Assortative Mating, Intergenerational Mobility, and Educational Inequality

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INTRODUCTION

How much husbands and wives resemble each other on socioeconomic status and other social traits has long been a concern for demographers and students of social stratification. For demography, assortative mating patterns govern the social makeup of marriages and families. For stratification, spousal resemblance is an indicator of openness or closure of the social order. When people tend to marry within their own social stratum, populations and societies are more closed when they commonly marry members of other socioeconomic groups. But more that just an indicator of family structure and social differentiation, assortative mating is also a part of the dynamic process by which populations and social hierarchies reproduce. The formation and persistence of marriages affects levels and patterns of fertility and thus population growth rates and age structures. The kinds of marriages that form between persons of varying social characteristics determine the "family backgrounds" for their offspring and thus affect the social characteristics of the next generation. At the family and individual levels, mothers and fathers affect the eventual social standing of their children. At the population and cohort levels, the joint distribution of parents' characteristics affect the level and distribution of offsprings' characteristics. Assortative mating, in short, is a key step in the reproduction of populations and social hierarchies from generation to generation and from one period to the next.

The importance of assortative marriage for long run changes in inequality in the United States is an issue that has recently come to light. Several recent studies have shown that the association between husband's and wife's educational attainments has increased over cohorts of marriages during the past several decades, a trend that has been the subject of diverse

interpretations.   Kalmijn (1991a, 1991b) has argued that this trend is part of a society-wide

evolution towards achievement-oriented principles of social organization and away from the

ascriptive ties of family background, religion, and race-ethnicity. A second interpretation is that

trends in assortative mating reflect changes in typical ages of leaving school and marrying. As

average ages of leaving school have increased, individuals are more likely to meet potential

spouses while in school or at a time when school-based ties still dominate their social lives. All

things being equal, this will drive up the correlation between the final educational attainments of

spouses (Mare 1991). Yet another interpretation stresses the increasing symmetry between men

and women in what they expect from a spouse. As both men and women come to expect that

women will do market work during most of their lives, men may increasingly adopt the same

market-based criteria for evaluating prospective wives that women have traditionally used for

husbands (Oppenheimer 1988). If education is a marker for potential economic success, this

change in expectations increases the correlation between spouse's educational attainment (Mare

1991). Although these interpretations may not lead to the same predictions about the future trend

in assortative mating, they are all broadly consistent with recent history.

Whatever the causes of increasing resemblance in educational attainment between

husbands and wives, there is reason for concern about its potential consequences. A closer

resemblance between husbands and wives on socioeconomic characteristics may increase the

variance among families and thus increase the variability in family backgrounds of the next

generation of children. This in turn has the potential for widening educational inequalities in the

next generation (Epstein and Guttman 1984; Mare 1991). A particularly extreme form of this

concern is expressed by Herrnstein and Murray (1994), who argue that increasing assortative

mating on educational attainment signifies the segregation of the "cognitive elite" from the remainder of the population, a trend that heightens the impact of (genetically determined) cognitive inequality on socioeconomic inequality.

Whereas the facts about the trend in assortative mating are reasonably clear, research on its consequences is scarce. Although the variance in the educational attainment of offspring, *ceteris paribus*, does indeed vary directly with the covariance of mother's and father's schooling, it is harder to establish that trends in assortative mating have exerted an upward force on educational inequality in the past or are likely to do so in the future. The effects of assortative mating work through a complex set of sociodemographic processes including intergenerational social mobility, fertility, and mortality. This paper reports an investigation of the effects of changes in marriage patterns on inequality of educational attainment in the United States. It is a progress report on a larger project, which investigates the implications of demographic processes for aggregate trends in educational attainment and other socioeconomic outcomes in the United States over the past 70 years. The project focuses on the combined effects of such factors as differential fertility and mortality, marriage, assortative mating, family structure, immigration, and intergenerational social mobility on the distribution of socioeconomic characteristics.

In recent papers (Mare 1996, in press) I have examined the effects of differential fertility on trends in average levels of and racial inequality in educational attainment. Using a one-sex model of population growth that allows for intergenerational educational mobility, I explored the degree to which the relatively low fertility rates of highly educated women has dampened the secular growth of educational attainment. This research speaks to recent concerns about the presumed "dysgenic" effect of differential fertility (Herrnstein and Murray 1994) and examines

the degree to which the effects of differential fertility depend upon the degree of intergenerational mobility (Lam 1993; Preston and Campbell 1993). Although high fertility of more poorly educated women has a long run negative effect on average educational attainment, this effect is very small. The effects are small in part partly because educational fertility differences are highly variable and often small. More important, the negative effect of differential fertility on average educational attainment is offset by relatively high rates of intergenerational mobility, which imply that many offspring of high fertility, poorly educated women achieve high levels of educational attainment.

In the present paper, I extend this work to examine the effects of assortative mating patterns on the distribution of schooling. The paper attempts to answer the following questions:

1. Are trends in educational inequality consistent with those that we would expect from increases in the correlation of mother's and father's schooling?

2. In the absence of offsetting factors, how large a change in educational inequality would be expected as a result of changes in assortative mating patterns?

3. Considering a wide hypothetical range of possible assortative mating regimes, what is the *potential* impact of assortative mating on educational inequality? How does this impact depend upon rates of intergenerational educational mobility?

The balance of the paper is as follows. The next section discusses the potential effects of marriage patterns on inequality and argues that a coherent demographic model is required to investigate these effects. Subsequent sections (1) review prior efforts at modelling the growth of populations that are heterogeneous in socioeconomic status (2) present a two-sex model of population growth that allows for differential fertility by educational attainment and varying

degrees of assortative mating by education; (3) describe the data used to carry out simulations of population growth under alternative assumptions about fertility, marriage, and intergenerational mobility; (4) describe trends in assortative mating during the past 50 years; (5) describe trends in educational inequality during the 20th century; (6) present the estimated effects of observed assortative mating patterns on educational inequality; (7) present the simulated effects of assortative mating under alternative hypothetical patterns of intergenerational educational mobility; and (8) summarize the findings and their implications.

ASSORTATIVE MATING AND EDUCATIONAL INEQUALITY

The potential effect of assortative mating on inequality can be seen by using a rudimentary model of the effects of parental educational attainment on offspring educational attainment. Let Y_i denote the years of school completed by a child in the i*th* family, M_i and F_i denote the educational attainments of the child's mother and father respectively. Then a simple linear model is

$$Y_i = \beta_0 + \beta_1 Mi + \beta_2 F_i + \varepsilon_i$$

where the β's denote parameters and ε_i denotes a random disturbance with a mean of zero. Under this model, the population average level of educational attainment of children is a function of population averages of mother's and father's schooling:

$$\mu_Y = \beta_0 + \beta_1 \mu_M + \beta_2 \mu_F$$

where the μ's denote the population averages of children's and parents' educational attainments. In addition, the variability of children's educational attainment is a function of the joint distribution of parents' educational attainments:

$$\sigma_Y^2 = \beta_1^2 \sigma_M^2 + \beta_2^2 \sigma_F^2 + 2\beta_1\beta_2 \; \sigma_{MF} + \sigma_\epsilon^2$$

where the σ^2's denote the variances of children's and parent's schooling and the random disturbance and σ_{FM} denotes the covariance of mother's and father's schooling. If mother's and father's schooling have equal effects on offspring's schooling, that is, $\beta_1 = \beta_2 = \beta$, then this expression simplifies to:

$$\sigma_Y^2 = \beta^2(\sigma_M^2 + \sigma_F^2) + 2\beta^2 \sigma_{MF} + \sigma_\epsilon^2$$

Thus the variance in offspring's schooling is affected by the variances of both mother's and father's schooling *and* by the covariance of parents' schooling. Over successive cohorts of children, an increase in the covariance of parents' educational attainments will increase the variance of children's schooling, assuming that the remaining parameters in the model remain constant.

Inasmuch as the inequality in a given trait in the offsprings' generation varies directly with inequality of the trait among fathers, with inequality among mothers, and with the covariance between fathers and mothers on the trait, one expects to see these effects in the population. The magnitude of these effects, however, must be determined empirically. To investigate these issues requires a two-sex model of population growth, which incorporates the education-specific fertility rates of male and female partners, patterns of assortative mating by educational attainment, and intergenerational educational mobility. The model jointly projects the socioeconomic distributions of men and women and take account of the interdependence of the male and female populations through marriage, as well as the processes of fertility, mortality, and intergenerational mobility that are included in the one-sex model.

