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THE DETERMINANTS OF MARITAL FERTILITY IN THE UNITED STATES, 1968–1970: INFERENCES FROM A DYNAMIC MODEL

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Abstract—Criticizing the static assumptions of previous socioeconomic and microeconomic models of marital fertility, particularly regarding the sequential and stochastic facets of family building, this paper advocates a dynamic perspective. Of particular concern is the assumption of equilibrium family size made by those who employ the static perspective. The equilibrium family size assumption implies that the parameters relating social and economic variables to fertility will be similar for all births, regardless of order. To test this assumption of constancy, a two-equation model of fertility and female employment is introduced. Contrary to the static perspective's implication of constant effects, substantial parity differences in the estimates of parameters for both equations are reported, as are several differences between blacks and whites. On the basis of this evidence, I conclude that the static decision-making framework should be replaced by a dynamic approach to marital fertility.

EQUILIBRIUM FAMILY SIZE OR SUCCESSIVE APPROXIMATIONS?

Most economic models of marital fertility address the impact of household economics on *completed* family size. Becker's (1960) seminal paper refers to equilibrium family size. Other models based on the "new home economics" approach also seek to explain equilibrium family size, typically indicated by children ever born (Becker and Lewis, 1973; DeTray, 1973; Willis, 1973). Easterlin (1969, 1973, 1975) diverges from Becker and other scholars in the specification of economic relationships and in his emphasis on the number of surviving children. But he agrees with Becker on the existence of an equilibrium choice. Turchi's (1975a) socioeconomic model of the demand for children incorporates noneconomic constraints on family size decisions that are suggested by a number of critics (Duesenberry, 1960; Blake, 1967, 1968; Easterlin, 1969, 1975; Namboodiri, 1972; Ryder, 1973a; Cochrane, 1974; Leibenstein, 1974) but maintains the assumption that a long-run equilibrium family size exists.

The attention to a long-term reproductive response follows from the static decision-making framework employed in most economic analyses of fertility, described below: "a one-period comparative static framework . . . in which a husband and wife of given ages and characteristics are considered to adopt, at the outset of marriage, a utility-maximizing lifetime plan for childbearing, for expenditures of time and money on children, and for other sources of parental satisfaction not related to children" (Willis, 1973, p. S17).

From an alternative perspective, the birth of each child may be viewed as an event that "alters the parents' perception in a way they cannot entirely anticipate" (Rosenzweig, 1976, p. 339). This alternative perspective suggests shifts in the impact of socioeconomic variables on subsequent births. The perfectly general utility function underlying the static model allows both nonlinearities and interactions in the effects of the socioeconomic and demographic variables related to completed family size. However, the one-period decision-making framework in which the model is cast precludes the possibility that exogenous variables may interact with parity in the determination of fertility, because to include parity, a lagged endogenous variable, would be incompatible with the one-period specification.

This paper follows the trend toward dynamic specifications of reproduction (Namboodiri, 1972; Rosenzweig, 1976) and labor force participation (Heckman and Willis, 1977) by drawing attention to the sequential and stochastic elements of family formation omitted by the one-period models. The dynamic perspective focuses on the birth interval as a decision point and regards completed family size as the sum of a sequence of outcomes culminating in the decision to have no more children. It emphasizes the likelihood of parity differences in socioeconomic effects (Namboodiri, 1974; Rosenzweig and Seiver, 1975). This paper demonstrates the significance of the hypothesized parity interactions in the determination of U.S. fertility between 1968 and 1970.

SEQUENCE AND UNCERTAINTY

Sequence is important for family formation because, for most couples, children are acquired one after the other, not in lots. Unless a couple plans to remain childless or to have exactly one child, desired family size cannot be achieved in a single period. For couples initially desiring more than one child, two or more periods are required for the achievement of desired family size (unless a multiple birth occurs). Couples not attaining desired family size in a single period have the opportunity to reconsider their decision during each successive birth interval until the desired number of children are born (or adopted). Given the opportunity for reconsideration, a proportion of couples will change their desires at each birth interval (see Westoff and Ryder, 1977). Furthermore, any couple that achieves their desired limit may subsequently raise their desires during the open interval between the achievement of previous desires and the onset of menopause. Namboodiri's (1972) fully dynamic model of marital fertility and Heckman and Willis's (1976) dynamic model of contraceptive strategies are based on similar considerations.

Of course, if a couple's fertility desires are nearly stable over time, the static decision-making framework's inattention to sequence may not produce an unreasonably inaccurate characterization, given the analytic power of the static assumption. However, panel studies of desired and expected fertility reveal considerable instability (Freedman et al., 1965; Bumpass and Westoff, 1970; Westoff and Ryder, 1977). A sequential decision-making framework is a logical alternative to the empirically unsupported, static one.

In addition to changes due to sequential decision making, the instability of desired fertility reflects temporal changes in the economic and noneconomic constraints faced by the couple and stochastic disturbances that affect fertility decisions throughout the childbearing years. Although the static approach permits changes in constraints, it assumes that equilibrium family size reflects the couple's "perfect foresight concerning all relevant demographic and economic variables over the course of their marriage" (Willis, 1973, p. S17). Furthermore, a single-period decision-making framework assumes away stochastic disturbances that affect fertility after the family size decision has been made (Willis, 1973, p. S17). Specifically, static models specify random disturbances that may cause couples to choose family sizes that deviate from what would be expected given their economic constraints, but nothing affects desired family size once an initial decision has been made.

Subsequent disturbances, including unplanned births and subfecundity, are important because they directly alter fertility expectations and indirectly affect desires through rationalization. Although unwanted fertility has declined since 1960, it remains a salient component of marital fertility in the United States (Ryder and Westoff, 1972; Westoff and Ryder, 1977). Furthermore, physiological impairments and unrealistic desires prevent some couples from bearing the desired number of children. When data on unplanned births or subfecundity are available, their impact on the couple's decision making can be evaluated from the dynamic perspective. Unfortunately, the data analyzed in this paper, a Public-Use Sample from the U.S. census of 1970, does not include such data, so this advantage of the dynamic perspective cannot be exploited here.

ANALYTICAL DIFFERENCES BETWEEN DYNAMIC AND STATIC PERSPECTIVES

Adoption of the dynamic perspective changes the analytical meaning of the interval between births. From the static perspective, each birth interval constitutes a pause in a process that is predetermined by the equilibirum family size that maximizes the one-period utility function. The dynamic perspective, on the other hand, treats each birth interval as a decision point, implying the possibility of different constraints on the fertility decision at each interval. ". . . (F)ertility decisions taken at different points in time and the success or failure in carrying out those decisions" are the foci of the dynamic perspective (Namboodiri, 1972, p. 298). In this paper, as in most, data limitations preclude the observation of decisions per se. Conclusions regarding couples' decision making are inferred from observed outcomes under the assumption that the outcomes adequately represent the decisions.

