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PATHWAYS TO EDUCATIONAL HOMOGAMY IN MARITAL AND COHABITING UNIONS

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PRELIMINARY DRAFT

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ABSTRACT

This study uses log-linear models and data from the National Longitudinal Survey of Youth (NLSY79) to compare the odds of educational homogamy in marital and cohabiting unions. Differences in the educational resemblance of married and cohabiting couples vary depending on the sample used. Cohabiting couples are less likely to be educationally homogamous than married couples using a sample of prevailing unions. Restricting the sample to newly formed unions, however, eliminates this difference. Furthermore, I find little support for the hypothesis that cohabiting couples who transition to marriage are more homogamous than cohabiting couples who separate, although these results vary by respondent's sex. My results suggest that differences in educational homogamy by union type in prevailing unions are driven by the accumulation of the most homogamous marriages over time rather than differences in sorting into unions or selection out of cohabiting unions into marriage.

INTRODUCTION

The dramatic increase in cohabitation in the United States over the past several decades has changed the social environment in which individuals make decisions about whether, when, and whom to marry. Differences in partner selection in cohabitation and marriage and the impact of cohabitation on the resemblance between spouses, however, are poorly understood. One common hypothesis about the possible effect of cohabitation on the resemblance between spouses is the "double selection" hypothesis, coined by Blackwell and Lichter (2000) but found elsewhere in the literature as well (e.g., Gwartney-Gibbs 1986:432; Sahib and Gu 2002). According to this hypothesis, married couples will be more likely to be homogamous than cohabiting couples because many of them have had two opportunities to accept or reject each other rather than just one. As cohabitors live together they gain new information about each other that they may not have otherwise had. This new information reduces uncertainty about the match and may lead to the dissolution of heterogamous matches thereby increasing the selectivity of marriage. Thus, the double selection hypothesis posits a "demographic winnowing process that successively selects individuals into cohabiting unions and then into marriages" (Blackwell and Lichter 2000:297).

Alternatively, the relative levels of homogamy in marital and cohabiting unions may vary depending on the characteristic in question. Some authors posit that married couples will be more homogamous on ascribed characteristics (such as race, age, and religion) but less homogamous on achieved characteristics (such as earnings and education) than cohabitors. Because cohabitors tend to value egalitarianism and are more tolerant of nontraditional family roles than married couples (e.g., Clarkberg, Stolzenberg, and Waite 1995), they may be more likely to match with someone with similar earnings and education levels but may be more tolerant of other differences among partners. Cohabiting unions may also be more heterogamous than marital unions if individuals choose to cohabit rather than marry when their relationships violate social norms against marriage across social boundaries (Capser and Bianchi 2002), pressures which may be stronger for ascribed characteristics such as race and religion than achieved characteristics such as education and earnings.

Economic models of marriage also predict that cohabitors will be less homogamous on ascribed characteristics but more homogamous on achieved characteristics than married couples. If the gains from marriage come from economic specialization and cohabitation is "looser bond" than marriage, then cohabitors may be more economically similar than married couples but less similar with respect to other characteristics (Jepsen and Jepsen 2002; Schoen and Weinick 1993). Such a process would tend to select the *less* homogamous cohabiting couples into marriage with respect to earnings and education and the *more* homogamous cohabiting couples into marriage with respect to other characteristics.

Past studies of assortative mating by union type have largely relied on cross-sectional data and thus have not provided a complete test of the double selection hypothesis (Blackwell and Lichter 2000; Jepsen and Jepsen 2002; Qian 1998; Qian and Preston 1993; Schoen and Weinick 1993). These studies generally find that cohabitors are less similar than married couples with respect to ascribed characteristics such as race/ethnicity, religious background, and age (Blackwell and Lichter 2000; Casper and Bianchi 2002; Gullickson 2004; Schoen and Weinick 1993; Jepsen and Jepsen 2002) and are more similar than married couples with respect to earnings and employment (Casper and Bianchi 2002; Brines and Joyner 1999). However, the findings with respect to education vary widely. Using data from the late 1980s and early 1990s, one study finds that cohabiting couples are more educationally similar than married couples

(Schoen and Weinick 1993), whereas another finds the opposite (Blackwell and Lichter 2000), and still others find no difference (Jepsen and Jepsen 2002; Qian 1998) or that the results vary by educational level (Blackwell and Lichter 2004). Differences in data sources, sample selection, methodology, and model choice may all explain these disparate results (Smock 2000).

Studies that have used longitudinal data to examine transitions out of cohabitation and into marriage cast doubt on the double-selection hypothesis with respect to education. Of the studies that examine the joint education characteristics of cohabitors, one finds that only cohabitating couples with large educational differences are more likely to separate than to marry (Smock and Manning 1997:337) while two others find no significant effect of educational differences on the likelihood of splitting up or marrying (Oppenheimer 2003:133; Sassler and McNally 2003:Table 3). Similarly, in a study of unmarried parents, Goldstein and Harknett (2004) find no evidence of the effects of educational differences on the likelihood of splitting up or marrying. These studies suggest that if marital unions are indeed more or less likely to be homogamous than cohabiting unions, it may be because of differences in the odds of homogamy between married couples who cohabit prior to marriage and those who marry without first cohabiting rather than because of differences in the resemblance of cohabitors who split up and those who marry. However, research on differences in educational homogamy between married couples who have and who have not cohabited prior to marriage is limited (Blackwell and Lichter 2004) and no study has examined the contribution of both selective exits from cohabitation and entry into marriage simultaneously.

This paper has two objectives. The first is to begin to resolve the disparate findings of past research on differences in educational homogamy by partner type. To do this, I use log-linear models and data from the National Longitudinal Survey of Youth (NLYS79). The

NLSY79 contains rich information on respondents' cohabitation and marital histories as well as spouse's and partner's educational characteristics for over a 20 year period. These data are comparatively new and research utilizing the information on the joint education characteristics of cohabiting and marital partners in these data is rare (Oppenheimer 2003). In addition to examining differences in the educational resemblance of couples by union type, I use the NLSY79 to determine the possible effects of differences in sample selection on the results of past studies. The samples used in past research have varied widely, ranging from analyses of all of the cohabitors and married couples in the population at a given time (Jepsen and Jepsen 2002; Spanier 1983), to cohabiting and married couples within a relatively narrow age range (Blackwell and Lichter 2000), to newly formed cohabiting and marital unions (Schoen and Weinick 1993). Because the NLSY79 contains information on multiple relationship transitions for the same individuals, it is ideal for assessing the effects of using these different samples on our conclusions about differences in partner resemblance by couple type. Finally, I corroborate my results using data from the Current Population Survey (CPS) (used by Qian 1998 and Spanier 1983) where possible.

Second, this paper uses the NLSY79 to directly test the hypothesis that couples who enter marriage via cohabitation are "doubly selected" on educational homogamy. I test both (1) whether cohabiting couples who split up are less likely to be homogamous than those who marry and (2) whether married couples who have cohabited prior to marriage are more likely to be homogamous than those who have not cohabited. Thus, this paper brings new longitudinal data to bear on the question of how differences in educational assortative mating vary by union type and the extent to which these differences may be the result of variation in the way in which cohabitors and married couples sort into and out of their relationships.

MEASURING ASSORTATIVE MATING IN COHABITATION AND MARRIAGE

Past research has generally relied on one of two approaches in comparing differences in assortative mating by union type. Either a sample of all couples in cohabiting and marital unions within a given age range, or *prevailing unions,* is used (Blackwell and Lichter 2000; Jepsen and Jepsen 2002; Spanier 1983) or researchers attempt to narrow the sample to *newly formed unions* (Qian 1998; Qian and Preston 1993; Schoen and Weinick 1993). The two types of samples may yield different results because of differences in the duration of cohabiting and marital unions, the selective dissolution of unions, and as a result of educational changes that occur after marriages and cohabitations begin (Schwartz and Mare 2003). Past scholars have found that the resemblance of married couples tends to increase as cohorts age due to the selective attrition of heterogamous couples from marriage (Schwartz and Mare 2003; Kalmijn 1991:505-506) and thus comparisons of homogamy between prevailing marriages and cohabiting unions may overstate differences that exist at the time of union entry.

Unfortunately, the data sources with the largest sample sizes are also those that do not contain information on the start dates of cohabiting and marital unions, which prevents the identification of new unions. The Census ceased collecting date of marriage information in 1980 and has never collected date of cohabitation onset. Thus, studies that use the 1990 and 2000 Census attempt to restrict their analyses to recently formed cohabiting and marital unions by examining relatively young couples (Blackwell and Lichter 2000; Gullickson 2004) or examine a wider cross-section of all unions (Jepsen and Jepsen 2002). As such, their comparisons may be affected by differences in the duration of cohabiting and marital unions, the selective survival of marital and cohabiting relationships, and educational upgrading after union formation.

Other studies compare the educational resemblance of married and cohabiting couples using data from the Current Population Survey (CPS). Although the Census included the "unmarried partner" category in its "relationship to the household" question in 1990, the CPS did not add this category until 1995. Research using the CPS in the pre-1995 period has therefore relied on indirect methods of identifying cohabitors (Qian 1998; Qian and Preston 1993; Spanier 1983). Nevertheless, those that use the June supplement of the CPS are able to restrict their analyses to newlyweds as these data contain marriage start date information through 1995 (Qian 1998; Qian and Preston 1993). Because cohabiting unions tend to be short-lived, a comparison of newlyweds and prevailing cohabiting unions may be relatively comparable to a comparison of newlyweds and newly formed cohabiting unions (Qian 1998:280).

Unlike other studies, Schoen and Weinick (1993) and Blackwell and Licther (2004) are able to measure the educational attainment of both partners at the beginning of their unions without relying on indirect methods of identifying cohabitors using the 1987-88 wave of the National Survey of Families and Households (NSFH) and the 1995 National Survey of Family Growth (NSFG), respectively. These studies are able to more precisely identify assortative mating into unions, but do not contain complete relationship histories with which identify the mechanisms by which married couples come to be more or less homogamous than cohabiting couples.[1] In sum, various data constraints have limited past researchers' ability to isolate the characteristics of spouses and cohabiting couples at the time at which they began their unions and to examine the contribution of transitions into and out of unions on the educational resemblance of couples.

