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MODELING AMERICAN MARRIAGE PATTERNS

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ABSTRACT

This paper investigates the application of the three-parameter, Coale-McNeil marriage model and some related hyper-parameterized specifications to data on the first marriage patterns of American women. Because the model is parametric, it can be used to estimate the parameters of the marriage process, free of censoring bias, for cohorts that have yet to complete their first marriage experience. Empirical evidence from three surveys is reported on the ability of the model to replicate and project observed marriage behavior. The results indicate that the model can be a useful tool for analyzing cohort marriage data and that recent cohorts are showing relatively strong proclivities to both delay and forego marriage. Consistent with earlier work, the results also indicate that education is a powerful covariate of the timing of first marriage and that race is a powerful covariate of its incidence.

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I. Introduction

Since the early 1970s, the rate of first marriages experienced by women aged fourteen and over has declined substantially in the United States (see Figure 1). This pattern, which has been characteristic of men as well, has been quite steady over time and goes hand in hand with the increasing proportion of young adults who are single in the population. According to some researchers, this trend reflects 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 the 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." Cherlin is joined in this speculation by the U.S. Bureau of the Census (see, e.g., Norton and Moorman, 1987), Glick (1984), Blau and Ferber (1986), and Bianchi and Spain (1986). Indeed, as Figure 2 shows, the median age at first marriage increased by more than one year for both males and females during the 1970s alone.

On the other hand, researchers such as Becker (1981) and Fuchs (1983) present theoretical arguments that suggest that the recent trends are potentially reflective of major 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 investigate a three-parameter model of the age distribution of women at first marriage. In particular, we study the application of the Coale-McNeil marriage model to survey data on the marriage patterns of successive birth cohorts of American women. Because the model is parametric, it can be used to compute estimates, which are free of censoring bias, of the mean age at marriage and the proportion ultimately marrying for cohorts that have yet to complete their first marriage experience. We also explore 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 examine the correlates of the timing and incidence of marriage for a succession of birth cohorts. Our empirical analysis, which is based on U.S. data that extend through the mid-1980s, also permits us to comment on the behavior that underlies the decline in overall U.S. marriage rates that began in the early 1970s.

Section II provides a 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 censored and non-censored data. Section III describes the data sets used in this study. Section IV addresses issues related to the use of sample weights in estimating the marriage model parameters. Section V presents and discusses the results of fitting various specifications of the model to cohort data in each of our data sets; this section also examines the sensitivity of our results to the degree of censoring. Section VI discusses our results and comments on their implications for the evolution of nuptiality patterns in the United States. We should note that all of our empirical efforts are focused on analyzing the marriage patterns of American women, as appropriate data for American men are of poor 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 first marriage for men (and women).)

II. The Model

Coale (1971) 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 15 and above age 50. Coale also observed that the differences in age-at-first 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, for example, age 50. As a basis for the application of these observations, Coale constructed a standard schedule of age at first marriage using data from Sweden, covering the period from 1865 to 1869.

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

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

This function can be related to any observed distribution by adjusting its location and dispersion, and its cumulative value as $x \rightarrow \infty$. 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 first 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 (i.e., equation (1)) arises as the convolution of an infinite number of mean-corrected exponential distributions whose parameters increase in arithmetic sequence. Moreover, Coale and McNeil showed that this distribution is 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 normal distribution describes the age at which women enter 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 closely replicate first marriage data (see, e.g., Ewbank, 1974; Rodriguez and Trussell, 1980; Trussell, 1980; Trussell and Bloom, 1983; and Grenier, Bloom, and Howland, 1987). Thus far, however, the suitability of the Coale-McNeil model to U.S. nuptiality data has not been carefully established.

To some extent, the success of the marriage model may be due to the flexibility 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.

Period factors, not modeled here, can worsen the fit of the cohort 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 periods to which our applications refer, as shown below in Section V.A. (Although we limit ourselves in this article to the analysis of a cohort marriage model, it should be noted that marriage patterns may also be usefully studied, under some circumstances and assumptions, using a standard cross-sectional life table approach.)

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:

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

where i denotes individual i, a_i^m is the age at first marriage for those individuals who have married (the set M), a_i^s is the age at the time of the 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 that will be present to the extent that not all women who ultimately do marry will have done so by the time of the survey.

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. We specify these relationships in linear form as follows:

$$\mu_i = X_i'\alpha$$

$$\sigma_i = Y_i'\beta$$

$$E_i = Z_i'\gamma \quad ,$$

where X_i, Y_i, and Z_i are the vector values of characteristics of an 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).

