Determinants of Long-Term Unions: Who Survives the "Seven Year Itch"?

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Abstract

Most studies of union formation focus on the probability of marrying, cohabiting, or divorcing in the next year. In this study, we take a long-term perspective by considering probabilities of forming unions by certain ages *and* maintaining them for eight or more years. We use NLSY79 data to estimate choice models for each stage of the union-forming process (single with no prior unions, cohabiting, first marriage, single with prior unions, non-first marriage) and, in turn, simulate women's union-related outcomes from ages 18 to 42. Based on simulated outcomes, we predict that a representative, 18 year-old with no prior unions has a 27% chance of entering a union within four years and maintaining it for at least eight years. This predicted probability changes very little for a 24 year-old with no prior unions or for a 30 year-old with prior unions. For older women, however, cohabitation becomes the modal form of entry into long-term unions. We also find that the likelihood of experiencing a long-term union varies with a woman's characteristics (especially race and family background), but is virtually invariant to factors that can be manipulated by public policy, including income tax laws, welfare benefits, and divorce laws.

I. Introduction

The Deficit Reduction Act of 2005 provides $150 million per year to promote healthy, stable marriage in the United States. This initiative is based on three key premises: that marriage has a causal effect on the well-being of children and their parents, that public policy can affect marriage decisions, and that long-term marriage is a desirable outcome. Despite having produced an extensive literature on the causes and consequences of marriage, social scientists know surprisingly little about one aspect of the federal Healthy Marriage Initiative: What is the likelihood that a single person in the U.S. will experience a long-term union, and how does this likelihood differ when the union is formed via different paths (cohabitation versus marriage), at different ages, and by individuals with different backgrounds? In studying union formation, analysts have consistently focused on the probability of entering a marriage, entering cohabitation, or exiting a union in the "next period." This concentration on short-term transitions does not directly identify factors affecting the probability that an individual will enter a union and *maintain* that union for many years.

In this paper, we contribute to the understanding of stable marriage by assessing long-term union probabilities. We use data from the 1979 National Longitudinal Survey of Youth (NLSY79) to estimate a series of sequential choice models in which (*a*) single women with no prior unions decide whether to remain single, cohabit, or marry; (*b*) cohabiting women decide whether to continue cohabiting, separate, or marry; (*c*) first-married women decide whether to divorce; and (*d*) women with prior unions advance through subsequent stages of being single, cohabiting, and married. Rather than focus on predicted short-term transition probabilities obtained from each stage-specific model, we use the estimates to simulate women's union formation histories from ages 18 to 42. These simulated outcomes allow us to predict the probability that a woman with a given marital history and observed characteristics will marry (or cohabit) by a given age and then remain with her partner for the long term, which we define as eight or more years.

Our study has three interrelated goals. The first is to assess the probability of forming and maintaining long-term unions entered through both cohabitation and marriage. We find that a representative, 18 year-old woman with no prior unions has a 28% chance of marrying in the next four years, a 72% chance of staying married for at least eight years conditional on marrying within four years and, therefore, a 20% chance of marrying in the next four years *and* staying married for the long-term. If we consider cohabitation along with marriage, this woman's chance of entering a union within four years increases to 45%, her conditional probability of maintaining the union falls to 61%, and her joint probability of forming a union *and* remaining with her partner increases to 27%. These findings corroborate long-standing evidence that cohabitation is a common form of union entry, but that cohabiting unions are less likely than marriages to last (Bumpass and Lu 2000; Bumpass *et al.* 1991; Manning and Smock 2002; Smock 2000). By taking a long-term perspective, however, we are able to demonstrate that the entry effect dominates the exit effect—that is, the cohabitation option significantly enhances the probability that a woman will have a relationship lasting eight or more years.

Our second goal is to compare the probabilities summarized above for women at different ages and with different union-forming histories. We consider 18 year-olds with no prior unions, 24 year-olds with no prior unions, and 30 year-olds with at least one prior union. Interestingly, we predict that all three "types" have approximately the same 27-28% chance of entering a union within four years and maintaining it for at least eight years. Compared to either the 18 year-old or the 24 year-old, however, the 30 year-old "re-single" woman is far less likely to marry in the next four years (13% versus 23% for the 24 year-old and 28% for the 18 year-old) and far more likely to cohabit (32% versus 17% for both the 18 and 24 year-old). In general, we find that cohabitation is far more common among women with prior unions than among women without prior unions—and that the majority of long-term second unions begin with cohabitation.

Our third goal is to identify the effects on long-term union formation of various exogenous factors. We consider a range of demographic and family background factors (race, ethnicity, household composition, *etc.*), a set of values-related factors (religion, family attitudes), a number of skill measures (cognitive and noncognitive test scores, schooling attainment), and various marriage market characteristics as well as measures of legally-conferred costs and benefits that vary with marital status (prevailing divorce laws, expected income tax obligations, and welfare and Medicaid benefits). We find that predicted probabilities of entering and maintaining unions are often highly sensitive to demographic, family background, and skill-related factors, but are generally invariant to changes in environmental factors that can potentially be manipulated by public policy. Among 18 year-old women with no prior unions, for example, the predicted probability of entering a union within four years and maintaining it for at least eight years is only 13% for a black with a disadvantaged background, versus 27% for a representative woman and 29% for the same representative woman in a highly "pro-marriage" environment.

