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Birth Rate Changes over Space and Time: A Study of Taiwan

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I. Introduction

A number of recent empirical investigations have concluded that economic variables account for a statistically significant share of cross-sectional variation in aggregate and individual fertility. Though the models underlying these studies differ in terms of analytical complexion and econometric complexity, their empirical findings are nonetheless similar.[1] Theories of household behavior derived from generally accepted economic tenets alone, however, do not yield many refutable propositions with regard to fertility, unless additional constraints are imposed. The choice of a formulation to the economic theory of fertility may therefore have to rely more heavily on the weight of empirical evidence as to how this decision-making process is constrained than other fields of applied economics, where joint production is less essential and consumption and production are more readily dis-

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[1] A review of empirical findings presently available to me is provided in appendix tables A1 and A2, where in tabular form the sign and statistical significance of the regression coefficients estimated for each class of economic variable are reported. It is, of course, crucial for the interpretation of these summarized findings to also know, in addition to the nature of the data, what other variables were included in the regression equation, and, in simultaneous-equations estimates, the estimation techniques used and the structure of the rest of the system. Table A1 reviews studies of medium- and low-income countries; table A2 is confined to studies of the U.S. population.

tinguished. Most current evidence on these constraints derives from analyses of single cross sections; change in reproductive behavior over time is undoubtedly the dimension of fertility that economists have studied least. It is, therefore, the focus of this paper.

Economic theories of fertility determination generally presuppose an equilibrium relationship between permanent or lifetime conditions and completed fertility, whereas empirical verification is often based on observations of either conditions prevailing sometime after the arrival of the last child or current conditions in conjunction with an incomplete stock or rate of addition to an unobserved stock of children. Moreover, unbiased estimates of cause and effect relations are difficult to obtain from a single cross section in which most economic and demographic aspects of the household are jointly dependent on prior conditions that have also influenced fertility. Thus, little is known about the time required for reproductive behavior to adapt to unanticipated change in environmental constraints, for example, how parents react to policy measures that seek to provide family-planning information and reduce the supply costs of modern birth control. To cope intelligently with rapid rates of population growth occurring in most low-income countries one needs an understanding of precisely this process of behavioral adjustment to demographic and economic change and the resulting disequilibrium that it introduces into the family formation process.

The factors affecting the timing of births over a lifetime may differ fundamentally from those affecting completed fertility. In high-income countries, short-run changes in completed fertility are generally moderate in comparison with changes in birth rates induced by variation in the timing of births (Ryder 1969). In low-income countries where completed fertility is falling rapidly but unevenly, changes in the timing of births seem of much less consequence in accounting for changes in birth rates over time. Economic models of change in fertility over time, whose primary implications relate to lifetime reproductive goals, may, therefore, be tested more accurately using data on birth rates if the data are drawn from this second environment. The choice is made here to study a low-income country for which fairly reliable demographic time series are available. To satisfy these conditions, I have had to sacrifice detail of an economic nature and deal with aggregate data when information on individuals would be preferable.

In this paper I investigate approaches to the interpretation of 6 years of aggregate economic and demographic data for the populations of 361 small administrative regions of Taiwan. My purpose is to obtain more information on the dynamic structure of a model of fertility variations, disequilibria, and changes that will make better sense of population problems and policy options in low-income countries. Although not wholly unexpected, given the scarcity of analytical tools and the shortage of data relevant to this task, few firm conclusions emerge.

The setting of Taiwan is sketched in Section II. In Section III I briefly review the implications of economics for reproductive behavior and specify a model within the data constraints for Taiwan that can test several of these implications. Cross-sectional estimates of this model are reported in Section IV, and several approaches for combining time series and cross sections are explored in Section V to obtain a more satisfactory picture of the time dimension of the responsiveness of birth rates in Taiwan to changing economic, demographic, and policy constraints.

II. The Setting: Taiwan, 1950–70

To aid in interpreting the subsequent econometric results, I shall outline some major features of Taiwan's recent past. Taiwan's postwar economic progress is notable by any standards: real national income doubled from 1951 to 1961 and doubled again by 1968; growth in per capita income deflated by consumer prices has annually averaged about 5 percent since 1962; the share of employment in agriculture fell from 57 percent of the male labor force in 1951 to 36 percent by 1967, while the number of persons engaged in agriculture remained practically constant. Until 1961 the number of women employed in agriculture also remained virtually constant, and little growth in employment for women occurred elsewhere in the economy. From 1961 to 1967, however, the number of women in agriculture decreased by 30 percent, and nonagricultural employment of women increased about 130 percent. In terms of the opportunity value of a woman's time in child rearing, almost three-fourths of the women in the agricultural labor force in 1967 were classified as engaged in housekeeping activities, which one might surmise need not have curtailed their family commitments. The shift in the allocation of women's time out of the home, therefore, coincides with the acceleration in the decline in crude birth rates after 1961 (see fig. 1).

Taiwan's population reached 13 million in 1966 and has grown at the rapid rate of about 3 percent per year for most of the postwar period (see fig. 1). Registered crude death rates declined after the war from 18.2 per 1,000 in 1947 to 11.5 by 1950, and decreased to 5.0 as the population became increasingly concentrated in low-mortality age groups. Registered crude birth rates increased after the postwar readjustment period, peaking in 1951 at 50.0 per 1,000; only after 1956 did a secular downward trend emerge clearly. This significant turning point is seen more clearly in figure 2, where the age composition of the population is held constant by examining birth rates for women of specific age groups. The decline in birth rates did not occur uniformly across age groups; it is accentuated among women over 29 years of age but not yet substantial among women less than 30.

According to the 1967 census, women 35 years of age or older averaged between five and six live births. For women aged 35–39, 93 percent of these

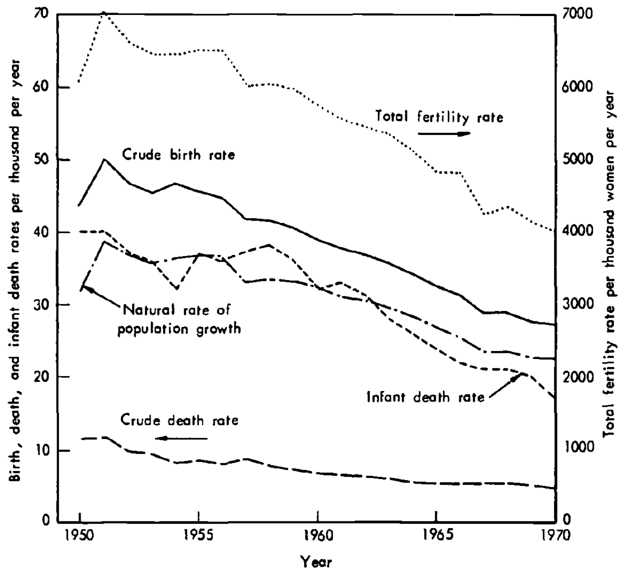

FIG. 1.—Annual registered vital statistics: Taiwan 1950–70 (source: Taiwan Dept. of Civil Affairs 1970, tables 1, 9, 10, and 17).

children were living, whereas only 80 percent of the children of women 20 years older were living. This difference in child-survival rates is in small part due to the greater age to which the older women's children have

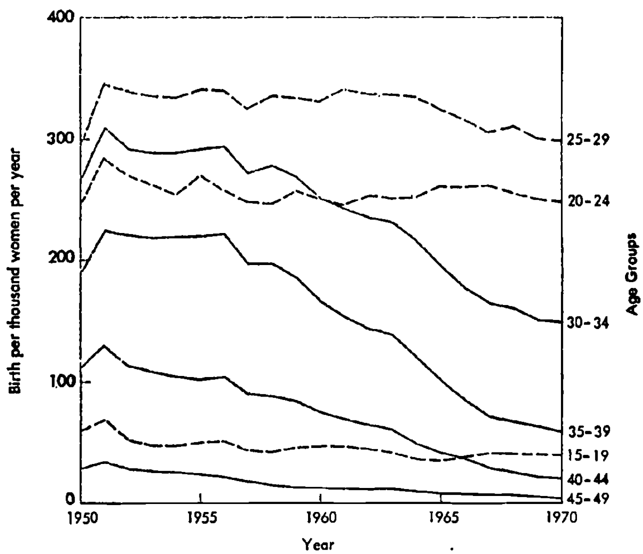

FIG. 2.—Age-specific birth rates: Taiwan (source: Taiwan Dept. of Civil Affairs 1970).

survived. More important, it is related to the decline in child mortality that has taken place in Taiwan and in most low-income countries during the last 20–40 years.

A large-scale family-planning program was launched on an experimental basis in Taichung City in 1963 (Freedman and Takeshita 1969) and in the following year in the majority of Taiwan's communities. Two types of fieldworkers were trained and coordinated to spread information about family planning and to promote the subsidized adoption of related services. The assignment of these workers across regions and over time is not obviously related to readily distinguished socioeconomic features of these regions or their resident populations.

III. Economic Hypotheses

According to the theory of consumer choice, quantity of a good demanded is constrained by available resources and relative prices, with associated "income" and "substitution" effects. When this framework is used to study factors influencing parent demand for children, however, well-known difficulties arise. First, parametrically given prices are not readily isolated or satisfactorily measured. Second, although additions to household resources relax the family's budget constraint, they also tend to change relative prices, depending upon their source, with offsetting effects on the demand for children (Schultz 1969). There are few instances, therefore, where either pure price or income effects can be identified and estimated from existing data.

A closely related problem stems from the interrelated nature of life-cycle choices, many of which bear upon reproductive behavior (Nerlove and Schultz 1970). Proxies that might at first appear useful as measures of the opportunity cost of a child, such as years of schooling completed by the child, must be treated as endogenous variables in a broader system of behavioral relationships. This household decision-making system also interacts with such public sector choices as the allocation of educational and family-planning expenditures. Simultaneous-equations estimation techniques are required to obtain asymptotically unbiased estimates of interactions among these classes of behavior; the consistency of ordinary least-squares estimates of a "demand" equation for fertility depends critically on the predetermined nature of all explanatory variables. Neither theory nor data are now adequate to identify and estimate the underlying structural demand and supply equations for both the number and the resource intensity of children.

Nonetheless, under a variety of assumptions explored in several papers in this volume, a reduced-form "demand" equation can be expressed for numbers of children that contains explanatory variables assumed to be outside of parents' control.[2] Since the implications of these alternative models do not differ with respect to the variables observed for Taiwan, no test of the adequacies of their underlying assumptions can be proposed. At least one dimension of public policy, however, will be explored as it might affect fertility, namely, the allocation of field personnel to family-planning activities.