As a practical matter, it is useful to perform these statistical tests. For example, Trussell and Bloom (1983) tested and confirmed the validity – in both statistical and substantive terms – of constraining σ to be constant across covariate cells in their analysis of first marriage and first birth data from the World Fertility Survey for Colombian women. Bloom and Trussell (1984) subsequently reached the same conclusion in their analysis of first birth data for women in the United States. Fay (1986) noted, however, that imposing this constraint could lead to misleading results if applied to first marriage data for the United States.

Trussell and Bloom (1983) and Sørensen and Sørensen (1984) research the use of various proportional and general hazard models for estimating the covariates of age at first marriage. These studies provide no evidence that these hazard models fit marriage data better than the extended Coale-McNeil model (which can itself be viewed as a hazard model). In addition, they highlight the fact that an important class of hazard models – nonparametric hazard models – are not readily applicable in contexts in which one is interested in projecting marriage rates from censored data (which is effectively what one must do to estimate parameters of the marriage process that are not conditional on the censoring information).

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. 5-9).

III. The Data

We use three independent data sets to explore the usefulness of the marriage models described above and 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 data sets used to estimate the marriage model parameters were derived from the June 1976 and June 1985 waves of the Current Population Survey, as well as from Cycle III of the National Survey of Family Growth, conducted in 1982.

The CPS is a nationwide sample 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 1985 CPS, the normal set of questions was supplemented with a set of retrospective marital and fertility history questions. Included on the supplementary survey instrument was a question on age at first marriage that was asked of all women aged 18 and above. Unfortunately, there are few variables recorded in the CPS that could sensibly be hypothesized to be associated with age at marriage. However, we have coded the following two variables: race (black, white) and educational attainment at the time of the survey (less than high school, high school graduate, more than high school). Although the CPS data set permits estimation of only two covariate effects, it is extremely useful because it refers to a nationally representative sample of all women and because it includes an exceptionally large number of observations. The June 1976 CPS was constructed according to a design that was similar to that of the June 1985 CPS.

Cycle III of the NSFG, conducted by the National Center for Health Statistics, consists of 7,969 personal interviews of women aged 15 to 44 of all marital statuses in which ever-married women were asked their age at first marriage. We analyze the NSFG primarily as a check on the quality of the Census Bureau data.

IV. Weighting Considerations

None of the data sets we analyze was generated by simple random sampling. The NSFG and the CPS are both based on a multi-stage area-cluster design, with an oversampling of black women in the NSFG. The complexity of these sample designs raises three statistical issues about the use of these data sets for estimating the parameters of the likelihood function in equation (3) and its hyper-parameterized form:

(1) Non-independence within clusters in the selection of respondents: Maximum likelihood estimates of the parameters in equation (3) that ignore the non-independence problem are consistent — both for estimation with and without covariates. However, the estimated standard errors will tend to understate the true standard errors that one would obtain if one modeled the non-independence. In principle, this problem could be directly addressed if the full structure of the data (i.e., the identity of the area-clusters) were known. Unfortunately, neither the CPS nor the NSFG reports information on clusters. Nonetheless, since there is little *a priori* reason to believe that the intra-cluster correlation in age at first marriage is substantial, there is little reason to believe that the bias in the estimated standard errors will be large (see Scott and Holt, 1982, for an analysis of this problem in the context of ordinary least squares estimation of the linear regression model);

(2) The treatment of sample weights in the model specified without covariates: In the CPS, the sample weights primarily incorporate information on age, race, sex, and area-cluster. Sample weights in the NSFG incorporate information about age, race, area-cluster, and marital status. Since we are interested in estimates that generalize to the overall population, the weights are necessary in estimating the version of the likelihood function that does not include covariates. Hoem (1985) shows that this procedure (i.e., using weights that account for the probability of selecting a particular respondent and receiving a usable survey response, and in the case of the NSFG, a poststratification adjustment based on CPS data) will result in consistent estimates of the parameters and of the variance-covariance matrix. (We also performed all computations separately using weighted and unweighted data and found little difference among either the estimates of the parameters or their standard errors.)

(3) The treatment of sample weights in the model that is specified with covariates: Estimates of the covariate effects will be consistent regardless of whether the sample weights are used in the estimation. However, if the correctness of the model specification is considered to be part of the maintained hypothesis, then there is an efficiency loss associated with the use of the weights and the standard errors will not be consistent. On the other hand, if it is not rigidly assumed that the model specification is correct, the sample weights should be used if they include information not captured by the right-hand side variables (i.e., if they are based on different information or capture nonlinearities). (See DuMouchel and Duncan, 1983, for a discussion of this issue in the context of the linear regression model.)

