6-A3 presents the ordinary least-squares and two-stage least-squares findings for Kreise, for both levels and changes. Looking at levels, we find that when we use ordinary least squares the estimated coefficient on infant mortality is negative and marginally significantly different from zero, but, using two-stage least squares the coefficient switches to positive (as we would expect from theory) but is still only marginally significant. Looking at changes, two-stage least-squares estimates of infant mortality are four times higher than in ordinary least squares and highly significantly different from zero.

In the analysis of 54 Prussian cities (Appendix Table 6-A4), we have an additional instrument for infant mortality. The variable sanitation represents cumulative municipal sanitation bond debt per capita[10] and represents a rough measure of the development of sanitation infrastructure. It is available only for cities. It is difficult to see any reason why this variable would have a direct effect on fertility (unless by raising the healthiness of women it increased their fecundity, which seems unlikely). Looking at levels, there is little difference between ordinary and two-stage least-squares estimates, with the estimated coefficient on infant mortality being negative and not significantly different from zero in both cases. In the more relevant estimates of change, the estimate on infant mortality is positive and highly significantly different from zero and quite a bit larger than the ordinary least-squares estimate.

To sum up, we find little evidence of any statistically important influence of infant mortality on fertility in our analysis of levels. On the other hand, infant mortality has a very strong and positive impact on fertility if we consider changes.

Estimation of the Infant Mortality Equation

The infant mortality models are shown in Appendix Table 6-A5. Instruments for the general marital fertility rate include proportion of workers employed in religious occupations, mining, and manufacturing; measures of the development of financial services; and the married sex ratio (a measure of spousal separation). All these are theoretically related to fertility, but probably have little effect on infant mortality. Using the two-stage least-squares regressions for Kreis levels, the effect of fertility on infant mortality is insignificant. However, in our regressions on changes, the estimate is positive and highly significant. Similar results are found in our analysis of cities (Appendix Table 6-A7).[11]

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[10]The variable's definition is based on a listing of municipal debt outstanding by purpose of loan (e.g., sanitation), date of loan, and amount of loan. The source was published in 1906, included loans through 1905, and covered all large cities in Prussia, including our 54. There was virtually no sewage construction in Germany before 1875, most loans were long term (over 30 years), and we assumed that it took about 5 years at most to complete construction on the project. Thus, the sanitation variable is cumulative municipal debt outstanding for sanitation loans per capita, calculated for each census period, and then lagged 5 years to allow for construction. For details see Galloway et al. (1996).

[11]Our theoretical models include both areal fixed effects and period fixed effects; however, we prefer to estimate the coefficients using only areal fixed effects. When only areal fixed effects are

Simultaneous Equation Bias

The most intuitive interpretation of simultaneous equation bias would be the following. A regression of fertility on mortality (and other variables) yields a positive coefficient. But perhaps this is because high fertility is causing high mortality, rather than the other way around. This would lead the estimated coefficient to overstate (be more positive than) the true effect of mortality on fertility; that is, it would have a positive bias. In this case, dealing with the simultaneity by means of instrumental variables or in some other way should reduce the size of the positive coefficient on mortality. But in fact, in our instrumental variable estimates, the size of the estimated coefficient on mortality gets larger, not smaller (i.e., the bias is apparently negative, not positive).

This unexpected outcome does not mean that something is necessarily wrong with our analysis, such as an inappropriate choice of instruments or a misspecified

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included, then the coefficients are estimated so as to explain optimally the pattern of changes over time within each Kreis, regardless of the overall levels of fertility and infant mortality in the Kreis. Therefore, the coefficients can be thought of as being based on the individual histories of each Kreis, with each taken as a case study. When both areal and period fixed effects are included in the estimation, the interpretation is somewhat changed. Now the coefficients are estimated so as to explain optimally the differences between changes over time in each Kreis and those changes that occurred nationally, or in the average of all the individual Kreise. Therefore, the overall national (or average across Kreise) trends in fertility and infant mortality have no effect on the estimates. Fertility could be rising in every Kreis at the same time that infant mortality was falling, to take an extreme example, but the estimated coefficients could still indicate a positive effect of one on the other. For this reason, our preferred estimates are those that include areal fixed effects, but not period fixed effects. Nonetheless, given that both fertility and mortality did decline in Prussia over this period (counter to the extreme example given above), some readers may believe that the model that includes period effects provides a more rigorous test and cleaner estimate than our preferred model.

We estimate the model with and without period effects. The effects of infant mortality on fertility differ as follows. The ordinary least-squares estimate using 407 Kreise with areal effects is 0.242 (Appendix Table 6-A3), with areal and period effects 0.140. The two-stage least-squares estimate using 407 Kreise with areal effects is 1.028 (Appendix Table 6-A3), with areal and period effects 0.709. The ordinary least-squares estimate using 54 cities with areal effects is 0.337 (Appendix Table 6-A4), with areal and period effects 0.165. The two-stage least-squares estimate using 54 cities with areal effects is 1.836 (Appendix Table 6-A4), with areal and period effects –0.036. All estimates are highly significant, except the last which is insignificant.

For the Kreise-based estimates, inclusion of period effects leaves the coefficients highly significantly greater than zero, but reduces their size by 30-40 percent, for both ordinary and two-stage least-squares estimates. For the smaller sample of cities, however, the changes are greater: The ordinary least-squares coefficient is reduced by a half, while remaining highly significantly greater than zero, but the two-stage least-squares coefficient now becomes very slightly negative (and its difference from zero is insignificant). All estimated coefficients for the effect of fertility on infant mortality (not shown) remain highly significantly greater than zero, but they also are decreased in size by 30 or 40 percent. Although our preferred estimates are those with only areal fixed effects, those with period effects also indicate a positive influence of infant mortality on fertility, on the whole.

model. The direction of bias is much more complicated than the intuitive interpretation would suggest.[12] If the true effect of mortality on fertility is α and of fertility on mortality is β, and if the errors in the fertility and mortality equations are ε and υ, respectively, then the sign of the bias is given by the sign of $\beta\sigma_{\varepsilon}^{2}/(1-\alpha\beta)+\sigma_{\varepsilon\upsilon}^{2}/(1-\alpha\beta)$ We strongly expect α and β to be positive. If there is no correlation of errors across the two equations, so that $\sigma_{\varepsilon v}^{2} = 0$, then the sign of the bias depends on whether $\alpha\beta$ is greater than 1, in which case it is negative, or less than 1, in which case it is positive. (If $\alpha\beta$ is greater than unity, then the system is dynamically unstable.) However, because potentially important variables are omitted from the equations, it is quite possible that $\sigma_{\varepsilon v}^{2}$ does not equal zero. Breastfeeding presumably reduces both fertility and mortality, and therefore its omission leads to a positive covariance. Variations in health and morbidity might reduce fertility and raise mortality, so their omission might lead to a negative covariance. For these reasons it is not clear a priori whether the correction for simultaneous equation bias should increase or reduce the estimated effect of mortality on fertility.

OVERVIEW OF PREVIOUS RESEARCH

Nearly all earlier studies of European historical fertility decline regress levels of fertility on levels of independent variables. Although this strategy does tell us something about fertility levels, it tells us little about fertility change.

We assess earlier studies on marital fertility[13] level and change in Europe using regional units of analysis (e.g., districts, counties, provinces) of countries or large portions of countries where some attempt has been made to employ multivariate techniques.[14] We restrict our overview to published research. Those studies that do not examine infant, child, or general mortality are of course

[12] The following discussion is based on a communication from Mark Montgomery, whom we thank.

[13] We review research that examines only marital fertility, with one exception. The study of Netherlands analyzes only crude birth rate, but we decided to include it because it is the only study of fertility change in the Netherlands. By focusing on marital fertility we are abstracting from any response of nuptiality to changed mortality. However, if nuptiality responded strongly positively to mortality change, so that the period of exposure to risk of childbearing within marriage did likewise, possibly even overcompensating, then focusing on marital fertility could be misleading. It might be useful to study the two together.

