ECONOMIC GROWTH CENTER

YALE UNIVERSITY

Box 1987, Yale Station
New Haven, Connecticut

CENTER DISCUSSION PAPER NO. 284

FERTILITY AND CHILD MORTALITY OVER THE LIFE CYCLE: AGGREGATE AND INDIVIDUAL EVIDENCE

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June 1978

Notes: Center Discussion Papers are preliminary materials circulated to stimulate discussion and ciritical comment. References in publications to Discussion Papers should be cleared with the author to protect the tentative character of these papers.

This work was partly supported by AID Contract otr-1432 and was facilitated by a grant from the Rockefeller Foundation for research in Economic Demography. I have benefitted from the comments of E. A. Hanushek, S. Rosen, and R. Willis and the computational assistance of Ruth Daniel. Remaining errors are my own responsibility.

Mortality has been markedly reduced in the poorer countries of the world in the last three to four decades; existing data typically indicate life expectancy increasing about one-half year per calendar year during this period (see G. J. Stolnitz). This decline of mortality has contributed to an increase in intrinsic rates of population growth of almost 1 percent per year. Further, since the most dramatic changes are observed for infants and children, the age composition of these populations has changed, contributing today to a transitory increase in the rate of population growth of another one-half of 1 percent per year.

In this unprecedented period of rapid population growth, a natural question to ask is whether fertility responds to the decline in mortality, and if so, then by how much and how fast, and whether this tendency toward demographic equilibrium dampening the rate of population increase across countries is also evident across economic classes within countries. There are some indications that the reduction in mortality may have been concentrated in lower income groups and though this would be appropriately construed as an egalitarian development (see Simon Kuznets), it also raises the possibility that the rate of natural increase of the poor may thereby differentially increase, widening in the next generation already large personal income and wealth differences.

Rolling back such a fundamental constraint on human life leads to rearrangements. Family decision making is affected most directly, for the family deals with the periods of economic dependency in the life cycle when the force of mortality is most heavy. Other functions of the family may be affected to a lesser degree: the transmission to heirs of a cultural heritage and the skills for a livelihood. Documentation that mortality change has in fact modified behavior has only recently gotten underway (see for example, Rati Ram and Theodore W. Schultz). The purpose of this paper is to present some evidence on one form of household response to mortality.

The tie between mortality and fertility is a complex one. On the way from being a traditional society to becoming modern, mortality and then fertility are generally observed to decline. This process, called the "demographic transition," remains imprecisely characterized in terms of underlying mechanisms, time dimension and relative magnitudes of change. Conflicting empirical evidence adds to the conceptual ambiguity: cross-sectional data suggests fertility and child mortality are positively related in low income countries (see the author, 1976b), whereas aggregate time-series in these countries show crude death rates falling for three decades before crude birth rates widely decline (see Kuznets).

Changes in general mortality levels are associated across populations with monotonic variation in mortality rates by age. Discussion focuses here on the effect of survival of a mother's own children on her fertility, because the causal relation is direct and obvious, child deaths are a substantial proportion of all deaths, and offspring survival is measured, though unfortunately subject to errors of recall. Concurrent change in child mortality outside of the nuclear family and adult mortality undoubtedly reinforce the incentives to modify reproductive goals, but adult mortality cannot be readily observed in the family and it is difficult to obtain other good proxies for relevant mortality conditions.

The paper is organized as follows. Section I reviews how economists have argued that child mortality may affect reproductive behavior, and discusses the unresolved problems of separating biological, behavioral and simultaneous relations. Though the economic framework yields plausible results, refutable predictions are scarce and empirical work is in order to focus theoretical developments. Section II summarizes aggregate intercountry evidence of fertility responses within a birth cohort to variation in child mortality. Section III reports similar estimates within families based on household survey data for urban Latin America and rural India.

I. Child Mortality and Reproductive Behavior and Motivations

There are four interrelated reasons why one expects fertility and child mortality to be associated; fertility may respond to expected or experienced child (and adult) mortality; child mortality may be influenced by fertility and the proportion of childbearing occurring to women at high risk, for example, the very young, the old, and the poor; both mortality and fertility may be affected by common observable factors, such as education; and finally, both may also be influenced by unobserved factors that generate a correlation between disturbances in equations determining fertility and child mortality.

To distinguish the above sources of covariation, a system of two equations determining fertility and child mortality must be estimated, in all likelihood using simultaneous equation techniques. Without a priori theoretical insights for identifying the two underlying structural relations, or panel data of time-series for couples, it seems premature to extract system estimates from a single cross section identified artificially. For example, only a few percent of the observed variation in child mortality rates across rural Indian families can be traced to a host of standard economic, demographic, or social variables. Furthermore, for use in policy or projection, one wants an estimate of how fertility responds to mortality variation that stems from both economic conditions facing the family, such as those related to family expenditures on food, and those conditions that are exogenous to family resources and market prices, such as public health programs that eradicate smallpox or control malaria.

In estimating the response of reproductive behavior here, the regime of mortality faced by the family is treated in the most simple fashion--predetermined. Such an estimate undoubtedly neglects some systematic, though I suspect small, feedback effect of fertility on child mortality, holding constant several conditioning factors, and may also be slightly biased by residual simultaneous sources of variation in vital rates. Future work will impose identifying restrictions and test for independence of residuals across structural equations (see De-Min Wu).