Since our model controls for sex (i.e., we look only at females) and for age and race, the major pieces of information that could be added by the weights relates to area-cluster in the CPS data. Since there is little reason to think that location is a relevant piece of information in the CPS, there is no compelling reason to use the weights in the analysis of that data set. Nonetheless, we have examined this issue empirically by comparing estimates of covariate models fit with both unweighted and weighted data. In most cases we find small differences between the weighted and unweighted results. However, in some cases, the standard errors were significantly larger when computed from the weighted data. This finding suggests that the sample weights do contain important conditioning information. Thus, all of the covariate estimates we report are based on models that use the sample weights (normalized to average 1.0).

V. Results

A. Estimates Computed without Covariates

We first fit the Coale-McNeil model without covariates to data from the NSFG and CPS 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, that is, μ, σ, and E are not allowed to depend on individual characteristics.

Table 1 presents maximum likelihood estimates of the marriage model parameters based on data from the 1985 CPS and the 1982 NSFG. Since the NSFG and CPS data were collected at points in time three years apart, we have defined age groups such that cohorts are matched across the two data sets. Our confidence in the estimates of μ, σ, and E would be enhanced if the estimates were similar for each cohort across data sets.

The estimates imply that the sizable increase in the median age at first marriage over time (illustrated in Figure 2) is due partly to an increase in the mean age at first marriage across cohorts and partly to a decline in the proportion ever-marrying across cohorts. For example, results from the CPS indicate that the mean age at first marriage, μ, has increased by about one and one-half years over cohorts born an average of 15 years apart. At the same time, the proportion ever-marrying has decreased by about seven percentage points. The results also indicate that only five to six percent of those aged 45 to 49 in 1985 will never marry, whereas twice that proportion will remain permanently unmarried among those born 20 years later. These results are consistent with those reported in Schoen, Urton, Woodrow, and Baj (1985), which indicate that the proportion ever marrying declined across cohorts born between 1940 and 1950.

In comparing the results from the NSFG and the CPS we see generally a high degree of consistency in the estimates of μ, σ, and E. Estimates of σ are similar across data sets for all cohorts but for those aged 30 to 34 in 1985 and are roughly constant across cohorts until the youngest cohort. Estimates of μ are essentially identical between data sets (within a range that allows for sampling variability).

It should be emphasized that the strong agreement among the results derived from the two data sets points toward the overall robustness of the estimates. The fact that the data analyzed were collected using different sampling schemes at different points in time adds to our overall confidence in the parameter estimates. Of course, it is possible that the model fits the data poorly, but 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 Figures 3 through 6 in relation to the estimated models.

Generally, the models based on equation (2) appear to replicate the data quite closely, and provide a satisfactory fit to the tails of the distributions. The most notable discrepancy between the observed and projected values of the g(a) function relates directly to the fact that laws and norms in the United States seem to "interfere" with what may be termed a more natural progression of events in the dating—courtship—marriage process (assuming that the Coale-McNeil structure provides a reasonably good representation of that process). Note that the marriage rates tend to fall short of the rates implicit in our estimates of the model distribution at the modal ages at first marriage and tend to exceed them at the teenage years prior to the mode. One might speculate that this residual pattern arises because American society observes either laws or cultural dictates that hinder marriage before the threshhold age of 18. Ewbank (1974) offers a similar interpretation based on his analysis of historical data referring to first marriages in Sweden. We might choose to model explicitly this behavioral pattern, but we do not in the interest of parsimony.

To test for the importance of period effects during the years under study, we estimated a set of contemporaneous correlations between residuals for different pairs of the five birth cohorts identified in the lower panel of Table 1. Both absolute and proportionate residuals were used to calculate these correlations. The results do not permit us to reject the view that period effects are unimportant in these data. Nine of the ten estimates of the correlation coefficients calculated using the absolute residuals are insignificantly different from zero, with three being positive and seven negative (the average coefficient is -.12). Examining the proportionate residuals, we also find that nine of the ten estimates of the correlation coefficients are insignificantly different from zero, with three being negative and seven positive (the average coefficient is .11).

B. Estimates Computed with Covariates

We now introduce covariates into the specification of the marriage model, for the cohorts aged 30 to 34 and above in 1985. We do not focus on younger cohorts in view of results presented in Bloom (1982) indicating that the Coale-McNeil model sometimes performs inadequately when the data are censored to within a few years above the estimated mean of the distribution.