In the present study, I use the NLSY79 data to bring new data to bear on the basic question of whether cohabiting or married couples are more educationally homogamous, determine how these results vary according to the sample used (i.e., prevailing unions vs. newly formed unions), and test whether cohabitors who split up are less homogamous than those who marry (the double-selection hypothesis) or whether couples who marry "cold" are more likely to be homogamous than those who enter marriage via cohabitation. In addition, because some previous studies have examined patterns of assortative mating in prevailing unions controlling for differences in their age distributions (Qian 1998; Schoen and Weinick 1993) whereas others have not (Blackwell and Lichter 2000, 2004), I examine the effects of controlling for differences in age distributions by union type throughout.

DATA

Overview

I use data from the National Longitudinal Survey of Youth (NLSY79) to examine differences in educational homogamy by union type and corroborate my results with the March Current Population Survey (CPS) where possible. The NLSY79 is a nationally representative sample of 12,686 American youth aged 14 to 21 as of December 31, 1978. Sample members in this cohort were interviewed yearly beginning in 1979 through 1994 and then every other year since then. This paper focuses on the period from 1979 to 2000.

The NLSY79 consists of three subsamples: a cross-sectional sample designed to be representative of American youth aged 14-21 as of December 31, 1978, an oversample of Hispanic, black, and poor non-black, non-Hispanic youths, and a military oversample. I exclude the poor non-black, non-Hispanic subsample and the military subsample from the analysis because they were not interviewed after 1990 and 1985, respectively, and thus their marital histories are truncated. There are 9,763 respondents in the cross-sectional sample and the black and Hispanic oversamples.

Advantages and Limitations of the NLSY79

The NLSY79 contains rich information on respondent's cohabitation and marital histories as well as spouse's and partner's educational characteristics throughout the interview period. These data make it possible to follow respondents' assortative mating "careers" over an extended period of time. Other commonly used data sets with rich cohabitation and marriage data are not nationally representative, such as the National Longitudinal Survey of the High School Class of 1972 (NLSHS72), have a limited follow-up period, such as the 1987-88 and 1990-94 waves of the National Survey of Families and Households (NSFH), or gather retrospective relationship histories which may be susceptible to duration bias and recall error, such as the 1995 National Survey of Family Growth (NSFG). By contrast, the NLSY79 is a prospective study in which respondents have been interviewed for over 20 years. This makes it possible to follow individuals through multiple cohabitation and marital transitions. Furthermore, the NLSY79 cohabitation data has been found to correspond well to data from other sources (Haurin 1994:21).

A disadvantage of the NLSY79 is that marriages and cohabiting unions that begin and end between interview years are missed because data on spouse's and partner's education and on respondent's cohabitation status is only consistently available at the time of the interview.[2] Although short-term cohabiting and marital unions are present in the data if they correspond with the survey date, they will be underrepresented relative to cohabitations and marriages of longer duration. This problem is likely to be more severe for cohabiting than marital unions as they are typically of shorter duration. The undercount of short-term cohabiting unions will affect my results if short-term cohabiting unions have significantly different assortative mating patterns than longer-term unions. This may be especially problematic in years in which the survey was administered every other year (1994-2000).[3]

Data on cohabitation start dates available beginning in 1990 provide information on the percentage of short-term cohabitations ending in marriage that are missed, but provide no information on the percentage of cohabitations ending in separation that are missed. Using these data, I find that about 30% of marriages between the ages of 18 and 37 that begin with cohabitation last for less than one interview year, which is similar to estimates from the 1987-88 NSFH (Bumpass and Sweet 1989: Table 4). As Oppenheimer (2003:133) argues, however, very short-term cohabitations are likely to be engagements and "the determinants of these marriages were probably very similar to the direct transitions to marriage by noncohabitors, which are being analyzed." Thus, the process of assortative mating into engagements may be very similar to assortative mating into marriage among non-cohabitors.[4]

Sample Selection

NLSY79. I use three samples from the NLSY79 in the current study. First, to determine differences in educational homogamy among *prevailing unions* by union type, I select a sample of marriages and cohabiting unions in which both partners are between 18 and 37 years of age.[5] Transforming the respondent-level data into person-years results in a sample of 214,786 observations (9,763*22 years). Restricting the data to couples in unions in which both partners are between the ages of 18 to 37 reduces the sample to 72,186, of which 63,559 are married couple-years and 8,627 are cohabiting couple years. I classify education into four categories according to the number of years of schooling completed. These are: less than 12 years of schooling; 12 years of schooling; 13 to 15 years of schooling; and 16 or more years of schooling. Couple-years with missing education information are dropped, resulting in a final sample size to 71,666 couple-years, of which 63,244 (88%) are married couple-years and 8,422 (22%) are cohabiting couple-years. These couple-years are contributed by 7,797 persons, which implies that the average respondent contributes 8.1 years of marriage and 1.1 years of cohabitation to the sample.

Second, to examine differences in *transitions* into and out of cohabitation and into marriage, I select only those couples in the first and last years of their cohabiting unions and couples in first year of their marriages. I identify 5 transition points ($T = 1,\ldots,5$) where $t = 1$ in the first couple-year in which a new cohabitation is reported; $t = 2$ in the last couple-year of a cohabiting union that ends in union dissolution; $t = 3$ in the last couple-year of a cohabiting union that ends in marriage; $t = 4$ in the first couple-year in which a new marriage (either a first or a later marriage) is reported and in which the respondent cohabited with his/her partner in the previous year; and $t = 5$ in the first couple-year in which a new marriage is reported but in which the respondent was not cohabiting with his/her spouse in the previous year. I identify 14,997 such transitions. I exclude couple-years for respondents who have transitioned from cohabitation to marriage but for whom the first year of their marriage or cohabiting spell is unobserved so that my comparisons of transitions are not affected by differences in sample characteristics. This reduces the transition sample to 14,662 (or by 2.2%) of which 3,869 (26%) are new cohabiting unions, 1,931 (13%) are cohabitation dissolutions, 1,495 (10%) are cohabitation exits to marriage, 5,823 (40%) are "cold" marriages, and 1,495 (10%) are marriage entries via cohabitation. These transitions are contributed by 7,092 respondents, which implies that the average respondent contributes 2.1 transitions to the data. Third, my sample of *new unions* is a subsample of the transition sample. It contains couple-years in which a new cohabitation occurred (3,869) or in which either a new marriage via cohabitation or a new "cold" marriage occurred (5,823+1,495=7,318).

Table 1 shows the percent distribution of couples exiting cohabitation and entering marriage by transition type. About half of cohabiting unions dissolve and half marry across most of the age range examined here. The exception is 34 to 37 year old women for whom about 60% of cohabiting unions end in separation rather than marriage. By contrast, the proportion of marriages that are preceded by at least one interview year of cohabitation increases by age at marriage. The percentage of marriages preceded by cohabitation rises from 8.6% among women aged 18 to 21 to more than 35% among women over age 30. Overall, about 22% of marriages are preceded by cohabitation, much lower than the 39% estimated from NSFH data for marriages contacted between 1980 and 1984 (Bumpass and Sweet 1989). Again, this difference is likely due to the fact that cohabiting unions in the NLSY79 can only be identified at the time of the interview.

March CPS. Finally, I use data from the March CPS to corroborate the results of my analysis of prevailing unions. To do this, I select a sample which matches the NLSY79 sample as closely as possible. I use data from the period in which NLSY79 sample members were interviewed (1979-2000) and select cohabiting and married couples in which both partners are between the ages of 18 and 37 and in which either the male or the female partner was 14 to 22 in 1979. Cohabitation must be inferred from individuals' marital status and living arrangements prior to 1995, which was the first year in which the CPS included an "unmarried partner" category in its "relationship to the householder" question. I do so for the entire period using the

POSSLQ methods ("Partners of the Opposite Sex Sharing Living Quarters") outlined by Casper and Cohen (2000). POSSLQ defines cohabiting households as those that "contain two and only two adults (age 15+) who are unrelated and of the opposite sex" (Casper and Cohen 2000:237). Using these methods, I identify 6,634 cohabiting couple-years and 58,745 married couple-years.[6]

METHODS

I examine differences in educational homogamy by union type using log-linear models for contingency tables. My contingency table for prevailing unions is produced by cross-classifying couple-years by female partner's education (<12, 12, 13-15, ≥16), male partner's education (<12, 12, 13-15, ≥16), female partner's age (18-21, 22-25, 26-29, 30-33, 34-37), and union type (marriage, cohabitation), which results in a 4 X 4 X 5 X 2 = 160 cell table. To model differences by type of union transition, I select couple-years in which a transition occurs and cross-classify couple years by female partner's education, male partner's education, female partner's age, and transition type (T = new cohabitation, exit to marriage, exit to dissolution, entry from cohabitation, cold marriage), which results in a 4 X 4 X 5 X 5 = 400 cell table.[7]

I focus primarily on a single measure of the difference between assortative mating in marital and cohabiting union: the odds of homogamy. Homogamy parameters describe the association between couples' education in terms of the odds that male and female partners have the same rather than different levels of education. Although homogamy models may not fit the data as well as more complex representations of the association, my goal is to provide a straightforward and easily interpretable measure of the difference between the two groups. I supplement my analysis of homogamy differences in two ways. First, I use crossings models to determine whether the odds of educational intermarriage vary by education level. Crossings models represent the association between partners' education as a series of barriers union formation between educational groups, or in terms of the relative permeability of boundaries between adjacent educational groups (e.g., Johnson 1980; Mare 1991). I also add terms that capture asymmetries in the association between male and female partner's education by sex. Past research has generally found a tendency for husbands to "marry down" with respect to education (hypergamy) but have found that this pattern is weaker or reversed in cohabiting unions (Blackwell and Lichter 2000, 2004; Casper and Bianchi 2002; Schoen and Weinick 1993; but see Qian 1998). I test whether these patterns are evident in the NLSY79.

