University of Wisconsin-Madison

Institute for Research on Poverty

Discussion Papers

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DP #593-79

An Analysis of the Labor Supply and Fertility of Married Women with Grouped Data from the 1970 U.S. Census

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January 1980

This paper is based upon my doctoral dissertation for which support was provided by the NIMH, Colby College, and the Center for Demography and Ecology and the Institute for Research on Poverty at the University of Wisconsin-Madison. I am greatly indebted to Glen G. Cain for assistance throughout the work on this project. Valuable comments were also provided by Randall S. Brown, David L. Crawford, Arthur S. Goldberger, V. Joseph Hotz, Stanley H. Masters and James A. Sweet.

ABSTRACT

The major objective of this study is to evaluate the results of previous studies of the labor supply and completed fertility of married women which used group (by SMSA) cross-sectional data from the 1940, 1950, and 1960 U.S. Censuses. Generalized Least Squares is used to estimate a two-equation model for five-year, age-of-wife cohorts (between ages 30 and 54) for both the non-Spanish white and black populations. Two data sets are used: SMSA averages from the 1970 Census published data, similar to those employed in previous studies, and averages calculated from an extract of the 1970 Public Use Sample. The latter permits more accurate measurement and the use of superior estimation methods, including corrections for sample selection bias.

In general, coefficient estimates with the 1970 published averages closely approximate those estimated with data from the three preceding censuses. Most importantly, the uncompensated wife's wage and husband's income elasticities are quite similar to previous estimates, except in the case of the white fertility equation. Continued evidence is also provided for a dominant discouraged-worker effect of transitory labor market conditions.

The findings with the 1970 Public Use Sample averages differ substantially from those obtained with the published averages. Most importantly, the effects of the wife's wage upon labor supply are insignificant and of varying sign. Correction for sample selection bias fails to

reverse this pattern. The estimated labor supply effects of husband's income are significantly negative, as expected, for the whites, but insignificant for the blacks, as are the nonlabor income coefficients for both races. Evidence of a dominant discouraged-worker effect persists but is weaker than in studies with published data. The estimates of the fertility equation for whites are as equally weak as those obtained with the published data. For the blacks, only the wife's wage and the proportion rural of the population have significant effects of the expected sign in contrast to significant effects of all the economic variables with the published data.

An Analysis of the Labor Supply and Fertility of Married Women with Grouped Data from the 1970 U.S. Census

1. INTRODUCTION

Several path-breaking studies of the labor supply of married women used grouped cross-sectional data from the U.S. Census (Mincer 1962, Cain 1966, Bowen and Finegan 1969, and Ashenfelter and Heckman 1974). This same data source has been used to estimate a model of the completed fertility of ever married women (Cain and Weininger 1973). The units of observation have been Standard Metropolitan Statistical Areas (henceforth SMSA) or cities from the 1940, 1950, and 1960 censuses. Recently, the 1970 SMSA data has been used by Fields (1976) to estimate the model used by Bowen and Finegan.

These data offered several advantages aside from their ready availability prior to the advent of several national surveys. Grouping is one method of obtaining consistent parameter estimates of income and wage effects in light of possible biases due to measurement error. Grouping by SMSA also afforded the opportunity to estimate the cyclical labor supply response to variation in employment and unemployment in the local labor market.

The estimates obtained in these studies have generally been statistically significant, stable over various censuses and consistent with theoretical predictions, particularly with respect to the effects of the husband's income and the wife's earning capacity. Another important result was the apparent procyclical labor supply response of married women to unemployment; that is, a dominant "discouraged worker effect."

Several deficiencies, however, marked these studies, most of which were readily acknowledged by the authors. First, the model specification did not reflect adequately more recent theoretical developments concerning endogenous wage rates and fertility. Second, there were several shortcomings in the estimation techniques, in particular the neglect of sample selection bias generated by the use of wage measures for current labor force participants. Third, the SMSA averages published by the Census Bureau (henceforth, published averages) often contained inappropriate observations, for example, average income for all males instead of for husbands or average nonlabor income for all families. In addition, measures of labor supply were limited to labor force participation rates and wage measures were based on earnings data that were contaminated with variation in hours of work.

The purpose of this study is to extend and evaluate this body of previous research using data from the 1970 U.S. Census. The theoretical shortcomings referred to above are not fully rectified. There are problems in reconciling theoretical advances with the limitations of all existing data sets, including the Census data. However, the econometric and data deficiencies of the earlier studies are more remediable. Most importantly, the 1970 Public Use Sample, unlike its 1960 predecessor, permits the identification of households by SMSA. Hence, SMSA averages of the desired variables may be calculated from an extract of appropriate households and families. These data also facilitate the use of more desirable estimation procedures, including recently developed corrections for sample selection bias. Estimates of a model of labor supply and completed fertility using such SMSA averages from the Public Use Sample may be compared to those obtained in previous studies using the published averages from the 1970 Census.

The second section of this paper contains a brief discussion of previous models and theoretical issues. The model, data and estimation methods are described in the third section. In the final section, we compare the estimates of prior studies with those obtained from the 1970 data.

2. THEORETICAL ISSUES

Equation (1) is a labor supply function which is representative of those estimated by Mincer, Cain, and Bowen and Finegan with SMSA data from the 1940, 1950, and 1960 Censuses.[1]

$$L = \alpha_0 + \alpha_1 W + \alpha_2 H + \alpha_3 N + \alpha_4 E + \alpha_5 F + \alpha_6 S + \alpha_7 U + \varepsilon \qquad (1)$$

where

L = the labor force participation rate of married women, husband present, age 14+, during the survey week.

W = the median income of females who worked 50-52 weeks in the pre-census year.

H = the median income of male family heads, spouse present in the pre-census year.

N = a measure of average non-employment income in the pre-census year.

E = median years of school completed by females age 25+.

F = a measure of the number or presence of children.

S = a dummy variable indicating whether the SMSA is South or non-South.

U = the civilian male unemployment rate during the survey week.

ε = disturbance term.

These studies produced relatively stable (across censuses) and statistically significant estimates of the above coefficients. W was intended to measure differences in the permanent earnings capacity of married women. H and N measured variation in family income from sources assumed to be unrelated to the wife's labor supply, thus, permitting the estimation of pure income effects. Estimates of these variables' coefficients yielded the anticipated positive wage and negative income effects with few exceptions.

A measure of current labor market conditions, U, was included to control for the timing impact of transitory variation in wages and income common to families within each SMSA.[2] The consistently negative estimates of α_7 indicate a dominant discouraged-worker (vs. added-worker) effect, although this interpretation has been challenged.[3] Education generally had a positive coefficient in accord with its role as a measure of tastes for market work and access to jobs with nonpecuniary advantages. The negative impact of the number of children or the presence of preschool age children was interpreted as either a long run or temporary variation in labor supply depending on the measure of F. The dummy variable for southern SMSAs had a positive sign in 1940 and 1950, but by 1960 this coefficient was insignificant.

Theoretical criticisms of the model in (1) can be grouped into two classes. First are the strong assumptions needed to interpret the estimates of α_1 and α_2 (or α_3) as the usual (uncompensated) wage and income effects respectively, even granted the exogeneity of these variables (see Heckman 1978). Second, there exist ample theoretical bases for questioning the exogeneity of most of the independent variables in (1). More specifically, each variable potentially reflects a decision determined in part by the unobservable factors represented in the disturbance, ε. Our discussion will focus on the latter problem.

The assumption of a predetermined allocation of the husband's time was relaxed by Ashenfelter and Heckman (1974) in a two-equation model of the couple's labor supply decisions. Some support is provided for the specification in (1) by their finding of an income effect (mainly from the husband's earnings) on the wife's labor supply which was substantially larger (in absolute value) than the cross-substitution effect of the husband's wage.

Current property income, as reflected in N, is most likely related to previous market work decisions. The exogeneity of the male unemployment rate is also suspect given that the wife's earnings may affect the cost, and hence the length, of the husband's job search. Insufficient data, however, has limited the empirical efforts to relax these assumptions.[4]

The role of individual choice in the determination of market wages has, of course, been explored extensively in the human capital literature. Most relevant to (1), Mincer and Polachek (1974) have confirmed the strong relationship between the wages of married women and prior work experience representing human capital accumulation. They also point out the identification problem inherent in such a procedure. That is, do such coefficient estimates measure the impact of experience on wages? Or, is experience a measure of labor supply that is influenced by permanent variation in earnings capacity. Finally, does some third factor, such as unobservable tastes for market goods, influence both work and investment throughout the life cycle.[5] This last possibility raises doubts concerning the interpretation of α_1 in (1) as the effect of exogenously imposed price variation.

The difficulty of using cross-sectional data to estimate wage effects is most clearly indicated by recent life cycle models of time allocation and human capital investment such as Heckman (1976). In this model, the optimal levels of work, consumption and investment throughout the life cycle are functions of a common set of variables: permanent individual characteristics, such as tastes and endowments of financial and human capital; the technology of human capital production; and market prices, such as the rental rate of human capital, that is, the price of quality constant labor. Average wages throughout the life cycle (permanent wages) reflect a series of decisions made jointly with, and subject to the same determinants as, life cycle labor supply.

Most cross-sectional labor supply analyses attempt to estimate the market work impact of permanent wage variation through the use of "auxiliary regressions" to predict wages for each observation.[6] However, most of the wage determinants commonly employed, e.g., education, experience, or residence (such as studies with SMSA data), represent decisions which from the life cycle perspective are influenced by such unobserved labor supply determinants as tastes and home productivity. The above estimation procedure assumes that such unobservable variables are distributed independently across time periods, conditional on available "control" variables, such as age and number of children. If not, the wage measures will be correlated with the current labor supply disturbance and the estimated wage coefficient will measure some mixture of price effects and those of any permanent unmeasured determinants of work and investment.[7]

The assumptions of perfect foresight and planning in the life cycle models are extreme. However, such a perspective highlights the equally tenuous assumptions underlying many cross-sectional analyses, that is, work and investment as a sequence of decisions whose unmeasured determinants are independently distributed over the life cycle. In short, the chronological order of related decisions, such as human capital investment and subsequent labor supply, does not justify the presumption of causal relationships.

These considerations also provide a basis for questioning the specification in (1) of fertility behavior as an exogenous determinant of labor supply. Modern economic analysis of fertility behavior can be traced to Leibenstein (1957) and Becker (1960). Willis (1974) provides the most elaborate theoretical framework for the consideration of fertility and labor supply as related decisions.[8] Of particular relevance for this study is the empirical analysis of completed fertility by Cain and Weininger (1973), using 1940 and 1960 SMSA data. Their model is presented in (2) along with variable definitions for the 1960 Census. Differences between

(1) and (2) in the wage and income measures are due to the estimation of age-segregated regressions in the latter study.[9]

$$F = \beta_0 + \beta_1 W + \beta_2 H + \beta_3 E' + \beta_4 S + \beta_5 C + u \qquad (2)$$

where

F = children ever born per 1,000 women ever married.

W = female wage defined as the median income in the pre-census year of females divided by the average number of weeks worked by females who worked, multiplied by 52.

H = mean male income in the pre-census year.

E'= percentage of females with less than 5 years of schooling.

S = dummy variable indicating whether the SMSA is in the South or not.

C = proportion of school age population in private elementary and secondary schools as a measure of the proportion Catholic of the population.

u = disturbance term.

As with labor supply, the focus of this fertility analysis was on wage and income effects. The estimates of β_1 and β_2 were generally significant with the expected negative and positive signs. E' was included as a measure of contraceptive efficacy and had the anticipated positive effect. The estimated impact of C, the measure of the proportion Catholic of the SMSAs population, was insignificant and of unstable sign. The coefficient for S was negative and highly significant.

There are numerous problems with the specification in (2). The estimates may be biased due to the failure to account for such factors as imperfect fertility control (Heckman and Willis 1974), tastes (Easterlin 1977) and childbearing costs (Turchi 1975). The resolution of these shortcomings exceeds the objectives (and data) of the present study and the reader is referred to the original article by Cain and Weininger for further discussion of these points. Of most relevance for our present purposes is the support from the SMSA data itself for the characterization of fertility as a resource allocation decision, sensitive to many of the same factors as labor supply. As such, it may be most appropriate to view their (labor supply and fertility) association as one of related decisions rather than the simple causal relationship assumed in (1).

Variation in the numbers and/or ages of children undoubtedly does have some causal effect upon the timing and permanent levels of labor supply. Its estimate would be useful in predicting the eventual labor supply impact of any variable whose influence on fertility might be known a priori, for example, changes in contraceptive technology. However, a consistent estimate of α_6 in (1) with a single equation specification requires the assumption of either unplanned fertility or an independent distribution (over time) of the unmeasured determinants of time allocation between home and market. Otherwise, fertility measures will be correlated with the labor supply disturbance, and the estimates of α_6 in (1) will measure some mixture of the effects of fertility and that of any permanent unmeasured determinants of time allocation, such as tastes for market goods. The problem is identical to that which complicates the estimation of wage effects upon labor supply, that is, the identification of the net effect of constraints imposed by either the market or previous household decisions.

An analysis similar to that of the labor supply function would also raise doubts concerning the interpretation of β_2 and β_3 in (2) as estimates of exogenously imposed wage and income levels. Estimates of the former may be biased due to joint determination of the wife's human capital investment and fertility decisions, and the latter by dependence of the husband's labor supply upon family size.

The previous discussion has emphasized problems of estimating theoretically congruent models of the labor supply and fertility of married women. The sizeable gap between the demands of recent theoretical developments and the capacity of existing data sets, especially that of the Census, calls for compromise and the use of models whose specification reflects both research goals and the nature of the available data.

3. THE MODEL, DATA AND ESTIMATION METHODS

A. A Joint Model of Labor Supply and Fertility

Equations (3) and (4) specify the bivariate model of labor supply and completed fertility estimates for this study.

$$L = \alpha_0 + \alpha_1 W + \alpha_2 H + \alpha_3 N + \alpha_4 E + \alpha_5 R + \alpha_6 C + \alpha_7 U + \varepsilon_1 \tag{3}$$

$$F = \beta_0 + \beta_1 W + \beta_2 H + \beta_3 N + \beta_4 E + \beta_5 R + \beta_6 C + \varepsilon_2 \tag{4}$$

Two sets of SMSA data from the 1970 Census were used along with some noncensus measures. (See Section 3 B and Appendix A for further information.) Due to the wide variety of data employed, only general variable definitions are presented below. Most variables are age and race specific.

L = labor supply of married women.