Empirical estimates of the response of fertility to child mortality also embody both voluntary modification of behavior and involuntary biological processes that constrain reproductive potential. The survival of breast-fed infants can lengthen their mothers' period of sterility following their birth, and thereby delay subsequent births. Under extreme assumptions the biological reduction in births due to the reduction in infant deaths is less than one-third, whereas more realistic exposure parameters for Latin America would suggest a maximum biological response of one-tenth (see S. H. Preston, p. 13). If empirical estimates of the derivative of births with respect to child deaths exceed .1 or at most .2, the excess is likely to arise from voluntary response patterns.

Economists have only recently begun to describe how fertility goals might adjust to the sequential and partially expected incidence of child mortality. First, leaving aside uncertainty, knowledge that a particular fraction of offspring will die before reaching a specific mature age has two offsetting effects; it increases the cost per survivor, and increases the number of births required to obtain a survivor. If the desired number of survivors is insensitive to or inelastic with respect to their cost and costs are inversely proportional to survival probabilities, the number of births sought will vary directly with mortality. Alternatively, if parent demands for survivors are cost elastic, reducing the heavy costs of child mortality encourages parents to have more births. However, to speak of a "survivor" as the relevant metric for framing parental goals neglects changes in the composition of benefits and opportunity costs as a child ages and the dependence of this net benefit stream on family size and child spacing.

Moreover, the resources parents forego to have a child need not be the same across families or within families, even if the prices of relevant market inputs and opportunities for child labor are identical. If the resource intensity of childrearing is viewed by parents as in part a long-term investment in their offspring, an exogenous reduction in mortality encourages more intensive child investments, such as schooling, migration, and more health investments, probably as a substitute for additional children (see Donald O'Hara).

When uncertainty due to mortality is explicitly considered, issues of hedging, insurance, and risk aversion enter; parents may modify their reproductive target in response to uncertainty according to their preferences with respect to family size and the distribution of opportunity costs associated with an excess or shortfall in survivors (the author, 1969). Research measuring preferences and opportunity costs with respect to family size has progressed slowly and as yet has not dealt directly with how these measures interact with mortality in the determination of reproductive behavior (see for example, J.M. Roberts, R.F. Strand, and E. Burmeister; L. Coombs; P.H. Lindert). The most satisfactory treatment of uncertainty is that developed by Yoram Ben Porath and Finis Welch, though its empirical application has lagged.

Given the concentration of child mortality in the initial years of life and the relatively long period of childbearing, the need for parents to hedge against child mortality appears low. Sequential decision making permits parents, for the most part, to replace deceased infants rather than bear (or withhold) additional children as a hedge against expected but uncertain future mortality. Adjustment to declines in child mortality can therefore be largely accomplished _ex post_, if parents are able to refrain from replacing offspring who (unexpectedly) survive. This capacity to adjust fertility initially through _replacement_ rather than according to _expectations_ creates the potential for a short lag between mortality and fertility. In the longer run, as mortality expectations adjust, an entirely new age pattern of reproductive behavior and child investment may emerge, conforming to perceived benefits and costs of birth spacing, family portfolios, and life cycle investment schedules.

A decade may nonetheless elapse between the initial decline in child mortality before replacement fertility begins to decline in the family formation process. Even then, the change in reproductive behavior may go unnoticed because of the small numerical importance of births to women in their late 30's and 40's. At the aggregate level, changes in age composition induced by the age pattern of mortality declines first depress crude birth rates slightly below their age standardized path for about a decade, and then increase them notably thereafter.[1] Crude birth rates 25 years after the onset of the mortality decline may thus rise, even as age standardized birth rates subside. Finally, resource and price constraints facing households may also change, raising or lowering reproductive goals independently of mortality. At this time, therefore, crude birth and death rates cannot directly clarify how fertility is adapting at the family level to child mortality mainly because changes in age composition are often unobserved.

In sum, economic logic has not yet described in a refutable form how fertility should be related to child mortality in equilibrium. In addition, the inaccessibility and cost of sufficiently reliable and acceptable birth control methods could, in many parts of the world, introduce another indeterminant innovational lag. The balance of the paper assembles some evidence of the actual empirical relation.

II. Aggregate Evidence on Fertility-Child Mortality Relationships

Widely available aggregate data permit testing of only a few questions about the relation between fertility and mortality. The minimum of data needed to estimate replacement response is the number of children born to comparably situated women who experienced different rates of child mortality. There are large parts of the world where no census information of this nature exists, including the United States.

As an example of the type of research that can be conducted with Census data, it is possible to test if the number of living offspring of women of a specific age is roughly constant in a particular year, irregardless of child mortality. The requisite information is available for some eighty countries for rural/urban or total populations (N = 95). By observing that the average number of children alive per woman A, equals the number born alive C, times a survival rate p = 1 - (D/C), (where D = C - A), the logarithm of C can be regressed on the logarithm of 1/p and the calendar year, t, to which the data pertain as in:

$$(1) \quad \ln C_i = \alpha_o + \alpha_1 t_i + \alpha_2 \ln(1/(1- (D_i/C_i))) \qquad (i = 1,\ldots,95)$$

If estimates of α_2 equal one, within age groups of women, the cross-sectional variation in cumulative fertility is simply replacement.