II. Background

Our analysis has three distinguishing characteristics. The first is that we model every stage of the union formation process—that is, we estimate transitions (*a*) from single to cohabiting or marriage, separately for women with and without prior unions; (*b*) from cohabiting to marriage or dissolution; and (*c*) from married to divorced, separately for first and subsequent marriages. A second characteristic is that our covariates include an unusually broad array of exogenous measures of family background, religious affiliation, earnings potential, marriage market characteristics, and legal/policy factors. Most significantly, our study is characterized by the manner in which we draw inferences. Whereas most analysts focus on estimated marginal effects of individual factors on short-term transitions, we compute simulation-based predicted probabilities of entering unions *and* maintaining them for many years. Because our broad-based approach links our analysis to virtually every existing study of the determinants of union formation, we do not attempt a comprehensive overview of the relevant literature. Instead, we point to select studies to illustrate how the current analysis can enhance our understanding of union formation.

Our long-term (or lifecycle) approach is not without precedent in the union formation literature. Light and Omori (2008) jointly model a multi-stage, union formation process similar to what we use here, but limit attention to first cohabitation spells and first marriages. Blau and van der Klaauw (2010) jointly model sequential transitions into and out of cohabitation and marriage along with transitions defined by the conception of a child. Van der Klaauw (1996) uses a dynamic programming model to estimate transitions into and out of marriage (ignoring cohabitation) jointly with labor force participation decisions, while Keane and Wolpin (2010) also use a dynamic programming model of marriage (without cohabitation), divorce, fertility, school enrollment, welfare participation, and labor supply. Each of these lifecycle approaches can, in principle, be used to "build" predicted probabilities of forming and maintaining long-term unions, but only Light and Omori (2008) provide such estimates. In other respects, Blau and van der Klaauw (2010), Keane and Wolpin (2010) and van der Klaauw (1996) extend the current approach by modeling outcomes (labor force participation, fertility, *etc.*) that are determined jointly with union formation and, in the latter two studies, by estimating dynamic structural models. We believe our study represents a useful middle ground between orthodox models that focus on single-stage transitions in the "next period," and more stylized, structural models such as Keane and Wolpin (2010) and van der Klaauw (1996).[1]

As noted, the majority of studies in the union formation literature focus on the estimated effect of a *single* determinant (or class of determinants) on a *single* stage of the union formation process. For example, Blackburn (2000) and Grogger and Bronars (2001) identify effects of welfare benefits on transitions to marriage among single, never-married women; Smock and Manning (1997) and Wu and Pollard (2000) examine the effects of employment and earnings on cohabitors' transitions into marriage; and Friedberg (1998) and Wolfers (2006) consider the effects of divorce laws on the probability of terminating a marriage. While single-stage analyses are the norm, some authors consider two different stages—*e.g.,* Bitler *et al.* (2004) identify effects of welfare policy on both marriage and divorce probabilities, while Martin and Bumpass (1989) and Teachman (1986) model marriage-to-divorce transitions separately for first and second marriages. Studies of this nature can be credited with providing most of what we know about the determinants of union formation, yet they are limited by a singular focus on transitions in the "next period."

Consider existing, single-stage evidence on the effects of schooling attainment on union formation. Regardless of whether schooling is of primary interest to the analyst, it is often included as a determinant in union formation models, and it is a rare factor for which a consensus exists on the signs (if not magnitudes) of its estimated effects. There is ample evidence that

[1]Life-table analyses such as Bramlett and Mosher (2002) and McCarthy (1978) also take a lifecycle view, but do not provide predicted probabilities of following multi-stage paths.

highly-schooled women are more likely than their less-schooled counterparts to transition from single (never married) to married in the next period in their mid 20s and beyond (Blackburn 2000; Sweeney 2002; Thornton *et al.* 1995; Waite and Spitze 1981). At the same time, it appears that schooling lowers the predicted probability of single-to-cohabitation transitions (Landale and Forste 1991; Thornton *et al.* 1995; Xie *et al.* 2003), and subsequently raises (lowers) the predicted probability of transitions from cohabitation-to-marriage (cohabitation-to-single)—although schooling effects on transitions out of cohabitation are often imprecisely estimated (Lichter *et al*. 2006; Manning and Smock 1995, 2002; Wu and Pollard 2000). Among women in both first and second marriages, evidence abounds that schooling lowers the probability of divorce in the next period (Lehrer and Chiswick 1993; Martin and Bumpass 1989; Phillips and Sweeney 2005; Teachman 1986).

Putting aside the fact that each stage-specific study uses a different sample and model specification, which precludes a meaningful comparison of point estimates, it is unclear how the evidence adds up. If schooling raises the probability that a single woman marries and subsequently lowers the probability that a married woman divorces—presumably because education is associated with greater gains to marriage (Becker 1974; Becker *et al.* 1977)—it stands to reason that highly-schooled women are more likely than less-schooled women to marry within a given interval *and* maintain that marriage for however long we wish to define the long-term. However, existing estimates of short-run transition probabilities do not directly identify the predicted probability of a given (multi-stage) long-term outcome, let alone how it differs for women with "low" versus "high" schooling. Moreover, in light of evidence that highly-schooled women delay marriage, it is unclear whether they are more likely than less-schooled women to experience the given long-term outcome if they are currently, say, 18 years old. It is equally unclear how schooling affects the probability of entering a cohabiting union within a given interval *and* maintaining the union for the long-term, given that schooling has been found to deter both entry into *and* exit from cohabiting unions.

Regardless of which determinant we consider (schooling, race, parental marital status, religion, marriage market conditions, *etc*.) it is impossible to infer from existing evidence how it affects the probability of forming a union and maintaining it for the long-term union. The literature has focused almost single-mindedly on predicting the probability of a particular, short-term transition (*e.g.,* single-to-married). We extend this approach by predicting probabilities of making a single-to-married transition (for example) in the next period, *or* the next period, *or* the next period, *or* the period after that—and then *not* transitioning from married-to-divorced in any of the subsequent eight periods.