F = number of children ever born to married women.

W = wages (weekly or hourly) of females or married women.

H = annual income or earnings of married men.

N = non-work conditioned income of families or married couples.

E = years of schooling completed by females or married women.

R = proportion of rural residents in the SMSA.

C = measure of proportion Catholic of the SMSA.

U = measure of labor market conditions, usually civilian male unemployment rate.

$\varepsilon_1, \varepsilon_2$ = disturbances assumed to have a bivariate normal distribution with zero means and variance-covariance matrix = $\{\sigma_{ij}/n_k\}$ for i,j = 1,2 and n_k = the number of couples included in the averages for the kth SMSA.

The only variable which has not been discussed previously is R, the proportion of rural residents in the SMSA. Virtually all of the SMSAs included in our sample have some rural residents, with the proportion rising as high as 40% in some instances.[10] Numerous studies have confirmed the association between rural residence, and both labor supply and fertility (for example, see Sweet 1974 and Rindfuss and Sweet 1977), although the exogeneity of household location might be questioned. The dummy variable for southern SMSAs in (1) and (2) has been excluded from our model reflecting both the difficulty of interpreting this (and other regional) coefficients and the minimal impact of its exclusion in previous studies.

The bivariate model in (3) and (4) obviously does little to rectify the basic problem stressed in Section 2, that is, the potential endogeneity of the independent variables. The estimation of more complex models of household decision-making requires the use of more elaborate data than Census provides.

As indicated previously, the likelihood of endogeneity may be especially acute in the case of the wife's wage. Grouped data have been used in recognition of numerous factors which could cause current individual wages (and incomes) to depart from their permanent levels. This constitutes one form of the "auxiliary wage regression" method, referred to in the previous section, in that the use of weighted averages provides estimates identical to those which would be obtained by replacing the original household observations with SMSA averages, i.e., the predicted values of each variable resulting from the regression of the micro data on a series of dummy variables for SMSA residence. This estimation method, however, does not remedy the problem of wage endogeneity if residence is correlated with the disturbances in (3) and (4) due to selective migration, i.e., the determination of both residential and labor supply decisions by unmeasured variables such as tastes. Cain and Weininger (1973) have suggested that the residential choices of the couple may be only minimally responsive to the earnings opportunities of the wife. However, this remains an unresolved empirical question and additional factors may induce correlation between residence and the disturbances in (3) and (4). These factors include selective migration (and subsequent marriage) by single women and the fact that the preferences of husbands, whose job opportunities should influence migration, are assumed to be measured in ε_1 and ε_2. Hence, the use of SMSA residence for permanent wage and income predictions may give rise to the same bias induced by other commonly used indicators of human capital investment decisions, such as education and experience.[11]

An earlier study by Cain and Dooley (1976) did incorporate wages along with the labor supply and fertility of married women into a simultaneous, three-equation model. The absence of both a wage equation and the reciprocal effects of the dependent variables on each other distinguish (3) and (4) from this previous effort. Several factors were responsible for this modification.

First, the theoretical case for the estimation of reciprocal effects is less compelling here than in the case of traditional simultaneous models. Specifically, labor supply and fertility are related decisions of a single couple rather than behavioral functions for distinct economic agents, as with market demand and supply functions. Second, and as a consequence of the foregoing, the identifying restrictions employed to estimate such reciprocal effects are only weakly justifiable in the Cain-Dooley model. Third, data limitations did not permit the specification of a wage equation which would address the fundamental problems raised by the interrelation of labor supply and human capital decisions over the life cycle. Moreover, the assumption of a less than perfectly elastic demand for the labor of married women implies the need to incorporate demand and supply functions for substitute sources of labor. Such an elaborate extension of the model was unfeasible for both the Cain-Dooley and the present study. Finally, the empirical results of Cain and Dooley were disappointing; particularly the estimates for the wage equation and the unreasonably large wage and income elasticities for the labor supply function.[12]

The only restriction in (3) and (4) is the exclusion of the measure of transitory labor market conditions, U, from the fertility equation.[13] Current measures of several other variables are used, but their grouped values are assumed to measure "long-run" or permanent differences in constraints—an assumption which is crucial in the case of the fertility equation. One unusual feature of the model, which results from the bivariate specification, is the inclusion of a measure of the proportion Catholic, C, in the labor supply function.

The labor supply function in (3) also differs from that estimated by Bowen and Finegan (1969) in that several labor market variables used in the latter are excluded from our specification. Most of these excluded variables, e.g., indices of demand and the supply of substitute labor, call for a more elaborate model than that of a labor supply function for a single demographic subgroup.

B. The Data

Our model was estimated with two sets of SMSA data from the 1970 U.S. Census: the averages published by the Census Bureau, similar to the measures used in previous studies; and averages calculated from an extract of the 1/100 Public Use Sample. Published averages for 1970 were available for the 125 SMSAs with a population greater than 250,000. Data for all such SMSAs except Honolulu was included in the white, non-Spanish sample.[14] Separate data for the black population were provided in 79 SMSAs. Data for the Spanish population were not used extensively due to small sample size (38 SMSAs), very weak preliminary results, and the heterogeneity of this group's characteristics. Appendix A contains the variable definitions and descriptive statistics.

The major advantages of the Public Use Sample (henceforth PBUS) are the availability of unpublished variables, such as hourly wages and annual hours of work, and the ability to extract a theoretically appropriate sub-sample of households and families. Household identification by metropolitan area is restricted to those SMSAs or groups of adjacent SMSAs with population greater than 250,000. From these areas only nonfarm married couples were selected, both spouses present with a wife age 30-54, physically able to work and a husband, black or white of non-Spanish origin, currently in the civilian labor force. Sub-families were permitted in the sample, although their proportion is extremely small given the age restrictions. Adult married males not in the labor force commonly exhibit severe health problems and military personnel are generally insulated from local labor market conditions. The wages and labor supply of farm residents present serious measurement problems. The lower age limit reflects our focus upon completed (or nearly completed) fertility. Both upper and lower age limits result from the need to measure permanent differences in wages, income, and labor supply.

Only those SMSAs with at least fifty couples meeting the above specifications were used. This criterion was easily met for white, non-Spanish (henceforth white) population in all 134 possible SMSAs (or groups), but limited the black sample to 45 observation units.

The 5% sample was used for all variables except two indirect measures of the proportion Catholic of the SMSA which were derived from the 15% PBUS sample. The selection criteria for the 15% sample were almost identical to those listed above. (See Appendix A for definitions and descriptive statistics of all variables from the PBUS.)

C. Estimation Methods

The exclusion of the unemployment rate (U) from the fertility equation in our model can be exploited to obtain lower variance estimates of the labor supply function via the method of "seemingly unrelated regressions." Due to the model specification however, only the labor supply estimates differ from those which would be obtained with ordinary least squares and do so minimally in most instances. The data were also appropriately weighted to correct for the heteroscedasticity induced by grouping. Separate regressions were estimated for the white and black samples, and for five-year, age-of-wife cohorts.[15] The advantage of the age-segregated regressions was the possibility of variation in structural parameters given the diverse fertility behavior of the cohorts in our sample.[16] The 1970 values of a cost-of-living index were available for 61 of the SMSAs in our sample.[17] There were, however, generally very minor differences between the coefficients obtained with real versus nominal measures of wages and income when the model was estimated for this subset of SMSAs. The exceptions will be noted in the text.

Steps were also taken to deal with two additional possible sources of bias in previous SMSA studies: the linear labor force participation rate (henceforth LFPR) equation and sample selection bias.

Previous studies of LFPRs have most commonly used a linear function. An LFPR function, however, is the grouped data counterpart of a conditional probability function of labor force participation for which a nonlinear form is most appropriate. One commonly used conditional probability function is the Probit function where

$$\text{Prob}(y_i = 1|\underline{x}_i) = F(\underline{x}_i{}'\underline{\beta}), \tag{5}$$

where

$$y_i = \begin{cases} 1 \text{ if the woman is in the labor force} \\ 0 \text{ otherwise} \end{cases}$$

$\underline{x}_i$ = a vector of independent variables

$F(.)$ = the cumulative standard normal density function

Grouping <u>exactly</u> on individual values of $\underline{x}_i$,

$$p_i = F(\underline{x}_i{}'\underline{\beta}) + u_i = P_i + u_i \tag{6}$$

$$q_i = 1 - F(\underline{x}_i{}'\underline{\beta}) - u_i = Q_i - u_i \tag{7}$$

where

p_i = proportion of observations at $\underline{x}_i$ with $y = 1$

q_i = proportion of observations at $\underline{x}_i$ with $y = 0$

and u_i has a binomial distribution with

$$E(u_i) = 0$$

$$\text{Var}(u_i) = P_i Q_i / n_i$$

where

n_i = number of observations at each $\underline{x}_i$ for $i = 1, \ldots, N$.

It can be shown (Goldberger 1974) that the following normit relation holds approximately,

$$F^{-1}(p_i) = \underline{x}_i{}'\underline{\beta} + v_i \tag{8}$$

where

$$E(v_i) = 0, \text{ Var } (v_i) = \frac{1}{n_i P_i Q_i}$$

Hence, equation (8) provides the opportunity to estimate the parameters of an appropriately bounded conditional probability function using grouped data in a linear regression. Specifically, (8) can be estimated with SMSA data by substituting the observed LFPRs for the p_i and the real averages of the independent variables in (3) for the $\underline{x}_i$. The LFPRs also provide consistent estimates of the p_i and Q_i needed to adjust for heteroscedasticity.

The use of age-race specific SMSA averages obviously does not constitute exact grouping on the basis of the independent variables. Hence, possible bias may arise from the nonlinearity of the normit transformation, $F^{-1}(.)$. That is, the normit of the areal LFPR will generally provide a biased estimate of the SMSA average of the $F^{-1}(p_i)$ associated with each value of $\underline{x}_i$.[18] However, such measurement error would appear to be slight given the apparently narrow dispersion of the p_i <u>within</u> each SMSA.[19] Hence, LFPR functions were estimated with the PBUS data using both the normit transformation and the linear function in (3).

A more serious econometric problem with labor supply studies of groups such as married women is sample selection bias. Several articles (Heckman 1974, Gronau 1974) have shown that the use of data on labor force participants alone to predict market wages for the entire sample, as has been the procedure in previous SMSA studies, will lead to inconsistent estimates of wages and, therefore, of the labor supply function.

Heckman (1974) assumes the following recursive model for the determination of labor force participation and hours (or weeks) of market work. (Individual subscripts have been omitted for convenience.)

$$S = \alpha_0 + \alpha_1 L + \underline{x}'\underline{\alpha}_2 + \varepsilon_1 \tag{9}$$

$$W = \gamma_0 + \underline{z}'\underline{\gamma}_1 + \varepsilon_2 \tag{10}$$

where

S = the wife's shadow wage, i.e., the monetary value of an additional hour of home time.

W = the wife's market wage offer.

L = annual hours of market work.

$\underline{x},\underline{z}$ = vectors of exogenous variables which may share common elements.

The disturbances, ε_1, and ε_2, are assumed to have a bivariate normal distribution with zero mean and variance -covariance matrix $= \{\sigma_{ij}\}$ for $i,j = 1,2$. The condition for labor force participation is $W>S$ at $L = 0$ or

$$\gamma_0 + \underline{z}'\underline{\gamma}_1 - \alpha_0 - \underline{x}'\underline{\alpha}_2 > \varepsilon_1 - \varepsilon_2 \tag{11}$$

Condition (11) implies that the conditional probability function for labor force participation is cumulative normal as in equation (5) above, that is,

$$\Pr(L>0|\underline{x},\underline{z}) = F(\phi) \tag{12}$$

where

$F(.)$ = the cumulative standard normal density function

$$\phi = \{\gamma_0 + \underline{z}'\underline{\gamma}_1 - \alpha_0 - \underline{x}'\underline{\alpha}_2\} \;/\sigma^*$$

$$\sigma^* = (\sigma_{22} + \sigma_{11} - 2\sigma_{12})^{\frac{1}{2}}$$

If condition (11) holds, then hours of work adjust so that $W = S$. Observed hours of work are then

$$L = \frac{\gamma_0 + \underline{z}'\underline{\gamma}_1 - \alpha_0 - \underline{x}'\underline{\alpha}_2}{\alpha_1} + \frac{\varepsilon_2 - \varepsilon_1}{\alpha_1} \tag{13}$$

In a second paper, Heckman (1975) presents the conditional expectations of hours and wages in the working portion of the sample.

$$E(L|\underline{x},\underline{z},L>0) = \frac{\gamma_0 + \underline{z}'\underline{\gamma}_1 - \alpha_0 - \underline{x}'\underline{\alpha}_2}{\alpha_1} + \frac{\lambda\sigma*}{\alpha_1} \tag{14}$$

$$E(W|\underline{x},\underline{z},L>0) = \gamma_0 + \underline{z}'\underline{\gamma}_1 + \lambda\frac{\sigma_{22} - \sigma_{12}}{\sigma*} \tag{15}$$

where

$f(.)$ = the standard normal density function

$$\lambda(\underline{x},\underline{z}) = \frac{-f(\phi)}{F(\phi)}$$

Note that λ is a function of $\underline{x}$ and $\underline{z}$. Therefore, the use of data for labor force participants alone to estimate (10) will produce inconsistent coefficient estimates and wage predictions due to omitted variable bias, as shown by (15). Such wage predictions will also lead to inconsistent labor supply estimates if substituted (for $\gamma_0 + \underline{z}'\underline{\gamma}_1$) into (13). Note further that the use of SMSA data will not remeay this problem, since the SMSA averages of (15) will not be equal to those of (10). That is, average wages among working women will be an inconsistent estimate of average available wages in the full population. Hence, our estimates of (3) and (4) will be subject to this bias.

The general problem is that a common unobserved variable(s) determines both the behavioral relation of interest and selection into the sample used for estimation. In this specific case, ε_1 and ε_2 determine both wages and the participation decision. A variety of work-related characteristics used to delimit our PBUS extract, such as marital status and residence, may give rise to this problem. However, our corrective efforts were confined to wage bias.

Heckman has (1975, 1977) proposed a relatively simple remedy. First, probit analysis may be used to estimate (12) with data on $\underline{x}$ and $\underline{z}$ for the entire sample. These results provide consistent estimates of ϕ, and hence, λ. Finally, the estimates of λ may be used to estimate (14) and (15) with data from the working portion of the sample.