Empirical Formulation of the Model for Taiwan

Only the rudiments of a formal model of fertility determination can be tested against Taiwan data. The object is to specify a relation between the birth rate parents want and the price, income, and information constraints that are not themselves determined simultaneously with, or subsequently by, the objective number of births. A single reduced-form equation is therefore analyzed here: that between birth rate in year t, b_t, and the lagged values of (1) the reciprocal of the child-survival rate, D (*return* and price effects); (2) the proportion of the male labor force employed in agriculture, A (relative *price* effects); (3) male school attainment, E^m (*income*, price, and information effects); (4) female school attainment, E^f (*price*, income, and information effects); and (5) two classes of family-planning fieldworkers employed by the health and family-planning programs, H and F (*information* and price effects). These variables and the time lags τ and τ' are defined more precisely in table 1, where data sources are given. The correlations among the variables and their means and standard deviations are shown in table 2. Unfortunately, none of the explanatory variables is available on an age-specific basis. Assuming the relation is approximately linear,[3] the equation to be estimated is of the following form:

$$\begin{aligned} b_t = \alpha_0 + \alpha_1 D_{t-\tau} + \alpha_2 A_{t-\tau} \\ + \alpha_3 E_{t-\tau}{}^m + \alpha_4 E_{t-\tau}{}^f \\ + \alpha_5 H_{t-\tau'} + \alpha_6 F_{t-\tau'} + e_t. \end{aligned} \tag{1}$$

The expected signs and relative magnitudes of the parameters, α's, and the choice of lags are discussed below. The disturbance term, e_t, when the

[2] More formally, ordinary least-squares estimates of the parameters to the equation are asymptotically unbiased only when all explanatory variables are predetermined. Lagged values of endogenous variables cannot, in general, be treated as independent of the disturbances in relationships accounting for the same or related forms of current behavior. Omitted variables pertaining to both the "quality" demand equation and the "quality" and "quantity" supply equations must also be assumed uncorrelated with observed explanatory variables considered in the "quantity" demand equation.

[3] Alternative nonlinear formulations of similar models are discussed and estimated for the Taiwan data elsewhere (Schultz 1971*a*). Cross-sectional findings do not change substantially nor do the tests of hypotheses depend on which of these functional specifications of the model is adopted.

TABLE 1

DEFINITIONS OF VARIABLES FOR 361 ADMINISTRATIVE REGIONS OF TAIWAN

b_t = The number of births per thousand women of specific age group in year t.

D_t = The reciprocal of the probability of survival from birth to age 15 estimated from age-specific death rates in year t.

A_t = The proportion of the male labor force employed in the agricultural sector in year t.

E_t^m = The proportion of men 12 years and over with at least a primary school certificate in year t.

E_t^f = The proportion of women 12 years and over with at least a primary school certificate in year t.

H_t = Number of man-months of VHEN (Village Health Education Nurse) employed by the family-planning program per thousand women aged 15–49 from 1964 to year t.

F_t = Number of man-months of PPHW (Pre-Pregnancy Health Workers) employed by the family-planning program per thousand women aged 15–49 from 1964 to year t.

e_t = Independently, normally distributed random disturbance until assumed to be otherwise in Section V.

τ = Combined behavioral and biological lag between the change in an environmental constraint on desired fertility and the change in birth rates.

τ' = Biological lag between change in birth control practice and birth rate.

α's = Parameters of the model to be estimated.

SOURCES.—Taiwan Department of Civil Affairs, *The Taiwan Demographic Factbook*, various annual issues since 1961; Taiwan Population Studies Center, *The Demographic Reference: Taiwan*, various years since 1956; and Taiwan Provincial Institute of Family Planning, *Family Planning Reference Book*, vol. 1, 1969.

observations are appropriately weighted is assumed to be normally distributed with zero mean, constant variance, and uncorrelated with the explanatory variables.

The dependent variable in the fertility equations estimated in the next section is birth rates for an age group of women for the years 1964–69. This is not the dimension of fertility underlying the theoretical basis for equation (1), namely the number of children born to an individual women. Formally, several unverified assumptions are required to assure that the aggregate flow of births across groups of women of different ages responds in the same way to differences in the determinants of fertility as does the stock of children-ever-born across individual women. Yet if one is not to wait decades until today's younger parents in Taiwan have completed their childbearing years in order to estimate the effects of changing environmental constraints on their completed fertility, research must in the interim analyze current birth rates for evidence of the contribution of specific dimensions of economic change and population policies. Thus the closest approximation to the economically prescribed dependent variable is the birth rate among women at least 30 years of age. Because these women are frequently completing the formation of their families, they are pri-

TABLE 2

SIMPLE CORRELATIONS AMONG VARIABLES*

VARIABLE LEVELS	VARIABLE DIFFERENCES 1969–64													
		Age-Specific Birth Rates												
	Total Fertility Rate† (1)	15–19 (2)	25–24 (3)	25–29 (4)	30–34 (5)	35–39 (6)	40–44 (7)	45–49 (8)	Child-Death Adjust-ment (Ratio) (9)	Agricultural Composi-tion (Propor-tion) (10)	Male Educa-tion (Propor-tion) (11)	Female Educa-tion (Propor-tion) (12)	VHEN (13)	PPHW (14)
1966:														
(1)	...	.262	.214	.393	.611	.489	.351	.229	.084	—.037	—.041	—.095	—.076	—.003
(2)	.374	...	.427	—.131	—.149	—.159	—.119	—.017	—.074	—.064	.056	.114	.096	.026
(3)	.594	.465	...	.123	—.366	—.464	—.497	—.409	—.083	—.097	.169	.208	.084	.173
(4)	.699	—.195	.317	...	.050	—.155	—.216	—.268	—.047	.033	.108	—.004	—.028	—.006
(5)	.851	.127	.184	.594	...	.548	.403	.353	.098	.026	—.176	—.234	—.081	—.059
(6)	.890	.318	.295	.506	.855	...	.581	.369	.192	.028	—.205	—.212	—.154	—.124
(7)	.836	.345	.265	.424	.801	.878	...	.566	.151	.000	—.119	—.111	—.090	—.135
(8)	.416	.146	—.026	.120	.459	.518	.545	...	.054	—.020	—.073	—.147	—.068	.037
(9)	.583	.307	.366	.317	.493	.527	.553	.187	...	—.009	—.090	—.019	.023	—.087
(10)	.567	—.077	.384	.646	.429	.433	.397	.138	.379	...	—.118	—.128	—.122	—.058
(11)	—.735	—.123	—.555	—.646	—.579	—.582	—.498	—.198	—.404	—.659	...	.352	.138	.092
(12)	—.643	.063	—.507	—.719	—.507	—.451	—.362	—.078	—.321	—.673	.909	...	.079	.069
(13)	.046	—.075	.053	.158	.008	.004	—.011	—.126	.063	.215	—.120	—.173	...	—.270
(14)	—.113	.017	.160	.066	—.012	.000	—.003	.009	.070	.181	—.113	—.158	—.299	...
Mean	4,835.0	40.7	275.	328.	189.	90.8	37.7	5.89	1.06	0.440	0.694	0.484	0.256	1.36
Standard deviation ..	806.0	21.2	45.9	45.9	44.6	38.7	22.4	6.84	0.0242	0.291	0.104	0.146	0.330	1.38
1969–64 differences:														
Mean	—948.0	2.31	—8.11	—35.0	—61.6	—55.0	—28.3	—3.92	—0.0296	—0.0166	0.0628	0.0884	0.574	4.72
Standard deviation ..	366.0	17.2	45.0	35.9	33.0	31.1	20.7	10.3	0.0411	0.0248	0.0460	0.0555	0.601	2.66

NOTE.—Variable levels in 1966 appear below the diagonal; variable differences, 1969–64, appear above it.

* Weighted by the number of women between the ages of 15 and 49 in each observed region.

† The total fertility rate (which is five [the age interval] times the sum of the age-specific birth rates) represents a cross-sectional proxy for completed fertility expected per thousand women surviving to age 50. It is not a longitudinal measure of completed fertility for any specific group of women, but rather an expected measure of fertility that would occur if current age-specific birth rates persisted for all childbearing years of a surviving cohort.

marily engaged in the marginal childbearing decisions for which the economic model is relevant.

Child mortality appears to increase the cost of rearing a child to maturity. If we can assume that parents feel themselves better off, all things considered, the longer their child survives (O'Hara 1972), an overall decline in mortality should make childbearing more attractive. But if the marginal return from having additional surviving children diminishes and becomes negative for small increases in the size of surviving family, the relation between child mortality and fertility will embody two partially offsetting price changes—a decline in the cost of rearing a surviving child and a decline in the marginal return. Economics does not tell us the magnitude of the net effect or even its sign. The observed behavior of parents will provide information on the relative weight parents attach to the costs of rearing a child who dies compared with the elasticity of their demand for surviving children. Highly inelastic demand for surviving offspring implies that there should exist in equilibrium a positive association between birth rates and anticipated or experienced child-death rates (or the reciprocal of survival rates). Changes in this child-death adjustment ratio would also affect the returns to human capital investments in a child, increasing the probability that a positive association will exist between child-death and birth rates (Schultz 1971*c*; O'Hara 1972).

Agricultural composition of the male labor force is intended to capture relative price effects. The cost of rearing any specific number of children might be lower in rural than in urban areas (the prices of food and housing might be lower), while, in traditional agriculture, the rate of return to schooling children may be depressed relative to the return anticipated by parents in the more dynamic urban sectors of the economy. This disparity in the returns to education could arise from higher direct and opportunity costs of obtaining schooling for the dispersed farm population, or from lesser productivity gains for the better-educated in agricultural in comparison with nonagricultural activities. The reduced-form equation discussed here cannot attribute a partial association between economic sector and fertility to any particular class of differential input prices or to differential returns to schooling, but both sets of relative prices might account for higher fertility in rural regions of many low-income countries when other demographic and economic factors were held constant. However, it need not follow that marriage and childbearing would occur earlier in the agricultural sector, for quite to the contrary, the scarcity of land and capital has often been attributed a role in delaying marriage in the rural sector when the urban sector is prospering.

The value of the time of each spouse is assumed to be equalized at the margin between market and nonmarket work. This would be the case if each spouse is engaged in both market and nonmarket activity, or if the wife specializes entirely in nonmarket production she is assumed to produce within the household some perfect substitutes for market goods.[4] I will assume, in other words, that the implications of "corner solutions" considered by Gronau (1973) and Willis (1971) are not here an important factor in accounting for the household allocation of time between market and nonmarket activities.

The measure of schooling considered here approximates the median accomplishments of the adult population. If this variable adequately reflects the lifetime value of husband and wife time in market and nonmarket activities,[5] plausible assumptions regarding the relative value of sex-specific

[4] This may sound like a strong assumption, but I think it is more realistic than the converse that stresses "corner solutions" to the household time-allocation problem in which the woman allocates all of her time to nonmarket work that has no meaningful market substitute. First, in a cross section only a proportion of the population is in the paid labor force at any one time. Over a lifetime a much larger fraction would have entered the labor force, equalizing market and nonmarket opportunities at various points in their life cycles. Second, a wide range of nonmarket production directly substitutes for market goods, services, and labor; unpaid family labor can at times be replaced by hired labor; much of household production is virtually interchangeable with goods and services that could be purchased from or bartered for in the market. I conjecture that over a woman's lifetime, even if she is a most productive homemaker situated in a remote labor market that does not offer her an attractively paying job, she will nonetheless allocate her time according to the relative prices of market goods and services she can replace by home-produced ones. If her household develops a comparative advantage in the production of some of these fringe "market" goods and services, this may induce her to become a petty trader or participate in a "home industry," as occurs frequently throughout much of Africa and Southeast Asia. A continuum of market and nonmarket activity exists, and the overlap is sufficient to equilibrate market and nonmarket productivity in most instances. My hunch is that the "corner solution" may be of interest in some highly developed societies, where the comparative advantage of home production is being eroded except in final consumption activities, and perhaps in the Muslim Middle East where very few employment options are open to the woman outside of the home.

