NBER WORKING PAPER SERIES

WHAT ARE THE DETERMINANTS OF DELAYED CHILDBEARING AND PERMANENT CHILDLESSNESS IN THE UNITED STATES?

David E. Bloom

James Trussell

Working Paper No. 1140

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge MA 02138

June 1983

Paper presented at the 1983 annual meetings of the Population Association of America, Pittsburgh, Pennsylvania. The authors are indebted to Clint Cummins and Barbara Vaughan for skillful programming assistance and to Arland Thornton for his help in locating some of the data used in this study. This research was funded by NICHD Contract #N01-HD9-12824 (Bloom) and NIH Grant #R01-HD11720 (Trussell). The research reported here is part of the NBER's research program in Labor Studies. Any opinions expressed are those of the authors and not those of the National Bureau of Economic Research.

NBER Working Paper #1140
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ABSTRACT

This paper presents estimates of delayed childbearing and permanent childlessness in the United States and the determinants of those phenomena. The estimates are derived by fitting the Coale-McNeil marriage model to survey data on age at first birth and by letting the parameters of the model depend on covariates. Substantively, the results provide evidence that the low first birth fertility rates experienced in the 1970's were due to both delayed childbearing and to increasing levels of permanent childlessness. The results also indicate that (a) delayed childbearing is less prevalent among black women than among non-black women, (b) education and labor force participation are important determinants of delayed childbearing, (c) the influence of education and labor force participation on delayed childbearing seems to be increasing across cohorts, (d) education is positively associated with heterogeneity among women in their age at first birth, (d) the dispersion of age at first birth is increasing across cohorts, (f) race has an insignificant effect on childlessness, and (g) education is positively associated with childlessness, with the effect of education increasing and reaching strikingly high levels for the most recent cohorts.

David E. Bloom
Department of Economics
Harvard University
Cambridge, Massachusetts 02138

(617) 495-4690

James Trussell
Office of Population Research
Princeton University
Princeton, New Jersey 08544

(609) 452-4946

I. Introduction and Background

During the 1970's, the first birth rate of American women reached its lowest level since the Great Depression. At the same time, the number of women having first births in their late twenties and early thirties showed a dramatic increase over the previous decades' experience. Some researchers attempted to explain these somewhat paradoxical facts by suggesting that they reflected a tendency of recent cohorts to delay their childbearing relative to that of older cohorts (see, for example, Sklar and Berkov, 1975; and Blake, 1979). In other words, they advanced the view that these facts were primarily due to a change in the timing of first births -- and not to a change in their completed level. This seemed to be a reasonable view since the most plausible alternative hypothesis -- an increase in permanent childlessness -- could not explain the increasing numbers of first births experienced by older women. In addition, it carried an important and reassuring implication, namely, that the first birth rate would soon begin to rise as the delayers began to reach their desired age at first birth.

Stimulated by these facts and by a desire to determine whether they were the result of delayed childbearing, increasing childlessness, or both, a number of independent research studies were conducted which focused on measuring recent changes in the timing and frequency of first births (see, for example, Masnick, 1980a; Bloom 1982; Morgan and Rindfuss, 1982; Morgan, 1982; and Mosher and Bachrach, 1982). Although these studies vary greatly in terms of the data they analyze (e.g., vital statistics data, retrospective survey data, or fertility expectations data), their analytical framework (e.g., period or cohort analysis), their statistical approach (e.g., simple examinations of age-specific first birth rates, complex parametric models, etc.), and the populations to which they refer (e.g., all women or ever-married women), their results are remarkably consistent:

they all provide evidence of either increasing childlessness, an increasing tendency to delay childbearing, or both.

The purpose of this paper is to further the analysis of age at first birth in the United States in two ways. First, we shall present new evidence on the tendency of recent cohorts of American women to delay their childbearing or to remain forever childless. This evidence is derived from fitting the Coale-McNeil marriage model to survey data on age at first birth. Because of its parametric nature, the Coale-McNeil model is extremely useful in this application since many of the cohorts whose first birth fertility patterns are of interest have yet to complete their childbearing years; when fit to incomplete data, estimates of the model permit one to project the remainder of a cohort's first birth fertility and thereby its mean age at first birth and proportion forever childless. Moreover, recent studies have established that the Coale-McNeil model provides a good fit to first birth data derived from vital registrations statistics both in the U.S. (Bloom, 1982) as well as in other countries (Bloom, 1983). Recent studies have also developed statistical methods and computer software for fitting this model to individual and household survey data on age at first birth (Rodriguez and Trussell, 1980). In addition, illustrative analyses demonstrating the application of these methods to survey data on age at first birth have been prepared for many of the countries in which World Fertility Surveys were conducted (Casterline and Trussell, 1980; Hobcraft and Trussell, 1980; Trussell, 1980). However, the Coale-McNeil model has yet to be applied to survey data for the U.S. In this study we remedy this deficiency by fitting the Coale-McNeil model to data on age at first birth from three recent surveys of American women: (1) Cycle II of the National Survey of Family Growth (conducted in 1976); (2) the young women sample of the National Longitudinal Survey (conducted in 1978); and (3) the Census Bureau's Current Population Survey (conducted in June 1980).

The second objective of this paper is to estimate the determinants of age at first birth in the United States. Most previous work has approached this problem by estimating the parameters of first birth schedules constructed separately for individual classifications of one or more different variables (e.g., the mean age at first birth by race group and years of education; see Trussell, 1980; Wilkie, 1981; and Bloom 1982). However, because cell sizes rapidly diminish as the number of variables and classifications increase, such attempts are severely limited by the availability of data. Multiple regression analysis has also been used to estimate the determinants of age at first birth (see Hirschman and Rindfuss, 1980; Masnick, 1980b; and Rindfuss, Bumpass, and St. John, 1980). Trussell and Bloom (1983) have shown, however, that regression analysis yields biased results if applied to a sample of women who have yet to complete their childbearing years. Moreover, regression analysis is less than fully satisfactory because it fails to incorporate existing knowledge about the age pattern of women at first birth (see Trussell, Menken, and Coale, 1982; Bloom, 1982; and Bloom, 1983). To effectively deal with these problems, Trussell and Bloom have developed a model which combines elements of both the Coale-McNeil model and of regression analysis. It does this by assuming that the Coale-McNeil model describes the underlying pattern of age at first birth but that its parameters depend on covariates in a regression-like manner.

In this paper we apply the Trussell-Bloom extension of the Coale-McNeil model to first birth data contained in the three surveys named above. The variables whose effects on age at first birth we estimate are: race, religion, rural-urban childhood residence, education, and labor force participation prior to first birth. We test various hypotheses about the effects of these variables both within and across cohorts and compare the results derived from the different data sets.

Section II provides a brief description of the Coale-McNeil marriage model, its application to survey data on age at first birth, and its extension to include covariate effects. Section III describes the three data sets used in this study. Section IV presents and discusses the results of fitting various specifications of the extended Coale-McNeil model to cohort data on age at first birth in each of the three survey data sets. Section V summarizes the results of the paper and comments on them in relation to results presented in other studies of age at first birth. This section also speculates on the implications of the results for the evolution of American fertility as well as for future research on the subject of American fertility.