Education is defined as years of schooling at the time of the survey and not at the time of the first marriage because this variable can be constructed for all three data sets. In work not reported here based on the National Longitudinal Survey of Young Women, we found that using education at the time of first marriage instead of education at the time of the survey had almost no impact on the parameter estimates.

Incorporating covariates into the model adds to its explanatory power, as shown by the highly significant increase in the maximized log likelihood. The results in Table 2 reveal that, generally, both education and race relate significantly to the timing of a woman's first marriage and that race bears especially importantly on the propensity to marry. Among white women aged 30 to 34 in 1985, for example, those with more than a high-school degree can expect to marry 3.6 years later on average than those who never completed high school. Blacks tend to marry about a year later than their white counterparts, controlling for education. (See also Bennett, Bloom, and Craig, 1989, which reports some comparable estimates. That article takes the empirical validity of the marriage model as given and focuses its attention mainly on testing alternative hypotheses about the estimates of racial differentials in marriage patterns and their evolution across cohorts.)

The proportion of women expected to ever marry has declined considerably across cohorts for all groups of women. Most notable is the dramatic rise in the proportion of black women who are expected to never marry. For example, approximately 12 percent of the oldest cohort of black women in the sample will never marry. However, for the cohort 15 years younger at the time of the survey, our estimates suggest that 25 percent will never marry.

Cross-cohort comparisons of the estimated education effects may be somewhat biased by cross-cohort changes in mean educational attainment *within* the educational categories we use. For example, in the 1985 CPS, the average educational attainment 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 0.2 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.

Last, the estimates reveal that the overall trend across cohorts to delay marriage is not solely due to increased educational attainment, but also to a tendency for more-educated women to marry at later ages. For example, the mean age at first marriage for more-educated white women increased by about one year from the cohort aged 45 to 49 in 1985 to the cohort born 15 years later. In contrast, the mean age at first marriage for less-educated white women exhibits no trend across these cohorts.

C. Sensitivity Analysis with Covariates

Since the parameter estimates reported in Tables 1 and 2 are computed from data that are censored, their accuracy is heavily dependent upon the statistical structure 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 Figures 3 through 6.

The closeness of the parameter and hyper-parameter estimates derived from different data sets 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.

For those cohorts aged 35 to 39, 40 to 44, and 45 to 49 in 1985, we have fit the model with covariates to data from the June 1976 CPS. Parameter estimates for these cohorts, which are reported in Table 3, are based on nine fewer years of marriage experience relative to estimates for the same cohorts computed from the 1985 CPS. There are two possible reasons why the estimates derived from the two surveys might differ: (1) the existence of sampling variability between independent samples drawn from the same population or (2) cohort marriage patterns do not adhere stably to the marriage model.

A fairly consistent message emerges from a comparison of the results found in Tables 2 and 3: The estimates of $\underline{E}$ appear more robust to censoring than the estimates of μ and σ. For all but four of the 18 subgroups implicit in our specification, the estimates based on the 1985 data reveal that women have married somewhat later in life than we would have anticipated given the estimates based on the 1976 data. Furthermore, for all subgroups we find that the estimated standard deviations of age at first marriage derived from the 1985 data are greater than those derived from the 1976 data.

We have also analyzed 1985 CPS data in which first marriages after 1976 are artificially censored. These data tend to yield estimates of μ and σ that are lower than those that emerge from the non-censored 1985 data, for a majority of the subgroups. Based on this result, it appears that modeling error accounts for more of the discrepancy between the estimates in Tables 2 and 3 than does sampling variability.

The estimates of $\underline{E}$, however, are quite stable across data sets, with an average absolute deviation between the 1976 and 1985 based estimates of less than one percentage point across the 18 subgroups. With the heightened public sensitivity in recent years concerning whether women are foregoing marriage entirely or merely delaying marriage (see, e.g., Bennett and Bloom, 1986), this result fosters confidence in the ability of results based on cohort marriage models to contribute usefully to this debate. Cohort marriage patterns do not adhere to the marriage model in a perfectly stable manner over time, but they do conform closely enough for the model to be judged a useful analytical tool.

VI. Discussion and Conclusions

This analysis has investigated the application of the three-parameter, Coale-McNeil marriage model to survey data on the timing and incidence of first marriage. Empirical evidence, based on data that were not generated by simple random sampling, is reported on the ability of the model to replicate observed U.S. marriage behavior and to estimate the parameters of the marriage process accurately from censored data. Specifications that permit the parameters of the model to depend on individual-level social and economic characteristics have also been examined.

