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**THE EFFECTS OF MARRIAGE, MARITAL DISSOLUTION, AND
EDUCATIONAL UPGRADING ON EDUCATIONAL ASSORTATIVE MATING**

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ABSTRACT

We examine changes in educational homogamy across one cohort's life and identify the demographic sources of these changes. Using data from the National Longitudinal Survey of Youth, we decompose changes in the odds of homogamy in prevailing marriages into three components: changes due to (1) new marriages, (2) marital dissolutions, and (3) educational upgrades after marriage. The odds of homogamy increase substantially as the cohort ages. These patterns are primarily driven by changes in the odds of homogamy among newlyweds entering their first marriages. Marital dissolutions, educational upgrades after marriage, and remarriages have smaller effects.

INTRODUCTION

Whom one marries is often thought of as an individual decision based on romantic love and mutual compatibility. On an aggregate level, however, marriage decisions follow regular patterns. Social scientists have long noted the tendency for individuals with similar characteristics to marry one another and often use assortative mating patterns to make inferences about the “openness” of societies. Because marriage creates close ties between individuals and families, societies in which many marriages cross social boundaries may be more open than societies in which there are few (Kalmijn 1991a, 1991b; Smits, Ultee, Lammers 1998). Further, because spouses share resources, assortative mating has implications for the distribution of cultural and economic resources. Finally, assortative mating shapes the characteristics of families and, to the extent that social attributes are inherited or learned from parents, the population composition of the next generation.

Patterns of educational assortative mating are of particular interest because of the pivotal role that education plays in the intergenerational transmission of social position, in structuring marriage markets, and in determining socioeconomic and demographic outcomes. Past research on educational assortative mating has been primarily concerned with explaining historical trends (Kalmijn 1991a, 1991b; Mare 1991; Qian 1998; Qian and Preston 1993; Rockwell 1976) and variation across nations (Raymo and Xie 2000; Smits, et al. 1998; Ultee and Luijkx 1990). These studies use cross-sectional data that represent “snapshots” of marriages in a population at particular times and places. These snapshots, however, conceal the underlying demographic processes that determine assortative mating patterns. Cross-sectional marriage data are made up of multiple birth cohorts, that is, groups of married people of similar ages. Each cohort’s life is

structured in ways that potentially affect the degree to which spouses resemble each other. Variation in average educational attainment, marital dissolution rates, marriage timing, and cohabitation rates may all affect assortative mating patterns within a cohort. Thus, cross-sectional assortative mating patterns represent an aggregation over multiple cohorts' assortative mating patterns, each shaped by potentially different and changing demographic factors.

Past research has tried to control for the dynamic nature of assortative mating within cohorts by limiting analyses to newly married couples or couples in their first marriages so that selective marital disruption and educational changes after marriage do not bias the results (e.g., Blackwell 1998; Kalmijn 1991a, 1991b, 1994; Lewis and Oppenheimer 2000; Mare 1991; Qian 1998; Qian and Preston 1993). Studies of cross-sectional historical trends find that the resemblance between newly marrying spouses on educational attainment has increased in the United States over the past several decades (Kalmijn 1991a, 199b; Mare 1991; Qian and Preston 1993), a trend which is particularly pronounced for college graduates (Kalmijn 1991b; Mare 1991). More recent research has incorporated intracohort variation in the entry into first marriages. These studies examine how the probability of entering a homogamous marriage varies by age and how these age patterns have shifted over time in several European countries (e.g., Bernardi, forthcoming; Blossfeld and Timm 1998; Chan and Halpin, forthcoming; Henz and Jonsson, forthcoming).

Although studies that examine assortative mating among newlyweds in first marriages are informative, marital dissolutions and post-marital educational changes may affect the association between the educational characteristics of spouses and thus the conclusions we make about the social distance between groups. For example, the “closeness” of groups at the time of marriage may be offset by high divorce rates. Social groups may be brought closer together

through intermarriage, but high rates of marital dissolution may reinforce social boundaries (Kalmijn 1998). Similarly, the number of marriages that cross social boundaries will decline if people continue their education after marriage to catch up to their partners. Thus, the social distance between groups that one can infer from marriage patterns depends not only patterns of new marriages but also on which couples remain married and the extent to which spouses' education characteristics change after they marry. Although previous research has acknowledged the influence of factors other than first marriage on total assortative mating patterns, no attempt has been made to quantify their effects.

To determine how demographic factors affect assortative mating as people age in this paper, we compare the association between spouses' education characteristics at different points in their lives. In other words, we examine changes in assortative mating in the stock of marriages by age. Next, we examine how the flows into and out of marriage change the nature of the stock of existing marriages. All changes in the overall association between the spouses' characteristics in the stock of marriages can be decomposed into variation in three flows: (1) new marriages (first and later marriages), (2) marital dissolutions, and (3) post-marital educational upgrades. All other factors that influence assortative mating work through one or more of these components. For example, cohabitation will increase the homogamy of new marriages if dissimilar couples break up before rather than after they marry (Blackwell and Lichter 2000). Here, cohabitation may be the root cause of increased homogamy but it is reflected in a change in the degree to which spouses resemble each other at the time of their marriage.

To understand how the components of the assortative mating process combine to make up observed trends in the stock of marriages, we follow the marital careers of a sample of American youth who were 14 and 22 years old in 1979. Once we understand the dynamics of

assortative mating across one cohort's life cycle, we can show how these demographic changes have contributed to changes in assortative mating across time and space and what basic demographic trends portend for the future. This paper develops methods that are suitable for investigating these possibilities.

In the first part of our paper, we review past work that provides clues about how demographic factors affect assortative mating patterns as people age. After describing our data and methods, we examine how assortative mating patterns change across a cohort's life and decompose these changes into their proximate causes. We conclude by generalizing from our intracohort results to consider how demographic factors may affect historical trends in educational assortative mating in the United States.

THE EFFECTS OF MARITAL AND EDUCATIONAL STATUS CHANGES ON EDUCATIONAL ASSORTATIVE MATING

The relative impact of new marriages, marital dissolutions, and post-marital educational upgrading on the degree of resemblance between spouses as they age is determined by two components: (1) their effects on homogamy by age and (2) the frequency with which these events occur. In other words, the impact of each of the flows on the stock of marriages is determined by both their *direction* and *volume*. In this section, we discuss how assortative mating into first and later marriages, selective marriage dissolution, and post-marital educational upgrades may affect the similarity of spouses in the stock of marriages.

First marriages. Prior theory and evidence suggests that the odds of educational homogamy by age may follow an inverted “U” shape in which the odds are low among those

who marry early, high among those who marry in their mid-20s, and low among those who marry late.

The odds of homogamy may be lower among those who marry at older ages because of the changing nature of the marriage market as people age (Lewis and Oppenheimer 2000; Lichter 1990; Mare 1991). As young people leave educational institutions and move into the labor market they may be more likely to encounter potential spouses who do not share their educational attainment (Mare 1991). Educational homogamy may also be lower among people who marry late because of the shrinking availability of potential partners (Lewis and Oppenheimer 2000; Lichter 1990). Still-single men and women may be forced to redefine what constitutes an acceptable match as potential mates are siphoned into marriage.

By contrast, the odds of homogamy among those who marry young may be lower than among those who marry in their mid-20s if young people are more likely match on expected rather than completed education. The odds of crossing an educational boundary may be higher among those who marry young by virtue of the fact that young people are more likely to be in school. Such a result would also be compatible with the hypothesis that the gap between the average age of school completion and first marriage is negatively related to the odds of homogamy (Mare 1991). High school sweethearts who marry after one partner has graduated while the other is still in school may cross an educational boundary when they marry but will be homogamous once the other partner graduates. The opportunities for heterogamy on completed education but homogamy on expected education decline as people age and have completed their schooling. Young people may also be less likely marry homogamously if they have a higher tolerance for educational heterogamy. Young people have shorter work and educational histories than do older adults and thus may be less attached to their own career/life style trajectories when

they marry, which may thereby increase the opportunity for “postmarital socialization” (Oppenheimer 1988). The potential for young spouses to influence each other’s future may decrease the importance of pre-marital sorting on traits such as education compared with individuals with well-established careers and life styles (Oppenheimer 1988:583).

Evidence from the United States suggest that multiple mechanisms may be at work in determining age patterns of educational assortative mating into first marriages. Past research on assortative mating in the U.S. indicates that the odds of homogamy are lower in marriages that are formed after the age of 30 (Lichter 1990), that they decrease with the availability of potential spouses (Lewis and Oppenheimer 2000), and that they are negatively related to the gap between school completion and age at first marriage (Mare 1991). None of these findings are inconsistent with the notion that the odds of homogamy increase among those who marry young. Indeed, other evidence from the U.S. suggests that homogamy increases with age (Qian 1998: 290-91). A potential reconciliation of these findings is that age patterns of homogamy are non-linear and follow an inverted “U” pattern, with low levels of homogamy among those who marry young, higher levels of homogamy at ages at which most people have recently completed their education, and low levels of homogamy among those who marry late.

Because first marriages are far more prevalent than marital dissolutions, remarriages, or post-marital educational upgrades, first marriages may make up the bulk of the change in the stock of marriages by age. However, first marriages may explain less of the age pattern of assortative mating at later ages as the frequency of first marriages drops and as remarriages and dissolutions constitute a larger portion of marital events.

Marital dissolutions. Studies of the effects of education on the odds of marital dissolution provide some insight into the possible impact of marital dissolutions on educational

assortative mating in the stock of still-existing marriages. These studies find that spouses who are less alike are more likely to separate (Bumpass and Sweet 1972; Bumpass, Castro Martin, and Sweet, 1991).¹ If educational differences increase the odds of marital dissolution, then marital dissolutions will increase the resemblance of still-married couples. When divorce is common the positive effects of marital dissolution on homogamy may be substantial.

Remarriages. Remarriages tend to be less homogamous than first marriages (Dean and Gurak 1978; Jacobs and Furstenberg 1986). This appears to partially be the result of a selection effect in which women who have had only one marriage have higher homogamy rates than do remarried women in either of their marriages. Although remarriages tend to be less homogamous than first marriages, age also plays a role. Women who remarry at younger ages tend to have higher levels of homogamy than those who remarry at older ages (Jacobs and Furstenberg 1986).