II. Background on the Coale-McNeil Marriage Model and its Use in Estimating the Covariates of Age at First Birth [1]

The Coale-McNeil marriage model is based on the observation by Coale (1971) that a common structure underlies age distributions of first marriages in different populations. As shown by Coale and further supported by numerous other studies inspired by Coale's work, this distribution is smooth, unimodal, skewed to the right, and is close to zero below age fifteen and above age fifty. Furthermore, Coale observed that the differences in age-at-marriage distributions across female populations are almost entirely accounted for by differences in their means, their standard deviations, and their cumulative values at the older ages, e.g., age fifty. To facilitate the application of this finding, Coale

1. For further details, see the following series of papers: Coale (1971), Coale and McNeil (1972), Trussell, Menken, and Coale (1982), Bloom (1982, 1983), Rodriguez and Trussell (1980), Casterline and Trussell (1980), Hobcraft and Trussell (1980), Trussell (1980), and Trussell and Bloom (1983).

constructed a standard schedule of age at first marriage using data for Sweden, 1865-1869. In later work, Coale and McNeil (1972) developed a closed-form expression which closely replicated this Swedish standard (and many other observed marriage distributions, after suitably transforming their means, standard deviations, and cumulative values at age fifty). The mathematics leading to this expression also provided an appealing behavioral interpretation of the social process underlying entry into first marriage. According to this interpretation, age at marriage is viewed as the sum of a series of random variables, the first describing the age at which a woman first becomes marriageable (assumed to be normally distributed) and the others measuring the successive delays between becoming marriageable and meeting one's first spouse, meeting one's first spouse and becoming engaged, and becoming engaged and getting married (with these random variables all assumed to be exponentially distributed with parameters in arithmetic sequence).

Subsequent research has done little either to confirm or deny the behavioral interpretation of the Coale-McNeil model. However, the interpretation does suggest that the marriage model can also be applied to distributions of age at first birth. This conclusion hinges essentially on the assumption of an exponential delay between first marriage and first birth, which would be true if there were no childbearing outside of marriage, if all women were equally fecund, and if fecundability did not decline with age.[2] Recent empirical studies have confirmed the ability of the Coale-McNeil model to replicate first birth distributions and have demonstrated its usefulness in their analysis (see Trussell, Menken, and Coale, 1982; Bloom, 1982, 1983; Rodriguez and Trussell, 1980; Casterline and

2. The conclusion follows because the convolution of a normal and <u>four</u> exponential variables can be very closely approximated by the convolution of a normal and <u>three</u> exponential variables (see Coale and McNeil, 1972).

Trussell, 1980; Hobcraft and Trussell, 1980; and Trussell, 1980).[3]

In formal terms, the Coale-McNeil model can be expressed as:

$$g(a) = \frac{E}{\sigma}\ 1.2813\ \exp\ [-1.145\ (\frac{a-\mu}{\sigma} + .805) - \exp\{1.896\ (\frac{a-\mu}{\sigma} + .805)\}] \qquad (1)$$

where g(a) is the proportion of women having their first birth at age a in the observed population and μ, σ, and E are, respectively, the mean and standard deviation of age at first birth (for those who ever have a first birth), and the proportion ever having a first birth.[4] Rodriguez and Trussell (1980) have derived the likelihood function associated with this model and have developed a computer program to estimate its parameters from survey data drawn either from a sample of all women or from a sample of women who had a birth prior to the survey date.[5] In the latter case, only the parameters μ and σ are estimated; E must be set at unity.

Trussell and Bloom extend this formulation by deriving the likelihood function which allows each of the (two or) three parameters to depend on covariates. For simplicity, they assume a linear relationship.

3. All of these studies conclude that the marriage model provides a good fit to first birth data, with the exception of the studies by Casterline and Trussell and Hobcraft and Trussell. However, it is likely that the "negative" results reported in those two studies were caused by age misstatement, sampling error, and period-related irregularities in the WFS data analyzed. Since similar problems may plague the present analysis we shall proceed cautiously and compare our results across data sets and with results based on aggregate data (which are less subject to such problems).

4. This form of the marriage model is a reparameterization of the original form presented in Coale and McNeil (1972). It was derived by Rodriguez and Trussell (1980) and is used here because it expresses the model in terms of parameters that are intuitively easier to understand than Coale's a_0, K, and C (although E = C).

5. The program is entitled NUPTIAL and is available from the World Fertility Survey (in London) at a nominal cost.

$$\mu_i = X_i'\beta \tag{2}$$
$$\sigma_i = Y_i'\gamma$$
$$E_i = W_i'\alpha$$

where the index i denotes individual i; X_i, Y_i, and W_i are the vector values of characteristics of that individual that determine respectively μ_i, σ_i, and E_i and β, γ, and α are the associated parameter vectors. As noted by the authors, the covariate vectors may or may not be different; however, in all cases, standard statistical tests can be used to draw inferences about the parameters. The authors also develop a computer program which computes maximum likelihood estimates of the parameter vectors β, γ, and α. The program uses the routine DFP in the numerical optimization package GQOPT.[6] For computational ease, the program requires the covariates to be categorical in nature.[7,8]

III. The Data

As noted in Section I, this study uses three independent data sets to estimate the age patterns of American women at first birth and their covariates. The use of multiple data sets is prompted by the fact that no one data set is uniquely well-suited to the tasks at hand. In addition, we feel that the

6. The routine DFP is described in Goldfeld and Quandt (1972, pp. 5-9). The package GQOPT is available from the Econometric Research Program, Department of Economics, Princeton University.

7. A program to estimate the extended Coale-McNeil model is available from the authors upon request (and at cost). It is a modified version of NUPTIAL which is much easier to use than the program used to compute the estimates in Trussell and Bloom (1983).

8. Trussell and Bloom (1983) also propose and investigate the use of a proportional hazards model in estimating the covariates of age at first birth. However, that model is not used in this study because (a) it can only be fit to data from an all-woman sample, (b) it cannot be used to project, and (c) empirically, it performed no better than the extended Coale-McNeil model in illustrative analyses presented in Trussell and Bloom (1983).

consistency of results derived from different sources of information is an important indication of their strength. The remainder of this section provides a brief description of each of the three data sets.

A. National Survey of Family Growth (NSFG), Cycle II [9]

Cycle II of the NSFG was conducted in 1976 by the National Center for Health Statistics through personal interviews with 8611 women aged 15-44 years. To be eligible for interview the women had to be either currently married, previously married, or never-married mothers with offspring living in the same household. Thus, the NSFG is a representative sample of ever-married women and never-married women with children present in their household. It is not a representative sample of never-married women who have had no children or of never-married mothers whose children do not live in their household.

For the purposes of this study, the NSFG is useful because it contains information on age at first birth along with several other retrospective socioeconomic variables that presumably influence the age at first birth. These variables and the coding scheme adopted for them are: race (black or not-black), religion (Catholic or non-Catholic), childhood residence (rural or urban), education at time of survey (less than high school, high school, greater than high school), and employment history prior to first birth (did or did not ever work). All women aged 25-44 at the time of the survey who had a first birth between ages 12 and 44 are included in our data file. Because we do not have information on women who never had a first birth, we cannot estimate the parameter E (i.e., the proportion ever having a first birth) from this sample; nor can we estimate its covariates. Observations were counted more or less heavily depending on their

9. A comprehensive publication detailing the design of the NSFG (Cycle II) is provided by U.S. Department of Health and Human Services (1981).

sample weights, with the weights adjusted to have mean unity.

