NBER WORKING PAPER SERIES

MARRIAGE PATTERNS IN THE UNITED STATES

David E. Bloom

Neil G. Bennett

Working Paper No. 1701

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
September 1985

An earlier version of this paper was presented at the 1984 Annual Meetings of the Population Association of America in Minneapolis. The authors are indebted to Clint Cummins and Cecilia Rouse for expert computational assistance. 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 #1701
September 1985

Marriage Patterns in the
United States

ABSTRACT

This paper analyzes cohort marriage patterns in the United States in order to determine whether declining rates of first marriage are due to changes in the timing of marriage, the incidence of marriage, or both. Parametric models, which are well-suited to the analysis of censored or truncated data, are fit separately to information on age at first marriage derived from three data sets which were collected independently and at different points in time. Extended versions of the models are also estimated in which the parameters of the model distributions are allowed to depend on social and, economic variables. The results provide evidence that the incidence of first marriage is declining and that there is only a slight tendency for women to delay marriage. In addition, education is the most important correlate of decisions about the timing of first marriage whereas race is the most important correlate of decisions about its incidence.

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

Neil G. Bennett
Department of Sociology
Yale University
New Haven, Connecticut 06520

1. Introduction

Since the late 1960s, the rate of first marriages experienced by individuals aged fourteen and over has declined substantially in the United States (see Figure I). This pattern, which has been characteristic of both men and women and has been quite steady over time, has contributed to the increasing proportion of single young adults in the population. According to some researchers, these facts reflect changes in the timing of marriage, and not changes in its ultimate incidence. For example, according to Cherlin (1981, p. 11), "The higher proportion of single young adults in the 1970s and early 1980s suggests only that they are marrying later, not foregoing marriage. It is unlikely that their lifetime proportions marrying will fall below the historical minimum of 90 percent." Indeed, as Figure II shows, the median age at first marriage increased by more than one year for both males and females during the 1970s.

On the other hand, other researchers such as Becker (1981) present theoretical models which suggest that the recent trends are primarily reflective of changes in the incidence of marriage since the rising economic status of women leaves them with less incentive to enter traditional marriages. These researchers are also quick to point out that a secular increase in the median age at first marriage is consistent with a decline in the proportion of individuals who ever marry, and not only with the phenomenon of delayed marriage.

Implicit in both of these views are projections of the future time series of marriage rates. For example, if marriage rates have declined mainly because of an increasing tendency to delay marriage, the rates should soon begin to rise as the delayers reach their desired ages of first marriage. Alternatively, if the decline is mostly the result of an increasing proportion of women deciding to (or, by default, just happening to) forego marriage, then marriage rates will tend to remain depressed in the future.

The purpose of this paper is to analyze recent nuptiality patterns in the United States in an attempt to distinguish between these alternative views of recent marriage trends. We do this by using a parametric model to analyze survey data on age at first marriage for successive birth cohorts. Because the model is parametric, it allows us to project marriage rates for cohorts which have yet to complete their first marriage experience and to thereby estimate their mean age at marriage and the proportion ultimately marrying. We also estimate an extended version of this model in which the parameters are allowed to depend on social and economic variables such as race and education. In this way, we investigate the correlates of the timing and incidence of marriage for a succession of birth cohorts.

Section II provides a brief description of the parametric model we use to represent the underlying pattern of age at first marriage; this section also discusses both the extension of the model to incorporate covariate effects and maximum likelihood estimation from truncated and non-truncated data sets. Section III describes the three data sets used in this study. Section IV presents and discusses the results of fitting various specifications of the model to cohort data in each of these data sets; this section also examines the sensitivity of our results to the degree of censoring. Section V discusses our results and comments on their implications for the evolution of nuptiality patterns in the United States.[1]

II. The Model

In 1971, Ansley Coale observed that age distributions of first marriages are structurally similar in different populations. As shown by Coale, these distributions tend to be smooth, unimodal, skewed to the right, and have density close to zero below age fifteen and above age fifty. Coale also observed that the differences in age-at-marriage distributions across female populations are largely accounted for by differences in their means, their standard deviations, and their cumulative values at the older ages, e.g., age fifty. As a basis for the application of these observations, Coale constructed a standard schedule of age at first marriage using data from Sweden, 1865-69.

Coale and McNeil (1972) subsequently developed a closed-form expression which closely replicates the reference distribution presented by Coale (1971):

$$g_s(x) = 0.1946\exp\{-.174(x-6.06)-\exp[-2.881(x-6.06)]\} \quad . \qquad (1)$$

This function can be related to any observed distribution by adjusting its location and dispersion, and its cumulative value as $x \to \infty$. The particular form of the model that we shall use, which characterizes any observed distribution, was derived by Rodriguez and Trussell (1980):

$$g(a) = \frac{E}{\sigma}1.2813\exp[-1.145(\frac{a-\mu}{\sigma}+0.805)-\exp\{-1.896(\frac{a-\mu}{\sigma}+0.805)\}], \qquad (2)$$

where g(a) is the proportion marrying at age a in the observed population and μ, σ, and E are, respectively, the mean and standard deviation of age at marriage (for those who ever marry) and the proportion ever marrying.

It is interesting to note that Coale and McNeil's model distribution of first marriage by age (e.g., equation (1)) arises as the convolution of an infinite number of mean-corrected exponential distributions whose parameters increase in arithmetic sequence. Moveover, Coale and McNeil have shown that this distribution is very closely approximated by the convolution of the three exponential distributions with the largest exponents (in the infinite sequence) and a normal distribution. This latter property of the Coale-McNeil model gives rise to an appealing behavioral interpretation of the model. According to this interpretation, each of the three exponential distributions characterizes the waiting time between two premarital stages (i.e., between the commencement of dating and ultimately meeting one's spouse, between meeting the spouse and engagement, and between engagement and marriage); the norm.1 distribution describes the age of entry of women into the marriage market. This interpretation received some empirical support in the original paper by Coale and McNeil in a direct test using data on the length of time that a sample of French husbands and wives knew each other before marrying. Subsequent research, however, has done little to confirm or deny the behavioral interpretation of the model. Nevertheless, a number of studies have provided additional support for the ability of the model to fit first marriage data (see, e.g., Ewbank, 1974; Rodriguez and Trussell, 1980; Trussell, 1980; and Trussell and Bloom, 1983). To some extent, the good fit my be due to the flexihility of three-parameter models to fit distributions that are smooth, unimodal, and skewed to the right. It is also likely that the Coale-McNeil model performs well because it is based on the marriage rates for an actual population. In other words, even though the true model generating a given distribution of marriage rates is unknown, the Coale-McNeil model may fit well (and better than a purely theoretical model such as that due to Hernes (1972) or a purely ad hoc empirical model such as that due to Keeley (1979)) because the true model is captured implicitly in the rates on which it (i.e., the Coale-McNeil model) is based.2