[14] F. van de Walle (1986:225-227) examined fertility levels and changes in relation to infant mortality levels and changes in historical European countries. She shows bivariate correlations of ${}_1q_0$ with I_g and I_f and I_m across countries of Europe and also across provinces within Europe. She also looks at bivariate correlations of changes in ${}_1q_0$ and I_g and claims these to be inconclusive, but in fact every significant correlation is positive in both periods (first period is 1870-1900 or so; the second is 1900-1930 or so). Cross-period correlations, for which there are no apparent justifications, sometimes have perverse signs. In another study, Fialová et al. (1990:102) present only bivariate correlations between levels of independent variables and marital fertility level in their analysis of

excluded. Urban and rural differences are shown where available. We differentiate between analyses of level effects and change effects.[15] Generally, we examine only those periods before World War II.

Tables 6-1 and 6-2 list summaries of previous research according to level and change, respectively, in alphabetical order by country. In all cases but one, marital fertility is the dependent variable. The author, date of publication, fertil-

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fertility in Czechoslovakia. Bivariate correlations are at best suggestive, and useful conclusions for the analysis of theoretically complex models cannot be drawn from them.

Some examine fertility decline using time series techniques applied to one region (i.e., only one unit of analysis). For example, Lutz (1987:84) examines annual estimates of the Coale-Trussell index of family limitation *m* for Finland from 1873 to 1917 along with age at marriage, education, gross domestic product, marriage rate, life expectancy at age 5, and infant mortality (which showed a positive, but insignificant association with *m*). Crafts (1984b:583) examined GMFR in England and Wales from 1877 to 1938 along with measures of wages, income, prices, illegitimate fertility, and child mortality (which was positively related to GMFR, but not significant). Haynes et al. (1985:560-565) examined annual crude birth rate and crude death rate data, but no other variables, in four countries. Because these studies use annual data, important factors such as sectoral employment, female labor force participation, urbanization, religion, language, and other census-derived variables are usually missing from the models. Furthermore, it is difficult to know what to say about long-term fertility decline when looking at the results from short-run time series analysis where the long-term trend has been necessarily removed. See Galloway (1994) for a review of the literature on short-run fluctuation analyses and the various interactions among annual variations in fertility, mortality, nuptiality, migration, wages, and weather.

There are many articles about fertility and infant mortality based on data from family reconstitution or genealogical studies, but few attempt to estimate the influence of infant or child mortality on fertility (Knodel, 1978). Furthermore, most of these are about one parish, most use little data beyond the demographic indices generated by family reconstitution, and very few cover the period of the fertility transition. Most are plagued by the usual problems of family reconstitution: no subsequent information on persons who leave the parish, no information on those who lived in the parish but who were not born or married or did not die there, a lack of total population counts and age structure for the entire parish population, possible selectivity based on the fact that the only registers analyzed are those that survived. See Flinn (1981) for an old but still useful overview of family reconstitution studies. For more recent analyses of both fertility and infant mortality using micro-level data see Knodel (1988).

[15]The few attempts to derive a measure of the "onset" of long-term fertility decline are particularly problematic. The Princeton European Fertility Project's definition based on a 10 percent decline in fertility seems arbitrary, and the plateau from which fertility is supposed to have declined is typically based on only a few observations. Teitelbaum (1984:178-179) looked at his estimate of onset of fertility decline in Great Britain using the period around 1851-1931, along with income, urbanization, females not in the labor force, sectoral employment, religion, ethnicity, and infant mortality rate (which was significantly negatively associated with onset of fertility decline). Knodel (1974:238-239) examined fertility decline using his estimate of secular fertility decline in 71 provinces of Germany for the period 1875-1910, along with levels of religion, bank account, sectoral employment, literacy, and the infant mortality rate (which was sometimes positively and sometimes negatively associated with I_g decline, depending on which measures of fertility decline from onset that Knodel used).

In some studies, we find analyses of fertility change in relation to levels of the independent variables (Knodel, 1974; Lesthaeghe, 1977; Benavente, 1989; Reher and Iriso-Napal, 1989; Haines, 1989). These results are difficult to interpret.

ity measure, number of units of analysis, periods analyzed, independent variables, and sign and significance of the mortality variable are discussed. In the following overview, the estimated mortality coefficient is considered statistically significant, that is, statistically different from zero, if the *t* statistic probability is 5 percent or less.[16] It should be understood that it may be difficult to compare findings directly because of differences in definitions of variables, number and type of control variables, and methods used.

Level of Fertility and Levels of Independent Variables

Belgium

Lesthaeghe (1977:213) examined I_g in the 22 arrondissements of Flanders in four periods between 1880 and 1910 along with language, industrialization, literacy, political affiliation, and infant mortality rate. He did the same for the 19 arrondissements of Wallonia. Infant mortality rate was significantly positively associated with I_g only in Flanders in 1880 and was insignificant otherwise, perhaps a result of the small sample sizes involved. We note that seven of the eight estimated infant mortality coefficients were positive.

England and Wales

Haines (1979:68) examined marital fertility in a random sample of 125 registration districts in England and Wales in 1851, 1861, and 1871 along with sectoral employment, female employment, urbanization, net migration, sex ratio, and infant mortality rate. He found that infant mortality was negatively associated with fertility in all three periods, but insignificant in 1851. Haines notes, "Although several adjustments were tried to correct for underreporting of infant deaths, it was felt safest simply to divide uncorrected infant deaths by uncorrected live-births to obtain the infant mortality rate" (1979:60-62). It seems likely that the degree of underreporting of infant deaths and live births varied independently, and perhaps substantially, from district to district. As a consequence, it is difficult to interpret these infant mortality regression estimates.

Crafts (1984a:94,98) studied the general marital fertility rate (GMFR) in 619 urban districts of England and Wales for four periods from 1871 to 1911, along with single woman labor force participation rate, income, migration, literacy, sectoral employment, and child mortality (which was significantly positively associated with the GMFR in 1911, but negative and insignificant in 1871, 1881,

[16]In much of the earlier research (Livi Bacci, 1971, on Portugal, Knodel, 1974, on Germany, and Coale et al., 1979, on Russia) partial correlation coefficients were published, but not *t* statistics. We have calculated the *t* statistics using the published partial correlation coefficients (Wonnacott and Wonnacott, 1979:180) so as to obtain some idea of the statistical significance of the estimates.

TABLE 6-1 Sign and Statistical Significance of Estimates, Elasticity, and (d*F*/d*q*)/*F* from Regressions of Marital Fertility Level on Infant or Child Mortality Level in Multivariate Studies of European Fertility Decline

Country and Region	Number of Districts	Method	Year: 1851	1860	1870
Belgium					
Flanders	22	TSLS			
Wallonia	19	TSLS			
England and Wales					
National	125	OLS	(neg) (−0.03)	(NEG) (−0.06)	(NEG) (−0.06)
Urban	619	OLS			(neg)
National	590	OLS			
Urban	222	OLS			
Rural	368	OLS			
Towns	101	OLS			
National	600	OLS			
France	81	OLS		POS	
Germany					
National	71	OLS			
National	71	OLS			
Prussian Kreise	407	OLS			
Prussian Kreise	407	TSLS			
Prussian cities	54	OLS			
Prussian cities	54	TSLS			
Italy					
North and central	53	OLS			
South	34	OLS			
Veneto	57	OLS			
Netherlands	375	OLS		POS	POS
Portugal	18	OLS			
Russia					
Rural	50	OLS			
Urban	50	OLS			

1880	1890	1900	1910	1920	1930	Source
POS	pos	pos	neg			1
0.17 *0.88*	0.10 *0.52*	0.07 *0.37*	0.03 *−0.18*			
pos	pos	pos	pos			1
0.21 *1.62*	0.13 *1.00*	0.57 *4.63*	0.30 *3.09*			
						2
(neg)	(neg)		POS			3
			0.07			
	(NEG)		(NEG)			4
	(NEG)		(POS)			4
	(NEG)		(NEG)			4
			POS			5
POS						6
						7
POS						8
	POS					9
	0.05 *0.25*					
	neg					10
	−0.04 *−0.22*					
	pos					11
	0.17 *0.95*					
	neg					12
	−0.04 *−0.20*					
	neg					13
	−0.16 *−0.82*					
neg			POS		POS	14
NEG			POS		POS	14
(neg)						15
POS	POS					16
			pos		pos	17
		pos			POS	18
		pos			POS	18

continued on next page

TABLE 6-1 *(continued)*

Country and Region	Number of Districts	Method	Year 1851	1860	1870
Spain					
Catalonia	84	OLS		(neg)	
Rural	50	OLS			
Capital cities	50	OLS			
Sweden					
National	25	OLS			
Rural	25	OLS			
Urban	25	OLS			