Post-marital educational upgrading. Educational changes may result in greater homogamy if people who marry before completing their education match on expected rather than completed education. These individuals may be heterogamous at the time of marriage but may become homogamous as the result of later educational changes. Educational changes may also increase homogamy if spouses positively affect one another's educational attainment. If individuals prefer homogamy to heterogamy, partners with less education in heterogamous marriages may feel compelled to increase their education after marriage to match their spouses' levels. Educational changes may result in lower levels of homogamy, however, if couples in homogamous marriages send one partner back to school while the other works as a strategy to increase the overall financial well-being of the family. Alternatively, if educational upgrades

occurred at random, they would reduce the homogamy of homogamous populations of marriages but increase the homogamy of heterogamous populations of marriages.

Part of the impact of educational changes on assortative mating may come through its indirect effect on marital disruption. Educational changes within marriage increase the odds of marital disruption (Davis and Bumpass 1976; Tzeng and Mare 1995). Educational changes may be disruptive in and of themselves or, alternatively, individuals may increase their education in anticipation of divorce. In either case, the effects of post-marital educational changes and marital dissolutions on assortative mating may not be independent. Although these studies alert us to possible interaction effects, they do not reveal whether educational changes result in more or less similarity between spouses.

These strands of research suggest that opposing forces may be at work in the assortative mating process. Marital dissolutions may increase educational homogamy in prevailing marriages, but remarriages may have the opposite effect. The effects of first marriages on the similarity of prevailing marriages may be non-linear. New marriages may increase or decrease the odds of homogamy by age at marriage depending on the timing of marriage and school departure of both spouses and the changing nature of marriage markets. Finally, empirical research to guide our hypotheses about the impact of educational changes is absent. Educational upgrading may increase the odds of homogamy in prevailing marriages if couples who marry heterogamously on expected education become homogamous after they complete their education or they may decrease the odds of homogamy if the bulk of educational upgrades occur among couples who are already homogamous.

DATA AND MEASUREMENT

We use the National Longitudinal Survey of Youth (NLSY79), a panel study of 12,686 men and women aged 14-21 as of December 31, 1978, to conduct our analysis. We restrict our analysis to the cross-sectional sample, which is designed to be representative of the population in the United States aged 14-21 as of December 31, 1978 and is made up of 6,111 respondents. Men and women in this cohort were first interviewed in 1979 and were re-interviewed yearly through 1994 and then biennially through 2000. The NLSY79 provides detailed histories of respondent's marital status and the educational characteristics of both partners. These data make it possible to trace changes in educational assortative mating by age to changes in the types of new marriages that occur as people age, to marital dissolutions, and to post-marital educational changes.

Important asymmetries exist in the NLSY79 marriage data that affect our sample selection and necessitate that we anchor our analysis to the age of NLSY79 respondents. Although it is possible to think about assortative mating as a function of “cohort age” because of the relatively high correlation between spouses' ages, marriages typically include two people who are likely to have different ages. Because the spouses of NLSY79 respondents may not be part of the cohort of youth aged 14-21 as of December 31, 1978 it is necessary to track respondents to preserve the cohort interpretation of the analysis. Furthermore, some of our analyses require information on marriage parity, which is available for NLSY79 respondents but not their spouses. Therefore, to analyze assortative mating among “husbands” and “wives,” rather than among “respondents” and “spouses,” we analyze male and female respondents separately. We refer to “respondents” when discussing the entire sample and “male respondents” and “female respondents” when discussing sex-specific results.²

Because respondent's age is our "clock", we select our sample based on respondent's age and do not place restrictions on spouse's age. We follow male and female respondents from the time they are age 18 until they are 37. Each year that a respondent is married constitutes one observation. For example, a respondent who is married from 18 to 37 appears 20 times in our data whereas a person who was first married at 37 only appears once. Respondents who marry, divorce, and then remarry are not present in the data when they are divorced but are present for each year they are married or remarried. Transforming the respondent-level data into respondent-year data expands our data to 114,279 respondent-years contributed by 6,111 respondents.³ Restricting the data to those years in which respondents were married reduces the sample to 50,888 couple-years, of which male respondents contribute 22,380 and female respondents contribute 28,508. Each person contributes an average of 10.4 married years to the data and 1.3 marriages spells. Marriage spells begin when a respondent either (1) moves from an unmarried state to a married state or (2) moves from an unknown marital state to a married state. Marriage spells end when a respondent either (1) moves from a married state to an unmarried state or (2) moves from a married state to an unknown marital state.

We classify education into five categories based on the number of years of schooling individuals have completed. They are: less than 10 years schooling, between 10 and 11 years, 12 years, between 13 and 15 years, and 16 or more years. Marital status and education information were imputed whenever possible for years in which respondents were not interviewed.⁴ We drop marriage spells in which we were unable to determine the educational attainment of either of the partners.⁵ This reduces our sample size to 22,045 couple-years contributed by men and 27,538 couple-years contributed by women, or by 1.5% and 3.4% respectively. We further delete marriage spells in which either partner reported an education category decline or if either partner

reported that his/her education increased by more than 1 category per year, with the exception of moves from less than 10 years of schooling to 12 years of schooling. We do this to allow for the possibility that these individuals received a GED. This reduces our sample to 45,760 couple-years, of which male respondents contribute 20,278 and female respondents contribute 25,482, or by 8.0% and 7.5% respectively.

Next, we identify couple-years in which new marriages, marital dissolutions, or post-marital educational changes occur. Couple-years with new marriages are those in which respondents reported moving into a first or later marriage or the year in which they reported reuniting with a former spouse.⁶ We separately identify first and later marriages from the stock of all new marriages. We tag the interview year prior to the year that respondents report the death of a spouse, a marital separation, or a divorce as a couple-year in which a marital dissolution occurs. Finally, couple-years with educational changes are those in which the education category of one or both partners is higher than that of the previous year.

In many couple-years, no marital events occur. On average, one marital event occurs in every 5.4 couple-years. Figure 1 shows the frequency distribution of marital events by respondent's age category, each of which are of width two, and respondent's sex (see Appendix Table 1 for the data that produced this figure). It shows that first marriages are far more common in our sample than other events until female respondents are in their late 20s and until male respondents reach their early 30s. Marital dissolutions are the next most common marital event. Educational changes are more frequent than remarriages until respondents are in their early to mid 20s at which point the frequency of remarriage exceeds that of educational change. There are a total of 5,160 marriages and remarriages in the sample and 2,193 marital dissolutions, implying that an average of 43% of marriages in our sample dissolved prior to 2000, at which

point respondents were between 35 and 37 years old. There are 1,131 educational changes in the sample, which implies that 20% of couples experience a joint-educational category increase, on average. Figure 1 also shows that whereas the modal marriage age category for both male respondents and female respondents is 22-23, the frequency of marriage among female respondents between the ages of 18-19 and 22-23 is far higher than frequency of marriage between these ages for male respondents. Furthermore, the incidence of marital dissolutions is much higher among young women than among young men reflecting their earlier age at marriage.

STATISTICAL METHODS

We examine age patterns of educational assortative mating using log-linear models for contingency tables. Our contingency table is produced by cross-classifying couple-years by respondent's education, spouse's education, and respondent's age and sex. We use log-linear models because they allow us to examine the association between couples' education while controlling for changes in the age distribution of education. This feature is particularly important for this analysis because there are large shifts in the education distribution of married couples across the age range we examine. Both the mean and the variance of husband's and wife's education increase with respondent's age.

We use homogamy models of marriage to examine trends in the association between husband's and wife's education by age. These models describe the association between couples' educational characteristics in terms of the odds that husbands and wives have the same rather than different levels of education.⁷ Homogamy models allow us to describe and decompose

trends in assortative mating by age using a single intuitive measure. Homogamy models thus represent “average” assortative mating patterns. More complex models that use multiple parameters to describe assortative mating patterns may provide a better fit to the data, but have limited use for our analysis because they produce age patterns of educational assortative mating that are highly influenced by patterns of age homogamy.⁸

If Y_R and Y_S denote the education category of the highest year of schooling completed for respondents and spouses respectively, homogamy is defined as,

$$d^H = 1 \text{ if } Y_R = Y_S, \text{ otherwise } d^H = 0.$$

Further, let

$$d_i^R = 1 \text{ if respondent's education is in category } i \text{ (} i = <10, 10-11, 12, 13-15, 16+\text{), otherwise } d_i^R = 0,$$

$$d_j^S = 1 \text{ if spouse's education is in category } j \text{ (} j = <10, 10-11, 12, 13-15, 16+\text{), otherwise } d_j^S = 0, \text{ and}$$

$$d_a^A = 1 \text{ if respondents are in age category } a \text{ (} a = 18-19, 20-21, 22-23, 24-25, 26-27, 28-29, 30-31, 32-33, 34-35, 36-37\text{), otherwise } d_a^A = 0.$$

Then a log-linear homogamy model is:

$$\begin{aligned} \log m_{ija} = & \beta + \sum_i \beta_i^R d_i^R + \sum_j \beta_j^S d_j^S + \sum_a \beta_a^A d_a^A \\ & + \sum_{ia} \beta_{ia}^{RA} d_i^R d_a^A + \sum_{ja} \beta_{ja}^{SA} d_j^S d_a^A \\ & + \beta^H d^H + \sum_a \beta_a^{HA} d^H d_a^A \end{aligned} \quad (1)$$

where m_{ija} is the expected number of marriages between respondents in education category i and spouses in education category j by respondent's age category a , and the β s are the parameters to be estimated. This model corresponds to the hypothesis that the odds of homogamy vary with

age net of age-specific variation in the marginal distributions of respondent's and spouse's education as represented by β_{ia}^{RA} and β_{ja}^{SA} .

Although we estimate assortative mating parameters using log-linear models, the test statistics from these models are not valid because our dependent variables, i.e., the cell frequencies of marriages, contain multiple observations per respondent. Our data consist of three hierarchical levels. Each couple-year observation (level 1) is nested within a marriage (level 2) and each of these marriages is nested within a respondent (level 3). Ideally, we would take into account the correlation structure between each of these levels. However, we ignore level 2 clustering in this paper.⁹ To correct for the respondent-level (level 3) clustering it is necessary to use individual-level rather than grouped data. We use multinomial logit models that yield the same coefficients as equation (1) but in which the units are couple-years rather than cell frequencies. This allows us to use the robust cluster option in STATA to correct for the clustering of errors around respondents (see Appendix B for technical details).