For the estimated equations the R^2's range from .2 to .3. The point estimates of α_2 (followed by their standard errors) among older women, for whom replacement is more nearly complete, cluster in the vicinity of one: age 30-34, 1.10 (.26); 35-39, 1.06 (.26); 40-49, .98 (.25); 50 or more, .77 (.16) (see the author, 1976b, Table 8.3). The hypothesis that α_2 equals one cannot be rejected; however, these estimates are undoubtedly biased upwards (see the author, 1976b).

Other cross-sectional estimates of this and other models based on regional variation in cumulative fertility and mortality rates suggest fully half of the variation in fertility is offset by differences in child mortality (see the author, 1976b). Unfortunately, there are only nine developing countries with relevant census data for two points in time.[2] Although in no way representative, this handful of cases shows that fertility declines compensated for about one-half of the concurrent declines in cohort child mortality for women aged 40-49.

III. Individual Analysis of Fertility and Own Child Mortality

Individual data provides a richer test of the interaction between fertility and child mortality. Analysis deals here with several representative samples of women age 30-49, with one or more births whose husbands are present, drawn from three Latin American urban surveys from 1964 conducted by Centro Latino-Americano de Demografiá (CELADE), and a rural Indian survey from 1970 conducted by the National Council of Applied Economic Research (NCAER).[3] Various theories have stressed different factors influencing fertility: these are reviewed elsewhere (see the author, 1976a). With these data, it is possible to hold constant at least a few price, income, and origin variables in order to assess the partial association between cumulative fertility and cumulative child mortality. Specifically, the education of the husband is included as an income effect. The education of the wife captures not only an income but also a dominant substitution effect, which may be attributed to the opportunity cost of her time in child rearing. Permanent income cannot be measured identically across samples, in urban areas it is the logarithm of household monthly expenditures exclusive of housing, and in rural areas it is total family income. The income variable is not consistently associated with fertility, controlling for husband's education, but it is inversely related to child mortality in all samples. Age of wife is included as a control for life cycle or birth cohort differences. Migrant origins and duration of city residence are held constant in the Latin American samples, while the presence of village health and educational institutions are included as controls in the rural Indian sample to capture local access to health and schooling services. The number of deceased children is normalized as a fraction of those predicted based on the woman's number of births, the age pattern of fertility, and an appropriate life table.[4]

Table 1 shows the results from the regression of children ever born on the normalized child mortality rate and the aforementioned variables. The regression coefficient on mortality and its t-ratio are shown in column (1). Columns (6) and (7) provide means of the fertility and normalized child mortality variables; the derivative of births with respect to deaths is shown in column (3) (see fn. b. Table 1). In all thirteen samples the level of fertility is positively associated with child mortality, and in all cases except women 40-49 in Mexico City, the associations are statistically significant (10 percent level). The derivative of births with respect to child deaths ranges widely, however, from .8-1.4 in Rio de Janeiro, to .4-.8 in San José, to .2-.3 in Mexico City, and in rural India between .3 and .5.

Why should fertility responses to child mortality show such substantial variation across populations? I propose the hypothesis that couples react to their child mortality experience by changing their reproductive performance to the extent that they are aware of a general downtrend in mortality in their segment of society. The absolute levels of child mortality among women 40-49 are initially similar (not reported) in the three city samples. They declined by 10 percent in Mexico City in the thirteen years spanned by these data (approximately 1945-58), 30 percent in Rio, and 40 percent in San José (col. (7)). Fertility decreased little in Mexico across these birth cohorts, whereas it declined notably in Rio and San José, and offsetting individualistic responses of fertility to own child mortality were substantial in the latter cities but not in the former.

The Indian sample is divided into cultivators working their own land and landless rural laborers, but fertility, mortality and response coefficients are not notably different (5 percent level) between these subsamples. Mortality is greater in India than in urban Latin America, and the secular downtrend across cohorts is moderate after a drop in the early 1950's. The compensatory response derivative of fertility with respect to child mortality increases to about one-half among Indian women at the end of their childbearing years, age 40-49.

To explore how fertility responds to child mortality across economic classes within a society, two income groups are defined (based not on observed income which might be endogenous, but on an instrumental variable prediction of income or expenditures derived from husband's age, education, and origins). Column (2) reports the regression coefficient on a dummy variable interacted with the mortality variable for the lower income group and its t-ratio. The null hypothesis of coefficient equality is

Table 1
Association Between Cumulative Fertility and Child Mortality in Selected Samples of Households: Urban Latin America 1964 and Rural India 1970[1]