NOTES: OLS, ordinary least squares. TSLS, two-stage least squares. POS, estimated coefficient is positive and significant within 5 percent. pos, estimated coefficient is positive but not significant within 5 percent. NEG, estimated coefficient is negative and significant within 5 percent. neg, estimated coefficient is negative but not significant within 5 percent. Parentheses mean that the estimated coefficient is difficult to interpret. See text for explanation. The number below the sign is the elasticity, defined as the estimated coefficient multiplied by the mean of infant or child mortality divided by the mean of marital fertility. It is calculated wherever possible. The number in italics next to the elasticity is $(dF/dq)/F$ where F is a measure of marital fertility and q is infant mortality. It is calculated wherever possible. For clarity in the table, I used the opposite sign of the author's estimate where the infant or child mortality measure is life expectancy at birth or some measure of survivorship.

and 1891). However the 1871-1891 regressions suffer from important data limitations, using 1911 data for four variables (Crafts, 1984a:97-99). In a later study, Crafts (1989:332-333) restricted his analysis to 1911 data in 101 towns using similar variables along with age of wife at first marriage and found that births per woman was significantly and positively associated with infant mortality.

Woods (1987:302) studied I_g in 590 districts of England and Wales for 1891 and 1911 along with coal miners, farm servants, females employed in textiles and as servants, literacy rate, and the probability of a child surviving from age 1 to 10 in 1861, defined by Woods as l_{10}/l_1 (1987:301). He used this 1861 child survivorship measure (which excludes infant mortality) as his mortality variable in his 1891 and 1911 fertility regressions (Woods 1987:301). It seems risky to try to estimate fertility using a child mortality measure lagged 30 and 40 years. It seems likely that districts experienced varying rates of mortality decline during

1880	1890	1900	1910	1920	1930	Source
						19
	(POS)	(pos)		POS		20
	(0.28)	(0.18)		0.31		
	(neg)	(NEG)		POS		20
	(−0.19)	(−0.61)		0.53		
			POS	POS		21
			0.45 *6.80*	0.39 *7.62*		
		POS				21
		0.36 *4.47*				
		pos				21
		0.06 *0.52*				

SOURCES: 1: Lesthaege (1977:106, 172, 213); 2: Haines (1979:63, 68); 3: Crafts (1984a:94, 98); 4: Woods (1987:302-304); 5: Crafts (1989:332-333); 6: Friedlander et al. (1991:341); 7: van de Walle (1978:287); 8: Knodel (1974:238); 9: Richards (1977:543-546). Results are based on a regression of pooled data for 1880, 1885, 1890, 1900, and 1910; 10: Galloway et al. (1994:143-152). Results are based on regressions using the average of data over the periods 1875, 1880, 1885, 1890, 1895, 1900, 1905, 1910. Also Appendix Table 6-A3, Equation 1, level; 11: Appendix Table 6-A3; 12: Galloway et al. (1995:38-39). Results are based on regressions using the average of data over the periods 1875, 1880, 1885, 1890, 1895, 1900, 1905, 1910. Also Appendix Table 6-A4, Equation 2, level; 13: Appendix Table 6A-4, Equation 2, level; 14: Livi Bacci (1977:194, 198); 15: Castiglioni et al. (1991:114); 16: Boonstra and van der Woude (1984:23, 34); 17: Livi Bacci (1971:122); 18: Coale et al. (1979:65); 19: Benavente (1989:229); 20: Reher and Iriso-Napal (1989:408, 412, 419, 420); 21: Mosk (1983:186, 252, 254, 257) and Sweden Central Bureau of Statistics (1969:115).

this period of 30-40 years. In both 1891 and 1911, Woods' 1861 measure of child survivorship was positively and significantly associated with I_g, although it is difficult to know what to make of this finding.

Woods (1987:303) examined 222 urban districts using the above data and found that the estimated coefficient on 1861 child survivorship ages 1-10 was significantly positive in 1891 but shifted to significantly negative in 1911. Woods (1987:304) analyzed 368 rural districts using the above data and found a significantly positive association between I_g and 1861 child survivorship ages 1-10 for both periods.

Friedlander et al. (1991:341) examined I_g in 600 districts of England and Wales for the period around 1870-1890 along with density, sectoral employment, females not in the labor force, urban proximity, and life expectancy at birth (which was significantly and negatively associated with I_g). We expect varia-

TABLE 6-2 Sign and Statistical Significance of Estimates, Elasticity, and $(dF/dq)/F$ from Regressions of Marital Fertility Change on Infant Mortality Change in Multivariate Studies of European Fertility Decline

Place	Number of Districts	Method	Period	Estimate	Elasticity	$(dF/dq)/F$	Source
Germany							
Prussian Kreise	71	OLS	1880-1910	POS	0.39	1.96	Richards (1977:543, 545)
Prussian Kreise	407	OLS	1875-1910	POS	0.17	0.95	Galloway et al. (1994:143, 152) and Appendix Table 6-A3, Equation 1, change
Prussian Kreise	407	TSLS	1875-1910	POS	0.70	3.92	Appendix Table 6-A3, Equation 1, change
Prussian cities	54	OLS	1875-1910	POS	0.29	1.49	Galloway et al. (1995:38, 39) and Appendix Table 6-A4, Equation 2, change
Prussian cities	54	TSLS	1875-1910	POS	1.55	7.97	Appendix Table 6-A4, Equation 2, change
Italy							
North and central	53	OLS	1880-1910	pos			Livi Bacci (1977:194, 199)
North and central	53	OLS	1911-1931	POS			Livi Bacci (1977:194, 199)
South	34	OLS	1880-1910	pos			Livi Bacci (1977:194, 199)
South	34	OLS	1911-1931	pos			Livi Bacci (1977:194, 199)
Netherlands	375	OLS	1871-1890	POS			Boonstra and van der Wonde (1984:36)
Sweden							
Urban	25	OLS	1900-1920	POS	0.09	1.15	Mosk (1983:257)
Urban	25	OLS	1900-1930	POS	0.15	2.00	Mosk (1983:254, 257) and Sweden Central Bureau of Statistics (1969:115)

NOTES: OLS, ordinary least squares. TSLS, two-stage least squares. POS, estimated coefficient is positive and significant within 5 percent. pos, estimated coefficient is positive but not significant within 5 percent. Elasticity is the estimated coefficient multiplied by the mean of infant mortality divided by the mean of marital fertility. It is calculated wherever possible. In the expression $(dF/dq)/F$, F is a measure of marital fertility and q is infant mortality. It is calculated wherever possible. Richards' results for Germany are based on a regression of pooled data for 1880, 1885, 1890, 1900, and 1910 using fixed effects. Galloway et al.'s results for Prussian Kreise and Prussian cities are based on regressions of pooled data for 1875, 1880, 1885, 1890, 1895, 1900, 1905, and 1910 using fixed effects. The Netherlands' regressions use crude death rate and crude birth rate instead of infant mortality and marital fertility.

tions in life expectancy at birth to be dominated by variations in overall child mortality, although a measure of overall child mortality would have been preferred.

At first glance, the picture for England and Wales appears confusing. However, it seems appropriate to be highly skeptical about Haines' and Woods' interpretations of their mortality estimates. Half of the variables in Crafts' 1871-1891 regressions used 1911 data. Thus, we are left with Crafts' analysis of 1911 data and Friedlander et al.'s study. Both find the expected significant and positive relationship between overall child mortality and fertility.

France

Van de Walle (1978:287) examined I_g in 81 departments in France for five periods from 1841 to 1851 along with religion, rural land revenue per capita, urbanization, and life expectancy at birth which was probably significantly negatively[17] associated with I_g. Watkins (1991:161) and Lesthaeghe (1992:275-317) did some research along these lines but neither included any measure of mortality in their analyses.

Germany

Knodel (1974:238-239) studied I_g in 71 provinces for the one period 1875-1910 along with religion in 1880, bank accounts in 1900, sectoral employment in 1882, literacy in 1875, and infant mortality rate in 1875 (which was positively and significantly associated with I_g). Knodel used maximum I_g in the interval as his dependent variable so caution must be used in interpreting these results.