The above procedure could possibly be used to estimate an annual hours (or weeks) function with SMSA average data from our PBUS extract. However, the first step, the estimation of an LFPR function for the entire sample, would require good measures of the determinants of inter-SMSA wage differentials. The disappointing results for the wage equation in Cain-Dooley (1976) indicates that such data is not readily available and, therefore, an alternative approach is proposed below.

Substituting (10) into (13) provides an alternative expression for the hours of work equation.

$$L = \frac{W - \alpha_0 - \underline{x}'\underline{\alpha}_2}{\alpha_1} - \frac{\varepsilon_1}{\alpha_1} \tag{16}$$

Taking expectations,

$$E(L|\underline{x},\underline{z},L>0) = \frac{E(W|\underline{x},\underline{z},L>0) - \alpha_0 - \underline{x}'\underline{\alpha}_2}{\alpha_1} - \frac{\lambda\delta}{\alpha_1} \tag{17}$$

Using results from Johnson and Kotz (1972), it can be shown that <u>among labor force participants</u>,

$$L = E(L|\underline{x},\underline{z},L>0) + u \tag{18}$$

where

$$E(u) = 0$$

$$V(u) = \frac{\sigma_{11}}{\alpha_1^2} \{\delta^2(1-\phi\lambda-\lambda^2) + 1 - \delta^2\}$$

$$\delta = (\sigma_{11} + \sigma_{12}) \;/\; \sigma^*(\sigma_{11})^{\frac{1}{2}}$$

Average wages among working wives in the PBUS extract are known and provide a consistent estimate of the SMSA mean of $E(W|\underline{x},\underline{z},L>0)$. Hence, given an estimate of the SMSA mean of λ, average data among working wives could be used to estimate the labor supply coefficients in (18), that is, α_0, α_1 and $\underline{\alpha}_2$.

Assume the LFPR to be a function of the SMSA means of $\underline{x}$, $\underline{z}$, ε_1 and ε_2. If ε_1 and ε_2 are mean independent of SMSA residence, as must be assumed to obtain consistent coefficient estimates of the LFPR equation, then the observed LFPR will be asymptotically uncorrelated with ε_1 and ε_2. More precisely, the observed LFPR will be asymptotically equal to its conditional mean in the population. The observed LFPR will then provide an estimate of the SMSA mean of $F(\phi)$. This estimate, in turn, can be used to estimate the SMSA mean of λ. This method is similar to that used by Gronan (1974) and completes the information necessary to obtain consistent estimates of (18). Of course, steps must be taken to correct for the heteroscedastic disturbance variance which, with grouped data, would be equal to $V(u)/n_i$ where n_i is the number of individual observations in the ith SMSA.

One characteristic of census data, however, complicates the above procedure. The Heckman method assumes that the sample observations can be divided into two groups: labor force participants and nonparticipants during a given period, usually a year. In contrast, the 1970 PBUS provides measures of earnings and weeks worked for 1969 and hours worked for the survey week in 1970. Hence, annual hours of work and hourly wages can be calculated only for those wives who worked during both periods (henceforth referred to as survey week workers). This subsample will not provide consistent estimates of an annual hours (or weeks) equation, such as (18), even if it is assumed that behavior during the 1970 survey week is typical of any given week during 1969. This results from the fact that workers during a given week do not, in general, constitute a random sample of all annual participants. Rather, the probability of weekly participation is equal to the proportion of weeks worked, assuming labor supply to be distributed randomly throughout the year, and, hence, is not independent of annual labor supply.

Hanoch (1976) has derived a method for estimating an annual hours (or weeks) equation from such a sample, assuming that survey week behavior is typical of that in 1969. The available evidence from the PBUS indicates that this may be a reasonable assumption (see Dooley 1977, p. 95). The implementation of this technique is straightforward, but the formal proof of the consistency of the resulting estimates is quite complex. Hence, the reader is referred to the original Hanoch article (1976, Appendix) for such a proof.

Equation (18) is an annual hours (or weeks) function for <u>all</u> annual participants and, as such, would normally require observations for all 1969 workers. Hanoch demonstrates, however, that consistent estimates of (18) may be obtained with data for <u>survey week</u> workers alone if each observation is weighted by

$(L'_{wi}/L_{wi})^{\frac{1}{2}}$ where

L'_{wi} = the conditional expectation of weeks worked among all annual (1969) workers for the ith wife.

L_{wi} = actual weeks worked in 1969 by the ith wife.

The above method may be readily adopted for use with our grouped data from the PBUS. The SMSA average of L_{wi} (weeks worked by survey week workers) is directly observable from the data. Average weeks worked among <u>all</u> 1969 workers will provide a consistent estimate of the SMSA mean of L'_{wi}, if the unobservable determinants of weeks worked are assumed to be mean independent of SMSA residence. That is, <u>observed</u> average weeks worked will be asymptotically equal to their conditional expectation in the population.[20] Hanoch further shows that such weighting induces no heteroscedasticity beyond that already present in (18).

Following the above procedure, (18) was estimated for both annual hours and weeks using SMSA averages from the PBUS for couples where the wife worked in both 1969 and the 1970 survey week. The published 1970 SMSA averages do not provide the necessary data. (18) was <u>not</u> estimated jointly with a fertility equation. The incorporation of fertility would have complicated greatly the analysis of sample selection bias and the Heckman-Hanoch estimation method guarantees only consistent estimates in any event. However, the estimates of (18) can be compared with the labor supply estimates from our joint model, (3) and (4), to assess the extent of sample selection bias.

4. THE EMPIRICAL RESULTS

Three sets of estimates are presented and discussed in this section: the single equation labor supply and fertility results reported by Cain (1966) and Cain and Weininger (1973) using SMSA data from the 1940, 1950, and 1960 Censuses; the estimates of (3) and (4) obtained with the 1970 published SMSA averages; and the estimates provided by SMSA averages from the 1970 PBUS extract of both our joint model and (18), the labor supply function adjusted for sample selection bias.

A. Estimates from Previous Studies with SMSA Data

Equation (1) presents the basic labor supply function estimated by Cain (1966) along with the definition for most variables. Tables 1 and 2 contain Cain's results for the total and nonwhite samples respectively.[21] Table 3 presents the wage and income elasticities evaluated at the sample mean.[22]

Two aspects of these findings are particularly noteworthy. First is the significantly negative effect of unemployment (U) on labor supply, with the exception of the nonwhites in 1950. This appeared to provide a clear indication of the larger absolute size of the discouraged-worker, as opposed to the added-worker, effect of changes in employment opportunities over the business cycle. As we have previously noted, however, Mincer (1966) and others expressed reservations concerning this interpretation.[23] The reason for the large increase in the absolute size of this coefficient from 1950 to 1960 is not entirely clear, but Fields (1976) has suggested a decline in the added-worker effect due to growing unemployment compensation. Second, the estimates for the wife's wage (W), husband's income (H), and nonlabor income (N) are significant and of the predicted sign. Both H and N are assumed to measure pure income differences and, therefore, should have similar coefficients. The variation in their size for 1960 may be attributable to either poor measurement or the inadequacies of the single equation specification.

The decline in wage elasticities for both samples from 1950 to 1960 was attributed by Cain (1966, p. 60) to the increase in the effect of education (E): the relative increase in white collar employment and associated non-pecuniary compensation was causing education to assume a larger role as a proxy for the total wage effect. There was less interest in a comparison with the 1940 estimates due to differences in the available measures and in economic conditions between the pre-WWII and post-WWII periods. Finally, the predictions generated by the above estimates were approximately in accord with the historical pattern of the LFPRs of married women over the (then) preceding 30 to 40 years.

Ashenfelter and Heckman (1974) also used the 1960 SMSA data to estimate a two equation model of the LFPRs of both married men and women. Their estimate of the wife's uncompensated own wage elasticity was .87, much closer to Cain's figure for the 1950 data. Their estimate of the income coefficient (the elasticity was not reported) for the wife's labor supply is -.89 which is approximately the average of Cain's 1960 estimates for H and N in Table 1.

Table 1

Estimates of the Labor Supply Equation for Total Married Women,

Husband Present, Age 14+, 1940, 1950, and 1960.

(Labor Supply = LFPR in the Survey Week)

Year	Constant	W[a]	H[b]	N[c]	E	F[d]	S	U
		($00's)	($00's)	($00's)				
1940	42.9*	1.32*	-1.02*		.11	-.94*	4.55*	-.43*
1950	37.9*	1.04*	- .43*	- .50+	.73*	-.48*	3.41*	-.40*
1960[e]	33.7*	.39*	-.46*	-1.61*	1.2*	-.09*	0.34	-1.89*[f]

a - In 1940, median income of all females with income.

b - In 1940, median income of all males with income.

c - In 1950, median income, excluding wages and self-employment income, per recipient of such income, age 14+. In 1960, the same measure per recipient of any income, age 14+. In 1940, no suitable measure was available.

d - In 1940, number of children under 5 years of age per 100 women ever married. In 1950, % of husband wife families with children under 18 years of age. In 1960, number of children ever born per 1,000 women ever married.

e - These estimates are from Cain (1967).

f - This estimate is from Cain and Mincer (1969). A data error in the measurement of U in Cain (1967) had been detected after publication. The correction for this error had little impact on the coefficient estimates of the other variables.

* indicates t-ratio greater than 1.96.
~~+ indicates t-ratio~~ greater than 1.645 (but less than 1.96).
° indicates t-ratio greater than 1.282 (but less than 1.645).

Table 2

Estimates of the Labor Supply Equation for Nonwhite

Married Women, Age 14+, 1950 and 1960.

(Labor Supply = LFPR in the Survey Week)

Year	Constant	W^a	H^b	E	F^c	S	U
1950	45.1*	2.85*	-1.35*		-.25	6.99°	-.16
1960	47.4*	1.12*	-.97*	1.53°	.00	1.9	-.55*

a - In 1950, median income of nonwhite females with income.
In 1960, median earnings of the nonwhite female civilian labor force.

b - In 1950, median income of nonwhite males with income.
In 1960, median earnings of the nonwhite male civilian labor force.

c - In 1950, % of nonwhite husband-wife families with children less than 18 years of age.
In 1960, the number of children ever born per 1,000 nonwhite women ever married.

* indicates t-ratio greater than 1.96.
\+ indicates t-ratio greater than 1.645 (but less than 1.96).
° indicates t-ratio greater than 1.282 (but less than 1.645).

Table 3

Wage and Income Elasticities of the Labor Force

Particiation Equation for 1940, 1950 and 1960.

Total and Nonwhite Married Women.

Total Population	1940	1950	1960
W	.52	.91	.40
H	-.62	-.58	-.77
Nonwhite Population			
W		.69	.39
H		-.71	-.74

The estimates of Bowen and Finegan (1969) generally coincide with those of Cain and have recently been updated by Fields (1976) with 1970 SMSA data. The most important of Fields' findings, for our purposes, were the lack of significant coefficient estimates for either income measure and a continued decline in the size of the wage elasticity. Fields follows Bowen and Finegan in interpreting these phenomena as a reflection of nonlinearity in the labor supply function, i.e., increases in average labor supply leading to smaller impacts of wage and income variation.

The model of completed fertility estimated by Cain and Weininger (1973) with 1940 and 1960 SMSA data is presented in equation (2) along with variable definitions for the latter period. The results are presented in Tables 4 and 5.

The estimated wage and income effects were consistent with theoretical predictions and generally significant.[24] The average income inelasticities with 1960 data, evaluated at the mean, were .30 for whites and .08 for nonwhites, although this racial differential disappeared for the youngest age groups. The average uncompensated wage elasticities were -.37 for whites and -.29 for nonwhites in 1960. The 1940 data produced larger elasticities, averaging .49 in the case of income and -.81 for the wife's wage. As expected, the effect of the education variable (E') was significant and positive, confirming that very low levels of education lead to higher fertility. The nonsignificant impact of the proportion Catholic (C) was attributed to imprecise measurement or possibly the failure to control for

Table 4

Estimates of the Fertility Demand Function for White and Nonwhite Women, 1960

(Fertility = Children Ever Born per 1,000 Women Ever Married)

Age Group	Constant	W ($00's)	H (00's)	E'	S	C
White Women[a]						
30-34	3070*	-28.7*	14.9	3.5+	-79.5*	.4
35-44	3421*	-30.4*	13.0*	13.0	-161.9*	-.9
45-49	3540*	-28.0*	8.9+	38.0*	-225.5*	-3.3
Nonwhite Women						
30-34	4096*	-50.5*	24.4*	46.8*	-110.5	
35-44	4365*	-29.0	.2	30.5*	-213.7+	
45-49	4639*	-13.6	-15.9+	12.9*	-255.7*	

a - The regressions for both races included the log of the population and those for white women, a measure of the percent of the population of Mexican origin. The exclusion of either variable had little effect on the above estimates.

* - indicates t-ratio greater than 1.96.

+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Table 5

Estimates of the Fertility Demand Function for White Women

1940

(Fertility = Children Ever Born per 1,000 Women Ever Married)

Age Group[a]	Constant	W[b] ($00's)	H[c] ($00's)	E'	C[d]
30-34	1454*	- 66.8*	45.3[o]	22.4*	1.0
35-39	1842*	-135.5*	49.7[o]	33.6*	2.1
40-44	2083*	-232.7*	113.4*	38.3*	-.41

a - These regressions also included measures of the percentage of males and females in various white collar occupations, the exclusion of which did little to change the above results.

b - Median income in 1939 of females divided by average months worked in 1939, multiplied by 12.

c - Median earnings for males in experienced labor force with $900 or more income and not on public emergency work.

d - Measure of proportion Catholic from 1936 survey of religious affiliation (see Grabil, Kiser and Whelpton 1958).

* - indicates t-ratio greater than 1.96.

+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

the rural proportion of the population, i.e., the generally urban residences of Catholics serving to offset the positive effect of religious affiliation. However, unpublished results of this author failed to substantiate the latter explanation.

Cain and Weininger readily acknowledge that their linear specification cannot explain fully the actual post-WWII pattern of fertility behavior. However, these results do provide impressive empirical confirmation of economic hypotheses in light of the frequently disappointing estimates of other studies in this area (see T.P. Schultz 1974).

B. Estimates with Published SMSA Data from the 1970 Census

In this section, we report the estimates of (3) and (4) obtained with variable measures from the 1970 Census similar to those used in previous studies. The data and estimation methods are described in sections 3B and 3C and Appendix A.