III. Methods

A. Estimating Choice Models

We model the union formation process in five stages. In stage 1, single women with no prior marriage or cohabiting experience decide on an annual basis whether to stay single, cohabit, or marry; each woman in our sample begins the decision-making process in stage 1, which we initialize at age 18. Women who choose cohabitation as their first union advance to stage 2, in which cohabiters make annual decisions to continue cohabiting, dissolve their union, or marry. Women who transition from stage 1 or stage 2 into a first marriage advance to stage 3, where they decide whether to maintain their first marriage or divorce. Upon terminating their first (or a subsequent) cohabitation spell or marriage women enter stage 4, in which "re-single" women with prior marriages and/or cohabitations again decide each year whether to stay single, cohabit, or marry. Due to a relatively small number of cohabitation spells in our data, women who transit from stage 4 to cohabitation reenter stage 2. (That is, stage 2 consists of *all* cohabitation spells, rather than first cohabitation spells only.) Women who transition from stage 4 or stage 2 to (re)marriage enter stage 5, which consists of all marriages beyond the first.

More formally, we assume that in each 12-month interval, women choose the stage-specific alternative that maximizes their expected utility. We express the expected utility of alternative j for woman i in stage g at time t as a linear function of various observed and unobserved factors:

$$V_{igt}^{j} = \beta_{1g}^{j} X_{igt}^{j} + \beta_{2g}^{j} Y_{igt}^{j} + \beta_{3g}^{j} Z_{ig}^{j} + \varepsilon_{igt}^{j} \quad \text{for } j = s,c,m \quad \text{and } g = 1,2,3,4,5 \qquad (1)$$

where X_{igt}^{j} represents time-varying marital history factors (current spell duration, number of prior cohabitation spells, *etc*.), Y_{igt}^{j} represents other time-varying covariates (prevailing divorce laws and other environmental factors), Z_{ig}^{j} represents a host of time-invariant demographic, family background, and skill measures, and ε_{igt}^{j} represents unobserved factors affecting the value of alternative j for woman i in stage g at time t.[2] The model allows both observed and unobserved factors to vary across women, over time (within and between stages), and across alternatives, although a number of the factors are time-invariant. In addition, the parameters describing the effect of $X, Y,$ and Z on expected utility are allowed to vary across stages, given that current spell duration, divorce laws, and many other factors are likely to have a different effect on the value of, say, marriage if currently married versus marriage if currently single.

We assume the residuals (ε) are distributed according to the Type I Extreme Value Distribution, which means the stage 1, 2, and 4 models become multinomial logits and the stage 3 and 5 models are binomial logits. We assume the ε are independent across stages and across alternatives within each stage, but we compute standard errors to account for their within-stage correlation over time for a given woman i.

B. Simulating outcomes

We use maximum likelihood estimates of parameters $\beta_{1g}^{j}, \beta_{2g}^{j}$, and β_{3g}^{j} and their covariance matrix to simulate each woman's union-forming history from age 18 to age 42. For our baseline

[2]Note that alternative j=c (cohabitation) is unavailable when the current stage is g=3 or g=5 (marriage).

simulations, we assign each woman her actual, time-constant values of factors Z and her actual, stage 1 (g=1), age 18 (t=1) values for time-varying covariates X and Y. While covariates Z and Y are held constant at their fixed or initial values, we update history variables X on a period-by-period basis to reflect the woman's simulated outcome for that year. Because we only use actual covariate values observed at $t=1$, we can follow each woman over her entire 24-year history regardless of whether she participates in the NLSY79 for that long.[3]

In addition to simulations based entirely on *actual* covariate values, we conduct additional simulations after assigning each woman a uniform set of values for select covariates. For example, we assign each woman to be black (an element of covariate vector Z), or we assign each woman to have at least 16 years of school (an element of Y), while in both cases using actual values for all other covariates; this strategy allows us to predict outcomes for women of a certain type rather than for a representative sample of "actual" women. We describe the alternative types in detail in section IV.B.

For each set of covariate values that we select, each woman's history from age 18 to 42 is simulated for each of 150 random draws from the estimated distribution of the parameter estimates. The means and standard deviations of the simulated outcomes constitute our predicted probabilities of various long-term patterns: (*a*) transitioning from single (with no prior unions) to a first cohabiting union, or a first marriage, or either between ages 18 and 22, and maintaining that union for at least eight years; (*b*) making the same transitions between ages 24 and 28, and maintaining the union for at least eight years; and (*c*) transitioning from separated or divorced to a cohabiting union, or marriage, or either between ages 30 and 34—conditional on terminating a first union of either type at age 30—and maintaining the new union for at least eight years.

IV. Data

A. Sample Selection

We estimate the multi-stage choice models described in section III using data from the 1979 National Longitudinal Survey of Youth (NLSY79). The NLSY79 began in 1979 with a sample of 12,686 individuals born in 1957-1964. The original sample contains 6,283 women (49.5% of the sample), 2,002 Hispanics (15.8% of the sample), 3,174 blacks (25.0% of the sample) and 7,510 non-Hispanic, nonblacks ("whites"). Respondents were interviewed annually from 1979 to 1994 and biennially thereafter, although only 7,757 respondents remained in the survey by 2008 due, in part, to the intentional dropping of over-samples of military participants and low-income whites.

