Human Capital, Family Planning, and Their Effects on Population Growth

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Statistical evidence at the household level suggests that fertility and child mortality are related to factors specified by economic models of family resource allocation and behavior. Several of these factors, such as women's education and family planning, appear to decrease both fertility and child mortality, leaving in doubt what net effects these human-capital and social-welfare programs have had on the recent slowing of world population growth. The specific factors considered here include family-planning programs, human-capital endowments of women and men, natural resource wealth, other economic structural determinants of the costs and benefits of children, and the availability of nutritional inputs. Cross-sectional relationships are reported, and changes within countries are analyzed with fixed-effect methods using data for 68 low-income countries for the last two decades.

I. The Problem and Issues for Analysis

The acceleration in population growth in the last 300 years is due to mortality declining faster than fertility. The origin of this historic improvement in mortality is uncertain, but it is reasonably well documented in Europe from about 1700, and some would argue that it is linked to technical changes in agriculture which were followed by structural adjustments and new opportunities provided by industrialization. In this century, the mortality decline spread to the lower-income regions of the world, causing what has become known as the "population explosion." Since about 1960, total fertility rates have declined by one-half in Latin America and Asia.[1] Only reduced-form relationships for fertility, mortality, and population growth are estimated in this paper. A limitation of this reduced-form approach is that I am unable to distinguish the endogeneous mechanisms through which exogeneous factors affect fertility (e.g., delay of marriage or reduced marital fertility rates) or child mortality (e.g., immunizations or longer intervals between births), nor can I estimate the effect of mortality on fertility, or vice versa. To facilitate a decomposition of conditioning variables on population growth operating through fertility and mortality, I will select suitable functional forms for the dependent variables and age-standardize these vital rates to approximate cohort concepts of fertility and child mortality.

II. Measures of Vital Rates to Facilitate Decomposition

Fertility, mortality, and population growth can be either measured as crude rates or measured as age- and sex-standardized rates. The dependent variables for the first decomposition are the crude birth rate (B), crude death rate (D), and natural rate of population increase (P), all expressed per

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[1]This rapid change in reproductive behavior is hypothesized to be a response to the prior improvement in child survival. To identify statistically the historical impact of child mortality on fertility requires a better understanding of the factors originally triggering the decline in child mortality than is currently at hand. Elsewhere (Schultz, 1993), I have statistically identified the structural fertility equation by assuming that calorie availability affects child mortality but does not directly affect fertility. When child mortality is endogenized in cross-country regressions, the estimated effect of child mortality on fertility increases several-fold and remains highly significant statistically.

thousand population per year:

$$P = B - D. \tag{1}$$

Each vital rate is regressed on the same vector of reduced form (exogeneous) determinants, and consequently the partial derivative of any determinant on population growth can be linearly decomposed into its effects on birth rates minus its effect on death rates. However, because birth and death rates vary substantially by age, crude birth and death rates are sensitive to the age composition of the population, which is itself largely a function of past age-specific schedules of fertility and mortality. By assigning equal weight to the age-specific birth rates for women and age-specific death rates for children, the total fertility rate, F, and the child survival rate, S, are defined, ideally from birth to the mean age of childbearing or about age 26. The product of these variables is a net reproduction rate, $N = F \times S$, for a cohort. These three demographic rates in logarithmic form are then the dependent variables for the second decomposition of long-run cohort population growth as a sum of age-standardized fertility and child survival:

$$\ln(N) = \ln(F) + \ln(S). \tag{2}$$

I shall consider child survival only to the age of five, because in many low-income countries age-specific mortality is measured with less error before age five, and about three-quarters of mortality before age 26 occurs before the fifth birthday. Later child and adult mortality is moreover highly correlated (0.95 or higher) with the earlier preschool child mortality (A. J. Coale and Paul Demeny, 1965).

III. Hypotheses and Measurement of Explanatory Variables

When models of individual behavior are estimated for large aggregates (such as countries) from variation in average behavior and averaged conditioning variables, many tenuous aggregation assumptions must be maintained. In this case, moreover, nonlinear relationships are expected, which implies that higher moments than the means are needed within countries to describe expected values, and these distributional parameters are not generally known.