PREVIOUS LITERATURE

The research reported in this paper is based on a model for the growth of a population that is socioeconomically heterogeneous and that experiences intergenerational socioeconomic mobility. Only a small number of prior studies have developed models for the reproduction of such populations. Matras (1961, 1967) presented a projection model for a one-sex population differentiated by both age and occupation. Preston (1974) analyzed the implications of differential fertility for differences in occupation distributions and occupational mobility opportunities for black and white men. Johnson (1980) sketched a two-sex population model of assortative mating, differential fertility, and the intergenerational transmission of religious affiliation. Using a one-sex model Lam (1986) examined the effects of differential fertility and intergenerational mobility on the time path of income inequality in Brazil. Preston and Campbell (1993) developed a two-sex model for the intergenerational transmission of intelligence (I.Q.). They showed that, under broad conditions, fixed rates of fertility and assortative mating by level of I.Q. lead to a stable distribution of I.Q. Mare (1996, in press) used a one-sex model to examine the effects of education fertility differences among women with varying levels of schooling on trends in average levels of educational attainment. Kremer (1995) developed a rudimentary two-sex model of assortative mating and educational mobility. His study is a point of departure for the present investigation inasmuch as his study examines in the way in which the impact of assortative mating on inequality is modified by social mobility. The present study goes beyond Kremer's work in that it develops an explicit model of renewal in an age-structured population and allows for complex patterns of intergenerational mobility and assortative mating.

THE TWO-SEX MODEL

 The model used in this paper assumes that the socioeconomic characteristics of both parents affect the mobility chances of their offspring and that the creation of marriages depends on the supply and mutual attraction of members of both sexes who vary in both age and educational attainment. This model is an extension of the two-sex models developed by Pollak (1986, 1987, 1990) and Preston and Campbell (1993). The extension lies in allowing for a population that is differentiated by socioeconomic traits and age, has intergenerational mobility between levels of the traits, and allows for a relatively unrestricted pattern of assortative mating and differential fertility by age and socioeconomic traits. Pollak generalized the one-sex stable population growth model, which assumes fixed age-specific fertility and mortality rates. In Pollak's Birth Matrix Mating Rule (BMMR) model, fertility rates are not fixed for either sex. Rather, the rates jointly defined by the ages (and other traits) of married couples are fixed. Preston and Campbell [1993] have developed a two-sex model that allows for differentiation by socioeconomic traits, but their models, unlike mine, do not allow for age structure and employ restrictive and extreme assumptions about assortative mating.[1]

 Assumptions. I assume that educational attainment is determined at birth, that the mortality rates of partners are independent, and that both male and female fertility rates are 0 below age 15 and above age 59.[2] Furthermore, following Pollak (1990), I assume that marriages last only one period -- in this case 5 years.[3] This assumption simplifies the mathematics of the model and makes the model applicable to a much broader array of data. In particular, this enables one to use assortative mating, fertility, mortality, and intergenerational mobility rates that

are not conditioned on marital duration. Operationally, it means that one regards couples as reestablishing their marriages and having children in each period at rates that are independent of their past marital experiences. Marriage and fertility, however, are assumed to depend on the ages and educational attainments of both spouses.[4] A further assumption in this model is that fertility occurs only within marriage. This enables me to focus explicitly on the effects of assortative mating. It would nonetheless be a useful refinement of the model to consider marital and nonmarital fertility.

The Model. Let $\mathbf{P_t}$ denote a $2RA \times 1$ vector of persons at time t, with typical entry P_{rast}, which is the population in the r*th* education group, the a*th* age group, the x*th* sex, in the t*th* year $r = 1, ..., R; a = 0\text{-}4, 5\text{-}9, ..., 55\text{-}59; x = \text{f, m}; t = 1,..., T)$. In each year, persons are at risk to unions with persons of the opposite sex in each of the possible age-education categories. Let $\mathbf{N_t(P_t)}$ be a vector-valued function that maps $\mathbf{P_t}$ into $\mathbf{C_t}$, an $(RA)^2 \times 1$ vector of couples who are classified by the age and educational attainment of each partner. The function $\mathbf{N_t}$ represents the marriage function, which reflects the forces of attraction between persons of various combinations of age and education. The results reported in this paper uses the harmonic mean marriage function, although other functions are also available (e.g., Pollard 1973; Schoen 1988; Pollack 1990).[5] Under the harmonic mean model, the expected number of unions in year t between women in age group a and education group r and men in age group a' and education group r is:

$$C_{aa'rr't} = \alpha_{aa'rr't}P_{arft}P_{a'r'mt}/(P_{arft} + P_{a'r'mt})$$

where $\alpha_{aa'rr't}$ is the "force of attraction" between women aged a in education group r and men aged a' in education group r'. In practice the $\alpha_{aa'rr't}$ is estimated from observed numbers of unions between men and women with varying characteristics in a given year and single persons of each

sex with those characteristics (Qian and Preston 1993).

The age-sex-education-specific populations five years later are linked to the population of couples at time t by a projection equation, that is,

$$\mathbf{P}_{t+5} = \mathbf{M}_S^* \mathbf{P}_t + \mathbf{M}_B^* \mathbf{N}(\mathbf{P}_t).$$

$\mathbf{M}_S^*$ and $\mathbf{M}_B^*$ are $2RA \times (RA)^2$ matrices of transition probabilities for the survival and birth parts of the population projection respectively. $\mathbf{M}_S^*$ can be written

$$\mathbf{M}_S^* \quad = \quad \begin{matrix} \mathbf{M}_{S11f} & \cdots & \mathbf{M}_{SARf} \\ \mathbf{M}_{S11m} & \cdots & \mathbf{M}_{SARm} \end{matrix} \qquad ()$$

where $\mathbf{M}_{Sarf}$ and $\mathbf{M}_{Sarm}$ are $RA \times RA$ submatrices for the ath category of the woman's age and the rth category of the woman's educational attainment for female and male offspring respectively and the subscript for time has been suppressed. $\mathbf{M}_B^*$ can be written in analogous fashion. The typical submatrix of $\mathbf{M}_S^*$, $\mathbf{M}_{Sarx}$, has A^2 submatrices of dimension $R \times R$, which contain the joint rates of survival between adjacent 5-year age groups for men and women. For women aged a in education category r and offspring of sex x (suppressing the subscript t for time):

$$\mathbf{M}_{sarx} = \begin{matrix} \mathbf{0} & \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{S}_{ar0} & \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \mathbf{S}_{ar5} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{0} & \mathbf{0} \\ & & & \cdots & & & \\ \mathbf{0} & \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} & \mathbf{S}_{ar55} & \mathbf{0} \end{matrix} \qquad ()$$

where the $\mathbf{S}_{ar0}$, ..., $\mathbf{S}_{ar55}$ are $R \times R$ submatrices of 5-year age-specific joint survival probabilities of women and their male partners. The typical element of submatrix $\mathbf{S}_{ara'}$, $S_{ijai'j'a'}$ is the probability that a woman aged a to $a + 5$ in education group i and her partner aged a' to $a' + 5$ in group i' are living five years later in groups j and j' respectively $(i,i',j,j' = 1, ..., R)$. The $S_{ijai'j'a'}$ reflect the joint processes of educational-specific adult mortality and mobility. Given the assumption of no

intragenerational educational mobility, the $\mathbf{S_{ara'}}$ are all diagonal submatrices in which the $S_{ijai'j'a'}$ equal the age-education specific 5-year joint survival rates of men and women if $i = j$ and $i' = j'$ and equal zero otherwise. Specifically, given the assumption of independence of partners' survival rates, when $i = j$ and $i' = j'$, $S_{ijai'j'a'} = S_{iai'a'} = ({}_5L_{i,a+5}/{}_5L_{ia})({}_5L_{i',a'+5}/{}_5L_{i'a'})$, where ${}_5L_{iat}$ is the life table population between ages a and $a + 5$ for education group i in year t.

The typical submatrix of $\mathbf{M_B^*}$, $\mathbf{M_{Barx}}$, has A^2 submatrices of dimension $R \times R$, which contain birth and intergenerational mobility rates for given combinations of men's and women's ages and educational attainments. Likewise, for women aged a in education category r and offspring of sex x (suppressing the subscript t for time):

$$\mathbf{M_{barx}} = \begin{array}{cccccc} \mathbf{0} & \mathbf{0} & \mathbf{B_{ar10x}} & \mathbf{B_{ar15x}} & \cdots & \mathbf{B_{ar59x}} \\ \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} \\ \multicolumn{6}{c}{\cdots\cdots\cdots\cdots\cdots\cdots\cdots\cdots\cdots\cdots\cdots} \\ \multicolumn{6}{c}{\cdots\cdots\cdots\cdots\cdots\cdots\cdots\cdots\cdots\cdots\cdots} \\ \mathbf{0} & \mathbf{0} & \mathbf{0} & \mathbf{0} & \cdots & \mathbf{0} \end{array}$$

where the $\mathbf{B_{ar10x}}$, ..., $\mathbf{B_{ar59x}}$ are $R \times R$ submatrices of 5-year age-specific birth and mobility probabilities, adjusted for parental and childhood mortality.