RESULTS

Testing for Age Variation in Educational Homogamy in Prevailing Marriages

Table 1 shows a comparison of the relative fit of four homogamy models using multinomial logit models. Because the clustering of the data is not accounted for in model estimation, the usual log likelihood tests are not appropriate for model testing. We therefore perform Wald tests to examine relative model fit. Model 1 fits equation (1). Here, we describe the association between respondent's and spouse's education in terms of the odds of homogamy and do not differentiate by the sex of the respondent. We allow these odds to vary over respondent's age categories. The Wald test of the null hypothesis that the interactions between homogamy and respondent's age

are jointly equal to zero is rejected, which indicates that the change in homogamy by respondent's age is statistically significant.

Model 2 relaxes the assumption that homogamy patterns for male and female respondents are equal by including three-way interaction terms between homogamy, respondent's age, and respondent's sex. Because of small cell sizes we use a quadratic representation of age in the three-way interaction terms. Table 1 shows that the Wald test of the null hypothesis that there is no sex difference in age patterns of homogamy is rejected, which indicates that there are statistically significant differences between age patterns of homogamy among male and female respondents. In addition, we clearly reject the hypothesis of no age variation in educational homogamy within male and female respondent samples (models 3 and 4).

Assortative Mating Patterns in Prevailing Marriages by Respondent's Age

Figure 2 shows variation in the predicted odds of educational homogamy by respondent's age predicted from models 3 and 4. The odds of homogamy increase dramatically by age for both male and female respondents net of changes in the marginal distributions of education. This increase is especially steep between the ages of 18-19 and 22-23 for both sexes. These odds increase by 59% for male respondents between the ages of 18-19 to 22-23 and 58% for female respondents between these ages. Homogamy continues to increase among male respondents until they are 24-25 and then begins to decline gradually until the end of the age interval. By contrast, homogamy increases slowly for female respondents after age 22-23 and through the end of the age interval. Over the entire age span, male respondents move from being 2.0 times as likely to a little less than three times as likely to be in a homogamous marriage rather than a heterogamous

marriage and female respondents move from being 1.7 times as likely to also being a little less than three times as likely to be in a homogamous marriage rather than a heterogamous marriage.

In the next section, we determine whether new marriages, educational changes, and marital dissolutions have positive or negative effects on homogamy in the stock of marriages and how this changes as the cohort ages. Following this, we determine the relative contribution of each of the flows to changes in assortative mating patterns in the stock of marriages as respondents' age.

Changes in Assortative Mating Flows by Respondent's Age

All changes in the overall association between the spouses' characteristics in the stock of marriages can be decomposed into variation in three flows: (1) new marriages (first and later marriages), (2) marital dissolutions, and (3) post-marital educational upgrades. This process can be represented with the following accounting equation:

$$M_{ij(a+1)} = M_{ija} + W_{ija} - D_{ija} + R_{ija} \pm E_{ija} \quad (2)$$

where,

i = husband's education category ($i = <10, 10-11, 12, 13-15, 16+$),

j = wife's education category ($j = <10, 10-11, 12, 13-15, 16+$),

a = respondent's age category ($a = 18-19, 20-21, 22-23, 24-25, 26-27, 28-29, 30-31, 32-33, 34-35, 36-37$), and

M_{ija} = the number of prevailing marriages at age a between husbands of education i and wives of education j ,

W_{ija} = the number of weddings (first marriages) at age a between husbands of education i and wives of education j ,

D_{ija} = the number of marital dissolutions at age a between husbands of education i and wives of education j ,

R_{ija} = the number of remarriages at age a between husbands of education i and wives of education j , and

E_{ija} = the net increase or decrease in the number of marriages in joint education category ij due to educational upgrading.

Thus, prevailing marriages between husbands of education i and wives of education j at age $a + 1$ ($M_{ij(a+1)}$) are made up of the stock of marriages at age a (M_{ija}), plus the number of weddings between ages a and $a+1$ (W_{ija}), minus the number of couples that dissolve between ages a and $a+1$ (D_{ija}), plus the number of remarriages between ages a and $a+1$ (R_{ija}), and plus or minus the net migration of couples into/out of joint education category ij as a result of educational upgrading (E_{ija}).

The overall impact of each of the assortative mating flows on the odds of homogamy in the stock of marriages is determined by both the *direction* and *volume* of the flows. The volume of the flows by respondent's sex is shown in Figure 1. In this section, we examine the *direction* of each of the flows as respondents age by examining the odds of homogamy in each of the components on the right hand side of equation (2) separately. To determine how new marriages affect the odds of homogamy in prevailing marriages, we estimate models 3 and 4 for couples who are in their first year of marriage. The homogamy coefficients from these models indicate how the odds of homogamy vary by age at marriage. We do this separately for first and later marriages. Next, we determine the effects of marital dissolutions on the odds of homogamy in prevailing marriage by running models 3 and 4 from the sample of couples who are in their last year of marriage. The homogamy coefficients from these models reveal whether the couples leaving the stock of marriage are more or less homogamous than those that are currently married. Finally, we determine the effects of educational changes on the odds of homogamy in prevailing marriages by estimating the ratio of the odds of homogamy in the year of an educational upgrade

to the odds of homogamy in the year prior to the upgrade for the sample of couples that experienced an upgrade. A ratio of greater than 1 indicates that educational upgrading increases the odds of homogamy in prevailing marriages whereas a ratio of less than 1 indicates that it decreases the odds of homogamy in prevailing marriages.¹⁰

Figure 3 shows the age pattern of homogamy for male and female respondents for each component of educational assortative mating. Age patterns of homogamy into first marriages are inverted “U” shaped for both sexes. The odds of homogamy increase for male respondents between the ages of 18-21 and 22-25, plateau, and decline after age 30-33. By contrast, the odds of homogamy for female respondents increase between the ages of 18-21 and 26-29 and decline thereafter. These results are consistent with the notion that multiple mechanisms are at work in determining age patterns of educational homogamy into first marriages.

The odds of homogamy among female respondents entering into remarriages increase at young ages and then decrease sharply. Homogamy is lower among male respondents between the ages of 18-21 and 22-25 than among female respondents, although very few young men remarry at these ages.¹¹ What is most striking about patterns of remarriage for both sexes is that, with the exception of remarriages among female respondents in the 22-25 age category, the odds of homogamy are generally much lower than those of first marriages and prevailing marriages. These findings are consistent with previous research, which has shown that remarriages tend to be less homogamous than first marriages (Dean and Gurak 1978; Jacobs and Furstenberg 1986). Because respondents entering into the stock of marriages through remarriage have lower levels of homogamy than couples that are already there, remarriages will tend to decrease the odds of homogamy in prevailing marriages. The extent to which this occurs will be determined in the next section.

Figure 3 also shows that marital dissolutions increase the homogamy of prevailing marriages for both male and female respondents. Although age patterns of homogamy for male and female respondents in their last year of marriage are markedly different, the odds of homogamy for both sexes in this sample are generally lower than the odds of homogamy in prevailing marriages. Because those couples that leave the stock of marriages through marital dissolutions are less homogamous than couples in prevailing marriages, marital dissolutions will tend to increase the odds of homogamy in prevailing marriages. This is consistent with previous research, which finds that marital dissolutions are more likely to occur among couples that are heterogamous (Bumpass and Sweet 1972; Bumpass, et al. 1991).

Unlike first marriages, remarriages, and marital dissolutions, the direction of the impact of educational upgrading on homogamy as people age differs by the sex of the respondent. For male respondents, the ratio of the odds of homogamy in the year of the educational upgrade to the odds of homogamy in the year prior is consistently less than 1, which indicates that the odds of becoming homogamous as the result of an educational upgrade are lower than the odds of becoming heterogamous. Thus, educational upgrades will tend to decrease the homogamy of prevailing marriages for male respondents. By contrast, the ratio of these odds for female respondents is consistently greater than 1, which indicates that the odds of becoming homogamous as the result of an educational upgrade are higher than the odds of becoming heterogamous. Thus, educational upgrades will tend to increase the homogamy of prevailing marriages for female respondents.

The sex differences in the direction of the effects of educational upgrading may be the result of sex differences in the timing of educational completion and marriage but they could also be a reflection of sex differences in the odds of homogamy in prevailing marriages. If

educational changes within marriage occur at random, educational changes will reduce the odds of homogamy in homogamous marriage populations but will increase the odds of homogamy in heterogamous marriage populations. The negative effect of educational upgrades for male respondents and the positive effects for female respondents, then, may be partially the result of the higher homogamy levels of male respondents compared with female respondents.

In the next section, we examine how the direction and the volume of the flows combine to produce changes in the stock of marriages. This allows us determine the relative importance of new marriages, marital dissolutions, and educational upgrading for changes in educational assortative mating by age.

The Impact of Changes in Assortative Mating Flows on Assortative Mating Patterns in the Stock of Marriages

We decompose age patterns of assortative mating by conducting a series of simulations that alter the observed marriage data. The goal of the simulations is to obtain estimates of the impact of one marital event on changes in assortative mating in prevailing marriages that are independent of the other marital events. In doing so, we make the simplifying assumption that the components of change are additive.

We determine the impact of marital events on changes in assortative mating patterns in prevailing marriages by age by successively increasing the number of marital events that make up the stock of marriages. We perform four simulations to conduct our decomposition using our accounting equation for the stock of marriages (equation 2). If $M_{ij(a+1)}$ is the number of marriages between husbands of education i and wives of education j at respondent's age $a + 1$, then our simulations are as follows:

$$M_{ij(a+1)} = M_{ija} + W_{ija} \quad (S1)$$

$$M_{ij(a+1)} = M_{ija} + W_{ija} + R_{ija} \quad (S2)$$

$$M_{ij(a+1)} = M_{ija} + W_{ija} + R_{ija} \pm E_{ija} \quad (S3)$$

$$M_{ij(a+1)} = M_{ija} + W_{ija} - D_{ija} + R_{ija} \pm E_{ija} \quad (S4)$$

where all terms are as defined previously.