Table 6 contains the results for the white sample using the LFPR in the survey week as the measure of labor supply. As with previous studies, the estimated coefficients for the wife's wage (W) and husband's income (H) are consistently significant and of the predicted sign.[25] Hence, Fields' (1976) finding of an insignificant impact of H with 1970 data is not supported, although the results in Table 6 do show a decline in the income elasticity (see below). This may be due to specification differences or the absence of age segregated regressions in Fields' study. The uncompensated wage elasticities, evaluated at the mean, range from .28 to .77. The value for all ages combined is .67, approximately midway between the estimates of .91 and .40 found by Cain with 1950 and 1960 data respectively (see Table 3). The husband's income elasticities range from -.21 to -.50. The estimate for the entire sample is -.49 in contrast to Cain's findings of -.58 for 1950 and -.77 for 1960. Hence, Cain's observation of a secular decline in the wage relative to the income effect is not supported. This ratio fell from an absolute value of 1.6 to .52 in 1950-1960 interval, but rose to 1.4 in the subsequent decade.

The coefficient estimates for nonlabor income (N) in Table 6 are relatively stable across age and of the predicted sign, but significant in only three cases. The elasticity for all ages combined is only -.04, far less than that for husband's income as was found in previous studies. The wife's education (E) has invariably nonsignificant effects, thereby contradicting another pattern discerned by Cain. The effect of the proportion rural (R) is close to zero and insignificant. In the case of C, the proxy for the proportion Catholic, the estimates are quite strong with the expected negative effect. A more direct, noncensus measure of the proportion Catholic (see C' in Appendix A) produced similar coefficients and significance levels in this and all other regressions where the substitution was made.[26]

The strong negative impact of the male unemployment rate (U) provides continued support for a dominant discouraged-worker effect. An alternative measure of transitory labor market conditions (see EM in Appendix A) produced similar results. The effect of recent percentage <u>changes</u> in employment levels was significantly positive, thereby strengthening the case for a procyclical labor response.

Our model was also estimated using the LFPR (L_{69}) and average weeks worked by all wives (L_w) in 1969 as dependent variables. These results are presented in Tables B-1 and B-2 in Appendix B and are quite similar to those in Table 6, especially with regard to wage and income elasticities. For all ages combined, the wage elasticities are .67 (L_{70}), .52 (L_{69}) and .66 (L_w). For husbands' income, the elasticities are -.49 (L_{70}), -.37 (L_{69}) and -.44(L_w). This similarity is encouraging in that they provide support for the use of estimates obtained from LFPR equations to make inferences concerning the determination of continuous measures of labor supply, e.g., weeks or hours of work. Such use of the LFPR coefficient estimates has justly been questioned in the past by Ben-Porath (1973) and others.

The foregoing results may also be compared to those obtained by Cain and Dooley (1976) from a simultaneous model of labor supply, fertility and wages. The wage and husband's income elasticities from the latter study were at least twice the absolute size of our present estimates, reaching incredibly large magnitudes in the younger age groups.[27] The same relationship was found for the male unemployment rate; however, Cain and Dooley did find effects for nonlabor income and the wife's education more consistent with prior expectations.

Table 7 contains the labor supply estimates for the black sample. The wage and husband's income effects are significant only for the older age groups and for the sample as a whole.[28] The wage and husband's income elasticities for all ages combined are .20 and -.28 respectively, approximately one half the absolute size of the estimates with 1960 data. However,

Table 6

Estimates of the Labor Supply Function with Published SMSA Data for White Wives, 1970.

(Labor Supply = LFPR's in the Survey Week, L_{70})

Age Group	Constant	W ($00's)	H ($00's)	N ($00's)	E	R	C	U
30-34	44.6*	.30*	-.17*	-.13	.03	.06°	-.42*	-1.12*
35-39	42.2*	.26*	-.11*	-.17	.56	.05	-.22*	-1.09*
40-44	42.0*	.44*	-.20*	-.18°	.72	-.01	-.19*	-1.50*
45-49	51.3*	.37*	-.09°	-.16	-.66	.00	-.01	-1.08*
50-54	33.2*	.66*	-.21*	-.30+	.51	-.02	.04	-1.49*
30-54	39.7*	.57*	-.20*	-.22°	.30	.02	-.09°	-1.78*

* - indicates t-ratio greater than 1.96.

\+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 124 SMSA's.

Table 7

Estimates of the Labor Supply Function with Published SMSA Data for Black Wives, 1970.

(Labor Supply = LFPR in the Survey Week, L_{70})

Age Group	Constant	W ($00's)	H ($00's)	N ($00's)	E	R	U
30-34	57.3+	-.07	-.27	-.21	1.83	.19	-.26
35-39	46.2*	-.03	-.06	-.81	1.81	.27+	-.61°
40-44	36.3	-.02	-.22°	-1.64	3.59*	.27+	-.12
45-49	52.8*	.50*	-.50*	-1.19	1.72+	.05	-.60°
50-54	57.5*	.74*	-.61*	-3.30*	1.09	.02	-.17
30-54	46.5*	.29°	-.45*	-1.44	2.68°	.17	-.21

* - indicates t-ratio greater than 1.96.

\+ - indicates t-ratio greater than 1.645 (but less than 1.96).

° - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 79 SMSA's.

as with the white sample, Cain's observation of a secular decline in the ratio of wage to income elasticities is not substantiated. One racial differential noted by Cain is still apparent, i.e., a lower ratio of wage to income elasticities for the black population. This finding provided an explanation for the greater growth over time in the LFPRs of white married women.

Previous studies with SMSA data did not estimate age-segregated regressions and, hence, provide no precedent for the instability of wage and income coefficients across cohorts in Table 7. Cain has suggested changes in occupational composition as a possible explanation. The proportion of black women in domestic service has declined from 60% in 1940 to 16% in 1970, with a particularly rapid decrease among the young.[29] Such younger females with access to more satisfying occupations, may exhibit behavior less sensitive to purely monetary factors. This explanation is also consistent with the decrease in wage and income elasticities over time; however, it probably cannot account totally for the sharp differences in Table 7.

The coefficients for the remaining variables are only sporadically significant. Nonlabor income (N) and the proportion rural (R) were not included in previous studies for nonwhites. The weak results for education (E) and the unemployment rate (U) confirm prior findings and, in the case of the latter variable, provide continued evidence of a major racial difference in behavior. Somewhat surprisingly, our alternative measure of labor market conditions (EM) produced significantly positive coefficients for each age group thereby furnishing the only evidence of dominant discouraged-worker effect for blacks.

The use of two alternative labor supply measures, the LFPR and average weeks worked in 1969, provided estimates which are reassuringly similar to those for the survey week LFPR (see Tables B-3 and B-4 in Appendix B). The alternative variables did, however, manifest somewhat larger (in absolute size) wage and income elasticities, and higher significance levels for most coefficients.

A comparison of the black labor supply estimates for our model with those from the Cain-Dooley specification reveals a pattern similar to that found for the whites. Specifically, the simultaneous model provided wage and income elasticities larger than can be thought credible for most age groups, yet resulted in nonlabor income and educational effects more consistent with previous findings.

One final set of labor supply functions were estimated with the published data in order to evaluate the stability of previous results. Following Cain (1966) and others, equation (3) was estimated with our fertility measure (F) as an independent variable. (These results are not shown.) Number of children ever born had virtually no direct labor supply impact, unlike the findings in Table 1, and only minimally altered the remaining coefficients. This confirms the findings of Cain and Dooley but contrasts with Fields who found a continued negative effect of percentage of families with children less than age six.[30] This implies a persistent timing effect of young children upon labor supply, but a diminished long-run effect of family size per se. As with data from previous censuses, the 1970 black sample exhibited insignificant effects of fertility upon labor supply.

The general accord of 1970 estimates with those obtained from prior censuses is not sustained in the case of the completed fertility equation for the white population, as shown in Table 8. The estimated coefficients for both wages and husband's income are generally nonsignificant contrary to the findings of Cain and Weininger (1973). This result is particularly puzzling given that all 1960 age groups, save the oldest, are included in the 1970 sample. It is difficult to attribute such divergent estimates entirely to sampling error. Additional explanations may include the influence of timing effects for the younger cohorts in the 1960 sample or shifts in the population composition across SMSAs.

Cain and Weininger did not include a measure of nonlabor income in their regressions. Our results for N are highly significant, but of opposite sign to that commonly predicted by economic models. Of course, the exogeneity of this variable is highly suspect and, therefore, one possible explanation is a correlation between preferences for market goods and large asset holdings. The exclusion of N has little effect on the other coefficient estimates.

The effect of the wife's education is generally nonsignificant, in contrast to the 1960 data; however, the Cain-Weininger variable measured the incidence of functional illiteracy rather than average years of schooling completed. Both the proportion rural and Catholic of the population are ususally significant and of the expected sign. Furthermore, our noncensus measure of Catholic affiliation (C') produced significant coefficients for all age groups. Hence, it is the noneconomic determinants alone that appear to produce empirical results which confirm prior

Table 8

Estimates of the Fertility Equation with Published SMSA Data for White Wives, 1970.

(Fertility = Number of Children Ever Born per 1,000 Women Ever Married, F)

Age Group	Constant	W ($00's)	H ($00's)	N ($00's)	E	R	C
30-34	2698*	-4.71	-1.01	-25.8*	30.3	4.71*	5.74*
35-39	4927*	-2.06	-1.68	-24.5*	-155.3+	5.58+	5.73o
40-44	3909*	6.62o	- .34	-27.8*	- 97.0o	5.75*	7.51*
45-49	2658*	-1.13	.66	-27.5*	10.2	4.90+	1.86
50-54	2183*	-5.73	1.34	-28.2*	44.6	5.64*	-.72
30-54	2849*	.21	-.36	-29.5*	3.4	5.17*	3.65

* - indicates t-ratio greater than 1.96.

\+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 124 SMSAs.

expectations in the case of the white wives. Such was also the case with the white fertility estimates obtained by Cain and Dooley (1970).

The fertility equation estimates for black wives are presented in Table 9. The wage and income coefficients (for both H and N) are with few exceptions significant and consistent with prior expectations. Unfortunately, the exceptions pertain to income effects for the oldest age groups wherein the possibly confounding effect of timing differences are nonexistent. A similar pattern of declining income effects across age groups was found by Cain and Weininger. The husband's income elasticities for the three youngest age groups average .29 compared to a 1960 value of .25. The wage elasticity for all ages combined is -.42, quite similar to the estimate of -.29 with the 1960 data. The estimated impacts of both education and rural residence for blacks are also consistently significant and of the predicted sign, the former result contrasting with those for the white sample. Cain and Dooley found considerably larger wage and income elasticities, in addition to frequently nonsignificant coefficients for the remaining variables.

In summary, a majority of the most salient conclusions from previous studies with SMSA data have been confirmed with the 1970 published averages. In the case of labor supply, the wife's wage and husband's income effects are significant and provide elasticity estimates roughly comparable to prior findings. Continuing evidence of a procyclical labor supply response was produced by several measures of labor market conditions. However, the only additional variables with consistently significant estimates are the proportion Catholic of the population for whites and formal educational levels for the blacks. This pattern of results was obtained with three different measures of labor supply.

Table 9

Estimates of the Fertility Equation with Published SMSA Data for Black Wives, 1970.

(Fertility = Number of Children Ever Born per 1,000 Women Ever Married)

Age Group	Constant	W ($00's)	H ($00's)	N ($00's)	E	R
30-34	8615*	-39.4*	17.5*	164.2*	-422.8*	8.95°
35-39	7593*	-48.1*	21.2*	300.1*	-332.4*	14.5*
40-44	6007*	-40.5*	12.0°	169.0+	-172.2*	15.7*
45-49	5204*	-23.3*	.15	258.4*	-174.1*	25.0*
50-54	4015*	-22.1*	1.07	49.3	- 77.5°	24.0*
30-54	5911*	-37.5*	14.4*	167.9*	-213.9*	16.8*

* - indicates t-ratio greater than 1.96.

+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 79 SMSAs.

A major departure from previous findings was signalled by the nonsignificant effects of the wife's wage and husband's income on completed fertility for the white sample. However, the black fertility results did confirm prior estimates and expectations, manifesting significant coefficients across most all age groups and variables.

C. Estimates with SMSA Averages from the 1970 Public Use Sample

In this section, we report two sets of findings obtained with the PBUS extract: estimates of the bivariate model, (3) and (4), using data for all couples in the extract;[31] and estimates of the labor supply function adjusted for sample selection bias. The latter regressions employ observations for working wives only.

The selection criteria for the extract are discussed in section 3B, and the variable definitions are contained in Appendix A. Table B-5 in Appendix B presents the correlation coefficients between variable measures from the published data and the PBUS extract which, with few exceptions, are quite large.[32] The low coefficient for nonlabor income (N) is probably due to the age dependence of this variable combined with the lack of an age-specific measure from the published data. The lowest coefficients are those for the wife's wage in the white sample, especially among the younger cohorts. This partially results from differences in measurement: The published data provides only a weekly wage for all 1969 workers as opposed to the hourly rate from the PBUS for wives working in both 1969 and the survey week. Still, the degree of correlation is surprisingly weak, especially given the much higher black coefficients, and is reflected in the very different pattern of wage effects reported below.

A final preliminary comment concerns the estimation of the net labor supply effect of transitory variation in labor market conditions using unemployment rates. Some grouping is required since, with household data, the husband's employment status captures principally the added-worker effect as was found by Bowen and Finegan (1969, p. 150). In the PBUS extract, the number of black couples per SMSA does fall as low as 50. Hence, the model was estimated with both the male unemployment rate from the published data and the husbands' unemployment rate from the PBUS extract. The coefficients for the former variable are presented in the following tables. The substitution of the latter variable resulted in coefficients of smaller absolute size, most likely reflecting the stronger added-worker effect, but having the same sign and significance level.

1. Estimates with the full 1970 public use sample extract. Table 10 contains the labor supply estimates for the white sample which contrast in several ways with those obtained from the published data in Table 6. The coefficients for husband's income (H) continue to be significantly negative with one exception. Except for the younger cohorts, however, the elasticities are generally smaller in absolute size resulting in an estimate of -.15 for all ages combined versus a value of -.49 found with the published data.[33] The effects of nonlabor income (N) are invariably nonsignificant in contrast to the negative and frequently significant estimates reported previously.