B. National Longitudinal Survey (NLS) of Young Women, 14-24 [10]

This NLS survey has been conducted yearly since 1968 when it started with 5159 women aged 14-24. The main purpose of this survey is to gather information on the labor market experiences of young women. As a result, it is primarily oriented toward questions on a wide range of socio-economic variables. However, in 1978, a complete reinterview of the original sample of women was conducted and a question on age at first birth was asked. Thus, we have used the 1978 NLS tape to construct a data set on age at first birth for women aged 24-34 in 1978.[11] Sample weights were used in the creation of this data set after adjusting the weights so they average to one.

In comparison to the NSFG data, the NLS data are more useful because the sample refers to all women and because the data are more recent. On the other hand, the NLS data have a smaller sample size, they refer to a narrower group of ages, and they contain information on fewer socio-economic variables relevant to a study of age at first birth. The variables used are race (black or non-black), childhood residence (rural or urban), and education at time of survey (less than high school; equal to high school; greater than high school). In addition, the NLS data may be somewhat nonrepresentative because of sample attrition, although the 1978 reinterview includes 76 percent of the original participants.

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10. For further details on the NLS, see Center for Human Resource Research (1982).

11. A few observations on women aged 35 in 1978 are also included in our sample.

C. Current Population Survey (CPS) [12]

The CPS is a nationwide sample survey conducted monthly by the Bureau of the Census. It involves detailed personal interviews in about 60,000 households in which information on a variety of demographic, social, and economic variables is recorded. The unit of observation is the individual; the sample universe consists of all persons living in the surveyed households.

In the June, 1980 CPS, the normal set of questions was supplemented with a set of retrospective marital and fertility questions. Included on the supplementary survey instrument was a question on age at first birth which was asked for all women aged 18-75. Unfortunately, there are few retrospective covariates in the CPS which could sensibly be hypothesized to affect age at first birth. However, we have constructed the following two variables: race (black, not black) and education at time of survey (less than high school, high school, greater than high school).[13]

Although the CPS data set only permits estimation of two covariate effects, it is extremely useful in this study because (a) it refers to all women, (b) it includes an exceptionally large number of observations which permits parameter estimation for single-year cohorts, and (c) it is the most recent of the three data sets used in this study. As with the two other sets of data, sample weights were used in creating this data file after adjusting them so they average to one.

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12. For further details see U.S. Bureau of the Census (1980).

13. With the exception of education, all of the covariates used in this study measure individual characteristics at the time of first birth. We define education on the basis of years completed prior to the survey rather than years completed prior to the first birth because we believe the former measure is a (marginally) superior social indicator and because it can be constructed for all three data sets. However, empirical results differed insignificantly when we experimented with the two alternative measures on the NLS data.

IV. Results

A. Estimates Computed Without Covariates

Table I presents the results of fitting the Coale-McNeil model to the three sets of first birth data described in the previous section. Note that these results treat the estimated parameters μ and σ (and E) as constants, i.e., they are not allowed to depend on covariates. Note also that, in order to facilitate the detection of changes over time, separate estimates were computed for each of the age groups indicated. For the sake of comparability with the results discussed in Section IV.B, these age groups were chosen to satisfy sample size requirements for estimation with covariates. In addition, we were, in some cases, able to compute estimates for younger cohorts than those included in Table I. However, because those estimates suggested the data were truncated below the mean age at first birth, we have chosen not to report them.[14]

Substantively, the results in Table I exhibit three interesting patterns. First, all three data sets show an upward trend in the mean age at first birth (μ) across recent cohorts, with the increase ranging from about .3 years in the NLS data to about 1.5 years across a wide range of cohorts in the NSFG and CPS data sets. This trend provides some evidence of delayed childbearing among recent cohorts although the mean age at first birth is not necessarily the best indicator of that phenomenon (see Bloom, 1982, pages 365-6). However, examination of a better indicator (not reported in Table I) -- the projected proportion of women who have a first birth between ages 25 and 34 expressed as a fraction of those who ever have a birth -- also reveals an increase across cohorts from about .23, .26, and .25 to about .32, .28, and .32 for the three data sets, respectively. These trends provide somewhat stronger confirmation of the increasing

14. Bloom (1982, p. 355 and n. 10) concludes that such estimates are likely to be seriously misleading.

tendency of recent cohorts to delay childbearing.

Second, the results from all three data sets show an upward trend in the standard deviation of age at first birth across cohorts. This finding reflects increasing heterogeneity in the age at which women experience their first birth.

Third, the results computed for the NLS and CPS data provide strong evidence of an increase across cohorts in the level of permanent childlessness, i.e., 1.0 - E. More specifically, according to these results, the incidence of childlessness among the most recent cohorts of women included in this analysis will reach 20 to 25 percent, which represents a substantial increase over the 10 percent rate which prevailed (or is projected to prevail) among the older cohorts.

Before we turn to the next sub-section's discussion of covariate effects, two additional points deserve mention. First, the parameter estimates reported in Table I are remarkably consistent across data sets, both in terms of their levels and their trends (see Figures 1 and 2). This finding provides considerable support for the external validity of these estimates. Second, the estimates of μ and σ (and E) computed from the three survey data sets are also remarkably similar to estimates reported in Bloom (1982) which were based on aggregate vital statistics data. This observation provides support for the results presented in that earlier study and also enhances our confidence in the results presented herein.

B. Estimates Computed With Covariates

The results of fitting the extended Coale-McNeil model to survey data on age at first birth are presented in Tables II, III, and IV. The results we present are representative of the broader set of results we computed in the process of conducting this research. In order to facilitate hypothesis testing, the results presented also refer to specifications which are successively nested in each other. In addition, since the covariates entered are, in all cases, categorical, their effects must be interpreted relative to the appropriate reference category. Depending on the data set and specification, these reference categories are always, when appropriate (1) non-black, (2) Catholic, (3) urban childhood residence, (4) completed education less than high school, and (5) did not work prior to first birth. The covariates all have linear effects on the Coale-McNeil parameters although their effect on age at first birth is highly nonlinear.

At the outset it should be noted that aggregate trends in age at first birth can be affected by the covariates in two ways. First, the model can remain the same across cohorts but values of the covariates can change. For example, it might be found that one year of increased education always increases age at first birth by 1.25 years. If educational attainment increases for each successive cohort, age at first birth will, as a consequence, increase in the population. Alternatively, the model may change across cohorts. For example, the effect of an additional year of education on age at first birth may increase across cohorts from 1.0 years to 1.5 years. Such a change in the model will also affect the aggregate age at first birth. This effect is independent of the effect of changing educational attainment and can be discerned by estimation of the model we propose. Of course, in practice, it is likely that the two effects operate simultaneously although it is useful to disentangle them, which is what we do below.