Because the goal of this paper is to describe *differences* in the educational resemblance of couples by union type and by union transitions, I start with a baseline model that contains no terms for differences in assortative mating by couple type but saturates the lower order interactions. This model can thought of as the model of "conditional independence" from which I evaluate departures by union type (Raymo and Xie 2000; Xie 1998). Formally, the baseline model is:

$$\begin{aligned}\log(\mu_{ijkl} / t_{ijkl}) = \lambda + \lambda_i^M + \lambda_j^F + \lambda_k^A + \lambda_l^U + \lambda_{ik}^{MA} + \lambda_{jk}^{FA} + \lambda_{kl}^{AU} + \lambda_{il}^{MU} + \lambda_{jl}^{FU} + \\ \lambda_{ikl}^{MAU} + \lambda_{jkl}^{FAU} + \lambda_{ijk}^{MFA}\end{aligned} \quad (1)$$

where *M* denotes male partner's education ($i = 1,\ldots,4$), *F* is female partner's education ($j = 1,\ldots,4$), *A* is female partner's age ($k = 1,\ldots,5$), and *U* is union type ($l = 1,2$). Thus, μ_{ijkl} is the expected number of unions between males in education category *i* and females in education category *j* in female's age category *k* and union type *k*. This model captures variation in the distribution of male and female partner's education by female partner's age and union type

(λ_{ikl}^{MAU} and λ_{jkl}^{FAU}), includes unrestricted interaction terms between male and female partner's education and age (λ_{ijk}^{MFA}), and contains all lower order terms. I weight the results using the NLSY79 1979 weight and March CPS household weight by using an offset term (t_{ijkl}) that is equal to the inverse of average weight in the cell (Agresti 2002:391; Clogg and Eliason 1987). For transitions, I replace the union type terms in equation (1) ($\lambda_l^U, \lambda_{kl}^{AU}, \lambda_{il}^{MU}, \lambda_{jl}^{FU}, \lambda_{ikl}^{MAU}, \lambda_{jkl}^{FAU}$) with terms for the transition type ($\lambda_t^T, \lambda_{kt}^{AT}, \lambda_{it}^{MT}, \lambda_{jt}^{FT}, \lambda_{ikt}^{MAT}, \lambda_{jkt}^{FAT}$) where t (t = 1,…,5) indexes transition type.

I add homogamy, crossings, and hypergamy terms to the baseline model to estimate differences in assortative mating by union type. A homogamy model is:

$$\log(\mu_{ijkl} / t_{ijkl}) = \text{Baseline model} + \gamma_{ml}^{HU} \qquad (2)$$

where H = 1 if male partner's education category equals female partner's education category and 0 otherwise and γ_{ml}^{HU} estimates the difference in the log odds of homogamy in union type l relative to the omitted union type (cohabitation), respectively. The model for transitions replaces the interactions between homogamy and union type with interactions by cohabitation or marriage transition $\gamma_{mt'}^{HT'}$ where T' is a restricted version of T in which the odds of homogamy among couples exiting cohabitation via marriage and entering marriage via cohabitation are constrained to be equal (t = 1,…,4).[8] Crossings models replace the homogamy/union type interaction (γ_{ml}^{HU}) with interaction terms for whether a union crosses an educational barrier (C = <12/12, 12/13-15, 13-15/16 or more years of schooling) by union type. Hypergamy models replace this interaction with a term for the interaction between union type and whether or not men have more education than their female partners (P = 0,1).[9]

A complication of using NLSY79 data to examine assortative mating in this way is that respondents may appear in the data for as many years in which they are either in a cohabiting or marital union. Because the data contain multiple observations per respondent, test statistics assuming independent observations are invalid. To correct for the respondent-level clustering, however, it is necessary to use individual-level rather than grouped data. Thus, I use binomial and multinomial logit models that are equivalent to the equations above but in which the units are couple-years rather than cell frequencies. I use the robust cluster option in STATA to correct for the clustering of errors around respondents.[10]

RESULTS

Log-linear Models

Table 2 provides the model specifications and fit statistics for selected models of educational homogamy. Appendix Table 1 shows the fit statistics for models that further explore these relationships with crossings and hypergamy parameters. I provide the likelihood ratio and BIC statistics to assess model fit. More negative BIC statistics indicate a better fitting model (Raftery 1995). I primarily discuss the results in conjunction my presentation of the homogamy coefficients below but provide a brief overview of the models here.

Model 1 in Panels A through D show the fit statistics for the baseline model for each of the four samples used in this analysis. Because the baseline model is relatively unparsimonious, it fits the data fairly well from the outset, and in the case of new unions, produces expected frequencies that are statistically indistinguishable from the observed data. Model 2 demonstrates the effects of removing the age marginals from the model. As is evident from the G^2 and BIC statistics, this model fits far worse than other models in the table. Model 3 controls for differences in the age distributions of married and cohabiting couples and contains an interaction between homogamy and union or transition type. Finally, Model 4 allows the difference between the educational resemblance of cohabitors and married couples to vary by female partner's age.

Model 5 (Appendix Table 1) relaxes the assumption that differences in the educational resemblance of couples by union type can be characterized solely by differences in the log odds of homogamy by allowing for sex asymmetries in assortative mating (hypergamy). In no case does this significantly increase the fit of the model to the data, which suggests that differences in assortative mating by union type are symmetric with respect to sex. These results are consistent with Qian's (1998) findings from the 1990 CPS, but contradict most other previous research (Blackwell and Lichter 2000, 2004; Casper and Bianchi 2002; Schoen and Weinick 1993). Qian (1998) shows differences in the tendency for men to "partner down" with respect to education in marriage and in cohabitation declined from 1970 to 1990. My sample may contain a somewhat larger proportion of recent unions than have past studies using data from the late 1980s through the mid-1990s (Blackwell and Licther 2000, 2004; Schoen and Weinick 1993), which may explain these null findings. Alternatively, differences in modeling strategy and sample selection may explain these differences.

Models 6 through 8 include terms for differences in the log odds of crossing educational barriers by union type (Appendix Table 1). This parameterization improves the fit of the model in the prevailing union samples and marginally improves the fit in the new unions sample, which indicates that the homogamy parameters alone do not provide an adequate summary of variation in assortative mating by union type. Thus, in what follows I summarize trends using the homogamy parameters but turn to the crossings parameters to gain insight into which portions of the education distribution generate the differences in homogamy I observe.

Prevailing Marriages and Cohabiting Unions

I first turn to a comparison of the log odds of homogamy in prevailing marriages and cohabiting unions and examine how the difference in homogamy by relationship type is affected by differences in the age distributions of marital and cohabiting unions. The results shown in the top panel of Figure 1 are estimated from NLSY79 data and those in bottom panel are estimated for the NLSY79 cohort using 1979-2000 March CPS data. I present these findings in the log scale as it is generally more convenient to discuss "differences in log odds" than "ratios of odds" although I refer to both in the text. Figure 1 shows the log odds of homogamy in prevailing marital and cohabiting unions estimated from Models 2 through 4.[11] Panel A of Figure 1 shows that the log odds of educational homogamy in both cohabiting and marital unions in the NLSY79 are high. The odds of homogamy among married couples implied by Model 2 are 2.78 times the odds of heterogamy ($\exp\{1.022\}$) and the odds of homogamy among cohabiting couples are 2.4 times the odds of heterogamy ($\exp\{0.884\}$). Although the odds of homogamy in both union types are high, they are somewhat higher in marriages than among cohabitors. The odds of homogamy in prevailing marriages are 15% higher than those in prevailing cohabiting unions ($\exp\{0.14\}$). The log odds of homogamy are generally higher in the CPS than in the NLSY79, but the gross differences in the log odds of homogamy among prevailing marriages compared to those among prevailing cohabiting unions are similar (0.15 versus 0.14).

The data from both sources show relatively small changes in the ratio of the odds of homogamy when differences in the age distribution of cohabitors and married couples are taken into account (Model 3). Adding female partner's age to the models reduces the difference in the log odds of homogamy by 9% (0.14-0.13/0.14) in the NLSY79 but increases this difference by 10% (0.15-0.17/0.15) in the March CPS. This suggests that age differences in the distributions of cohabiting couples do not account differences in the log odds of homogamy between the two union types.

The results from Model 4 (the rightmost portions of Panels A and B) show the age patterns of resemblance in the two union types and why the differences remain after adjusting for differences in the age composition of the two groups. Although Table 2 (Panels A and B, Model 4) indicates that differences in the log odds of homogamy do not vary by female partner's age, I present the unrestricted patterns here for illustrative purposes. Figure 1 shows that age patterns of homogamy in both the NLSY79 and the March CPS increase from age 18-21 through age 26-29. After this point, the log odds of homogamy drop slightly in the March CPS data and more substantially in the NLSY79. In the NLSY79, trends in the odds of homogamy among both married couples and cohabitors follows a strong inverted "U" pattern whereas the March CPS shows hints of this pattern.[12] However, the log odds of homogamy among cohabitors are lower than those among married couples at each age and thus the difference in the log odds of homogamy is not eliminated by controlling for age. The general correspondence between the results from the March CPS and the NLSY79 is reassuring given the limitations of the NLSY79.[13] These results are consistent with past findings of greater levels of homogamy among marriages than among cohabiting unions using samples of prevailing unions (Blackwell and Lichter 2000).