A dynamic perspective is not synonymous with a short-range perspective. Asserting that parents reevaluate their plans after the birth of each child does not imply that the decision to bear an additional child is approached frivolously. Nor does it follow that short-term fluctuations in socioeconomic variables are more important than variation in permanent conditions in determining actual fertility. Indeed, since childrearing requires outlays over so many years, any model of fertility decision making, static or dynamic, should stress the importance of differentials in permanent earning capacity over transitory fluctuations in earnings. Addressing this point by comparing the effects of permanent and annual earnings on the period fertility of whites in the United States, Hout (1976), using a specification somewhat different from the one proposed below, reports stronger effects at each parity for permanent earnings.

The issue of the effects of permanent versus transitory changes in socioeconomic variables arises in another way as well. The static perspective defines a relevant dependent variable for analysis: children ever born. The adaptation to the dynamic perspective seems straightforward enough: analyze the parity progressions of women who are past childbearing age. Several studies have done so (Bean and Wood, 1974; Simon, 1975).

However, when one's interest is not confined to sequential fertility decisions but includes sequential labor force participation as well, the dearth of detailed work and childbearing histories directs attention to *period* fertility and labor force participation. This redirection poses a problem, because differentials in period fertility contain both timing and number elements (e.g., Kiser et al., 1968, pp. 255–264). The ambiguity is not as great when analyzing first and second births in contemporary United States. Effects on first and second births are predominantly effects on the timing of those births, because nearly all married women have had or expect to have at least one child, and nearly 90 percent of those in the cohorts analyzed below expect at least two (U.S. Bureau of the Census, 1976, Table 2). Interpreting effects on higher order births is more difficult because of the timing-number mix, e.g., does a positive income effect on third births mean poor couples postpone third births or forego them altogether?

What specific parity differences can be expected? Ryder's (1973b) characterization of family size norms in the United States suggests stronger effects of income on the period fertility of childless women and those with one child than on that of those who have two or more children: "These norms specify that all people are expected to marry and have two children as soon as, and provided that, their economic circumstances permit" (p. 61). This hypothesis is not negated by the lack of substantial socioeconomic effects on the birth expectations of couples with few children (Namboodiri, 1974), because the hypothesis refers to timing factors not tapped by expectations.

The effects of wife's employment and earning potential may or may not vary with parity. If nearly all women expect at least two children, as Blake (1968, p. 17), Ryder (1973b, p. 61), and others suggest, then the ratio of the opportunity costs of childbearing to the benefits of additional children should increase with parity. Employed women and women with high earning potential should be just as likely to have first and second births as other women, but they should be decreasingly likely to have higher order births. On the other hand, employed women and those with greater earning potential may delay first and second births in order to accumulate work experience (Rosenzweig, 1976). These conflicting perspectives reflect the confounding influences of long- and short-term effects and make it difficult to make firm predictions regarding parity differences in the effects of employment and potential earnings.

Since spacing preferences appear to be oriented more towards the desired duration of child care responsibilities than towards specific lengths for each interval (Bumpass and Westoff, 1970, pp. 32–37), age differences in the likelihood of a birth may be greater at lower than at higher parities.

Because childrearing and employment compete for the time of married women, this study includes a parity-specific analysis of married women's employment. Theory and previous research suggest parity differences in the relative effects of the causes of employment. Microeconomic theory predicts a negative effect of husband's income and a positive effect of wife's potential earnings on her labor force particpation (Mincer, 1962). Mincer hypothesizes that the substitution effect of husband's earnings increases with parity, because wives are less able to exchange labor market time and home time as child care responsibilities increase. Waite's (1976) analysis supports Mincer's hypothesis for the 1940–1960 period and suggests that changes in preferences (not self-selection—see p. 144 below) produce the parity differences in effects.

Another possible parity difference is in the effect of previous employment on current employment. Mott (1972) and Waite (1976) report that, as parity increases, women without previous experience become successively less likely to enter the labor market, at least while young children are present in the home.

RECURSIVE AND SIMULTANEOUS MODELS

Recursive Model

Two models are employed to evaluate the parity interactions. The first model consists of a pair of recursive equations. The period fertility of a white ($k = 1$) or black ($k = 2$) woman i who has previously borne j children (F_{ijk}) and her labor force participation during the same period (L_{ijk}) are race-and-parity-specific (linear) functions of her previous employment (E_{ijk}), duration of marriage (D_{ijk}), earning potential (W_{ijk}), and birth cohort (C_{ijkm}), where m = 1931–1935, 1936–1940, 1941–1945, 1946–1951, and husband's earning potential at age 40-$Y40_{ijk}$:

$$F = a_{jk}(E, D, W, C_m, Y40), \qquad (1)$$

and

$$L = b_{jk}(E, D, W, C_m, Y40) \qquad (2)$$

(redundant subscripts are omitted; variables are defined below). It should be noted that this model is strictly recursive, i.e., past fertility and labor force experience are assumed to be uncorrelated with the disturbances in equations (1) and (2). Under the usual assumptions (Johnston, 1972, pp. 121–123), ordinary least squares (OLS) is an acceptable, if inefficient, estimator of the parameters of equations involving dummy dependent variables, as equations (1) and (2) do. Customary significance tests for such equations are biased, however. In this paper, significance tests are computed using an adjustment derived by Ashenfelter (Bowen and Finegan, 1969, pp. 644–648).

Previous employment is surely related to the disturbances in the labor force equation. This questionable specification is used, because I suspected that the bias resulting from excluding previous employment from the model is greater than that resulting from including it.

In general, potential earnings should be considered endogenous to this type of model, since the wage a woman can command depends on her experience and current labor market activity. To avoid simultaneity bias, a set of exogenous instrumental variables is used to estimate the effects of potential earnings (see measurement details below).

The limitations inherent in census data force the exclusion of two potentially important variables: religion and childrearing costs. Religious differentials in fertility have long been noted (Kiser et al., 1968, pp. 229–234). Increased use of contraception by U.S. Catholics has resulted in reductions in religious differentials (Ryder and Westoff, 1972), however, indicating that the bias due to the exclusion of religion from the analysis of recent fertility is not as great now as it would have been for earlier periods. Because of the interaction between income and religion (Ryder and Westoff, 1971, pp. 77–80), the most likely bias is toward a positive income effect at higher parities.