In simulation (S1), the expected stock of marriages among husbands of education i and wives of education j at respondent's age $a + 1$ is the sum of the number of existing marriages at age a between husbands and wives in the ij^{th} education category and the number of first marriages that occur between husbands of education i and wives of education j between respondent's age a and age $a + 1$ (W_{ija}). This calculation creates a new table in which the stock of marriages is only comprised of new marriages. We then estimate models 3 and 4 using these simulated data. The coefficients from this model estimate the odds of homogamy if first marriages were the only component of change in age patterns of educational assortative mating. Another way of thinking about this is that simulation (S1) restricts changes in the stock of marriage to changes that are the result of one flow, i.e., first marriages.

In simulation (S2), the stock of marriages changes as a result of both first marriages (W_{ija}) and remarriages (R_{ija}). As in the first simulation, we estimate models 3 and 4 using the simulated data. The difference in the log odds of homogamy estimated from simulations (S2) and (S1) is the effect of remarriages on assortative mating patterns.

In simulation (S3), changes in the stock of marriages are composed of first and later marriages and educational upgrades. In this simulation, we calculate the net migration into/out of the ij^{th} education category due to educational upgrading (E_{ija}). After accounting for the redistribution due to educational upgrading between ages a and $a + 1$, we add new marriages and remarriages occurring between husbands of education i and wives of education j between the

ages of a and $a + 1$ to the stock of marriages. The difference between the log odds of homogamy estimated from the data from simulation ($S3$) and those estimated from simulation ($S2$) represents the effect of educational upgrading on assortative mating patterns.

Finally, in simulation ($S4$), the stock of marriages is composed of all of its components: first and later marriages, educational upgrades, and marital dissolutions. To conduct this simulation, we performed the same calculations as in simulation ($S3$), but subtract couples in the ij^{th} education category who dissolve their marriages between age a and $a + 1$. The difference in the homogamy parameters estimated from simulations ($S4$) and ($S3$) represents the effect of marital dissolutions on assortative mating patterns.

Figure 4 shows the impact of the flows on the odds of homogamy in the stock of marriages by respondent's age and sex. The trend for prevailing marriages is constructed from the homogamy coefficients from simulation ($S4$), which contains all of the parts of the assortative mating process.¹² Each of the other lines represents age patterns of educational homogamy had only one of these flows been responsible for all of the change in assortative mating. Table 2 shows the decomposition of the total change in the odds of homogamy that are accounted for by first marriages, dissolutions, remarriages, and educational changes. Column C of Table 2 shows how much higher/lower the odds of homogamy at age 36-37 would have been had only one marital event been responsible for the changes in the odds of homogamy relative to the odds of homogamy in prevailing marriages at age 36-37. Column F shows the percentage of the change in the log odds of homogamy in prevailing marriages across the age range we examine that are attributable to each of the marital events.

Figure 4 and Table 2 show that first marriages make up by far the largest part of the change in the odds of homogamy in prevailing marriage by age for both sexes. First marriages

make up 145% of the change in the log odds of homogamy in prevailing marriages for male respondents and 105% of the change for female respondents (Table 2, Column F). Had first marriages been the only flow into or out of marriage, the odds of homogamy would have been 18% higher than they were at the end of the age interval for male respondents and would have been essentially the same (3% higher) for female respondents (Table 2, Column C). This indicates that impact of the drop in the odds of homogamy among newlyweds entering their first marriages after the age of about 30 (Figure 3) on the odds of homogamy in prevailing marriages is muted by the relative infrequency of marriages that occur after this point (Figure 1).

Marital dissolutions, remarriages, and educational changes all have smaller effects on changes in the odds of homogamy. Marital dissolutions have a positive impact on the homogamy of the stock of marriages for both sexes. Column F of Table 2 shows that marital dissolutions account for 17% of the increase in the log odds of homogamy for male respondents but that marital dissolutions have a very small effect for female respondents. Remarriages have a larger impact on changes in homogamy by age than marital dissolutions for both male and female respondents. If remarriages were the only marital events to have occurred over this age interval, the log odds of homogamy would have decreased by 24% among male respondents and by 16% among female respondents over the age interval that we examine. For male respondents, educational changes have a relatively large negative impact on the change in homogamy. If educational changes were the only marital events to occur over this age interval, the log odds of homogamy among male respondents would have declined by 38%. By contrast, educational changes account for 10% of the increase in the log odds of homogamy among female respondents.

Figure 4 also shows that sex differences in age patterns of homogamy in prevailing marriages can be traced to sex differences in the odds of homogamy in first marriages. Figure 4 shows that the initial difference in the odds of homogamy between male and female respondents is due to differences in the homogamy of first marriages: male respondents who first marry at age 18-19 are more likely to enter a homogamous union than are female respondents who marry in this age category. The rapid increase in the odds of homogamy among male respondents after age 18-19 is almost exclusively due to changes in the homogamy of first marriages. By contrast, the increase in the odds of homogamy among female respondents and is primarily due to changes in the homogamy of first marriages but is also due to the positive effects of marital dissolutions. The odds of homogamy for male respondents are propelled upward by the steep increase in the odds of homogamy in first marriages and are kept higher than the odds of homogamy for female respondents by the momentum of the high levels of homogamy of earlier marriages. The odds of homogamy in prevailing marriages for male respondents, however, are steadily eroded by the negative impact of remarriages, educational changes, and marriages occurring at later ages until they reach parity with female respondents at age 34-35 (Figure 2). By contrast, the odds of homogamy for female respondents remain high as they age because the positive effects of educational changes and marital dissolutions are almost completely offset by the negative effects of remarriages.¹³

To summarize, the stock of marriages is made up of flows into and out of marriage of different directions and strengths. Age patterns of educational homogamy into first marriages make up the overwhelming majority of the trend in the odds of homogamy as the cohort ages. Although the odds of homogamy drop among men and women who marry at older ages, this does not have a large impact on the homogamy of stock of marriages because the majority of

marriages occur at younger ages with higher levels of homogamy. Remarriages offset the increasing odds of homogamy in the stock of marriages and marital dissolutions reinforce these trends whereas the impact of educational changes differs by sex.

DISCUSSION AND CONCLUSIONS

This analysis highlights the potential importance of demographic factors other than first marriage in explaining variation in total educational assortative mating patterns across time and place. Changes in educational assortative mating in the United States have previously been attributed to broad societal changes, i.e., growing individualism and changes in the meaning of marriage (Kalmijn 1991a, 1991b; Oppenheimer 1994), or to a more limited set of demographic factors, i.e., the gap between age at school completion and age at first marriage (Mare 1991). But the magnitude of *intracohort* variation is far larger than the largest historical changes in educational assortative mating. The most dramatic historical change in educational assortative mating over the past several decades has been the growing tendency of college graduates to marry each other. Between 1940 and 1987, the odds of marriage between college graduates and non-college graduates decreased by 20% in the United States (Mare 1991, Table 4). Trends in the odds of homogamy in first marriage over this period are less dramatic. Between 1940 and 1960 the odds of marrying homogamously dropped by 8% but increased by 8% from 1960 to 1985-1987 (calculations from Mare 1991, Table 4).¹⁴ The intracohort increase in the odds of homogamy by age is very large compared with both these historical trends. The odds of homogamy in prevailing marriages increase by 41% for male respondents and 67% for female respondents across the age interval we examine. Changes in the odds of homogamy among newlyweds are

also large. The odds of homogamy among newlyweds entering into their first marriages increase by 28% for male respondents and 69% for female respondents between the ages of 18-21 and 26-29.

The vast majority of the increase in the odds of homogamy by age is the result of the increasing odds of homogamy in first marriages at young ages. We find that the odds of homogamy among husbands and wives entering their first marriages follow an inverted “U” shape. These results are consistent with the hypothesis that homogamy is highest upon school completion and declines thereafter (Mare 1991). If respondents are more likely to be in school at younger ages, they may also be more likely to marry across educational boundaries.

Alternatively, homogamy may be lower at younger ages if young people have a higher tolerance for heterogamy because their career paths and lifestyles are less established and more uncertain than are older people’s (Oppenheimer 1988). The ages at which the odds of homogamy are highest are those at which most people have completed their schooling. After about the age of 30, the odds of homogamy begin to decline. This is consistent with the notion that people revise their preferences for homogamy as the number of eligible mates dwindles (Lewis and Oppenheimer 2000; Lichter 1990) and with the hypothesis that the odds of educational heterogamy increase as people’s marriage markets are increasingly structured by one’s occupation or interests rather than by educational institutions (Mare 1991).

Changes in the odds of homogamy in first marriages in this cohort provide grounds for speculating about intercohort trends in educational homogamy. Suppose that the inverted “U” shape of the odds of homogamy among first-married newlyweds is constant over time. Then as the average age at first marriage moves from 18-21 to 22-25, the average odds of homogamy in prevailing marriages and among newlyweds will increase. If the average age at marriage

increases past about age 30, however, average levels of homogamy in prevailing marriages and among newlyweds will decline. At present, the average age at first marriage is moving toward the downward slope of the inverted “U.” Marital status life tables for 1995 imply that the average age of first marriage is 28.6 for men and 26.6 for women (Schoen and Standish 2001). This suggests that assortative mating levels may decrease among cohorts to come.

Suppose, however, that the shape of the inverted “U” is not constant over time, but is affected by changes in marriage timing. This would occur if the number of eligible partners of a given age affects the odds of homogamy. For example, if the observed odds of homogamy decline once respondents are older than about 30 because the pool of eligible mates is increasingly sparse, we would expect that as the number of eligible mates increase at older ages, so would the odds of homogamy. This would shift the peak of the inverted “U” to the right. The shape of the inverted “U” could also be affected by the timing of educational completion. For example, if the observed odds of homogamy are low when respondents are young because they are more likely to be in school than older respondents, then a continued expansion of education will also shift the peak of the inverted “U” to the right.

While the majority of age patterns of assortative mating can be explained by age patterns of homogamy among couples entering their first marriages, the effects of marital dissolutions, educational changes, and remarriages are not trivial. Marital dissolutions have small positive effects on changes in the odds of homogamy, remarriages have larger negative effects, and educational changes have a substantial negative effect for male respondents but a smaller positive effect for female respondents. Increases in divorce and remarriage over the past several decades (Bumpass, Sweet, and Castro Martin 1990; Castro Martin and Bumpass 1989; Goldstein 1999) and growing school enrollment rates among married persons (Bumpass and Call 1989)

suggest that factors other than first marriages may be increasingly important in explaining the overall degree of resemblance between spouses' educational attainment.