	Coefficient on Child Mortality in Fertility Equation (t ratio)		Derivative of Births wrt Child Deaths at Sample Means[2] (Sample Size)			Overall Variable Means (Standard Deviation)		Ratio of upper to Lower Income Class Variable Means	
	Overall	Class Difference	Overall	Upper Class	Lower Class	Fertility	Mortality	Fertility	Mortality
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Rio de Janeiro, Brazil, 1964									
30-34	.304 (3.94)	-.113 (.43)	.752 (315)	.899 (122)	.650 (193)	3.14 (1.92)	.488 (1.31)	.746	.250
35-39	.617 (4.78)	.295 (1.16)	1.35 (279)	1.68 (120)	1.17 (159)	3.21 (2.20)	.370 (.973)	.697	.686
40-49	.541 (4.95)	.832 (3.87)	.941 (325)	1.10* (124)	.825* (201)	3.76 (2.70)	.681 (1.32)	.768	.604
San José, Costa Rica, 1964									
30-34	.206 (2.33)	.0554 (.19)	.561 (268)	2.95 (99)	.508 (169)	4.03 (2.06)	.539 (1.37)	.761	.221
35-39	.429 (3.67)	-.341 (1.15)	.827 (240)	1.18 (80)	.771 (160)	5.04 (2.90)	.867 (1.54)	.668	.446
40-49	.212 (1.85)	-.462 (1.42)	.397 (287)	.510 (81)	.362 (206)	5.35 (3.24)	.936 (1.64)	.709	.398
Mexico City, D.F., Mexico, 1964									
30-34	.123 (1.55)	.357 (2.30)	.252 (2.95)	.322* (116)	221* (179)	4.48 (2.27)	.939 (1.60)	.718	.635
35-39	.112 (1.97)	.0666 (.78)	.184 (239)	.217 (98)	.166 (141)	5.43 (2.86)	.925 (1.57)	.799	.438
40-49	.121 (1.19)	-.218 (1.02)	.174 (348)	.224 (136)	.152 (212)	5.57 (3.27)	1.07 (1.70)	.696	.447
India Rural ARIS, 1970									
30-39									
Land Owners	.265 (3.08)	-.0897 (.55)	.246 (832)	.240 (376)	.252 (456)	4.22 (1.80)	.351 (.708)	1.06	1.09
Landless	.408 (2.97)	-.920 (2.79)	.382 (349)	.344* (133)	.407* (216)	4.17 (1.82)	.295 (.673)	1.17	.703
40-49									
Land Owners	.768 (5.37)	-.415 (1.40)	.523 (621)	.502 (324)	.549 (297)	5.04 (2.15)	.320 (.582)	1.10	.634
Landless	.782 (4.36)	-.0163 (.97)	.539 (277)	.518 (134)	.559 (143)	4.88 (2.22)	.426 (.734)	1.29	.774

Notes: *If distinct regression coefficients are estimated for the two income classes in those cases where the t ratio exceeds 1.6, the derivative of births with respect to child deaths is .0663 and 1.25 for upper and lower class samples age 40-49 in Rio, -.231 and .474 for age 30-34 in Mexico City, and 1.13 and .209 for age 30-39 in the rural Indian landless sample, by income class.

[1] Derivation of columns described in text. Underlying regression results obtainable from author.

[2] Derivative $dC/dD = \beta(\widetilde{C/D})/(\widetilde{C}+\beta(\widetilde{C/D})(\overline{D/C}))$ where β is the regression coefficient of the normalized child death rate, $(\widetilde{C/D})$ is the reciprocal of the sample mean expected child death rate, $\overline{C}$ and $(\overline{D/C})$ the sample means of fertility and actual child mortality.

rejected (10 percent level) in only three out of thirteen cases, suggesting a common estimated coefficient may exist for both income classes. It should be stressed that given the small size of these samples and the relative infrequency of child mortality, little confidence can be placed on differences between these groups, except as a source of working hypotheses.

When derivatives are evaluated at subsample means using the common regression coefficient across income groups, the response patterns are similar for India but quite different in Latin America (cols. (4) and (5), largely because the urban fertility and child mortality are substantially lower at higher income levels. In addition, the large declines in child mortality in Rio and San José that were noted earlier appear to have most benefited the upper income classes (col. (9)). Consistent with my hypothesis, the derivative response of fertility with respect to mortality is also larger for these upper income classes. Recall also that household income and child mortality is negatively correlated in all samples. This may indicate that the widely accepted view that economic development plays only a minor role in the remarkable mortality transition in low income countries (see Stolnitz) needs reevaluation. Analysis is needed on how improvements in the economic environment of the family and its behavior influence prospects for child survival, and in turn impinge on fertility and other forms of human and physical capital investment within the family. In other words, the next step is to estimate the full structural equation model determining both fertility and child mortality at the family level.

IV. Conclusions

Across samples of urban and rural households in Latin America and India, statistically significant associations are reported between cumulative fertility and cumulative child mortality, holding constant age, education, income, and origins. Individual reproductive responses to child mortality increase to fully compensating levels only in those populations where child survival has markedly improved. Aggregate trends as well as individual child survival experience should be examined jointly in future efforts to understand individual reproductive behavior in low income countries. The economic determinants and consequences of mortality now warrant more study, given the magnitude of recorded change in life expectancy and our nearly complete ignorance of who has benefited by this significant process during economic development and why.

Footnotes

[1]The reversal in crude and age standardized fertility trends are seen in several countries of Latin America and Asia in the mid 1960s, where age specific vital rates are relatively reliable. Crude death rates, of course, fall much faster than age standardized mortality measures, increasing population growth rates beyond sustainable levels in the first several decades of mortality decline. Hence, the paradox that crude death rates in countries like Taiwan and Puerto Rico are about half of U.S. levels today.

[2]Data from U.N. Demographic Yearbooks available repeatedly for Bermuda, Brazil, Cyprus, Fiji, Gilbert and Ellice Islands, Hong Kong, Indonesia, Philippines and the Solomon Islands.