In selecting a sample for our analysis, we first confine our attention to the 6,283 women in the original NLSY79 sample. We eliminate men from our sample because the determinants of union

[3]For a similar use of simulated outcomes in different applications, see Angeles *et al.* (2005) and Blau and van der Klaauw (2010).

transitions are often found to differ for men and women (see, for example, Alm and Whittington 1999; Burgess *et al.* 2003) and a gender comparison is beyond the scope of our study. Next, we eliminate women who are not observed from age 18 onward—that is, we eliminate women who are six months or more beyond their 18th birthday when interviewed in 1979, or who permanently leave the survey prior to reaching age 18. We also drop women who marry or cohabit before age 18. These selection criteria reduce the sample to 2,859 women born in 1960-64 whose complete union-forming histories are observed from our chosen initialization point (age 18) onward. We choose age 18 as the starting date for union-forming decisions because relatively few individuals cohabit or marry prior to this age, and an earlier date would reduce our sample size dramatically. Finally, we eliminate 98 women for whom key covariates cannot be constructed, generally because the state or county of residence is unknown. These selection rules yield a sample of 2,761 women observed at age 18 with no prior unions.

To estimate the stage-specific choice models, we form samples with annual observations for these 2,761 women from age 18 until they are last observed.[4] For each person-year observation, we then construct regressors and variables identifying the woman's current state (single, cohabiting, or married) and the transition (if any) made within the next 12 months. The union status variables and most time-varying regressors are based on data collected in an event history format, so we can construct values at 12-month intervals regardless of whether an interview occurred in each year.[5] To identify current union status and transition dates, we do not rely solely on marital status at each interview date. Instead, we use "clean" starting and ending dates for each marriage provided by the survey, all available information on cohabitation starting and ending dates, and identifiers for cohabiting partners and spouses. This information allows us to determine, *e.g.,* that a single woman transitions into cohabitation within the next 12 months even if she reappears as single at the next interview, or that a cohabiting woman transitions to single within the next 12 months even if she is still cohabiting (but with a new partner) at the next interview.

We estimate separate binomial or multinomial logits for five stages. Stage 1 consists of initial single spells. Our stage 1 sample has 20,810 person-year observations for 2,761 women, all of whom begin the spell at our chosen initialization age of 18 and have no prior unions.[6] The stage 2 sample contains 4,721 person-year observations for 1,292 women who are observed cohabiting at any point during the observation period. Because this sample is substantially smaller than our

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[4]Women are observed until their last interview date, which is no later than 2006 (the last survey year available to us when we constructed the data).

[5]We describe the covariates in section IV.B. The only time-varying variables that are not based on event history data are county- and state-specific environmental variables. Residential location is generally only known at the time of each interview, so we assume unknown residential changes take place half way between successive interview dates.

[6]Table A-1 indicates sample sizes for all five stage-specific samples.

samples of single spells and marriages, we opt not to disaggregate stage 2 into first cohabitation spells and subsequent spells. The stage 3 sample contains 25,566 person-year observations for 2,178 women who are observed during their first marriage. Clearly, women who transition from single (no prior unions) to cohabiting to first marriage appear in stages 1, 2, and 3, while women who transition directly from single to first marriage appear only in stages 1 and 3. Stage 4 consists of all single spells experienced by women with prior cohabitation spells and/or marriages; this sample contains 10,850 person-year observations for 1,492 women. The stage 5 sample contains 6,092 observations for 724 women observed in second and third marriages.

B. Covariates

Aside from variables that track each woman's cohabitation and marriage history (current spell duration, number of past cohabitation spells, *etc*.), we focus on covariates that are exogenous to union-forming decisions. Rather than control for current employment status, cumulative labor market experience, household composition, school enrollment, and other factors that are determined jointly with union transitions, we use time-invariant demographic, family background, and skill measures in combination with an array of time-varying environmental factors. Table 1 contains means and standard deviations for the person-year samples for stages 1 and 3, as well as for the stage 1 simulation sample consisting of first observations for each woman.

The union-related history variables include current spell duration (in years) and its square to account for duration dependence in the probability of exiting each stage. We experimented with more flexible functional forms before determining that a quadratic adequately captures duration dependence for each stage. Our history covariates also include the age at which the current spell began, the number of prior cohabitation spells, the number of prior marriages, and a dummy variable indicating whether the woman cohabited with her husband prior to the current marriage. We include these history measures in light of evidence that transitions into and out of unions are significantly affected by prior cohabitation and marriages (Brien *et al.* 2006; Reinhold 2010; Svarer 2004; Teachman 2008). Aside from duration and its square, each covariate in this group is included in only a subset of stages—*e.g.*, age at which the spell began is included in stages 2-5 but is fixed at 18 for stage 1, while the "cohabited with spouse" indicator is only relevant to the marriage stages.

Family background variables (all of which are time-invariant) include indicators that the woman is black, Hispanic, and foreign born. We also control for the woman's mother's highest grade completed, and whether at age 14 the woman lived with her mother only, with her mother plus a stepfather, or with her mother and father; the omitted group is any living arrangement that excludes the mother. We also include an "access to reading materials" dummy variable that equals one if the woman reports that magazines, newspapers, and/or a library card was available in her home at age 14. We control for these factors because race, ethnicity, parental marital status and socioeconomic status have been established in the literature as important (and exogenous) determinants of union transitions (*e.g.,* Bennett *et al.* 1989; Manning and Smock 1995; Phillips and Sweeney 2005).