The coefficient estimates for U (male unemployment rate from the published data) continue to indicate a procyclical labor supply response;

Table 10

Estimates of the Labor Supply Function with Public Use Sample Data for White Wives, 1970 (Labor Supply = LFPR in Survey Week, L_{70})

Age Group	Constant	W	H ($00's)	N ($00's)	E	R	C	U
30-34	41.21*	.19	-.16*	.24	1.75	-.09	-.51*	-1.50*
35-39	43.51*	-.43	-.16*	-.04	2.32°	-.11°	-.35*	-1.21*
40-44	43.79*	-.46	-.04°	.05	1.11	.06	-.22*	-.42
45-49	10.96	2.80*	-.12*	.08	3.94*	.11	.10°	-1.61*
50-54	34.19*	.50	-.04	.08	1.54	.18+	-.03	-.69
30-54	28.94+	-.72	0.06°	.14	2.34°	.07	-.16*	-.93*

*indicates t-ratio greater than 1.96.

+indicates t-ratio greater than 1.645 (but less than 1.96).

°indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 134 SMSA's

however, the low t-ratios for two age groups are without precedent. In addition, the alternative measure of labor market conditions (EM) generally had nonsignificant coefficients thereby not reinforcing the evidence for a dominant discouraged worker effect.

The labor supply impacts of the proportions rural (R) and Catholic (C) are similar to those obtained with the published data. The estimates for the wife's education (E), though still weak, are more supportive of pre-1970 findings than was the case in Table 6.

The most striking contrast with prior results is provided by the nonsignificant, and often negative, coefficients for the wife's wage, W. (Note that W is now measured in dollars per hour rather than hundreds of dollars per year.) Even the estimate for the 45-49 age group loses significance when real wages are employed (see p. 14). The PBUS wage was constructed from the division of annual earnings by annual hours of work. Hence, errors in the measurement of labor supply may produce a spuriously negative wage coefficient, although this is difficult to prove unambiguously in multiple regression. Two factors offset this argument: first, the dependent variable is an LFPR rather than average hours and second, the weekly wage measure from the published data was constructed in a similar manner, yet exhibited strongly positive coefficients.[34] These differences in estimates were foreshadowed by the low correlation coefficients for the wife's wage in Table B-5. The published data has the advantage of large intra-SMSA sample size. On the other hand, the PBUS measures are hourly rather than weekly wages and hence do not reflect variation in labor supply. Both wage measures are subject to unknown degrees of sample selection bias. Exploratory regressions with several additional wage measures are discussed later in this section.[35]

Table B-6 in Appendix B contains estimates of the labor supply function using the normit transformation of the survey week LFPR as the dependent variable. Little comment is necessary since the signs and significance levels of these coefficients are identical to those in Table 10 except, of course, for the constant. The husband's income elasticity[36] for all ages combined was -.13 as compared to a value of -.15 obtained with the LFPR as the dependent variable. This similarity provides support for the use of a linear specification in this and previous analyses of labor force participation rates. As with the published data, the use of the LFPR and average weeks worked in 1969 as labor supply measures also results in only marginally different estimates, particularly with respect to wage and income variables (see Tables B-7 and B-8 in Appendix B).

Table 11 contains the labor supply estimates for the black sample. The divergence between these results and those for the published data in Table 7 is even greater than in the case of the white sample. None of the wage or income coefficients are both significant and of the expected sign. This is particularly puzzling given the relatively large correlation coefficients between measures of these variables from the published and PBUS data. The two data sets do differ substantially in the number of SMSAs, 79 in the published data versus 45 in the PBUS. However, labor

Table 11

Estimates of the Labor Supply Function with 1970 Public Use Sample for Black Wives (Labor Supply = LFPR in Survey Week, L_{70})

Age Group	Constant	W	H ($00's)	N ($00's)	E	R	U
30-39	.67	-4.75+	-.07	-.77	6.84*	.38+	-.74o
40-54	61.69*	.34	-.17	-.02	1.63	-.09	-1.33
30-54	24.78	-3.73	-.14	-.15	5.08+	.22	-.74

* - indicates t-ratio greater than 1.96.

\+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 45 SMSA's.

supply estimates, obtained with the published data for those 43 SMSAs common to both samples, closely approximated those for all 79 areas.[37]

The remaining coefficients are similar to those found with the published data: a positive and usually significant effect of the wife's education and weak results for both the proportion of rural residents and unemployment. However, the use of recent percentage changes in employment (EM) provided strong evidence of a net discouraged worker effect, just as with the black published data.

Table B-9 in Appendix B contains the black labor supply estimates using as dependent variables the normit transformation of the survey week LFPR, and the LFPR and average weeks worked in 1969. The pattern of coefficient estimates closely resembles the generally weak results in Table 11 with the exception of a significantly negative effect of unemployment upon average weeks worked.

The foregoing estimates for both races fail to support one of the major conclusions of all analyses performed with published census data, i.e., the existence of a strong, positive own wage effect. In the white sample, husband's income and various measures of labor market conditions provided continued, though weakened, support for previous findings, but such was generally not the case for black couples.

Table 12 contains the estimates for the white fertility equation which are as weak as those obtained with published data in Table 8. The wage coefficients are negative but significant in only one case. Both husband's and nonlabor income show significantly negative effects for most of the age groups. Of these two anomalous results, only the latter

Table 12

Estimates of the Fertility Equation with Public Use Sample Data for White Wives, 1970 (Fertility = Number of Children Ever Born per 1,000 Wives)

Age Group	Constant	W	H ($00's)	N ($00's)	E	R	C
30-34	4393*	-.90	-2.45	-10.23	-123.2+	2.41	4.21°
35-39	2250*	-10.04	-2.86°	-24.70*	91.41°	2.81	7.13*
40-44	1944*	-30.43	-3.14*	-12.21°	118.2°	2.67	5.45+
45-49	1534*	-39.84	-3.31*	-14.98*	140.8*	4.87°	2.72
50-54	1695*	-148.0*	.53	-6.25	81.72+	9.30*	4.25*
30-54	984.5°	-33.89	-2.66°	-43.18*	193.4*	4.18	5.59*

*indicates t-ratio greater than 1.96.

+indicates t-ratio greater than 1.645 (but less than 1.96).

°indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 134 SMSA's

was true of the published data. However, the use of the husband's real income reduces these estimates to statistical insignificance.

Most unusual are the significantly positive effects of the wife's education on fertility. In the author's knowledge, the only precedent for this finding concerns Catholics' education in parochial schools (Westoff and Potvin 1966). The coefficients for the proportion rural (R) and Catholic (C) generally conform to both predicted effects and those estimated with the published data.

Table 13 contains the estimates for the fertility equation with the black sample. The coefficients for the wife's wage are significantly negative and provide elasticity estimates similar to those previously obtained, specifically -.37 for all age groups combined versus -.46 with the published data. The slope estimates for both income measures, however, usually fail to reach statistical significance, although the husband's real income does provide coefficients similar to those obtained with published data. Of the two remaining variables, E and R, only the proportion rural confirms the results in the previous section.

In summary, the fertility estimates with PBUS data are less supportive of previous SMSA studies than were the results from the 1970 published data. For the white sample, neither set of estimates was particularly encouraging. However, the PBUS data produced significant coefficients for three variables (H, N and E) which were opposite in sign from that expected. The black PBUS sample did not produce any such anomalies, but yielded significant coefficients for only two variables (W and R), as opposed to the full set of estimates with the published data.

Table 13

Estimates of the Fertility Equation with 1970 Public Use Sample for Black Wives (Fertility = Number of Children Ever Born per 1,000 Wives)

Age Group	Constant	W	H ($00's)	N ($00's)	E	R
30-39	5450*	-296.4+	-4.70	-1.66	-103.5	15.39°
40-54	3910*	-382.2*	5.10	-53.44	-41.58	31.69*
30-54	4497*	-473.1*	6.85	-96.32°	-58.42	24.18*

*indicates t-ratio greater than 1.96.

+indicates t-ratio greater than 1.645 (but less than 1.96).

°indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 45 SMSAs

The most striking result reported thus far for the PBUS data is the absence of significantly positive effects of the wife's age on labor supply. This relationship has been one of the principal foci of previous studies. In order to evaluate this finding further, our model was reestimated using two alternative wage measures.

The first was a weekly wage measure from the PBUS similar to that used with the 1970 published data; that is, average earnings of 1969 participants divided by the average weeks worked among 1969 participants, all multiplied by 52. (These results are not shown.) These variables clearly reflect differences in earnings capacity <u>and</u> labor supply, but its use may indicate any estimation biases inherent in such a measure. The second variable was obtained from the occupational wage surveys of the Bureau of Labor Statistics in selected SMSAs. These data provide a very limited sample, due to limitations by occupation, industry, firm size and work schedule, but possess two distinct advantages over census data. First, they are reported by firms and may be less prone to measurement error. Second, they are available for several experience levels within well-defined occupations which are commonly held by married women. Therefore, they may provide better measures of <u>quality-constant</u> wage differences across SMSAs. (Appendix A contains a detailed description of the measures used.)

Neither measure had any noticeable impact on the estimates for the black sample, in particular the wage coefficient. Equally minimal alterations were obtained for the white labor supply function with the exception of a significantly positive effect of education across all age groups when the weekly wage was used. The impact on white fertility of each wage measure was negative and highly significant which contrasts with, rather than replicates, the weak results obtained with the published data. Most importantly however, both wage measures failed to illuminate the reasons for the divergent labor supply estimates provided by the 1970 published and PBUS data.[38]

2. Estimates of the labor supply function adjusted for sample selection bias. In this final section, we report the labor supply estimates obtained with PBUS data for those couples where the wife worked in both 1969 and the survey week. This permits use of the techniques developed by Heckman (1975) and Hanoch (1976) to correct for sample selection bias. Moreover, we shall be able to estimate a function for annual hours of work, a heretofore impossible task with SMSA data. The inclusion of only couples with working wives led to a relatively small number of families in certain SMSAs, especially in the black sample. Hence, only the data for whites was employed. Moreover, only a single-equation model was estimated due to the complexity of the econometric analysis and the minimal changes induced by the use of multi-equation estimation methods for our complete model.

The function estimated was presented in general form in equations (17) and (18), and is specified below with minor notational changes. Unless indicated otherwise, all variables refer to the SMSA average from the PBUS extract for couples where the wife worked in both 1969 and the survey week.

$$L = \psi_0 + \psi_1 W + \psi_2 H + \psi_3 N + \psi_4 E + \psi_5 A + \psi_6 P + \psi_7 U + \psi_8 \lambda + u \qquad (19)$$

where

L = annual hours (or weeks) of work.

W, H, N, E, A, P, U = independent variables previously used and defined in Appendix A.

$\lambda = \frac{-f(\phi)}{F(\phi)}$

$F(\phi)$ = proportion of wives in the population who worked in 1969, as consistently estimated by the 1969 sample LFPR.

ϕ = linear combination of determinants of annual labor force participation.

f(.), F(.) = the standard normal and cumulative standard normal density functions.

$E(u) = 0$

$$Var(u) = \frac{\sigma_{11}}{\alpha_1^2 n} \{\delta^2(1 - \phi\lambda - \lambda^2) + 1 - \delta^2\}$$

$$\delta = (\sigma_{22} + \sigma_{12}) \ / \ \{\sigma^*(\sigma_{11})^{\frac{1}{2}}\}$$

$$\sigma^* = (\sigma_{22} + \sigma_{11} - 2\sigma_{12})^{\frac{1}{2}}$$

$\sigma_{11}, \sigma_{22}, \sigma_{12}$ = elements of the covariance matrix of the disturbances from the shadow wage and market wage functions of the Heckman model (see section 3C).

n = the number of couples with working wives in each SMSA.

ψ_0, ψ_1, ψ_2 thru ψ_7 and ψ_8 in (19) are equivalent to: α_0/α_1, $1/\alpha_1$, $\underline{\alpha}_2/\alpha_1$ and δ/α_1 respectively in (17). These notational differences between (17) and (19) result from the fact that no attempt will be made to identify all the parameters of the Heckman model. Our interest is only in the labor supply effects of the independent variables.

There is no unambiguous theoretical prediction for the sign of ψ_8 above. α_1 in (17) is the uncompensated wage effect which has been found generally to be positive in the case of married women. The sign of δ depends on σ_{12} since both σ_{11} and σ^* are non-negative. σ_{12} is the covariance between the disturbances in the market and shadow wage equations and will depend most importantly upon the degree of similarity between the skills required for home versus market production. Conceivably δ, and hence α_8, could be zero if $\sigma_{11} = \sigma_{12}$, which would imply the absence of sample selection bias if (19) were to be estimated excluding λ.[39]

The unknown constant term in the disturbance variance, $(1-\delta^2)$, complicates the use of traditional weighting procedures to correct for heteroscedasticity. Hanoch (1976, p. 20) reports that his method for dealing with the same problem is quite cumbersome and quantitatively of minor significance. For convenience, it was decided to ignore this term. The appropriate weight for each observation, then would be $\{n/(1 - \hat{\phi}\hat{\lambda}-\hat{\lambda}^2)\}^{\frac{1}{2}}$ where $\hat{\phi}$ and $\hat{\lambda}$ are estimates of ϕ and λ derived from the 1969 sample LFPR.

Recall, however, that the subsample of working wives was self-selected on the basis of labor supply in both 1969 and the survey week. The Heckman estimation method assumes a single selection criterion. As indicated on p. 21, Hanoch (1976) has derived a straightforward weighting scheme for resolving this problem. Utilization of this technique with our data calls for the following weight for each observation:

$$\{nL_w'/(1 - \hat{\phi}\hat{\lambda} - \hat{\lambda}^2)L_w\}^{\frac{1}{2}}$$

where L_w' = average weeks worked in 1969 by all 1969 participants.

L_w = average weeks worked in 1969 by all wives who participated in <u>both</u> 1969 and the survey week.

Table 14 contains the estimates for (19) with annual hours of work as the dependent variable. Not only are most of the wage coefficients unexpectedly negative, but significantly so in four of the six age groups. This raises the previously mentioned possibility of bias due to the presence of annual hours in the denominator of the wage variable. However, such bias would have to be quite strong to account for the above results, particularly in the case of the 45-49 age group which had produced the only significantly positive wage effect in previous regressions with PBUS data.

The coefficients for husband's income (H) are still negative and significant in three instances. The elasticity for all ages combined is -.19 which closely approximates the estimates obtained with several labor supply measures for the full PBUS sample.