In choosing variables for inclusion as covariates, we were limited by the nature of the available data. Nevertheless, of those variables that were available in each data set, we chose covariates whose effect on fertility has been either suggested or demonstrated in other studies (see, especially, Waite and Stolzenberg, 1976; DeJong and Sell, 1977; Veevers, 1970; Westoff and Jones, 1979; Masnick, 1980b; Rindfuss, Bumpass, and St. John, 1980; Wilkie, 1981; Bloom 1982;

Bloom and Pebley, 1982; Mosher and Bachrach, 1982; Morgan and Rindfuss, 1982.) Thus, over _a priori_ expectations are (1) that μ is negatively related to being black and having an urban childhood residence, but positively related to years of education and participation in the labor force,[15] and (2) that E is negatively related to education, labor force participation, and urban childhood residence, but positively related to being Catholic and being black.[16]

We begin our discussion of substantive results with the estimates presented in Table II for the NSFG data. The first set of columns presents the estimates computed when μ and σ are _both_ modeled as linear functions of a constant and variables which measure race, religion, childhood residence, education, and labor force participation prior to first birth. The second set of columns presents estimates of the same model except that the covariate effects on σ are constrained to be zero. Both of these specifications are generalizations of the model whose estimates are reported in Table I in which covariate effects are constrained to be zero for both μ and σ.

The most notable result of Tables I and II is that the incorporation of covariates into the model adds significantly to the model's explanatory power. As can be easily verified by performing the appropriate likelihood ratio tests, this statement holds true for all cohort groupings when covariate effects are allowed for _both_ the mean and the standard deviation. Moreover, the pattern of covariate effects is basically consistent with our _a priori_ expectations,

15. We have no _a priori_ prediction of the effect of being Catholic on μ since the contraceptive practices of Catholics suggests a negative effect while the prohibition on sex before marriage suggests a positive effect.

16. Our statistical procedure makes no correction for simultaneity bias which may be introduced by the reciprocal effect of age at first birth on the covariates in equations (2), e.g., on education. However, we believe this limitation of our procedure is mitigated by the use of broad educational categories and by the findings of Waite and Stolzenberg (1976) and Masnick (1980b) which provide little evidence of such reciprocal effects.

although there are some surprises. First, education and labor force participation prior to first birth have positive and statistically significant effects on μ. These results indicate that more educated women and childless women who work (and ultimately bear children) are more likely to delay childbearing. In addition, the effect of labor force participation is greater for more recent cohorts. When coupled with the fact that labor force participation rates for (young) females have risen over time, this finding suggests that labor force participation is becoming an increasingly important factor underlying the aggregate trend to delay childbearing. Education appears to be another important determinant of this trend. Since the parameter estimates do not change much across cohorts, education influences age at first birth in the population because successive cohorts have higher levels of educational attainment.

Second, the effects of race, religion, and childhood residence on μ all tend to be small in magnitude, i.e., less than 1 year, and are often statistically insignificant. Of all these effects, perhaps the most surprising is the small race effect which is contrary to the significant negative effect found in most other studies (e.g., Wilkie, 1981; Bloom, 1982, and Morgan and Rindfuss, 1982). However, keep in mind that the race effects reported in those other studies are based on models that are univariate in nature, unlike the race effects reported in Table II, which hold other variables such as education and labor force participation, fixed. In fact, in comparison to the results in Table II, estimates (not reported here) of the race effect for specifications in which no other covariates are included are always larger and are often statistically significant. Thus, in contrast to other studies, the NSFG results suggest that the independent effect of race on age at first birth is small, although it appears that race does have an indirect effect on age at first birth which operates through its effect on other covariates which influence age at first birth, e.g., education.[17]

Third, the standard deviation of age at first birth appears to be relatively high for blacks, for women who work prior to their first birth, and for Catholics. Although these results are somewhat difficult to assess, they do suggest that women with these characteristics (who ultimately bear children) are (or will be) more heterogeneous in the timing of their first births than women without them. In addition, the effect of labor force participation on the degree of heterogeneity appears to be growing across cohorts.

Let us now consider the estimation results in Table III for the NLS data. The organization of this table is similar to that of Table II except that we now report estimates of the parameter E and its covariates although the number of covariates is reduced.

In general, the results presented in Table III strongly support the inclusion of covariates. The value of the log likelihood is significantly increased when we allow for covariate effects on μ or on μ and E.[18] Moreover, tests of significance performed for individual estimates suggest that race and education are important determinants of the mean age at first birth while residence and education are important determinants of the proportion ever having a first birth. More specifically, the effect on μ of being black is negative and significant, holding education and residence constant. Furthermore, the estimated race

17. On the basis of a multivariate regression analysis, Masnick (1980b) also finds that the negative effect on age at first birth of being black is attenuated by the inclusion of other variables, and especially by including an education variable.

18. Although we do not report the results here, we were also not able to reject the hypothesis that all three parameters depend on covariates. The results are not reported because including covariates for σ generally had little effect on estimates of the covariate effects for μ and E, and because the pattern of results for the covariates of σ are less interesting than the results for μ and E and were, in fact, similar to those computed from the NSFG data.

effects are about one year greater than those estimated from the NSFG and they are attenuated less by the inclusion of other covariates (although fewer covariates are actually included).[19] The race effects also increase across the two cohort groups, suggesting that the tendency to delay childbearing is less characteristic of black women than of non-black women (since the intercept also increases a little). Also increasing across cohorts are the effects of education which are positive and greater in magnitude than those computed for the NSFG data. Thus, the NLS results suggest that education, i.e., increasing educational attainment combined with the increasing education effect, is an (increasingly) important factor in the delay of childbearing.

The NLS results also provide interesting estimates of σ and of the determinants of E. First, the estimates of σ increase across the two cohort groups, providing evidence of increasing dispersion in age at first birth within covariate cells. Second, the effect of race on E is small and insignificantly different from zero while the residence effect is significant and operates to increase E by three to four percentage points for women with rural backgrounds. On the other hand, education has a negative effect on E, with the effect being small for women who do not continue their education past high school. However, women who do continue their educations past high school have substantially lower probabilities of ever having a first birth. Thus, education appears to be an important determinant of childlessness.

Finally, let us turn to the results computed from the June, 1980 CPS that are reported in Table IV. Like the results presented for the NLS data in Table III, the CPS results are for two separate specifications, one in which μ and E

19. These findings are essentially unchanged when we compare race effects estimated from identical specifications in the two data sets, i.e., when we drop the labor force participation variable and the religion variable from the NSFG model.

depend on covariates (i.e., σ is treated as a constant) and one in which only μ depends on covariates (i.e., σ and E are constants). Although the nature of the CPS data limits us to the estimation of just two covariate effects -- race and education -- the sample sizes are large enough to permit an analysis of results for single year cohorts. Thus, we may focus our attention more closely on cross-cohort changes in covariate effects.