Our findings suggest the usefulness of revisiting previous trend studies of educational assortative mating in the United States. Demographic factors such as marital dissolutions and post-marital educational changes may explain a significant portion of change in the degree to which spouses have resembled each other across time. Such an analysis may either strengthen or weaken the conclusions about the “openness” of society depending on how these demographic factors work together in affecting the degree of spousal resemblance in previous cohorts. This paper provides a framework for understanding how flows into, out of, and within marriage affect the degree of resemblance between spouses in the stock of all marriages. It provides a method for understanding how the intracohort components of assortative mating fit together to make up the assortative mating in prevailing marriages and thereby allows for the exploration of how changes in demographic factors such as divorce, remarriage, and education have affected historical trends in educational assortative mating.

APPENDIX A. DERIVATION OF MARITAL STATUS AND EDUCATIONAL CHARACTERISTICS FOR MISSING INTERVIEW YEARS

The NLSY79 contains information on current marital status as well as up to three marital status changes since the respondent's last interview. Both the type and the date of the change are recorded. We use the marital change information to "fill in" marital histories and educational characteristics of respondents and spouses in years in which respondents were not interviewed, including years in which the survey was not administered, i.e., 1995, 1997, and 1999.

Because spouse's education and age information is only available at the time of the current interview and not retrospectively since the last interview, it was necessary to develop assumptions to impute a couple's characteristics in years in which the respondent was not interviewed. First, when we imputed marital histories, we assumed that age increases one year for every missing survey year. Second, we assume that a couple's education in missing interview years is equal to their education at the first valid interview year after the missing interview spell if a new marriage occurred during the missing interview years. In other words, we impute backwards. Third, we assume that a couple's education in missing interview years is equal to their education at the last valid interview after the missing interview spell if a divorce occurred during the missing interview spell. In other words, we impute forward. The second and third assumptions introduce potential biases by ruling out the possibility of education change over the imputed period. This means we may underestimate the role of educational change in these marriages.

Fourth, we assume that a couple's education in missing interview years is equal to their education at the current interview year if the couple remained married from the last to the current

interview, or we impute backwards in time. We do this under the assumption that if educational changes occur, they are more likely to do so earlier rather than later in life. For these cases, we may be underestimating the age of the educational change. We could not impute the couple's characteristics for respondents who had a complete marital spell between interviews, i.e., a marriage and a dissolution. Thus, these marriages do not appear in our data.

In addition to imputing respondent's and spouse's education and marital history information for missing interview years, we implemented the following rules when dealing with missing or inconsistent spouse's education data. First, cases in which spouse's highest grade completed is reported as "ungraded" were assigned the previous highest grade completed reported. This follows the method used for respondent's education described in Appendix 8 of the NLSY79 Codebook Supplement. Second, if the respondent is married to the same spouse for three or more years and the spouse has valid education data before and after the year(s) with missing education data and the spouse's education is the same before and after the missing data spell, we replaced the missing years with the value for spouse's education in the adjacent years. Third, if the respondent is married to the same spouse and there is valid data on spouse's education for three consecutive years before a single year of missing data, then the missing year is equal to spouse's education in the previous year. Fourth, if a spouse's education decreases by more than 8 years in a single year but the adjacent two years are equal, then the errant year is replaced with the adjacent year's value.

To assess the impact of these decisions on our results we conducted our analysis without marriage spells in which we made potentially "controversial" imputations. We defined marital spells with controversial imputations as those in which we imputed two or more consecutive years of educational information and in which either the respondent's or the spouse's education

category was not the same at the beginning and end of the missing data spell. Dropping these spells does not alter the interpretation of the results report here (results available from authors upon request).

APPENDIX B. CORRESPONDENCE BETWEEN LOG-LINEAR AND MULTINOMIAL LOGIT HOMOLOGY MODELS

To correct for the respondent-level clustering it is necessary to use individual-level rather than grouped data. We use multinomial logit models that yield the same coefficients and standard errors as equation (1) but in which the units are couple-years rather than cell frequencies. This allows us to use the robust cluster option in STATA to correct for the clustering of errors around respondents.¹⁵ If Y_{RS} is the joint distribution of respondent's and spouse's education for respondent r with spouse s and Y_{RS} has k categories ($k = 1, \dots, 25$), then a multinomial logit model is:

$$\log \left[\frac{P(Y_{RS} = k | X)}{P(Y_{RS} = K | X)} \right] = \beta_k + \sum_a \beta_{ka}^A d_{ra}^A \quad (2)$$

where

$$d_{ra}^A = 1 \text{ for respondent } r \text{ in age category } a \text{ (} a = 18-19, 20-21, 22-23, 24-25, 26-27, 28-29, 30-31, 32-33, 34-35, 36-37\text{), otherwise } d_a^A = 0,$$

the k categories are defined as in Appendix Figure 1 and the β s are the parameters to be estimated.

Recall, equation (1) is the homology model:

$$\begin{aligned} \log m_{ija} = & \beta + \sum_i \beta_i^R d_i^R + \sum_j \beta_j^S d_j^S + \sum_a \beta_a^A d_a^A \\ & + \sum_{ia} \beta_{ia}^{RA} d_i^R d_a^A + \sum_{ja} \beta_{ja}^{SA} d_j^S d_a^A \\ & + \beta^H d^H + \sum_a \beta_a^{HA} d^H d_a^A \end{aligned} \quad (1)$$

The multinomial logit model in (2) is not equivalent to the model in (1) without imposing a series of constraints on the coefficients. Note that the log-linear model in (1) estimates 99 parameters:

1 constant,

4 for the respondent education marginals $(I - 1)$,

4 for the spouse education marginals $(J - 1)$,

9 for the respondent age marginals $(A - 1)$,

36 for the interaction between respondent's age and education $[(I - 1)(A - 1)]$,

36 for the interaction between spouse's age and education $[(J - 1)(A - 1)]$,

1 for the main effect of homogamy, and

9 for the interaction between homogamy and respondent's age $(A - 1)$.

= 99 parameters,

whereas the unconstrained multinomial logit model in (2) estimates a total of 240 parameters:

24 constants $(K - 1)$

216 age effects $[(K - 1)(A - 1)]$.

= 240 parameters.

To replicate the results of model (1) with model (2), we determine the structure of odds ratios in the homogamy model and apply it to the multinomial logit coefficients. The design matrix for the homogamy model is shown in Appendix Figure 2.

Using the design matrix in Appendix Figure 2, we can derive the constraints we need to obtain equivalence with model (1). Let a_{RS} equal the log of the odds ratio where RS denotes the lower right cell of the log odds ratio considered. Then a complete accounting of the log odds ratios in Figure 3 is:

1. $a_{22} = (1 + 1) - (0 + 0) = 2$
2. $a_{23} = (0 + 0) - (1 + 0) = -1$
3. $a_{24} = (0 + 0) - (0 + 0) = 0$
4. $a_{25} = (0 + 0) - (0 + 0) = 0$
5. $a_{32} = (0 + 0) - (0 + 1) = -1$
6. $a_{33} = (1 + 1) - (0 + 0) = 2$
7. $a_{34} = (0 + 0) - (1 + 0) = -1$
8. $a_{35} = (0 + 0) - (0 + 0) = 0$
9. $a_{42} = (0 + 0) - (0 + 0) = 0$
10. $a_{43} = (0 + 0) - (0 + 1) = -1$
11. $a_{44} = (1 + 1) - (0 + 0) = 2$
12. $a_{45} = (0 + 0) - (1 + 0) = -1$
13. $a_{52} = (0 + 0) - (0 + 0) = 0$
14. $a_{53} = (0 + 0) - (0 + 0) = 0$
15. $a_{54} = (0 + 0) - (0 + 1) = -1$
16. $a_{55} = (1 + 1) - (0 + 0) = 2$

The log odds ratios given above reveal the relationship between the log odds ratios in a homogamy model. We therefore constrain the log odds ratios such that:

1. $a_{22} = a_{33}$
2. $a_{22} = a_{44}$
3. $a_{22} = a_{55}$
4. $a_{23} = a_{32}$
5. $a_{23} = a_{34}$

6. $a_{23} = a_{43}$
7. $a_{23} = a_{45}$
8. $a_{23} = a_{54}$
9. $a_{24} = 0$
10. $a_{25} = 0$
11. $a_{35} = 0$
12. $a_{42} = 0$
13. $a_{52} = 0$
14. $a_{53} = 0$
15. $a_{22} = -2*a_{23}$

which gives us 15 constraints of 16 log odds ratios. The free parameter describes the association between respondent's and spouse's education, in this case, in terms of the likelihood of being on the diagonal versus off the diagonal, or the likelihood of being homogamous. The constraints above indicate that the log odds of homogamy are equal to $a_{22} \div 2$ or, equivalently, $-a_{23}$.

These constraints directly apply to the coefficients in equation (2), which are the log odds of being in a particular joint education category (Y_{RS}) with reference to the omitted category.