Nonlabor income (N) continues to exhibit generally nonsignificant effects and the impact of education (E) is weaker than in regressions with the entire PBUS sample. The proportion rural (R) is significant and carries the expected negative sign, a result not previously obtained with the 1970 data. The proportion Catholic (C) and the unemployment rate (U) exhibit estimates congruent with both prior expectations and findings. The support for a dominant discouraged-worker effect is, however, less consistent than has been the case with the published SMSA data.[40]

The majority of the coefficient estimates for λ are not significantly different from zero. Moreover when (19) was estimated excluding λ, the remaining coefficients (and their standard errors) changed very little. Both of these results provide limited evidence for a low level of sample selection bias.

Table 14

Estimates of the Labor Supply Function with 1970 Public Use Sample for White Wives and Correction for Sample Selection Bias (Labor Supply = Annual Hours of Work by Women Working in 1969 and the Survey Week, L_h)

Age Group	Constant	W	H ($00's)	N ($00's)	E	R	C	U	λ (percentage points)
30-34	1471*	-41.4*	- .39	-6.57	39.4	- .06	-1.62	- 8.4	3.84*
35-39	2406*	.95	-2.70*	1.51	-37.6^{o}	-5.13*	-5.40*	- 6.67	1.13
40-44	2230*	-20.9^{+}	-2.15*	4.83	-29.0	-2.80*	-5.61*	8.17	-.30
45-49	1675*	-59.1*	.12	4.35^{o}	17.3	-1.50	-3.30*	-18.3*	.31
50-54	1495*	-44.1*	- .86	4.88*	46.3*	-2.61*	-1.66*	-13.6^{+}	1.54*
30-54	1968*	10.4	-3.07*	5.27	2.34	-4.63*	-4.30*	-18.5*	.63

* - indicates t-ratio greater than 1.96.

\+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 134 SMSA's.

Equation (19) was also estimated with average weeks worked (among working women) as the dependent variable. The results are presented in Table B-10 of Appendix B and differ in only two respects from those in Table 14. First, the wage coefficients, while still negative, are not significantly different from zero. Second, evidence of a net discouraged-worker effect is totally absent. The coefficients for all measures of labor market conditions were insignificant and of unstable sign. The results for λ again provide no evidence of severe sample selection bias, i.e., nonsignificant own coefficients and a minimal effect of its (λ's) exclusion upon the other estimates. Moreover, the pattern of coefficients in Table B-10 does not differ greatly from those of the average weeks worked function estimated with all wives, except in the case of unemployment (see Table B-8).

In summary, this last set of estimates has not provided information substantially different from the regressions with the full PBUS sample. Most importantly, no evidence was provided of a positive wage effect. On the contrary, the estimated coefficient was usually negative and significantly so for annual hours of work . The impact of the husband's income continued to be negative with some attenuation, however, in both absolute size and significance levels. There were two contrasts with the results from both the 1970 published data <u>and</u> the entire PBUS sample. First, the impact of rural residence was negative and frequently significant. Second, the support for a dominant discouraged-worker effect was weak and, indeed, nonexistent for annual weeks of work.

A logical extension of this analysis would be to estimate our model, in particular the labor supply function, with individual household data from the 1970 Public Use Sample. A recent study (Morgenstern and Hamovitch 1976) provides such estimates for both annual hours and weeks of work, albeit with a somewhat different specification and sample than were used for this study.[41] These findings generally confirm both theoretical expectations and the work of Cain (1966) and Bowen and Finegan (1969) with the 1960 Public Use Sample. (Neither of these last two studies, unfortunately, directly estimated wage effects.) In particular, the wage and income coefficients are highly significant and translate into elasticity estimates of .35 and -.07 respectively for average weeks worked.[42] These are considerably smaller in absolute size than the values of .66 and -.44 estimated in this study with the 1970 published data. The low elasticity may be due to the combination of the husband's earnings and family nonlabor income into one variable by Morgenstern and Hamovitch.[43] Apparently, then, both individual household observations and highly aggregated data (the published averages) from the 1970 Census provide the anticipated labor supply estimates whereas data representing intermediate levels of aggregation (SMSA averages from the PBUS) do not.

5. SUMMARY AND CONCLUSION

The main objective of this study has been to extend and evaluate various analyses of the labor supply and completed fertility of married women which used SMSA data from the 1940, 1950, and 1960 U.S. Censuses. A bivariate model of labor supply and numbers of children ever born was estimated (by age and race) using two sets of data from the 1970 Census: the published SMSA averages, similar to those used in prior studies and

SMSA averages derived from an extract of the 1970 Public Use Sample. In addition, the normit transformation was used to derive an appropriately bounded labor force participation rate equation and recently developed adjustments for sample selection bias were adopted to a grouped data framework.

In general, coefficient estimates with the 1970 published averages closely approximate those estimated with data from the three preceding censuses. Most importantly, the uncompensated wife's wage and husband's income elasticities in the labor supply function are highly significant and comparable in size to previous estimates, except for the youngest black cohorts. Continued evidence of a net discouraged-worker effect is provided by several measures of transitory labor market conditions among whites and, for the first time, to a limited extent among blacks. The other labor supply determinants with consistently significant effects are the proportion Catholic of the white sample and the educational level of the black wives. The most disappointing results concerned the white fertility equation for which only rural residence and the proportion Catholic confirm prior expectations and findings. The black fertility estimates, on the contrary, are almost invariably significant and result in wage and income elasticities congruent with prior findings.

The findings with the 1970 Public Use Sample averages differ substantially from those cited in the previous paragraph. Most importantly, the effects of the wife's wage upon labor supply are insignificant and of varying sign for both races, thus contradicting one of the most stable and important results of previous studies. All other labor supply coefficients for the black sample, with the exception of the wife's education, are not significant. As with the published data for the whites, husband's income, proportion Catholic, and unemployment exhibit the expected negative labor supply effects, although the absolute size and significance levels of the estimates are somewhat attenuated.

The estimates of the white fertility equation with the Public Use Sample are as weak as those found with the published data, the only noticeable difference being a previously unencountered positive effect of the wife's education. The black fertility estimates for the wife's wage and proportion rural alone were significant and of the expected sign, unlike the full range of coefficients with the published data.

The use of various labor supply measures (including the normit transformation), several alternative wage measures and a cost-of-living adjustment did not yield substantially different estimates, except for the negative white fertility effects of a BLS wage variable.

Finally, no major improvement was obtained by the use of the working portion of the white sample to estimate the labor supply function adjusted for sample selection bias. On the contrary, this effort produced significantly negative wage effects on annual hours of work and the weakest evidence yet found of a net discouraged-worker effect.

The most important results of this study are the dissimilar labor supply estimates obtained with the two 1970 SMSA data sets, particularly for the wife's wage. The Public Use Sample average failed to yield a significantly positive wage elasticity, in contrast to the relatively stable estimates of published data from the 1970 and three preceding censuses, and exhibited generally weaker findings for all labor supply determinants. Unfortunately, there do not exist *a priori* grounds for assessing the relative validity of these estimates. The published averages offer much greater intra-SMSA sample size whereas the Public Use Sample averages permit more accurate variable measurement and superior estimation methods. The estimates obtained with household data from the 1970 Public Use Sample by Morgenstern and Hamovitch (1978), along with those from alternative micro data sources, do support our results from the published averages. As has been stressed, moreover, the restrictions which census data impose upon both model specification and econometric methods raise a host of possible questions concerning any set of estimates from this source. However, the Census, in particular the published SMSA data, offers one of the few sources of information concerning labor market and fertility behavior during the periods covered by the seminal works of Mincer, Cain, and Bowen and Finegan. The major contribution of this study is methodological: inferences and predictions from conventional economic models of labor supply and fertility of wives, using Census data, require attention to the inherent defects of the data, and the results are not robust to alternative specifications.

NOTES

[1]The measures of many variables varied across censuses. In addition, most of the models estimated by Cain and Bowen and Finegan contained one or more additional variables which are excluded from (1) and the model estimated for this study. The justifications for such exclusions will be presented in the next section.

[2]Grouped data (SMSA averages) were used to mitigate the influence of the many family-specific sources of transitory variation in measures of wages and income from a cross section.

[3]See Mincer (1966). An alternative measure of transitory labor market conditions, recent percentage changes in total employment, has been proposed by Fleisher and Rhodes (1976) and was used in this study.

[4]For studies of the relation between assets and labor supply, see Fleisher, Parsons and Porter (1973) and Smith (1975).

[5]Mincer and Polachek do report estimates for a simultaneous model of wages and labor supply; however, the sources of identification raise additional questions. See the comment by Duncan (1974).

[6]See the articles by Hall, Greenberg and Kosters, Garfinkel and others in Cain and Watts (1973) for examples.

[7]In the terminology of Heckman and Willis (1977), there is unobserved heterogeneity in the population.

[8]See T. P. Schultz (1974) for a summary of economic studies of fertility behavior.

[9]Several peripheral variables used by the authors have been excluded from (2) and the model estimated for this study.

[10]Rural households in the Public Use Sample extract are limited to non-farm residences.

[11]This reference to the potential endogeneity of most all human capital investment decisions raises the question of why the wife's education (E) was retained in our model. The basic reason was for comparison with the results of previous studies. Moreover, its (education) exclusion induced minimal change in the remaining coefficient estimates.

[12]We will, however, compare the Cain-Dooley results with those obtained in this study and, in addition, report estimates of a single-equation labor supply model with number of children as an independent variable.

[13]See footnote 3 concerning the appropriate measure for this factor.

[14]Following previous studies, Honolulu was excluded mainly for reasons of its unique racial composition.

[15]Due to the limited number of black couples for some SMSAs in the PBUS sample, separate regressions were estimated for only two age groups, 30-39 and 40-45.

[16]An additional rationale for age segregation is the possibility of variation in the shape, as well as height, of labor supply-age profiles in response to variation in the independent variables. Such a possibility has been shown in Heckman (1976) and greatly complicates the interpretation of cross sectional labor supply studies. Unfortunately, Heckman's analysis

fn 16 continued. . .

provides few general guidelines as to the most appropriate stage in the life cycle for the estimation of responses in the permanent level, as opposed to the timing, of labor supply with cross sectional data. Timing, of course, may also influence family size for the younger wives in our sample.

[17] I am most grateful to John Bishop for providing 1969 cost-of-living estimates for 61 of the SMSAs used in this study. These figures were based on Bureau of Labor Statistics estimates adjusted by Bishop to reflect different tax rates, transporation costs, and quality-constant housing costs.

[18] Specifically, $F^{-1}\{E(p_i)\} \lesseqgtr E\{F^{-1}(p_i)\}$ depending on the value of $\underline{x}_i$. The left hand side of this expression will be greater than (less than) the right hand side along the concave (convex) portion of the conditional probability function.

[19] Cross-tabulations (by wife's age, education and husband's income) for selected SMSAs indicate very few cells with p_i less than 30% or greater than 70%.

[20] A similar rationale was used above to justify the use of observed LFPRs to estimate the SMSA mean of λ.

[21] Although not presented here, the results of Bowen and Finegan (1969) were similar to Cain's despite the inclusion of several additional labor market variables in the former specification.

[22]Cain also estimated a log-linear model. The resulting elasticity estimates were generally similar to those in Table 3.

[23]Specifically, Mincer has suggested that inter-SMSA unemployment variation reflect long run structural differences among labor markets. Supportive evidence is provided by the high correlation among SMSA unemployment rates over time. However, see Appendix A for an alternative measure of transitory labor market conditions (EM) whose effects (see Section 3B) generally support the evidence for a dominant discouraged worker effect.

[24]The predicted sign of the income effect should, of course, be left open, given the longstanding debate concerning "true" versus "observed" income elasticities (see Becker and Lewis 1974, Becker and Tomes 1976).

[25]One surprising exception to this pattern of significant results was obtained with the median earnings of full year (50-52 weeks) white female workers. This wage measure is very similar to that used by Cain and avoids the possible bias in our normal wage measure of earnings divided by labor supply (see W in Appendix A). Unfortunately, this variable was unavailable by age and even the value for white workers had to be calculated in a less than satisfactory fashion, which may account for the anomalous estimates. The use of median earnings for total full year female workers, i.e., the measure identical to that used by Cain, resulted in strongly positive wage effects as in Table 6. See Dooley (1977 p. 121) for further discussion and the specific estimates.

[26]Due to the unconventional nature of including the proportion Catholic in the labor supply function, our model was also estimated with this variable confined to the fertility equation. The only noticeable effects were slightly larger (in absolute size) coefficients for the wife's wage and husband's income, particularly in the younger age categories.

[27]An alternative specification estimated by Cain and Dooley eliminated the reciprocal effects of labor supply and fertility. This produced wage and husband's income elasticities which were more stable across age groups, but still considerably larger than those estimated for this study. See Dooley (1977, pp. 132-136) for further discussion of these comparisons.

[28]As with the white sample, the use of median earnings of full year black female workers for all ages combined produced an insignificant wage coefficient.

[29]In 1970, the proportions of black female workers who were employed as domestic servants were the following: age 18-24, 5%; age 25-34, 9% age 35-54, 20%; and age 55-64, 38%. The sources of information on occupational composition were Cain (1966, p. 87) and the 1970 U.S. Census Subject Report PC(2)-8B, Earnings by Occupation and Education, Tables 8 and 11.

[30]Cain found a negative impact for both fertility measures with the 1960 data.

[31]The only exception is hourly wages which could be calculated only for those wives who worked both in 1969 and the survey week.

[32]Recall that the two data sets contain unequal numbers of SMSAs for both races. The correlations were based on the 124 (43) SMSAs common to the two white (black) samples. Differences in both intra-SMSA sample size and variable definitions account for the imperfect correlation.

[33]The measure of H from the PBUS data is based on the husband's labor income alone, however, to minimize terminological complexity, we shall continue to refer to this variable as husband's income.

[34]Of course, the larger number of observations per SMSA should lessen the degree of measurement error in the published data variables relative to the PBUS ones.

[35]The explanation for insignificant wage effects would appear not to lie in low levels of variation in the PBUS wage measure. The variance of the PBUS wage measure, relative to its mean, was greater than that for the wage measure from the published data. This result was observed for both races and most all age groups.

[36]Note that this elasticity measures the proportional change in the LFPR and not in the dependent variable, i.e., the normit transformation of the LFPR.

[37]Published data were not available for two of the groups of adjacent SMSAs in the black sample.

[38]Similar results were obtained with another wage measure from the published data, the median earnings of full-year female workers.

[39]Note, however, that this condition would not guarantee the absence of sample selection bias from estimates of the labor supply function with

fn 39 continued. . .

the full sample, which were presented in the first section of this chapter. The absence of sample selection in this case would require that the conditonal expectation of market wages among working women equal the conditional expectation among all women. A sufficient condition for this to hold is that $\sigma_{12} = \sigma_{22}$, as is shown by equation (15).