Richards (1977:546) examined I_g in 71 provinces for five periods from 1880 to 1910 along with net migration, religion, urbanization, sectoral employment, and infant mortality rate (which was significantly and positively associated with I_g).

Galloway et al. (1994:152) analyzed general marital fertility in 407 Kreise covering all of Prussia using the average of quinquennial data from 1875 to 1910 along with religion, ethnicity, education, health, female labor force participation, income, mining, urbanization, financial, insurance, communication, sex ratio, and legitimate infant mortality variables. Legitimate infant mortality was negatively, but insignificantly, associated with fertility. However, using two-stage least squares the sign shifted to positive (see Appendix Table 6-A3).

Using the same periods and variables, but excluding urbanization and includ-

[17]Van de Walle (1978:287) suggests, "On the face of it, we end up with an 'explanation' of the fertility decline along the line of population transition theory, with a major role played by the decline of mortality and with an independent influence of income."

ing manufacturing and population size variables, Galloway et al. (1995:39) found that legitimate infant mortality in the 54 largest cities of Prussia was negatively related to general marital fertility, but not statistically significant. The same results were found using two-stage least squares (see Appendix Table 6-A4).

Italy

Livi Bacci (1977:198) examined I_g in 92 provinces for the periods 1881, 1911, and 1931 along with urbanization, sectoral employment, literacy, proportion married, and infant mortality rate in north and central Italy and south Italy. In north and central Italy infant mortality rate was insignificant in 1881, but positively and significantly associated with I_g in 1911 and 1931. In south Italy infant mortality rate was significantly negatively associated with I_g in 1881, but positively and significantly associated with I_g in 1911 and 1931.

Castiglioni et al. (1991:114) examined I_g in 1881 in 57 districts in Veneto along with topography, occupation, females employed in agriculture, migration, and infant mortality rate. Infant mortality was negatively associated with I_g, but test statistics were not provided.

Netherlands

Boonstra and van der Woude (1984:34, 40, 44, 51) examined the crude birth rate in 375 districts for eight periods from 1851 to 1890 along with net migration, religion, density, literacy, soil type, and crude death rate. The crude death rate was significantly and positively associated with the crude birth rate in each of the eight periods. Unfortunately, more refined measures of fertility and infant or child mortality were not available (Boonstra and van der Woude, 1984:24), but because this appears to be the only study of Netherlands fertility decline using multivariate analysis, we decided to include it.

Portugal

Livi Bacci (1971:122) analyzed I_g in 18 provinces in 1911 and 1930 along with sectoral employment, literacy, and infant mortality rate, which was positively associated with I_g in both periods, but significant only in 1930.

Russia

Coale et al. (1979:65) examined I_g for 50 provinces in rural and urban sectors in 1897 and 1926 along with urbanization, sectoral employment, literacy, and infant mortality rate. In each case, infant mortality rate was positively associated with I_g, but the estimates were significant only in 1926.

Spain

Reher and Iriso-Napal (1989:412) examined I_g in the rural sector of 50 provinces for 1887, 1900, and 1920 along with sectoral employment, migration, urbanization, 1936 political election results, literacy, female nuptiality, and $_5q_0$, which was positively associated with I_g in all periods, but significant only in 1887 and 1920.

Reher and Iriso-Napal (1989:420) also examined I_g in the 50 provincial capital cities for the same three periods along with sectoral employment, females in labor force, migration, city size, 1936 political election results, literacy, female nuptiality, and $_5q_0$, which was insignificantly related to I_g in 1887, significantly negatively associated with I_g in 1900, and significantly positively associated with I_g in 1920.

In the 1887 and 1900 regressions, $_5q_0$ was the average of 1860 and 1900 data (Reher and Iriso-Napal, 1989:410). As a consequence it is difficult to know what to make of the regression estimates. In general it seems unwise to use only two data points, 1860 or 1900, to create a child mortality variable covering 40 years. Even if one had $_5q_0$ data for intervening years, a 40-year average centered around 1880 seems an inappropriate indicator of child mortality when trying to explain 1887 and 1900 fertility. Therefore, we have strong reservations about the $_5q_0$ estimates from the regressions for the first two periods.

Benavente (1989:229) studied I_g in 84 selected local areas, not necessarily representative, of Catalonia in 1857 along with nuptiality, sectoral employment, proximity to France, and child mortality, which was negatively associated with I_g, but insignificant. The child mortality measure is number of deaths per 1,000 for children under age 7 in 1837 (Benavente, 1989:227). It is not clear whether this measure included infant deaths, which are known to be severely under-registered during this period. Furthermore, the 84 local areas analyzed are those areas in which registration data have survived, suggesting additional caution in interpreting the regression finding.

Sweden

Mosk (1983:252) examined I_g in 25 counties in Sweden in 1910 and 1920 along with sectoral employment, wage, primary school attendance, and legitimate infant mortality rate, which was significantly positively associated with I_g.

Mosk (1983:256-257) also studied I_g in the rural sector of 25 Swedish counties in 1900 along with sectoral employment, agricultural wage, social structure, and legitimate infant mortality rate, which was significantly positively associated with I_g. He also examined I_g in the urban sector of 25 Swedish counties in 1900 along with nonagricultural wages and legitimate infant mortality rate. The estimate on infant mortality was positive, but not significantly associated with I_g.

Fertility Change and Changes in Independent Variables

Germany

Using a fixed-effects model, Richards (1977:545) examined I_g in 71 provinces for five periods from 1880 to 1910 along with province dummy variables, net migration, religion, urbanization, sectoral employment, and infant mortality rate, which was significantly and positively associated with I_g.

Galloway et al. (1994:152) analyzed GMFR in 407 Kreise covering all of Prussia using quinquennial data from 1875 to 1910. Independent variables included religion, ethnicity, education, health, female labor force participation, income, mining, urbanization, financial, insurance, communication, sex ratio, legitimate infant mortality, and 407 Kreis dummies. From this fixed-effects model, changes in legitimate infant mortality were found to be positively and significantly associated with changes in fertility. Among all the independent variables, legitimate infant mortality was the fourth most important in terms of contribution to predicted change in average GMFR from 1875 to 1910, just behind female labor force participation, insurance, and communication variables (Galloway et al., 1994:156). A two-stage least-squares model also yielded a positive and significant association between fertility change and infant mortality change (see Appendix Table 6-A3).

Using the same periods and variables, but excluding urbanization and the 407 Kreise dummies variables, and including manufacturing, population size, and 54 city dummy variables, Galloway et al. (1995:39) found that change in legitimate infant mortality in the 54 largest cities of Prussia was positively and significantly related to changes in the GMFR. In terms of components of predicted change in average GMFR from 1875 to 1910, legitimate infant mortality was second only to female labor force participation in importance among all the variables considered (Galloway et al., 1995:41). Our two-stage least-squares model also revealed a positive and significant association between changes in fertility and changes in infant mortality (see Appendix Table 6-A4).

Italy

Livi Bacci (1977:199) examined change in I_g in 92 provinces for the periods 1881-1911 and 1911-1931 along with changes in urbanization, sectoral employment, literacy, proportion married, and changes in infant mortality rate in north-central and southern Italy. Changes in infant mortality were not significantly associated with changes in I_g in either region from 1881 to 1911, insignificant in the south from 1911 to 1931, but significantly positively associated with I_g in north-central Italy from 1911 to 1931. Each of the four estimated infant mortality change coefficients was positive.

Netherlands

Boonstra and van der Woude (1984:36) examined change in the crude birth rate in 375 districts for the period from 1871 to 1890 along with change in density; change in crude death rate; and levels of net migration, religion, density, literacy, and soil type. Change in the crude death rate was significantly and positively associated with change in the crude birth rate.

Sweden

Mosk (1983:257) examined changes in I_g in the urban sector of 25 Swedish counties from 1900 to 1920 and 1900 to 1930 along with changes in nonagricultural wages and changes in legitimate infant mortality, which were positively and significantly associated with changes in I_g in both periods.

Interpretation Problems and Some Suggestions

Interpretation problems in the analyses of level effects generally revolve around questionable definitions of the infant and child mortality variables, the use of explanatory variables in inappropriate time frames, and reliance on strictly bivariate correlations. We suspect that most of these problems result from data limitations in the original source material. Nonetheless, we found it necessary to address these problems to obtain a reasonably convincing overall picture of previous findings.