[5] Observed wages, however, may differ from the opportunity cost of time because the market wage rate is net of investment in on-the-job training (Mincer 1962*a*); because depreciation in specific types of human capital may be modified by shifting time between market and nonmarket activities (Michael and Lazear 1971); and because the market wage may include a premium for the utility (or disutility) of market work. For example, the wife who allocates her time to nonmarket activities may accelerate depreciation on her stock of market-specific human capital. The downward bias on market wages as a measure of the current price of time would tend to be serious during the first decade after school completion for men, but it is not clear whether this is also true for women. More generally, the variable that is predetermined with respect to the fertility decision is not the current potential market wage but a potential lifetime wage stream (Mincer 1974*b*). How people redistribute that wage stream by *all* forms of human capital investments will be jointly determined with their desired number and timing of births. Only this preinvestment potential wage can be viewed as uncontaminated by life-cycle goals and unobserved factors, such as tastes. If this interpretation has merit, the contemporaneous association between wages and completed fertility contains a serious simultaneous-equations bias that is very difficult to isolate completely. The first step would appear to be to estimate wage functions separately and use them to predict potential lifetime wage opportunities as explanatory variables in the derived demand-for-children equation. For these reasons, schooling might be preferable to current market wages as a proxy for potential lifetime wages.

time inputs into the rearing of numbers of children[6] imply $\alpha_4 < \alpha_3$, and past studies of fertility might lead one to conjecture that $\alpha_3 \geqslant 0$, and $\alpha_4 < 0$. A serious practical problem with joint estimation of the partial association between birth rates and male and female schooling is the collinearity between the sex-specific schooling rates.[7]

Information about the nature and local availability of modern birth control methods is provided by the two classes of family-planning field personnel; this should reduce birth rates among older women by reducing the information costs of avoiding additional unwanted or ill-timed births. But the reverse effect might occur among younger women, who may adopt a more concentrated pattern of childbearing when they are provided with a reliable means of birth control (Keyfitz 1971). Activities that seek to spread a given set of information to a fixed population of potential demanders are likely eventually to experience decreasing returns to effort. This anticipated nonlinearity in the relation between family-planning activity and its effect on birth rates is documented in another study (Schultz 1971*a*) but is neglected here for simplicity. The regression coefficients α_5 and α_6 should, therefore, be interpreted with caution as only approximations for the *average* effectiveness of a personnel man-month of local activity per 1,000 women of childbearing age.

The impact of personnel allocated to family planning on completed fertility will depend upon at least three factors: (1) the elasticity of parent demand for numbers of children; (2) the advantage of the program's new birth control technology compared with that indigenously available, in terms of individual search and use costs; and (3) the distributional efficiency of the program imparting new knowledge of birth control and subsidizing its use among those segments of the population with a more elastic demand for children and least initial access to adequate (modern) birth control technology.

Temporal Specification

Anticipations about the future are not always fulfilled; it takes an uncertain amount of time to have a child, and children die unpredictably.

[6] This conclusion is derived rigorously in the model presented in this volume by Ben-Porath, in which the variable resource intensity of children is omitted. It is implied as well by Gardner (1972) and assumed by Willis (1971). In De Tray's (1972*a*) model the female time-value intensity is assumed greater than the male's for numbers (C) of children, leading to the prediction of an algebraically smaller coefficient for the female than for the male wage variable in the equation for derived demand for children.

[7] The simple correlation (weighted by number of women aged 15–49 in the relevant region and time period) between the male and female primary school variables is very high, .91. Between the change variables the correlation is, of course, less, .35 (see table 4). I have not uncovered any good reasons why the regional sex-specific schooling variables should be so highly correlated in a country, such as Taiwan, that is experiencing substantial migration and other rapid changes, such as those in educational attainment.

Consequently, lags and adjustments of a biological and behavioral nature interpose themselves between the observed environment characterized in the model and the observed birth rate the model seeks to predict. In other words, what should be the values of τ and τ' in equation (1)? Most environmental conditions that modify parents' preferences for children change slowly, at least at the regional level considered here. Death rates, on the contrary, are subject to substantial year-to-year variation and cannot be predicted accurately within a small community, to say nothing of the family unit. Short-run reaction to replace recent child losses may therefore permit identification of the speed and magnitude of parent reaction to unanticipated changes.

Some sequential models of fertility determination would imply a mechanism by which the unanticipated death of a child to a still-fertile mother may influence her to have a somewhat higher birth rate than she would otherwise have had. Although the death of a child may motivate the mother to seek an additional offspring, the effect on her actual reproductive behavior is difficult to distinguish in the *short run* if she is young, for members of her age cohort will probably continue having additional desired children for some time, regardless of their personal experience of child mortality. If, on the other hand, the mother of the deceased child is older, say in her late thirties, when a sizable proportion of her cohort intends to avoid further births, her seeking another birth will sharply distinguish her from others in her cohort within a relatively short time. Experimentation with various discrete (see τ in eq. [1]) and simple distributed lags for Taiwan suggests the association between regional child-death rates and subsequent birth rates is statistically strongest after the lapse of 2–4 years, particularly among women 30–44 years of age (Schultz 1971*a*). These findings, that a pronounced 2–4-year average behavioral-biological lag exists between child deaths and subsequent births, are generally consistent with other evidence on the mean length of closed birth intervals for women not practicing modern contraception (Potter 1963). Lacking better evidence, not only the child-death ratio but also the schooling and agricultural variables that must affect birth rates with a behavioral-biological lag are lagged 3 ($\tau = 3$). The bias introduced by this approximation for what undoubtedly should be a more complex distributed-lag formulation is discussed in Section V.

Family-planning program inputs represent a still more complex problem of measurement and dynamic model specification. The temporal link between environmental conditions and fertility and that between birth control information and fertility should differ. Current practice of birth control influences birth rates 9 months hence; the lag between adoption of more effective birth control methods and their maximum impact on birth rates is likely to be no more than 1 year, not the 2 or 3 required, on average, for a woman to deliver a child. Therefore, program inputs are lagged only 1 year ($\tau' = 1$). Furthermore, the program is likely to have a persisting effect on contraceptive knowledge in the community, slowly "depreciating" as people forget or move away and as new generations of parents grow up or move in. Thus, all past program inputs must be considered as possible determinants (with perhaps exponentially diminishing weights) of the target population's current knowledge of birth control methods. Time series as yet are too short for an empirical exploration of the declining time dimension of program input effectiveness. A first approximation is simply to ignore the depreciation effect and treat the total of all past program inputs as a determinant of current effective knowledge of birth control in the community.

IV. Cross-sectional Evidence

Ordinary least-squares estimates on equation (1) are reported for 1965 and 1969 in table 3, for which the moment matrices were weighted by the observed number of women in the specific age group in the region;[8] only selected implications of these results are discussed below.

In each year from 1964 to 1969 (not shown), between 48 and 65 percent of the interregional variation in total fertility rates is accounted for by equation (1). There is substantial variation in the explanatory power of the model across years and age groups and in the size and statistical significance of the regression coefficients of each explanatory variable.[9]

The coefficient for the child-survival factor is positive and substantial relative to the magnitude of birth rates in every year and age group, although this pattern is most pronounced among older women and teenagers. For example, for the age group 40–44 the ratio of the coefficient of the child-survival factor to the mean value of the birth rate approaches 15 by 1969, and even for the total fertility rate this ratio increases to more than three by 1969. These estimates imply parent demand for surviving children is inelastic.

[8] Disturbances in the birth-rate equation are likely to have differing variances. The expected variance of the estimate of a population mean (the frequency of births in this case) is inversely proportional to the size of the sample population. Therefore, the observations on which the regression is based are weighted by the square root of the regional population of women of the specific age group in the relevant year. Neglect of heteroscedasticity of disturbances asymptotically biases upward the estimated variance of the parameters, reducing their t-ratios. However, in the case at hand, it seems likely that the variance of the disturbances may not be independent of other variables in the model, which implies the ordinary least-squares estimates of the parameter means may also be biased. Similar reasoning often leads to respecifications of estimation equations to obtain approximately constant-variance disturbances (Malinvaud 1970, pp. 302–6).

[9] The t-ratios reported in parentheses beneath each regression coefficient in all of the tables are used to test statistical confidence in rejecting the null hypothesis that there appears to be no particular association between the variable and the birth rate. Since most of the parameters have a predicted sign, the one-tailed test is generally appropriate. For a large n (degrees of freedom), a t-ratio exceeding 1.65 satisfies a confidence level of 95 percent, and one exceeding 2.34 satisfies a confidence level of 99 percent.

TABLE 3

CROSS-SECTIONAL REGRESSIONS ON BIRTH-RATE LEVELS

Dependent Variable	Constant Term	Child-Death Adjustment (Ratio)	Agricultural Composition (Proportion)	Male Education (Proportion)	Female Education (Proportion)	Family-planning Program Man-Months (Cumulative Inputs to Prior Year): VHEN	Family-planning Program Man-Months (Cumulative Inputs to Prior Year): PPHW	R^2/SEE
Total fertility rate:*								
1965	2148. (2.70)	4567. (7.15)	535. (3.76)	—4156. (6.53)	913. (1.95)	—93.7 (0.71)	—136. (2.58)	.529 (546.)
1969	—699. (4.35)	13,150. (9.20)	189. (1.67)	—4944. (7.78)	1372. (3.29)	190. (4.39)	18.7 (1.89)	.588 (455.)
Age-specific birth rates for ages:								
15–19								
1965	—65.1 (2.56)	135. (6.57)	—11.0 (2.41)	—116. (5.69)	90.1 (6.04)	—7.92 (1.90)	—2.78 (1.66)	.271 (17.5)
1969	—232. (4.22)	327. (6.70)	—11.9 (3.15)	—168. (8.00)	98.5 (7.10)	5.12 (3.55)	.542 (1.65)	.317 (15.2)
20–24								
1965	169. (2.94)	172. (3.68)	—10.1 (0.98)	—136. (3.03)	6.47 (0.20)	6.00 (0.64)	—2.64 (0.70)	.159 (38.9)
1969	—373. (2.55)	739. (5.63)	—.858 (0.09)	—244. (4.51)	7.01 (0.20)	14.8 (3.97)	2.95 (3.41)	.453 (38.3)
25–29								
1965	255. (5.81)	92.8 (2.63)	49.2 (6.29)	—5.87 (0.17)	—97.3 (3.79)	12.1 (1.67)	—3.14 (1.08)	.492 (29.7)

* See definition note to table 2.