When cell 1 of Figure 2 ($k = 1$) is the omitted equation, the constraints given above translate to the following 15 constraints on the model's constant terms:

1. $\beta_7 - (\beta_6 + \beta_2) = (\beta_{13} + \beta_7) - (\beta_{12} + \beta_8)$
2. $\beta_7 - (\beta_6 + \beta_2) = (\beta_{19} + \beta_{13}) - (\beta_{18} + \beta_{14})$
3. $\beta_7 - (\beta_6 + \beta_2) = (\beta_{25} + \beta_{19}) - (\beta_{24} + \beta_{20})$
4. $(\beta_8 + \beta_2) - (\beta_7 + \beta_3) = (\beta_{12} + \beta_6) - (\beta_{11} + \beta_7)$
5. $(\beta_8 + \beta_2) - (\beta_7 + \beta_3) = (\beta_{18} + \beta_{12}) - (\beta_{17} + \beta_{13})$
6. $(\beta_8 + \beta_2) - (\beta_7 + \beta_3) = (\beta_{24} + \beta_{18}) - (\beta_{23} + \beta_{19})$
7. $(\beta_8 + \beta_2) - (\beta_7 + \beta_3) = (\beta_{20} + \beta_{14}) - (\beta_{19} + \beta_{15})$
8. $(\beta_8 + \beta_2) - (\beta_7 + \beta_3) = (\beta_{24} + \beta_{18}) - (\beta_{23} + \beta_{19})$
9. $(\beta_9 + \beta_3) - (\beta_8 + \beta_4) = 0$
10. $(\beta_{10} + \beta_4) - (\beta_9 + \beta_5) = 0$
11. $(\beta_{15} + \beta_9) - (\beta_{14} + \beta_{10}) = 0$

12. $(\beta_{17} + \beta_{11}) - (\beta_{16} + \beta_{12}) = 0$
13. $(\beta_{22} + \beta_{16}) - (\beta_{21} + \beta_{17}) = 0$
14. $(\beta_{23} + \beta_{17}) - (\beta_{22} + \beta_{18}) = 0$
15. $\beta_7 - (\beta_6 + \beta_2) = -2 * [(\beta_8 + \beta_2) - (\beta_7 + \beta_3)]$

We impose the same set of constraints on the age coefficients, i.e., the β_{ka}^A s, to allow the homogamy parameter to vary by respondent's age. The homogamy parameter for the omitted age category is equal to $[\beta_7 - (\beta_6 + \beta_2)] \div 2$ or $-[(\beta_8 + \beta_2) - (\beta_7 + \beta_3)]$.

We apply these 15 constraints to the multinomial logit model's constants, and 135 constraints to the respondent's age coefficients $[15*(A - 1)]$ for a total of 150 constraints. This yields a multinomial logit model with 90 parameters $(240-150 = 90)$. As mentioned above, we estimate 99 parameters in the log-linear model. However, 9 of these parameters represent the coefficients for the marginal effects of respondent's age, which have no effect on Y_{RS} and are thus not relevant to the multinomial logit model. Thus, applying these constraints to the multinomial logit model in (2) gives us the same estimates of assortative mating as the log-linear model in (1) but allows us to adjust the standard errors for the level 3 clustering.

We estimate these models using the "ml" command for maximum likelihood estimation in STATA. At present, the packaged multinomial logit command "mlogit" cannot handle complex constraints on the constant terms.

ENDNOTES

¹ Tzeng and Mare (1995) do not find an effect of educational differences on marital disruption.

These results are not necessarily inconsistent with other work. Tzeng and Mare measure educational differences in years whereas Bumpass and Sweet (1972) and Bumpass, Castro Martin, and Sweet (1991) measure educational differences in terms of levels, e.g., college graduate versus non-graduate or high school graduate versus non-graduate. Thus, the negative effects of larger educational differences could be hidden by Tzeng and Mare's assumption of linearity.

² Although we examine assortative mating processes among both male and female respondents and each of these processes involves both a husband and a wife, our results for male and female respondents are not merely two sides of the same coin. The results from the female sample are the mirror image of the male sample for two primary reasons. First, our male and female respondent samples are independent. Thus, one source of asymmetry between the male and female respondent samples may stem from sex differences in the way in which men and women respond to questions about their own and their spouses' education. Further, sampling variation alone may lead to asymmetries between the two samples. Second, as discussed below, we do not restrict spouses' ages in either sample. Doing so would be undesirable because it may increase age homogamy at the tails of the age interval we examine, which could artificially increase educational homogamy at these ages. Thus, asymmetry between the male and female respondent samples may also stem from sex differences in spouse's age. Because men tend to marry younger women, many of the young marriages included in the male respondent sample do not have a counterpart in the female respondent sample and many of the older marriages in the female

respondent sample do not have a counterpart in the male respondent sample. Even when both partners are between the ages of 18 and 37, however, age patterns of educational homogamy may differ by sex of respondent because of sex differences in the timing of educational completion and marriage. These issues are discussed in reference to the results in endnote 13 below.

³ The number of respondent-years is not equal to $6,111 \times 20$ (one respondent year for each year between the ages of 18 and 37) because some respondents were older than 18 at the beginning of the interview period and some respondents were younger than 37 at the end of the period. The full NLSY79 sample is made up of individuals who were between 14 and 22 in 1979, or who were born between 1957 and 1965. Restricting the sample to respondents aged 18 to 37 means that we do not observe the full cohort of NLSY79 respondents for all ages. Specifically:

- the 18 year olds in our analysis were born between 1961 and 1965,
- the 19 year olds were born between 1960 and 1965,
- the 20 years olds were born between 1959 and 1965, and
- the 21 year olds were born between 1958 and 1965.

Therefore, our analysis of assortative mating among respondents between the ages of 18 to 21 is weighted towards older sample members. Similarly, our analysis of assortative mating among respondents who are 36 and 37 is weighted towards younger sample members. Specifically:

- the 36 year olds in our analysis were born between 1957 and 1964, and
- the 37 year olds were born between 1957 and 1963.

This introduces the possibility that differences in assortative mating by birth cohort may affect our results. To test this possibility, we ran a model analogous to model 2 presented in Table 1 but replaced the male respondent dummy variable with a cohort dummy variable. We define an

“old cohort” as those born between 1957 and 1960 and a “young cohort” as those born between 1961 and 1965. We find no evidence of a three-way association between birth cohort, respondent’s age, and the odds of marital homogamy (Wald Chi-square = 0.29, d.f. = 2).

⁴ Appendix A gives an overview of how we determined the marital status and education levels of respondents and spouses for years in which respondents were not interviewed.

⁵ The NLSY79 main data file contains revised versions of respondent’s education, which are used here. The revised versions are created by using additional survey information to correct grade “reversals” and to fill in ambiguous or incomplete education histories (see NLSY79 Supplement, Appendix 8). No such revised variables exist for spouse’s education and many of the additional variables to help correct inconsistencies or ambiguities for respondents are not available. We implemented a series of decision rules for imputing/recoding inconsistent or missing spouse’s education values (see Appendix A for details).

⁶ It is possible to separately identify reunions and formal remarriages although we do not do so here. Reunions account for only 17.4% of remarriages.

⁷ Recent studies of marriage in several European countries, which are similar in spirit to ours although examine only first marriages, use a different methodology (e.g., Bernardi, forthcoming; Blossfeld and Timm 1998; Chan and Halpin, forthcoming; Henz and Jonsson, forthcoming). These studies take a competing risks approach to the transition to marriage in which singlehood ends in one of three ways: in a homogamous marriage, in a heterogamous marriage to someone with more education, or in a heterogamous to someone with less education. Thus, for example,

these models calculate the probability of entering a homogamous marriage rather than remaining single at a given age. Age patterns in the probability of entering a homogamous marriage are affected by two factors: (1) the odds of marriage at that age and (2) the odds of homogamy at that age given marriage. Our goal is to examine the latter, not the former, and therefore log-linear models of married couples are appropriate for the present endeavor.

⁸ For example, crossings models are a common alternative to homogamy models and have been found to fit marriage data well (e.g., Blackwell 1998; Johnson 1980; Mare 1991). These models represent the association between husband's and wife's education as a series of barriers to marriage between educational groups; that is, they measure the odds of marriage across an educational boundary. However, patterns of age homogamy strongly affect changes in crossings parameters by age. For instance, it is far less likely that an 18-year-old high school graduate will marry someone who has completed college than it is that a 28-year-old high school graduate will marry someone who has completed college. This is because of age homogamy: the 18-year-old is likely to marry someone else around the age of 18 who is also likely not to have completed college whereas the 28-year-old is likely to marry someone else around the age of 28 who has a far greater chance than the 18-year-old of having completed college. A way to solve this problem is to control for assortative mating on age. This is undesirable in the present analysis, however, because our goal is to examine trends in the overall association between couples' education characteristics as they age, not the trend in this association net of assortative mating on age.

⁹ See Appendix B for a discussion of our reasons for ignoring level 2 clustering.

¹⁰ Running separate analyses for each flow greatly reduces our sample sizes (see Appendix Table 1 for sample sizes). These samples cannot support the detailed age classification used in the previous section. In this portion of the analysis, therefore, we use the log-linear model in equation (1), but collapse age into five categories ($a = 18-21, 22-25, 26-29, 30-33, 34-37$).

¹¹ Because there are only 4 remarriages among male respondents between the ages of 18 and 21 (see Appendix Table 1), we were unable to run the full age classification for the male respondent remarriage sample. Thus, the results presented in Figure 3 for the male respondent remarriage sample are run from a model in which the first two age categories have been collapsed, i.e., $a = 18-25, 26-29, 30-33, 34-37$.

¹² The data produced by simulation (S4) are not equivalent to the observed data. The observed data contain marriage spells for which we were unable to determine the year in which the marriage started or ended. We were unable to determine the start date of a respondent's marriage spell if missing interview years preceded it and inadequate information was given about the start date of the marriage. Likewise, we were unable to determine the end date of a respondent's marriage spell if missing interview years followed it and inadequate information was given about the end date of the marriage or if the respondent dropped out of the survey. Of the 45,760 couple-years in our sample, 174 or 0.4% are couple-years in which we do not have the marriage start date of the marriage spell (excluding the first year of the interview period) and 3,414 or 7.5% are couple-years in which we do have marriage end dates (excluding the last year of the interview period). The odds of homogamy in prevailing marriages in Figure 2 are calculated

from the observed data tables whereas the odds of homogamy in prevailing marriages in Figure 4 are calculated from the data constructed in simulation (*S4*). It was necessary to use the data constructed from simulation (*S4*) in the decomposition portion of our analysis to ensure that the sum of the change in the odds of homogamy of the components equals the total change in prevailing marriages. However, the odds of homogamy calculated from the observed data are very similar to those calculated from simulation (*S4*). An alternative approach would be to expand our accounting equation (equation 2) to include censored entry and exit from marriage.