The parameters of equation (2) may be estimated in a variety of ways depending on the nature of the available data (see Rodriguez and Trussell (1980) for further details). In the present application we shall work with survey data on age at first marriage for individual women and will use a maximum likelihood estimator. Thus, for a sample of all-women (i.e., a random sample of ever-married and never-married women in some population or cohort), we will estimate μ, σ, and E by maximizing the following log likelihood function:

$$\log L_A = \sum_{i \varepsilon M} \log g(a_i^m \mid \mu,\sigma,E) + \sum_{i \varepsilon \overline{M}} \log[1-G(a_i^s \mid \mu,\sigma,E)], \qquad (3)$$

where a_i^m is age at first marriage for those individuals who have married (the set M), a_i^s is age at the time of survey for never-married individuals (the set $\overline{M}$), and $G(\cdot)$ is the cumulative distribution function for the density function $g(\cdot)$ expressed in equation (2). Observe that the second summation on the right hand side of equation (3) accounts for censoring which will be present to the extent that not all women who ultimately do marry will have done so by the time of the survey.

Alternatively, for a sample of ever-married women we employ a conditional likelihood function which is constructed from the likelihoods of individuals' marrying at particular ages (a_i^m) conditional on their having married by their age on the date of the interview:

$$\log L_{EM} = \sum_{i \in M} [\log g(a_i^m) - \log G(a_i^s)]. \tag{4}$$

Observe that this function does not depend on the parameter E, because E is a proportionality factor in both $g(\cdot)$ and $G(\cdot)$ and therefore cancels when the conditional likelihood (i.e., $g(a_i^m)/G(a_i^s)$) is expressed. This formulation therefore corrects for the truncation of never-married women, although the parameter E is not estimable from the truncated data.

Following Trussell and Bloom (1983), we extend this model to allow for covariate effects by specifying a functional relationship between the parameters of the model distribution and a set of covariates. For example, we may specify these relationships in linear form as follows:

$$\mu_i = X_i'\alpha$$

$$\sigma_i = Y_i'\beta$$

$$E_i = W_i\gamma ,$$

where 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 α, β, and γ, are the associated parameter vectors to be estimated. Because of the model's inherent nonlinearity, the parameters are identified even if all of the covariate vectors are the same. Standard statistical tests (t-tests and likelihood ratio tests) can, however, be used to assess the validity of different exclusion restrictions (e.g., $\sigma_i = \sigma$ and $E_i = E$ for all i).[3]

All of the maximum likelihood estimates presented in this paper were computed using the Davidon-Fletcher-Powell routine contained in the numerical optimization package GQOPT. This routine is described in Goldfeld and Quandt (1972, pp. 5-P).

III. The Data

As noted in Section I, this study uses three independent data sets to investigate the marriage patterns of American women. The use of multiple data sets is prompted by the fact that no single data set is uniquely well-suited to the tasks at hand. In addition, we feel that the consistency of results derived from different sources of information, collected at different points in time, is an important indication of their strength. The remainder of this section provides descriptions of the three data sets.

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

Cycle II of the NSFG was conducted in 1976 by the National Center for Health Statistics through personal interviews with 8,611 women aged 15-44 years. For the purposes of this study, the NSFG is useful because it provides data on a representative sample of ever-married women with information on age at first marriage along with several socioeconomic variables which presumably influence age at marriage. These variables and the coding scheme adopted for them are: race (black or non-black), religion (Catholic or non-Catholic), childhood residence (rural or urban), and education at time of survey (less than high school, high school, greater than high school). All women aged 20-44 at the time of the survey who first married between ages 12 and 44 are included in our data file. Because we do not have information on a representative sample of never-married women, we cannot estimate the parameter E (i.e., the proportion ever-marrying) from this sample; nor can we estimate its covariates. However, as discussed above, consistent estimates of the parameters μ and σ and their covariates can still be computed provided this sample selection rule is explicitly incorporated in the estimation procedure (which we do when we analyze the NSFG data). Observations were counted more or less heavily depending on their sample weights, with the weights adjusted to have mean unity.

B. Current Population Survey (CPS)

The CPS is a nationwide sample survey conducted monthly by the Bureau of the Census. It involves detailed personal interviews in about 70,000 households during 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, 1982 CPS, the normal set of questions was supplemented with a. set of retrospective marital history questions. Included on the supplementary survey instrument was a question on age at first marriage which was asked of all women aged 18-75. Unfortunately, there are few retrospective covariates in the CPS that could sensibly be hypothesized to affect age at marriage. However, we have constructed the following two variables: race (black, non-black) and education at time of survey (less than high school, high school, greater than high school).

Although the CPS data set permits estimation of only 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, and (c) it is very recent. As with the NSFG, sample weights were used in creating this data file after adjusting them so they average to one.

C. National Longitudinal Survey (NLS) of Young Women

This survey has been conducted annually since 1968 when it started with 5,159 women aged 14-24. The main purpose of this survey has been to gather longitudinal information on a wide range of socioeconomic variables for use in analyzing the labor market experiences of young women. In 1978, a complete reinterview of the original sample of women was conducted which included a question on age at first marriage. We have used this information, along with information on number of marriages and marital status from each prior wave of this survey (for ever-married women who failed to report their age at first marriage in 1978), to construct a data set on age at first marriage for women aged 24-35 in 1978.[4]

In comparison to the NSFG and CPS data, the NLS data are more useful because they contain information on two socioeconomic variables relevant to a study of age at first marriage that are not available in the 'other two data sets. This information refers to the occupation of the respondent's father (i.e., blue collar or not) and to the structure of the respondent's family when she was age 14 (i.e., both parents present or not). In addition, the NLS includes information on 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). The NLS does, however, have several weaknesses in that it does not include information about religion, it has a smaller sample size than the CPS, and it refers to a narrower group of ages than either the NSFG or the CPS. The NLS data may also be a nonrepresentative sample of the population because of sample attrition, although the 1978 file includes information on 79 percent of the original NLS participants.