TABLE 3 (*Continued*)

Dependent Variable	Constant Term	Child-Death Adjustment (Ratio)	Agricultural Composition (Proportion)	Male Education (Proportion)	Female Education (Proportion)	Family-planning Program Man-Months (Cumulative Inputs to Prior Year): VHEN	Family-planning Program Man-Months (Cumulative Inputs to Prior Year): PPHW	R^2/SEE
1969	121. (1.05)	215. (2.10)	37.3 (4.71)	.124 (0.00)	−127. (4.37)	5.38 (1.77)	.794 (1.13)	.480 (31.0)
30–34								
1965	101. (1.97)	181. (4.43)	27.4 (2.98)	−188. (4.57)	38.0 (1.26)	−5.20 (0.62)	−5.76 (1.71)	.363 (35.0)
1969	−312. (2.81)	539. (5.50)	6.77 (0.83)	−192. (4.22)	67.0 (2.25)	6.61 (2.14)	−.146 (0.20)	.278 (32.3)
35–39								
1965	−19.1 (0.43)	216. (6.00)	30.5 (3.80)	−233. (6.49)	79.6 (3.02)	−14.0 (1.89)	−9.31 (3.14)	.462 (30.7)
1969	−438. (5.33)	568. (7.81)	2.19 (3.63)	−229. (6.85)	134. (6.12)	4.55 (2.04)	−.0819 (0.16)	.346 (24.0)
40–44								
1965	−26.0 (0.98)	115. (5.43)	16.7 (3.46)	−120. (5.56)	41.6 (2.65)	−6.94 (1.55)	−3.93 (2.18)	.397 (18.5)
1969	−251. (5.36)	301. (7.26)	6.21 (1.83)	−113. (5.94)	71.9 (5.83)	2.18 (1.72)	−.127 (0.43)	.323 (13.5)
45–49								
1965	12.0 (1.23)	7.70 (1.00)	3.17 (1.81)	−43.3 (5.35)	30.7 (5.19)	−3.23 (1.93)	.0613 (0.09)	.132 (6.94)
1969	9.09 (0.51)	5.65 (0.36)	−.637 (0.49)	−35.9 (4.69)	28.3 (5.53)	−.125 (0.24)	.115 (0.99)	.093 (5.58)

In the short run, it was anticipated that the link between child mortality and fertility would be more pronounced among older mothers; this seemed plausible because they were more likely already to have the number of surviving offspring they wanted and would, therefore, weigh heavily the survival or death of these earlier children in their decision whether to have an additional child. The strong statistical association between child-death rates and birth rates of women aged 30 and over confirms this hypothesis; the strength of the association with teen-age birth rates suggests two subsidiary tentative hypotheses.

Life-cycle commitments of a mother may depend to a substantial degree upon the survival of her first born. The dislocation of her life may be somewhat less if a subsequent child dies later in the formation of her family. It may be assumed that teen-age mothers more frequently experience the loss of their first born than other age groups and are thus motivated more strongly, on average, to shorten the interval to their next birth.[10] This logic could account for a strong replacement link at the family level between child mortality and the teen-age birth rate.[11]

Societies also traditionally adjust to the heavy incidence of childhood mortality by promoting early marriage, and though this pattern may stem from desired fertility goals, it is likely to become institutionalized and only slowly adapt to sudden changes in death rates or in other constraints on the family formation process. The aggregate association observed in Taiwan between child-death rates and subsequent teen-age birth rates may, therefore, represent *long-term* regional differences in the proportion of married teen-age women (see Schultz 1971*b*). Examination (in the next section) of time series may help to determine whether this relation observed in the cross section of aggregates is due to misspecification of the model's temporal dimensions or neglect of intermediary causal relations involving, for example, the timing of first marriage.

Local reliance on agricultural employment is often positively associated with total fertility rates, except in 1967, but *t*-ratios are not substantial. Differences in the age pattern of fertility in agricultural regions are more pronounced, however. Birth rates among teen-age women are decidedly lower in agricultural areas, but distinctly higher among women aged 25–29, and somewhat higher during the first half of the period among all women over age 29. These findings do not confirm the importance of

[10] First, it is a medical fact that infant mortality is greater among the children of teen-age mothers and among firstborn. It is also true that a greater proportion of a teen-age mother's children at risk of dying are firstborn, compared with older groups of mothers. Since the child-death rates cannot be linked to the age of the child's mother, the child-death rate is likely to reflect disproportionately the family mortality experience of the youngest group of mothers.

[11] A similar pattern of stronger reproductive replacement response was observed among teen-age married women (and married women between the ages of 30 and 39) to the loss of their own child in a study of household survey data from Bangladesh (Schultz and DaVanzo 1970, table 12).

relative price effects operating between the agricultural-nonagricultural sectors of Taiwan, but they do suggest that the timing of marriage and births may be modified by the availability of complementary agricultural resources.

A puzzling result from these cross-sectional regressions is the consistently negative coefficient of male schooling and the positive coefficient of female schooling. Although collinearity may be called upon to excuse these unexpected results, and the schooling variable may be criticized as a proxy for the value of parent-time, the results nonetheless cast doubts on the adequacy of static cross-sectional evidence of the economic determinants of fertility in a rapidly changing environment.

Except for birth rates among women aged 25–29, the negative partial association with male schooling is statistically notable in every year and age group. Increasing the proportion of men with a primary school certificate from 50 to 60 percent in a community is associated with approximately a 10 percent reduction in births, with the greatest proportionate declines occurring among women aged 15–19 and over 30.

Increased schooling for women is associated with lower teen-age birth rates, but noticeably higher birth rates among women over the age of 35. Although the teen-age pattern might have been predicted, the expectation from economic theory is for educated women to have fewer births after age 30. As industrialization has proceeded, the historic tendency has been for the value of women's time to increase relative to market goods and for women's participation in the paid labor force to increase. If proper controls were available for husband's wage rate, nonhuman wealth, and child-survival rates in a regression model accounting for variation in fertility, better-educated women would tend to have fewer births in their lifetimes. But in all likelihood, the better-educated women would concentrate these fewer births in the first decade of their marriages and possibly exhibit higher birth rates in their mid- and late twenties compared with less educated women. These expectations are not confirmed by the cross-sectional estimates based on the levels of birth rates in Taiwan.

The nationwide family-planning program initiated in 1964 appears to have reduced the total fertility rate in 1965 and 1966, but thereafter program efforts are *positively* associated with total fertility rates. An explanation for this weakening and eventually reversed relationship between program information inputs and the level of total fertility is found in the sharply diminishing returns to scale experienced by the program and the differential effects of the program on women of different ages. The former attribute of the program's effectiveness has been treated elsewhere, and an unnecessary digression would be required here to introduce nonlinear production functions for family-planning activities (see Schultz 1971*a*). Suffice it to say that the marginal effect of program personnel at average employment levels diminished from 1965 to 1969 as the cumulative level of program activity increased, while in each single year the marginal effect of personnel was relatively less in regions where past program activity was relatively intense. The differential age effects, however, are visible, even from the simple linear estimates of the family-planning program reported in table 3.

Program inputs are most consistently linked to lower birth rates among women 30 years old and over. Among these prime candidates for family-planning assistance, the distribution of field personnel is strongly associated with substantial declines in age-specific birth rates.

A very different pattern of program effectiveness emerges from the study of birth rates among women less than 30 years old. In the first year or two of the national program, field personnel elicited several reductions in these younger age-specific birth rates. After 1966 the tendency became increasingly statistically significant for birth rates among women aged 15–29 to be *higher* in regions intensively canvassed by family-planning program fieldworkers. This evidence is consistent with the hypothesis advanced earlier that where reliable birth control methods are understood and made more accessible, marriage may not be further delayed; and child spacing, if more widely practiced at all, may produce oscillations in birth rates among younger women.

Another implication of these results that is seen more clearly in the nonlinear formulation of the model (Schultz 1971*a*) is that one class of fieldworker—the Village Health Education Nurse (VHEN)—was more effective in reducing birth rates in the early years of the program than the other class of fieldworker—the Pre-Pregnancy Health Worker (PPHW). (The PPHW is trained in the family-planning program to systematically contact and recruit mothers, generally in their thirties, with a recently recorded birth; the VHEN, on the other hand, is instructed to disseminate the principles and practices of home economics, family sanitation and hygiene, and family planning to the entire village population. The observed pattern of differing age-specific effectiveness of these two classes of fieldworker is, therefore, not implausible.)

The major puzzle embodied in these cross-sectional findings attaches to the estimates for the male and female schooling variables. The paradox is not resolved by indirectly adjusting the variables for regional differences in age composition, nor by adopting alternative measures of educational attainment. Inclusion in the model of a proxy for male income, derived from the regional structure of male employment and national estimates by industry of value added per worker, reduces the size and significance of the coefficient on the male schooling variable but does not change the signs on male or female schooling. Reversal of this unexpected finding is discussed in the next section, where analysis deals with time series and cross sections together.

V. Time Series of Cross Sections

Economic models of the determinants of fertility tend to be formulated and tested in static terms. Parents are viewed as deciding in a single period on the appropriate number of births needed to yield them an optimal lifetime number of children. Though these abstractions have proved a powerful generalizing device where none has existed before, little attention has been given to the question of what economic theory and statistical techniques can say about models of dynamic behavior that might be confirmed or refuted by empirical evidence (Nerlove 1972*a*). Reproductive behavior occurs sequentially, and the constraints on child-bearing exert diverse influences on many other areas of economic and demographic decision making in the household sector. Exploration of the time dimension of this process and its complex ramifications on other household choices is warranted.

In the preceding section, cross-sectional variation in the level of *birth rates* was analyzed, although the implicit economic model set forth earlier was framed in terms of completed lifetime fertility. This unavoidable change in the measurement of the dependent variable provides certain advantages, however, when the temporal dimension of parent reaction to changing economic and demographic constraints is the object of analysis. Discrete lags between fertility and the explanatory variables were introduced to approximate the average time for reproduction to respond and for birth control information to take effect. But the stochastic nature of the reproductive process and the numerous neglected features of the individual that could affect reaction times suggest that a *distributed* lag would be more appropriate to the study of changes in fertility. Yet identification and estimation of these lag structures are difficult because of the limited availability of time-series information and the strong positive serial correlation of such relevant characteristics of regional populations as wages, nonhuman wealth, industrial structure, and schooling.

The cross-sectional findings in Section IV are, as a result, not to be interpreted as estimates of the impact of slowly changing environmental constraints on birth rates during the 3-year lag interval. For example, the systematic portion of the regime of mortality is determined by such slowly changing factors as long-term investments in public health, sanitation, water supplies, transportation, geography, climate, and socioeconomic characteristics of the population; thus interregional differences in child mortality contain a relatively stable component over time. High positive serial correlation in regional differences in mortality implies that cross-sectional observations on mortality in any single time period contain substantial information about the interregional differences that existed 5, 10, and perhaps even 20 years earlier (see Griliches [1961] for a discussion of this type of bias).