Substantively, we find that the age at first marriage has increased by about one and one-half years across cohorts born between the early 1940s and the late 1950s. The proportion never-marrying has also changed substantially over time, more than doubling for women aged 25 to 29 in 1985 (12-13 percent) relative to those born 20 years earlier (5-6 percent). We also find that educational attainment has a strong positive association with the age at which women first marry, given that they marry. In addition, race is found to be a large and increasingly important correlate of a woman's propensity to marry. For example, only 80 percent of black women aged 35 to 39 in 1985 who had not graduated high school can be expected to marry, as compared with 92 percent of their white counterparts. We also note that the increased propensity (across cohorts) to delay marriage is due not only to increased educational attainment, which is traditionally associated with later age at marriage, but also to the tendency for highly educated women to marry at increasingly older ages. Finally, our analysis suggests that the time-series decline in U.S. marriage rates that began in the early 1970s is associated with a growing tendency across cohorts to forego marriage, and for those who do marry, to marry at later ages.

Table 1. Parameter estimates of the simple Coale-McNeil marriage model, with asymptotic standard errors reported in parentheses.

Data Set	Age in 1985	μ	σ	E
NSFG (1982)	25-29	23.06 (.30)	4.87 (.24)	.868 (.034)
	30-34	21.69 (.13)	3.70 (.11)	.886 (.011)
	35-39	21.83 (.12)	3.98 (.10)	.909 (.009)
	40-44	21.56 (.14)	3.94 (.11)	.953 (.007)
CPS (1985)	25-29	22.71 (.10)	5.00 (.09)	.877 (.008)
	30-34	21.92 (.07)	4.35 (.07)	.886 (.005)
	35-39	21.65 (.06)	4.15 (.05)	.922 (.004)
	40-44	21.39 (.06)	4.06 (.05)	.947 (.003)
	45-49	21.07 (.07)	4.12 (.06)	.944 (.004)

Table 2. Parameter estimates, based on the June 1985 Current Population Survey, of the hyper-parameterized Coale-McNeil marriage model, with asymptotic standard errors reported in parentheses.

		Age in 1985			
		30-34	35-39	40-44	45-49
μ	Constant	19.68 (.15)	19.94 (.15)	19.66 (.14)	19.98 (.14)
	Black	1.06 (.25)	0.62 (.21)	1.96 (.29)	0.93 (.28)
	Ed = HS	1.23 (.16)	0.91 (.16)	1.12 (.16)	0.65 (.17)
	Ed > HS	3.60 (.18)	2.90 (.18)	2.80 (.17)	2.37 (.20)
σ	Constant	3.45 (.13)	3.80 (.13)	3.59 (.12)	3.87 (.13)
	Black	1.21 (.22)	0.63 (.18)	2.07 (.25)	1.54 (.25)
	Ed = HS	0.07 (.14)	-0.41 (.15)	-0.26 (.14)	-0.44 (.15)
	Ed > HS	1.10 (.16)	0.58 (.16)	0.51 (.15)	0.66 (.17)
E	Constant	.907 (.012)	.923 (.012)	.944 (.009)	.948 (.009)
	Black	-.160 (.017)	-.126 (.017)	-.086 (.016)	-.070 (.016)
	Ed = HS	.023 (.013)	.034 (.013)	.026 (.009)	.021 (.009)
	Ed > HS	-.020 (.014)	.002 (.014)	.002 (.011)	-.021 (.012)

Table 3. Parameter estimates, based on the June 1976 Current Population Survey, of the hyper-parameterized Coale-McNeil marriage model, with asymptotic standard errors reported in parentheses.

		Age in 1985		
		35-39	40-44	45-49
μ	Constant	18.96 (.11)	19.23 (.12)	19.40 (.12)
	Black	0.70 (.25)	0.71 (.21)	0.42 (.29)
	Ed = HS	1.62 (.13)	1.40 (.14)	1.13 (.14)
	Ed > HS	4.16 (.07)	3.42 (.17)	3.04 (.18)
σ	Constant	2.96 (.10)	3.40 (.11)	3.39 (.10)
	Black	0.78 (.19)	1.21 (.23)	1.19 (.23)
	Ed = HS	0.03 (.11)	-0.26 (.12)	-0.21 (.12)
	Ed > HS	1.03 (.10)	0.60 (.15)	0.47 (.15)
E	Constant	.945 (.009)	.945 (.009)	.963 (.007)
	Black	-.130 (.019)	-.071 (.019)	-.063 (.016)
	Ed = HS	.009 (.011)	.018 (.010)	.001 (.008)
	Ed > HS	-.022 (.013)	-.011 (.012)	-.035 (.011)