A more serious exclusion is that of childrearing costs. Although childrearing expenditures are somewhat discretionary (Becker and Lewis, 1973; Willis, 1973), social class determines costs to a significant degree (Duesenberry, 1960; Easterlin, 1969, 1973; Namboodiri, 1972; Turchi, 1975a, 1975b). The income effects estimated in this paper are total effects in the sense that they include both positive direct effects and negative indirect effects via costs (of course, if income is the only source of variance in costs, the direct-indirect distinction is unimportant).

Husband's and wife's educations are excluded from the equations, but their impacts on fertility and labor force participation are captured in the model by two composite variables: husband's projected earnings and wife's potential earnings (measurement details are given below). Including the spouses' educations in composites constrains their effects (see Hauser and Goldberger, 1971), but including them in the model along with the composites introduces severe multicollinearity.

Although the sample size is quite large, it is not sufficient to examine cohort differences in the models' parameters.

These exclusions inject some bias into the estimation, but they do not invalidate the analysis. Parity differences in the effects of social and economic variables are inconsistent with the static perspective, whether the differences reflect the effects of observed or unobserved variables. The analysis is invalid only if biases produce spurious parity differences.

Simultaneous Effects Model

If childbearing and working are substitutes as assumed, the disturbances of equations (1) and (2) will be negatively correlated. In order to evaluate the extent to which that correlation results from simultaneous influence between the endogenous variables, a simultaneous effects model is also considered.

Without restrictions on some of the effects of exogenous variables, the simultaneous effects model would be underidentified. Identification is difficult, because demographic theory provides only weak justification for restricting the effects of any of the exogenous variables.

Although uncertain identifying restrictions may distort results, the potential for addressing three important issues makes the risk tolerable, particularly since estimates from a less questionable model [equations (1) and (2)] are available for comparison. First, since many women participate in the labor force both before and after bearing children, the effects of past fertility on current labor force participation and labor force experience on current fertility are likely to be weaker than the simultaneous effects. If we ignore simultaneous effects, we will underestimate the magnitudes of the constraints and trade-offs involved. Second, Mincer (1962) hypothesizes that the wife's earning potential accounts for the association between fertility and labor force participation. That hypothesis can be rejected (as a short-run prediction, at least) if the reciprocal effects of current fertility and labor force participation are significant, net of potential earnings. Third, research on the simultaneous effects of birth expectations and labor force plans has revealed asymmetry; the effect of labor force plans on the expected fertility of young women is stronger than the reciprocal effect (Waite and Stolzenberg, 1976). It is important to compare the results for expectations with the effects of actual births and employment.

Restriction of at least one parameter in each equation is a necessary condition for identification (Johnston, 1972, p. 358). To that end, previous employment is deleted from the fertility equations, and duration of marriage is deleted from the labor force equation. Excluding a variable from an equation does not imply that the effect of the excluded variable is inconsequential; it does constrain the effect of the excluded variable. Excluding previous employment from the fertility equation is a matter of some consequence in light of Rosenzweig's (1976, p. 340) hypothesis that the "effects of socio-economic characteristics on birth expectations or fertility may appear to differ significantly by parity and age . . . or work status . . . when differences in female work histories are neglected." The implications of the findings of the models presented here for Rosenzweig's hypothesis are discussed below.

The parameters of the simultaneous model are estimated by two-stage least squares (2SLS). The use of dichotomous endogenous variables limits the efficiency of 2SLS by violating the homoscedasticity assumption. The consistency of 2SLS does not depend on homoscedasticity (Johnston, 1972, p. 383), so 2SLS was chosen over more efficient estimators that cannot be used to estimate asymmetrical simultaneous effects (Goodman, 1973; Schmidt and Strauss, 1975). Because the endogenous variables are dichotomous, significance tests using standard errors computed in the usual way would be inconsistent. Lacking a more rigorous procedure, I applied the Ashenfelter formula, applied to the OLS estimates for the recursive model, to the 2SLS estimates for the simultaneous model.

Although these problems of identification and estimation hamper the data analysis, they do not invalidate it as long as they are borne in mind while interpreting the results, and as long as the results are compatible with the results from the recursive model.

Selectivity is an estimation problem of some consequence. By affecting births to lower parity women, the exogenous variables screen the entrants to the higher parity subsamples. Consequently, it is unlikely that the disturbances of the higher parity equations are independent of the exogenous variables. Heckman and Willis (1976) have recently proposed an estimator for this type of model. Unfortunately, it was not available at the time this analysis was performed.

DATA SOURCE

For applying the dynamic perspective to empirical data, the ideal data set would be complete fertility histories acquired over the childbearing years of a cohort.

No such data exist, of course, but the perspective can be useful in analyzing inferior data, e.g., census data. In this paper, the data analyzed are short-term fertility histories obtained from the household enumerations in the 1970 U.S. census Public-Use Samples (PUS) (U.S. Bureau of Census, 1972). To make the sampling variability of the black and white samples more comparable, the black households were selected from the 1/100 PUS and the white households from the 1/1000 PUS.

MEASUREMENT

Fertility

Period fertility is indicated by the number of own *children under two years old* living with married-once, spouse-present women who were born between 1931 and 1951 and married before 1968. The marital status restrictions ensure a uniform family context for the measurement of past and current fertility. Although annual fertility (children under one year of age) may hypothetically be a more interesting variable, small timing failures (e.g., a failure to conceive as soon after discontinuing contraception as planned), the spacing effect of lactation, and the temporary separation of spouses introduce too many biases into annual data. The two-year interval was chosen because it was long enough for these biases to iron out but not long enough to obscure the simultaneous relationship between fertility and labor force participation.

Inferring fertility histories from census enumerations of own children would be infallible if all children born between April 1968 and April 1970 were enumerated, lived with their mothers, and survived to April 1970. None of these conditions is perfectly satisfied, however (Siegel, 1974; Rindfuss, 1976; U.S. Bureau of the Census, 1975, Table 84). Annual data for women of all marital statuses presented by Rindfuss (1976, Tables 3–4) indicate that errors from all sources cause the own children method to underestimate the fertility rates of white women 20 to 39 years old by an average of 1.5 percent in 1969 and 6.1 percent in 1968. The underestimates for black women 20 to 39 years old are more significant; they average 17.0 percent in 1969 and 16.6 percent in 1968. Although sufficient data are lacking, it seems reasonable to assume that data for women in intact marriages are more accurate than the data reported by Rindfuss. The racial difference in the accuracy of the own children data precludes quantitative comparisons between blacks and whites, but racial differences in the patterns of effects are tentatively investigated below.