The estimates from cross sections using discretely lagged levels will tend to understate long-run responses because a single annual observation approximates the appropriate weighted distribution of past observations with substantial error. On the other hand, a discrete 3-year lag overstates short-run responses and creates the erroneous impression that the estimated response would occur in 3 years, whereas in all likelihood it would take longer. To improve the temporal specification and estimation of the dynamic behavioral relationship accounting for reproductive behavior, one needs the combined information of time series and cross sections. Below, I report estimates from the pooled time series of cross sections, assuming a Nerlovian two-component model of the stochastic structure of the disturbance (Nerlove 1971*a*). Then, first-differences of the cross section are estimated and related to disequilibrium in the family formation process.

A Two-Component Model of Disturbances

The statistical properties of the estimates of equation (1) are determined by the nature of the disturbances in the model, e_{it}, where i refers to region. The disturbance term presents both the net effects of numerous factors that have unavoidably been omitted from the analysis and errors of measurement and approximation in the form of the behavioral relationship. It is reasonable to assume that many of these effects are specific to regions and relatively time-invariant. As a first approximation, then, the disturbance term might be decomposed following Nerlove (1971*b*) into two independent elements, a region-specific time-invariant effect, μ_i, and a region- and time independent effect, v_{it}. The stochastic structure for the disturbance in equation (1) might then be expressed:

$$E\, e_{it}e_{i't'} = \begin{cases} \sigma^2 = \sigma_\mu{}^2 + \sigma_v{}^2 & , i = i', t = t' \\ \sigma_\mu{}^2 & , i = i', t \neq t' \\ 0 & , \text{otherwise,} \end{cases}$$

$$E\, e_{it} = 0, \text{ all } i \text{ and } t.$$

Let the parameter $\rho = \sigma_\mu{}^2/\sigma^2$ be defined as the proportion of the variance of the disturbances accounted for by the region-specific component. It may be shown (Nerlove 1971*a*) that generalized least squares for a model with this form of variance-covariance matrix amounts to using transformed values of the variables, which are a weighted combination of the original observations and the deviations from regional means. These weights can be expressed as a simple function of ρ. Several methods for estimating ρ have been considered; the two-stage method used here appears to show least bias, least mean-square error, and greatest overall robustness against specification error (Nerlove 1971*b*). The procedure has been used by Schultz (1967) and Nerlove and Schultz (1970) in the study of birth rates in Puerto Rico.

Fitting equation (1) to the pooled time series of cross sections, weighted estimates are reported in table 4, part A for the levels of the original variables, in table 4, part B for the deviations of these variables from regional means, and in table 4, part C for the transformed variables based on the values of ρ reported in the last column of the table.

Between 57 and 84 percent of the variation in the residuals from the pooled regression on the original variable levels is attributable to the region-specific component (see values of ρ in table 4, part C). The estimates based directly on the variable levels are similar whether analysis is limited to individual years (table 3) or the pooled timed series of cross sections (table 4, part A). Additional information contained in the time-series dimension of these data is extracted only when the relative importance of the two disturbance components is estimated and used to obtain the transformed variable estimates of the model parameters. In the total fertility equation based on the transformed variables, the coefficients for the child-survival factor and male schooling remain significantly different from zero and of the appropriate sign, but their size is about half of those implied by the analysis of levels. This confirms that *direct cross-sectional estimates of short-run responses are seriously biased upward.* The total fertility effect from agricultural composition is no longer significant, although there is a tendency for agricultural regions to have higher birth rates for women between the ages of 25 and 29 and somewhat lower birth rates thereafter.

Of greater importance is the shift in sign of the coefficients estimated for female schooling. In the total-fertility-rate equation, female schooling now depresses fertility as does male schooling, but elasticities calculated at regression means are still greater for male than for female schooling (—.32 versus —.17). In all but the teen-age birth-rate equation, female schooling is associated with lower birth rates. Although the puzzling behavior of the estimates for female schooling has been partially resolved by the analysis of pooled time series of cross sections, the economic prediction that the coefficient on women's schooling, as a proxy for the value of their time, should algebraically be less than the coefficient on men's schooling is confirmed by the transformed estimates *only* for the birth-rate equation for women aged 25–29.

The transformed variable estimates based on the time series of cross sections also indicate that family-planning information has twice the effect on birth rates as that implied by the estimates based on the levels, reducing total fertility rates 8 percent rather than 4 percent. This should also have been anticipated, for these program inputs are not subject to the same dynamic specification errors that biased upward the coefficients on gradually changing environmental constraints. Among all women over the age of 24, birth rates are significantly inversely related to the allocation of both classes of family-planning field personnel. Among the younger women,

TABLE 4

REGRESSION ON POOLED TIME SERIES OF CROSS SECTIONS

A. LEVELS, 1964–69

DEPENDENT VARIABLE	CONSTANT TERM	CHILD-DEATH ADJUSTMENT (RATIO)	AGRICULTURAL COMPOSITION (PROPORTION)	MALE EDUCATION (PROPORTION)	FEMALE EDUCATION (PROPORTION)	FAMILY-PLANNING PROGRAM MAN-MONTHS (CUMULATIVE INPUTS TO PRIOR YEAR) VHEN	PPHW	R^2/SEE
Total fertility rate	−2,246. (4.40)	9,067. (20.7)	234. (4.17)	−4,500. (16.5)	714. (3.74)	−23.9 (0.94)	−38.3 (7.86)	.567 (544.)
Age-specific birth rates for ages:								
15–19	−87.0 (5.04)	182. (12.3)	−14.2 (7.72)	−171. (18.9)	110. (17.4)	6.14 (7.64)	1.27 (8.23)	.237 (17.8)
20–24	56.4 (1.45)	303. (9.03)	−5.88 (1.36)	−182. (8.95)	4.66 (0.34)	19.9 (5.05)	−.0592 (0.07)	.242 (41.2)
25–29	116. (3.71)	220. (8.22)	37.3 (10.8)	53.0 (3.19)	−167. (14.3)	−1.48 (0.94)	−1.34 (4.41)	.482 (32.8)
30–34	−129. (3.91)	408. (14.5)	10.7 (2.90)	−188. (10.4)	23.8 (1.90)	−7.53 (4.48)	−4.18 (13.0)	.424 (35.6)
35–39	−243. (8.61)	432. (17.9)	10.3 (3.26)	−247. (16.2)	98.2 (9.23)	−8.42 (5.98)	−3.53 (13.0)	.461 (30.4)
40–44	−148. (8.97)	232. (16.6)	9.27 (5.03)	−130. (14.5)	57.1 (9.04)	−9.18 (5.71)	−3.51 (10.1)	.414 (17.9)
45–49	−6.09 (1.04)	24.7 (4.97)	−.740 (1.18)	−33.9 (10.4)	20.2 (8.91)	−.999 (3.30)	−.175 (3.12)	.092 (6.47)

TABLE 4 (*Continued*)

B. Deviations from Regional Means, 1964–69*

Dependent Variable	Child-Death Adjustment (Ratio)	Agricultural Composition (Proportion)	Male Education (Proportion)	Female Education (Proportion)	Family-planning Program Man-Months (Cumulative Inputs to prior Year): VHEN	Family-planning Program Man-Months (Cumulative Inputs to prior Year): PPHW	R^2/SEE
Total fertility rate	2,142. (6.53)	−918. (2.19)	−2,159. (5.62)	−2,875. (9.68)	−219. (9.85)	−56.1 (12.3)	.568 (271.)
Age-specific birth rate for ages:							
15–19	−16.5 (1.43)	−14.4 (1.01)	5.41 (0.40)	23.1 (2.21)	.494 (0.64)	.113 (0.71)	.026 (9.29)
20–24	−23.3 (0.73)	−56.4 (1.37)	−9.44 (0.26)	1.68 (0.06)	−.250 (0.12)	−1.33 (2.95)	.008 (26.3)
25–29	112. (3.87)	−5.81 (0.16)	−4.84 (0.14)	−140. (5.44)	−9.71 (4.98)	−1.90 (4.75)	.192 (23.5)
30–34	152. (6.33)	−58.9 (1.87)	−159. (5.56)	−210. (9.58)	−13.5 (8.20)	−3.31 (9.90)	.519 (20.0)
35–39	132. (6.47)	−23.2 (0.89)	−185. (7.70)	−136. (7.30)	−14.1 (10.3)	−3.02 (10.7)	.549 (16.9)
40–44	67.4 (5.43)	−16.3 (1.01)	−65.7 (4.48)	−72.5 (6.39)	−6.66 (7.91)	−1.87 (10.8)	.447 (10.4)
45–49	10.9 (1.96)	−5.46 (0.74)	−12.5 (1.85)	−24.9 (4.73)	−.316 (0.82)	.0102 (0.13)	.071 (4.73)

* Intercept suppressed.

TABLE 4 (*Continued*)

C. TRANSFORMED VARIABLES, 1964–69

Dependent Variable	Constant Term	Child-Death Adjustment (Ratio)	Agricultural Composition (Proportion)	Male Education (Proportion)	Female Education (Proportion)	Family-planning Program Man-Months (Cumulative Inputs to Prior Year) VHEN	PPHW	R^2/SEE	$\hat{\rho}$
Total fertility rate ..	4,471. (9.83)	2,748. (8.06)	74.4 (0.53)	−2,454. (6.78)	−1,860. (6.91)	−218. (10.1)	−60.6 (14.5)	.553 (288.)	.840
Age-specific birth rates for ages:									
15–19	55.5 (3.52)	−2.22 (0.19)	−3.42 (0.76)	−45.1 (3.62)	35.4 (3.83)	1.85 (2.50)	.450 (3.19)	.027 (9.90)	.815
20–24	321. (7.70)	1.35 (0.04)	17.0 (1.50)	−88.1 (2.77)	−23.8 (1.01)	4.13 (2.04)	−.149 (0.38)	.027 (27.4)	.769
25–29	223. (6.35)	136. (4.76)	38.9 (5.47)	18.9 (0.74)	−151. (8.12)	−7.71 (4.45)	−1.81 (5.64)	.271 (24.8)	.566
30–34	144. (4.43)	192. (7.79)	−13.3 (1.47)	−136. (5.24)	−110. (5.79)	−15.6 (9.92)	−4.37 (14.6)	.494 (21.2)	.788
35–39	64.4 (2.32)	172. (8.19)	−8.80 (1.10)	−174. (7.88)	−45.4 (2.78)	−16.1 (12.2)	−3.92 (15.4)	.523 (17.9)	.809
40–44	2.10 (0.13)	93.2 (7.33)	3.14 (0.71)	−73.3 (5.59)	−18.6 (1.94)	−7.46 (9.28)	−2.25 (14.7)	.426 (11.0)	.763
45–49	1.10 (0.16)	16.5 (3.02)	−2.67 (1.84)	−17.0 (3.18)	−1.15 (0.30)	−.978 (2.81)	−.205 (3.20)	.058 (4.96)	.619

the reverse is indicated. On balance, the program's effect has been to reduce the number of births even though it may accelerate the "tempo" of fertility among some younger women, at least in the short run.