[3]Survey data was kindly provided by two research centers. The Centro Latino-americano de Demografiá, Santiago, Chile, coordinated under Carmen Miro's direction a series of urban comparative fertility surveys each of about 2000 women of childbearing age. The Indian Additional Rural Income Survey of about 6000 rural households was directed by M.T.R. Sarma during 1968-71 at the National Council of Applied Economic Research, New Delhi.

[4]Coale and Trussel (1974) birth schedules by age are first scaled down to yield the cumulative fertility reported by each woman in the survey. This imputed flow of births is then subjected to national age specific mortality rates to obtain an expected number of child deaths. Costa Rican and Mexican life tables are for 1966, the Indian 1961, and lacking a Brazilian table, the one for Colombia in 1965 was substituted as reasonable for health conditions in Rio de Janeiro (Keyfitz and Flieger, 1968).

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DATA APPENDIX

Tables A-1 through A-4 report the regressions on children ever born and Tables A-5 and A-6 present descriptive statistics on the urban Latin American and rural Indian samples considered here. The underlying regression results warrant a brief set of comments, aside from the coefficients on child mortality and the implied response elasticities that are discussed in the text of this paper.

Education of the wife is associated with diminished fertility in the great majority of urban Latin American samples, i.e., the joint F ratio test of the statistical significance of the set of dummy variables satisfied the 5% level or better. The husband's education is also inversely associated with fertility in Latin America, holding constant for the unstable but generally positive effect of household permanent income (monthly expenditure level) on fertility. Particularly in the older cohort, age 40-49, when childbearing is complete, and the effect of education delaying marriage and the initiation of childbearing has diminished in magnitude, household income has a positive impact on fertility in both Costa Rica and Mexico City. In rural India, family income again has a positive effect on fertility, as do other indicators of wealth, such as the amount of land owned among the landowning classes, but higher fertility is noted for couples where the husband has obtained or exceeded the primary level of schooling. The joint F test for the set of male education dummy variables is statistically significant for all four samples, whereas the inverse effect of the wife's education on fertility is statistically significant only for the landowning samples, and notably only for wives who are 30 to 39 and may not yet have completed their childbearing.

Other variables do not contribute much to the explanation of fertility in these sample. The rural origin of the wife (or husband--not reported) in the Latin American samples helps to explain somewhat higher fertility among rural migrant women 40-49 in Rio de Janeiro, but is otherwise insignificant in the other groups and occasionally of the wrong sign. Residence in an Indian village that has an educational institution is associated with slightly lower fertility among landless wives age 30-39, whereas this variable obtains the opposite sign in the other three groups. The provision of schooling does not promote a substitution of schooling for family size, nor does it clearly subsidize fertility directly. The availability of a local health center is positively associated with fertility, but is statistically significant only for the younger wives.

As indicated in the text of the paper, the strongest and most persistent correlate of fertility is child mortality. As discussed in the paper, it would seem proper to view child mortality as itself being influenced by the resources available to the household and the allocative and behavioral decisions made by household members. Exploratory studies were undertaken, therefore, to determine how much of the cross-family variation in child mortality could be accounted for by economic variables. In the case of the urban Latin American samples a strong inverse association is noted between child mortality and household expenditure level, suggesting that regardless of migrant status, rural origins and parent education, the current availability of market resources in the household is a strong predictor of the couple's past experience of child mortality. This is not true for rural India, where there is little evidence that family income or landholdings are associated inversely with child mortality. Aside from differences associated with regions in India (climate?), little of the variance in child mortality is explained by the regressions. It is concluded that variables that might logically have identified a structural equation model accounting for both fertility and child mortality simultaneously are not an adequate basis for restricting and estimating the model. For example, if region and the local availability of a health center are used to identify the child mortality variable in the Indian fertility equation, (i.e., these two variables do not directly enter the fertility equation), the instrumental variable estimate of the coefficient on child mortality is generally positive but unstable and statistically insignificant by conventional standards. Although household income had a more noted association with child mortality in urban Latin America that is opposite in sign from that between income and fertility, the structural equation estimates for the Latin American were no more satisfactory than were those for rural India where the simple explanatory power of the economic variables is lower.

Variation across families in child mortality that led to responsive changes in fertility is largely unassociated with measured socioeconomic characteristics of the village or household members. Though probably not entirely random, our existing knowledge of the systematic factors influencing child mortality is limited, and is probably unable to support at this time a structural equation approach to the problem. Consequently, the estimates reported in the paper treat child mortality as predetermined from the point of view of the family's fertility decisionmaking process. These results are offered not as a satisfactory conceptual approach to the problem, but as a working hypothesis that probably yields a reasonable empirical first approximation for the responsiveness of fertility to cross sectional variation in child mortality.

Table A-1

Regressions on Children Ever Born by Age of Mother: Urban Latin America, 1964

(beneath regression coefficients are the absolute value of t ratios)