IV. Results

A. Estimates Computed Without Covariates

We first fit the Coale-McNeil model without covariates to data from the NSFG, CPS, and NLS in order to ascertain the general trends in marriage patterns across cohorts. The fact that we do not include covariates in the estimation procedure implies that we treat the parameters μ, σ, and E as constants, i.e., μ, σ, and E are not allowed to depend on individual characteristics.

These preliminary results derived from the three data sets are presented in Table I. Since the NSFG and CPS data were collected at points in time six years apart, we have created age groups such that cohorts can be followed over the six-year period. Our confidence in the estimates of μ and σ would be enhanced if the estimates were similar for each cohort across data sets.

Despite conventional wisdom that age at first marriage has been increasing dramatically in this era of increased labor force participation and careerism among women, we find that the mean age at first marriage has remained quite stable across cohorts. Results from the NSFG indicate that μ has increased by only one year over cohorts born 20 years apart. Estimates of μ derived from the CPS and NLS indicate perhaps even smaller increases over time.

In comparing the results from the NSFG and the CPS we see a remarkable consistency in the estimates of μ and σ. Estimates of σ are very similar across data sets and are essentially constant across cohorts. Estimates of μ are virtually identical between data sets for the younger cohorts and diverge slightly for the older cohorts. This divergence may be accounted for in part by sampling variability since the estimated standard error of μ for the 46-50 year olds in the CPS is .07 and for the 40-44 years olds in the NSFG, .11).

Another prominent feature of Table I are the estimates of E derived from the CPS data. The parameter E cannot, of course, be estimated from the NSFG because the sample consists only of ever-married women. From the CPS results, however, we can clearly see a monotonic, downward trend across cohorts in the proportion of women who will ultimately marry. It appears that only 4 percent of those who were in their late 40's in 1982 will never marry, whereas as many as 13 percent of those in their late 20's will remain unmarried. The estimates of E also decline across the two NLS cohorts. Moreover, they are extremely close to the CPS estimates for roughly the same cohorts.

It should be emphasized that the strong agreement among the results derived from the three data sets-points toward the overall robustness of the estimates. The fact that different data were used, obtained at different points in time, and that estimates were derived using somewhat different models (i.e., with regard to the different likelihood functions employed in order to account for the differing nature of the two samples -- one of all women and one of only ever-married women) adds to our overall confidence in the parameter estimates. Of course, it is possible that the model fits the data poorly, hut in roughly the same way across data sets. To examine this possibility, we have calculated observed marriage rates by age for the four oldest cohorts in the CPS data and have plotted these in Figure III in relation to the estimated models. Although the models tend to underpredict the proportion of marriages occurring at the modal age at marriage, they do correspond to the data quite closely, and especially in the tails of the distributions. Thus, it appears that equation (2) does indeed provide a satisfactory fit to the data.

B. Estimates Computed with Covariates

We now introduce covariates into the specification of μ. Table II, which reports estimates computed from the 1976 NSFG, reveals that the effects of three covariates--Black, Catholic, and Rural--are statistically significant, but substantively trivial. In contrast, the impact of education on age at first marriage is substantial. Women who are high school graduates with no further education marry approximately two years later, on average, than women with less education (controlling for race, religion, and childhood residence). The mean age at first marriage of women with education beyond high school is nearly four years higher than that of women who are not high school graduates.

Table III reports parameter and hyperparameter estimates computed using data from the 1982 CPS. In the first model we estimate, we allow μ to depend on covariates while σ and E are assumed to be constant across all individuals in the sample. The results generated by fitting this first model are qualitatively similar to those obtained using data from the NSFG. Given that a women eventually marries, whether she is black has little bearing in and of itself on when she marries. However, educational attainment has a significant positive effect on age at msrriage.[5]

In addition to allowing μ to vary among subgroups of the population, in our second model we permit E to depend on covariates. The striking result is that while μ is strongly associated with educational level, E is not: With the exception of the most recent cohorts, education is either statistically insignificant in its association with proportions ever-marrying or the association is of small enough magnitude to be of little substantive interest.

On the other hand, among younger cohorts of women, race is strongly correlated with the probabilty of ever-marrying. The correlation was minimal for the oldest cohort in our study. Thus, race has been of increasing importance in differentiating those who will marry from those who will not.

Results derived from the 1978 NLS, as shown in Table IV, while somewhat difficult to compare with the results in Tables II and III, reveal patterns of nuptiality that are similar to those revealed by estimates computed from the other data sets. Education is positively associated with the age at which a woman marries and negatively associated with her propensity to marry. Here, too, we see that being black is associated with a substantially reduced probabilty of ever-marrying, yet has a trivial association with the age at which one marries (conditioned upon marrying).[6]

The additional variables incorporated in the more complex model are substantively (and in many cases, statistically) insignificant. For example, it appears that variables which measure parental background (e.g., the blue collar - white collar variable) are minor factors in the determination of age at marriage. Moreover, the fact that the marriage of a woman's parents dissolved sometime prior to her adolescence has only a small negative association with the age at which she marries. We might speculate that, under the (debatable) assumption that young women with single parents find their home life less pleasant than it would he otherwise, these women are motivated to leave home earlier than their counterparts who have parents with intact marriages. Traditionally, marriage has been one mechanism by which a woman can leave home at an early age, although marriage is increasingly less necessary in recent years for her to do so.

C. Sensitivity Analysis

Since many of the parameter estimates reported in Tables II-IV are computed from data that are either truncated, censored, or both, their reliability is heavily dependent on the statistical structure which we have imposed on the data. To some extent, the underlying structure is supported by the reasonably close fits of the model to the data as shown in Figure III. The closeness of the parameter and hyperparameter estimates derived from different datasets collected at different points in time provides further support for the model. However, one additional test of the adequacy of the model seems appropriate and has been conducted. This test essentially involves censoring information on first marriages that took place in the last ten years of the data and fitting the model to the artificially censored data to see how well the estimates predict actual experience.

We have carried out this experiment using the CPS data. Estimates were computed for cohorts aged 36-40, 41-45, and 46-50 in 1982, using data on their first marriage experience as of 1972. These estimates are reported in Table V for what we found to be the strictest form of this test of the model: estimates of the extended version of equation (2) which allow for covariate effects.