The CPS results are similar to the NSFG and NLS results in several ways. First, likelihood ratio tests do not permit us to reject the hypothesis that μ separately, or μ and E together, depend on covariates. On the other hand, for about one half of the cohorts we were able to reject the hypothesis that σ depends on covariates (when μ and E both allow for covariate effects).[20] Second, the CPS results show that being black has a significant negative effect on μ, with the estimated effect being closer in magnitude to the effect estimated from the NLS data than to the effect estimated from the NSFG data (even when comparable models are estimated). Moreover, the negative race effect seems to be increasing in absolute value across cohorts, a finding which provides further evidence that delayed childbearing is primarily a phenomenon that is associated with non-black women (since the intercept also increases slightly). In addition, results not reported here show that the race effect is attenuated by the inclusion of education as a covariate. Third, the CPS results show that education has a significant positive effect on μ with the magnitude of the estimated effect being roughly similar to that estimated from the NSFG and NLS data. However, the increase across cohorts in the magnitude of the education effect is particularly striking and provides strong evidence that education is an important determinant of delayed childbearing (see Figure 3).

20. We computed, but do not report, the estimates necessary to confirm this statement.

The final results of interest in Table IV relate to the covariates of E. In general, the results provide little evidence of a race effect with the coefficient on the race variable usually being small in magnitude and statistically insignificant. Thus, like the NLS results, the CPS results also provide no evidence that race is an important determinant of permanent childlessness. On the other hand, education does appear to be an important determinant of childlessness. The coefficients on education are generally negative and significant with magnitudes that are particularly large for women who continue their educations beyond high school. Moreover, the education effects show fairly dramatic increases across cohorts, ranging from essentially zero in the cohorts aged 35 and over to nearly twenty-five percent in the youngest cohorts (see Figure 4). Thus, not only is education an important determinant of childlessness, it is also a determinant whose importance appears to be growing.

V. Summary and Conclusions

This paper has presented estimates of delayed childbearing and permanent childlessness in the United States and of the determinants of those phenomena. The estimates of delayed childbearing and permanent childlessness were derived by fitting the Coale-McNeil marriage model to survey data on age at first birth. The determinants of those phenomena were derived by estimating the extended version of the model proposed by Trussell and Bloom (1983) in which the parameters of the model are allowed to depend on covariates. The covariates of the parameter E (i.e., the proportion of women ever having a first birth) are interpreted as covariates of permanent childlessness (after reversing their signs). The covariates of the parameter μ (i.e., the mean age at first birth) are interpreted as covariates of delayed childbearing. We also discuss the covariates of σ

(i.e., the standard deviation of age at first birth) since that parameter also relates to the phenomenon of delayed childbearing.

Estimates are computed for cohorts of women covered by three sets of data: the National Survey of Family Growth (1976), the National Longitudinal Survey of Young Women (1978), and the Current Population Survey (June, 1980). The first set of estimates refer only to women who ever have a first birth (i.e., we do not estimate E or its covariates) while the second and third sets of estimates refer to all women. Since the underlying pattern of age at first birth is represented by a parametric model, we are able to compute consistent estimates of parameters and covariates even for cohorts that have not yet completed their childbearing years. This is an important feature of our study since existing folklore on delayed childbearing and increasing childlessness suggest that they are both phenomena which refer primarily to the fertility of recent cohorts.

The results of this study provide new evidence that the fertility behavior of recent cohorts of American women is characterized by both delayed childbearing and increasing childlessness. Because our results are based on survey data, they complement those presented in Bloom (1982) which support similar conclusions using comparable methods, but with aggregate data. The results also provide strong support for the extension of the Coale-McNeil model to include covariate effects. In virtually every specification we estimated, the explanatory power of the model was significantly increased by adding covariates. Moreover, estimates of the effects of different covariates reveal that (a) delayed childbearing is less prevalent among black women than among non-black women, (b) education and labor force participation are important determinants of delayed childbearing, (c) the influence of education and labor force participation on delayed childbearing seems to be increasing across cohorts, (d) education is positively associated with heterogeneity among women in their age at first birth, (e) the dispersion of age at first birth is increasing across cohorts, even after controlling for the effect of different covariates on μ and E, (f) race has an insignificant effect on childlessness, and (g) education is positively associated with childlessness, with the effect of education increasing across cohorts and reaching strikingly high levels for women in recent cohorts who continue their educations beyond high school.

Before concluding this paper, we comment briefly on the significance and implications of these findings. First, the results presented provide strong evidence of changing cohort fertility patterns and determinants. This finding highlights the importance of adopting a cohort approach to the study of initiation of childbearing. In addition, it suggests that attempts to project incomplete cohort fertility by reference to the completed fertility of older cohorts may be misleading because of the likelihood that substantially different models are generating the two patterns.

Second, the results of this study are consistent with some of the results of other studies of the determinants of delayed childbearing and permanent childlessness. For example, our results are consistent with the results of Masnick (1980b) and Wilkie (1981) on the direction of the effects of education and race on age at first birth. Our results also conform to Masnick's (1980b) finding on the insignificance of childhood residence. On the other hand, our results suggest that being Catholic has an insignificant effect on age at first birth, unlike the result in Masnick. Finally, our results on the determinants of childlessness are similar to those of DeJong and Sell (1977) who conclude that education and labor force participation have positive effects on the incidence of childlessness, and to those of Mosher and Bachrach (1982) who find an important education effect.

Finally, the results of this study strongly suggest that cohort fertility patterns are becoming increasingly heterogeneous. For example, recent cohorts show much greater differences in the incidence and timing of their first birth fertility than do older cohorts. Moreover, the differences are not solely the result of the changing distribution of individual characteristics across cohorts, e.g., increasing educational attainment and labor force participation for a substantial fraction of the cohort. Rather, the differences also seem to be the result of particular characteristics having greater effects on first birth fertility. Thus, it appears that women's fertility patterns will, to a greater extent than ever before, be differentiated on the basis of observable characteristics. Certainly, the results of this study provide evidence that race, education, and labor force participation are important indicators of those differences. Nevertheless, it seems apparent that other variables which we do not control for are also having an impact. Thus, we recommend further application of the models used here to data sets which will permit richer covariate specifications. We also recommend that demographic surveys include more retrospective questions relating to social, economic, demographic, and attitudinal variables which may be related to first birth decisions. Greater use of longitudinal survey designs is also desirable. We already have suitable analytical constructs and some indication that fertility decisions will increasingly depend on observable information. What we need now are richer data sets so that future research can explore the determinants of age at first birth more fully.

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Table I - Estimates of the Coale-McNeil Model Without Covariates*

Data Set	Cohort	μ	σ	E	-ln L**	N
NSFG	25-29	23.9	5.4	--	3581.2	1530
(1976)	30-34	22.9	4.9	--	3938.0	1489
	35-39	22.4	4.6	--	3556.5	1304
	40-44	22.4	4.5	--	3345.1	1221
NLS	24-34+	23.2	5.3	.86	10393.3	4275
(1978)	24-29	23.1	5.6	.80	5455.7	2437
	30-34+	22.8	4.5	.89	4860.6	1838
CPS	25	23.6	5.6	.75	2997.4	1536
(1980)	26	24.1	6.2	.79	3274.4	1561
	27	23.8	5.8	.81	3278.4	1445
	28	24.7	6.4	.87	3520.9	1474
	29	24.1	5.8	.85	3556.8	1426
	30	23.3	5.2	.83	3617.5	1418
	31	23.8	5.4	.84	3654.3	1400
	32	23.4	5.3	.86	3792.0	1416
	33	23.8	5.7	.90	4056.6	1462
	34	23.3	5.0	.86	3116.3	1137
	35	22.8	4.8	.85	2986.4	1089
	36	22.7	4.9	.87	3228.4	1164
	37	23.0	5.0	.92	3496.2	1226
	38	22.5	4.6	.90	3020.8	1081
	39	22.3	4.4	.90	2963.3	1068
	40	22.5	4.6	.90	2859.0	1019
	41	22.9	5.0	.88	2681.2	942
	42	22.4	4.8	.92	2577.5	907
	43	22.6	4.6	.90	2529.5	897
	44	22.6	4.7	.91	2675.6	944
	45	23.1	5.0	.89	2854.6	991
	46	22.6	4.8	.91	2390.4	836
	47	23.1	4.7	.91	2384.5	834
	48	23.1	5.0	.89	2599.3	903
	49	23.1	5.0	.87	2509.7	876
	50	23.4	5.3	.89	2776.9	947

*All estimates are significant at the .01 level.