The typical element of submatrix $\mathbf{B_{ara'x}}$, $B_{ijai'j'a'x}$, is the probability that a woman aged a to $a + 5$ in education group i and a her partner aged $a' + 5$ in group i' have a child of sex x who is living five years later and ultimately enters education group j. The $B_{ijai'j'a'x}$ incorporate information on the fertility of women aged a to $a + 5$ and their partners aged a' to $a' + 5$, the survival of both parents during that age interval, the sex-specific mortality of children during the first five years of life, and the probabilities that women in education group i and their partners in education group i' have daughters and sons who attain education group j. That is,

$$B_{ijai'j'a'x} = {}_5L_{ii'0x}(F_{ii'aa'x} + S_{iai'a'}F_{ii',a+5,a'+5,x})M_{ii'jx},$$

where $_5L_{ii'0x}$ denotes the life table population between ages 0 and 5 for the offspring with sex x of mothers in education group i and fathers in group i', $F_{iai'a'x}$ denotes the rate of births of sex x to women aged a to $a + 5$ in education group i and their partners aged a' to $a' + 5$, $S_{iai'a'}$ denotes the age-education specific 5-year joint survival rates for women aged a to $a + 5$ in education group i and their partners aged a' to $a' + 5$ in education group i', and $M_{ii'jx}$ denotes the probability that a couple in which the woman is in education group i and the man is in education group i' has a child of sex x who reaches education group j.

Equilibrium Properties. Pollak (1990) has proven the existence of a nontrivial equilibrium for the BMMR model when marriage and fertility rates are constant. Preston and Campbell (1993) investigated the equilibrium properties of their models with fixed marriage, fertility, and intergenerational mobility rates. They demonstrate the existence of a unique equilibrium in models with completely endogamous mating and, under some conditions, for models with random mating. The equilibrium properties of the model used in this paper are thus far unknown. In practice, however, I have found no cases where long run projection using the model fails to achieve an equilibrium age and education distribution. In all cases that I have examined, moreover, these equilibria do not depend on the initial population. This suggests that under a range of realistic fixed mobility, marriage, and fertility regimes, the two-sex model yields a unique equilibrium. Nonetheless, further research on this issue is needed.

Method of Analysis. The ingredients of the projection model are the age-education-specific two-sex fertility rates ($F_{ii'aa's}$) and survival rates ($S_{iai'a'}$), the two-sex educational mobility rates ($M_{ii'jx}$), and the forces of marital attraction ($\alpha_{aa'rr't}$). By making alternative assumptions about how these quantities vary, one can explore the effects of demographic and mobility regimes on

education distributions. In the present investigation, I carry out two types of analysis using the two-sex model.

First, I compute the equilibrium age-education distributions implied by the assortative mating and fertility rates observed at the dates of the 1940, 1960, 1970, 1980, and 1990 censuses.[6] In all of these calculations I assume a constant pattern of intergenerational educational mobility implied by several large bodies of survey data (see below). This is in keeping with the relative stability of intergenerational mobility patterns throughout the 20th century (Mare 1981; Hout, Raftery, and Bell 1993). Changes over time in the equilibrium distribution of educational attainment show the long run effects of changes in assortative mating patterns on educational inequality. As discussed further below, I carry out these calculations under alternative assumptions about fertility differences by educational attainment and alternative statistical models of assortative mating.

Second, I compute hypothetical equilibrium age-education distributions under alternative assumptions about the strength of association between husband's and wife's schooling and between parents' and offspring's schooling. The purpose of this analysis is to show how the effect of assortative mating on inequality depends on the degree of intergenerational mobility. I calibrate the strength of assortative mating and intergenerational mobility using measures based on loglinear models for frequency data. These models are discussed below.

DATA

To carry out these analyses I use Census data for the computation of fertility rates and two-sex marriage rates; mortality data from both published and survey microdata sources; and

survey data on intergenerational educational mobility

   Fertility. I estimate fertility rates specific to mother's and father's 5-year age interval,

mother's and father's educational attainments, and sex of child using microdata from the

decennial censuses. In the broader project of which the present study is a part, I estimate

quinquennial age-race-education-specific fertility rates by the "own children in the household"

method for 1925 through 1990 from the 1940, 1960, 1970, 1980 and 1990 Censuses (Cho,

Grabill, and Bogue 1970; Grabill and Cho 1965; Rindfuss and Sweet 1977).[7]   For the present

paper, I use estimates for the five decennial census dates. This method uses information on the

number of children aged k of mothers aged a and fathers aged a' at census year t to infer the

number of births that occurred to wives aged $a - k$ and husbands aged $a' - k$ in year $t - k$. The

estimated number of children, mothers, and fathers are adjusted for childhood and adult mortality

during the period from $t - k$ to t, using the data on socioeconomic mortality differences that are

discussed below. The own-children estimates may be distorted with the propensity of children to

live apart from their parents. This bias, however, can be minimized by basing estimates on the

numbers of relatively young children in the household -- that is, children aged 0 to 14 in each

census. The 1940 and 1960 Public Use files are one percent random samples of the U. S.

Population; the 1970 file is a six percent sample, and the 1980 and 1990 files are five percent

samples.   To offset the instability of estimates for single years from the own children method, I

average the fertility rates over up to five adjacent years. The years for which fertility rates are

estimated, the band of years used to estimate the rates, and the census that is used for

constructing the estimates are as follows:

Year Estimated	Time Band	Census
1940	1938-40	1940
1960	1958-62	1970
1970	1968-72	1980
1980	1978-82	1990
1990	1988-90	1990

Intergenerational Mobility. To estimate intergenerational educational mobility rates, I use three large national cross sectional samples that represent the mobility experience during the study period and contain measures of family background and socioeconomic achievements of the respondents. The surveys are:

1. 1987-88 National Survey of Families and Households (NSFH) (Sweet, Bumpass, and Call 1988). This sample contains 13011 persons aged 18 and over, including a cross-section sample of 9643 households and a double sampling of blacks, Puerto Ricans, Mexican American, single-parent families, families with stepchildren, and persons in marriages of short duration. I restrict the NSFH sample to persons aged 20-69.

2. General Social Surveys (GSS) from 1972 to 1994 (Davis and Smith 1992). This self-weighting sample contains approximately 35,000 persons aged 18 and over who were interviewed in one of the 23 repeated national cross sectional surveys from 1972 through 1994. I restrict the GSS sample to persons aged 20-69.

3. Wave 2 data from the 1986, 1987, and 1988 Survey of Income and Program Participation (SIPP). This self-weighting sample contains 43,446 persons aged 25 to 64, who were administered a survey module on family background and intergenerational mobility (Hauser and Jordan 1992). Although some respondents were outside the 25-64 age range, parents' educational attainment is not measured for persons outside this age range.

Each survey is used to construct outflow rates of intergenerational educational mobility (that is, proportions of persons in each category of educational attainment within categories of mother's and father's educational attainment). In each survey, the respondent is sampled from a cohort of *offspring* and parental information is obtained retrospectively. Thus the samples provide distributions of parents of sampled offspring, rather than the progeny of a sample of adults of childbearing age. To take this feature of the design into account, I weight the NSFH, GSS, and SIPP observations inversely proportional to their reported number of siblings plus one. When these samples are restricted to persons in the age ranges described above and observations with missing data on parents' schooling, offspring's schooling, or number of siblings are eliminated, the resulting sample size is 59,652 which includes 28,996 observations from SIPP, 21,554 from GSS, and 9,102 from NSFH.

The raw outflow matrix contains the distributions of offspring's schooling conditional on parents' schooling. Because of the secular increase in educational attainment throughout the period covered by the mobility data, this matrix implies substantial educational upgrading from parent to offspring. If this matrix were used as a basis for projecting population change over many generations, the implied equilibrium distribution of schooling would be absorbed into the highest category of educational attainment. This would indicate a very low level of inequality for any marriage pattern, simply as a result of the nonstationarity of the intergenerational mobility matrix. To avoid this problem, I adjust the mobility matrix to have symmetric origin and destination distributions. That is, before computing outflow rates, I equate the marginal distributions of mother's and daughter's and of father's and son's educational attainment while preserving the odds ratios for association between parents' and offspring's schooling. This

results in an approximately stationary mobility matrix.[8] The outflow matrix is shown in Table

A1.

      Marriage and Assortative Mating. Data on assortative mating are drawn from the 1940

and 1960-90 Censuses. Counts of marriages are obtained specific to husband's and wife's

educational attainments and ages for couples in which both partners are aged 15-59.

Denominators for marriage rates are also obtained for the corresponding sex-age-education

groups from each census. These denominators are based on the harmonic mean formula for

marriage rates. In the projection models, marriage rates are based on the total stock of currently

existing marriages, rather that on couples who are newly married.

      Mortality. I use two sources of data on socioeconomic differences in adult mortality:

the 1960 Mortality Followback Survey, based on deaths to persons in the 1960 Census from May

through August 1960 (Kitagawa and Hauser 1973); and the National Longitudinal Mortality

Survey (NLMS) (Preston and Elo 1994; Elo and Preston 1995). Data from the 1960 study are

only available in published form and are limited to age-sex-education-specific death rates for

whites. The NLMS covers the years 1979-85 and gives estimates by age, sex, race, and

educational attainment.