To explore the effects of religious affiliation and attitudes on union formation and dissolution, we control for a set of dummy variables indicating whether the woman was raised Baptist, Catholic, another Christian denomination (Methodist, Lutheran, *etc.*), or any other religion; women who claim no religion form the omitted group. We also control for whether the woman reports attending church at least once a week, or (if not weekly) at least once a month; the omitted group is infrequent or no church attendance. To control for whether the woman has traditional values, we count the number of times she agrees or strongly agrees with such statements as "a woman's place is in the home, not in the office or shop," and "women are much happier if they stay at home and take care of their children." We use seven such questions on women's roles, so scores range from zero (liberal) to seven (traditional). Each variable in this group is based on responses provided in 1979, when the women in our sample were 14-18. We control for this set of variables because religiosity and traditional values have been shown to influence union formation (Clarkberg *et al*. 1995; Lehrer and Chiswick 1993; Thornton *et al*. 1992), presumably because they reflect a distaste for cohabitation and divorce.

We control for each woman's skill level with the following set of time-invariant variables: an age-adjusted score for the Armed Forces Qualifications Test (AFQT), an age-adjusted score on the 10-item Rosenberg Self Esteem Index, and dummy variables indicating whether the woman's highest grade completed at age 35 is less than 12, 13-15, or 16 or more; a highest grade completed of 12 is the omitted group. The AFQT score is derived from scores on the Armed Services Vocational Aptitude Battery, which was administered to NLSY79 respondents in 1980, and the Rosenberg score is derived from a 10-item scale administered during the 1980 interview. Raw scores for both tests are summarized in table 1, but our regressions and simulations use residuals obtained by regressing raw scores on a set of birth-year indicators. We include AFQT scores and schooling attainment in our models as exogenous measures of earnings potential and financial independence; these factors are widely acknowledged to be important determinants of union formation for women (Oppenheimer 2000; Xie *et al*. 2003). Measures of noncognitive skill and personality traits are not often included in union entry and exit models (see Light and Ahn 2010; Lundberg 2010; and Schmidt 2008 for exceptions), but we use self-esteem as a measure of nonfinancial independence, or nonfinancial attractiveness to potential marriage partners.

Our final set of controls is intended to capture characteristics of marriage markets and exogenous, policy-driven costs and benefits associated with marriage. Following Lichter *et al.* (1991, 2002), we control for the percent of the woman's county population that is male, the percent that shares her race/ethnicity (black, white, or Hispanic), and the county's population density. To construct these variables, we use data from the City and County Data Book (U.S. Census Bureau) for the county of residence corresponding to each person-year observation; City and County Data Books are not available on an annual basis, so we use the closest available year for each observation.

We also control for the prevailing divorce laws with five dummy variables that characterize the ease with which divorce can be obtained, ranging from the most liberal environment (where divorces are granted and property settlements are determined without the need to establish fault) to the most conservative (where fault must be established for all divorces). Our environmental controls also include the maximum, monthly AFDC or TANF benefit available for a family of four; the average Medicaid expenditure for a family of four; and the expected state income tax this woman and her (expected) partner would pay if they were married net of their expected, joint tax obligation if they were single or cohabiting. Each of these policy variables is specific to the state of residence and calendar year corresponding to the person-year observation, and is strictly exogenous in the sense of reflecting the legal climate rather than the woman's family income, family size, or benefit eligibility.[7] The effects of state divorce law on union transitions have been explored by Friedberg (1998), Peters (1986), Wolfers (2006), and many others. Bitler *et al.* (2004), Blackburn (2000), Grogger and Bronars (2001), and Yelowitz (1998) analyze marriage incentives in welfare and Medicaid programs, while Alm and Whittington (1999) and Whittington and Alm (1997) have assessed effects on union transitions of marriage penalties (or bonuses) implicit in the *federal* income tax code. Blau and van der Klaauw (2010) and Light and Omori (2008) consider this entire array of policy factors in multi-stage models of union formation.

We use the covariate groups described above to define several distinct "types" of women for whom we simulate cohabitation and marriage outcomes. Our first set of simulations assigns each woman her actual covariate values.[8] We then assign each woman to be black and non-Hispanic while maintaining her actual values for all other covariates. To assess further the effects of family background, we make all women a disadvantaged black by assigning Black=1, Hispanic=0, foreign born=0, mother's highest grade=9, lived with single mother at age 14=1, and

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[7]We provide details on the construction of these variables in the appendix. Our exogeneity claim relies on the assumption that women do not choose their state of residence *in conjunction with their marital status* to lower divorce costs, reduce income taxes, or increase welfare or Medicaid benefits. Short of modeling migration decisions, the only alternative to assuming state of residence is exogenous is to include state fixed effects in our models. We opt not to use this identification strategy because within-state (intertemporal) variation in each factor is fairly systematic: over time, divorce laws move towards "no fault," tax law becomes more marriage neutral, and welfare-related costs of marriage are reduced. Because we cannot separate the effects of these temporal trends from aging effects—and because most of the variation in each factor is between states—we prefer to rely on cross-state variation in our data.

[8]As noted in section III, we use the "period one" value for time-varying variables, and update values of history variables (spell duration, number of prior marriages, *etc.*) on the basis of each period's simulated outcome.

access to reading materials at age 14=0. Turning to the religion/values variables, we first assign all women weekly church attendance, and then assign a full set of religion factors by making weekly church attendance=1, Baptist=1 (given that the Baptist indicator generally has a bigger estimated effect on union transitions than other religion categories), and traditional values=4, to correspond to the 90th percentile for the full sample of 2,761 women. Next, we consider two types of high-skill women: we assign all women a highest grade completed of 16 or more years, and then assign all women 16+ years of schooling plus AFQT and self esteem scores equal to the 90th percentile for the full sample of 2,761 women. Finally, we place all women in a favorable marriage market by assigning a county sex ratio corresponding to the 90th percentile for the full sample of 2,761 women. We then consider a more broadly defined pro-marriage environment by assigning all women the 90th percentile county sex ratio, plus a state without unilateral divorce (the omitted group for our divorce law categories), and a state income tax marriage penalty corresponding to the 10th percentile in the overall distribution.