+This cohort also includes some data for women aged 35.

**-Log Likelihood

NOTE: μ is an estimate of the cohort's mean age at first birth;
σ is an estimate of the standard deviation of age at first birth for the cohort;
E is an estimate of the proportion of women in the cohort ever having a first birth.

Table II - Estimates of the Coale-McNeil Model with Covariates, 1976 NSFG Data*, **

		COHORT							
	Variable	25-29	30-34	35-39	40-44	25-29	30-34	35-39	40-44
μ	Constant	21.613	22.205	21.087	21.758	20.972	21.239	20.503	21.163
	Black	0.357*	0.146*	-0.153	-0.187*	-0.552	-0.517	-0.391	-1.202
	Non-Catholic	-0.144*	-1.299	-0.666	-0.730	-0.268	-0.136*	0.135*	-0.221*
	Rural	-0.238*	0.581	0.382	-0.309*	0.113*	-0.203*	-0.001*	0.036*
	Ed = High School	1.227	1.295	1.833	1.824	2.043	2.009	2.163	2.073
	Ed > High School	4.020	4.837	4.525	3.057	3.532	3.546	3.558	3.080
	Worked before FB	3.416	2.961	2.258	2.106	1.914	1.701	1.494	1.395
σ	Constant	4.168	4.716	3.965	4.312	3.282	3.720	3.510	3.523
	Black	0.862	0.793	0.532	1.011				
	Non-Catholic	-0.035*	-1.118	-0.753	-0.552				
	Rural	-0.225*	0.609	0.411	-0.582				
	Ed = High School	-1.035	-0.861	-0.288*	-0.343				
	Ed > High School	0.251*	1.160	1.082	-0.204*				
	Worked before FB	1.812	1.577	1.001	0.944				
	-Log Likelihood	3175.4	3540.6	3248.8	3087.5	3224.8	3634.1	3290.9	3117.0

*Coefficient not significant at the 10% level, two-tailed test.

**Sample sizes are provided in Table I.

Table III
Estimates of the Coale-McNeil Model With Covariates, 1978 NLS Data*, **

		COHORT				
	Variable	24-34++	30-34++	24-34++	24-29	30-34++
μ	Constant	20.833	20.728	21.134	21.009	20.778
	Black	-1.836	-1.277	-1.965	-2.340	-1.289
	Rural	-0.194	0.030*	-0.291	-0.543	0.006*
	Ed = HS	2.226	1.975	2.287	2.511	1.968
	Ed > HS	4.121	3.561	4.698	5.234	3.652
σ	Constant	4.305	3.833	4.635	4.781	3.886
E	Constant	0.920	0.923	0.884	0.796+	0.887
	Black	0.003*	-0.028*			
	Rural	0.044	0.033			
	Ed = HS	-0.036	-0.020*			
	Ed > HS	-0.268	-0.140			
	-Log Likelihood	9713.05	4626.14	9823.38	5112.56	4653.41

*Coefficient not significant at the 10% level, two-tailed test.

**Sample sizes are provided in Table I.

+E was fixed at .796 in this run because its estimated value in unconstrained estimation was implausibly high. In fixing E this way, we follow the advice of Rodriguez and Trussell (1980) and Trussell and Bloom (1983).

++This cohort also includes some data for women aged 35.

Table IV
Estimates of the Coale-McNeil Model With Covariates, 1980 CPS Data*, **

	μ				σ	E				
COHORT	Constant	Black	Ed=HS	Ed>HS	Constant	Constant	Black	Ed=HS	Ed>HS	-ln L
25	21.34	-2.40	2.08	3.69	4.76	0.95	0.03*	-0.10	-0.47+	2787.2
26	21.72	-2.78	2.12	4.15	5.23	0.96	0.06	-0.11	-0.41+	3061.9
27	21.12	-2.74	2.38	3.94	4.77	0.95	-0.00*	-0.09	-0.34+	3084.8
28	21.26	-1.90	2.38	4.75	5.08	0.95	-0.03*	-0.03*	-0.28+	3323.1
29	21.60	-1.94	1.94	3.91	4.97	0.92	0.03*	-0.05	-0.18	3408.3
30	21.23	-1.71	1.94	3.39	4.48	0.93	0.02*	-0.06	-0.24	3457.8
31	21.89	-2.27	1.36	3.32	4.72	0.93	-0.01*	-0.07	-0.18	3525.7
32	21.09	-2.01	2.28	3.38	4.67	0.93	0.01*	-0.07	-0.15	3677.0
33	21.71	-2.31	1.83	3.61	5.06	0.94	0.00*	-0.05	-0.09	3933.7
34	21.47	-1.88	2.00	3.09	4.57	0.92	0.05	-0.02*	-0.16	3007.8
35	20.91	-1.46	1.57	3.79	4.26	0.88	-0.00*	0.02*	-0.13	2862.9
36	20.67	-2.12	2.00	3.53	4.19	0.90	0.05	-0.03*	-0.06	3105.3
37	21.30	-1.31	1.93	3.01	4.65	0.91	0.02*	0.03*	-0.04	3411.4
38	20.98	-1.12	1.51	2.92	4.30	0.91	0.02*	0.05	-0.08	2930.4
39	20.87	-1.82	1.43	2.84	3.92	0.94	-0.06	-0.02*	-0.06	2864.4
40	21.56	-1.75	0.94	2.36	4.24	0.94	-0.03*	-0.03*	-0.08	2787.4
41	21.42	-1.82	1.72	2.81	4.52	0.89	-0.01*	0.02*	-0.05	2614.3
42	20.75	-1.11	1.86	3.09	4.29	0.93	0.01*	-0.01*	-0.02*	2502.4
43	21.32	-1.15	1.28	2.77	4.26	0.93	0.05	-0.03*	-0.06	2471.0
44	21.15	-0.88	1.53	2.97	4.32	0.94	-0.05*	-0.03*	-0.04*	2610.1
45	21.75	-1.32	1.31	2.72	4.66	0.88	0.02*	0.02*	-0.01*	2805.3
46	21.31	-1.09	1.55	2.93	4.44	0.89	0.07	0.02*	-0.01*	2333.1
47	22.22	-1.97	1.10	1.87	4.42	0.94	-0.04*	-0.02*	-0.05	2346.0
48	21.86	-1.62	1.49	2.47	4.58	0.89	-0.04*	0.03*	-0.04*	2545.5
49	21.64	-1.20	2.00	2.84	4.62	0.89	-0.06*	-0.02*	-0.03*	2454.5
50	21.79	-1.79	2.20	3.13	4.76	0.91	-0.01*	-0.02*	-0.06	2703.6
25	21.30	-2.76	2.18	4.89	4.96	0.75++				2837.8
26	21.66	-3.12	2.27	5.24	5.43	0.79++				3102.7
27	21.25	-2.98	2.47	4.71	5.08	0.81++				3121.0
28	21.48	-1.98	2.42	5.49	5.47	0.87++				3350.4
29	21.77	-2.04	2.00	4.34	5.26	0.86				3422.7
30	21.34	-1.77	1.98	3.72	4.69	0.83				3489.5
31	21.93	-2.28	1.40	3.52	4.83	0.84				3541.4
32	21.11	-2.05	2.31	3.51	4.74	0.85				3690.2
33	21.72	-2.32	1.86	3.71	5.12	0.89				3939.0
34	21.51	-1.92	2.00	3.20	4.63	0.86				3031.7
35	20.93	-1.48	1.56	3.85	4.30	0.85				2880.2
36	20.67	-2.12	2.00	3.55	4.20	0.87				3109.2
37	21.31	-1.32	1.91	3.03	4.66	0.92				3419.4
38	20.99	-1.12	1.50	2.94	4.31	0.90				2949.2
39	20.87	-1.82	1.43	2.84	3.93	0.90				2869.0
40	21.56	-1.75	0.94	2.37	4.24	0.90				2792.8
41	21.43	-1.82	1.72	2.82	4.53	0.88				2618.3
42	20.75	-1.11	1.86	3.09	4.29	0.92				2502.8
43	21.32	-1.15	1.28	2.77	4.26	0.90				2474.5
44	21.15	-0.87	1.53	2.97	4.32	0.91				2612.5
45	21.75	-1.32	1.31	2.72	4.66	0.89				2806.3
46	21.31	-1.10	1.56	2.94	4.44	0.91				2337.9
47	22.22	-1.97	1.10	1.87	4.42	0.91				2348.4
48	21.86	-1.62	1.49	2.47	4.58	0.89				2549.4
49	21.64	-1.20	2.00	2.84	4.62	0.87				2456.3
50	21.79	-1.79	2.20	3.13	4.76	0.89				2705.6