The estimates in Table V may be compared to those presented in the last three columns of Table III. In general, the estimates computed for individual cohorts as of 1972 tend to be quite close to those computed using ten years of additional marriage experience. Indeed, the estimates of the covariate effects on μ and their standard errors are extremely close and would support identical substantive conclusions. The estimates of σ and of the covariate effects on E are a bit farther apart, although not seriously so given the small absolute magnitudes of what appear to be statistically significant differences. Thus, on balance, we believe the results of this test provide further support for the ability of the model to fit censored data.

V. Discussion and Conclusion

Changes in the marriage process can be decomposed into two distinguishable phenomena: changes in the timing of marriage and changes in its overall incidence. Period or cross-sectional data relating to these phenomena-whether first marriage rates, the proportion ever-married in a particular age group, or the mean age at marriage, for example--are often misleading in their implications. It would be desirable to interpret the various changes we find in marriage statistics as reflective of changes in life cycle patterns of women. In a period of potentially unstable nuptiality patterns, the only valid interpretations are those inferred from cohort-based data. In this analysis we have examined nuptiality patterns of cohorts of American women using data from the 1976 NSFG, the 1982 CPS, and the 1978 NLS. Implementing a parametric model, we can project the currently incomplete marriage experience of cohorts. In this fashion, we can estimate the mean age at marriage and the proportion ever-marrying in young cohorts today. Thus, we can resolve in good part some of the arguments in the literature concerning the current and future trends in the timing and incidence of first marriages in the United States.

We have found that age at first marriage has been quite stable across birth cohorts spanning twenty years. However, the proportions ever-marrying have changed substantially over time: The proportion of women ultimately never-marrying will be three times as high for women 26 to 30 years of age in 1982 as for those 46 to 50 in that year.

Several additional major findings emerge in our analysis when we fit an extended version of the nuptiality model to the three data sets. Educational attainment has a strong positive association with the age at which women marry, given that they marry. Further, higher education is increasingly negatively associated with the probability of ever-marrying among recent cohorts. In addition, race was found to be a large and increasingly important correlate of a woman's propensity to marry. For example, only 66 percent of black women aged 26 to 30 in 1982 who had not graduated high school can be expected to marry, as compared with 91 percent of their white counterparts.

Perhaps the most interesting conclusion that can be drawn from this analysis relates to the divergence of marriage and fertility patterns that is now underway. For example, Bloom (1982) and Bloom and Trussell (1984) report strong evidence of increasing permanent childlessness across cohorts of American women. Indeed, the gap between proportions never marrying and proportions never bearing children has increased from roughly 6 percent to 15 percent across the cohorts analyzed in this study. Thus, it appears that marriage is being displaced as the major choice variable used to control fertility, in favor of effective contraception and abortion. I" other words, marriage continues to be the major institution bringing substantial numbers of couples together, although childbearing appears to be declining in its importance as a motive for the formation of those unions.

Footnotes

1. All of our empirical efforts are focused on analyzing the marriage patterns of American women, as appropriate data for American men are of insufficient quality (see, e.g., Pendleton, McCarthy, and Cherlin, 1984). See Rodgers and Thornton (1985) for the results of an attempt to fit parametric models to survey data on age at marriage for men (and women).

2. Period factors, not modelled here, can worsen the fit of the model to the data and increase the variance of projection errors by generating irregularities in the uncensored portion of the first marriage distribution. However, period factors do not seem to be of substantial importance during the time period under consideration.

3. Trussell and Bloom (1983) and Sørensen and Sørensen (1984) research the use of proportional hazard and general hazard models for estimating the covariates of age at first marriage. However, hazard models are not used in this investigation because (1) these earlier studies provide no evidence that they fit marriage data better than the extended Coale-McNeil model, (2) hazard models cannot be fit to data for a sample of ever-married women, and (3) hazard models cannot be used to project the marriage experience of young cohorts.

4. Some of the 24-year olds in 1968 had reached age 35 by the time of the 1978 survey.

5. With the exception of education, all of the covariates used in this study are measures which refer to the time of first marriage. Education is defined as years of schooling at the time of the survey and not at the time of

the first marriage because we believe that the former measure is a (marginally) superior social indicator and because it can be constructed for all three data sets. In addition, in experiments conducted using the NLS data, we discovered that parameter estimates differed trivially using the two alternative measures of education. This finding should not be surprising since few women return to school after their first marriage and since, of those women who do return to school, very few shift across the broad educational categories which we have defined.

6. Cross-cohort comparisons of the estimated education effects may be somewhat biased by cross-cohort changes in mean educational attainment within the education categories we use. For example, in the 1982 CPS, mean years of education was (by definition) unchanged across the cohorts we analyze for the =HS category, but increased by 1.1 years for the <HS category and by .2 years for the >HS category. Thus, the modest increase in estimated education effects across cohorts is likely to underestimate the true increase since cross-cohort growth in educational attainment within the reference category exceeded that in the two other education categories.

References

Becker, Gary S. 1981. A Treatise on the Family. Cambridge: Harvard University Press.

Bloom, David E. 1982. "What's Happening to the Age at First Birth in the United States? A Study of Recent Cohorts." Demography 19: 351-370.

Bloom, David E. and James Trussell. 1984. "What are the Determinants of Delayed Childbearing and Permanent Childlessness in the United States?" Demography 21: 591-612.

Cherlin, Andrew. 1981. Marriage, Divorce, Remarriage. Cambridge: Harvard University Press.

Code, Ansley J. 1971. "Age Patterns of Marriage." Population Studies 25: 193-214.

______ and Donald R. McNeil. 1972. "The Distribution by Age at First Marriage in a Female Cohort." Journal of the American Statistical Association 67: 743-749.

Ewbank, Douglas. 1974. "An Examination of Several Applications of the Standard Pattern of Age at First Marriage." Unpublished Ph.D. Dissertation, Princeton University.

Goldfeld, Stephen M. and Richard E. Quandt. 1972. Nonlinear Methods in Econometrics. North Holland Publishers: New York.

Hernes, Gudmund. 1972. "The Process of Entry into First Marriage." American Sociological Review 37: 173-82.

Keeley, Michael C. 1979. "An Analysis of the Age Pattern of First Marriage." International Economic Review 20 (2): 527-544.

Pendleton, Audrey, James McCarthy, and Andrew Cherlin. 1984. "Assessing the Quality of Retrospective Marriage Histories: The June 1980 Current Population Survey." Unpublished paper.

Rodgers, Willard C. and Arland Thornton. 1985. "Changing Patterns of First Marriage in the United States." Demography, forthcoming.