V. Findings

A. Estimated Marginal Effects for One-Year Transitions

Maximum likelihood estimates for all five stage-specific choice models appear in table 2. Before turning to the simulated, long-term outcomes based on these estimates (section V.B), we consider conventional, short-term marginal effects. Table 3 shows the estimated marginal effects for a select group of covariates that play a key role in our simulations. These marginal effects are computed from the stage 1-3 estimates in table 2, setting all other factors equal to stage-specific sample means.

For the most part, the estimated marginal effects in table 3 have the expected sign. Single black women with no prior unions are 4.1 and 4.9 percentage points less likely to enter cohabitation and marriage, respectively, than are their observationally equivalent nonblack counterparts; if these women *do* cohabit, they are 8.2 percentage points less likely than nonblacks to marry their partner and an imprecisely estimated 1.2 percentage points more likely to separate from their partner. Women who are raised Baptist, who attend church regularly, or who hold traditional family attitudes are generally less likely than others to cohabit and more likely to marry—although, interestingly, a 1-point increase in traditional views is associated with a 0.2 percentage point decrease in the predicted probability of a single-to-cohabitation (SC) transition in stage 1, but also a 1.2 percentage point reduction in the probability of a cohabitation-to-marriage (CM) transition in stage 2. A ten percentage point increase in the county male sex ratio is predicted to raise the stage 1 likelihood of cohabiting by 2.8 percentage points; its estimated effect on entry into marriage is smaller in magnitude and statistically insignificant. Laws that allow no-fault divorce and property settlement (which can be viewed as lowering the cost of divorce) are associated with an increased probability of union dissolution (although the estimated marginal effects are imprecisely estimated), while an increase in the cost of marriage via a larger income tax marriage penalty is predicted to deter entry into marriage among single women (stage 1) and cohabiting women (stage 2).

While the estimates summarized in table 3 contain few surprises, it is noteworthy that only two covariates prove to be consistently union enhancing or detracting—that is, to have a negative estimated effect on all transitions *into* unions and a positive estimated effect on all transitions *out of* unions if detracting (identified in table 3 with a shaded cell), or to have the opposite estimated effects if enhancing. The estimated marginal effects of "black" and "lived with single mother" are, unsurprisingly, consistently union detracting, but all other covariates are found to have inconsistent estimated effects on union formation. For example, single women with 16+ years of schooling are predicted to be at least two percentage points less likely than other women to *enter* cohabitation or marriage (stage 1), but also at least one percentage point less likely to *exit* cohabitation or marriage (stages 2-3); this corresponds to other authors' findings that we summarized in section II.[9] Clearly, these estimates leave us unable to predict qualitatively whether highly-schooled, single women are more or less likely than their counterparts to form a union and maintain it for eight years. Our simulations are designed to facilitate this type of long-term inference.

Even with clear-cut determinants such as "black," our simulations produce new evidence that cannot be gleaned from the stage-specific estimates shown in tables 2 and 3. From the stage 1 marginal effects in table 3, one might crudely estimate that an 18 year-old, single black woman is 16 percentage points (0.04x4) less likely than a representative woman to cohabit or marry by age 22 and that if she marries, she is equally likely to remain married for eight or more years. While these back-of-the-envelope calculations are roughly consistent with the long-term predictions presented in section V.B, the table 3 estimates do not reveal the bottom line: What is the likelihood that an 18 year-old, single black woman will form a union by age 22 and maintain it for at least eight years? Moreover, a focus on estimated marginal effects of individual covariates does not allow us to assess, say, the effect of being black *and* being non-Hispanic *and* having been raised by a single mother, *and* having any number of other characteristics.

B. Simulation-Based Predicted Probabilities of Long-Term Unions

In table 4, we present predicted probabilities of entering early first unions (by age 22) and maintaining those unions for at least eight years for a sample of women who are single (with no prior unions) at age 18. After discussing the patterns seen in table 4, we proceed to tables 5-6, which consider later first unions formed between ages 24 and 28, and second unions formed between ages 30 and 34; in both cases, we condition on being single at the initial age of 24 (table 5) or 30 (table 6). The estimates in all three tables are based on the same simulated outcomes from age 18 to 42 for a uniform sample of 2,761 women—although the table 4 estimates only use simulated outcomes from age 18 to 30 (eight years past the latest union entry age), and the

[9] We estimate a negative relationship between schooling and single-to-marriage transitions because we do not focus exclusively on older and/or nonenrolled women.

table 5 estimates only use outcomes through age 36. Note that all sample means presented in tables 4-6 are statistically significant at 5% levels.

The column A estimates in table 4 are based on simulations in which each woman is assigned her actual, first period covariate values; history variables are updated on the basis of each period's predicted outcome. Rows a-c reveal that the predicted probability that an 18 year-old, single woman with no prior unions will marry (without cohabiting) by age 22 is 0.28; the predicted probability that she first cohabits by age 22 is 0.17; and the predicted probability that she forms a union of either type is 0.45 (0.28+0.17).[10] The conditional probabilities in rows a′-c′ indicate that among women who make a single-to-married transition by age 22, 72% are predicted to remain married to the same spouse for at least eight years. Among women making a single-to-cohabitation transition, 42% are expected to remain with the same partner for eight or more years; this includes women who convert their cohabiting union to marriage.[11] All told, 61% of women who form a union by age 22 can be expected to maintain the union for at least eight years. Finally, the joint probabilities in rows a″-c″ reveal that 27% (0.45·0.61) of single, 18 year-old women are expected to form a union by age 22 *and* maintain that union for at least eight years; these women have a 20% chance of marrying (without prior cohabitation) and remaining with their spouse for at least eight years, but only a 7% chance of cohabiting and remaining with their partner for that long.