¹³ There are several reasons why the odds of educational homogamy may be higher at younger ages for male respondents than for female respondents. First, sex differences in assortative mating by age may account for the sex differences in educational assortative mating we observe. Men traditionally marry younger women, but there is a floor to this tendency as state laws and social customs forbid persons younger than a certain age to marry. This means that an 18-year-old male has a smaller range of socially acceptable ages from which to pick a spouse than does an 18-year-old woman. Indeed, results from our sample show that 18 and 19 year old men marry women who are slightly older or the same ages as themselves, whereas 18 and 19 year old women marry men who are approximately 3.5 years older than themselves. To the extent that age homogamy translates into educational homogamy, sex differences in age homogamy may account for sex differences in educational homogamy. To investigate this possibility, we ran our log-linear models controlling for assortative mating by age using several representations of assortative mating on age. In all of these models, sex differences in age patterns of educational assortative mating in prevailing marriages persist virtually unchanged, which suggests that sex

differences in assortative mating on age do not account for sex differences in educational assortative mating in our sample.

Another possibility is that sex differences in sample attrition may explain the sex differences in assortative mating in prevailing marriages we observe. The results presented here are not weighted to correct for sample attrition. However, when we weight the sample sex differences in assortative mating in prevailing marriages remain. The sex differences also do not appear to be the result of decisions we made about imputing incomplete educational histories (see Appendix A) or our decision to drop marital spells with incomplete or invalid education data. Sex differences in age patterns of educational homogamy are smaller but persistent in a model run off of the original data.

Finally, there is suggestive evidence that differences in how men and women report their own and their spouses' educational attainment may contribute to the sex differences in educational assortative mating we observe in prevailing marriages. The NLSY79 sample contains a small number of couples in which both the husband and the wife are survey respondents. Our sample contains 60 couples and 667 couple-years in which both husbands and wives report their own and their spouse's education. Our analysis of these data reveals that husbands are somewhat more likely to report that their marriages are homogamous than are their wives, which implies that reporting differences may be responsible for some portion of the sex differences in age patterns of educational homogamy. (All results available from authors upon request.)

¹⁴ The odds of homogamy were 3.3 times the odds of heterogamy in 1940, decreased to 3.0 in 1960, and then increased monotonically in 1970 and 1980 until reaching 3.3 in 1985-87.

¹⁵ As mentioned in the text, our data consist of three hierarchical levels. Each couple-year observation (level 1) is nested within a marriage (level 2), and each of these marriages is nested within a respondent (level 3). In this analysis, we ignore level 2 clustering for several reasons. First, the robust cluster option in STATA cannot handle more than one level of clustering. To circumvent this pitfall, it is theoretically possible to estimate a random effects multinomial logit model using STATA's program to fit Generalized Linear Latent and Mixed Models, or "GLLAMM." GLLAMM allows for the inclusion of random effects at multiple levels, which would account for both the level 2 and the level 3 clustering. However, this would require estimating 2 random effects for each of the 24 equations (K-1) in our multinomial logit model and their covariances (see text below for an accounting of the multinomial logit parameters). Given our present computing capabilities, this task is untenable. Another alternative would be to include fixed effects for marriage parity in our model. This approach is conceptually inappropriate because it would prevent us from decomposing age trends into parts due to first and later marriages. In any case, the level 2 clustering should not have a large effect on our standard errors because 85% of the couple-years in our sample are years in which respondents are in their first marriages. We chose to correct for clustering at level 3 rather than at level 2 under the assumption that it is more egregious to assume that the errors across marriages within the same individual are independent than it is to assume that respondents share a common error across marriages.

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**TABLE 1. GOODNESS-OF-FIT STATISTICS FOR SELECTED MULTINOMIAL LOGIT MODELS, NLSY79
(STANDARD ERRORS ADJUSTED FOR CLUSTERING AROUND RESPONDENTS)**

Model	N	Pseudo-Log Likelihood	Wald test	Chi-square (d.f.)	p-value
1. RA, SA, HA	45,760	-113147	HA = 0	33.92 (9)	0.0001
2. RA, RM, AM, SA, SM, RA _Q M, SA _Q M, HA _Q M	45,760	-112857	HA _Q M = 0	6.55 (2)	0.0379
3. For male respondents: RA, SA, HA	20,278	-50255	HA = 0	19.98 (9)	0.0180
4. For female respondents: RA, SA, HA	25,482	-62571	HA = 0	31.36 (9)	0.0003

NOTES: R = respondent's education category; S = spouse's education category; A = Respondent's age category, H = homogamy dummy variable, M = male respondent dummy variable, A_Q = a linear and quadratic term for respondent's age category.

TABLE 2. DECOMPOSITION OF CHANGES IN HOMOGAMY BY RESPONDENT'S AGE AND SEX, NLSY79

Respondent's Sex and Marital Event	Odds			Log odds		
	(A)	(B)	(C)	(D)	(E)	(F)
	Age category		Odds at 36-37 by Marital Event ÷ Odds at 36-37 in Prevailing Marriages	Age category		Change in Log Odds by Marital Event ÷ Change in Prevailing Marriages
	18-19	36-37		18-19	36-37	
Male respondent						
Prevailing marriages	2.00	2.90	1.00	0.69	1.07	1.00
First marriages	2.00	3.44	1.18	0.69	1.23	1.45
Remarriages	2.00	1.83	0.63	0.69	0.60	-0.24
Marital dissolutions	2.00	2.13	0.73	0.69	0.76	0.17
Educational changes	2.00	1.74	0.60	0.69	0.55	-0.38
Female respondent						
Prevailing marriages	1.69	2.82	1.00	0.52	1.04	1.00
First marriages	1.69	2.89	1.03	0.52	1.06	1.05
Remarriages	1.69	1.55	0.55	0.52	0.44	-0.16
Marital dissolutions	1.69	1.70	0.60	0.52	0.53	0.02
Educational changes	1.69	1.77	0.63	0.52	0.57	0.10