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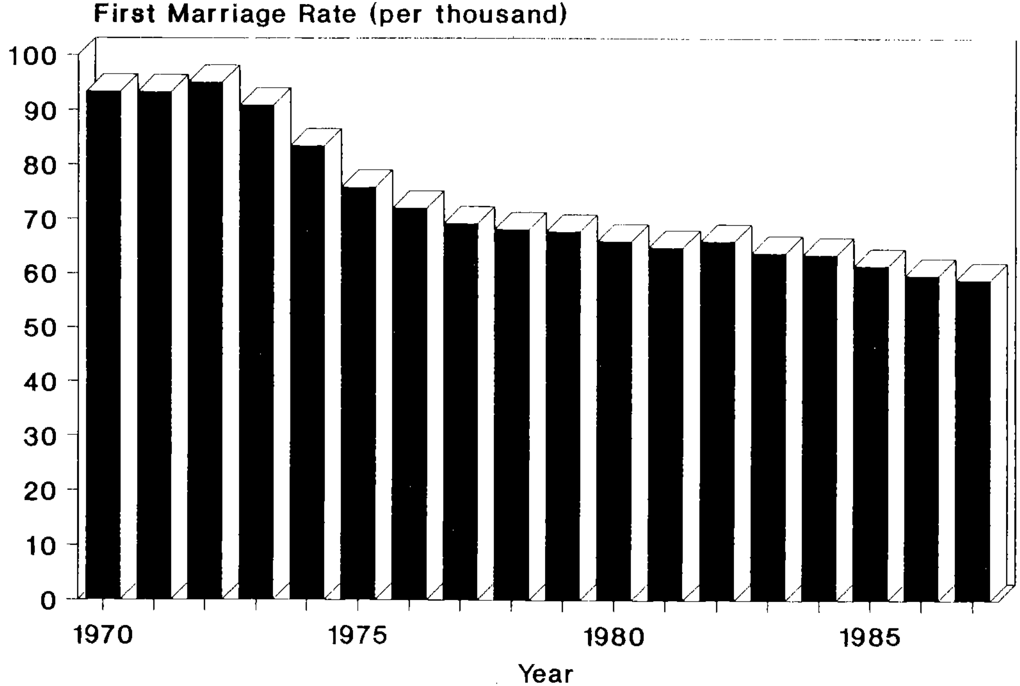
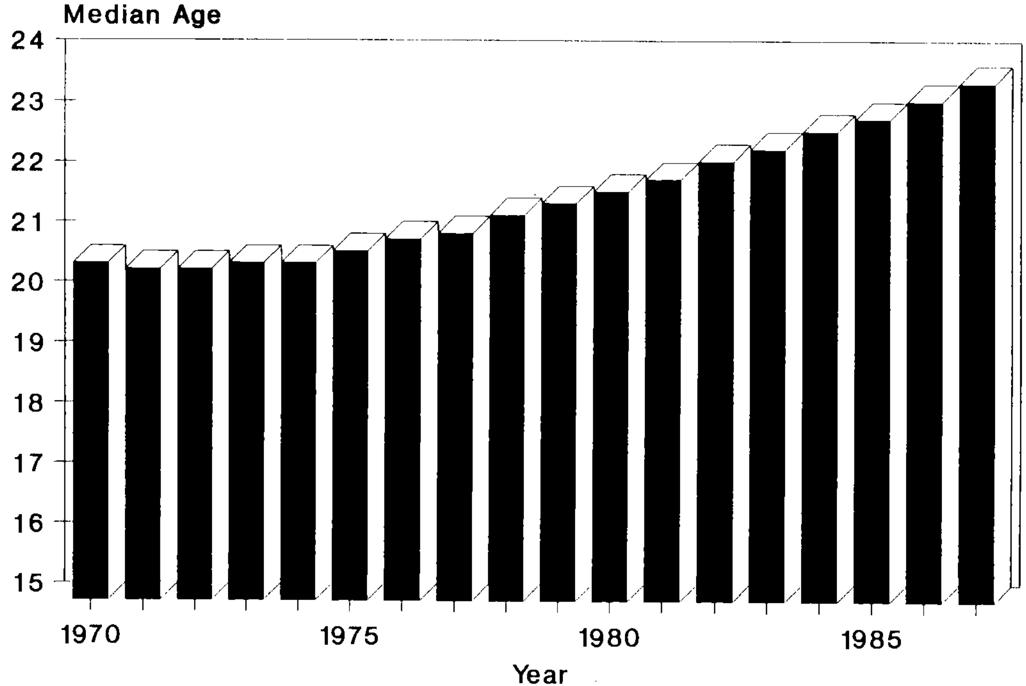
Figure 1: First Marriage Rates of
Women Aged 15 and Above, 1970-1987