First-Difference Model

Another approach to the analysis of time series where omitted variables are thought to bias estimates of short-run response is to assume that region-specific, relatively time-invariant effects can be removed by first-differencing the basic model over time. Assume a model of the form

$$b_{it} = \alpha_0 + \sum_{k=1}^{n} \alpha_k X_{ikt} + \sum_{j=1}^{m} \beta_j Y_{ijt} + e_{it}, \qquad (2)$$

where the X_{ikt} are n observed short-run determinants of birth rates and the Y_{ijt} are m unobserved long-run determinants of birth rates, and e_{it} is a normal random error independently distributed with respect to time and region. If the Y_{ijt} do not change, say within a 5-year period, but tend to be correlated with specific X_{ikt}, which do change, then the cross-sectional regression on levels will spuriously attribute the effect of Y_j's to X_k's and bias estimates of the short-run response of birth rates to the observed short-run factors.

Absolute differences between cross sections several years (δ) apart provide, then, another test of the model's specification and a further procedure for evaluating how rapidly reproductive behavior responds to specific short-run changes in economic and demographic determinants of the desired number of children. The estimated equation becomes

$$b_{it} - b_{i.t-\delta} = \sum_{k=1}^{n} \alpha_k (X_{ik.t-\tau} - X_{ik.t-\tau-\delta}) + V_i,$$

where $V_i = e_{it} - e_{i.t-\delta}$.

Using the maximum available value of δ for the Taiwan data of 5 years, this reformulation of equation (1) states that only changes experienced in the explanatory variables from 1961 to 1966 (or to 1968 for family-planning inputs) affect changes in birth rates from 1964 to 1969.[12] This

[12] Because of greater year-to-year variability of a stochastic nature in child-death rates estimated for small communities, the first-differenced child-death adjustment factor is based on 2-year averages at the beginning and end of the 5-year period, i.e., $[(D_{1961} + D_{1962}) - (D_{1966} + D_{1967})]/2$. Death rates were first published in 1961.

procedure sharply reduces multicollinearity among the explanatory variables (see table 2) and the statistical significance of regression coefficients for the reasons mentioned earlier (see also Fisher 1962). Comparisons across methods of estimation may rely on the standard error of estimate (SEE), given the maintained model specification, since when the intercept is forced through the origin no obvious interpretation attaches to R^2.[13]

The weighted estimates of the first-differenced form of equation (1) are reported in table 5, part A. The child-survival factor is positively associated with birth rates to a statistically significant extent *only* among women aged 30–44—those in the groups thought to be most responsive, in the short run, to such changes in the regime of child mortality. Changes in agricultural composition of the labor force are not apparently related to changes in birth rates. For the first time the coefficient of male schooling is positive for birth rates between ages 20 and 29, which is consistent with expectations that the income effect would exceed the substitution effect for the value of male time. Women's schooling may increase birth rates from ages 15 to 24, but it substantially reduces those rates thereafter. From 25 years of age and over, the algebraic value of the coefficient on women's schooling is less than that on men's schooling, as expected on economic grounds. The family-planning personnel may contribute to a slight increase in birth rates among women aged 15–19, but reduce birth rates substantially among women aged 25–44, as anticipated.

The predictions of the general economic model that can be tested with these data for Taiwan are confirmed when the initial static model is differenced over time, although the magnitudes of the response estimates are smaller than those based on the individual cross sections and similar to those based on the combined time series of cross sections (see table 6). Since first-differencing an economic behavioral relationship is often a severe econometric test for specification error, I conclude from this evidence that reproductive behavior does systematically respond in the *short*

[13] Indeed, in half of the regressions reported in table 5, the standard error of estimate exceeds the standard deviation of the dependent variable. This may appear to cast doubt on the maintained hypothesis that the intercept of the first-differenced equation equals zero or, in other words, that the birth-rate equation does not shift over time independently of the six specified explanatory variables. Relaxing this constraint and permitting the intercept to assume a nonzero value tends to reduce somewhat the size of the regression coefficients and their t-statistics, but not their signs or relative magnitudes. If the hypothesis of the zero intercept is maintained, a summary statistic in place of R^2 might be defined with respect to both the variance of the (first-differenced) dependent variable and the variance of the level of the appropriate birth rate. Hence, $S^2 = 1 - (\Sigma V_i^2)/\{N\ [\mathrm{var}\ (b_{it} - b_{i,t-\delta}) + \mathrm{var}\ (b_{it})]\}$, where S is the new summary statistic, N is the number of observations, and var() is the variance of the respective variables. In this case, for example, the first-differenced relationship estimated over the total sample for the birth rate of women aged 30–34 yields an $S^2 = 0.592$, where the $R^2 = -.062$ (see table 5, part A). Relaxing the maintained hypothesis that the intercept is zero, the standard error of estimate falls to 31.97 and $S^2 = 0.647$ and $R^2 = .0812$.

TABLE 5

REGRESSION ON ABSOLUTE DIFFERENCES OF BIRTH RATES: 1969 LESS 1964

Dependent Variable	Child-Death Adjustment (Ratio)	Agricultural Composition (Proportion)	Male Education (Proportion)	Female Education (Proportion)	Family-planning Program Man-Months (Cumulative Inputs to Prior Year) VHEN	PPHW	R^2/SEE
				A. Total Sample			
Total fertility rate	4,748. (5.14)	−680. (1.15)	−922. (1.65)	−2,482. (5.72)	−248. (6.41)	−66.0 (8.63)	−.519 (455.)
Age-specific birth rates for ages:							
15–19	−27.7 (0.80)	−14.8 (0.68)	−10.9 (0.51)	24.2 (1.45)	1.67 (1.18)	−.0638 (0.22)	.022 (17.0)
20–24	92.1 (0.95)	−68.4 (1.12)	18.1 (0.33)	23.1 (0.55)	−3.85 (0.96)	−.753 (0.95)	−.024 (46.5)
25–29	102. (1.30)	19.5 (0.40)	65.5 (1.43)	−102. (2.87)	−12.1 (3.75)	−3.46 (5.41)	−.101 (37.1)
30–34	284. (4.05)	−25.4 (0.55)	−104. (2.46)	−196. (6.02)	−14.2 (4.84)	−3.71 (6.49)	−.062 (34.3)
35–39	330. (5.33)	−25.7 (0.65)	−103. (2.73)	−141. (4.80)	−13.7 (5.37)	−3.46 (6.82)	.052 (30.7)
40–44	168. (4.08)	−16.2 (0.59)	−41.6 (1.64)	−56.2 (2.87)	−7.30 (4.21)	−2.19 (6.38)	.025 (20.7)
45–49	5.30 (0.28)	−8.88 (0.66)	−7.49 (0.59)	−29.2 (2.93)	−1.32 (1.53)	.0500 (0.29)	.027 (10.2)

TABLE 5 (*Continued*)

Dependent Variable	Child-Death Adjustment (Ratio)	Agricultural Composition (Proportion)	Male Education (Proportion)	Female Education (Proportion)	Family-planning Program Man-Months (Cumulative Inputs to prior Year)		R^2/SEE
					VHEN	PPHW	
B. Subsample for Which 1964 Birth Rates Tended to be Less than Predicted by the Cross-sectional Model							
Total fertility rate	3,895. (2.77)	−1,219. (1.33)	−74.3 (0.08)	−3,297. (4.65)	−161. (2.79)	−65.4 (5.63)	−.795 (432.)
Age-specific birth rate for ages:							
15–19	−82.1 (2.46)	−57.1 (2.75)	−47.9 (2.42)	32.0 (1.89)	2.94 (2.36)	.775 (2.95)	.155 (10.3)
20–24	84.7 (0.64)	−249. (2.54)	−52.8 (0.68)	32.0 (0.54)	14.6 (2.44)	.0157 (0.01)	.095 (46.3)
25–29	−11.8 (0.13)	−11.9 (0.15)	4.35 (0.07)	−155. (2.81)	.780 (0.16)	−.967 (1.02)	−.057 (34.2)
30–34	270. (2.90)	−6.60 (0.11)	−141. (2.59)	−157. (3.43)	−9.14 (2.52)	−2.10 (2.71)	−.062 (30.9)
35–39	236. (3.80)	44.1 (1.11)	−80.2 (2.13)	−160. (4.78)	−9.52 (3.89)	−2.24 (4.61)	.309 (20.5)
40–44	112. (3.60)	−13.5 (0.60)	−46.6 (2.42)	−46.9 (2.99)	−3.91 (3.30)	−1.42 (5.34)	.132 (10.3)
45–49	−16.3 (1.87)	−1.81 (0.23)	1.37 (0.19)	2.33 (0.39)	−.332 (0.70)	−.125 (1.36)	.020 (3.57)

TABLE 5 (*Continued*)

Dependent Variable	Child-Death Adjustment (Ratio)	Agricultural Composition (Proportion)	Male Education (Proportion)	Female Education (Proportion)	Family-planning Program Man-Months (Cumulative Inputs to prior Year) VHEN	PPHW	R^2/SEE
	C. Subsample for Which 1964 Birth Rates Tended to be More than Predicted by the Cross-sectional Model						
Total fertility rate	5,905. (4.87)	—54.7 (0.07)	—2,298. (3.09)	—1,807 (3.27)	—302. (5.68)	—63.6 (6.03)	—.336 (464.)
Age-specific birth rate for ages:							
15–19	29.5 (0.53)	11.0 (0.31)	—22.3 (0.61)	39.1 (1.48)	.884 (0.36)	—.884 (1.90)	.020 (20.8)
20–24	185. (1.53)	12.8 (0.19)	—34.7 (0.52)	—11.4 (0.22)	—14.6 (3.12)	—.283 (0.31)	—.021 (39.4)
25–29	485. (4.19)	10.9 (0.20)	91.4 (1.63)	—112. (2.78)	—15.2 (3.96)	—4.15 (5.47)	—.124 (33.1)
30–34	356. (3.92)	—53.5 (0.88)	—77.6 (1.39)	—241. (5.91)	—16.8 (4.04)	—5.61 (7.61)	—.020 (32.3)
35–39	460. (4.62)	—42.6 (0.66)	—187. (2.95)	—108. (2.45)	—13.6 (3.15)	—4.97 (5.66)	—.328 (36.4)
40–44	237. (3.49)	21.7 (0.50)	—95.4 (2.23)	—60.1 (1.92)	—9.37 (2.92)	—2.64 (4.68)	.001 (25.5)
45–49	68.8 (1.81)	1.00 (0.05)	—32.9 (1.55)	—28.3 (1.78)	—1.99 (1.38)	—.0119 (0.04)	.090 (12.8)

TABLE 6

ELASTICITY OF BIRTH RATES WITH RESPECT TO EXPLANATORY VARIABLES

		Estimation Technique and Sample Composition				
		Time Series of Cross Sections		Cross Section of First-Differences		
Dependent Variable	Independent Variable	Levels (1)	Transformed (2)	Total (3)	"Equilibrium" (4)	"Disequilibrium" (5)
Total fertility rate	Child-death adjustment*	1.9	.58	1.0	.85	1.2
Total fertility rate	Female schooling	.064	—.17	—.22	—.30	—.15
Total fertility rate	Male schooling	—.59	—.32	—.12	—.01	—.27
Birth rate, 35–39†	Child-death adjustment*	3.9	1.5	3.0	2.6	3.3
Birth rate, 35–39†	Female schooling	.37	—.17	—.53	—.76	—.32
Birth rate, 35–39†	Male schooling	—1.4	—.98	—.58	—.57	—.81

SOURCES.—Evaluated at regression means in 1964. Coefficients for col. 1 from table 4, part A; col. 2, table 4, part C; col. 3, table 5, part A; col. 4, table 5, part B; and col. 5, table 5, part C.