Rodriguez, German and James Trussell. 1980. "Maximum Likelihood Estimation of the Parameters of Coale's Model Nuptiality Schedule from Survey Data." Technical Bulletin No. 7. World Fertility Survey: London.

Sørensen, Annemette and Aage B. Sørensen. 1985. "An Event History Analysis of the Process of Entry into First Marriage." in Kertzer, D.I., Ed., Family Relations in Life Course Perspectives, New York: JAI Press.

Trussell, James. 1980. "Illustrative Analysis: Age at First. Marriage in Sri Lanka and Thailand." Scientific Reports No. 13. World Fertility Survey: London.

Trussell, James and David E. Bloom. 1983. "Estimating the Covariates of Age at Marriage and First Birth." Population Studies 37 (3): 403-416.

Table I

Estimates of the Coale-McNeil Model Without Covariates*

Data set	Cohort	μ	σ	E	$-\ln L$	N	P(EM)
NSFG (1976)	20-24	21.80 (.24)	4.30 (.18)	-	2707.5	1359	(1.0)
	25-29	21.49 (.11)	3.79 (.09)	-	4304.1	1837	(1.0)
	30-34	21.16 (.11)	4.08 (.09)	-	4180.6	1622	(1.0)
	35-39	20.89 (.11)	3.95 (.09)	-	3574.2	1363	(1.0)
	40-44	20.80 (.11)	4.02 (.09)	-	3447.4	1303	(1.0)
CPS (1982)	26-30	21.81 (.07)	4.28 (.06)	.868 (.006)	17226.8	5532	.797
	31-35	21.63 (.06)	4.08 (.05)	.917 (.004)	17099.1	5776	.900
	36-40	21.33 (.06)	4.09 (.05)	.942 (.003)	13603.1	4681	.938
	41-45	21.20 (.06)	4.02 (.05)	.952 (.003)	11424.0	3969	.951
	46-50	21.39 (.07)	4.13 (.06)	.958 (.003)	10727.1	3703	.957
NLS (1978)	24-29	21.26 (.11)	3.77 (.09)	.918 (.009)	5597.2	2344	.837
	30-35	20.98 (.09)	3.43 (.07)	.948 (.006)	4426.3	1756	.937

*Estimated standard errors are reported in parentheses below parameter estimates.

μ represents the cohort's mean age at first marriage;
σ represents the standard deviation of age at first marriage for the cohort;
E represents the proportion of women in the cohort who ever marry;
P(EM) is the observed proportion of the cohort (of size N) who are ever-married at the time of the survey.

Table II

Estimates of the Coale-McNeil Model with Covariates: NSFG (1976)**

	Variable	20-24	25-29	30-34	35-39	40-44
μ	Constant	18.22 (.12)	18.37 (.12)	19.01 (.13)	19.02 (.15)	19.23 (.16)
	Black	-.31* (.16)	.03* (.15)	-.06* (.19)	.03* (.23)	-.83 (.22)
	Ed=HS	2.39 (.11)	2.30 (.12)	2.04 (.14)	1.89 (.16)	2.00 (.16)
	Ed>HS	3.74 (.13)	3.87 (.13)	3.84 (.16)	3.73 (.20)	3.11 (.20)
	Catholic	.24 (.11)	.53 (.10)	.75 (.14)	.16* (.16)	.29* (.15)
	Rural	-.23* (.12)	.06* (.12)	-.23* (.15)	.01* (.17)	-.16* (.16)
σ	constant	2.42 (.08)	2.64 (.06)	3.41 (.07)	3.35 (.08)	3.47 (.08)
	-an L	2378.3	3894.7	3893.4	3385.8	3292.1

*Coefficient not significantly different from zero at the .01 level.
**Estimated standard errors are reported in parentheses below parameter estimates.

Table III

Estimates of the Coale-McNeil Model with Covariates: CPS (1982)**

		COHORT									
	Variable	26-30	31-35	36-40	41-45	46-50	26-30	31-35	36-40	41-45	46-50
μ	Constant	20.04	19.90	19.98	20.02	20.16	20.01	19.90	19.97	20.02	20.16
		(.10)	(.10)	(.10)	(.10)	(.10)	(.10)	(.10)	(.10)	(.10)	(.10)
	Black	-.05*	-.26*	.13*	.11*	.02*	-.29	-.30	.12*	.11*	.02*
		(.12)	(.12)	(.13)	(.14)	(.14)	(.12)	(.12)	(.13)	(.14)	(.14)
	Ed=HS	1.23	1.34	1.14	1.08	1.36	1.27	1.34	1.14	1.09	1.36
		(.10)	(.10)	(.11)	(.11)	(.11)	(.11)	(.10)	(.11)	(.11)	(.10)
	Ed>HS	2.86	2.72	2.32	2.12	2.18	2.82	2.70	2.33	2.12	2.18
		(.11)	(.10)	(.11)	(.12)	(.12)	(.11)	(.11)	(.11)	(.12)	(.12)
σ	constant	3.96	3.78	3.91	3.85	3.92	3.92	3.77	3.91	3.85	3.92
		(.06)	(.04)	(.05)	(.05)	(.06)	(.05)	(.04)	(.05)	(.05)	(.05)
E	Constant	.863	.914	.941	.952	.957	.910	.949	.935	.956	.962
		(.006)	(.004)	(.003)	(.003)	(.003)	(.012)	(.008)	(.009)	(.008)	(.007)
	Black						-.252	-.141	-.087	-.075	-.030
							(.019)	(.016)	(.015)	(.015)	(.012)
	Ed=HS						.027*	.002*	.036*	.018*	.017*
							(.014)	(.010)	(.010)	(.008)	(.007)
	Ed>HS						-.074	-.051	.001*	-.014*	-.032
							(.015)	(.011)	(.011)	(.010)	(.010)
	-an L	16839.6	16712.5	13387.7	11260.8	10570.6	16685.1	16639.8	13346.2	11231.4	10548.1

*Coefficient not significantly different from zero at the .01 level.