National Center for Health Statistics

Figure 2: Median Age at First Marriage for Women, 1970-1987



National Center for Health Statistics

Observed vs. Projected Marriage Rates for Women Aged 30 to 34 in 1985

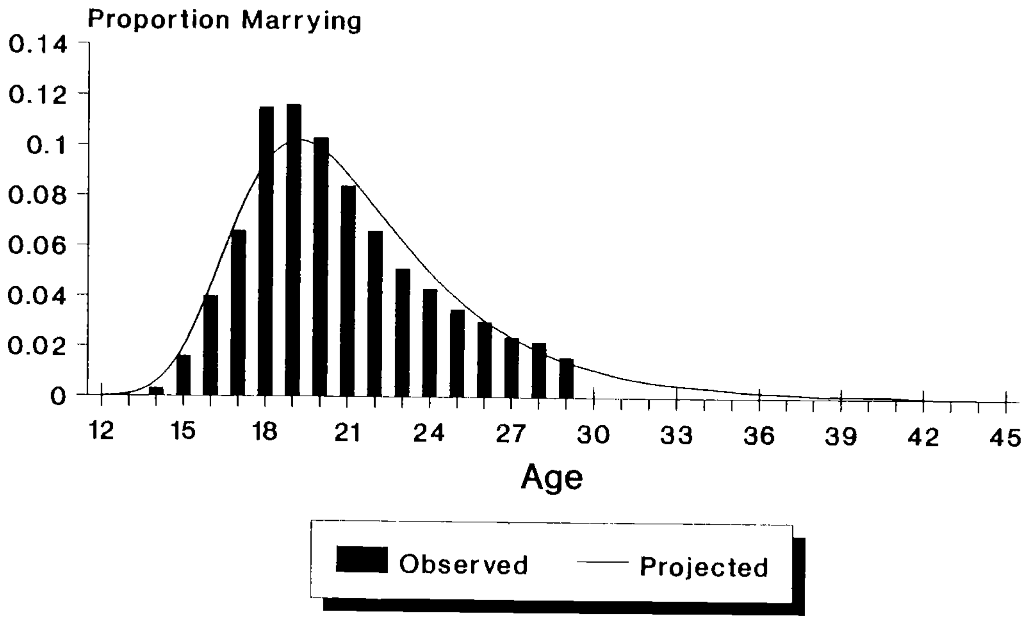


Figure 3

Observed vs. Projected Marriage Rates for Women Aged 35 to 39 in 1985