It would be useful if future researchers would include a detailed discussion of the quality of the data, especially of infant mortality, and would provide the reader with an unambiguous explanation of each variable's construction. In many cases we simply were not told what period was covered by the mortality variable. For example, was an infant mortality variable for 1890 based on the average of annual rates from 1888 to 1892, the average from 1889 to 1891, or simply the one year 1890? If the latter, then there are serious questions about the variable's relevance because we know that infant mortality rates can vary substantially from year to year.

Rather than divide a data set into two groups, and then run regressions for each group separately as some authors have done, it might be more instructive simply to create a dummy variable for each group and interact each independent variable with the dummy. This would preserve sample size and enable the researcher to determine whether the estimates for a given group are significantly different from the estimates for another group. Of course the best approach is to operationalize the variable that theoretically distinguishes the groups from each other, rather than rely on a dummy.

We found that using appropriate instruments for the infant mortality variable actually shifted the estimated impact of infant mortality on fertility from negative in our ordinary least-squares regression to positive using two-stage least squares in our analysis of Kreis levels (Appendix Table 6-A3). There was no difference in sign or significance in the ordinary least-squares and two-stage least-squares estimates in terms of change (Appendix Tables 6-A3 and 6-A4). Theory suggests that it is necessary to use statistical methods, such as instrumental variables, which can deal with problem of simultaneity.

Finally, we believe that the appropriate model for analyzing fertility decline involves the regression of changes in fertility on changes in the explanatory variables. We use a fixed-effects model. Although this places greater demands on the data set, both in terms of quality and quantity, it yields the most relevant results.

FINDINGS

Signs and Significance of Estimates

Table 6-1 summarizes the results of the regressions of the infant mortality level on the fertility level. If we discount the estimates that are difficult to interpret, and use our two-stage least-squares estimates for Prussian Kreise and cities instead of our ordinary least-squares estimates, we see that for levels nearly all the estimates are positive (34 of 38), as theory would predict. Of the 24 statistically significant estimates, 23 have positive signs. Of the 14 estimates that are not statistically significant, 11 are positive. Of these 11, 8 had relatively small samples sizes (n less than 26).

Although the study of fertility level may be interesting, analysis of fertility change is most relevant to demographic transition theory. Table 6-2 summarizes the ten regressions that examine the effects of changes in infant and child mortality on changes in fertility. Again we use the two-stage least-squares estimates instead of the ordinary least-squares estimates for Prussian Kreise and cities, although in this case it makes no difference because they have the same signs and significance. For changes the estimates are positive in every case (ten of ten), and significant in seven of them.

Elasticities

To compare the magnitude of the impact of infant mortality on fertility across countries, we calculated, where possible, the elasticity of fertility with respect to mortality. Figure 6-2 presents the distribution of elasticities for level effects for all interpretable estimates.[18] The average elasticity for levels is 0.21 (0.13 if we include the seven estimates that are difficult to interpret). In other

[18] The two-stage least-squares estimates, rather than the ordinary least-squares estimates, are used for Prussia in this and all subsequent discussion.

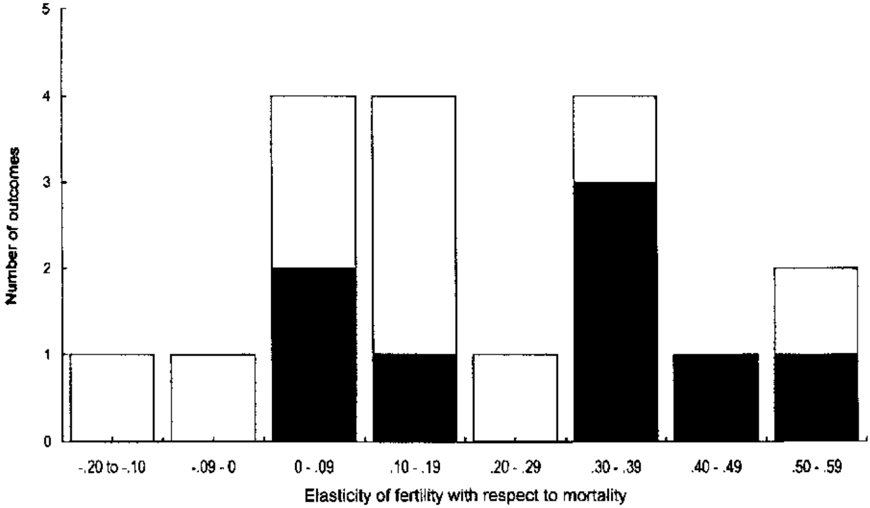

FIGURE 6-2 Distribution of elasticities of fertility with respect to mortality, level effects. NOTES: Only interpretable estimates are included. Mean is 0.21. The dark shading indicates statistically significant estimates. The light shading indicates statistically insignificant estimates. Only the two-stage least-squares estimates are used for Prussia.

words, on average a 10 percent decrease in infant mortality leads to about a 2.1 percent decline in fertility.

We are much more interested in the elasticities derived from studies that examine the effect of changes in infant mortality on fertility. We have such elasticity data for Germany, Prussian Kreise, Prussian cities, and urban Sweden (Table 6-2). The average elasticity is 0.58 and the distribution is shown in Figure 6-3. Elasticities of infant mortality and fertility change models appear to be much higher than in the level models. The elasticity in Germany was 0.39, 0.79 in Prussian Kreise, and 1.55 in Prussian cities. This suggests that the fertility of urban populations in Prussia was much more responsive to mortality changes than that of rural populations. A possible explanation for the high urban elasticity might be that urban couples may have become relatively more aware of infant mortality decline and its relation to infant and child survivorship, which may in turn have caused them to reduce fertility at a relatively faster rate (as measured by elasticity). This increased awareness may have been a result of a greater access to information about infant mortality decline from newspapers, pamphlets, books, organizations, peers, or health workers found in the cities.

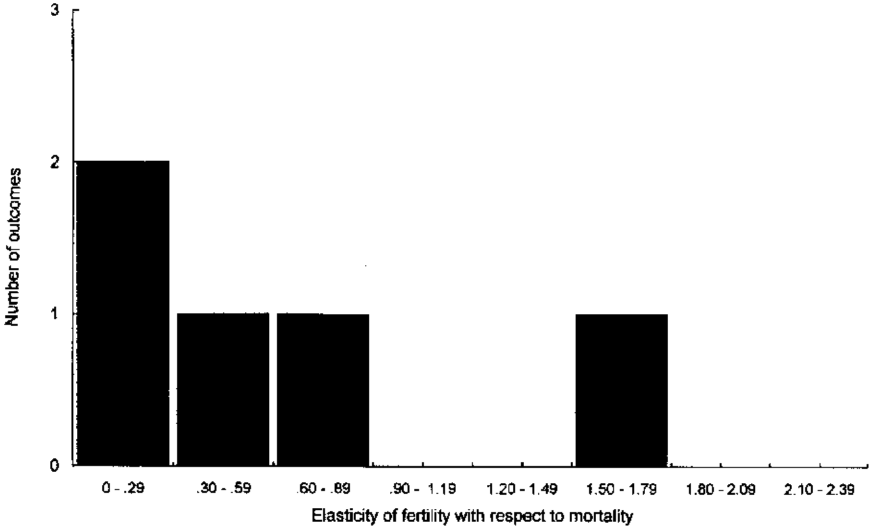

FIGURE 6-3 Distribution of elasticities of fertility with respect to mortality, change effects. NOTES: Only interpretable estimates are included. Mean is 0.58. All estimates are statistically significant. Only the two-stage least-squares estimates are used for Prussia.

Child Replacement and Lactation Interruption

There are two important points of reference for interpreting the size of the estimated effects. First, how large a response of fertility to infant mortality would be required to leave the population growth rate unaffected? And second, how large an effect can be expected from lactation interruption alone?