* The average child-death adjustment ratio in 1961 (3-year lag appropriate for the 1964 birth-rate regression) was about 1.073, which implies a child-death rate to age 15 of about 68 per 1,000 live births, i.e., child-death adjustment ratio = [1/(1-child-death rate)]. The elasticity of birth rates with respect to the child-death rate would be 0.073 times the elasticity with respect to the child-death adjustment ratio reported above.

† Age group of women.

run to the evolving configuration of economic and demographic constraints that I could observe in Taiwan.

Disequilibrium and Change in Birth Rates

If elements of an economic theory of fertility have some empirical validity in predicting long-run desired equilibrium levels of fertility toward which parents gravitate, then unexplained deviations of birth rates from those predicted by the theory should contain information about the magnitude of reproductive disequilibrium present across populations. In a period of secularly decreasing fertility, which has been pronounced in Taiwan among women over the age of 30 since the mid-1950s (see fig. 2), positive residuals in a cross-sectional regression of birth-rate levels should constitute evidence of disequilibrium and perhaps imply a relative inability to control reproductive behavior at a tolerable individual cost, to correspond to the rapidly changing environment. More specifically, let me propose that in regions in which actual birth rates (particularly among older women) exceed the model's predicted birth rates, the information disseminated by the family-planning program is likely to be in greater demand. These disequilibrium regions are likely to experience more substantial declines in birth rates as a function of the subsequent allocation of family-planning personnel.

A second, more tentative, hypothesis would suggest that where positive disequilibrium was substantial initially, a greater response of birth rates could be expected to subsequent changes associated with a further reduction in desired birth rates. This conjecture would be consistent with the assumption that the rate of adjustment in birth rates was a positive function of the extent of disequilibrium between actual birth rates in the previous period and desired birth rates in the current period.

To test these simple concepts of disequilibrium and change in birth rates, the 361 regions of Taiwan were divided in half, based on the algebraic size of their residuals from regressions on 1964 birth-rate levels (not reported). The absolute differences in birth rates between 1964 and 1969 were then regressed on the differenced explanatory variables. The regression results for the half of the sample with primarily negative residuals (below the regression plane), which I will call the "equilibrium" regions, and those for the sample with primarily positive 1964 residuals (above the regression plane),[14] which will be called the "disequilibrium" regions, are shown in table 5, parts B and C, respectively.

The first hypothesis is confirmed; the regression coefficients for the family-planning personnel variables are algebraically smaller for the disequilibrium regions than for those regions presumed to be closer to equilibrium. In six of the 10 possible comparisons for the birth-rate equations for women between the ages of 20 and 44, the regression coefficients differ significantly in the anticipated direction between the two subsamples, based on a one-tailed t-test at the 5 percent confidence level. It is also interesting to note that the effect of the program personnel *to increase* the teen-age birth rates is also greater across the disequilibrium regions.

The second hypothesis is more difficult to assess, largely because the data are less appropriate. To test the stock-adjustment model, information is required on the initial stocks or number of living children of women of specific ages. Since these data are not published by small regions in Taiwan, I have assumed that the regional variation in birth rates in 1964 closely parallels the regional variation in completed fertility at that date, and positive discrepancies between observed and predicted birth rates reflect in part "unwanted" births. Changes in child mortality elicit a 50 percent greater response in birth rates in the disequilibrium regions than they do across the equilibrium regions (see table 6). Male and female schooling coefficients for the equilibrium regions conform to the pattern observed in most cross-sectional studies of fertility; female schooling is the more significant and sizable deterrent to high birth rates among women

[14] Of course, to obtain subsamples of approximately equal size, i.e., 180 and 181, the dividing line between the two groups of residuals does not turn out to be exactly zero.

aged 25–44. In the total-fertility equation, the schooling of men is hardly associated with birth rates, whereas the elasticity of total fertility rates with respect to women's schooling is —0.30. In the disequilibrium regions, the implied effects of changes in male schooling achievement are greater than those of changes in female schooling; the elasticity of total fertility with respect to male schooling is —0.27 and for female schooling, —0.15. This reversal in the relative importance of schooling of men and women on fertility between the equilibrium and disequilibrium subsamples is also evident in the older age-specific birth-rate equations, as illustrated in table 6 with reference to women 35–39 years old. Although these tentative findings may be rationalized, they raise more questions in my mind than they answer. Explicit stochastic models of dynamic behavior are now needed to make sense of these results and to proceed with the analysis of other time-series evidence on reproductive behavior that will permit economic analysis to be beyond the static notion of long-run equilibrium.

VI. Conclusions

Age-specific birth rates for a cross section of small communities over time were analyzed to infer the responsiveness of birth rates in Taiwan to changing economic, demographic, and policy constraints. There are sound reasons to also consider information on individual reproductive behavior over time, since most models of fertility determination presuppose a relation between fertility and the number of living children parents have and the immediate and lifetime circumstances of the parents. Lacking longitudinal data for individuals, I have explored two methods for extracting information about the dynamic process of birth-rate determination from an integrated treatment of a time series of aggregate cross sections.

Two issues were raised. The first is the traditional concern of economists with separating short-run reactions from long-run adjustments toward equilibrium. Evidence from the analysis of the time series of cross sections, assuming a two-component variance model and the first-differenced cross section over time, implied that estimates based on a single cross section are seriously biased. Slowly changing constraints in the parents' environment, such as the regime of child mortality, agricultural composition, and male schooling, are attributed, in the cross section, an exaggerated and distorted role in affecting birth rates. Both approaches to time series and cross-sectional variation in birth rates, however, continue to imply that child mortality and adult schooling exert strong and statistically significant effects on birth rates within the relatively short 3–5-year time horizon examined here.

More curious, the effect of female schooling is obscured entirely in the cross-sectional estimates, whereas either approach to the time series provides indications that increased schooling for women in Taiwan is associated with markedly lower birth rates, especially among older women who are concluding their childbearing. Also, the effect of the family-planning program on birth rates is biased downward by about 50 percent in cross-sectional estimates compared with those obtained from the time series. Such important shifts in the model's parameters do not give one confidence in analyses that stop with estimates from a single cross section.

The second issue is how to treat disequilibrium in reproductive behavior within an economic framework, especially where this phenomenon must be quantitatively important, as in low-income countries in which fertility has been high and birth rates are beginning to decline rapidly. Evidence was presented that a rudimentary theory of variation in birth rates might help to isolate communities where birth rates were atypically high given the configuration of environmental constraints. Allocations of family-planning field personnel were shown to have exerted twice the effect in reducing birth rates among older women in these "disequilibrium" communities than in the "equilibrium" communities. Evidence of "disequilibrium" in reproductive behavior derived from analyses of residuals might prove to be a useful method, therefore, for the stratification and study of change in fertility over time, and as a guide for the efficient allocation of policy resources.

Priority should be given to extending the static equilibrium theory of household time allocation and decision making to allow the introduction of dynamic elements of innovation, search and information costs, the biological and behavioral constraints on the supply of children, and the longitudinal complexities of intergenerational savings and transfers. Theoretical progress in these directions and more extensive empirical analyses of aggregate and individual time series might improve substantially our understanding of both the economic and demographic behavior of the household sector during the process of economic development.

TABLE A1

SUMMARY OF EMPIRICAL RESULTS FROM STUDIES OF ECONOMIC DETERMINANTS OF FERTILITY IN LESS DEVELOPED COUNTRIES

Country/Time	Equation/ Data Type	Dependent Variable	Independent Variables: Adult Education			Wage Rate		Family Income	Child Schooling	Child Labor	Child Mortality	Rural or Agriculture	Sample Size
			Male (1)	Female (2)	All (3)	Male (4)	Female (5)	(6)	(7)	(8)	(9)	(10)	
1. Puerto Rico, 1950–57	Single aggregate	Crude birth rate	...	...	—**	...	...	—*	—**	+**	+**	+	75
2. Egypt, 1960	Reduced form aggregate	Child/woman ratio	...	—[a]	...	...	...	...	...	+[a]	...	—[a]	41
3. Israel,[b] 1961													
Jewish non-kibbutz	Single aggregate	Adjusted birth rate	±	—**	...	...	...	...	...	...	...	=	431
Kibbutzim	Single aggregate	Adjusted birth rate	+	+	...	...	...	...	...	...	...	...	180
Non-Jewish	Single aggregate	Adjusted birth rate	+	—**	...	...	...	...	...	...	...	—**	133
4. Philippines, 1968	Structural micro	Children born to women, 35–39[c]	...	—*	...	...	...	...	...	+	++**	+**	250

SOURCES.—Study 1, Schultz 1969, p. 175; study 2, Schultz 1970, p. 44; study 3, Ben-Porath 1970*a*, p. 30; study 4, Harman 1970, p. 29-D-1; study 5, Nerlove and Schultz 1970, p. 45; study 6, Schultz 1971*a*, p. 61; study 7, Da Vanzo 1971, p. 78; study 8, Maurer, Ratajczak, and Schultz 1972, table 9; and study 9, Schultz 1972*b*.

[a] No confidence intervals are reported for reduced-form coefficients derived from estimates of structural equations.

[b] Treated as endogenous and estimated with simultaneous-equations techniques.

[c] Fertility for women aged 35–39 or 35–44 is used for comparative purposes.

* Regression coefficient exceeded 1.65 times its standard error, i.e., 5 percent one-tailed *t*-test of statistical significance.

** Regression coefficient exceeded 2.37 times its standard error, i.e., 1 percent one-tailed *t*-test of statistical significance.