[40]The substitution of our alternative measures for both proportion Catholic (C') and labor market conditions (EM) provided similar results.

[41]Morgenstern and Hamovitch limit their sample to white women age 16-65, once married, spouse present, not currently enrolled in school, who were recorded as being a member of one of the census's three digit occupational categories. The last criterion limited the sample to individuals who held a job in the 1960-1970 decade--70 percent of the sample fulfilling the remaining criteria. Separate labor supply functions were estimated for women in part-time and full-time occupations, as defined by average weekly hours of work (for the occupation) of less than or greater than 32 hours. Such sample splitting based upon a labor supply related decision (occupational choice) raises the possibility of sample selection bias. Although the magnitudes of the estimated coefficients vary somewhat between regressions, the pattern of signs and significance levels is quite similar.

[42]These are weighted averages of estimates from the full-time and part-time regressions and, of course, are not necessarily equal to elasticities which would result from a pooled sample.

[43]Other micro data sources have provided theoretically congruent wage and income elasticity estimates, e.g., see Heckman (1974) and Schultz (1975).

References

Ashenfelter, O. and Heckman, J. 1974. The estimation of income and substitution effects in a model of family labor supply. Econometrica 42, 73-85.

Becker, G. S. 1960. An economic analysis of fertility. In Demographic and economic change in developed countries. Universities - National Bureau Committee for Economic Research. Princeton: Princeton University Press.

Becker, G. S. and Lewis, H. G. 1974. On the interaction between quantity and quality of children. In T.W. Schultz (ed.), The economics of the family: Marriage, children and human capital. Chicago: University of Chicago Press.

Becker, G. S. and Tomes, N. 1976. Child endowments and the quantity and quality of children. Journal of Political Economy, 84, S143-S163.

Ben-Porath, Y. 1973. Labor force participation rates and the supply of labor. Journal of Political Economy, 81, 697-704.

Bowen, W. G. and Finegah, T. A. 1969. The economics of labor force participation. Princeton: Princeton University Press.

Cain, G. G. 1966. Married women in the labor force: An economic analysis. Chicago: University of Chicago Press.

______. 1967. Unemployment and the labor force participation of secondary workers. Industrial and Labor Relations Review, 20, 275-297.

______ and Dooley, M. D. 1976. Estimation of a model of labor supply, fertility, and wages of married women. Journal of Political Economy, 84, S179-S201.

______ and Mincer, J. 1969. Urban poverty and labor force participation: Comment. American Economic Review, 59, 185-194.

______ and Watts, H. W., eds. 1973. Income maintenance and labor supply. New York: Academic Press.

______ and Weininger, A. 1973. Economic determinants of fertility: Results using cross-sectional aggregate data. Demography, 10, 205-233.

Dooley, M. D. 1977. An analysis of the labor supply and fertility of married women with grouped data from the 1970 U.S. census. Unpublished Ph.D. dissertation, University of Wisconsin-Madison.

Duncan, O. D. Comment. 1974. In T.W. Schultz (ed.), The economics of the family: Marriage, children and human capital. Chicago: University of Chicago Press.

Easterlin, R. A. 1977. Relative economic status and the American fertility swing. In E.B. Sheldon (ed.), Social structure, family life styles, and economic behavior. Philadelphia: J.P. Lippincott.

Fields, J. 1976. A comparison of intercity differences in the labor force participation of married women in 1970 with 1940, 1950 and 1960. Journal of Human Resources, 11, 568-577.

Fleisher, B. M., Parsons, D. O., and Porter, R. D. 1973. Asset adjustment and labor supply of older workers. In G.G. Cain and H.W. Watts (eds.), Income maintenance and labor supply. New York: Academic Press.

Fleisher, B. and Rhodes, G. 1976. Unemployment and the labor force participation of married women. Review of Economics and Statistics, 58, 398-406.

Garfinkel, I. 1973. On estimating the labor-supply effects of a negative income tax. In G.G. Cain and H.W. Watts (eds.), Income maintenance and labor supply. New York: Academic Press.

Goldberger, A. S. 1974. Classnotes. Economics, 718.

Grabill, W., Kiser, C. and Whelpton, P. 1958. The fertility of American women. New York: Wiley.

Greenberg, D. H. and Kosters, M. 1973. Income guarantees and the working poor: The effect of income-maintenance programs on the hours of work of male family heads. In G.G. Cain and H.W. Watts (eds.), Income maintenance and labor supply. New York: Academic Press.

Gronau, R. 1974. Wage comparisons -- A selectivity bias. Journal of Political Economy, 82, 1119-1144.

Hall, R. E. 1973. Wages, income, and hours of work in the U.S. labor force. In G.G. Cain and H.W. Watts (eds.), Income maintenance and labor supply. New York: Academic Press.

Hanoch, G. 1980. A multivariate model of labor supply: Methodology for estimation. In J.P. Smith (ed.), Female labor supply. Princeton: Princeton University Press.

Heckman, J. J. 1974. Shadow Prices, market prices, and labor supply. Econometrica, 42, 679-694.

______. 1975. Shadow, market wages and labor supply revisted: Some computational and conceptual simplications and revised estimates. Unpublished manuscript, University of Chicago.

______. 1976. A life-cycle model of earnings, learning, and consumption. Journal of Political Economy, 84, S11-S44.

______. 1980. Sample selection bias as a specification error. In J.P. Smith (ed.), Female labor supply. Princeton: Princeton University Press.

______. 1978. A partial survey of recent research on the labor supply of women. American Economic Review, 68, 200-207.

______ and Willis, R. G. 1975. Estimation of a stochastic model of reproduction: An econometric approach. In N.E. Terleckyj (ed.), Household production and consumption. New York: Columbia University Press.

______. 1977. A beta-logistic model for the analysis of sequential labor force participation by married women. Journal of Political Economy, 85, 27-58.

Johnson, N. and Kotz, S. 1972. Distribution in statistics: Continuous multivariate distributions. New York: Wiley.

Leibenstein, H. 1957. Economic backwardness and economic growth. New York: Wiley.

Mincer, J. 1962. Labor force participation of married women. In H.G. Lewis (ed.), Aspects of labor economics. Universities-National Bureau Conf. Series no. 14. Princeton: Princeton University Press

______. 1966. Labor force participation and unemployment: A review of recent evidence. In R.A. Gordon and M. S. Gordon (eds.), Prosperity and unemployment. New York: Wiley.

______ and Polachek, S. 1974. Family investments in human capital: Earnings of women. In T.W. Schultz (ed.), Economics of the family: Marriage, children and human capital. Chicago: University of Chicago Press.

Morgenstern, R. D. and Hamovitch, W. 1976. Labor supply of married women in part-time and full-time occupations. Industrial and Labor Relations Review, 30, 59-67.

Rindfuss, R. R. and Sweet, J. A. 1977. Postwar fertility trends and differentials in the United States. New York: Academic Press.

Schultz, T. P. 1980. Fertility determinants: A theory, evidence, and an application to policy evaluation. In J.P. Smith (ed.), Female labor supply. Princeton: Princeton University Press.

Smith, J. P. 1980. Assets and labor supply. In J.P. Smith (ed.), Female labor supply. Princeton: Princeton University Press.

Smith, J. P. ed. 1980. Female labor supply. Princeton: Princeton University Press.

Sweet, J. A. 1974. The employment of rural farm wives: 1970. Center for Demography and Ecology Working Paper 74-22. University of Wisconsin-Madison.

Turchi, B. 1975. The demand for children: The economics of fertility in the U.S. Cambridge: Ballinger Publishing Co.

Westoff, C. and Potvin, R. 1966. Higher education, religion and women's family size orientations. American Sociological Review, 31, 489-496.

Willis, R. J. 1974. Economic theory of fertility behavior. In T.W. Schultz (ed.), Economics of the family: Marriage, children and human capital. Chicago: University of Chicago Press.

Appendix A

Variable Definitions, Descriptive Statistics and Data Sources

Published Census Data

These variables were obtained for non-Spanish whites and blacks and the following age groups unless stated otherwise: 30-34, 35-39, 40-44, 45-49, 50-54, and 30-54.

F: number of children ever born per 1,000 women ever married as of census week.

L_{70}: proportion of married females, husband present, who were in the labor force during the census week. The LFPR for those age 35-44 was used for age groups 35-39 and 40-44.

L_{69}: proportion of females who performed some market work in 1969.

L_w: average weeks worked by all females in 1969.

H: mean income of males with income in 1969.

W: mean income of females with income in 1969 divided by average number of weeks worked of those who worked in 1969 and multiplied by 52.

U: the proportion unemployed of males in the civilian labor force during the census week.

E: female median years of school completed.

N: mean "other income" of families in 1969, available by race but not age of head. Other income is a residual category which excludes wage and salary, nonfarm self-employment, farm self-employment, social security, and public assistance.

R: proportion of the population living in rural areas available by race but not age.

C: proportion of the elementary and high school population enrolled in parochial schools as of the census week, available only by SMSA.

Public Use Sample Data

These variables were obtained for the same age groups as the published data for the white non-Spanish population. For the black population, the age groups were: 30-39, 40-54, and 30-54.

Five Per Cent Sample

F: number of children ever born to wives.

L_{70}: proportion of wives in the labor force during the survey week.

NL_{70}: $F^{-1}(L_{70})$ where F(.) is the cumulative standard normal density function.

L_{69}: proportion of wives who performed some market work in 1969.

L_w: mean weeks worked by all wives in 1969.

L'_w: mean weeks worked in 1969 by wives who worked in 1969.

L_h: mean annual hours of work by wives who worked for money both in 1969 and in the pre-survey week.

W: mean hourly wages of wives for same group as L_h.

H: mean earnings of husbands in 1969.

N: mean "other income" of the couple in 1969. See definition of N for published data.

E: mean years of formal education for wives.

U: the proportion of husbands unemployed during the survey week.

R: the proportion of (non-farm) couples residing in rural areas.

Fifteen Per Cent Sample

C: proportion of white children enrolled in grades 1-12 who were enrolled in a parochial school as of the survey week. An alternative measure - the proportion of white couples, with children in grades 1-12, who had at least one child enrolled in a parochial school - provided coefficient estimates virtually identical to the above measure.

Non-Census Data

EM: percentage change in the total number of employees in the SMSA between mid-March 1968 and mid-March 1970.

C': proportion Roman Catholic of the white non-Spanish population. See Table A-1 for source of the total number of Catholics. It was assumed that all Spanish-Americans and no blacks are counted as Catholic.

BLS Wage Measures: Two measures were derived from the 1970 Wage Surveys. The first (W_{BLS1}), available for 77 SMSAs, is the overall mean of the mean wages of females in the following occupations: File Clerk, Class C; Office Girl; Typist, Class B; and Janitor. The second, (W_{BLS2}), available for 44 SMSA's, is the overall mean of the mean wages of inexperienced female typists and inexperienced female clerical workers.

Table A-1

Descriptive Statistics and Sources of Variables Used in Regression Analyses

Published Census Data

Variable and Age Group	White Mean	White SD	Black Mean	Black SD	Source Followed by Table Number
F:					
30-54	2,745	226	3,387	375	
30-34	2,679	218	3,347	348	
35-39	3,014	259	3,776	414	a 161
40-44	2,972	252	3,621	482	
45-49	2,682	230	3,227	443	
50-54	2,426	235	2,847	464	
L_{70}:					
30-54	44.3	4.6	59.3	6.3	
30-34	36.4	5.0	60.5	6.8	
35-44	45.1	4.4	61.0	6.5	a165
45-49	48.2	4.9	57.4	5.8	
50-54	46.2	5.5	55.2	7.6	
L_{69}:					
30-54	55.7	4.3	66.5	6.1	
30-34	50.5	4.9	68.5	7.0	
35-39	54.0	4.7	68.0	6.8	a167
40-44	57.5	4.4	67.0	6.9	
45-49	57.8	5.0	64.3	6.0	
50-54	57.8	5.0	64.3	6.0	
L_W:					
30-54	22.3	2.0	27.4	3.0	
30-34	18.4	2.2	27.2	3.5	
35-39	20.8	2.1	28.0	3.4	a167
40-44	23.3	2.1	27.8	3.3	
45-49	24.4	2.3	27.0	2.9	
50-54	24.4	2.3	27.0	2.9	
W:					
30-54	5,102	535	4,131	767	
30-34	4,886	501	4,422	801	
35-39	4,923	518	4,339	766	
40-44	5,086	572	4,189	827	a193, 167
45-49	5,227	607	3,953	789	
50-54	5,371	649	3,634	758	

Table A-1 (cont.)

Variable and Age Group	White Mean	White SD	Black Mean	Black SD	Source Followed by Table Number
H: 30-54	10,915	1,166	6,017	1,061	
30-34	9,772	957	5,997	972	
35-39	10,885	1,162	6,240	1,067	
40-44	11,382	1,239	6,156	1,087	a193
45-49	11,491	1,321	5,959	1,158	
50-54	11,022	1,316	5,654	1,129	
E: 30-54	12.5	0.39	10.9	0.77	
30-34	12.5	0.38	11.8	0.49	
35-39	12.6	0.35	11.4	0.64	
40-44	12.5	0.40	10.9	0.86	a148
45-49	12.5	0.41	10.3	1.07	
50-54	12.2	0.66	9.5	1.09	
U: 30-54	2.2	1.0	4.4	1.9	
30-34	2.2	1.0	4.9	2.7	
35-39	2.0	1.0	4.5	2.2	
40-44	2.1	1.0	4.4	2.1	a164, 193
45-49	2.2	1.1	4.0	1.7	
50-54	2.3	1.2	4.1	2.4	
N	877	372	196	61	a205
R	163	11.1	13.3	9.2	b14
C	9.3	6.3	--	--	a83
Public Use Sample Data					
F: 30-54	2.75	.22	3.36	.49	5% Sample
30-34	2.68	.24			
35-39	3.01	.31	3.40	.49	
40-44	2.94	.30			
45-49	2.69	.28			
50-54	2.42	.28	3.31	.62	
L_{70}: 30-54	46.1	5.3	62.5	7.9	5% Sample
30-34	37.5	8.1			
35-39	44.4	7.7	62.3	8.5	
40-44	48.6	6.5			
45-49	50.2	8.2			
50-54	50.0	7.9	63.0	9.0	

Table A-1 (cont.)