Differences in educational homogamy between prevailing cohabiting and marital unions may arise from several sources. Homogamy in marriage in each subsequent age category is made up of the accumulation of all of the marriages that have survived through previous periods, minus all the divorces that occur, plus any new marriages, and plus or minus the net effects of any educational changes that occur within marriage (Schwartz and Mare 2003). This is also true of cohabiting unions except the effects of the accumulation of past cohabitors and of educational changes will be far lower for cohabitors because of the short duration of the majority of cohabiting spells. In the present sample, the median number of cohabitation interview years respondents contribute to the data is 1 and the mean is 2.1. By contrast, the median and the mean number of married interview years per marriage respondents contribute to the data are 5 and 5.9, respectively. Thus, although I examine the "stock" of both cohabiting and marital unions here, the stock of cohabitors is likely to closely resemble its "flow" whereas the stock of marriages contains a high proportion of "build up" from past flows, especially at older ages.

New Cohabiting and Marital Unions

Next, I assess the extent to which differences in the log odds of homogamy between prevailing cohabiting and marital unions are due to differences in the way in which couples sort into unions. Figure 2 shows the results from Models 2 through 4 using a sample of newly formed marriages and cohabiting unions from the NLSY79. Figure 2 shows that among newly formed unions the differences in the log odds of homogamy shown in Figure 1 are eliminated. Both the total and net differences in the log odds of homogamy as well as differences at each age interval are statistically insignificant when newly formed unions are compared. These results indicate that differences in the log odds of homogamy in prevailing marriages and cohabiting unions are not attributable to differences in homogamy at the time of union formation in this cohort.[14]

Again as in Figure 1, the inverted "U" shape in the log odds of homogamy for both union types is evident. These age patterns are consistent with several hypotheses about the role of marriage markets in choosing a mate. First, the decline in the odds of homogamy among couples beginning their unions at older ages is consistent with the hypothesis that the shrinking availability of potential partners forces still-single men and women to redefine what constitutes an acceptable match (Lewis and Oppenheimer 2000; Lichter 1990). Alternatively, as young people leave educational institutions and move into the labor market they may be more likely to encounter potential partners who do not share their educational attainment (Mare 1991). By contrast, the odds of homogamy among those who marry young may be lower than among those who marry at older ages if young people are more likely match on expected rather than completed education. The similarity in these trends by union type suggests that whether or not individuals decide to cohabit or marry they may be subject to similar marriage market pressures as they age.

Exiting Cohabitors and Entering Newlyweds by Transition Type

Although I find no difference in the odds of homogamy by union type among newly formed cohabiting and marital unions, these findings do not preclude differences in selection out of cohabitation and into marriage. Differences in selection out of cohabitation via separation or marriage and into marriage via cohabitation or "cold" marriage may offset each other if cohabitors who transition to marriage are more likely to be homogamous than cohabitors who separate but if those heterogamous cohabitors who separate are replace by equally heterogamous couples who marry without first cohabiting. Figure 3 shows the odds of homogamy among cohabitors exiting their relationships via separation, those who transition to marriage, and those who are in their first year of a "cold" marriage. Figure 3 reveals that there do not appear to be offsetting effects of transitions. Instead, I find no differences in the odds of homogamy between those who exit cohabitation via marriage, those who exit via separation, or those who enter marriage without first cohabiting either in the overall estimates, the estimates controlling for each, and at each age interval.

However, these results appear to be sensitive to the sex of the respondent. Male respondents who transition to marriage from cohabitation are more likely to be homogamous than those who separate or those who marry "cold" and female respondents who transition to marriage from cohabitation are less likely to be homogamous than those who separate (Appendix Figure 1). Variation by respondent's sex could be the result of differences in the composition of the male and female-respondent samples, measurement error, or sample attrition.[15] These results must be regarded as tentative because of the differences I find by respondent's sex, but they nevertheless suggest that the differences in the log odds of homogamy by union type observed in Figure 1 are neither due to the selection of the most homogamous couples from cohabitation to marriage or to greater levels of homogamy among "cold" marriages than among couples who marry via cohabitation.

To investigate this hypothesis further, I estimate differences in the log odds of homogamy by union type among couples in their last year of marriage (who exit the sample either because of divorce or separation) and among couples in the last year of cohabitation (who exit either because of marriage or separation) compared with married and cohabiting couples that are not in their last year of a union. These results are shown in Figure 4. This figure shows that couples whose marriages dissolve tend to be somewhat less homogamous than couples in the remaining stock of marriages. The odds of homogamy in the remaining stock of prevailing marriages are about 11% (exp{1.037-0.934}) higher than the odds of homogamy among couples about to dissolve their marriages. By contrast, the log odds of homogamy among exiting cohabitors are slightly higher than among cohabitors in the remaining stock of unions, although this difference is statistically insignificant. These results provide further evidence that differences in the log odds of homogamy by union type may largely driven by the accumulation of the most homogamous marriages over time rather than by differences in their homogamy at their entry or transitions between cohabitation and marriage. Future research should pursue a more complete decomposition of the difference in the log odds of homogamy in prevailing unions.

Variation in Educational Intermarriage by Union Type

Describing differences in assortative mating by union type on the basis of homogamy alone may mask significant variation by educational level. To determine whether cohabitors or married couples are more or less likely to cross particular educational boundaries, I turn to the results of the crossings models. The crossings models reveal that patterns of educational assortative mating by union type vary across the education distribution (Appendix Table 1). Among prevailing unions, cohabitors are somewhat more likely married couples to cross the lowest and highest educational barriers but are less likely to cross the middle barrier (Appendix Table 2). The odds of crossing the less than 12/12 years of schooling barrier and the "some college"/college barrier relative to the odds of homogamy are 27% and 11% higher among cohabiting couples than among married couples, respectively, although the latter difference is not statistically significant. By contrast, the odds of crossing the 12 years of schooling/"some college" barrier is 18% higher among married couples than among cohabitors. The combination of these differences results in higher levels of homogamy in marital than in cohabiting unions. These results, however, are not consistent across data sources. Data for the NLSY79 cohort from the March CPS indicate that cohabitors are less likely to cross each of the three educational barriers. Data from the 1995 NSFG are consistent with results from the NLSY79 (Blackwell and Lichter 2004).

Although there appear to be almost no differences in the log odds of homogamy by union type among couples entering new unions (Figure 2), there are significant differences in the odds of crossing educational barriers by union type. The odds of crossing the 12/13-15 years of schooling barrier are significantly higher among newly married couples than among newly cohabiting couples. By contrast, the odds of crossing the lowest educational barrier (less than 12/12 years of schooling) are higher among cohabitors than among married couples but this ratio is not statistically significant. These two differences offset each other and lead to no difference in the log odds of educational homogamy by union type. These results imply that there *are* differences in the way in which cohabitors and married couples sort into marriage but that they do not translate into higher or lower odds of educational homogamy among newly formed cohabiting and marital unions.

SUMMARY AND CONCLUSIONS

I find that differences in the educational resemblance of married and cohabiting couples vary depending on the sample used, the point at which assortative mating patterns are measured, and the measures employed. In analyses conducted internal to the NLSY79, I show how differences in sampling and modeling strategies produce alternate interpretations of assortative mating in cohabiting and marital unions.   Although a host of other data sources would need to be analyzed to pin down the reasons for the disparate findings of past research, my findings provide insight into the possible effects of differences in the ways in which these differences are conceptualized and measured.

First, like past research using data on prevailing cohabiting and marital unions, I find that cohabitating couples are somewhat less likely to be educationally homogamous than married couples (Blackwell and Lichter 2000; Gullickson 2004; but see Jepsen and Jepsen 2002). These differences, however, are not apparent at the time at which couples form their unions.   Rather, couples entering new marriages and cohabitating unions have very similar odds of educational homogamy across the age interval I examine. This suggests that the accumulation of the most homogamous marriages over time, rather than differences in union formation patterns by relationship type, account for differences in the odds of homogamy by union type among prevailing unions. These findings also suggest that past studies that have not restricted their analyses to couples in newly formed unions may somewhat overstate differences in the educational resemblance of married and cohabiting couples. Even among relatively young couples, differences in the odds of homogamy by union type are around 10 to 15% larger among prevailing unions than among newly formed unions.

Second, although I find no difference in the odds of homogamy by union type among newly formed unions, assortative mating into cohabiting and marital unions do differ by the extent to which they cross educational barriers. The odds of union formation between those with 12 years of schooling and those with "some college" are higher for married couples than among cohabitors, whereas the odds of union formation across the less than 12/12 years of schooling barrier are lower for married couples than for cohabitors. These results are consistent with estimates from the 1995 NSFG (Blackwell and Lichter 2004) but not with the March CPS. The differences in assortative mating by educational barrier in the NLSY79, however, offset each other and lead to no overall difference in the odds of homogamy by union type.

Third, I find little support for the hypothesis that marriages are more homogamous than cohabiting unions because couples are "doubly selected" into marriage via cohabitation. Using a pooled sample of male- and female-respondents, I find no difference in the odds of homogamy between cohabitors who exit their unions via cohabitation and those that exit via marriage. In this respect, my results are consistent with past longitudinal studies of the effects of the joint educational attainment of cohabitors on the probability of transitioning to marriage or separating (Goldstein and Harknett 2004; Oppenheimer 2003:133; Sassler and McNally 2003). Furthermore, the pooled sample reveals no differences in the odds of homogamy among couples who enter marriage via cohabitation and those who were not cohabiting in the previous interview year. However, these results vary by the sex of the respondent for reasons that are not yet fully understood and therefore should be regarded as tentative. The sensitivity of this portion of the analysis to the sex of the respondent points to the usefulness of investigating compositional differences between the two samples as well as the impact of measurement error and survey non-response in future research (Sassler and McNally 2003). Despite these caveats, my results suggest that differences in the odds of homogamy in prevailing marriages and cohabiting unions are not the result of the selection of the most homogamous cohabitating couples into marriage. The accumulation of the most homogamous marriages over time and selective marital dissolution are more likely explanations. Future research should perform a more complete decomposition of these differences.