*Coefficient not significant at the 10% level, two-tailed test.
**Sample sizes are provided in Table I.
+These estimates should be interpreted cautiously since the data are truncated near (or below) the estimated mean for this education group.
++E was fixed in these runs because unconstrained estimates of E were implausibly high (see Rodriguez and Trussell, 1980, and Trussell and Bloom, 1983).

FIG 1 - MEAN AGE AT FIRST BIRTH

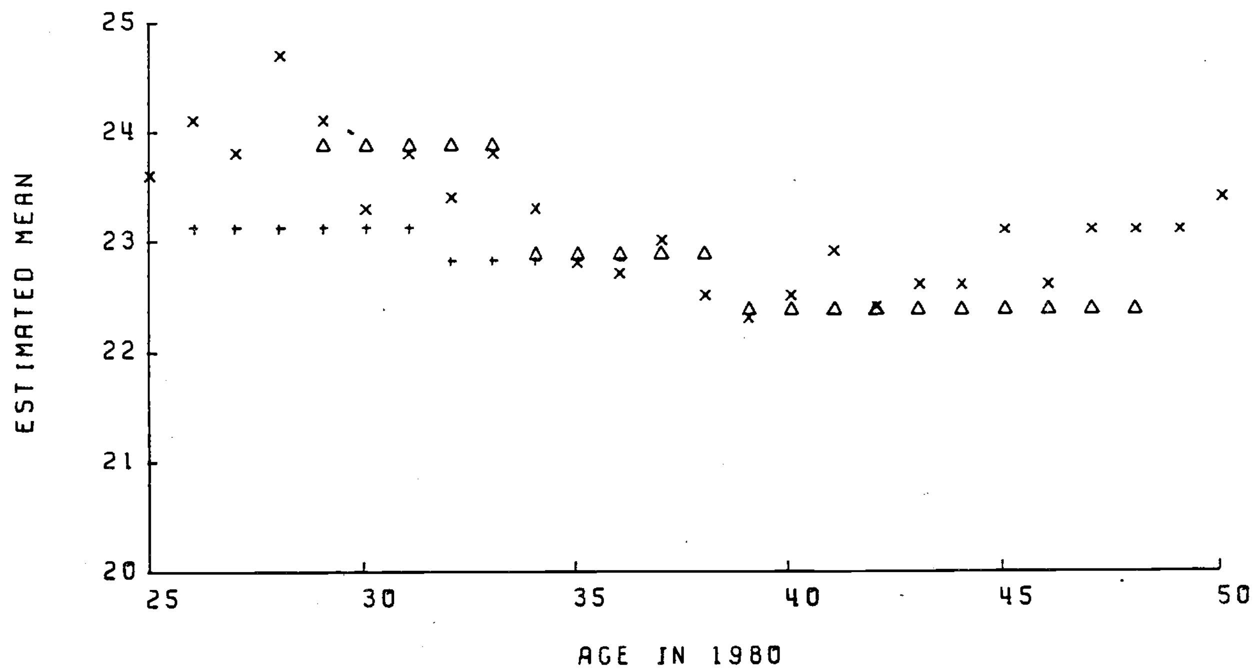
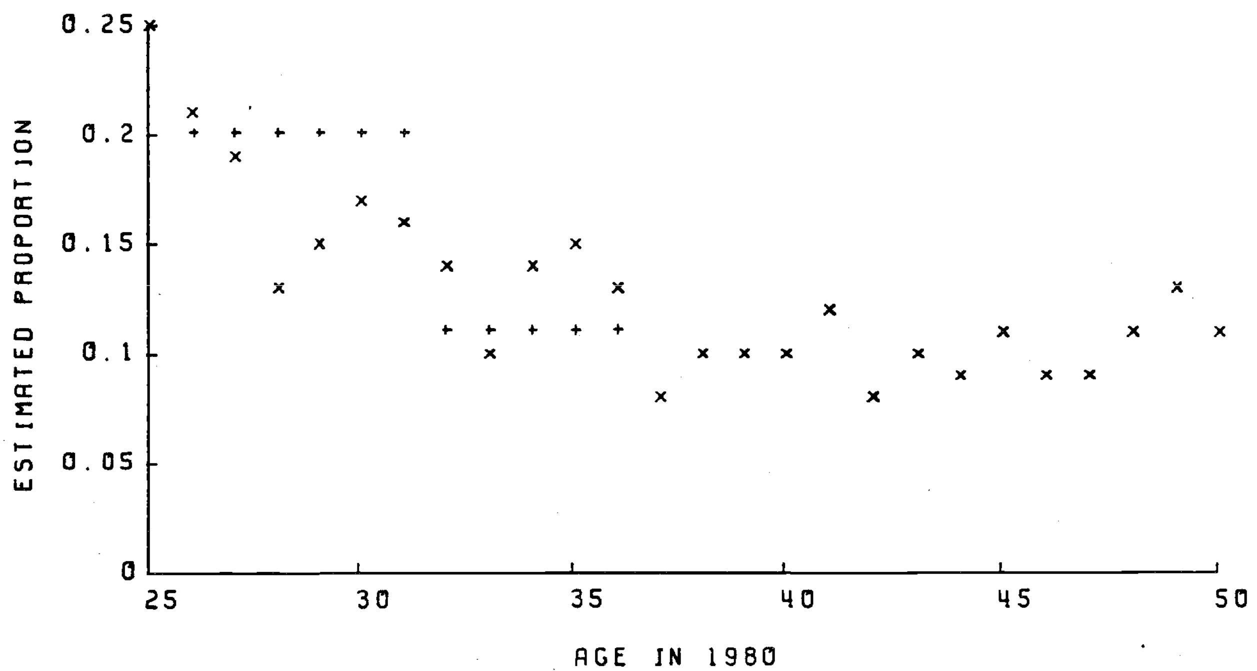