TABLE A1 (*Continued*)

			Independent Variables										
			Adult Education			Wage Rate							
Country/Time	Equation/ Data Type	Dependent Variable	Male (1)	Female (2)	All (3)	Male (4)	Female (5)	Family Income (6)	Child Schooling (7)	Child Labor (8)	Child Mortality (9)	Rural or Agriculture (10)	Sample Size
5. Puerto Rico, 1950–60	Structural aggregate	Crude birth rate	...	...	−**	...	...	+	−	−**	+**	...	825
6. Taiwan, 1964–68	Single aggregate	Birth rate for women, 35–39[c]	...	±	...	...	...	...	−**	...	+**	±	361
7. Chile, 1960	Structural aggregate	Children born to women, 35–39[c]	...	...	+	−*[b]	...	...	...	+*[b]	+**	...	50
8. Thailand, 1960	Reduced form aggregate	Children born to women, 35–39[c]	+[a]	−[a]	...	−[a]	...	...	−[a]	+[a]	...	...	71
9. Bogota, Colombia, 1965	Structural micro	Children present, women, 30–34	−	+*	−[b]	−*[b]	+	...	...	...	...	...	63

TABLE A2

SUMMARY OF EMPIRICAL RESULTS FROM STUDIES OF ECONOMIC DETERMINANTS OF FERTILITY IN UNITED STATES

			Sample/Time							
			Adult Education		Wage or Earnings					
Independent Variables	Equation/ Data Type	Dependent Variable	Male (1)	Female (2)	Male (3)	Female (4)	Family Income (5)	Nonhuman Wealth (6)	Interaction (2)·(3) (7)	Sample Size
1. 1960 census, SMSA:										
White	Single	Children born per		—*a	+*	—**	...	...	...	100
Nonwhite	aggregate	married woman,		—**a	+*	—**	...	...	...	
		35–44								
2. 1960 census grouped:										
Married 14–21	Single	Children born per		—**	—	...	...	...	...	480
Married 22 years or more ...	aggregate[b]	married woman,		—**	—**	...	...	...	...	
		35–44								
3. 1940 and 1960, grouped census:										
	Single	Children born per								
	aggregate	married woman,								
Without interaction	1940	40–44[c]		—**	+*	...	...	...	...	35
Without interaction	1960	35–44[c]		—**	+	...	...	...	...	98
With interaction	1940	40–44[c]		—**	—**	...	...	...	...	35
With interaction	1960	35–44[c]		—**	—**	...	...	...	+**	98
4. 1960 census states:										
Rural-farm	Single	Children born per	+**	—**	...	—**	+	...	...	40
Urban	aggregate	married woman,	+	—	...	—**	+**	...	...	
		40–44								

SOURCES.—Study 1, Cain and Weininger 1967 (rev. 1971), tables 3 and 4; study 2, Willis 1971, p. 61a; study 3, Sanderson and Willis 1971, tables 2 and 3; study 4, Gardner 1972, table 2; study 5, De Tray 1972*a*, p. 36; study 6, Michael 1974; study 7, Schultz 1972*a*.

[a] The education variable was the percentage of women 25 years of age and over with *less than* 5 years of schooling. The sign of the regression coefficient is reversed here for comparative purposes.

[b] Equation is fit in logarithmic form in terms of dependent fertility variable, and those explanatory variables measured in money terms, i.e., incomes.

[c] Women's potential wage was constructed from 1959 earnings of employed wives adjusted for hours worked per week (see Willis 1971, pp. 117 ff.).

[d] Median value of housing used as a proxy for nonhuman wealth (see De Tray 1972*a*, p. 29).

[e] South and nonsouth residence at age 16 was used to divide black population according to the quality of schooling received.

[f] Treated as endogenous and estimated with simultaneous-equations techniques.

* Regression coefficient exceeded 1.65 times its standard error, i.e., 5 percent one-tailed t-test of statistical significance.

** Regression coefficient exceeded 2.37 times its standard error, i.e., 1 percent one-tailed t-test of statistical significance.

TABLE A2 (*Continued*)

			Sample/Time							
			Adult Education		Wage or Earnings					
Independent Variables	Equation/ Data Type	Dependent Variable	Male (1)	Female (2)	Male (3)	Female (4)	Family Income (5)	Non-human Wealth (6)	Interaction (2)·(3) (7)	Sample Size
5. 1960 census counties	Single aggregate[b]	Children born per married woman, 35–44	+	—**	+*	—**	...	+*[d]	...	519
6. 1968 suburban household survey	Single micro	Children born per married woman, 35–39	—	—*	...	...	+*	...	...	513
7. 1967 survey of economic opportunity:		Children born per married woman,								
White married, husband present	Structural micro	35–39	—**	—	+**[f]	—**[f]	...	—	...	1,098
Black married, husband present	Structural micro	35–39	—**	—	—[f]	—**[f]	...	—	...	475
Black not resident in South at age 16[e]	Structural micro	35–39	—**	+**	+**[f]	—**[f]	...	—	...	158
White married, husband present	Structural micro	40–44	—**	+	+**[f]	—**[f]	...	+	...	1,138
Black married, husband present	Structural micro	40–44	+	+	—**[f]	—[f]	...	—	...	441
Black not resident in South at age 16[e]	Structural micro	40–44	—**	+**	+**[f]	—**[f]	...	—	...	116

Comment

James Tobin

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Paul Schultz's paper provides a very useful summary of current economic hypotheses about human fertility, with succinct commonsense versions of their theoretical rationales. I was grateful also for his account of past empirical attempts to test these hypotheses and estimate their parameters. The paper then turns to the regional time series available for Taiwan, which Schultz analyzes with great ingenuity and methodological sophistication.

Over the period 1964–69 in Taiwan, fertility declined about 20 percent. Age-specific birth rates fell in all age brackets except the youngest, 15–19. The declines, in percentage of the 1966 levels, were systematically related to age, ranging from 3 percent for 20–24-year-old women to 75 percent for 40–44-year-old women. The association of larger relative declines with higher age is virtually monotonic.

How are these facts to be explained? Certain possible explanatory variables have been trending in the "right" directions. Public birth control program inputs have greatly risen. Child-survival rates have increased, diminishing the occasion for replacement births. The agricultural proportion of the population has diminished, and the conventional hypothesis is that industrialization and urbanization bring lower fertility. The educational attainments of women have increased; the hypothesis is that education raises the opportunity cost of childbearing and child rearing. On the other hand, male educational attainment has also grown, though by less than female, and the new standard hypothesis is that this trend should increase the birth rate.

Paul Schultz's study may be viewed as an attempt to see whether these possible explanatory trends do in fact account for the observed declines in birth rates. If they do, he reasons, the fertility declines should be most pronounced in those geographical areas where the trends are most pronounced. Schultz has districtwide data for 361 administrative regions.

His presumption that these are useful data is consistent with current population microeconomics. The presumption is that aggregate fertility is the summation of individual behavior and decision.

An alternative hypothesis would be that individual fertility responds to changing national, even international, cultural norms, which are not to be understood by studying individual or regional differences. I find it hard to imagine that differences between U.S. states or counties would do much to help us understand the decline in age-specific birth rates in this country in the 1960s, or their rise in the 1940s. But as an amateur in the field, I should eschew such speculations and return to considering Schultz's paper on its own grounds.

Unfortunately, table 5, part A, indicates that interregional differences in the listed explanatory trends did little to explain interregional differences in birth-rate declines. The standard errors of estimate of the regressions are as often as not larger than the standard deviations of the dependent variable. However, the signs of the coefficients generally conform to the hypotheses.

The most encouraging statistical results are those reported in table 5, parts B and C. Schultz divides the regional observations for each regression evenly between those for which birth rates were high in 1964 and those for which they were low in 1964. Of high 1964 birth-rate regions, those with the stronger birth control programs had the largest declines. Moreover, birth control inputs made more of a difference where the need and opportunity for them was greater, in the "high" rather than the "low" regions. With respect to the other explanatory variables, the partitioning seems to be of little significance.

I am sure Paul Schultz has already thought of better ways of testing the hypotheses that led him to this partitioning. It does make sense that the absolute change in birth rate depends on the initial rate—perhaps the initial rate for the next younger age group would be better—in order to capture cohort effects. It does make sense that the marginal efficacy of family-planning programs depends on the interaction of initial birth rates and program inputs. The initial levels of the independent variables, as well as that of the dependent variable, would enter if one had in mind a stock-adjustment model. But I am not sure I see the logic of using the residuals from the 1964 cross section as an interactive factor with *all* the independent variables in the regressions of 1964–69 differences.

A stock-adjustment model is what one would come up with if he really allowed himself to be infected by the spirit of this meeting, namely that children are durable goods that yield utility-generating services in amounts that depend on certain input flows. But before we ask Dale Jorgensen or his equivalent to apply neoclassical investment theory to these durable goods, we might remind him of some of their peculiar properties. They come in discrete integral lumps; they cannot be bought or sold in the used-child market or scrapped at will; the rental market is highly imperfect; delivery time is more than normally uncertain; their qualities are very uncertain ex ante, and ex post control of quality is quite limited; their own requirements and tastes alter the household utility function, and so on. Suppressing these concerns and looking at the problem as one of durable goods investment, I am led to some other questions about Schultz's specifications.

The first concerns the treatment of the survival rate. The rate used is the theoretical probability of survival to age 15, calculated from specific age-mortality rates for each year. Choice of this variable seems to imply that families calculate changes of survival ex ante, and beget extra children in advance to allow for expected attrition. Given that the mean survival rate appears to be 0.94 in 1966 and given the discreteness of choices available to an individual family, I find this description of behavior implausible.

In the text Schultz suggests a different scenario—specific replacement of lost children by subsequent births. The opposite hypothesis would be discouragement, in the literal sense of loss of heart. This matter could be tested directly if individual observations were available, but here Schultz has only regional aggregates. A low survival rate in the past may have contributed to an aggregate stock of children below desired stocks, but unexpected or regretted births may have contributed to the opposite. In any case, the logical variable would seem to be the number of living children per 1,000 mothers in the region, parceled out by age of mother if possible.

If the replacement scenario is correct, one could expect the elasticity of the birth rate with respect to the survival rate to be greater the older the mother. An older woman has less time to make the replacement. Schultz's tables do disclose some tendency in this direction.

My second question concerns the interpretation of the regional proportions of persons in agriculture or with education. Movements over time in these proportions occur mainly through the young, who adopt or attain different characteristics from their elders. How does the fact that the region is becoming less agricultural and better educated affect the behavior of families whose occupation and education were already fixed long ago? I would have more confidence in estimates of the effects of these variables on birth-rate declines if the significant coefficients were systematically concentrated on the younger age brackets.

In the level cross-section regressions, these proportions presumably represent mainly long-standing differences among regions, differences which may well affect family size targets. The appropriate dependent variable is then the *stock* of children per 1,000 women, age-corrected, not the birth rate. Clearly an interregional difference in such a proportion is not the same thing as an intertemporal change in the proportion for a given region. If region A is and always has been better educated than region B, the difference is diffused over all age brackets. If region B becomes better educated in 1969 than it was in 1964, the difference is concentrated in the younger brackets.

In a similar vein, it is one thing to find that educated women have fewer children than their less-educated contemporaries. It is another thing to expect that this difference will predict how much a general increase over time in women's education in a country or a region will diminish overall fertility. The second effect will be weaker than the first. This follows on purely economic grounds, since a general advance in educational attainment cannot increase women's wages as much as a similar advance concentrated on a few women puts them ahead of their peers. In any case, the opportunity cost of child rearing depends not only on the market value of women's time but also on the time required for child rearing, and education may diminish the latter at the same time that it raises the former.

We must be careful not to prove too much. Human reproduction will continue, I suspect, even if all women are college educated. At least equal in importance to the calculations of the new home economics, it seems to me, is the general social definition of the appropriate role of women, determining simultaneously how much schooling they get, how much work they do outside the home, and how many children they have.