Overall, differences in educational assortative mating in cohabiting and marital unions are small. The picture that emerges from these results is one of a similar process of educational assortative mating for both cohabiting and marital unions governed more by age at unions formation than by differences by union type. I find that for both cohabitors and married couples, the odds of educational homogamy rise among couples making transitions at young ages, peak among couples making transitions in their mid- to late-20s, and fall among those making transitions at older ages. A fruitful avenue for future research would be to examine the role of the timing of school enrollment and union formation on age patterns in the odds of homogamy.

ENDNOTES

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[1] With the release of the 1992-94 wave, it became possible to use the NSFH to study transitions from cohabitation to marriage. However, these data were not available at the time at which Schoen and Weinick (1993) were conducting their analysis.

[2] Beginning in 1990, married respondents were asked whether they lived with their partner before marriage and if they had, the date at which they began living together and whether they lived with their partner continuously. I use these data where possible to impute partner information for years in which respondents missed interviews or in years in which the interview was not administered. I also use the detailed marital status change information to impute respondents' marital status and educational characteristics in missing interview years. I follow the procedures outlined in Appendix A of Schwartz and Mare (2003) in handling missing data for marriages. Similar procedures were used for cohabitation. Sensitivity tests show that my results are robust to these procedures.

[3] To assess the possible effects of the switch to the biennial interview schedule, I replicate my analysis using data from 1990 to 1994 but in which I delete cohabitation data from 1991 and 1993. This simulates what my results would have been had respondents been interviewed biennially in this period. I then compare these results to those using all of the available data over this period. Dropping these years does not affect the difference by union type in either the prevailing union or new unions samples (the results are within 3% of each other). Dropping these years has larger effects on differences in the odds of making transitions. This switch occurred relatively late in the survey period, at which point NLSY79 respondents were 29 and older. Therefore, my results for transitions among older couples should be interpreted with caution. There are two possible ways of addressing this problem in future research. One

solution is to use the simulated 1990-1994 results to adjust my findings for the 1994-2000 period. This method assumes that the bias in the 1990-1994 period is constant over the 1994-2000 period. Another would be to use the detailed information on the start and end dates from the first and second waves of the NSFH to calculate a similar adjustment factor.

[4] The hypothesis that the resemblance among couples in marriages preceded by very short-term cohabitations are similar to those that are not preceded by cohabitation is testable using NLSY79 data from 1990-2000. An analysis of the relationship between cohabitation duration and marital homogamy is an important avenue of future research.

[5] Restricting the sample to cases in which both partners are between 18 and 37 years of age effectively doubles my sample size and allows me to pool the female-respondent and male-respondent samples. This is desirable because of the relative rarity of some of the union transitions I examine at the youngest and oldest ages. Thus, this sample is representative of couples in which one partner was between 14 and 22 in 1979 and in which both partners are between the ages of 18 and 37 between 1979 and 2000.

A less restrictive way of defining the sample is to select all female-respondents in unions from age 18 to 37, and all male-respondents who have female partners age 18 to 37 without restricting male partner's ages. Such a sample would allow me to examine assortative mating patterns by female partner's age, but also allows for more variation in male partner's age. However, this sample is somewhat lopsided because, although the ages of the male partners of female-respondents vary freely, the ages of the male-respondents are restricted to those born in the NLSY79 cohort. Because of the awkwardness of this alternative sample, I chose to restrict my analysis sample to couples in which both partners are between the ages of 18 and 37. Restricting the age of partners in this way tends to affect the oldest and youngest age categories slightly. I discuss these effects in conjunction with the results. An additional complication of pooling the male- and female-respondent samples is that some of my results are sensitive to the sex of the respodent. These differences are also discussed in the context of the results.

[6] Because of the sample rotation scheme in the CPS, up to half of sample members in any given year may have been interviewed in the previous year. To avoid duplicate observations, I examine the outgoing rotation groups (month-in-sample 4 to 8) only.

A drawback of the POSSLQ method is that cohabiting partners who live in households with other adults are not counted but households which are made up of two roommates of the opposite-sex are. A test of the difference in the odds of homogamy between self-identified and indirectly identified cohabiting couples reveals that these limitations do not appear to substantially affect the odds of homogamy among members of the sample for whom both measures are available. Using data from 1995 through 2000 (at which point sample members were 30 to 37), I find that the odds of educational homogamy are 3% higher among self-identified cohabitors than POSSLQs but that this difference is statistically insignificant (results available upon request).

[7] Because cohabitation exits occur in the last year of cohabiting unions and new marriages occur the first year the marriage is reported, there is no overlap between observations exiting cohabitation and those entering marriage. However, couples can both enter and exit cohabitation in the same year. Because the unit of observation is a transition, these cases contribute two observations to the data.

[8] The odds of homogamy among couples exiting cohabitation via marriage and those among couples entering marriage via cohabitation may not be equal if one or both partners increase his or her education between the last year of the cohabitation and the first year of the

marriage. A test of this restriction indicates that a model with the restricted parameterization fits the data as well as a model that does not impose this restriction.

[9] A model that includes both hypergamy and crossings is:

$$\log(\mu_{ijkl} / t_{ijkl}) = \text{Baseline model} + \gamma_{nl}^{PU} + \gamma_{ijl}^{HWU}$$

where $P = 1$ if male partner's education category is greater than female partner's education category and 0 otherwise, and γ_{nl}^{PU} estimates the difference in the log odds of homogamy relative to the baseline union type (cohabiting unions) and

$$\gamma_{ijl}^{HWU} = \begin{cases} \sum_{q=j}^{i-1} \gamma_{ql} & for\ i > j, \\ \sum_{q=i}^{j-1} \gamma_{ql} & for\ i < j, \\ 0 & for\ i = j. \end{cases}$$

where γ_{ql} represents the change in the difficulty of crossing educational barrier q in union type l relative to the baseline union type (cohabitation).

[10] The models used here are convenient because they are easily replicated using binary logit and multinomial logit models. For example, equation (2) can be estimated using the following logit model with union type as the dependent variable:

$$\text{logit}[P(U = 1 \mid M = \text{i}, F = j, A = k, H = m)] = \alpha + \beta_i^M + \beta_j^F + \beta_k^A + \beta_{ik}^{MA} + \beta_{jk}^{FA} + \delta_m^H .$$

In this case, $\delta_m^H = \gamma_{ml}^{HU}$ from equation (2). When the data are weighted this correspondence is not exact. Likewise, a multinomial logit model for transitions equivalent to equation (2) is:

$$\log \frac{\pi_t(x)}{\pi_T(x)} = \alpha_t + \beta_{ti}^M + \beta_{tj}^F + \beta_{tk}^A + \beta_{tik}^{MA} + \beta_{tjk}^{FA} + \delta_{tm}^H$$

where $\pi_t(x) = [P(T = t \mid M = \text{i}, F = j, A = k, H = m)]$ and where $\delta^H_{2m} = \delta^H_{4m}$, that is, the odds of homogamy are constrained to be equal across couples exiting cohabitation into marriage and entering marriage via cohabitation. See Agresti (2002:330) for details.

[11] The models used here do not produce interpretable coefficients for the odds of homogamy for the omitted union type (cohabitation) because of the inclusion of the interaction terms between male and female partner's education (MF), which control for a general pattern of assortative mating across relationship types. Rather than choosing an arbitrary point of comparison, I estimate the odds of homogamy for cohabitors using modified versions of Models 2 and 4 in which I replace the MF terms with homogamy (H) terms. The difference by union type parameters are estimated from Models 2 to 4 and are added to the estimates for cohabitation from the restricted models.

[12] The inverted "U" shape of patterns of educational homogamy by female partner's age in the CPS is more apparent using a sample in which the male partner's age is unrestricted. Restricting the sample to couples in which both partners are 18 to 37 years old results in slightly higher odds of homogamy and slightly larger difference by union type among 34 to 37 year old female partners in the CPS. The age patterns of assortative mating and differences by union types in prevailing unions are largely invariant to the restriction on male partner's age in the NLSY79. Differences in the log odds of homogamy are smaller for female-respondents than for male-respondents in both the NLSY79 and the CPS but these differences are not statistically significant.

[13] Comparisons of educational homogamy among "unmarried partners" and marriages for the NLSY79 cohort using the 1990 and 2000 Census also exhibit similar trends although

differences in the log odds of homogamy by union type are greater in the Census than in the March CPS (results available upon request).

[14] The June CPS contains information on marriage start dates through 1995, but it is not possible to replicate my analysis of newly formed cohabiting and marital unions using the CPS because it lacks information on the start dates of cohabiting unions. A rough estimate is possible through a comparison of newly contracted first marriages from the June CPS and never-married POSSLQs (identified here from the March CPS), assuming that most cohabitors split up or marry relatively quickly after union formation. When these samples are compared for the NLSY79 cohort, the differences shown in Panel B of Figure 2 are small and statistically insignificant.

The differences in the log odds of homogamy and age patterns of assortative mating by union type shown in Figure 2 are very similar to those using a sample in which male partner's age has not been restricted. Furthermore, although age patterns of assortative mating by union type vary somewhat by the sex of the respondent, the general result—that differences in the log odds of homogamy by union type are not apparent when couples form their unions—is robust to the sex of respondent.

[15]There are several possible avenues for future research to address this issue. First, although I control for the age for the female partner in these models, I do not control for the age of the male partner or assortative mating on age. It is possible that the male and female-respondent samples may have different distributions of male partner's ages and/or different patterns of assortative mating on age that account for these respondent-sex differences. Partner's age and assortative mating on age should be included in future versions of the models.

Second, these differences may be due to selective attrition from the sample or reporting differences by respondent's sex. Sassler and McNally (2003) find that sample attrition and missing partner data produce biased estimates of the effects of partners' characteristics on transitions from cohabitation using the first and second waves of the NSFH, but that the effects of educational homogamy on the odds of making union transitions are insignificant both using a data set corrected for sample attrition and missing partner data and an uncorrected sample. Although insignificant in both cases, the changes in the homogamy coefficients were not trivial. These findings point to the need to carefully assess the effects of sample attrition and missing partner data on the results presented here.