Independent Variables	Rio de Janeiro, Brazil			San José, Costa Rica			Mexico City, DF, Mexico		
	30-34	35-39	40-49	30-34	35-39	40-49	30-34	35-39	40-49
=1 if wife's birthplace is rural	.260	.564	.342	-.136	-.483	.643	.0419	.474	.153
	(.46)	(.79)	(3.49)	(.47)	(1.17)	(1.48)	(.14)	(1.06)	(.39)
No education, wife	.899	.364	.332	1.17	1.56	.305	.521	1.27	1.86
	(2.19)	(.62)	(.54)	(1.36)	(1.57)	(.25)	(1.06)	(1.73)	(2.70)
Did not complete primary education, wife	.340	.326	.661	.0070	.414	2.11	.871	.672	.598
	(1.23)	(.94)	(1.76)	(.02)	(.85)	(4.36)	(2.51)	(1.34)	(1.26)
Secondary education, wife	-.538	-.206	-.158	-.757	-.897	-.0459	-.277	-.381	-1.78
	(1.84)	(.53)	(.38)	(1.99)	(1.66)	(.08)	(.74)	(.71)	(3.52)
University education, wife	.252	-1.31	-.489	-.887	-1.47	1.25	-.715	.582	-1.23
	(.30)	(1.30)	(.51)	(1.60)	(1.62)	(1.46)	(.90)	(.55)	(1.42)
No education, husband	1.65	1.34	1.42	1.97	1.94	.0699	.916	.313	.177
	(3.59)	(2.36)	(2.32)	(2.38)	(.99)	(.05)	(1.73)	(.40)	(.27)
Did not complete primary education, husband	.651	.153	.450	.473	.272	1.05	.681	.527	1.08
	(2.21)	(.42)	(1.09)	(1.49)	(.58)	(2.32)	(1.95)	(1.01)	(2.34)
Secondary education, husband	.180	-.371	-.267	-.365	-1.26	-.320	-.249	-.436	-.622
	(.65)	(1.02)	(.68)	(1.03)	(2.30)	(.60)	(.65)	(.76)	(1.21)
University education, husband	-.399	-.351	-.0539	-.376	-.780	-.498	-.730	-1.06	-1.47
	(.82)	(.69)	(.09)	(.83)	(1.10)	(.68)	(1.58)	(1.82)	(2.57)
Age of wife	.143	.039	-.0199	.261	-.0324	-.0438	.243	.307	-.0713
	(2.04)	(.43)	(.43)	(3.19)	(.26)	(.74)	(2.87)	(2.49)	(1.34)
Log of expenditures (monthly)	.339	-.0120	.206	.0758	.336	.917	.0801	.248	1.05
	(1.63)	(.08)	(.79)	(.38)	(1.08)	(2.89)	(.37)	(1.37)	(3.92)
Child mortality, normalized	.304	.617	.541	.206	.429	.212	.123	.112	.121
	(3.94)	(4.78)	(4.95)	(2.33)	(3.67)	(1.85)	(1.55)	(.97)	(1.19)
Intercept	-5.62	1.63	1.61	-4.73	4.08	-.593	4.34	-8.13	.987
	(1.74)	(1.44)	(.46)	(1.70)	(.84)	(.18)	(1.44)	(1.71)	(.33)
R^2	.2134	.1891	.1969	.2070	.1874	.2138	.2015	.1524	.2150
S.E.E.	(1.740)	(2.028)	(2.471)	(1.878)	(2.687)	(2.944)	(2.073)	(2.703)	(2.956)
Sample size	315	279	325	268	240	287	295	238	348

Table A-2

Regressions on Children Ever Born
by Age of Mother and Landowning Status:
Rural India, 1970
(beneath regression coefficients are the absolute value of t ratios)

	30-39		40-49	
	Land Owners	Landless	Land Owners	Landless
Primary education, wife	-.0502 (.21)	-.444 (1.52)	-.693 (1.57)	-.0610 (.12)
Above primary, but not matriculate, wife	-.311 (1.02)	-.680 (1.45)	-.710 (1.23)	-.491 (.71)
Matriculate, wife	-1.65 (3.39)	-.616 (1.07)	-1.12 (1.51)	-.482 (.48)
Primary education, husband	.464 (3.00)	.662 (2.70)	1.06 (4.39)	.814 (2.21)
Above primary but not matriculate, husband	.494 (3.08)	1.16 (4.14)	.982 (4.49)	.685 (1.80)
Matriculate, husband	.331 (1.56)	.628 (1.83)	.752 (2.83)	.773 (1.89)
=1 if educational institution exists in village	.0695 (.32)	-.708 (1.66)	.212 (.68)	.330 (.56)
=1 if health center exists in village	.409 (2.69)	-.0697 (.32)	.423 (1.99)	.390 (1.27)
Age of wife	.149 (7.32)	.108 (3.27)	.0151 (.55)	.012 (.28)
Child mortality, normalized	.265 (3.10)	.408 (2.97)	.768 (5.37)	.782 (4.36)
Family income (in 10,000 Rs)	.392 (3.01)	.661 (2.49)	.697 (3.38)	.478 (1.23)
Intercept	-1.48 (2.01)	.478 (.40)	2.87 (2.37)	2.89 (1.51)
R^2 S.E.E.	.1166 (1.702)	.1524 (1.708)	.1185 (2.041)	.1058 (2.150)
Sample size	832	349	621	277

Table A-3

Regressions on Children Ever Born with Child Mortality by Income Class by Age of Mother: Urban Latin America, 1964

(beneath regression coefficients are the absolute value of t ratios)