Labor Force Participation

In this study of marital fertility between 1968 and 1970, the labor force variable cannot be restricted to 1970 experience; it must include experience throughout the period. Census data on work histories are sketchy, but the questions asked make it possible to separate the sample into two groups: those who were employed full or part time at any time between January 1968 and April 1970 (scored 1) and those who were not (scored 0).

Previous Parity and Employment

Measuring previous parity is straightforward given census data. It is defined simply as children ever born minus children under two years of age.

Ideally, previous employment should be measured in months or years of actual experience, but census data do not include such detailed information. The best indicator available is a dummy variable distinguishing those employed in 1965 (scored 1) from those unemployed or out of the labor force (scored 0) at that time.

Potential Earnings

A simple approach to measuring potential earnings would be to assume that the earnings potential of nonworking wives (who have no observed earnings) is equal to that of working wives with similar social and demographic characteristics. That assumption would justify an instrumental variable regression, using coefficients estimated for working wives to estimate the potential earnings of nonworking wives. Gronau (1973) and Heckman (1974) argue against this assumption. They reason that nonworking wives must value their time at home more highly than working wives, or they too would be working. This suggests that the instrumental variable approach would underestimate the potential earnings of nonworking wives. On the other hand, Turchi (1975a, pp. 98–101) and Waite (1976) argue that assuming equality between workers and nonworkers overestimates the nonworkers' potential, because their rates of return for education and other human capital investments are lowered by their lack of experience and on-the-job training.

Lacking an estimator that simultaneously adjusts for both the positive and negative biases of the instrumental variable method, I have made the tenuous assumption that the *net* bias is close enough to zero to be considered negligible. Fligstein and Wolf's (1976) finding of no net bias with respect to occupational status offers indirect support for this assumption.

Wife's potential earnings were constructed by regressing the natural log of working women's reported earnings for 1969 on weeks worked (scored to the midpoints of reported intervals), years of school completed, region (South = 1), metropolitan residence (SMSA = 1), birth cohort (dummy variables for five-year intervals), and work limitation (disabled or handicapped = 1), by race. Hours of work per week is excluded from the prediction equation for two reasons related to the changeability of women's work schedules. First, earnings and weeks of work are reported for 1969; hours are reported for the week prior to the census. Given the variability of many women's work schedules, hours as reported are a poor indicator of average hours per week worked in 1969. Second, many women who worked in 1969 reported no hours worked during the week prior to the census. The exclusion of hours worked introduces an unknown amount of bias into the calculations.

Expected log earnings were computed for all women (working or not) using the coefficients in the first two columns of Table 1. To avoid simultaneity bias in estimating the effect of potential earnings on fertility and labor force participation, only the exogenous variables were permitted to take their observed values. All women were assigned 52 as the value for weeks worked. Thus, potential earnings represent the annual earnings expected, assuming a full year's work (variance in hours worked is not controlled).

Husband's Earnings

Husband's earnings are projected to age 40 to adjust for life cycle effects, transitory fluctuations in earnings, and differences between permanent and annual earnings. The projection was accomplished in four steps.

1. The occupational SEI (Duncan, 1961; Featherman et al., 1975) of all employed males 18 years old and over in the 1/1000 PUS was regressed on SMSA residence, race, education, industry, disability, and age group.
2. The projected SEI for husbands was computed using the regression coefficients in column 3 of Table 1, assigning all husbands an age of 40 years.
3. The natural log of the 1969 earnings of men aged 18 and over was regressed on current SEI, SMSA residence, race, education, industry, disability, and age group.
4. Husband's earnings at age 40 were computed using the coefficients in column 4 of Table 1, replacing current SEI with projected SEI and assigning all husbands an age of 40 years.

Other Variables

Duration of marriage, measured in single-year intervals, was computed from wife's current age and her age at first marriage, as reported to the census. Likewise, wife's birth cohort, measured in five-year intervals, was obtained from her reported

Table 1.—Coefficients Used to Compute Wife's Potential Earnings and Husband's Projected Earnings Derived from Regressions on Samples of Married-Once, Spouse-Present Women Married Before 1968 and Employed Males Aged 18 Years and Older: United States, 1970

Independent Variables	Computed Variable: Potential Earnings of White Wives (N=4,751)[a]	Potential Earnings of Black Wives (N=7,288)[b]	Projected SEI of Husbands (N=44,350)[a]	Projected Earnings of Husbands (N=44,350)[a]
Weeks worked, 1969	.050	.026	---	---
Education (years)	.059	.128	3.350	.047
Region (South=1)	.066	-.217	---	---
Residence (SMSA=1)	.106	.197	2.258	.153
Disability (yes=1)	.010	-.299	-.872	-.185
Occupation (SEI)	---	---	---	.007
Birth cohort				
1946-1951	---	---	---	---
1941-1945	.197	.233	---	---
1936-1940	.210	.282	---	---
1931-1935	.210	.334	---	---
Industry				
Agriculture, forestry, fishing	---	---	-21.318	-.421
Mining	---	---	-8.300	.183
Construction	---	---	-8.769	.146
Durable manufacture	---	---	-7.397	.165
Other manufacture	---	---	-6.152	.070
Transportation	---	---	-9.641	.122
Communication	---	---	6.175	.077
Utilities	---	---	-6.660	.108
Retail trade	---	---	-5.577	.142
Wholesale trade	---	---	-.914	.071
Business, insurance and real estate	---	---	7.931	-.016
Business and repair services	---	---	-7.315	-.044
Personal services	---	---	-11.503	-.242
Entertainment services	---	---	-3.446	-.286
Professional services	---	---	3.915	-.194
Public administration	---	---	---	---
Race (black=1)	---	---	-8.046	-.258
Constant	.190	.501	7.022[a]	3.493[c]
R^2	.538	.330	.434	.373

a- The source for this sample is the 1/1000 Public Use Sample, 1970.
b- The source for this sample is the 1/100 Public Use Sample, 1970.
c- This term is the sum of the estimated regression constant and the coefficient for the group aged 40-44.

age. Race and current marital status are reported directly in the census data. Only black and white married-once, spouse-present women born between 1931 and 1951 and married before 1968 are included in the analysis.

The means specific for race and parity are presented in Appendix Table 1.