      The projection models and fertility estimates also require estimates of educational

differences in infant and child mortality. I use two sources of data. One is the 1964-66 National

Sample Surveys of Natality and Mortality (National Center for Health Statistics 1970), which are

based on a 1 in 1000 sample of registered births to married women occurring between 1964 and

1966 and a 1 in 36 sample of registered infant deaths during the same period. The other data are

the 1985 linked birth and infant death records, which are based on all registered births and infant

deaths in that year (National Center for Health Statistics 1990).[9]

Using these data I calculate age-education-specific mortality estimates for infants and adults.[10] These data provide information neither for children and teenagers nor for mortality rates for most of the period from 1940 to 1990. I obtain temporal information using published decennial age-race-sex-specific death rates based on vital statistics and census data. To get education-specific mortality estimates for children and teenagers and for each census year from 1940 to 1990, I fit a third order polynomial to the log odds of mortality, using the available survey and vital statistics based rates as data points. This polynomial incorporates the assumptions that mortality varies with educational attainment, age, sex, and time, but that educational mortality differences do not vary by race, sex, age, and time.[11] I use the estimated polynomial to interpolate rates for ages and years for which data are not available and to obtain smoothed estimates for all combinations of age, sex, educational attainment, and year. By construction, the estimated differences in mortality by educational attainment are time-invariant, although the mortality levels do vary over time. Although these estimation procedures are not ideal, alternative realistic assumptions that about mortality are unlikely to affect estimated mortality very much or to affect the results of the analyses reported here.

TRENDS IN EDUCATIONAL ASSORTATIVE MATING

As discussed above, several studies have documented an increase in the association between the educational attainments of husbands and wives in the United States during the past several decades. Whether these changes will bring about increasing in inequality for subsequent generations depends in part on how large these increases have been.   In this section, I review

these trends in assortative mating during the past 50 years and show that one's conclusions about

the trend depend upon the specific measures of assortative mating that one chooses and, to a

lesser extent on the population of marriages that is examined. .

In the study of assortative mating per se, it is important to focus on the flow of new

marriages rather than the stock of all existing marriages. New marriages reflect directly the

impact of marriage market pressures that may affect the resemblance of spouses. In contrast to

the stock of all existing marriages, moreover, new marriages are not depleted by selective

divorce, mortality, or changes in spouses' characteristics after the marriage has taken place (Mare

1991; Kalmijn 1991a, 1991b). In the study of the intergenerational impact of assortative mating,

however, it is important to focus on the stock of existing marriages rather than just weddings.

The effects of assortative mating on the socioeconomic makeup of later generations work

through fertility and intergenerational mobility. Subject to the constraints of age, all couples --

not just newlyweds -- are the population at risk to fertility and the intergenerational transmission

of inequality. In this context it is only useful to look at the flow of new marriages insofar as they

are a "leading indicator" of change in the stock of all marriages.

Table 1 reports trends in assortative mating using several measures and for alternative

populations of marriages observed in the decennial censuses. All estimates are based on couples

in which both partners were aged 15 to 59. The zero-order Pearson correlation between years of

school completed by husbands and wives does not show a clear trend between 1940 and 1990,

declining somewhat from 1940 to 1970 and increasing only slightly since then.[12] This apparent

absence of trend, however, reflects both that couples in a given year are members of many

marriage cohorts and also the limitations of correlation-based measures for understanding

assortative mating trends. For couples married during the 5 years prior and 5 to 10 years prior to each census, there is somewhat stronger evidence of trend. The correlation was relatively high in 1940, drops in marriages observed in 1960 and trends upwards thereafter.[13]

The second through fifth columns of Table 1 report canonical correlations between the sets of categorical variables that denote husband's and wife's educational attainment. These correlations avoid the scaling assumptions required by the Pearson measure and provide evidence of nonlinear and multidimensional change. Although the first canonical correlation have little systematic trends, the second through fourth correlations all tend to trend upward in all three populations either for the entire period or from 1960 on. Although the sizes of these correlations are hard to interpret, they indicate that the Pearson correlations obscure the nonlinear associations between spouses' educational attainments which may change systematically over time.

A preferable way to assess the association between husband's and wife's schooling is to use standard methods for the analysis of frequency data that are based on odds ratios. These measures are not affected by changes in the univariate distributions of schooling which were considerable during this period. The sixth column of Table 1 reports the log odds ratio estimated from a model of uniform association between husband's and wife's schooling. The log odds ratio assumes a uniform pattern of linear by linear association throughout the assortative mating table.[14] These measures increase monotonically from 1940 to 1990, indicating a marked strengthening of the association between spouses' attainments. For all marriages, the log odds ratio increased from .566 to .715 from 1940 to 1990, an increase of about 25 percent. The percentage increases are somewhat larger over the 40 year period from 1940 to 1980 for new marriages, suggesting that the change for marriages as a whole has been driven by more recent

marriage behavior.

An even more revealing set of measures for trends in assortative mating takes account of variations in the trends across levels of educational attainment. "Crossings" models show how difficult it is to marry someone above or below one's own educational level (e.g., Johnson 1980; Mare 1991). For example, in 1940 the log odds of a couple having one spouse with less than 12 years of schooling and the other having 12 or more years relative to both partners being above or below this educational threshold is -.728. As these parameters become more negative they show that marrying across a given barrier is becoming more difficult. Although the general drift of the crossings parameters in Table 1 is in this direction, the growth of educational barriers to marriage is larger and more sustained at the highest level of schooling. For the barrier between college graduates and other persons, the crossings parameters increased by more than 30 percent over the 50 year period for all marriages and over the 40 year period for new marriages.

These analyses confirm that the association between spouses' educational attainments has indeed increased markedly over the past fifty years, although our capacity to detect this change depends rather strikingly on the tools that we use to measure it. They also show, however, that the trend for existing marriages, which is the relevant population for the analysis of the inequality-producing effects of assortative mating, is more gradual than for weddings and recent marriages.

TRENDS IN EDUCATIONAL ATTAINMENT AND INEQUALITY

Table 2 and Figure 1 document trends in educational attainment and educational inequality in the United States for cohorts born between 1880 and 1970.[15] Table 2 shows the

estimates for the total population and separately by for men and women. Because the trends are similar for men and women, only the trends for both sexes combined are discussed here.[16] The 20th century witnessed extraordinary growth in formal education. Over the 90 year span of cohorts shown here, average years of school completed increased by approximately 5.5 years, from about 7.5 years for cohorts born at the end of the 19th century to about 13 years for cohorts born in the late 1950s and 1960s. Although educational attainment has continued to increase for some racial and ethnic minorities (Mare 1995), for the population as a whole average educational attainment peaked for cohorts born in the early 1950s and has been at a high plateau for cohorts born since then.[17]

This long run rise in average educational attainment has been accompanied by equally dramatic changes in the dispersion of the schooling distribution. In the 19th century, the population was concentrated at what, by today's standards, was a very low level of schooling. The long run S-shaped trend in average education shown in the first panel of Figure 1 has resulted in a population that is increasingly concentrated at a high level of schooling. Thus "inequality" of educational attainment increased over the first few cohorts shown in Table 2 and then decreased dramatically. The table shows trends in three measures of dispersion, the standard deviation (S.D.); the coefficient of variation, which is the ratio of the standard deviation to the mean; and the entropy (H) of the schooling distribution.[18] By all three measures, inequality of educational attainment fell dramatically across cohorts born between the early years of this century and 1970. The standard deviation fell approximately 30 percent over this period, the coefficient of variation fell by approximately 50 percent and *H* fell by about 10 percent. As we consider the effects of social and demographic forces that may have increased inequality,

therefore, it is important to recognize that these effects have occurred in the context of a massive *decrease* in inequality.

ASSORTATIVE MATING AND EQUILIBRIUM LEVELS OF INEQUALITY

To see the implied effects of changing marriage patterns on educational inequality, I estimate the equilibrium distribution of educational attainment that would arise if the patterns of assortative mating observed in each census were to continue indefinitely. A comparison of these estimates over censuses indicates the size of the effect of assortative mating. For each of the three measures of educational inequality discussed in the previous section, I carry out four sets of projections to equilibrium. For two sets of projections I allow the patterns of differential fertility by educational attainment that are observed in each period to persist indefinitely into the future. For the other two sets, I assume no differential fertility by educational attainment. Within each of these pairs of sets, I do one set of projections that assumes that assortative mating follows a simple pattern of uniform association, as described above;[19] and another set of projections based on observed patterns of assortative mating.

Table 3 reports the equilibrium inequality measures for these four sets of projections. In general, these simulations indicate that the effects of changes in assortative mating patterns are very small. Whether differential fertility is assumed or not, under uniform association, the standard deviation of schooling increases only modestly from 1940 to 1980 and then somewhat more after that. The overall increase during the 50 year period, however, is only about 10 percent despite a greater than 30 percent increase in the uniform association parameter for assortative mating over this same period (see Table 1). Changes in the coefficient of variation are somewhat

larger, especially during the 1980-90 interval than during the previous 40 years. Over the 50 year

period, this measure increases by about 25 percent. No systematic trends can be seen in the

entropy measure.

Although uniform association provides a convenient one-parameter summary of

assortative mating patterns, it does obscure variation in mobility patterns across levels of

schooling and in various parts of the assortative mating table. As shown in Table 1, increases in

marital homogamy occurred throughout the education distribution from 1940 to 1990, but these

changes were larger and more consistent at the highest schooling level. The full effect of trends

in assortative mating patterns can be seen in the lower panel of Table 3, which uses observed

mobility patterns rather than those imposed by a specific model. These projections provide even

less evidence of a systematic effect of trends in assortative mating on educational inequality than

those based on the uniform association model. Between 1940 and 1990, the standard deviation

of educational attainment does not change at all, although there is some evidence of a decline

between 1940 and 1980 before a modest increase in 1990. The coefficient of variation indicates

no change in inequality between 1940 and 1980 and then a sharp increase in 1990. Again, the

entropy measure shows no evidence of trend.