We can draw three broad inferences from the column A estimates. First, cohabitation is a common form of union entry that raises the predicted probability of forming an early first union by 17 percentage points, or by 59% relative to the predicted probability of forming an early first marriage. Second, an early marriage (without prior cohabitation) has a 72% chance of lasting eight years, which is 1.7 times greater than the likelihood that a union formed via cohabitation will last that long. Third, the entry effect dominates the exit effect in the sense that cohabitation increases the predicted joint probability by seven percentage points, or 35% (0.27 versus 0.20). In short, cohabitation is less likely than marriage to lead to a long-term union, yet is sufficiently common as a form of entry that it substantially increases the overall chance of experiencing a long-term union.

Turning to the estimates in columns B-E of table 4, an unsurprising finding is that black women and especially disadvantaged black women are substantially less likely than others to enter first unions of either type (rows a-b), to maintain first unions conditional on forming them (rows a'-

[10]To clarify the interpretation of these estimates, 28% of simulated paths from age 18 to 22 have the form SM*, SSM*, SSSM*, or SSSSM*, where the asterisk represents the fact that simulated outcomes beyond the initial single-to-married transition are irrelevant for this computation. Similarly, 17% have simulated paths of the type SC*, SSC*, SSSC*, or SSSSC*.

[11]Unsurprisingly, the majority of women who remain with their cohabiting partner for at least eight years convert the cohabitation to marriage. Only 5% of the simulated paths underlying the row b' prediction reveal eight consecutive years of cohabitation.

b'), or to enter *and maintain* a long-term union of either type (rows a″-b″). For example, we predict that poor blacks (column B') have a 12.5% chance of forming *and* maintaining a long-term union entered via cohabitation *or* marriage, which is less than half as large as the corresponding column A estimate. Table 4 also reveals that high skill women (columns D and D') are the next-least likely of any type to enter early first unions: the predicted probability of entry into any union by age 22 (row c) is more than ten percentage points lower than the column A estimate, albeit higher than the estimates for blacks. However, the conditional probability of remaining with one's partner for eight years is estimated to be higher for skilled women (0.78-0.85 for unions entered via marriage, 0.47-0.51 for unions entered via cohabitation) than for any other type, so the predicted joint probabilities of entering *and* maintaining an early first union are only slightly lower than what is seen in column A (0.23 versus 0.27). More surprising is the finding that the remaining types considered in table 4 are predicted to differ only slightly from the representative women in column A. Women who attend church regularly (column C) and women with a more broadly-defined set of traditional values (column C′) are somewhat more likely than their column A counterparts to marry by age 22 (30-34% versus 28%) and slightly less likely to cohabit (13-15% versus 17%), but their overall chance of entering a union by age 22 remains at about 45%. The simulated outcomes for women in a favorable marriage market (column E) and women in a more broadly-defined "pro-marriage" environment (column E′) are virtually identical to the column A estimates.

To summarize, the findings in table 4 corroborate well-established evidence that blacks have lower probabilities of forming unions than nonblacks, and that highly-schooled women tend to delay marriage. Among our new findings is the evidence in row c": the predicted probability of entering a union (via cohabitation or marriage) by age 22 and remaining with the same partner for at least eight years ranges from 0.12-0.15 for blacks to 0.23 for high skill women to 0.27-0.30 for all other types of women that we consider. Looking across columns, the predicted probabilities in row c" are 2-8 percentage points (20-40%) higher than the estimates in row a". We consistently find that single-to-cohabitation transitions occur frequently enough to raise substantially the predicted probability that a woman will enter and maintain a long-term, early union.

To learn how the likelihood of long-term unions varies with the age of entry, in table 5 we consider later first unions that are formed between ages 24 and 28.[12] We find a number of striking differences between early and later first unions. First, the predicted probability of marrying (row a) is 3-7 percentage points lower (depending on type) for later unions than for early unions, while the predicted probability of cohabiting (row b) is one percentage point higher for all types; as a result, the predicted probability of entering *any* union falls slightly with age for

[12]In table 5 we condition on women who remain single (with no prior unions) to age 24; *i.e.,* women whose simulated outcome from age 19 to 24 is SSSSSS. We then require that women enter a union by age 28 to correspond to the four-year "at risk" window used for table 4.

all types. For example, in column A of table 5 the predicted probabilities of marrying or forming any union are 0.23 and 0.40, respectively, versus 0.28 and 0.45 in table 4. Second, all predicted conditional probabilities (rows a'-c') are higher in table 5 than in table 4: first unions are more likely to last for at least eight years when formed late rather than early. Third, the resulting effect is to leave the predicted joint probability of forming *and maintaining* any union (row c'') virtually unchanged for all types. For example, we predict that women in a “pro-marriage” environment (columns E-E') have a 28-29% chance of experiencing an early long-term union and a 28% chance of experiencing a later long-term union.

To follow up on an issue emphasized earlier, table 5 reveals that cohabitation plays an even more prominent role in the formation of later long-term unions than it does for early unions. The predicted probabilities of forming and maintaining a late union (row c'' of table 5) are 3-10 percentage points (33-58%) higher than the predicted probabilities of forming and maintaining a late *marriage* (row a''). In table 4, the row c'' estimates are only 20-40% higher than the corresponding estimates in row a''. Cohabitation contributes more to the overall probability of experiencing a long-term *later* union because it becomes a more common form of union entry with age (row b), and because unions formed via cohabitation are much more likely to last for eight years (row b') when formed late versus early.