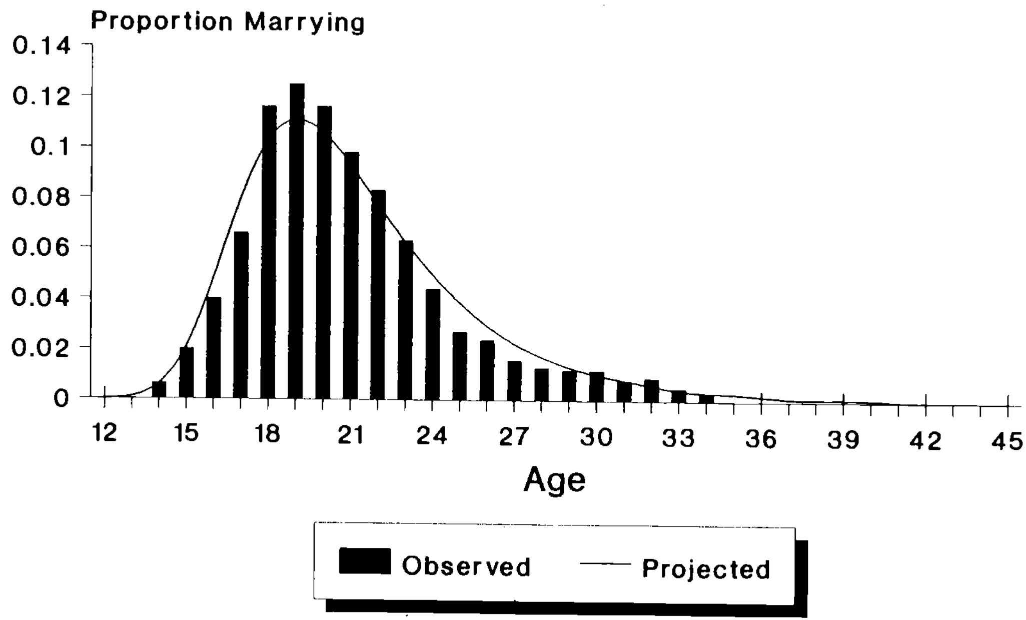


Figure 4

Observed vs. Projected Marriage Rates for Women Aged 40 to 44 in 1985

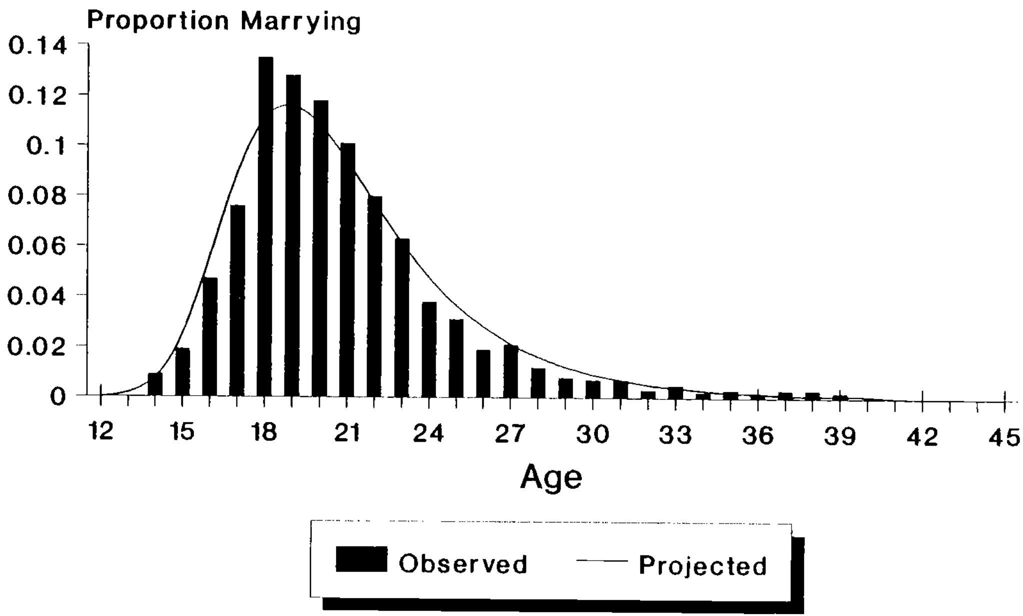


Figure 5

Observed vs. Projected Marriage Rates for Women Aged 45 to 49 in 1985

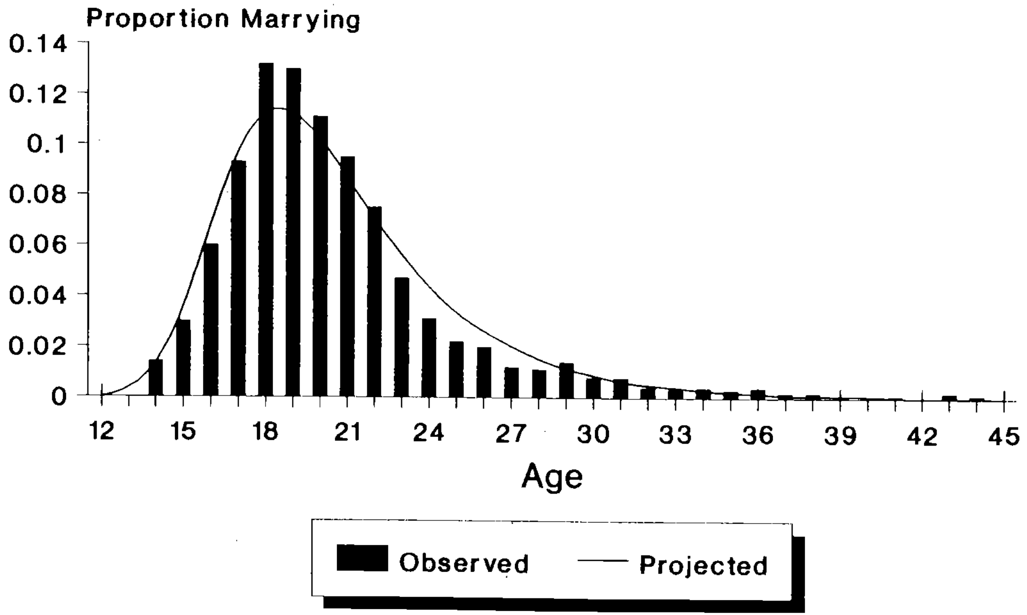


Figure 6