On balance, the changes in educational inequality implied by trends in assortative mating

from 1940 to 1990 are modest. Only the change in the coefficient of variation between 1980 and

1990 provides any strong evidence of increased inequality. This change, in conjunction with the

absence of trend before 1980 is not congruent with pattern of changes in actual assortative

mating shown in Table 1. That is, the association between husband's and wife's schooling in the

stock of marriages did not increase any more between 1980 and 1990 than it did during the

previous decades. Whatever the source of the increase in the coefficient of variation (and the much more modest increase in the standard deviation) between 1980 and 1990, it is not attributable to especially big changes in assortative marriage during this period.[20]

HOW MOBILITY ALTERS THE EFFECTS OF ASSORTATIVE MATING

To understand the weak and inconsistent effects of assortative mating on educational inequality, it is necessary to examine more closely the relationship between assortative mating, intergenerational mobility, and educational inequality. The effects of assortative mating on educational inequality depend upon the degree to which parents' educational attainments are associated with those of their offspring. In the simple linear model of schooling presented earlier in this paper, the contribution of the covariance of parents' educational attainments to the inequality of offspring's attainment is weighted by (the product of) the coefficients for the effects of the level of parents' schooling on the level of offspring's schooling. This simple model illustrates that if parents do not affect their offspring's schooling, the association between parents' educational attainments has no bearing on educational inequality. Even a change to a regime of completely endogamous marriage would have no affect on the distribution of offspring's schooling in that case.

To see how assortative mating affects educational inequality when educational inheritance is positive, I estimated equilibrium distributions of educational attainment under a wide range of assumed mobility and assortative mating regimes. In particular, I assumed a pattern of uniform association for both assortative mating and for intergenerational mobility. In the simulations, the uniform association parameter for mother's and father's schooling varies

from 0 to 1.5 in intervals of .1. The parameter for intergenerational mobility varies from 0 to 1.0

under a model in which mothers and fathers have an equal effect on offspring's schooling. That

is, if f_{ijk} denotes the number of families in which the mother (M) is in the ith education category,

the father (F) is in the jth category and the son or daughter (O) is in the kth category, the model

is:

$$\log f_{ijk} = \lambda + \lambda_i^M + \lambda_j^F + \lambda_k^O + \lambda_{ij}^{MF} + \lambda^{(M+F)O}(ik + jk)$$

where the λ's are parameters. Alternative levels of mobility (immobility) are specified by

varying $\lambda^{(M+F)O}$.[21]

Figures 2 - 4 show the simulated levels of educational inequality for alternative regimes

of assortative mating and educational mobility. For each of the three measures of inequality, the

strength of intergenerational inheritance has a major impact on the effect of assortative mating

and inequality. Inequality increases with the level of assortative mating for all mobility regimes

except where $\lambda^{(M+F)O} = 0$ (that is, in the figures, MobLOR = 0). The gradient for the assortative

mating effect, however, steepens markedly with the effect of parental schooling on offspring's

schooling. For example, over the full range of hypothetical marriage regimes, the standard

deviation varies by about by .16 (approximately 5 percent) when $\lambda^{(M+F)O} = .2$; by about .5

(approximately 15 percent) when $\lambda^{(M+F)O} = .5$; and by about .75 (approximately 20 percent) when

$\lambda^{(M+F)O} = .8$. Thus in populations that have a high degree of intergenerational immobility, changes

in assortative mating can potentially have a large effect on educational inequality. In populations

with only modest degrees of educational inheritance, by contrast, the effects of changes in

marriages regimes are very small.

In evaluating the results in Figures 2-4 one should compare the range of possible values

of the uniform association parameters for mobility and marriage included in the simulations to

the empirically observed values. As shown in Table 1, during the past 50 years, $\lambda^{(M+F)O}$ has varied

from .57 to .72, which implies a small change in educational inequality under any of the mobility

regimes considered in Figures 2-4. In fact, however, the uniform association parameter for

educational inheritance during this period falls well toward the bottom end of the range of values

considered here. For the pooled mobility data used in the present analyses, the parameter is

approximately .22.[22] At this level of mobility, the observed change in educational assortative

mating over the past 50 years can only have a trivial effect. Thus, even under extremely liberal

assumptions about the future changes in assortative mating, the impact of educational inequality

will be very, very small.

CONCLUSIONS

This paper has examined the conjecture that changes in patterns of assortative mating,

consisting of increasing resemblance between husbands and wives on educational attainment, are

a source of long-run increases in educational inequality. There is no doubt that the association

between husband's and wife's educational attainment has increased significantly during the past

half century. Measures of assortative mating which control for changes in the univariate

distributions of schooling suggest that the degree of resemblance between spouses increased by

25 to 30 percent over this period. The impact of this change on educational inequality, however,

is extremely small. We have presented three pieces of evidence in support of this conclusion.

First, the trend in educational inequality over cohorts born since the turn of the century has been

monotonically *downward*, despite increasing parental resemblance on educational attainment. Educational inequality shows no sign of being deflected from its downward trajectory by the patterns of assortative mating observed during the past five decades.

Of course, it is possible that the changes in educational assortative mating are too recent to have much impact on educational inequality, inasmuch as many of the cohorts born during the era of increasing spousal resemblance have not yet finished their schooling. One might think that increasing educational inequality will inevitably emerge. However, other results presented here indicate that if educational inequality does increase, it will not be due to assortative marriage. The second piece of evidence bearing on this issue comes from the equilibrium distributions of educational attainment that are implied by the patterns of assortative mating over the past five decades. These distributions show no systematic trend toward increased educational inequality even though these calculations have eliminated all factors that could offset the effects of changes in assortative mating.

The third piece of evidence comes from simulations which show that the potential effects of assortative mating on educational inequality can only be large when intergenerational educational immobility is much larger than has been observed during the past century. Under the existing regime of educational mobility, the potential effects of moving to a much higher degree of spousal resemblance than currently occurs are still likely to be very small indeed. Only with a major increase in educational inheritance can further increases in assortative mating have much of an effect.

Although assortative mating has a trivial effect on educational inequality, there is room for further investigation of its effects on other forms of inequality. The long run trend toward

equality of formal educational attainment has not been accompanied by any reduction in earnings and income inequality. Indeed, the latter inequalities have increased markedly during the past two decades. Accompanying these increases has been a sharp increase in economic inequality among persons receiving varying amounts of schooling. This suggests that it is important to consider schooling not only in the traditional units of years completed but also in the units of economic and other rewards that schooling brings. I am considering alternative metrics of schooling in my ongoing work.

While spousal resemblance on educational attainment has increased, marriages between members of different racial and ethnic groups have increased (Qian 1997). This suggests that changing marriage patterns may ameliorate racial and ethnic inequalities in economic status. Moreover, because race-ethnicity is more "heritable" than educational attainment, the effects of changing assortative mating may be larger than those described in this paper (Kremer 1995). Of course, in this case racial and ethnic intermarriage yields offspring who do in fact experience "mobility" from one race-ethnic group to a mixed racial-ethnic status. Before one can assess the effects of increasing intermarriage on inequality one must resolve the conceptual and definitional issues that are created by racial-ethnic exogamy.

FOOTNOTES

1. In particular, Preston and Campbell consider the cases of "complete endogamy"-- that is, marriage only within one's own group -- and random mating -- that is, marriage probabilities proportional to the size of each group. Using a particular functional form for the marriage function (see below), I allow for an arbitrary degree of marital endogamy.

2. In assuming that educational attainment is "determined at birth," the model does not focus on the accumulation of educational attainment during a lifetime. Rather, it assumes that individuals are given their final educational statuses at birth (with varying probabilities depending on their parents' educational attainments) and that these predestined statuses govern their future marital and reproductive behavior. This simplification enables one to avoid modelling the timing of educational attainment, fertility, and marriage within individuals' lives.

3. Pollak (1987) also develops a model that allows marriages to last more than one period. This model is more complex mathematically and has much more severe data requirements because it requires all rates to depend on both age and marital duration. Preston and Campbell (1993) also assume that marriages last only only period. Because their model does not incorporate age, they assume that marriages last one generation instead of one age interval as is assumed in Pollak (1990) and the present study.

4.   A further assumption of this analysis is that the distribution of educational attainment is mainly determined by family resources and preferences for schooling and is relatively unconstrained by the supply of "slots" in the educational system. The models used in this paper assume that social mobility and marriage are separable at the aggregate level. They do not allow for the possibility that patterns of marriage may interact with the structure of available socioeconomic outcomes to force changes in patterns of social mobility. The latter type of effect occurs when socioeconomic distributions are determined by a relatively fixed distribution of socioeconomic opportunities which are, in turn, established by the demand for labor of various types. In the 20th century United States, however, it is likely that educational opportunities have not been seriously constrained by supply.