A static model of the lifetime demand for surviving children has been outlined on other occasions and generates a number of empirical predictions (Gary S. Becker, 1960; Schultz, 1981): education of women raises the cost of childbearing and reduces fertility; education of men may increase or decrease fertility, but in either case it will reduce fertility less (algebraically) than will the education of women; national income per adult from physical capital and natural resources is expected to increase the demand for children, assuming children are a normal good. It is also widely conjectured that the net cost of childrearing is greater for parents residing in urban than in rural areas, and the opportunities for parents to monitor their children in productive work is greater in an agricultural setting than in a nonagricultural one. The costs of vocational training of children also appear greater for parents in urban than in rural environments, because training is more likely to occur within family enterprises in rural areas. With regard to mortality, nutrition, summarized by available calories, is associated with decreases in mortality, especially at very low levels of income (Robert W. Fogel, 1990; John Strauss, 1993). The most important predictor of reduced child mortality in both rural and urban areas is the formal education of the mother. Agricultural employment is also associated with higher child mortality, perhaps because of the interaction of animal and human diseases and the greater time cost of obtaining modern medical care on the farm (Schultz, 1981).

There are alternative measures for several of these determinants of fertility, mortality, and population growth. Estimates of non-human-capital stocks are not widely available for low-income countries. I have therefore used net exports of mineral fuels as a share of GDP (Robert Summers and Alan Heston, 1991) to approximate natural-resource rents.[2] Calories are from three FAO country surveys. Education is the average years of education of men and women over the age of 15 as estimated by The World Bank for the *1991 World Development Report*. Robert Barro and Jong-Wha Lee (1993) have also estimated the average years of education of adults age 25 and over from UNESCO tabulations of census educational attainment distributions. Because one-half of childbearing occurs before women reach age 25, and as many as one-third of the women over 15 are between 15 and 25, the Barro-Lee education figures are not used here.[3]

The monetary, time, and psychic costs of birth-control technology can be reduced by organized family-planning programs. Activities of family-planning programs are commonly summarized by an index available for only three years on a comparable basis: 1972, 1982, and 1989 (W. Parker Mauldin and B. Berelson, 1978; John A. Ross et al., 1992 table 11). Some components of the index, however, are closely related to the prevalence of contraceptive use and, hence, do not distinguish between program-subsidized *supply* of birth-control services from behavioral changes that may reflect unobserved factors affecting parent *demand* for children. Thus, this family-planning variable is not a random experimental treatment. It is rather a complex social outcome that reflects both exogenous public initiatives to subsidize the supply and diffusion of contraception and the endogenous (to my model) response of parents who demand better means to avoid unwanted births. Treating this family-planning index as an exogenous determinant of population growth is, consequently, likely to overstate the policy effect of a supplement to the budget of any randomly selected family-planning program, due to simultaneous-equation bias. By subsequently fitting the model to changes over time within countries, I attempt to correct for this obvious bias due to omitted fixed, country-specific factors that could explain both levels of family-planning demands and fertility (Chang Hsiao, 1986).

IV. Empirical Results from Pooling Three Cross Sections: 1972–1989

Table 1 reports the cross-country reduced-form regressions on the six dependent variables represented in equations (1) and (2). The data are from all low-income countries in 1972, 1982, and 1989. The sample is restricted to countries for which there are data for at least two years. A sample of 204 country-years is obtained from 68 countries. (Sources of all data and the sample are described in Schultz [1993].)

To evaluate the estimated relationships, it may be useful to compare how a standard-deviation change in each of the explanatory variables (column (vii) of Table 1) is related to the dependent variables. An increase in family planning (e.g., 25 units such as shifting from Colombia to China in 1989) is associated with a reduction in birth rates of 2.5 and a reduction in crude death rates of 1.2, for a net impact on population growth of -1.2 per thousand. Increasing women's schooling by 2.2 years is associated with birth rates 4.7 lower and death rates 2.0 lower, for a net effect of reducing population growth by 2.8. More schooling of men has the opposite but smaller demographic effect to that of women. A uniform expansion of education for men and women is, therefore, associated with a substantial decline in population growth. The ratio of fuel exports to GDP is positively associated with birth rates, as expected for a pure wealth effect, and raises population growth by 1.3 per thousand. An increase of 370 calories is associated with a decrease in death rates of 0.9 and also, more surprisingly, with a decrease in birth rates of 1.9, for a net decline in population growth by 1.0. Agriculture has roughly offsetting effects, raising both birth

[2] GDP per adult has also been specified as a proxy for this variable with similar pronatal results (Schultz, 1993).

[3] Barro and Lee (1993) are, however, concerned with the education stock embodied in the labor force, and hence the age interval 25 or more is better designed for their needs than for this investigation. The conclusions I stress here were not affected by using the Barro-Lee education series.

and death rates, with no significant effect on crude population growth. Urbanization is not significantly associated with any of the crude vital rates.