FINDINGS

White Wives

Recursive model. The results for white wives are reported first. The OLS coefficients for the reduced form equations are in Table 2. Previous employment was expected to affect current fertility negatively

Table 2.—Reduced Form Coefficients for Two-Equation Model of Period Fertility and Employment of Married-Once, Spouse-Present White Women Married Before 1968, By Previous Parity: United States, 1968–1970

Independent Variables	Previous Parity									
	0	1	2	3	4+	0	1	2	3	4+
	Fertility, 1968-1970					*Employment, 1968-1970*				
Employment, 1965	.035	-.006	-.005	-.010	-.022	.182[a]	.189[a]	.379[a]	.453[a]	.467[a]
Potential earnings	-.116	-.013	-.051	-.136[a]	-.127[a]	.542[a]	.280[a]	.222[a]	.213[a]	.244[a]
Duration of marriage	-.156[a]	-.241[a]	-.247[a]	-.174[a]	-.201[a]	-.076[a]	.166[a]	.242[a]	.189[a]	.296[a]
Husband's earnings, age 40	.130[a]	.129[a]	-.023	-.020	-.020	-.108[a]	-.207[a]	-.093[a]	-.140[a]	-.042
Cohort										
1946-1951	.052	.018	.010	.030	-.019[c]	.130	.112	.229	.101	.168[c]
1941-1945	.021	.052	-.005	.028	-.019[c]	-.046	-.046	-.022	.041	.168[c]
1936-1940	-.055	-.077	-.014	-.017	-.005	-.148	-.045	-.028	.011	.021
1931-1935	-.174	-.124	.019	-.003	.009	-.108	-.143	-.099	-.048	-.061
Constant	.318[a]	.064	.748[a]	.962[a]	.999[a]	-.794[a]	.243[a]	-.278[a]	-.033	-.802[a]
R^2	.081	.099	.083	.065	.046	.117	.052	.148	.168	.151
Error variance	.219	.219	.138	.108	.111	.140	.233	.213	.208	.209
Fertility-employment error covariance	-.027	-.035	-.021	-.009	-.012					
Number of women	1,991	2,760	3,535	2,560	2,417					

a- $|t| \geq 1.96$.
b- Coefficients multiplied by 10.
c- Coefficients constrained to be equal.

Source: 1/1000 Public-Use Sample, 1970.

(no hypothesis regarding parity differences was advanced). This expectation is not borne out in the data; none of the coefficients is significantly different from zero. This finding casts doubt on Rosenzweig's (1976) hypothesis that findings of parity differences in socioeconomic effects for U.S. women [as reported by Namboodiri (1974) and Rosenzweig and Seiver (1975)] are spurious results due to the exclusion of previous employment. A full discussion of the relationship between fertility and employment will be undertaken after the results for the simultaneous model are presented.

Negative effects on fertility were expected for wife's potential earnings. The coefficients are negative at each parity, but they are significantly less than zero at the highest two parities only. Two perspectives on parity differences in the effects of potential earnings were discussed above. If women with high earning potential conform to the two-child norm, the effect of potential earnings would be weaker at low parities. On the other hand, a desire to accumulate experience may cause women with high earning potential to delay births, a condition under which no parity differences would appear. The coefficients in Table 2 suggest that both processes operate. While the only significant coefficients are those for women with three or more births prior to April 1968, the coefficient for previously childless women is nearly as large as the significant coefficients and is nearly significant itself ($p = .11$).

The coefficients for duration of marriage are uniformly negative, as expected. Although no parity differences were hypothesized, the coefficients for previous parities 1 and 2 are significantly larger than the others (two-tailed t-tests, $p < .05$). Duration of marriage is an indicator of the passage of time. Following that interpretation, we find that the peaked pattern of the coefficients indicates that second and third births are more sensitive to the passage of time than are other births. Further research is needed to illuminate this interesting pattern.

The impact of husband's projected earnings was expected to decline with increases in parity. The results form impressive support for the hypothesis that the timing of first and second births is highly sensitive to economic opportunity, while the timing of subsequent births is not subject to economic influence of the same magnitude. The step-like pattern of the coefficients may reflect preferences of higher income couples who already have two children to use their money for goods and services unrelated to childrearing, as argued by Blake (1968) and Namboodiri (1972), or to invest their earnings in the children they already have as Becker (1960; Becker and Lewis, 1973) argues; census data are inadequate for a definitive test (Cochrane, 1974). Nevertheless, the results do support Ryder's (1973b, p. 61) characterization of family formation norms, particularly his stress on timing: ". . . as soon as, and provided that, their economic circumstances permit."

Age-cohort differences are significant (F-ratios for the set of dummy variables, $p < .05$) at previous parities 0 and 1 only. The pattern of smaller differences was expected, because the length of time from first to last birth does not vary as much as does completed fertility (Bumpass and Westoff, 1971, pp. 32–37).

The proportion of the variance explained by the exogenous variables declines with each successive birth after the second, indicating the disturbing influence of involuntary childlessness and the increase in unwanted births as a proportion of all births as parity increases (Westoff, 1976).

The results for the white wives' employment equation are much as expected. The coefficients for previous employment are all positive and significant, and they increase monotonically with parity. The coefficients for wife's potential earnings are consistently positive; the coefficients for husband's projected earnings are negative (and significant at each but the highest parity).

The parity differences in the effects of husband's and wife's earnings are not as

Mincer (1962) hypothesized, however. The effect of wife's earnings was expected to approach zero as parity increased, and the effect of the husband's earnings was expected to be strongest at the highest parities. The divergence from expectation is due in part to the inclusion of previous employment in the equation; the dependence of previous employment on husband's and wife's earnings mediates their effects on current employment (Hout, 1976).

The significant positive effects of duration of marriage on the employment of wives with one or more children at home is an indirect effect: the product of the negative effect of duration of marriage on fertility and the negative effect of fertility on employment. The negative effect among previously childless women is puzzling. The same result appears as a positive coefficient for current fertility at parity 0 in Table 3 below. Perhaps recently married women delay first births to accumulate work experience (Rosenzweig, 1976) or simply to earn money to defray the anticipated expenses.