Variable and Age Group	White Mean	White SD	Black Mean	Black SD	Source Followed by Table Number
NL_{70}:					
30-54	-9.85	13.6	32.3	21.3	5% Sample
30-34	-32.6	21.8			
35-39	-14.3	20.1	31.2	23.0	
40-44	-3.7	16.6			
45-49	.40	21.1			
50-54	-.05	20.3	34.1	24.0	
L_{69}:					
30-54	52.2	5.7	67.7	7.9	5% Sample
30-34	45.5	8.4			
35-39	50.9	7.8	68.7	8.6	
40-44	54.7	7.2			
45-49	55.4	8.2			
50-54	54.0	8.6	66.9	8.9	
L_W:					
30-54	20.7	2.6	28.6	3.8	5% Sample
30-34	16.3	3.8			
35-39	19.3	3.6	28.2	4.4	
40-44	21.8	3.1			
45-49	23.1	3.8			
50-54	23.1	3.9	28.3	4.0	
L'_W:					
30-54	43.6	1.1			5% Sample
30-34	40.9	3.0			
35-39	42.0	2.6			
40-44	43.5	2.2			
45-49	45.1	1.8			
50-54	45.5	2.2			
L_h:					
30-54	1547	96			5% Sample
30-34	1402	193			
35-39	1452	153			
40-44	1547	162			
45-49	1630	129			
50-54	1657	124			

Table A-1 (cont.)

Variable and Age Group	White Mean	White SD	Black Mean	Black SD	Source Followed by Table Number
W:					
30-54	3.01	.50	2.63	.51	5% Sample
30-34	3.08	.99			
35-39	3.03	1.13	2.73	.61	
40-44	3.14	1.47			
45-49	2.89	.45			
50-54	2.93	.52	2.55	.60	
H:					
30-54	11,437	1,328	6,596	1,160	5% Sample
30-34	10,909	1,344			
35-39	11,659	1,605	6,744	1,088	
40-44	11,905	1,675			
45-49	11,596	1,565			
50-54	11,089	1,783	6,476	1,284	
N:					
30-54	592	245	112	73	5% Sample
30-34	308	227			
35-39	436	273	97	82	
40-44	599	395			
45-49	732	441			
50-54	937	572	126	107	
E:					
30-54	11.9	.44	10.8	.60	5% Sample
30-34	12.2	.42			
35-39	12.1	.46	11.4	.48	
40-44	11.9	.46			
45-49	11.7	.54			
50-54	11.5	.60	10.3	.79	
U:					
30-54	1.82	1.16	2.56	1.74	5% Sample
30-34	1.57	1.55			
35-39	1.80	1.79	2.39	2.15	
40-44	1.80	2.01			
45-49	1.84	1.83			
50-54	2.09	1.65	2.66	2.45	

Table A-1 (cont.)

Variable and Age Group	White Mean	White SD	Black Mean	Black SD	Source Followed by Table Number
R:					
30-54	14.3	6.9	14.7	7.7	5% Sample
30-34	15.6	8.4			
35-39	14.6	7.8	14.7	8.2	
40-55	13.9	7.7			
45-49	14.0	8.0			
50-54	13.2	7.1	15.2	8.8	
C:					
30-54	10.9	7.3			15% Sample
30-34	9.4	7.0			
35-39	10.1	7.6			
40-44	11.8	8.3			
45-49	12.7	8.9			
50-54	13.4	11.1			
<u>Non Census Data</u>					
EM	5.6	5.7	6.5	4.1	c
C'	20.1	16.4			b, d
W_{BLS1}	2.07	.19	2.07	.19	e
W_{BLS2}	1.92	.16	1.92	.16	e

Notes: a. 1970 census, <u>Characteristics of the Population</u>, vol. 2-51, Chapter D.

b. 1970 census, <u>Characteristics of the Population</u>, vol. 2-51, Chapter A.

c. Bureau of the Census, <u>U.S. County Business Patterns: 1968 and 1970</u>, Table 3A.

d. Johnson, D., Richard, P. and Quinn, B. <u>Churches and Church Membership in the U.S.: An Enumeration by Region, State and County</u>. Washington, D.C.: Glemnary Research Center, 1974.

e. Bureau of Labor Statistics, selected <u>Area Wage Surveys</u>, 1970, Tables A-1, A-5 and B-1.

Appendix B

Supplementary Regressions

Table B-1

Estimates of the Labor Supply Function with Published SMSA Data for White Wives, 1970.

(Labor Supply = LFPR in 1969, L_{69})

Age Group	Constant	W ($00's)	H ($00's)	N ($00's)	E	R	C	U
30-34	42.27*	.42*	-.18*	-.07	.55	.01	-.37*	-1.22*
35-39	42.08*	.28*	-.13*	-.01	1.40	.02	-.33*	- .92*
40-44	42.75*	.46*	-.17*	-.01	1.30	-.04	-.15*	-1.33*
45-49	52.90*	.34*	-.05	-.09	.23	-.02	-.05	-1.09*
50-54	50.84*	.53*	-.16	-.14	.11	-.04	-.05	-1.11*
30-54	45.4*	.57*	-.19*	-.13	.65	-.01	-.13*	-1.60*

* - indicates t-ratio greater than 1.96.

\+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 124 SMSAs.

Table B-2

Estimates of the Labor Supply Function with Published SMSA Data for White Wives, 1970.

(Labor Supply = Average Weeks Worked in 1969, L_w)

Age Group	Constant	W ($00's)	H ($00's)	N ($00's)	E	R	C	U
30-34	16.05*	.18*	-.07*	.03	.24	.01	-.13*	-.53*
35-39	13.43°	.13*	-.06*	-.04	.78	.01	-.12*	-.48*
40-44	7.66	.10+	-.05+	.04	1.43+	-.02	-.04	-.59*
45-49	25.30*	.15*	-.01	-.06	-.41	-.01	-.01	-.63*
50-54	22.22*	.26*	-.07*	-.10°	-.11	-.02	-.01	-.65*
30-54	17.3*	.29*	-.09*	-.07	.27	-.01	-.03	-.95*

* - indicates t-ratio greater than 1.96.

\+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 124 SMSAs.

Table B-3

Estimates of the Labor Supply Function with Published

SMSA Data for Black Wives, 1970.

(Labor Supply = LFPR in 1969, L_{69})

Age Group	Constant	W ($00's)	H ($00's)	N ($00's)	E	R	U
30-34	43.44°	-.24	-.23	-1.07	4.47°	.07	-.57+
35-39	67.61*	-.17	-.09	-1.33	1.49	.13	-.75+
40-44	46.69*	.34+	-.56*	-2.05	3.85*	.14	.11
45-49	63.59*	.54*	-.63*	-2.21°	2.27*	-.14°	-.76*
50-54	66.29*	.59*	-.60*	-2.63+	1.69+	-.10	-.27
30-54	56.07*	.31+	-.56*	-2.13°	3.27*	.02	-.30

* - indicates t-ratio greater than 1.96.

+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 79 SMSAs.

Table B-4

Estimates of the Labor Supply Function with Published SMSA Data for Black Wives, 1970.

(Labor Supply = Average Weeks Worked in 1969, L_w)

Age Group	Constant	W ($00's)	H ($00's)	N ($00's)	E	R	U
30-34	9.35	.05	-.16+	- .91	2.38+	.06	-.43*
35-39	25.34*	.06	-.12o	-1.13	.94	.04	-.41*
40-44	15.04*	.26*	-.27*	- .95	1.90*	.08	-.18
45-49	25.45*	.36*	-.30*	-1.43*	.99*	-.07o	-.54*
50-54	26.83*	.37*	-.27*	-1.48*	.61o	-.03	-.27+
30-54	20.33*	.28*	-.30*	-1.24+	1.56*	.02	-.37*

* - indicates t-ratio greater than 1.96.

+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 79 SMSAs.

Table B-5

Correlations Between Variable Measures from the 1970 Public Use Sample and Published Census Data

Age	F	L_{70}	L_{69}	L_w	W	H	N	E	R	C	U
					Whites						
30-34	.77	.74	.68	.66	.33	.76	.26	.28	.82	.87	.56
35-39	.81	.71	.71	.72	.02	.87	.24	.54	.81	.87	.60
40-44	.82	.61	.61	.57	.04	.74	.37	.53	.75	.87	.65
45-49	.71	.71	.68	.70	.53	.78	.53	.65	.75	.84	.48
50-54	.67	.67	.68	.64	.46	.71	.29	.68	.65	.75	.61
30-54	.92	.88	.84	.82	.50	.90	.54	.73	.85	.94	.85
					Blacks						
30-39	.69	.66	.64	.62	.70	.95	.32	.68	.77		.54
40-54	.85	.75	.69	.70	.59	.91	.48	.77	.81		.42
30-54	.88	.78	.79	.77	.74	.95	.55	.84	.88		.65

Table B-6

Estimates of the Labor Supply Function with Public Use Sample Data for White Wives, 1970 (Labor Supply = Normit of LFPR in Survey Week, NL_{70})[a]

Age Group	Constant	W	H ($00's)	N ($00's)	E	R	C	U
30-34	-18.75	.49	-.45*	.65	4.69	-.25	-.14*	-3.72*
35-39	-12.31	-1.20	-.41*	-.11	5.65°	-.27°	-.88*	-2.87*
40-44	-14.49	-1.21	-.11°	.16	2.73	.15	-.56*	-1.02
45-49	-95.08*	6.82*	-.31*	.17	9.65*	.27	.24°	-3.92*
50-54	-39.76	1.40	-.09	.19	3.84	.43+	-.07	-1.70
30-54	-51.18°	-1.95	-.15°	.32	5.81°	.17	-.41*	-2.28*

* - indicates t-ratio greater than 1.96.
\+ - indicates t-ratio greater than 1.645 (but less than 1.96).
o - indicates t-ratio greater than 1.282 (but less than 1.645).
Sample size: 134 SMSAs.

[a]The normit transformation is the inverse of the cumulative standard normal density function. Hence, the dependent variable is a standard normal variable. The dependent variable has been multiplied by 100 for convenience in presenting and reading the coefficient estimates. In addition, the weighting scheme was adjusted to correct for the heteroscedasticity beyond that caused by variation in SMSA sample size. See equation (8), p. 16.

Table B-7

Estimates of the Labor Supply Function with Public Use Sample Data for White Wives, 1970 (Labor Supply = LFPR in 1969, L_{69})

Age Group	Constant	W	H ($00's)	N ($00's)	E	R	C	U
30-34	68.26*	.62	-.19*	.36°	.56	-.17+	-.58*	-1.52*
35-39	57.68*	-.23	-.16*	-.04	1.98°	-.24*	-.46*	-1.40*
40-44	41.19*	-.70°	-.07+	.02	2.31°	-.02	-.29*	-.31
45-49	9.86	3.37*	-.16*	.19	4.79*	.06	.01	-1.82*
50-54	34.33+	-.40	-.01	-.01	2.03	.12	-.05	-.59
30-54	37.56*	-1.06	-.08*	.21	2.62*	-.03	-.25*	-1.07*

*indicates t-ratio greater than 1.96.

+indicates t-ratio greater than 1.645 (but less than 1.96).

°indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 134 SMSAs.

Table B-8

Estimates of the Labor Supply Function with Public Use Sample Data for White Wives, 1970 (Labor Supply = Average Weeks Worked in 1969)

Age Group	Constant	W	H ($00's)	N ($00's)	E	R	C	U
30-34	24.32*	-.05	-.07*	.15°	.30	-.06°	-.23*	-.62*
35-39	24.90*	-.23	-.08*	.00	.72	-.08*	-.19*	-.69*
40-44	21.63*	-.32°	-.03+	.01	.55	-.004	-.14*	-.33°
45-49	7.87	1.43*	-.06*	.06	1.66*	.03	-.003	-.91*
50-54	15.45+	-.23	-.01	.06	.86	.06	-.01	-.48+
30-54	18.31*	-.35	-.04+	.16°	.80	-.003	-.10*	-.60*

*indicates t-ratio greater than 1.96.

+indicates t-ratio greater than 1.645 (but less than 1.96).

°indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 134 SMSAs.

Table B-9

Estimates of the Labor Supply Function with 1970 Public Use Sample Data for Black Wives (Labor Supply = Normit of LFPR in Survey Week, NL_{70}; LFPR in 1969, L_{69}; and Average Weeks Worked in 1969, L_w)

Age Group	Constant	W	H ($00's)	N ($00's)	E	R	U
				NL_{70}			
30-39	-131.6^{+}	-11.77^{o}	-.20	-1.86	17.89*	1.01^{+}	-1.96^{o}
40-54	27.33	1.08	-.45	.21	4.30	-.21	-3.33
30-54	-66.65	-7.77	-.37	-.07	13.04^{o}	.59	-1.89
				L_{69}			
30-39	16.06	-5.10^{+}	-.05	-1.17	5.89^{+}	.46*	- .46
40-54	55.24*	- .57	-.09	- .69	2.05	.05	- .80
30-54	24.85	-4.53^{o}	-.09	-1.12	5.39^{+}	$.28^{o}$	- .30
				L_w			
30-39	2.84	-1.81^{o}	-.04	- .89	2.93*	.19*	- .57*
40-45	17.50^{o}	- .07	-.04	- .40	1.61	-.02	$- .76^{o}$
30-54	6.47	-1.25	-.04	- .68	2.66*	.11	$- .51^{o}$

* - indicates t-ratio greater than 1.96.

\+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 45 SMSAs

Table B-10

Estimates of the Labor Supply Function with 1970 Public Use Sample for White Wives and Correction for Sample Selection Bias (Labor Supply = Average Weeks Worked by Women Working in 1969 and the Survey Week, L_w')

Age Group	Constant	W	H ($00's)	N ($00's)	E	R	C	U	λ (percentage points)
30-34	35.7*	-.58+	.004	-.18*	.91+	-.01	.05+	.06	.06*
35-39	51.9*	-.17	-.04*	.07	-.35	-.05*	-.04*	.01	.01
40-44	53.8*	-.13	-.02o	.09o	-.67o	-.02	-.03*	.03	-.003
45-49	45.6*	-.40	.03*	.04	-.24	.03o	-.02*	-.03	-.002
50-54	45.4*	-.41	-.001	.06o	.27	-.02	.01	-.16	2.36+
30-54	46.9*	.14	-.02+	.09o	-.07	-.02o	-.01	-.12	1.07

* - indicates t-ratio greater than 1.96.

+ - indicates t-ratio greater than 1.645 (but less than 1.96).

o - indicates t-ratio greater than 1.282 (but less than 1.645).

Sample size: 134 SMSAs.