5. In future work, I will explore the robustness of my results to the use of alternative marriage functions.

6. In practice, from an initial population, I project the population for 500 years or about 20 generations. In most instances, however, the distribution of the population by age and education fails to change by one tenth of one percent after about 250 years.

7. Although one can, in principle, compute own children estimates from the 1950 Census, the Public Use Sample includes too few women to provide reliable estimates. The long form of the 1950 Census was administered to a sample of *persons* rather than households. Because educational attainment was a long form item, educational attainment is recorded for only a fraction of women of childbearing ages in the 1950 Public Use Sample.

8. Some degree of nonstationarity is introduced into the projection by differential fertility of couples with varying levels and combinations of educational attainment. Differential fertility, however, has only a modest long run effect on average educational attainment (Mare 1996, in press), and thus is not a major source of nonstationarity.

9. Records from the 1985 data for California, Texas, and Washington are excluded from the analysis because they do not include information on maternal educational attainment.

10. The adult mortality rates are specific to persons' own educational attainments; the infant mortality rates are specific to their mothers' educational attainments.

11. The NLMS enables one to estimate separate patterns of education differences in mortality by sex and age. Because the estimated interactions between education and the other variables are not statistically significant, I retain only additive effects of schooling in the polynomial used in the analysis.

12. All measures in Table 1 are based on a five-category classification of husband's and wife's educational attainment (0-8, 9-11, 12, 13-15, and 16 or more years completed). The Pearson correlations assume scalar values of 6, 10, 12, 14, and 16 for these five categories respectively.

13. For new marriages no estimates are available for 1990 because the 1990 census obtained no information on age at marriage or marital duration.

14. More specifically, the model assumes that the local odds ratios formed by any 2 x 2 tables formed by pairs of adjacent rows and pairs of adjacent columns in the table are equal (Agresti 1990).

15. The trends reported in Table 2 and Figure 1 are based on data from the decennial censuses of 1940 and 1960-90 and have been constructed so that, with the exception of the 1965-69 cohort, estimates are based on persons between the ages of 25 and 59. The note to Table 2 indicates which census was used for each cohort. To make the estimates comparable to those based on the analyses reported later in the paper, I initially grouped the data into five categories of years of schooling (0-8, 9-11, 12, 13-15, and 16 or more) and assigned scores to these categories (6, 10, 12, 14, and 16 respectively). This tends to lower the estimated upward trend in mean educational attainment and to lower the standard deviation and coefficient of variation within each cohort compared to estimates based on ungrouped data. Thus, the estimates reported here may differ somewhat from those reported elsewhere.   The schooling data from the 1990 census were recoded as follows: Persons who reported a highest level of schooling completed of less than 9th grade were coded as "0-8;" persons who reported "9th grade," "10th grade," "11th grade," or "12th grade, no diploma" were coded as "9-11;" persons who reported "High School Graduate" were codes as "12;" persons who reported "Some college but no degree"or "Associate Degree in College" were coded "13-15;" all others were coded "16+."

16. Average levels of schooling have been similar for men and women throughout the 20th century. At the beginning of the century, however, men were overrepresented at the bottom and the top of the education distribution, evidently a result of the low educational requirements for

manual labor and lack of access to the professions for women. Over the course of the century, sex differences in the distributions of schooling have grown smaller.

17. For more detailed discussion of trends in rates of completing specific levels of schooling, see Mare (1995).

18. The entropy measure in computed over a five-category classification of schooling (0-8, 9-11, 12, 13-15, and 16 or more years of school completed) according to the formula:

$$H = \Sigma_{k=1}^{5} p_k \log(1/p_k)$$

where k indexes the education category and p_k denotes the proportion of the population in the kth education category. For a five category distribution H varies from 0 (maximum concentration) to approximately 1.609 (maximum dispersion). Because H is substantially influenced by categories that have very low relative frequencies, this measure shows unusually low dispersion for late 19th century cohorts in which relatively few persons completed college.

19. That is, these calculations assume that the number of couples of varying combinations of husband's and wife's educational attainment follow a model of uniform association with parameters equal to the log odds ratios shown in the first panel of Table 1. These numbers of couples form the numerators for the marital attraction parameters in the two-sex model.

20. Obviously, the trend in the coefficient of variation is affected by trends in both the standard deviation rather than the mean. Changes in the equilibrium means of educational attainment between 1940 and 1990 indicate that the mobility matrix that I have used is not purely stationary. This is a problem that I will address in ongoing work.

21. In practice, the model also includes terms for assortative mating by age, which are held constant across simulations.

22. When mother's and father's education effects are allowed to differ, their estimates are .227 and .212 respectively.

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TABLE 1
TRENDS IN SELECTED MEASURES OF EDUCATIONAL ASSORTATIVE MATING

	Pearson Correlation	Canonical Correlations				Log Odds Ratio	Crossings Parameters			
		1	2	3	4		9+/<9	12+/<12	13+/<13	16+/<16
All Marriages										
1940	0.606	0.613	0.338	0.179	0.104	0.566	-0.943	-0.728	-0.672	-0.695
1960	0.590	0.603	0.366	0.147	0.090	0.599	-0.820	-0.693	-0.782	-0.783
1970	0.583	0.606	0.381	0.159	0.100	0.654	-0.773	-0.748	-0.839	-0.831
1980	0.597	0.615	0.397	0.171	0.124	0.707	-0.832	-0.792	-0.827	-0.895
1990	0.596	0.601	0.412	0.211	0.157	0.715	-1.077	-0.781	-0.725	-0.961
New Marriages										
<5 Years										
1940	0.597	0.609	0.380	0.155	0.097	0.630	-0.874	-0.695	-0.874	-0.715
1960	0.570	0.603	0.364	0.157	0.105	0.689	-0.808	-0.670	-0.948	-0.828
1970	0.589	0.623	0.374	0.186	0.126	0.797	-0.844	-0.765	-0.924	-0.938
1980	0.596	0.617	0.391	0.205	0.138	0.812	-0.975	-0.793	-0.855	-0.952
5-10 Years										
1940	0.589	0.600	0.341	0.163	0.075	0.561	-0.844	-0.708	-0.738	-0.683
1960	0.565	0.594	0.352	0.134	0.080	0.624	-0.758	-0.661	-0.871	-0.806
1970	0.576	0.615	0.369	0.150	0.091	0.727	-0.758	-0.743	-0.887	-0.912
1980	0.601	0.621	0.398	0.187	0.127	0.786	-0.903	-0.819	-0.838	-0.948

NOTE: -- Measures are for couples in which both husbands and wives are aged 15-59.

TABLE 2
TRENDS IN LEVEL AND DISPERSION OF EDUCATIONAL ATTAINMENT BY SEX FOR COHORTS BORN FROM 1880 TO 1970

Cohort	Mean	S.D.	Coef. Var.	H	Mean	S.D.	Coef. Var.	H	Mean	S.D.	Coef. Var.	H
	Both Sexes				Women				Men			
1880-84	7.55	3.05	0.404	0.939	7.77	3.23	0.415	0.986	7.55	3.23	0.429	0.889
1885-89	7.70	3.14	0.408	0.993	7.91	3.31	0.419	1.034	7.70	3.34	0.434	0.949
1890-94	7.93	3.24	0.409	1.074	8.15	3.43	0.421	1.112	7.93	3.45	0.435	1.033
1895-99	8.23	3.35	0.407	1.162	8.41	3.52	0.419	1.185	8.23	3.58	0.434	1.133
1900-04	8.54	3.45	0.404	1.249	8.76	3.62	0.413	1.273	8.54	3.70	0.433	1.217
1905-09	9.08	3.49	0.385	1.365	9.30	3.65	0.392	1.378	9.08	3.76	0.414	1.344
1910-14	9.42	3.42	0.362	1.401	9.61	3.53	0.368	1.398	9.42	3.71	0.394	1.396
1915-19	10.25	3.38	0.330	1.478	10.31	3.43	0.333	1.448	10.25	3.74	0.365	1.500
1920-24	10.76	3.29	0.306	1.497	10.70	3.25	0.304	1.437	10.76	3.72	0.346	1.539
1925-29	10.92	3.26	0.299	1.515	10.91	3.18	0.292	1.448	10.92	3.73	0.341	1.558
1930-34	11.25	3.13	0.278	1.505	11.13	3.06	0.275	1.436	11.25	3.57	0.317	1.556
1935-39	11.87	2.96	0.250	1.482	11.60	2.93	0.253	1.426	11.87	3.35	0.282	1.525
1940-44	12.22	2.82	0.231	1.469	11.97	2.84	0.237	1.425	12.22	3.14	0.256	1.505
1945-49	13.07	2.72	0.208	1.447	12.67	2.76	0.218	1.421	13.07	2.98	0.228	1.457
1950-54	12.91	2.56	0.198	1.425	12.79	2.65	0.207	1.407	12.91	2.78	0.215	1.440
1955-59	13.03	2.51	0.193	1.440	13.13	2.60	0.198	1.422	13.03	2.74	0.210	1.457
1960-64	12.91	2.49	0.193	1.444	13.12	2.57	0.196	1.424	12.91	2.71	0.210	1.461
1965-69	12.60	2.26	0.179	1.360	12.91	2.33	0.181	1.340	12.60	2.44	0.194	1.373

NOTE. -- Data are from U.S. Censuses of 1940 (1880-1914), 1960 (1915-34), 1970 (1935-44), 1980 (1945-55), and 1990 (1955-69).