The pattern of the age-cohort coeffi-

Table 3.—Metric Two-Stage Least Squares Coefficients for Two-Equation Model of Period Fertility and Employment of Married-Once, Spouse-Present White Women Married Before 1968, by Previous Parity: United States, 1968–1970

Independent Variables	Previous Parity 0	1	2	3	4+
			Fertility, 1968-1970		
Employment, 1968-1970	.193	-.033	-.013	-.023	-.048
Potential earnings	-.221[a]	-.004	-.048	-.131[a]	-.116[a]
Duration of marriage[b]	-.141[a]	-.236[a]	-.244[a]	-.170[a]	-.187[a]
Husband's earnings, age 40	.151[a]	.122[a]	-.024	-.024	-.022
Cohort					
1946-1951	.027	.022	.013	.032	-.006[c]
1941-1945	.030	.050	-.005	.029	-.006[c]
1936-1940	-.027	-.079	-.014	-.017	-.001
1931-1935	-.153	-.129	.018	.004	.005
Constant	.472	.072	.744[a]	.965[a]	.964[a]
Error variance	.235	.217	.138	.107	.111
			Employment, 1968-1970		
Fertility, 1968-1970	.721[a]	-.144[a]	-.477[a]	-.584[a]	-.930[a]
Potential earnings	.375[a]	.042	-.046	.066	.031
Employment, 1965	.131[a]	.174[a]	.381[a]	.447[a]	.446[a]
Husband's earnings, age 40	-.167[a]	-.163[a]	-.080[a]	-.158[a]	-.065[a]
Constant	-.202	1.117[a]	1.025[a]	.924[a]	.665[a]
Error variance	.296	.230	.228	.235	.286
Error covariance	-.215	.004	.048	.059	.102
Number of women	1,991	2,760	3,535	2,560	2,417

a- $|t| \geq 1.96$.
b- Coefficients multiplied by 10.
c- Coefficients constrained to be equal.

Source: 1/1000 Public Use Sample, 1970.

cients is not surprising. The truncated age range of this sample corresponds to the downward slope of the U-shaped age-employment relationship common to cross-sectional analyses (Oppenheimer, 1972).

Simultaneous model. The main issue in analyzing the simultaneous model is the nature of the fertility-employment relationship. The relationship is decidely asymmetrical. Fertility affects employment strongly at each parity, but employment does not affect fertility at any parity. This contrasts sharply with the asymmetry between the expected fertility and labor force participation of young women. Waite and Stolzenberg (1976) report an effect of labor force participation plans on expected fertility that is much stronger than the reciprocal effect. This study analyzes period fertility; Waite and Stolzenberg analyze the long-term plans of young women. The differences in perspectives and findings suggest that, in the short run, the discomforts of pregnancy and the demands of newborns decrease labor force participation and thereby account for the negative association between fertility and employment while, in the long run, fertility is curtailed to accommodate career commitments. Mincer's hypothesis that the negative association is attributable to the wife's wages is not supported.

The 2SLS estimates are questionable at parities 0 and 4+. The coefficients for current fertility in the employment equations at those two parities are particularly puzzling. The sensitivity of these estimates is examined below.

The coefficients for the exogenous variables in the fertility equation are not substantially different from the corresponding coefficients in the recursive model. The impact of potential earnings on the fertility of previously childless women appears stronger in the simultaneous model than in the recursive model. The effects of duration of marriage, husband's earnings, and age-cohort on fertility are nearly identical in the two models. In the employment equation, estimates of the effects of wife's and husband's earnings differ somewhat, further obscuring the hypothesized parity differences. The pattern of parity differences in the coefficients for previous employment is as expected. Age-cohort differences in employment were not significant at any parity, so the cohort dummies were dropped from the employment equation.

Income Adequacy

Although the pattern of parity differences in the effect of husband's earnings on fertility has been interpreted as evidence that fertility norms accelerate the timing of the first and second births of those most able to afford children, the need to share economic resources among family members suggests an alternative explanation. Viewed on a per capita basis, an increase in husband's earnings goes farther in a small family than in a larger one. In other words, the adequacy of any given income is inversely proportional to family size. Thus, the parity differences in the effect of husband's earnings on fertility may be the result of an inappropriate metric being applied. The problem may be further compounded by differences in family composition with parity groups.

To adjust the metric of husband's earnings, I used poverty cut-offs by family size and composition (U.S. Bureau of the Census, 1972, p. 122). To avoid confounding the correction factor with fertility, poverty cut-offs were computed using number of children in the household minus children under age two. The index of income adequacy was obtained by dividing husband's total personal income in 1969 by the computed poverty cut-off.

Income adequacy is not a monotonic transformation of personal income within parity groups, because it adjusts for differences in family composition. For example, compare two households: the first composed of a husband who earned \$12,000 in 1969, a wife, a six-year-old child, and the wife's mother; the second composed of a husband who earned \$11,000, a wife, a six-year-old child, and an infant. Both households would be included in the previous parity 1 subsample. A "regression" of children under two on earnings using these two cases would produce a negative slope. Computing income adequacy adjusts for the presence of the wife's mother in the first household and reverses the order of the two households on the income variable: the first household's income adequacy is 3.12; the second's income adequacy is 3.71. A "regression" of children under two on income adequacy using these two cases would produce a positive slope.

The results obtained by substituting husband's income adequacy for his projected earnings are in Table 4. Although small positive coefficients appear at parities 2 and 3, the overall pattern of parity differences is unchanged: the income effects are substantially stronger among couples who had fewer than two children prior to April 1968. Accounting for differential income adequacy does not refute the two-child norm explanation of parity differences advanced above. The coefficients for the other variables in the model are changed little by the change of income indicator.

Instability of parameter estimates. The estimates for the effect of fertility on employment and of the covariance between the disturbances of the two equations are

Table 4.—Metric Two-Stage Least Squares Coefficients for Revised Model of Period Fertility and Employment of Married-Once, Spouse-Present White Women Married Before 1968, by Previous Parity: United States, 1968–1970

Independent Variables	Previous Parity 0	1	2	3	4+
	Fertility, 1968-1970				
Employment, 1968-1970	.179	-.013	.003	-.010	-.035
Potential earnings[b]	-.144[a]	.041	-.124[a]	-.201[a]	-.182[a]
Duration of marriage[b]	-.150[a]	-.243[a]	-.245[a]	-.169[a]	-.189[a]
Income adequacy[b]	.281[a]	.229[a]	.091[a]	.067[a]	.056[a]
Cohort					
1946-1951	.046	.033[a]	.005[a]	.028[a]	-.007[c]
1941-1945	.027	.047	-.001	.032	-.007[c]
1936-1940	-.050	-.087	-.014	-.016	-.003
1931-1935	-.179	-.142	.016	-.006	.004
Constant	.784[a]	.365	.879[a]	1.086[a]	1.054[a]
Error variance	.232	.217	.138	.107	.111
	Employment, 1968-1970				
Fertility, 1968-1970	.590[a]	-.196[a]	-.520[a]	-.641[a]	-.949[a]
Potential earnings	.362[a]	.012	-.021	.032	.093
Employment, 1965	.144[a]	.171[a]	.373[a]	.436[a]	.433[a]
Income adequacy[b]	-.439[a]	-.275[a]	-.217[a]	-.254[a]	-.220[a]
Constant	-.715[a]	.619[a]	.664[a]	.465[a]	.239[a]
Error variance	.246	.229	.231	.240	.287
Error covariance	-.181	.012	.051	.063	.102

a- $|t| \geq 1.96$.
b- Coefficients multiplied by 10.
c- Coefficients constrained to be equal.