**Estimated standard errors are reported in parentheses below parameter estimates.

Table IV

Estimates of the Coale-McNeil Model with Covariates: NLS (1978)

	Variable	24-29	30-35	24-29	30-35
μ	Constant	18.77 (.11)	19.14 (.13)	19.28 (.15)	19.66 (.17)
	Black	-.24* (.16)	.21* (.18)	-.16* (.16)	.40 (.18)
	Ed=HS	2.06 (.12)	1.71 (.14)	1.91 (.12)	1.57 (.14)
	Ed>HS	3.54 (.12)	3.02 (.15)	3.28 (.13)	2.76 (.15)
	Rural			-.29 (.09)	-.12* (.11)
	Dadblue			-.15* (.10)	-.41 (.12)
	BrokenHH			-.68 (.13)	-.51 (.15)
σ	Constant	2.90 (.06)	3.05 (.06)	2.85 (.06)	3.00 (.06)
[illegible]	constant	.976 (.010)	.976 (.009)	.978 (.011)	.976 (.009)
	Black	-.191 (.029)	-.117 (.027)	-.192 (.030)	-.117 (.026)
	Ed=HS	-.027* (.014)	.002* (.011)	-.031 (.014)	.001* (.011)
	Ed>HS	-.144 (.019)	-.056 (.014)	-.148 (.019)	-.056 (.015)
	$-\ell n$ L	5173.0	4190.2	5152.6	4178.3

*Coefficient not significantly different from zero at the .01 level.
**Estimated standard errors are reported in parentheses below parameter estimates.

Table V

Estimates of the Coale-McNeil Model with Covariates Computed from Artificially Censored Data: CPS (1982) Censored to 1972**

	Variable	Age in 1972: 26-30	31-35	36-40
μ	constant	19.84 (.10)	19.90 (.10)	20.00 (.10)
	Black	-.01* (.13)	.01* (.13)	-.04* (.14)
	Ed=HS	1.19 (.11)	1.10 (.11)	1.36 (.10)
	Ed>HS	2.31 (.11)	2.11 (.12)	2.16 (.12)
σ	constant	3.78 (.05)	3.72 (.05)	3.75 (.05)
E	constant	.964 (.009)	.963 (.007)	.962 (.006)
	Black	-0.100 (.017)	-.088 (.016)	-.038 (.012)
	Ed=HS	.046 (.010)	.020 (.008)	.017 (.007)
	Ed>HS	.054 (.010)	.008* (.009)	.001* (.057)
	$-\ln L$	11647.4	10608.5	10085.6

*Coefficient not significantly different from zero at the .01 level.
**Estimated standard errors are reported in parentheses below parameter estimates.

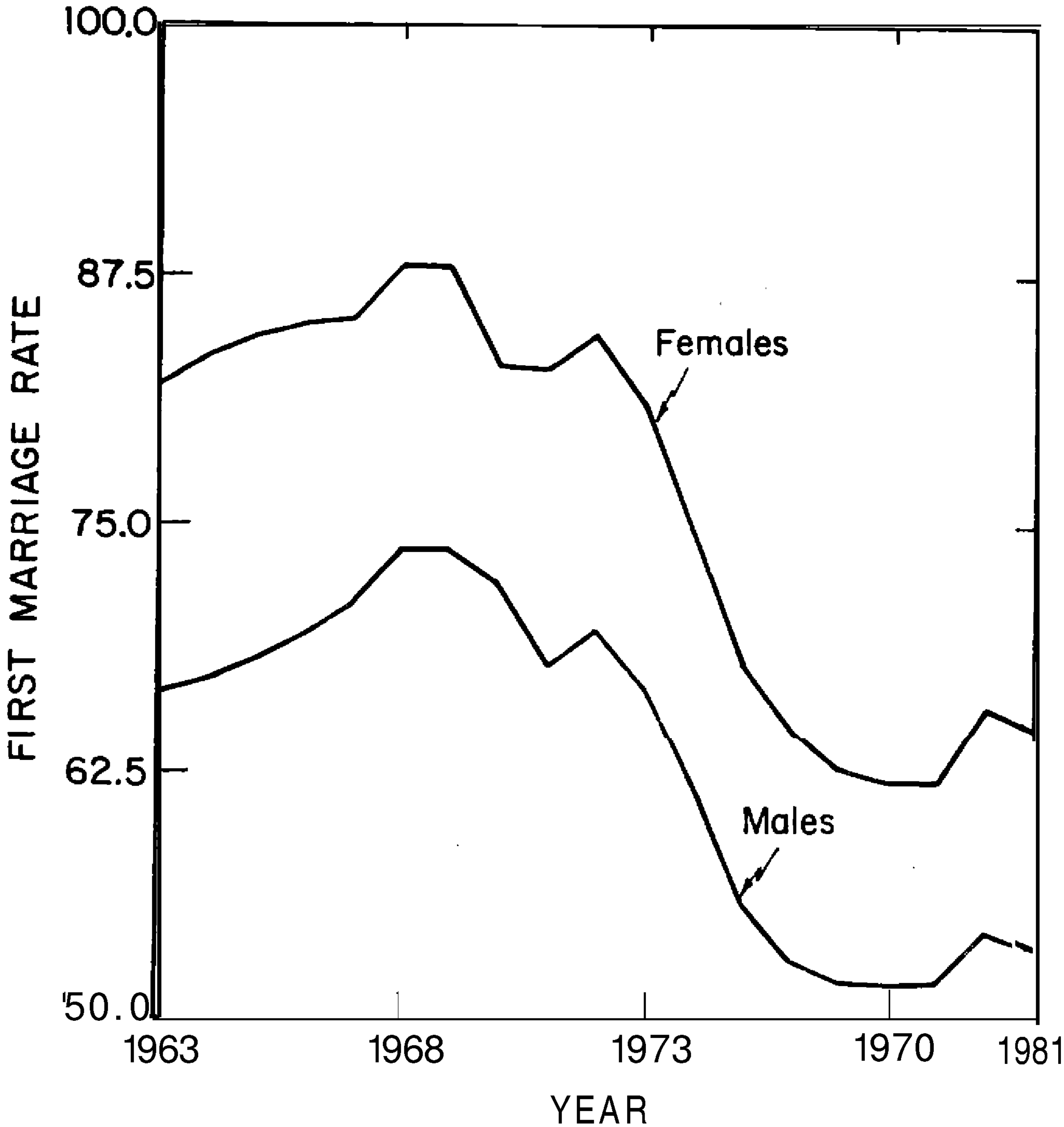


Figure I-First Marriage Rate by Sex, 1963-1981 (number of first marriages per 1000 never married individuals aged 14 and above, for each sex)