Let the number of children surviving to some age, say 15 years, be $SF = (1 - Q)F$, where F is the total number of births born over the reproductive years of a woman, and S is the survival probability from birth to age 15. Whereas in earlier equations M referred to general child mortality, here we use $Q = {}_{15}q_0$ and $q = {}_1q_0$ for specific measures, and $S = 1 - {}_{15}q_0$ and $s = 1 - {}_1q_0$ for their complements. The effect on the number of surviving children of a variation in mortality before this age, Q, is $d(SF)/dQ = dF/dQ\ (1 - Q) - F$. There is said to be complete replacement when this is zero, and this corresponds roughly to the situation in which a change in mortality does not affect the long-term population growth rate. In our data we observe marital fertility, not total fertility F, but let us for now ignore this fact. In our data we also observe the infant mortality rate q rather than mortality to age 15, Q. Fortunately, q and Q are fairly closely linked, so we may use q as a proxy for Q. The question then becomes: How sensitively must F respond to q for SF, the number of surviving children, to be constant when q varies? Or what is the benchmark value of $(dF/dq)/F$, call it $[(dF/dq)/F)]^*$, such that $d(SF)/dq$ is 0? Differentiating and solving, we find that $[(dF/dq)/F)]^* = (dQ/dq)/(1 - Q)$, which is the desired benchmark.

To attach an actual number to this benchmark we need to know dQ, dq, and Q. In principle, the relation of Q to q could be anything, and indeed over relatively short periods there can be a great deal of variation in the relation of changes in the two measures. For example, $(dQ/dq)/(1 - Q)$ in Germany is 2.30 in the 1870s, 8.22 in the 1880s, 2.65 in the 1890s, 2.12 in the 1900s, and 1.87 from 1910 to 1925 (Statistischen Reichsamt, 1930:168). However, over a series of longer periods, or for general tendencies as expressed in model life tables, the relation is much more regular. An examination of the Coale and Demeny (1983) Model West Female life tables shows that $\ln(1 - Q)$ is nearly linear in q, with a slope of about 2.5; that is, $\ln(1 - Q) = k - 2.5q$. Differentiating this expression, we find that $(dQ/dq)/(1 - Q) = 2.5$, which is the break-even point. The significance of this convenient empirical fact is that the break-even value is roughly constant over the course of the demographic transition and does not vary with the general level of fertility or mortality. This value of 2.5 agrees well with the value of 2.44 calculated over the entire period 1870-1925 for Germany (Statistischen Reichsamt, 1930:168). Table 6-1 provides $(dF/dq)/F$ for levels, where calculable, and these are plotted in Figure 6-4. The average is about 2.11. The

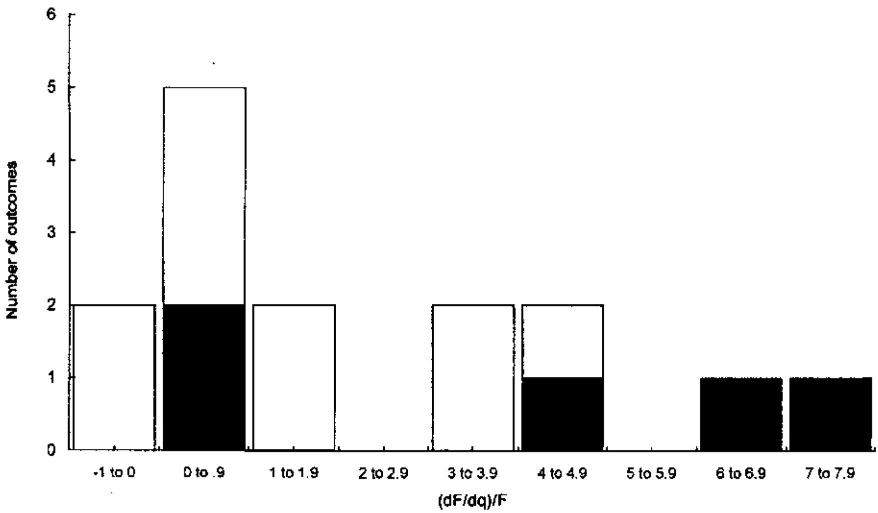

FIGURE 6.4 Distribution of sensitivities of fertility to mortality measured by $(dF/dq)/F$ level effects. NOTES: Only interpretable estimates are included. Mean is 2.11 The dark shading indicates statistically significant estimates. The light shading indicates statistically insignificant estimates. Only the two-stage least-squares estimates are used for Prussia.

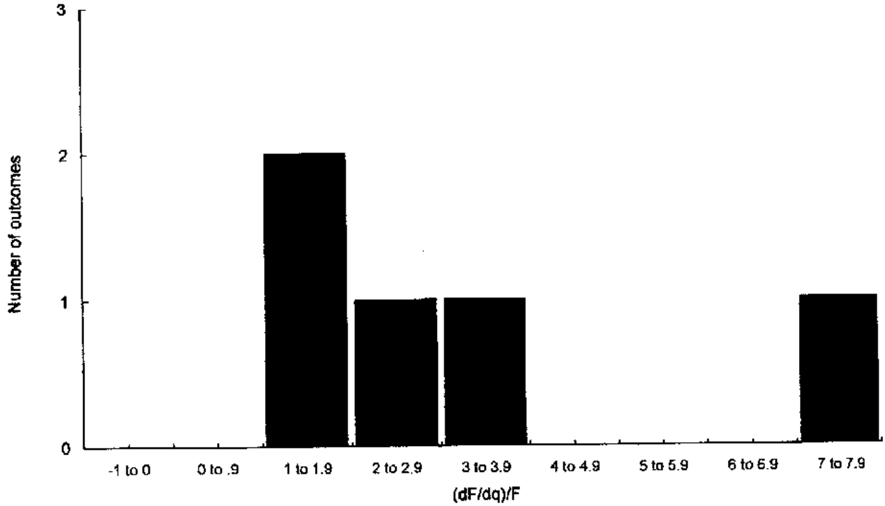

FIGURE 6-5 Distribution of sensitivities of fertility to mortality measured by (d*F*/d*q*)/*F* change effects. NOTES: Only interpretable estimates are included. Mean is 3.40. All estimates are statistically significant. Only the two-stage least-squares estimates are used for Prussia.

distribution of (d*F*/d*q*)/*F* for changes is shown in Figure 6-5 which is based on data in Table 6-2. For changes, the mean of (d*F*/d*q*)/*F* is about 3.40.

It is worth considering the possible role of the lactation interruption effect in bringing about a positive causal effect of infant mortality on marital fertility. Preston (1978:8) gives a useful analysis of this question. He concludes that $(dF/dq)/F = (P2 - P1)/I$, where *P*1 is the average period of sterility following a live birth that results in an infant death, and *P*2 is the average period of sterility following a live birth that results in a surviving infant, while *I* is the overall average length of a birth interval. *P*1 and *P*2 will differ due to the lactation interruption effect and also because of the effect of practices and taboos such as prescription of intercourse while breastfeeding, or spending time following delivery with the mother's family. In the context of Europe, it is mainly the lactation interruption effect that is likely to have mattered.

Knodel (1978:25) provides a useful summary of evidence from European micro-level family reconstitution studies. The lactation interruption effect ($P2 - P1$) varies widely from 12 months to no months, apparently due to wide differences in breastfeeding practices. Average birth intervals (*I*) are about 30 months. The values of (d*F*/d*q*)/*F* will thus range from 0 to 0.4. We use our calculation of (d*F*/d*q*)/*F* from German data discussed above, and let $P2 - P1$ be on average about 6 months and I about 30 months, all reasonable averages for Prussia. Recall from above that the value of $(dF/dq)/F$ for complete offset is 2.44. The value of $(dF/dq)/F$ from lactation interruption alone is 0.2. Thus, lactation interruption accounts for only one-twelfth of the effect needed for full replacement. This is far too low to account for the size of response of fertility to infant mortality change that we observe in our data, and which has typically been found in studies of historical Europe.

Dynamic Stability

If the product of the estimated structural coefficients indicating the effect of fertility on infant mortality, and indicating the effect of infant mortality on fertility, is greater than unity, then the system they describe is dynamically unstable. In this case, an external shock that raised mortality, for example, would lead to an increase in fertility, which would lead to an increase in mortality larger than the first, and initiate an explosive spiral. One might take the view that because history has not been explosive, the product of the true coefficients must be less than unity. This would also imply that the first term in the bias expression given earlier is positive, and thereby remove one possible source of negative bias, making it more difficult to explain the direction of bias we have found in our estimates. However, one might equally take the view that in fact history has been explosive. Wolpin (in this volume) describes such a view, suggesting that an exogenous decrease in infant and child mortality may have set off a downward spiral of fertility and mortality. This would imply that the first term in the bias expression is negative, making it easier to explain the direction of bias we found. In our actual two-stage least-squares estimates, we find that this product in Prussian Kreise is 0.53 (Appendix tables 6-A3 and 6-A6) and in Prussian cities 0.95 (Appendix tables 6-A4 and 6-A7), suggesting dynamic stability in both areas, but with the cities bordering on instability.