Source: 1/1000 Public Use Sample, 1970.

very large and of opposite sign for white women in the highest and lowest parity groups. These unrealistic estimates signal an instability in the model. To assess the consequences of that instability for substantive conclusions, a method of constraining the error covariance to zero is necessary. Such a constraint is impossible using a single-equation estimator like 2SLS. Jöreskog (1973) has developed a multiequation estimator that is equivalent to full-information maximum likelihood (FIML) when the estimates are unconstrained. The advantage of Jöreskog's ML estimator is the ease with which constraints can be incorporated in the model (Jöreskog and Van Thillo, 1972). Constrained and unconstrained parameter estimates for women in the highest and lowest parity groups are presented in Table 5. Note that FIML estimates are more sensitive to specification errors than 2SLS (Johnston, 1972) and were not used throughout this analysis for that reason.

The results in Table 5 show that, whatever the source of the instability, it is the estimates of the reciprocal effects of employment and fertility that are most affected. Estimates of the substitution and income effects in both equations are somewhat sensitive to estimated error covariance. Though more reasonable than the coefficients for highest and longest parity in Table 4, the estimates in Table 5 cannot be taken as final. An error covariance of exactly zero is as unrealistic as a very large one. Thus, the true parameters must lie somewhere between the constrained and

Table 5.—Full-Information Maximum Likelihood Coefficients for Constrained and Unconstrained Models of Period Fertility and Employment for Married-Once, Spouse-Present White Women Married Before 1968, by Previous Parity: United States, 1968–1970

Independent Variables	Previous Parity			
	0		4+	
	Unconstrained	Constrained	Unconstrained	Constrained
	Fertility, 1968-1970			
Employment, 1968-1970	.271	.734[a]	-.041	.104[a]
Potential earnings	-.054	.320[a]	-.200[a]	-.230[a]
Duration of marriage[b]	-.110[a]	-.229[a]	-.211[a]	-.232[a]
Income adequacy[b]	.301[a]	.021	.048	.103[a]
Cohort				
1946-1951	.090	.119	-.060[c]	-.030[c]
1941-1945	-.015	.010	-.060[c]	-.030[c]
1936-1940	-.100	-.150	-.007	-.004
1931-1935	-.133	-.247	.022	.011
Constant	.343	-1.338[a]	1.195[a]	1.255[a]
Error variance	.243	.262	.111	.116
	Employment, 1968-1970			
Fertility, 1968-1970	.691[a]	-.414[a]	-1.100[a]	-.315[a]
Potential earnings	.382[a]	.327[a]	.072[a]	.181[a]
Employment, 1965	.141[a]	.149[a]	.427[a]	.458[a]
Income adequacy[b]	.453[a]	-.414[a]	-.212[a]	-.254[a]
Constant	-.823[a]	-.202[a]	.333[a]	-.166[a]
Error variance	.278	.198	.318	.213
Error covariance	-.219	0.[c]	.120	0.[c]

a- $|t| \geq 1.96$.
b- Coefficients multiplied by 10.
c- Constrained coefficients.

Source: 1/1000 Public Use Sample, 1970.

Table 6.—Reduced Form Coefficients for Two-Equation Model of Period Fertility and Employment of Married-Once, Spouse-Present Black Women Married Before 1968, by Previous Parity: United States, 1968–1970

Independent Variables	Previous Parity 0	1	2	3	4+	0	1	2	3	4+
	Fertility, 1968–1970					*Employment, 1968–1970*				
Employment, 1965	.052[a]	-.009	-.010	-.055[a]	-.025	.302[a]	.280[a]	.344[a]	.351[a]	.411[a]
Potential earnings	.038	.005	-.061[a]	.015	-.038[a]	.123[a]	.104[a]	.079[a]	.016	.044[a]
Duration of marriage[b]	-.100[a]	-.158[a]	-.144[a]	-.125[a]	-.077[a]	-.071[a]	.008	.008	.013	-.001
Husband's earnings, age 40	.033	.042	.003	-.092[a]	-.063[a]	.026	-.056[a]	-.036	.050	.039
Cohort										
1946–1951	.158	.095	.122	.196	.090[c]	.070	.050	.050	.031	.011[c]
1941–1945	.029	.018	-.011	.020	.090[c]	-.007	-.004	.024	-.002	.011[c]
1936–1940	-.073	-.099	-.040	-.043	-.027	-.054	-.015	-.020	.006	.005
1931–1935	-.161	-.121	-.057	-.061	-.030	-.019	-.080	-.063	-.018	-.012
Constant	.193	.329[a]	.690[a]	.876[a]	.953[a]	.226	.542[a]	.483[a]	.339[a]	.318[a]
R^2	.114	.101	.097	.106	.060	.174	.107	.137	.135	.154
Error variance	.169	.189	.157	.146	.153	.136	.166	.177	.180	.195
Fertility-employment error covariance	-.007	-.011	-.019	-.014	-.017					
Number of women	1,401	2,324	2,372	1,786	3,635					

a- $|t| \geq 1.96$.
b- Coefficients multiplied by 10.
c- Coefficients constrained to be equal.

Source: 1/100 Public Use Sample, 1970.

unconstrained estimates. The constraint appears to be more acceptable at parity 4+ than at parity 0. It is reassuring that the sign of only one parameter estimate is significantly different in the constrained and unconstrained solutions. In conclusion, while inferences regarding the exact magnitudes of the effects for women in the extreme parity groups must remain somewhat tentative, the directions of the effects are stable.