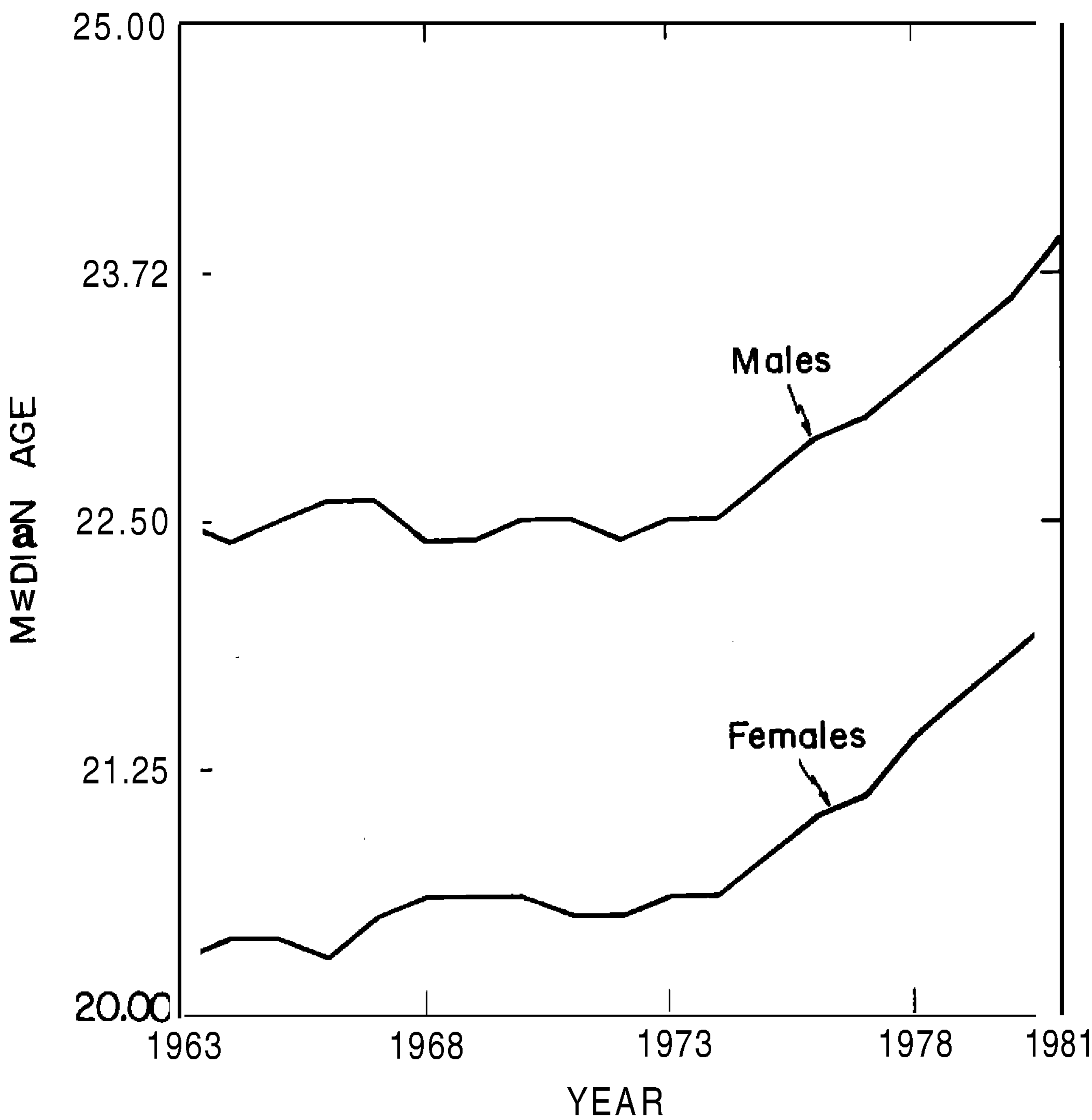
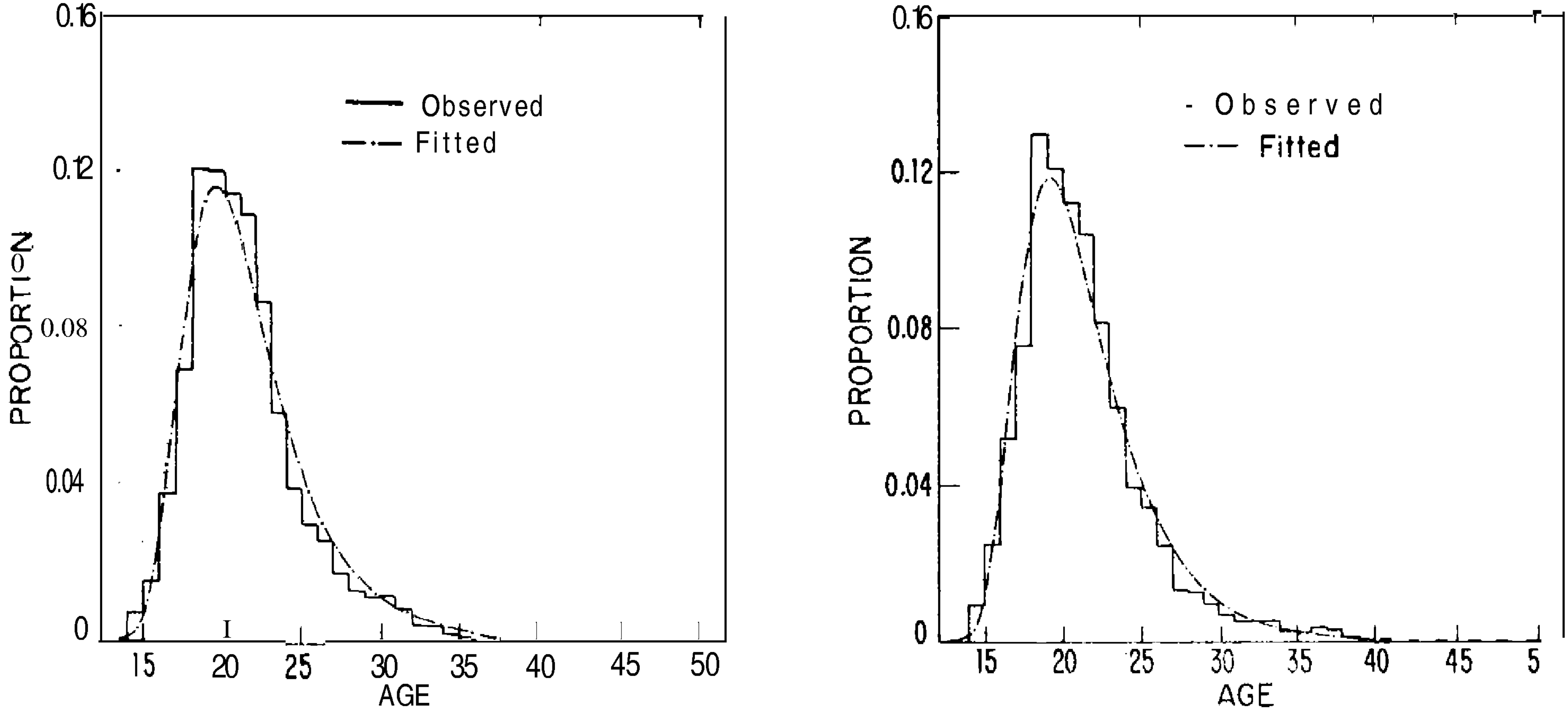