TABLE 3
EQUILIBRIUM MEASURES OF EDUCATIONAL INEQUALITY
IMPLIED BY ASSORTATIVE MATING PATTERNS, 1940-90

Year	Period-Specific Differential Fertility			No Differential Fertility		
	S.D.	Coef. Var.	H	S.D.	Coef. Var.	H
Uniform Association						
1940	2.90	0.230	1.562	2.89	0.231	1.566
1960	2.89	0.235	1.579	2.90	0.235	1.573
1970	2.91	0.240	1.581	2.92	0.240	1.577
1980	2.98	0.251	1.585	3.00	0.252	1.584
1990	3.17	0.294	1.562	3.10	0.281	1.575
Observed Assortative Mating						
1940	3.25	0.263	1.583	3.15	0.256	1.588
1960	3.19	0.268	1.605	3.04	0.250	1.591
1970	3.07	0.255	1.599	2.99	0.245	1.591
1980	3.07	0.258	1.602	3.05	0.255	1.598
1990	3.28	0.304	1.578	3.23	0.286	1.598

TABLE A1
OUTFLOW MATRICES FOR TWO-SEX
INTERGENERATIONAL EDUCATIONAL MOBILITY

| Mother's Schooling	Father's Schooling	0-8	9-11	12	13-15	16+	Total
		Son's and Daughter's Schooling					
		Sons					
0-8	0-8	0.611	0.215	0.110	0.042	0.022	1.000
	9-11	0.372	0.328	0.179	0.066	0.054	1.000
	12	0.276	0.260	0.246	0.121	0.097	1.000
	13-15	0.277	0.271	0.178	0.138	0.135	1.000
	16+	0.102	0.129	0.169	0.274	0.326	1.000
9-11	0-8	0.266	0.394	0.204	0.088	0.048	1.000
	9-11	0.146	0.417	0.248	0.132	0.057	1.000
	12	0.134	0.295	0.289	0.190	0.091	1.000
	13-15	0.079	0.204	0.327	0.200	0.191	1.000
	16+	0.045	0.409	0.214	0.175	0.157	1.000
12	0-8	0.206	0.274	0.291	0.139	0.089	1.000
	9-11	0.066	0.310	0.324	0.191	0.108	1.000
	12	0.048	0.193	0.360	0.240	0.158	1.000
	13-15	0.025	0.165	0.243	0.322	0.245	1.000
	16+	0.080	0.119	0.180	0.283	0.338	1.000
13-15	0-8	0.072	0.271	0.274	0.218	0.165	1.000
	9-11	0.104	0.256	0.233	0.237	0.171	1.000
	12	0.097	0.181	0.227	0.294	0.202	1.000
	13-15	0.170	0.133	0.171	0.265	0.262	1.000
	16+	0.000	0.073	0.140	0.294	0.492	1.000
16+	0-8	0.099	0.316	0.197	0.206	0.182	1.000