Independent Variables	Rio de Janeiro, Brazil			San José, Costa Rica			Mexico City, DF, Mexico		
	30-34	35-39	40-49	30-34	35-39	40-49	30-34	35-39	40-49
=1 if wife's birthplace is rural	.258	.549	3.29	-.140	-.488	.649	-.0092	.456	.175
	(.46)	(77)	(3.44)	(.48)	(1.18)	(1.50)	(.03)	(1.01)	(.44)
No education, wife	.902	.334	.322	1.17	1.56	.419	.494	1.28	1.88
	(2.19)	(.56)	(.53)	(1.36)	(1.57)	(.34)	(1.01)	(1.73)	(2.37)
Did not complete primary education, wife	.343	.288	.707	.0036	.347	2.12	.903	.673	.584
	(1.23)	(.83)	(1.92)	(.01)	(.71)	(4.39)	(2.62)	(1.34)	(1.23)
Secondary education, wife	-.530	-.218	-.317	-.763	-.882	-.0598	-.227	-.364	-1.79
	(1.81)	(.56)	(.77)	(2.00)	(1.63)	(.10)	(.61)	(.68)	(3.55)
University education, wife	.259	-1.11	-.769	-.883	-1.50	-1.17	-.691	.574	-1.25
	(.31)	(1.09)	(.83)	(1.59)	(1.66)	(1.36)	(.87)	(.54)	(1.44)
No education, husband	1.65	1.28	1.12	1.96	1.91	.104	.845	.309	.219
	(3.59)	(2.24)	(1.84)	(2.36)	(.98)	(.08)	(1.60)	(.40)	(.33)
Did not complete primary education, husband	.650	.139	.398	.469	.304	1.09	.659	.532	1.10
	(2.20)	(.39)	(.99)	(1.48)	(.65)	(2.41)	(1.90)	(1.02)	(2.37)
Secondary education, husband	.151	-.337	.221	-.352	-1.52	-.595	.0035	-.422	-.804
	(53)	(93)	(.55)	(.98)	(2.57)	(1.05)	(.01)	(.73)	(1.48)
University education, husband	-.402	-.298	.371	-.370	-.917	-.674	-.486	-1.02	-1.58
	(83)	(.58)	(.66)	(.81)	(1.28)	(.90)	(1.03)	(1.71)	(2.71)
Age of wife	.143	.0352	-.0126	.262	-.0233	-.0504	.245	.305	-.0691
	(2.03)	(.39)	(.28)	(3.19)	(.19)	(.85)	(2.91)	(2.46)	(1.30)
Log of expenditures (monthly)	.331	-.0019	.282	.0718	.327	.950	.114	.251	1.05
	(1.58)	(.01)	(1.11)	(.36)	(1.05)	(3.00)	(.53)	(1.38)	(3.94)
Child mortality, normalized	.406	.453	.0303	.156	.705	.616	.0850	.0679	.281
	(1.64)	(2.37)	(.18)	(.57)	(2.63)	(2.01)	(.71)	(.35)	(1.51)
Child mortality, normalized for lower class	-.113	.295	.832	.0555	-.341	-.462	.357	.0666	-.218
	(.43)	(1.16)	(3.88)	(.19)	(1.15)	(1.42)	(2.30)	(.28)	(1.02)
Intercept	-5.52	1.64	.248	-4.73	3.885	-.492	-4.78	-8.09	.899
	(1.70)	(.45)	(.07)	(1.69)	(80)	(.15)	(1.59)	(1.70)	(.30)
R^2	.2139	.1931	.2339	.2071	.1921	.2196	.2162	.1527	.2175
S.E.E.	(1.742)	(2.026)	(2.417)	(1.881)	(2.686)	(2.939)	(2.058)	(2.709)	(2.956)

Table A-4

Regressions on Children Ever Born with Child Mortality by Income Class by Age of Mother and Landowning Status: Rural India, 1970

(beneath regression coefficients are the absolute value of t ratios)

	30-39		40-49	
	Land Owners	Landless	Land Owners	Landless
Primary education, wife	-.0499 (.21)	-.397 (1.37)	-.679 (1.53)	-.0580 (.12)
Above primary but not matriculate, wife	-.312 (1.02)	-.682 (1.47)	-.647 (1.12)	-.488 (.70)
Matriculate, wife	-1.65 (3.38)	-.557 (.97)	-1.07 (1.44)	-.481 (.48)
Primary education, husband	.452 (2.90)	.637 (2.63)	1.06 (4.38)	.813 (2.20)
Above primary, but not matriculate, husband	.464 (2.74)	.852 (2.86)	.853 (3.59)	.677 (1.59)
Matriculate, husband	.306 (1.41)	.381 (1.09)	.638 (2.29)	.765 (1.72)
=1 if educational institution exists in village	.0732 (.34)	-.713 (1.69)	.228 (.73)	.329 (.56)
=1 if health center exists in village	.416 (2.72)	-.0158 (.07)	.424 (1.99)	.389 (1.26)
Age of wife	.150 (7.33)	.0971 (2.95)	.0127 (.46)	.0122 (.29)
Child mortality, normalized	.318 (2.48)	1.13 (3.87)	1.04 (4.29)	.793 (2.65)
Child mortality, normalized for lower class	-.0897 (.55)	-.920 (2.79)	-.415 (1.40)	-.0163 (.04)
Family Income (in 10,000 Rs)	.389 (2.98)	.694 (2.64)	.696 (3.35)	.479 (1.23)
Intercept	-1.52 (2.05)	.895 (.75)	3.01 (2.49)	2.89 (1.50)
R^2 S.E.E.	.1169 (1.703)	.1716 (1.691)	.1213 (2.039)	.1058 (2.151)