More confidence may be placed in the second age-standardized decomposition of cohort surviving fertility (columns (iv)–(vi) in Table 1), from which the same patterns emerge. The same increase in family planning is associated with a 7.8-percent decrease in fertility (i.e., $25 \times (-0.0031)$) and a 0.9-percent increase in child survival, for a net decrease in surviving fertility of 6.8 percent. A standard-deviation increase in women's schooling is linked to a 18.4-percent decline in fertility and a 3.2-percent increase in survival, for a net effect on surviving fertility of -15.1 percent. Male schooling again exerts its smaller effect on fertility in the opposite direction. Countries that depend more on fuel exports for their income have more rapid population growth. Calorie availability is connected to 2-percent higher child survival, 5-percent lower fertility, and 3-percent slower cohort population growth. The larger proportionate declines in fertility than in child mortality suggest that the demand for surviving children is price-inelastic (Schultz, 1981). Age-standardized measures of vital rates confirm, as expected, that urbanization is associated in the cross section with lower fertility, and perhaps slightly lower survival, for a marked impact in terms of slowing cohort growth.

V. Empirical Results from Fixed-Effect Estimates, 1972–1989

The above estimates, based on a pooling of cross sections, may be inefficient and biased by the nonindependence of repeated observations on the same countries and by the omission of variables that affect fertility and mortality but that are correlated with the included regressors. It would be convenient to assume that these omitted country effects are random, whereas I expect that these omitted factors will be persistent and can then be treated as fixed for the short periods analyzed in this paper. Breusch-Pagan Lagrange multiplier tests (see Hsiao, 1986) reject the null hypothesis of homoscedasticity across the country-year

TABLE 1—REGRESSIONS ON POPULATION GROWTH COMPONENTS IN POOLED CROSS SECTIONS

A.

Explanatory exogenous variables (EEV)	Crude vital rates: Birth, B (i)	Death, D (ii)	Population growth, P (iii)
(a) Family-planning activity score	-0.0980 (5.13)	-0.0496 (5.61)	-0.0484 (2.44)
(b) Adult female years of schooling	-2.13 (5.68)	-0.875 (5.03)	-1.26 (3.22)
(c) Adult male years of schooling	1.11 (4.42)	0.182 (1.56)	0.928 (3.56)
(d) (Exports − imports of fuels)/GDP	8.86 (3.09)	-1.58 (1.20)	10.4 (3.50)
(e) Agriculture, percentage of labor force	0.102 (4.27)	0.0657 (5.97)	0.360 (1.45)
(f) Percentage urban	-0.0281 (0.99)	-0.0135 (1.03)	-0.0146 (0.50)
(g) Calories per capita (kcal/day)	-5.20 (3.89)	-2.41 (3.89)	-2.79 (2.01)
R^2:	0.759	0.833	0.418
Mean of dependent variable (SD):	38.9 (9.42)	12.4 (5.24)	26.6 (6.31)

B.

EEV	Age-standardized vital rates: Total fertility ln(F) (iv)	Child survival, ln(S) (v)	Survivors' fertility, ln(N) (vi)	Means (SD) variables in sample (vii)
(a)	0.0031 (4.92)	0.00037 (2.18)	-0.00273 (3.97)	35.0 (25.0)
(b)	-0.0824 (6.40)	0.0145 (4.35)	-0.0679 (5.03)	3.41 (2.23)
(c)	0.0365 (4.24)	-0.00134 (0.60)	0.0352 (3.90)	5.24 (2.48)
(d)	0.521 (5.30)	0.00956 (0.38)	0.530 (5.15)	0.243 (0.130)
(e)	0.00247 (3.02)	-0.00139 (6.60)	0.0011 (1.26)	50.3 (25.2)
(f)	-0.00296 (3.06)	-0.0001 (0.60)	-0.0031 (3.07)	38.6 (21.5)
(g)	-0.136 (2.97)	0.0544 (4.58)	-0.0819 (1.70)	2.38 (0.370)
R^2:	0.795	0.794	0.678	
Mean (SD):	1.62 (0.350)	-0.150 (0.0904)	1.47 (0.293)	

Notes: The absolute values of t ratios are reported in parentheses beneath ordinary least-squares coefficients. To save space, coefficients for controls of Catholic, Protestant, and Islamic religions, year dummies for 1982 and 1988, and a constant are not reported.

observations, under which the previous ordinary least-squares estimates are preferred. Specification tests also reject the null hypothesis that the country fixed effects are uncorrelated with the regressors (Hsiao, 1986). Consequently, unbiased fixed-effect (FE) estimates are reported in Table 2. However, FE estimates exclude information from the cross sections, while providing a robustness check on the more common estimated relationships across countries at different levels of economic development (Table 1), assuming the country effects are persistent.