Black and White Parameter Estimates Compared

Only a limited comparison of black and white marital fertility is possible using the

Table 7.—Metric Two-Stage Least Squares Coefficients for Two-Equation Model of Period Fertility and Employment of Married-Once, Spouse-Present Black Women Married Before 1968, by Previous Parity: United States, 1968–1970

Independent Variables	Previous Parity 0	1	2	3	4+
			Fertility, 1968-1970		
Employment, 1968-1970	.171[a]	-.032	-.030	-.157[a]	-.068[a]
Potential earnings	.017	.008	-.059[a]	.017	-.052[a]
Duration of marriage[b]	-.087[a]	-.158[a]	-.144[a]	-.122[a]	-.093[a]
Husband's earnings, age 40	.029	.040	.002	-.084[a]	-.054[a]
Cohort					
1946-1951	.146	.097	.123	.200	.082[c]
1941-1945	.030	.018	-.010	.020	.082[c]
1936-1940	-.064	-.100	-.040	-.043	-.027
1931-1935	-.158	-.123	-.059	-.064	-.025
Constant	.008	.249	.582[a]	.729[a]	.767[a]
Error variance	.175	.189	.156	.146	.153
			Employment, 1968-1970		
Fertility, 1968-1970	.071	-.051	-.052	-.106	.179
Potential earnings	.096[a]	.104[a]	.076[a]	.046[a]	.047[a]
Employment, 1965	.265[a]	.279[a]	.343[a]	.345[a]	.415[a]
Husband's earnings, age 40	.002	-.054	-.036	.041	.052
Cohort					
1946-1951	-.042	.054	.056	.052	-.007[c]
1941-1945	-.028	-.003	.024	.000	-.007[c]
1936-1940	-.002	-.020	-.022	.001	.010
1931-1935	.095	-.087	-.066	-.025	-.005
Constant	.131	.505[a]	.463[a]	.380[a]	.122
Error variance	.231	.165	.175	.179	.207
Error covariance	-.151	.004	-.005	.029	-.031
Number of women	1,401	2,324	2,372	1,786	3,635

a- $|t| \geq 1.96$.

b- Coefficients multiplied by 10.

c- Coefficients constrained to be equal.

Source: 1/100 Public Use Sample, 1970.

Table 8.—Metric Two-Stage Least Squares Coefficients for Revised Model of Period Fertility and Employment of Married-Once, Spouse-Present Black Women Married Before 1968, by Previous Parity: United States, 1968–1970

Independent Variables	Previous Parity 0	1	2	3	4+
			Fertility, 1968-1970		
Employment, 1968-1970	.160[a]	-.032	-.028	-.152[a]	-.063
Potential earnings[b]	.001	.019	-.037	.011	-.055[a]
Duration of marriage[b]	-.081[a]	-.159[a]	-.141[a]	-.120[a]	-.089[a]
Income adequacy[b]	.276[a]	.041	-.231[a]	-.321[a]	-.298[a]
Cohort					
1946-1951	.148	.100	.125	.195	.081[c]
1941-1945	.035	.017	-.011	.017	.081[c]
1936-1940	-.065	-.102	-.041	-.042	-.028
1931-1935	-.166	-.126	-.058	-.058	-.023
Constant	.122	.366[a]	.549[a]	.454[a]	.589[a]
Error variance	.174	.189	.156	.146	.153
			Employment, 1968-1970		
Fertility, 1968-1970	.077	-.087	-.072	-.155	.200
Potential earnings	.118[a]	.104[a]	.075[a]	.063[a]	.066[a]
Employment, 1965	.265[a]	.280[a]	.343[a]	.342[a]	.415[a]
Income adequacy[b]	-.232[a]	-.197[a]	-.145[a]	-.354[a]	.055
Cohort					
1946-1951	-.050	.055	.057	.069	-.007[c]
1941-1945	-.033	-.002	.022	-.001	-.007[c]
1936-1940	.002	-.022	-.023	-.003	.010
1931-1935	.108	-.089	-.065	-.026	-.006
Constant	.102	.337[a]	.351[a]	.448[a]	.249
Error variance	.174	.165	.175	.179	.209
Error covariance	-.159	.011	-.003	.034	-.035

a- $|t| \geq 1.96$.
b- Coefficients multiplied by 10.
c- Coefficients constrained to be equal.

Source: 1/100 Public Use Sample, 1970

PUS data because, as stated above, the own children method used to measure period fertility is less reliable for blacks. Nevertheless, differences in the pattern of parameter estimates by parity can be meaningfully compared. Four appreciable racial differences are evident in Tables 6 7, and 8, and they are cataloged here.

1. Among blacks the effect of fertility on employment is not significant at any parity; among whites it is the reciprocal effect of employment on fertility that is insignificant.

2. The racial difference in the effect of potential earnings on fertility and employment is similar to the difference in reciprocal effects: the fertility of black women and the employment of white women are affected by potential earnings at only one parity, while the effect of potential earnings on the other endogenous variable is significant at each parity.

3. The pattern of successively less positive income adequacy effects holds for both blacks and whites, but the income effect itself is stronger among whites. Whether the difference stems from attenuation of the income effect among blacks due to their lesser contraceptive efficacy via the relationship of efficacy to socioeconomic status (Ryder and Westoff, 1971), socioeconomic or racial differences in the cost of children (Turchi, 1975a, pp. 117–162), or a direct minority group effect (Bean and Wood, 1974) cannot be ascertained in this study.

4. Both blacks and whites exhibit a curvilinear pattern of duration of marriage effects with the strongest effects at parities 1 and 2, but at each parity the effects are stronger for whites.

As Bean and Wood (1974) argue, racial differences in the parameters of the fertility equation may result from the insecurities of minority group status (Goldscheider and Uhlenberg, 1969; Goldscheider, 1971, pp. 289ff). Bean and Wood's argument applies most directly to their measure of male income relative to that of men in the same ethnic group. Lacking such an index, this study cannot explore their line of inference.

CONCLUSION

In place of the static decision-making framework, this paper offers a dynamic perspective similar to Namboodiri's (1972). The theoretical critique offered by the dynamic perspective is that couples do not select a desired family size early in marriage and orient their behavior to that goal throughout marriage. They may have some number in mind, but the behavior-orienting decision is the choice of one more child. When and if that child is born, their preferred family size is reevaluated and a decision reached about a subsequent child. So it goes until the marriage is dissolved, one of the marriage partners becomes sterile, or the couple decides to stop having children. This perspective is consistent with the sequential way in which children are acquired, the periodic disturbances (such as unplanned births) that affect fertility, and with the parity differences in the effects of social and economic factors presented in this paper.

Of crucial importance to the dynamic perspective are the parity differences in the effects of husband's earnings and income adequacy on recent fertility. In summary, data for the period April 1968 to April 1970 show that among white couples with *fewer than two children* prior to April 1968, couples in which the husband's projected earnings were high used their economic advantage to increase the size of their families; couples lacking economic resources were less likely to have additional children. But among white couples with *two or more children,* the economically advantaged used their money for goods and services unrelated to childbearing (though possibly related to childrearing); they were neither more nor less likely than couples with less income to experience additional births.

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