Perhaps it is more interesting to examine the ultimate effect on fertility of an exogenous change in mortality, taking the coefficient estimates at face value. Let fertility F be a linear function of infant mortality q with coefficient b, and let q be a linear function of f with coefficient d. Let intercepts be a and c, respectively. Then c is the shifter to represent an exogenous change in q. We can solve for the equilibrium level of F, say F^*, after all changes have worked their way through the system. This is

$$F^* = (a + bc)/(1 - bd).$$

The estimated coefficient b gives the initial impact of q on F, while the factor $1/(1 - bd)$ accounts for the feedback cycle. With two-stage least-squares estimates, b is 1.03 in Kreise and 1.84 in cities, d is 0.51 in Kreise and 0.52 in cities, so that bd is 0.53 in the Kreise and 0.95 in the cities, as mentioned in the preceding paragraph. The multiplier $1/(1 - bd)$ in Kreise is about 2 and in the cities about 20. This suggests that the cities are very near to dynamic instability, whereas in the more rural areas the feedback is also important, but not so much as in the cities. With ordinary least-squares estimates the feedback multipliers are about 1.1 for both Kreis and cities, which seems more realistic.

We do not want to make too much of these results because our analysis lacks dynamic structure. The important point is that estimates of impacts do not tell the entire story. The full effects may be larger by a lot or by a little as the feedbacks play out.

CONCLUSIONS

According to theory, fertility and infant mortality should affect each other simultaneously, and these effects should be positive. A review and assessment of published research on European marital fertility and infant mortality suggest that there is a generally consistent, significant, and positive association between infant mortality level and marital fertility level. The evidence for this positive association of levels is stronger than has been realized. However, we argue that this association of levels is largely irrelevant for the policy issues of interest.

The evidence appears even stronger in studies that examine the association of changes in infant mortality with changes in marital fertility, which is more appropriate for examining secular fertility decline and most relevant for policy issues. In every case, changes in infant mortality are positively associated with changes in fertility and most are significant. However, there are important issues of causality that must be resolved before drawing conclusions. The few studies that attempted to disentangle the direction of causality using instrumental variables estimation found, as we did, that important causality was operating in both directions.

If we take these estimates of association at face value, then they are mostly far larger than could be explained by lactation interruption. They are also substantial relative to the size necessary to bring about a complete offset of fertility when mortality changes, leaving growth rates unchanged. Some are smaller than this offset level, some about equal to it, and some are larger than it. We are suspicious of estimates that indicate more than completely offsetting changes in fertility, which we find in our own two-stage least-squares estimates.

In general, parameter estimates using instrumental variables techniques are sensitive to the specific choice of instruments. Although we cannot be sure that the instruments we used in our estimation are the most appropriate, we believe they are the best given the data that are available. Furthermore, we cannot completely rule out the possibility that the estimated associations of fertility and mortality, even when using instrumental variables and fixed-effects methods, actually reflect a spurious association induced by unobserved variables that influence both fertility and mortality and that change over time. These variables might include breastfeeding, health conditions, nutrition, or unobserved aspects of economic development and modernization.

Nonetheless, we believe that the repeated estimation of positive associations, particularly with instrumental variables and fixed-effects models, likely does reflect a true and substantial effect of mortality change on fertility change. In the case of Prussia, we have been able to include an unusually extensive array of variables measuring differing aspects of socioeconomic change, and we still find strong positive effects of mortality change on change in fertility. Uncertainty arising from possible unobserved time-varying factors is a problem when making inferences from any time series analysis, not just this one, and there are corresponding problems with any cross-sectional analysis. On balance, then, we believe that there is substantial evidence that mortality decline was an important cause of fertility decline in Europe.

APPENDIX

Appendix tables begin on the following page.

TABLE 6-A1 Definitions of Variables Used in the Analysis

Variable	Definition
GMFR	General marital fertility rate (legitimate births per 1,000 married females 15-49).
Catholic	Catholics per 100 total population.
Slav	Slavic speakers per 100 total population.
Church	Employees in religious occupations per 100 population over age 20.
Education	Teaching employees per 100 population aged 6-13.
Health	Health employees per 100 total population.
FLFPR	Female labor force participation rate (employed females per 100 female population aged 20-69) (excludes agriculture and service).
Income	Average real income of male elementary school teachers in Deutsche marks as of 1900.
Mining	Mining employees per 100 employed persons.
Manufacturing	Manufacturing employees per 100 employed persons. Used only in the city model
Urban	Urban population per 100 total population. Used only in the Kreis model.
Bank	Banking employees per 100 population over age 20.
Insurance	Insurance employees per 100 population over age 20.
Communications	Post, telegraph, and railway employees per 100 population over age 20.
Population	Population, in thousands. Used only in the city model.
Infant mortality	Legitimate infant mortality rate (legitimate deaths under age one per 1,000 legitimate births).
Married sex ratio	Married males/married females.
Kreis born	Population born in Kreis per 100 total population. Used only in the Kreis model.
City born	Population born in city per 100 total population. Used only in the city model.
Sanitation	Cumulative municipal sanitation bond debt per capita in Deutsche marks. Available only for cities.
ASDR	Age-specific death rate for males aged 30-34. Data are available only for Regierungsbezirke.

NOTES: For details and sources see Galloway et al. (1994, 1995, 1996). Data are available quinquennially from 1875 to 1910. Vital registration variables are based on 5-year average centered around each quinquennial year. Stillbirths are excluded throughout. There are 36 Regierungsbezirke and 407 Kreise in Prussia. We also examined 54 cities. In general, the German occupational censuses do not lend themselves to calculation of economic sector variables because of a peculiar redefinition of female agricultural laborers that leads to an improbable 2 million increase in the category between 1895 and 1907 (Tipton, 1976:153-158). However, city populations were probably not affected by this problem because there were few agricultural workers in the cities. This is the reason the variable Manufacturing is available only in the cities. Mining is available for both Kreise and cities because virtually all miners were men.

TABLE 6-A2 Models Used in the Fertility Analysis

Variable	Kreis Fertility Model (equation (1))	City Fertility Model (equation (2))	Expected Sign
Dependent	GMFR	GMFR	
Independent	Catholic	Catholic	+
	Slav	Slav	+
	Church	Church	+
	Education	Education	–
	Health	Health	–
	FLFPR	FLFPR	–
	Income	Income	–
	Mining	Mining	+
		Manufacturing	+
	Urban		–
	Bank	Bank	–
	Insurance	Insurance	–
	Communications	Communications	–
		Population	–
	Infant mortality	Infant mortality	+
	Married sex ratio	Married sex ratio	+

NOTES: In equation (1) two-stage least-squares age specific death rate (ASDR) for males aged 30-34 is used as an instrument for Infant Mortality. In equation (2) two-stage least-squares Sanitation and ASDR are used as instruments for Infant Mortality.