In the fixed-effect estimates, family-planning activity is no longer significantly related statistically to any of the six demographic dependent variables, and point estimates are a small fraction of those reported in Table 1. Despite the controversial treatment of family planning as exogenous, it does not explain *changes* in demographic rates over time. The fuel-export share of GDP and agricultural share of the labor force remain significantly related to fertility, but the magnitudes are about one-half of those cross-sectionally estimated. Changes in urbanization no longer explain changes in vital rates, for this variable is undoubtedly extrapolated crudely by trends. The estimated effects of women's and men's schooling remain significant, and even increase overall in magnitude. Calorie availability is not statistically significant in explaining crude vital rates but continues to be negatively related to the preferred age-standardized measure of fertility.

The coefficients in the fertility equation, aside from family planning and urbanization, are similar in sign and rough magnitude in the cross section and FE time series. Narrowing the gap between women's and men's human capital and increasing employment outside of agriculture are the best available predictors of declines in fertility and population growth in low-income countries, both measured across countries at different stages in their development and over time within these same countries. Deriving national income from oil exports is associated with higher fertility and population growth. The FE estimates reconfirm that increasing women's schooling by a year

TABLE 2—FIXED-EFFECT ESTIMATES OF POPULATION-GROWTH COMPONENTS

A.

Explanatory exogenous variables (EEV)	Crude vital rates		
	Birth, B (i)	Death, D (ii)	Population growth, P (iii)
(a) Family-planning activity score	−0.00993 (0.51)	−0.0112 (1.27)	0.00127 (0.062)
(b) Adult female years of schooling	−2.05 (3.55)	0.896 (3.45)	−2.94 (4.85)
(c) Adult males years of schooling	2.10 (4.13)	−0.578 (2.53)	2.68 (5.01)
(d) (Exports − imports of fuel)/GDP	−0.467 (0.19)	−0.918 (0.84)	0.451 (0.18)
(e) Agriculture, percentage of labor force	0.999 (2.79)	0.0174 (1.08)	0.0825 (2.19)
(f) Percentage urban	−0.0062 (0.10)	−0.0118 (0.44)	0.0056 (0.09)
(g) Calories per capita (kcal/day)	−2.18 (1.78)	−0.147 (0.27)	−2.04 (1.58)
$F_{[9, 118]}$:	15.2	57.4	4.95
Lagrange multiplier, Breusch-Pagan test:[a]	61.6	67.4	62.4

B.

EEV	Age-standardized vital rates		
	Total fertility, $\ln(F)$ (iv)	Child survival, $\ln(S)$ (v)	Survivors' fertility, $\ln(N)$ (vi)
(a)	−0.00029 (0.34)	−0.00011 (0.80)	−0.00039 (0.46)
(b)	−0.131 (5.33)	−0.0206 (5.28)	−0.152 (6.00)
(c)	0.0877 (4.05)	0.0164 (4.77)	0.104 (4.68)
(d)	0.274 (2.65)	−0.0025 (0.15)	0.271 (2.55)
(e)	0.00418 (2.73)	−0.00028 (1.17)	0.00390 (2.48)
(f)	−0.0009 (0.33)	0.0000 (0.03)	−0.0008 (0.32)
(g)	−0.136 (2.60)	0.0050 (0.60)	−0.131 (2.44)
$F_{[9, 118]}$:	19.4	55.2	12.7
Lagrange multiplier, Breusch-Pagan test:[a]	58.9	87.5	58.0

Notes: The absolute values of t ratios are reported in parentheses beneath ordinary least-squares coefficients. To save space, coefficients for controls of Catholic, Protestant, and Islamic religions, year dummies for 1982 and 1988, and a constant are not reported.

[a] Null hypothesis: Homoscedastic ordinary least squares.

is associated with a 12-percent reduction of total fertility and a marked decline in crude or cohort population growth. The only puzzle in the FE estimates is the reversal in signs of the coefficients on men's and women's schooling in the death-rate and child-survival equations, which suggests to me that these reduced-form equations are probably better specified for fertility than they are for mortality. In the future, the determinants of mortality should receive more attention by economists to rectify this situation.

VI. Conclusion

The age-standardized decomposition of population growth provides the preferred basis for this analysis, though the results do not often differ qualitatively from those using crude vital rates. The fixed-effect estimates are less likely to be biased than those from pooled cross sections due to the omission of unobserved variables, most of which are probably persistent within a country. According to these fixed-effect estimates, increasing the schooling of women is the best predictor for reducing fertility and curbing population growth, whereas family planning does not exhibit a significant effect. Careful studies of household/community samples from several countries have isolated a significant impact of family planning on fertility, but the aggregate data for countries examined in this paper are not well designed to evaluate such social programs involving both public subsidies and private demands (Schultz, 1993). Previous studies that have used these country-level data and have shown large effects of family planning on fertility and population growth need to be reappraised.

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