	9-11	0.176	0.034	0.200	0.242	0.348	1.000
	12	0.167	0.078	0.229	0.299	0.227	1.000
	13-15	0.000	0.021	0.152	0.417	0.410	1.000
	16+	0.096	0.035	0.099	0.265	0.504	1.000
		Daughters					
0-8	0-8	0.591	0.259	0.107	0.028	0.015	1.000
	9-11	0.278	0.381	0.214	0.087	0.040	1.000
	12	0.411	0.178	0.265	0.096	0.050	1.000
	13-15	0.329	0.169	0.238	0.164	0.100	1.000
	16+	0.269	0.214	0.185	0.180	0.153	1.000
9-11	0-8	0.282	0.379	0.224	0.070	0.044	1.000
	9-11	0.159	0.422	0.265	0.101	0.053	1.000
	12	0.033	0.354	0.360	0.167	0.085	1.000
	13-15	0.063	0.247	0.272	0.249	0.170	1.000
	16+	0.000	0.293	0.245	0.253	0.209	1.000
12	0-8	0.132	0.301	0.333	0.152	0.082	1.000
	9-11	0.081	0.239	0.353	0.211	0.116	1.000
	12	0.078	0.188	0.372	0.224	0.138	1.000
	13-15	0.131	0.134	0.246	0.288	0.201	1.000
	16+	0.096	0.141	0.184	0.280	0.299	1.000
13-15	0-8	0.124	0.210	0.306	0.241	0.119	1.000
	9-11	0.055	0.259	0.241	0.259	0.186	1.000
	12	0.061	0.175	0.241	0.311	0.211	1.000
	13-15	0.137	0.086	0.150	0.344	0.284	1.000
	16+	0.095	0.061	0.102	0.344	0.399	1.000
16+	0-8	0.124	0.086	0.224	0.294	0.271	1.000
	9-11	0.000	0.318	0.256	0.177	0.248	1.000
	12	0.042	0.070	0.238	0.320	0.331	1.000
	13-15	0.094	0.086	0.107	0.307	0.406	1.000
	16+	0.086	0.031	0.061	0.258	0.565	1.000
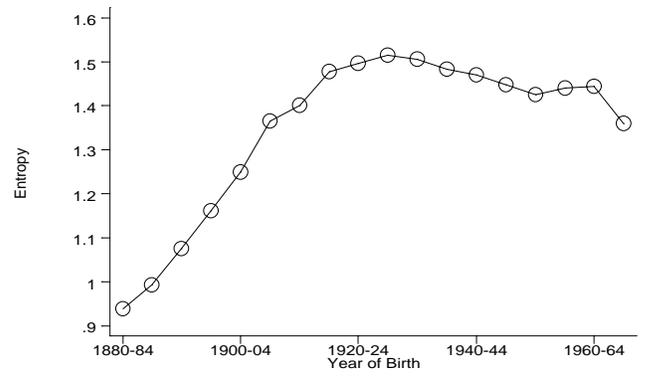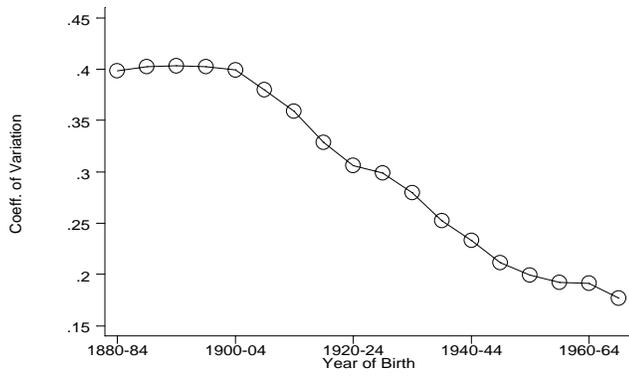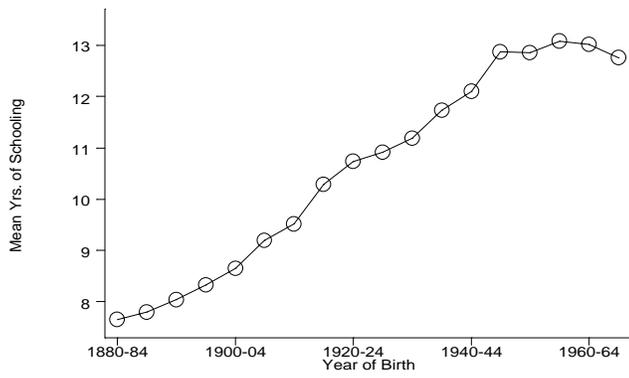
FIG. 1. -- Trends in Level and Distribution of Schooling
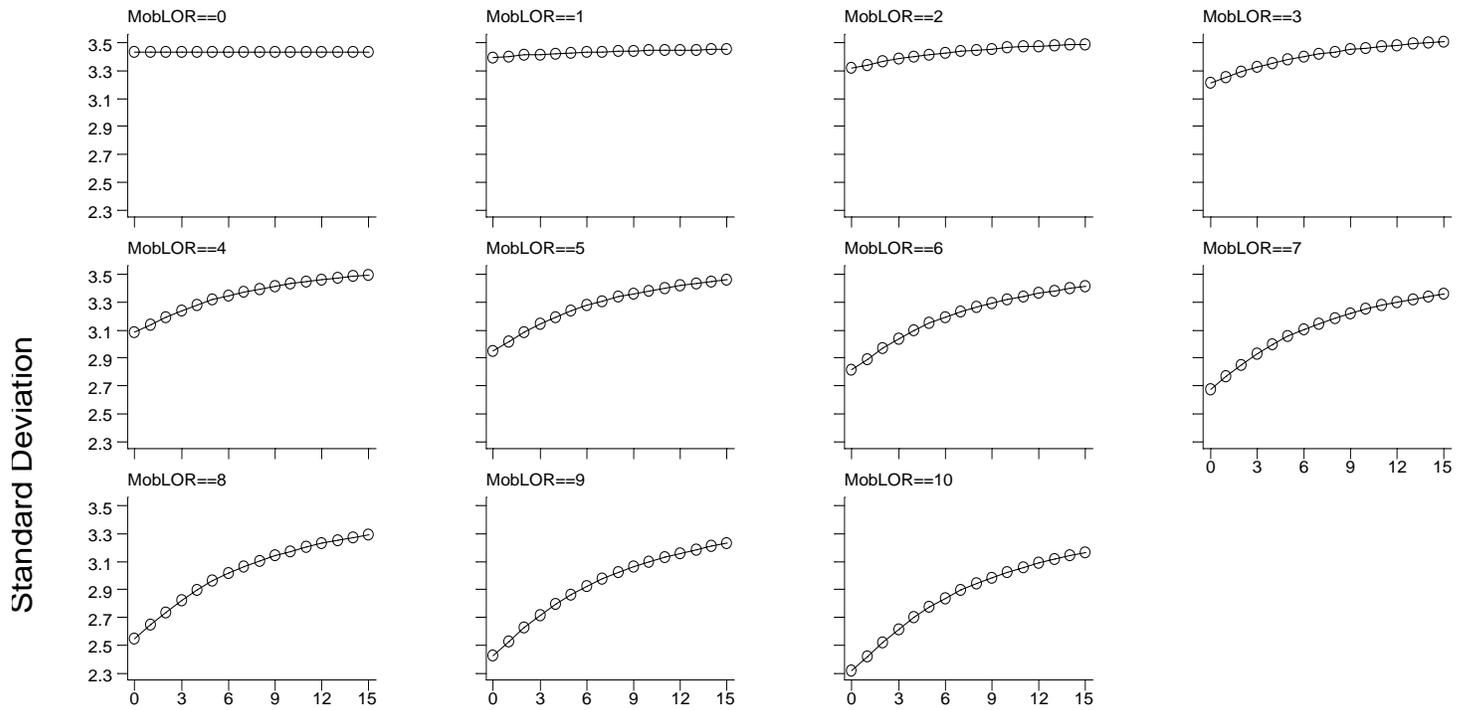
FIG 2: -- Effects of Assortative Mating on St. Dev. of Education
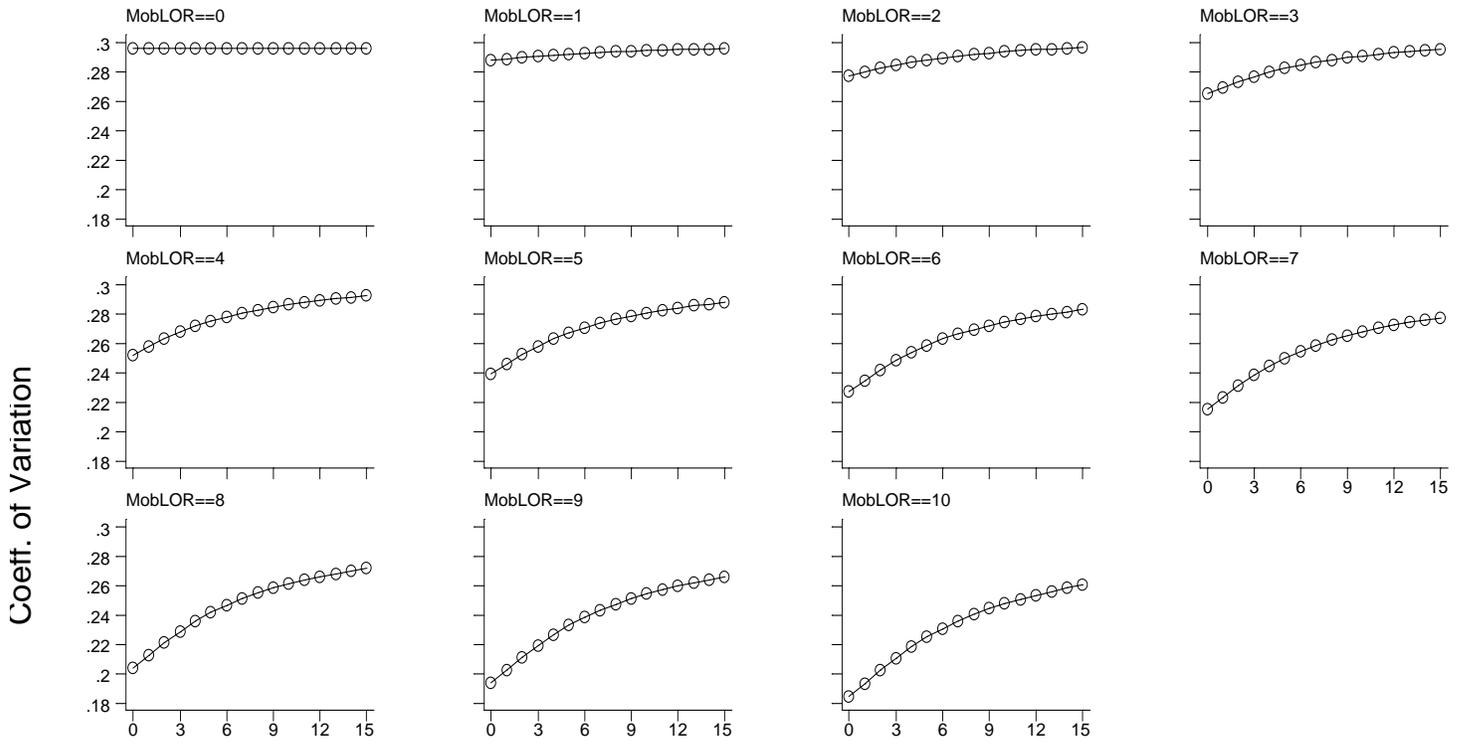
FIG 3: -- Effects of Assortative Mating on C.V. of Education
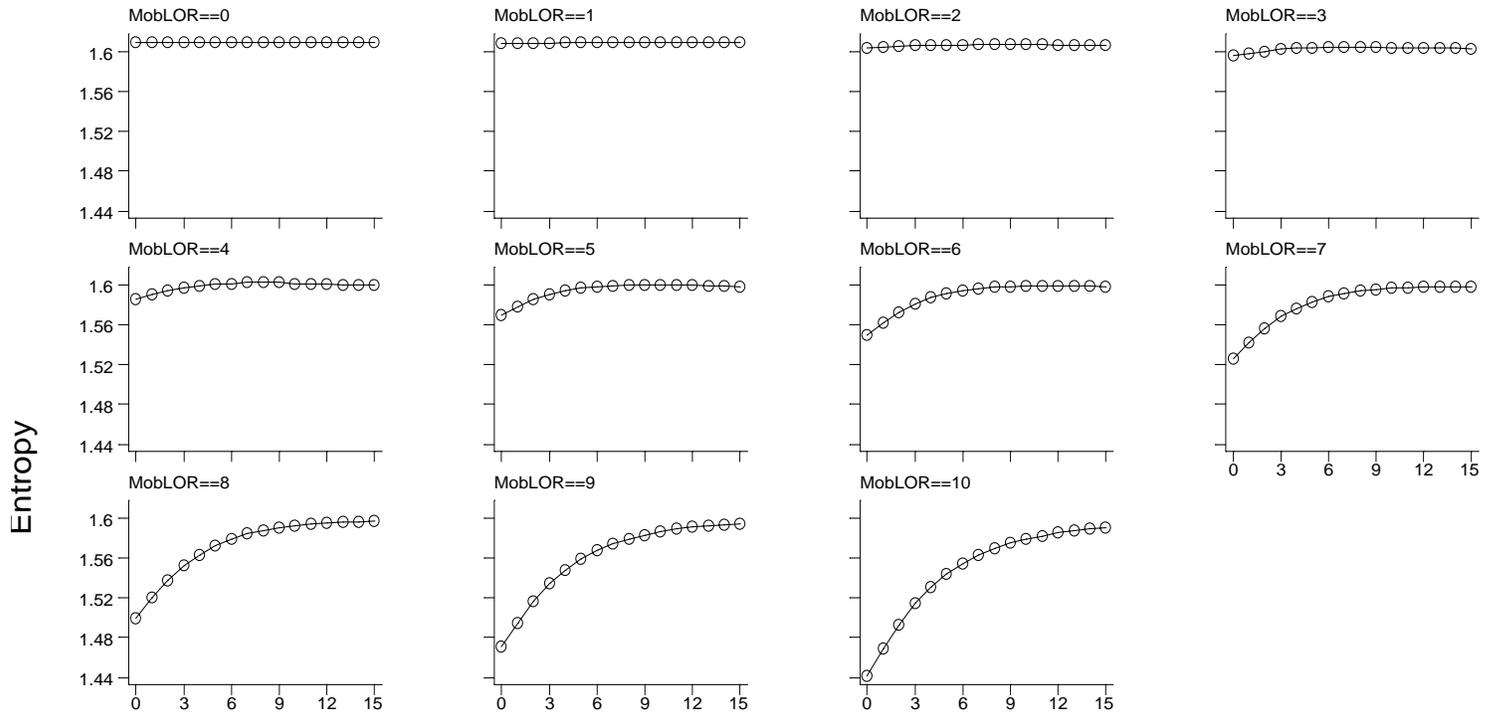
FIG 4: -- Effects of Assortative Mating on Entropy of Education