TABLE 6-A3 Equation (1): Summary of Ordinary and Two-Stage Least-Squares Fertility Regression Results for Kreise in Prussia, 1875-1910 (dependent variable is GMFR)

Variable	Expected Sign	Level OLS	Level TSLS	Change OLS	Change TSLS
Constant		189.374**	258.026**		
Catholic	+	0.693**	0.716**	2.138**	−1.654**
Slav	+	0.348**	0.137	−0.283	−1.161**
Church	+	1.013	10.586	23.231**	38.721**
Education	.	−5.488‡	−5.816	−9.075**	−7.164**
Health	−	−32.456*	−42.767*	−6.596	24.430‡
FLFPR	−	−0.529‡	−2.296*	−1.235**	−1.153**
Income	−	−0.014	0.032	0.002	0.004
Mining	+	1.032**	1.251**	0.757**	0.430
Urban	−	0.034	−0.089	0.107	0.385‡
Bank		−36.020	1.850	−55.325**	−26.903‡
Insurance	−	29.251	−48.872	−133.466**	−66.125*
Communications	−	−0.453	5.708	−7.333**	−1.799
Infant mortality	+	−0.052‡	0.478‡	0.242**	1.028**
Married sex ratio	+	91.828‡	−107.824	40.674*	−57.928

NOTES: OLS, ordinary least squares. TSLS, two-stage least squares. The unit of analysis is the Kreise. The level regressions use averages of each variable over eight quinquennial periods from 1875 to 1910. The change regressions are fixed-effects models using data for eight quinquennial periods from 1875 to 1910. Estimates for the 407 Kreis dummy variables are omitted. In the level regressions $n = 407$ and OLS $R^2 = 0.681$. In the change regressions $n = 3{,}256$ and OLS $R^2 = 0.920$. **, *, ‡ indicate that the coefficient is statistically significant at the 1 percent, 5 percent, and 10 percent levels, respectively, two-tailed test. Age-specific death rate for males aged 30-34 is used as an instrument for Infant Mortality in the two-stage least-squares regressions for both level and change. The two-stage least-squares t statistics are based on the structural residuals (Hall et al., 1992:133-134) and are asymptotically correct. The ordinary least-squares results are discussed at length in Galloway et al. (1994).

TABLE 6-A4 Equation (2): Summary of Ordinary and Two-Stage Least-Squares Fertility Regression Results for Cities in Prussia, 1875-1910 (dependent variable is GMFR)

Variable	Expected Sign	Level		Change	
		OLS	TSLS	OLS	TSLS
Constant		422.365*	448.292*		
Catholic	+	0.581**	0.628**	–0.044	2.354‡
Slav	+	0.275	0.218	2.010*	-1.990
Church	+	25.989	18.053	–2.113	25.009
Education	–	–4.064	–4.589	–4.585**	–4.809‡
Health	–	–1.314	–1.618	7.693	96.812**
FLFPR	–	–1.264*	–0.976	–3.571**	–0.919
Income	–	0.004	–0.004	–0.022**	–0.030**
Mining	+	1.828**	1.616*	0.161	–1.221
Manufacturing	+	0.761*	0.468	0.357	–0.221
Bank	–	–44.682*	–53.788*	–11.139	54.861‡
Insurance	–	17.901	23.651	–43.001**	11.970
Communications	–	–5.276*	–6.122‡	–5.959*	–0.045
Population	–	–0.007	0.002	–0.041**	0.056
Infant mortality	+	–0.042	–0.187	0.337**	1.836**
Married sex ratio	+	–193.165	–163.150	354.038**	194.328

NOTES: OLS, ordinary least squares. TSLS, two-stage least squares. The unit of analysis is the city. The level regressions use averages of each variable over eight quinquennial periods from 1875 to 1910. The change regressions are fixed-effects models using data for eight quinquennial periods from 1875 to 1910. Estimates for the 54 city dummy variables are omitted. In level regressions n = 54 and OLS R^2 = 0.888. In the change regressions n = 432 and OLS R^2 = 0.896. **, *, ‡ indicate that coefficient is statistically significant at the 1 percent, 5 percent, and 10 percent levels, respectively, two-tailed test. Sanitation and age-specific death rate for males aged 30-34 are used as instruments for Infant Mortality in the two-stage least-squares regressions for both level and change. The two-stage least-squares t statistics are based on the structural residuals (Hall et al., 1992:133-134) and are asymptotically correct. The ordinary least-squares results are discussed at length in Galloway et al. (1994).

TABLE 6-A5 Models Used in the Infant Mortality Analysis

Variable	Infant Mortality Model: Kreis (equation (3))	Infant Mortality Model: City (equation (4))	Expected Sign
Dependent	Infant mortality	Infant mortality	
Independent	Catholic	Catholic	+
	Slav	Slav	+
	Education	Education	−
	Health	Health	−
	FLFPR	FLFPR	+
	Income	Income	−
	Urban		+
	Communications	Communications	−
		Population	+
	GMFR	GMFR	+
	Kreis born	City born	−
		Sanitation	−

NOTES: In equation (3) two-stage least squares, Church, Mining, Bank, Insurance, and Married sex ratio are used as instruments for GMFR. In equation (4) two-stage least squares, Church, Mining, Manufacturing, Bank, Insurance, and Married sex ratio are used as instruments for GMFR.

TABLE 6-A6 Equation (3): Summary of Ordinary and Two-Stage Least-Squares Infant Mortality Regression Results for Kreise in Prussia, 1875-1910 (dependent variable is infant mortality rate)

Variable	Expected Sign	Level		Change	
		OLS	TSLS	OLS	TSLS
Constant		656.315**	723.051**		
Catholic	+	0.325**	0.475	0.109	0.639**
Slav	+	0.087	0.132	0.898**	0.925**
Education	–	–15.705**	–18.071**	–0.483	1.892
Health	–	–7.386	–14.945	–43.048**	–35.017**
FLFPR	+	2.567**	2.364**	0.050	0.362‡
Income	–	–0.110**	–0.111**	–0.009**	–0.007*
Urban	+	–0.256	–0.286	0.366**	0.392**
Communications	–	–12.640**	–12.368**	–5.862**	–3.967**
GMFR	+	–0.304**	–0.503	0.297**	0.512**
Kreis born	–	–3.418**	–3.559**	–0.333**	4.10**

NOTES: OLS, ordinary least squares. TSLS, two-stage least squares. The unit of analysis is the Kreis. The level regressions use averages of each variable over eight quinquennial periods from 1875 to 1910. The change regressions are fixed-effects models using data for eight quinquennial periods from 1875 to 1910. Estimates for the 407 Kreis dummy variables are omitted. In the level regressions n = 407 and ordinary least-squares R^2 = 0.432. In the change regressions n = 3,256 and R^2 = 0.922. **, *, ‡ indicate that the coefficient is statistically significant at the 1 percent, 5 percent, and 10 percent levels, respectively, two-tailed test. Church, Mining, Bank, Insurance, and Married sex ratio are used as instruments for GMFR in the two-stage least-squares regressions for both level and change. The two-stage least-squares t statistics are based on the structural residuals (Hall et al., 1992:133-134) and are asympototically correct. These results are discussed at length in Galloway et al. (1996).

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TABLE 6-A7 Equation (4): Summary of Ordinary and Two-Stage Least-Squares ~~Fertility~~ Regression Results for Cities in Prussia, 1875-1910 (dependent variable is infant mortality rate)

Variable	Expected Sign	Level		Change	
		OLS	TSLS	OLS	TSLS
Constant		459.350**	455.027**		
Catholic	+	0.621‡	0.609	−1.345**	−1.296*
Slav	+	−0.737	−0.743	1.983**	1.287
Education	−	−0.082	0.104	0.207	1.954
Health	−	−51.765	−51.793	−58.008**	−52.045**
FLFPR	+	0.977	1.002	−0.547	0.614
Income	−	−0.073**	−0.073**	0.009*	0.014**
Communications	−	−7.184	−7.084	−7.551**	−3.649
Population	+	0.053*	0.053*	−0.053**	0.033‡
GMFR	+	−0.198	−0.183	0.247**	0.517**
City born	−	−2.215**	2.220**	−0.852*	−0.289
Sanitation	−	−0.906	−0.910	−0.434**	−0.287*

NOTES: OLS, ordinary least squares. TSLS, two-stage least squares. The unit of analysis is the city. The level regressions use averages of each variable over eight quinquennial periods from 1875 to 1910. The change regressions are fixed-effects models using data for eight quinquennial periods from 1875 to 1910. Estimates for the 54 city dummy variables are omitted. In the level regressions $n = 54$ and ordinary least-squares $R^2 = 0.412$. In the change regressions $n = 432$ and $R^2 = 0.903$. **, *, ‡ indicate that the coefficient is statistically significant at the 1 percent, 5 percent, and 10 percent levels, respectively, two-tailed test. Church, Mining, Manufacturing, Bank, Insurance, and Married sex ratio are used as instruments for GMFR in the two-level least-squares regressions for both level and change. The two-stage least-squares t statistics are based on the structural residuals (Hall et al., 1992:133-134) and are asympototically correct. These results are discussed at length in Galloway et al. (1996).

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