FIG 2 - PROPORTION CHILDLESS



x x x x x 1980 CPS

\+ + + + + 1978 NLS

Δ Δ Δ Δ Δ 1976 NSFG

Source: Table I

Note that results for the 1976 NSFG and 1978 NLS have been translated to apply to cohort ages in 1980.

FIG 3 - FACTORS AFFECTING MEAN AGE AT FIRST BIRTH

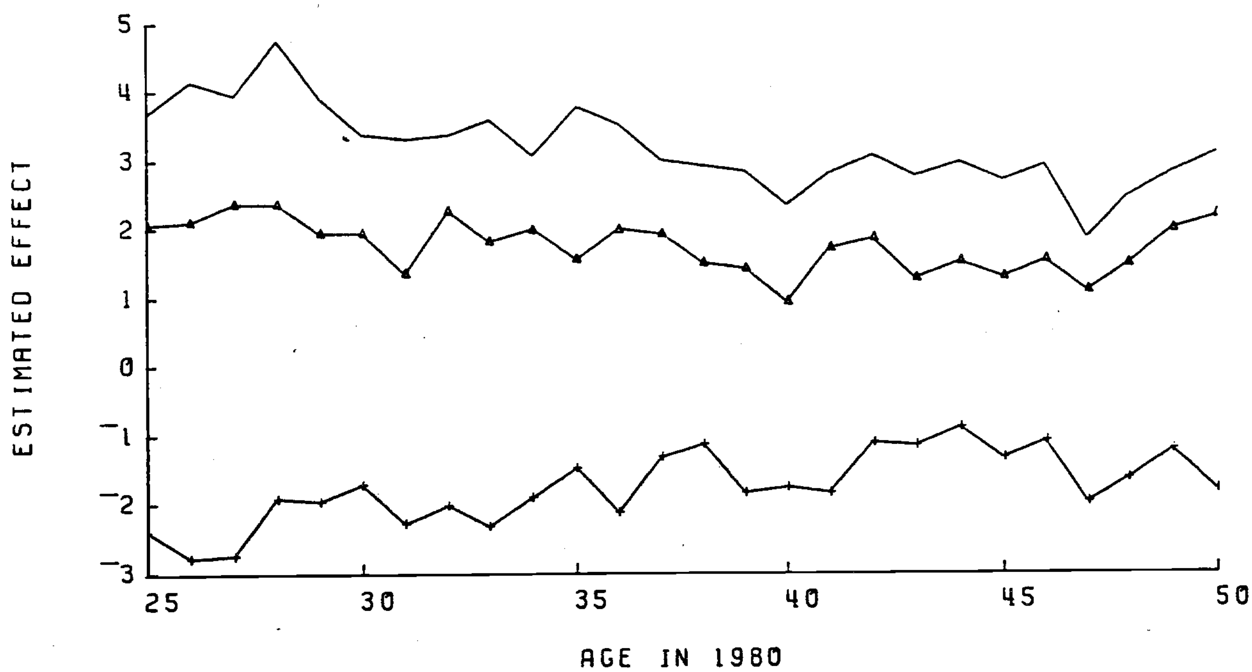
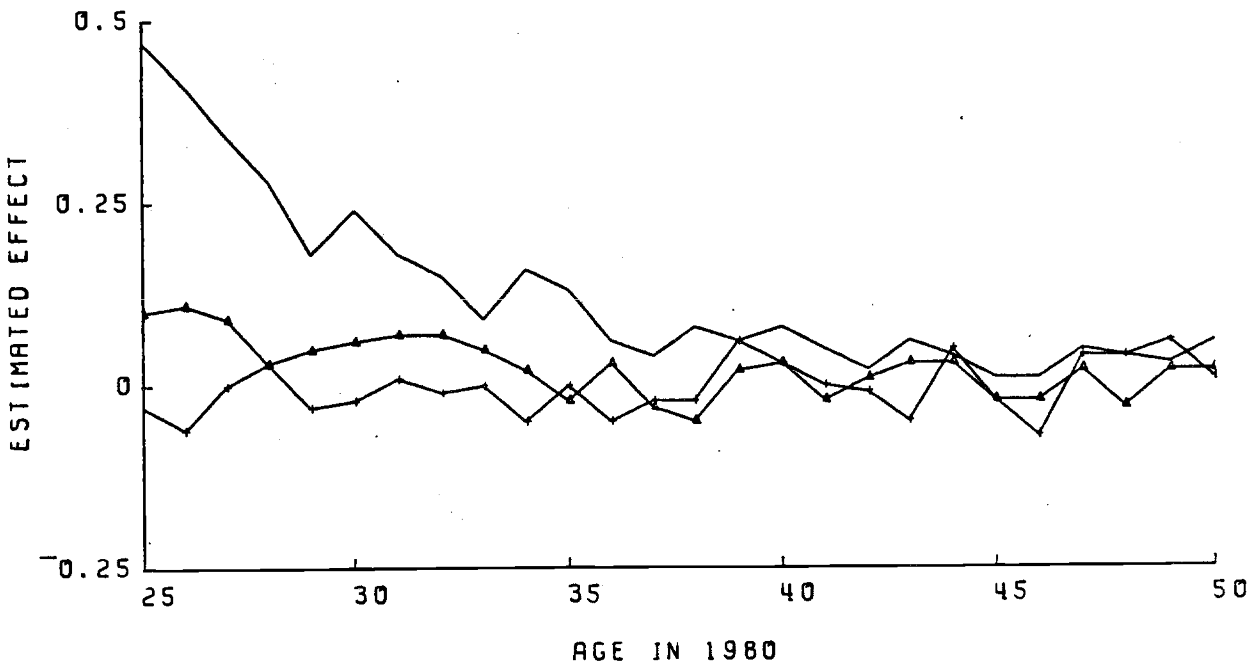


FIG 4 - FACTORS AFFECTING PROPORTION CHILDLESS



————————	Education >HS
Δ Δ Δ Δ Δ	Education = HS
+ + + + +	Race = Black

Source: 1980 CPS, Table IV, top panel