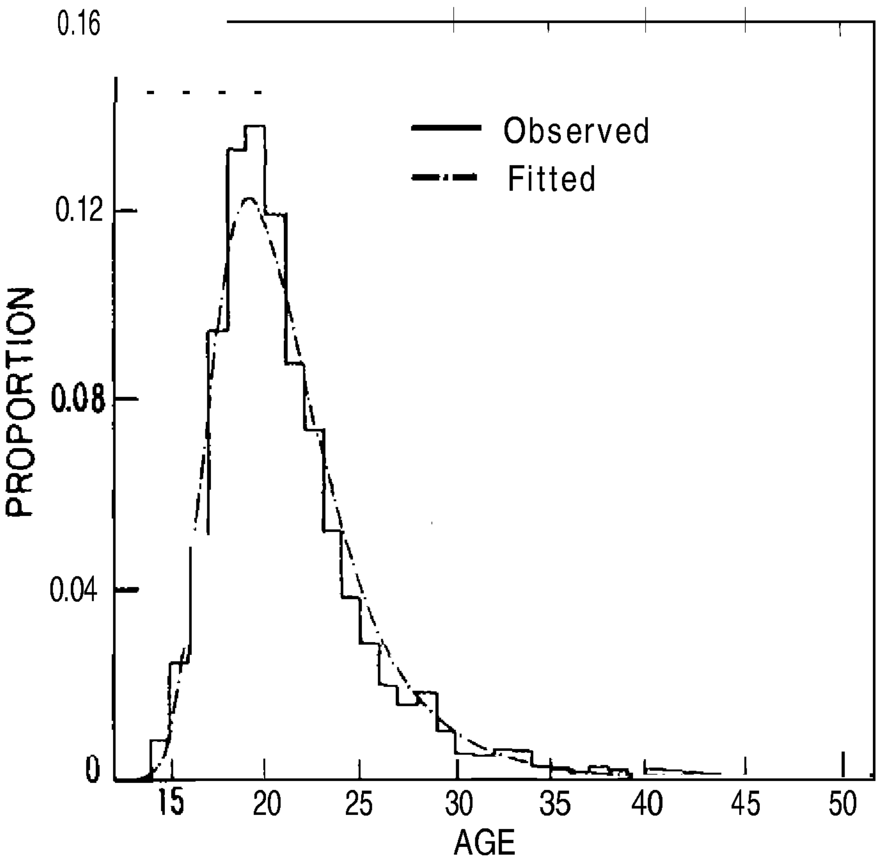
Figure II- Median Age at First Marriage by Sex, 1963- 1981



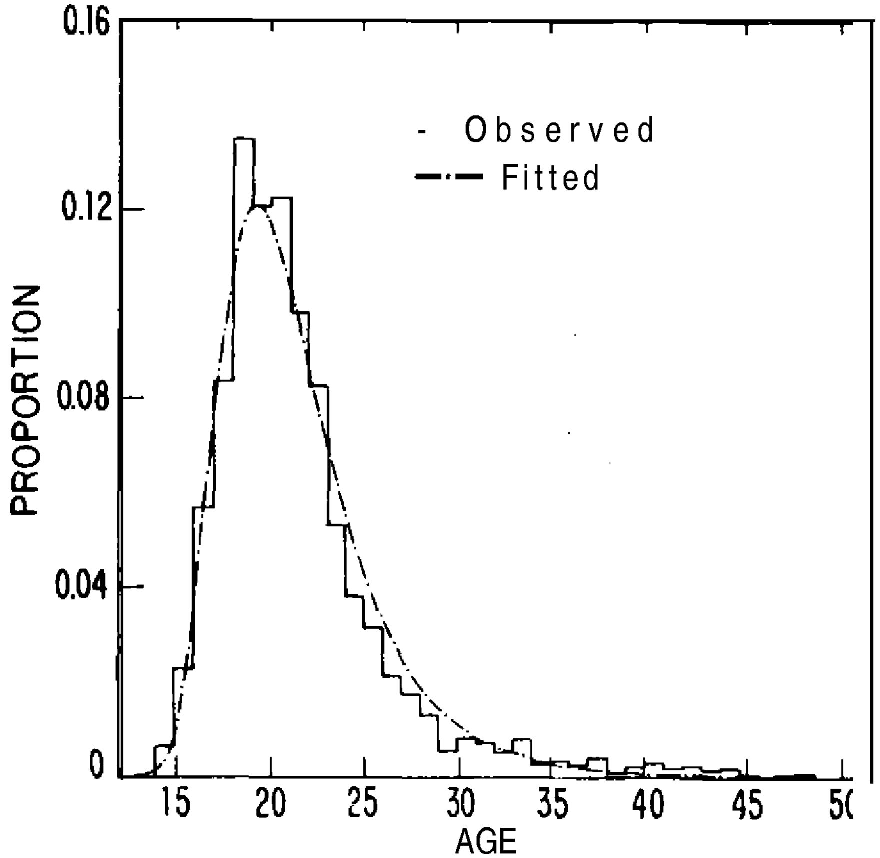
(a) **31-35** year olds

(b) 36-40 year olds

Figure III -- First **Marriage** Distributions, All Women, 1982 CPS



(cj 41-45 **year** olds



(d) 46-50 year olds

Figure III (continued) -- First Marriage **Distributions,** All Women, 1982 CPS