FIGURE 1. FREQUENCY OF MARITAL EVENTS BY RESPONDENT'S AGE AND SEX, NLSY79

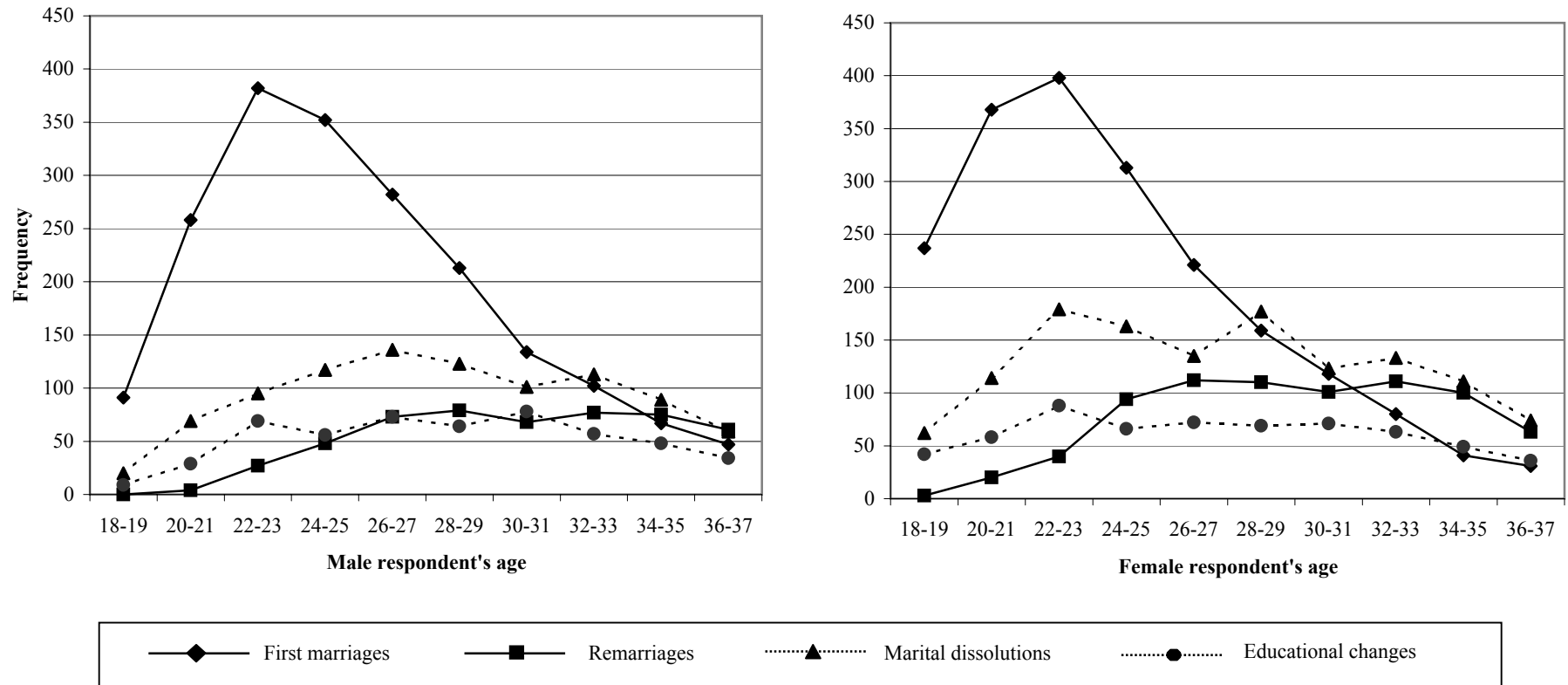


FIGURE 2. ODDS OF HOMOGAMY IN PREVAILING MARRIAGES BY RESPONDENT'S AGE AND SEX, NLSY79

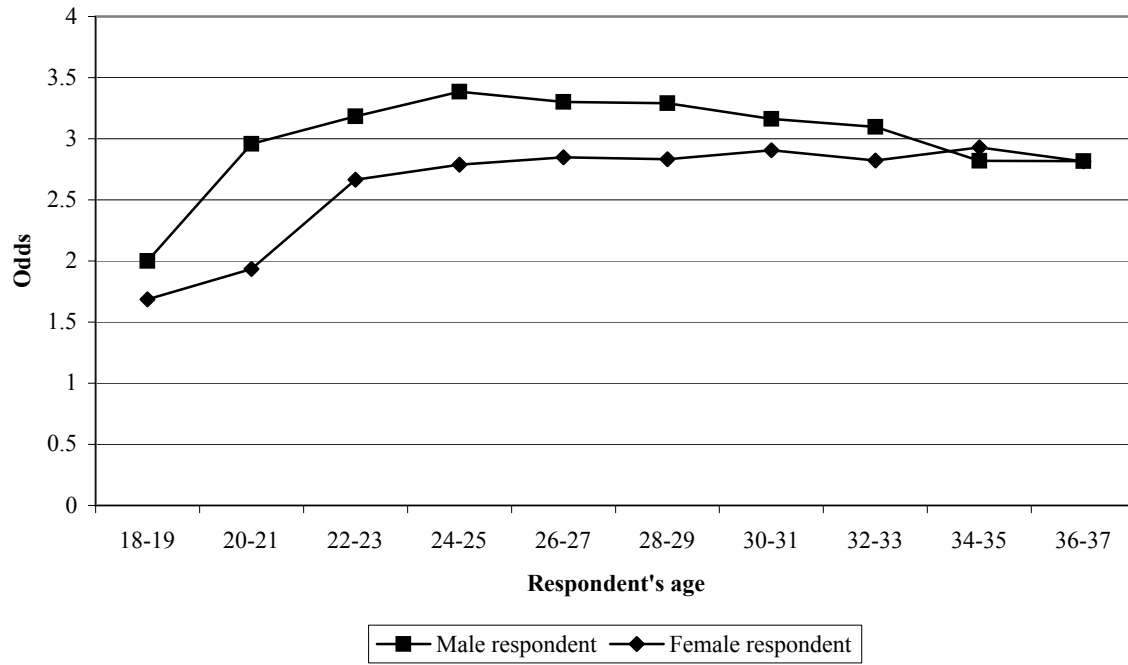


FIGURE 3. ODDS OF HOMOGAMY BY MARITAL EVENT AND RESPONDENT'S AGE AND SEX, NLSY79

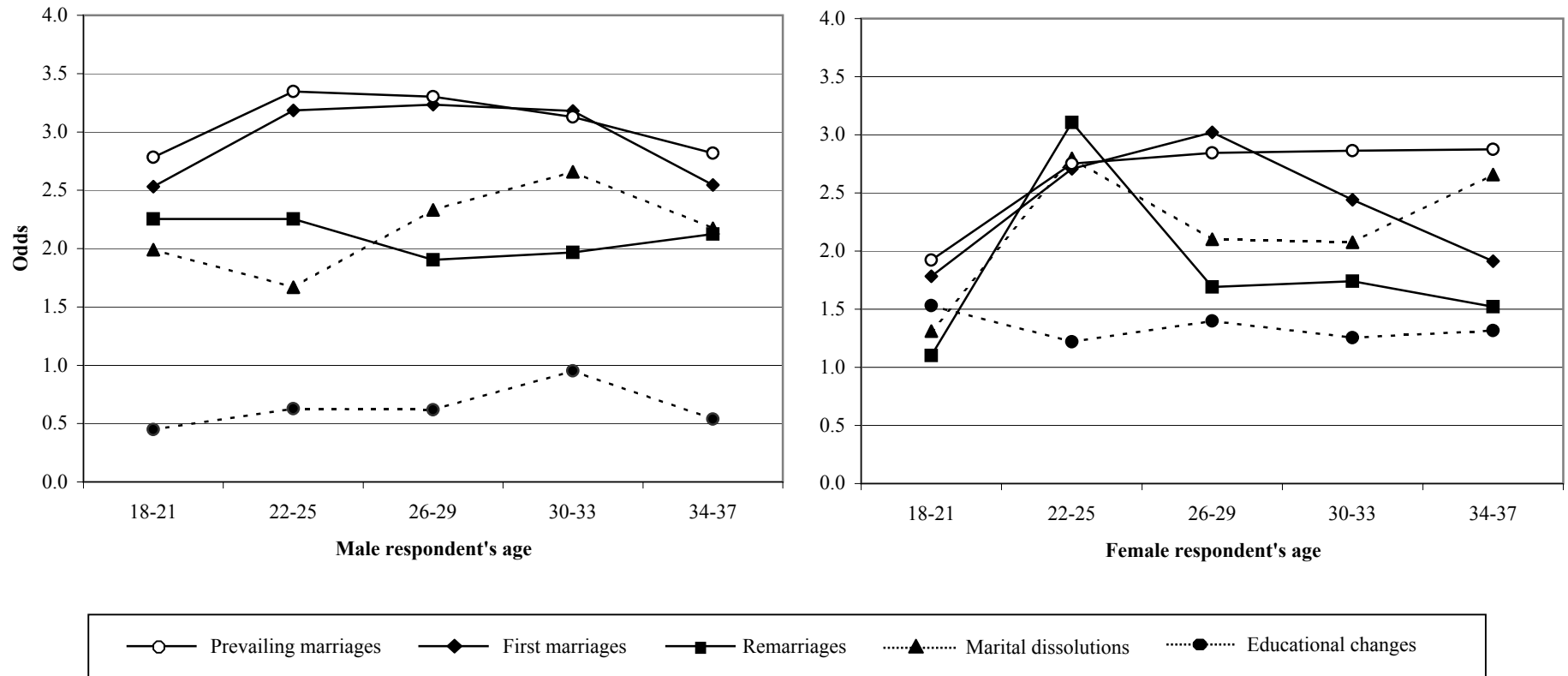
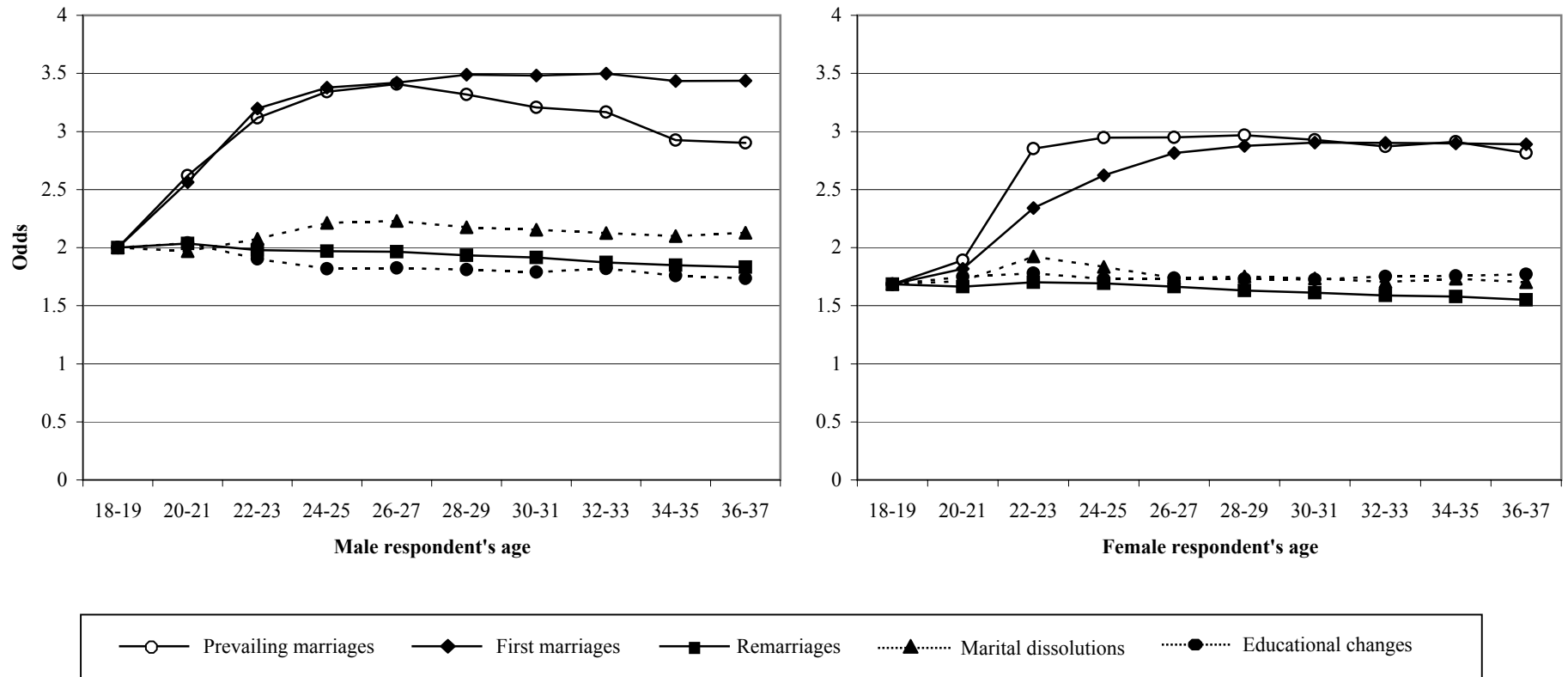


FIGURE 4. EXPECTED ODDS OF HOMOGAMY UNDER ALTERNATIVE ASSUMPTIONS BY RESPONDENT'S AGE AND SEX, NLSY79



APPENDIX TABLE 1. DISTRIBUTION OF MARITAL EVENTS BY RESPONDENT'S AGE AND SEX, NLSY79

Marital Event	Respondent's Age										Total
	18-19	20-21	22-23	24-25	26-27	28-29	30-31	32-33	34-35	36-37	
Male respondents											
First marriage	91	258	382	352	282	213	134	102	67	47	1,928
Remarriage	0	4	27	48	73	79	68	77	75	61	512
Marital dissolution	20	69	95	117	136	123	101	113	89	59	922
Post-marital Educational change	9	29	69	56	73	64	78	57	48	34	517
Total	120	360	573	573	564	479	381	349	279	201	3,879
Female respondents											
First marriage	237	368	398	313	221	159	118	80	41	31	1,966
Remarriage	3	20	40	94	112	110	101	111	100	63	754
Marital dissolution	62	114	179	163	135	177	123	133	111	74	1,271
Post-marital Educational change	42	58	88	66	72	69	71	63	49	36	614
Total	344	560	705	636	540	515	413	387	301	204	4,605

APPENDIX FIGURE 1. DEFINITION OF Y_{RS}

	Y_S				
Y_R	1	2	3	4	5
1	1	2	3	4	5
2	6	7	8	9	10
3	11	12	13	14	15
4	16	17	18	19	20
5	21	22	23	24	25

APPENDIX FIGURE 2. DESIGN MATRIX FOR HOMOGAMY MODEL

	Y_S				
Y_R	1	2	3	4	5
1	1	0	0	0	0
2	0	1	0	0	0
3	0	0	1	0	0
4	0	0	0	1	0
5	0	0	0	0	1