Finally, these differences may be the result of systematic measurement error by respondent's sex. The second wave of the NSFH administered the same questionnaire to both respondents and their spouses and partners. It would be useful to examine whether differences by respondent's sex seen in the NLSY79 are also found in the NSFH or whether these differences are due to measurement error. These are important avenues for future research.

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Table 1.Transitions Out of Cohabitation and Into Marriage by Transition Type and Female Partner's Age.

Female	Cohabitation Exits			Marriage Entry		
Partner's Age	via Separation	via Marriage	Total	via Cohabitation	Cold Marraiges	Total
Total	50.0	50.0	100.0 (3,426)	21.8	78.2	100.0 (7,318)
18-21	53.7	46.3	100.0 (651)	8.6	91.4	100.0 (1,976)
22-25	49.6	50.4	100.0 (1,071)	19.3	80.7	100.0 (2,484)
26-29	46.6	53.4	100.0 (855)	29.3	70.7	100.0 (1,575)
30-33	46.9	53.1	100.0 (577)	37.3	62.7	100.0 (865)
34-37	61.3	38.7	100.0 (272)	36.1	63.9	100.0 (418)

SOURCE: National Longitudinal Survey of Youth (NLSY79).
NOTES: Sample sizes are shown in parentheses. Data are weighted using 1979 weights to correct for oversampling and for survey non-response.

Table 2. Goodness of Fit Statistics for Selected Models of Educational Homogamy in Cohabiting and Marital Unions.

								Adjusted for Person-Level Clustering		
					Unadjusted					
Data Source, Sample, and Model	df	G^2	BIC	H_0	G^2	df	p-value	Wald test	df	p-value
Panel A. NLSY79 Prevailing Unions										
1. MUA, FUA, MFA	45	179.4	-326	--	--	--	--	--	--	--
2. MU, FU, MF, HU	136	15,678.7	14,158	HU = 0	32.0	1	0.000	4.19	1	0.041
3. MUA, FUA, MFA, HU	44	153.4	-339	HU = 0	26.1	1	0.000	3.75	1	0.053
4. MUA, FUA, MFA, HUA	40	152.1	-295	HUA = 0	1.3	4	0.869	0.44	4	0.979
(N = 160 cells; 71,666 couple years)										
Panel B. CPS Prevailing Unions, NLSY79 Cohort										
1. MUA, FUA, MFA	45	147.5	-357	--	--	--	--	--	--	--
2. MU, FU, MF, HU	136	15,480.1	13,955	HU = 0	34.7	1	0.000	--	--	--
3. MUA, FUA, MFA, HU	44	108.9	-384	HU = 0	38.7	1	0.000	--	--	--
4. MUA, FUA, MFA, HUA	40	105.2	-343	HUA = 0	3.7	4	0.447	--	--	--
(N = 160 cells; 65,379 couple years)										
Panel C. NLSY79 New Unions										
1. MUA, FUA, MFA	45	39.1 *	-380	--	--	--	--	--	--	--
2. MU, FU, MF, HU	136	4,676.3	3,408	HU = 0	0.0	1	0.936	0.08	1	0.775
3. MUA, FUA, MFA, HU	44	39.1 *	-371	HU = 0	0.1	1	0.817	0.12	1	0.732
4. MUA, FUA, MFA, HUA	40	37.6 *	-335	HUA = 0	1.4	4	0.837	0.99	4	0.732
(N = 160 cells; 11,187 couple years)										
Panel D. NLSY79 Transitions										
1. MTA, FTA, MFA	180	234.3	-1,492	--	--	--	--	--	--	--
2. MT, FT, MF, HT'	353	6,549.2	3,164	HT' = 0	1.6	3	0.661	0.24	3	0.970
3. MTA, FTA, MFA, HT'	177	233.6	-1,464	HT' = 0	0.7	3	0.874	0.87	3	0.870
4. MTA, FTA, MFA, HT'A	165	228.3	-1,354	HT'A = 0	5.4	12	0.945	8.33	12	0.951
(N = 400 cells; 14,613 couple years)										

SOURCES: National Longitudinal Survey of Youth (NLSY79); March Current Population Survey (CPS) 1979-2000.

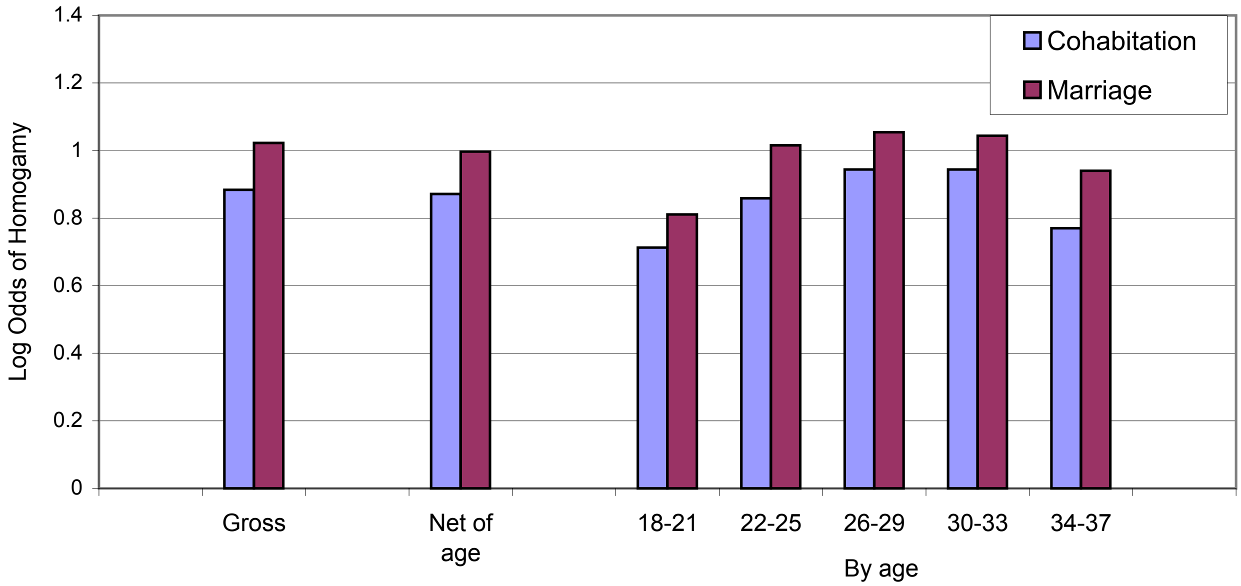
NOTES: Cohabiting couples are self-identified in the NLSY79 and are identified using POSSLQ methods in the CPS following Casper and Cohen (2000). Models adjusted for person-level clustering are estimated using binomial or multinomial logistic regression equivalents to log-linear models where union type or transition type is the dependent variable. The NLSY79 results are weighted using 1979 probability weights. The CPS results are weighted using household weights.

Model terms are as follows (degrees of freedom in parentheses): M = Male partner's education (3); F = Female partner's education (3); A = Female partner's age (4); U = Union type (1); H = Homogamy (1); T = Transition type: New cohabitation (omitted); Cohabitation exit via dissolution; Cohabitation exit via marriage; Marriage entry via cohabitation; Marriage entry without cohabitation (4); T' = Transition type: New cohabitation (omitted); Cohabitation exit via dissolution; Cohabitation to marriage; Marriage entry without cohabitation (3).

*Model fits the data at p > .10.

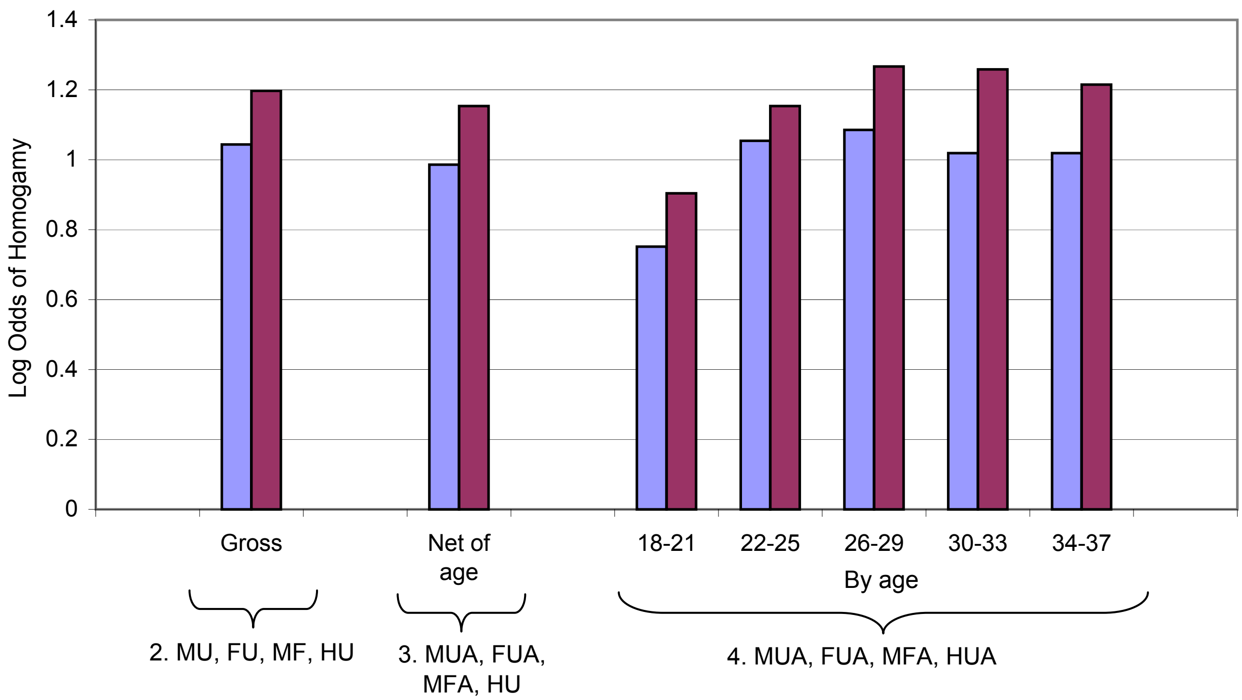
Figure 1. Log Odds of Educational Homogamy in Prevailing Unions by Union Type, Selected Models.