Table A-5a

Means and Standard Deviations of Variables

From Urban Latin American Surveys - 1964

City/Variable Name	Age of Wife				
	25-29	30-34	35-39	40-49	25-49
I. San Juan, Costa Rica:					
Children Ever Born	3.16 (1.69)	4.03 (2.06)	5.04 (2.90)	5.36 (3.26)	4.42 (2.71)
Child Death Rate	.0325 (.0988)	.0478 (.121)	.0788 (.140)	.0897 (.158)	.0627 (.134)
Normalized Death Rate	.382 (1.173)	.539 (1.368)	.867 (1.539)	.927 (1.636)	.682 (1.46)
Log of Expenditures (monthly)	6.47 (.741)	6.58 (.759)	6.68 (.777)	6.83 (.720)	6.65 (.761)
Husband's Schooling (in years)	7.24 (4.08)	7.02 (4.13)	7.02 (4.47)	6.53 (3.97)	6.94 (4.17)
Wife's Schooling (in years)	6.36 (3.50)	6.15 (3.64)	5.75 (3.39)	5.64 (3.35)	5.97 (3.48)
Age at Marriage	19.04 (3.05)	20.12 (3.79)	20.14 (3.98)	21.34 (4.82)	20.20 (4.07)
Normalized Marital Fertility Rate	.917 (.367)	.805 (.348)	.721 (.356)	.679 (.357)	.778 (.368)
Sample Size	251	268	240	284	1043

Table A-5b

City/Variable Name	Age of Wife				
	25-29	30-34	35-39	40-49	25-49
II. Rio de Janiero, Brazil:					
Children Ever Born	2.42	3.14	3.21	3.74	3.16
	(1.36)	(1.92)	(2.20)	(2.70)	(2.18)
Child Death Rate	.0352	.0619	.0493	.0980	.0629
	(.106)	(.162)	(.129)	(.189)	(.154)
Normalized Death Rate	.286	.488	.370	.680	.468
	(.868)	(1.31)	(.973)	(1.32)	(1.16)
Log of Expenditures (monthly)	10.70	10.76	10.87	11.01	10.84
	(.683)	(.602)	(.912)	(.669)	(.731)
Husband's Schooling (in years)	7.63	7.05	7.27	7.78	7.43
	(4.54)	(4.59)	(4.90)	(4.96)	(4.77)
Wife's Schooling (in years)	6.52	5.77	5.86	5.72	5.94
	(3.92)	(3.63)	(3.73)	(3.70)	(3.75)
Age at Marriage	20.37	21.09	21.53	20.99	21.01
	(2.89)	(3.76)	(4.29)	(4.53)	(3.97)
Normalized Marital Fertility Rate	.858	.690	.539	.475	.632
	(.422)	(.346)	(.341)	(.326)	(.386)
Sample Size	259	315	279	321	1174

Table A-5c

City/Variable Name	Age of Wife 25-29	30-34	35-39	40-49	25-49
III. Mexico City, Mexico:					
Children Ever Born	3.71 (1.70)	4.48 (2.27)	5.45 (2.85)	5.57 (3.27)	4.79 (2.71)
Child Death Rate	.0611 (.132)	.0997 (.169)	.0988 (.159)	.130 (.207)	.0982 (.173)
Normalized Death Rate	.598 (1.30)	.939 (1.60)	.887 (1.46)	.1065 (1.70)	.875 (1.54)
Log of Expenditures (monthly)	6.85 (.870)	6.97 (.801)	7.00 (1.11)	7.18 (.702)	7.01 (.896)
Husband's Schooling (in years)	6.60 (4.84)	6.96 (4.96)	7.09 (5.16)	6.85 (5.01)	6.86 (5.00)
Wife's Schooling (in years)	5.36 (3.49)	5.16 (3.48)	5.15 (3.46)	5.13 (3.94)	5.20 (3.62)
Age at Marriage	18.82 (3.09)	20.30 (4.07)	19.58 (4.04)	20.04 (4.75)	19.69 (4.13)
Normalized Marital Fertility Rate	1.057 (.390)	.934 (.416)	.768 (.368)	.673 (.372)	.858 (.416)
Sample Size	316	295	238	348	1197

Table A-6

Means and Standard Deviations of Variables
From Rural India Survey, 1970

	Age 30-39		Age 40-49	
Variable Name	Land Owners	Landless	Land Owners	Landless
Children Ever Born	4.22 (1.80)	4.17 (1.82)	5.04 (2.15)	4.88 (2.22)
Child Death Rate	.0877 (.176)	.0737 (.167)	.0889 (.161)	.118 (.203)
Normalized Death Rate	.351 (.708)	.295 (.673)	.320 (.582)	.426 (.734)
Family Income (in 10,000 Rs)	.557 (.469)	.363 (.388)	.512 (.417)	.382 (.368)
Husband's Schooling (in years)	4.64 (4.81)	4.87 (5.00)	6.16 (4.96)	5.82 (5.13)
Wife's Schooling (in years)	.832 (2.53)	1.25 (2.99)	.481 (2.03)	.895 (2.56)
Age at Marriage	17.2 (3.24)	17.6 (3.36)	17.5 (3.55)	18.4 (3.73)
Normalized Marital Fertility Rate	.608 (.243)	.615 (.245)	.541 (.266)	.539 (.234)
Sample Size	832	349	621	277