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	Family-planning Program Man-Months (Cumulative Inputs to Prior Year): 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	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)		R^2/SEE	$\hat{\rho}$
						VHEN	PPHW		
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

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

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)		R^2/SEE
					VHEN	PPHW	
				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) VHEN	PPHW	R^2/SEE
	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)		R^2/SEE
					VHEN	PPHW	
	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

			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
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*)

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)	
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)	Non-human 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 aggregate	Children born per 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*)

Independent Variables	Equation/ Data Type	Dependent Variable	Sample/Time: Adult Education, Male (1)	Adult Education, Female (2)	Wage or Earnings, Male (3)	Wage or Earnings, 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:	Structural micro	Children born per married woman,								
White married, husband present		35–39	—**	—	+**[f]	—**[f]	...	—	...	1,098
Black married, husband present		35–39	—**	—	—[f]	—**[f]	...	—	...	475
Black not resident in South at age 16[e]		35–39	—**	+**	+**[f]	—**[f]	...	—	...	158
White married, husband present		40–44	—**	+	+**[f]	—**[f]	...	+	...	1,138
Black married, husband present		40–44	+	+	—**[f]	—[f]	...	—	...	441
Black not resident in South at age 16[e]		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.