Panel A. NLSY79



Mar - Cohab:	0.14**	0.13*	0.10	0.16	0.11	0.10	0.17

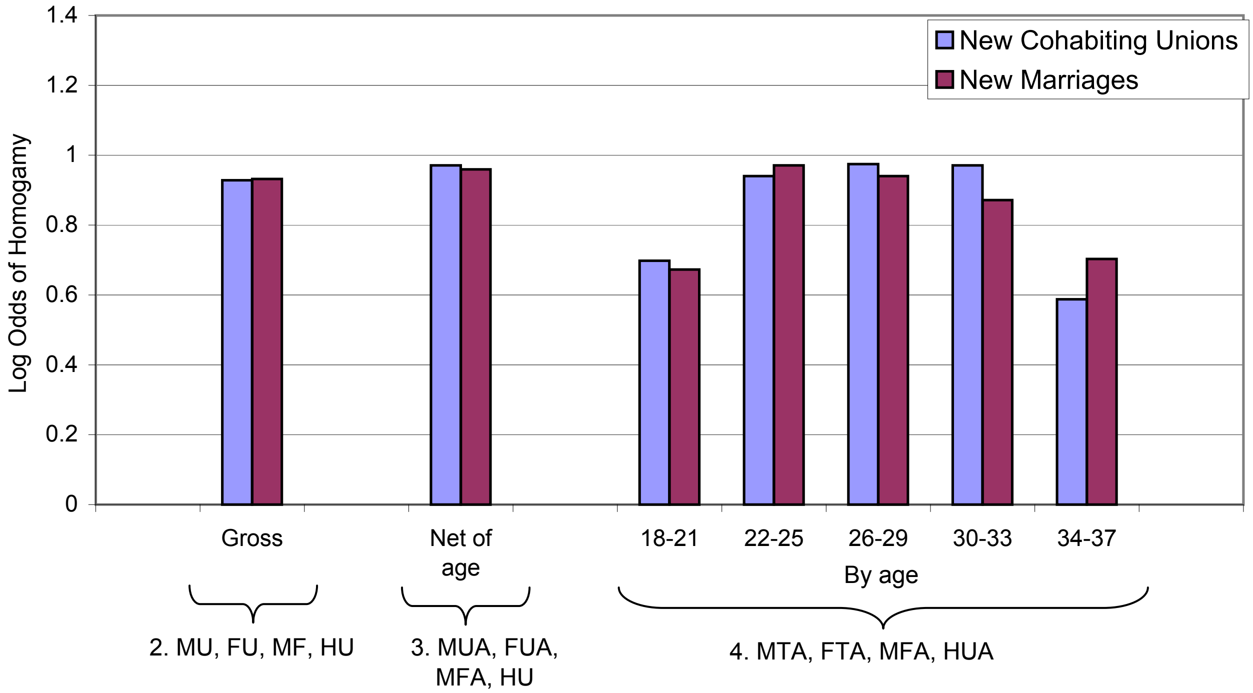
Panel B. March CPS, NLSY79 Cohort



Mar - Cohab:	0.15***	0.17***	0.15**	0.10**	0.18***	0.24***	0.20**

NOTES: ***p < .001; **p < .05; *p < .01. Significance tests were performed using Wald tests and adjusted for respondent-level clustering in the NLSY79.

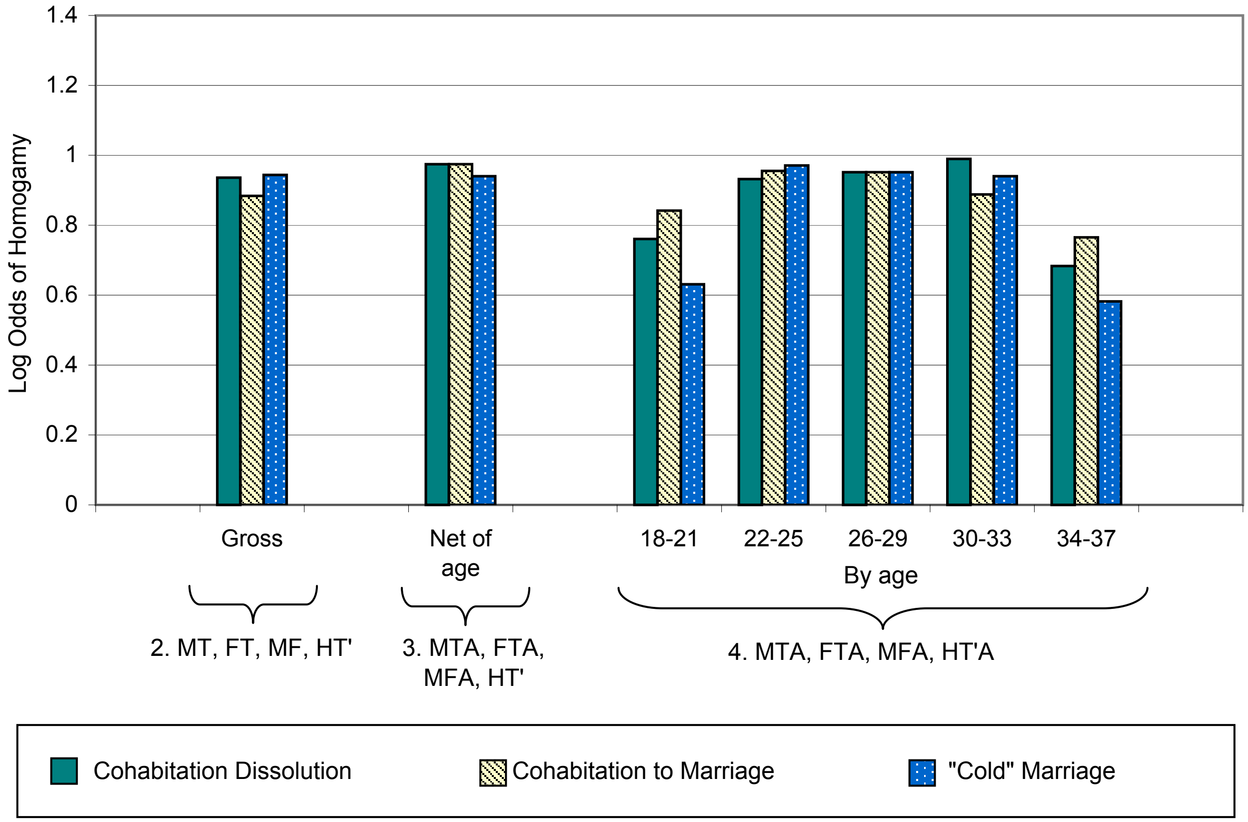
Figure 2. Log Odds of Educational Homogamy among New Cohabiting and Marital Unions, Selected Models, NLSY79.



Mar - Cohab:	0.00	-0.01	-0.03	0.03	-0.03	0.10	0.12

NOTES: ***p < .001; **p < .05; *p < .01. Significance tests were performed using Wald tests and adjusted for respondent-level clustering in the NLSY79.

Figure 3. Log Odds of Educational Homogamy among Exiting Cohabitors and Newlyweds by Transition Type, Selected Models, NLSY79.

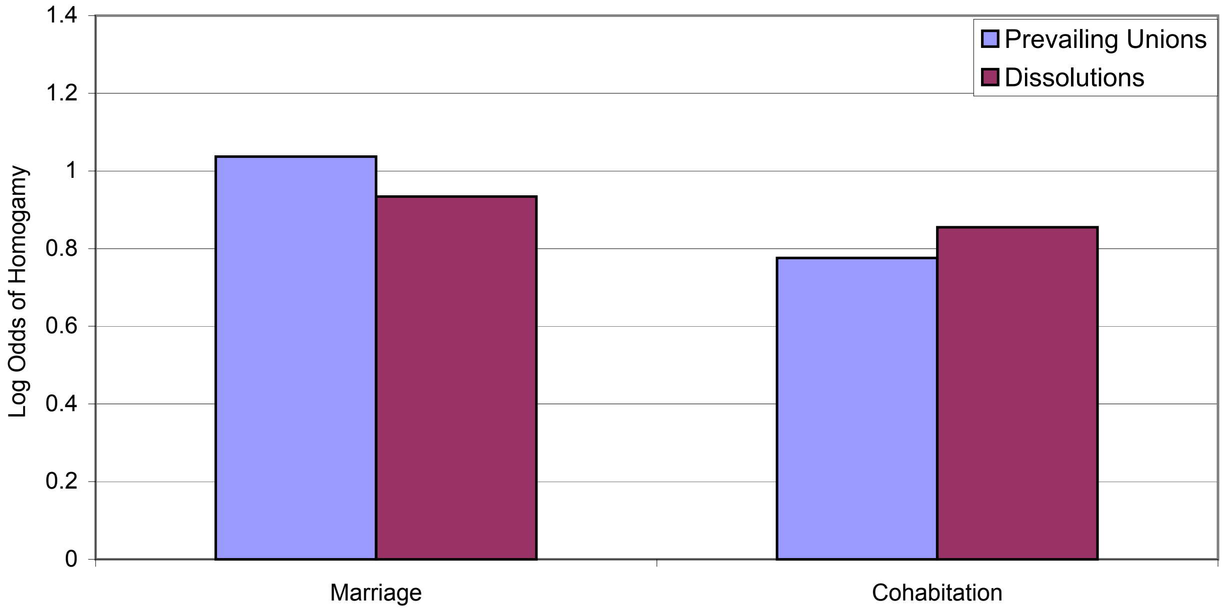


Differences:

CtoM-Diss	-0.05	0.00	0.08	0.02	0.00	-0.10	0.08
Cold-CtoM	0.06	-0.03	-0.21	0.02	0.00	0.05	-0.18
Cold-Diss	0.01	-0.03	-0.13	0.04	0.00	-0.05	-0.10

NOTES: CtoM = Cohabitation to Marriage; Diss = Cohabitation Dissolution; Cold = Marriage not preceded by Cohabitation.
***p < .001; **p < .05; *p < .01. Significance tests were performed using Wald tests and adjusted for respondent-level clustering in the NLSY79.

Figure 4. Log Odds of Educational Homogamy for Prevailing Unions and Union Dissolutions by Union Type, NLSY79.



Differences:		
Prevailing - Dissolutions	0.10**	-0.08

NOTES: ***p < .001; **p < .05; *p < .01. Significance tests were performed using Wald tests and adjusted for respondent-level clustering in the NLSY79.

Appendix Table 1. Goodness of Fit Statistics for Additional Models of Educational Assortative Mating in Cohabiting and Marital Unions.

Data Source, Sample, and Model	df	G^2	BIC	H_0	Unadjusted G^2	Unadjusted df	Unadjusted p-value	Adjusted for Person-Level Clustering Wald test	Adjusted df	Adjusted p-value
Panel A. NLSY79 Prevailing Unions										
5. MUA, FUA, MFA, HU, PU	43	152.0	-329	PU = 0	1.37	1	0.241	0.31	1	0.575
6. MUA, FUA, MFA, HU, CU	41	62.6	-396	HU = 0	1.51	1	0.219	0.38	1	0.540
7. MUA, FUA, MFA, CU	42	64.1	-405	CU = 0	115.3	3	0.000	20.05	3	0.000
8. MUA, FUA, MFA, CUA	30	49.1	-286	CUA=0	15.00	12	0.242	6.64	12	0.881
(N = 160 cells; 71,666 couple years)										
Panel B. CPS Prevailing Unions, NLSY79 Cohort										
5. MUA, FUA, MFA, HU, PU	43	106.5	-376	PU = 0	2.34	1	0.126	--	--	--
6. MUA, FUA, MFA, HU, CU	41	67.4	-392	HU = 0	0.12	1	0.728	--	--	--
7. MUA, FUA, MFA, CU	42	67.5	-403	CU = 0	80.0	3	0.000	--	--	--
8. MUA, FUA, MFA, CUA	30	47.0	-289	CUA=0	59.5	11	0.060	--	--	--
(N = 160 cells; 65,379 couple years)										
Panel C. NLSY79 New Unions										
5. MUA, FUA, MFA, HU, PU	43	39.1 *	-362	PU = 0	0.03	1	0.864	--	--	--
6. MUA, FUA, MFA, HU, CU	41	32.8 *	-349	HU = 0	0.03	1	0.874	0.07	1	0.795
7. MUA, FUA, MFA, CU	42	32.8 *	-359	CU = 0	6.34	3	0.096	6.30	3	0.098
8. MUA, FUA, MFA, CUA	30	29.2 *	-250	CUA=0	3.62	12	0.989	2.73	12	0.997
(N = 160 cells; 11,187 couple years)										
Panel D. NLSY79 Transitions										
5. MTA, FTA, MFA, HT', PT'	174	263.0	-1,406	PT' = 0	0.71	3	0.872	--	--	--
6. MTA, FTA, MFA, HT', CT'	168	251.5	-1,360	HT' = 0	1.36	3	0.715	0.94	3	0.816
7. MTA, FTA, MFA, CT'	171	252.8	-1,387	CT' = 0	11.69	9	0.231	8.09	9	0.525
8. MTA, FTA, MFA, CT'A	135	203.3	-1,091	CT'A=0	49.48	36	0.067	34.02	36	0.563
(N = 400 cells; 14,613 couple years)										

SOURCES: National Longitudinal Survey of Youth (NLSY79); March Current Population Survey (CPS) 1979-2000.
NOTES: Cohabiting couples are self-identified in the NLSY79 and are identified using POSSLQ methods in the CPS following Casper and Cohen (2000). Models adjusted for person-level clustering are estimated using binomial or multinomial logistic regression equivalents to log-linear models where union type or transition type is the dependent variable. The NLSY79 results are weighted using 1979 probability weights. The CPS results are weighted using household weights.

Model terms are as follows (degrees of freedom in parentheses): M = Male partner's education (3); F = Female partner's education (3); A = Female partner's age (4); U = Union type (1); H = Homogamy (1); P = Hypergamy (1); C = Crossings (3); T = Transition type: New cohabitation (omitted); Cohabitation exit via dissolution; Cohabitation exit via marriage; Marriage entry via cohabitation; Marriage entry without cohabitation (4); T' = Transition type: New cohabitation (omitted); Cohabitation exit via dissolution; Cohabitation to marriage; Marriage entry without cohabitation (3).
*Model fits the data at p > .10.

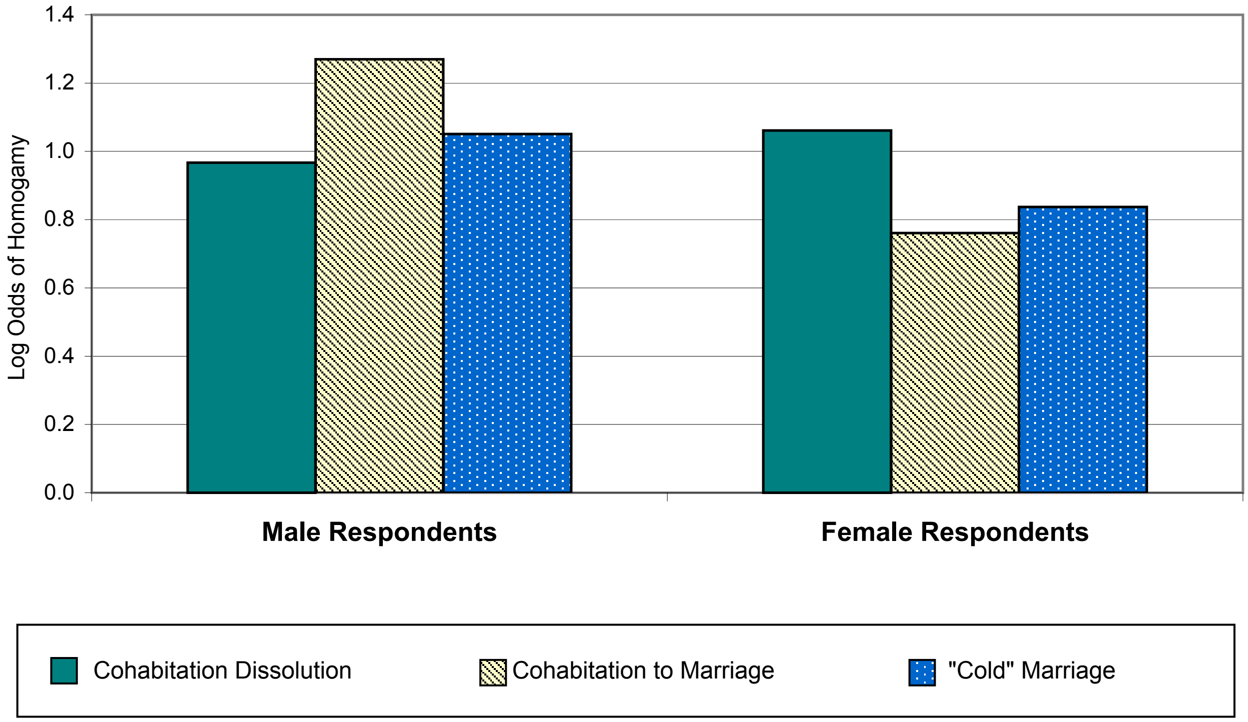
Appendix Table 2. Odds of Crossing an Educational Barrier vs. Odds of Homogamy by Union or Transition Type

	Educational Barrier		
Data Source and Sample	lt 12/12	12/13-15	13-15/ge 16
Panel A. NLSY79 Prevailing Unions			
Marriage	0.41	0.45	0.34
Cohabitation	0.56	0.38	0.38
Ratio (M:C)	0.73 **	1.18 *	0.89
Panel B. CPS Prevailing Unions, NLSY Cohort			
Marriage	0.36	0.42	0.35
Cohabitation	0.51	0.45	0.36
Ratio (M:C)	0.71 ***	0.93 **	0.97
Panel C. NLSY79 New Unions			
Marriage	0.48	0.45	0.39
Cohabitation	0.53	0.41	0.39
Ratio (M:C)	0.91	1.10 **	1.00
Panel D. NLSY79 Transitions			
Cohabitation Dissolution	0.54	0.40	0.32
Cohabitation to Marriage Transition	0.54	0.39	0.40
"Cold" Marriage	0.48	0.43	0.37
Ratio (Cohab Diss:Cohab to Mar)	1.00	1.03	0.80
Ratio (Cohab to Mar:"Cold" Mar)	1.13	0.91	1.08
Ratio (Cohab Diss:"Cold" Mar)	1.13	0.93	0.86

SOURCES: National Longitudinal Survey of Youth (NLSY79); March Current Population Survey (1979-2000).
NOTES: Results are estimated from Model 7 in Panels A thorough D of Appendix Table 1.

Significance tests for the ratio between crossings parameters by union or transition type were performed using Wald tests and adjusted for respondent-level clustering in the NLSY79 and are indicated as: ***p < .001; **p < .05; *p < .01.

Appendix Figure 1. Log Odds of Educational Homogamy among Exiting Cohabitors and Newlyweds by Transition Type and Respondent's Sex (Results Estimated from Model 7), NLSY79.



Differences:		
CtoM-Diss	0.30**	-0.30**
Cold-CtoM	-0.22*	0.08
Cold-Diss	0.08	-0.22

NOTES: CtoM = Cohabitation to Marriage; Diss = Cohabitation Dissolution; Cold = Marriage not preceded by Cohabitation.
***p < .001; **p < .05; *p < .01. Significance tests were performed using Wald tests and adjusted for respondent-level clustering in the NLSY79.