Finally, we consider how the likelihood of long-term unions differs for women who have separated from a previous partner. To compute the estimates in table 6, we condition on women whose simulated outcome has them forming a first union via cohabitation or marriage and separating/divorcing at age 30. We then consider the probability of reentering cohabitation or marriage within the next four years and maintaining the second union for at least eight years.

The estimates in table 6 highlight the increased prominence of cohabitation over the lifecycle as a pathway to long-term unions. For every type of woman considered, the predicted probability of a single-to-marriage transition (row a) and the predicted joint probability of entering and maintaining a marriage (row a'') is lower for “re-single” women who divorce/separate at age 30 than it is for single women forming both early and later first unions. For example, in column A of table 6 we estimate that women have a 13% chance of re-marrying by age 34 and a 10% chance of remarrying and remaining remarried for at least eight years; the comparable estimates are 28% and 20% in table 4, and 23% and 18% in table 5. When it comes to cohabitation, however, the predicted entry, conditional, and joint probabilities (rows b, b', and b'') are much higher for women with prior unions than what we saw in tables 4-5. For example, a divorced/separated woman in a pro-marriage environment is predicted to have a 34% chance of cohabiting within four years (row b), a 55% chance of remaining with that partner for eight years (row b') and a 18% chance of entering and maintaining a union formed via cohabitation (row b''); the comparable estimates in table 5 are 19%, 49% and 10%.

The bottom line revealed by table 6 is that 30 year-old women with prior unions do *not* differ much from 18 year-old or 24 year-old women with no prior unions in the predicted probability of entering and maintaining a union: the estimates in row c'' of table 6 range from 0.14 for poor blacks to 0.17 for nonpoor blacks to around 0.30 for every nonblack type, while the row c'' estimates in table 4 range from 0.13 for poor blacks to 0.30 for women with traditional values. However, 30 year-olds with prior unions are more likely to form a new union via cohabitation than via marriage (predicted probabilities are greater in row b than in row a), and their likelihood of maintaining a union entered via cohabitation is remarkably high (row b'). As a result, their predicted probability of entering and maintaining *any* union for 8+ years (row c'') is two to three times higher than their predicted probability of entering and maintaining a marriage. Moreover, cohabitation makes the *largest* absolute and relative contribution to the probability of experiencing a long-term union (0.047) for disadvantaged blacks, which is the group with the *lowest* chance of experiencing a long-term union.

VI. Conclusions

Most analysts who study the determinants of union formation focus on the estimated effects of individual covariates on short-term transition probabilities—*e.g.,* they consider the estimated effect of a particular religious affiliation on the probability that a single woman cohabits or marries in the next year, or the estimated effect of divorce law on the probability that a married couple dissolves their union in the next year. Studies of this nature form a literature of undeniable breadth and influence, yet the consistent focus on entry or exit fails to identify determinants of long-term unions. Our contribution is to estimate a series of stage-specific transition models, use the estimates to simulate women's union-related outcomes from age 18 to 42, and then predict the probability of various long-term paths. Specifically, we consider the probability of forming a union (via cohabitation and/or marriage) in the next four years, the probability of maintaining that union for at least eight years, and the joint probability of forming the union and maintaining it for the long-term. Our analytic strategy can be viewed as a tractable middle ground between estimation of single-stage, short-term outcomes and more stylized, structural estimation.

Some of our findings corroborate well-established evidence that black women are less likely than others to form unions, that college-educated women tend to delay union formation, and that cohabiting unions are less likely than marriages to endure. Even here, however, we provide new insights: because of its emphasis on year-to-year transitions, existing research has not revealed that, for example, 18 year-old black women are given a 27% chance of entering a union (via cohabitation or marriage) within four years, but only a 15% chance of forming a union and maintaining it for at least eight years.

Other findings are more surprising—especially those that highlight the increased role that cohabitation plays in long-term union formation as women age. We predict that a representative, 18-year old woman with no prior unions has a 28% chance of marrying by age 22, and a 45% chance of marrying *or* cohabiting; the cohabitation option raises the predicted chance of union entry by 17 percentage points, or 60% relative to the (marriage-only) baseline. We predict that this same woman has a 20% chance of marrying by age 22 *and* remaining with her husband for at least eight years, and a 27% chance of forming any union (via cohabitation or marriage) *and* maintaining it for at least eight years; because cohabitation tend to be short-lived, it raises the predicted probability of entering and maintaining an early union by only 35%. However, for a woman who ends her first union at age 30, cohabitation raises the predicted probability of union entry by 240%, and raises the predicted probability of entering *and* maintaining a second union by 180%. Cohabitation spells are far less likely than marriages to last for the long-term, yet the sheer number of unions that begin with cohabitation (especially among older women) leads to a significant increase in the chance that a woman will experience a long-term union. In our view, the substantial impact of cohabitation on long-term union formation has not previously been fully understood.

Another key finding is that the predicted probability of a long-term union is sensitive to a number of observed factors—but not to factors that can potentially be manipulated by public policy. We consistently find that black women (especially those with a disadvantaged background) are much less likely than a representative woman to form a long-term union. However, none of the policy factors that we considered (divorce law, expected income tax obligations, expected Medicaid benefits) proved to have an important effect on the predicted probability of long-term unions. Our study remains silent on the issue of whether public policy *should* be used to promote union formation, and on the link between long-term unions and healthy unions—but our findings suggest that incentives provided by tax policy, divorce law, and welfare benefits are unlikely to have an important effect on women's decisions to enter unions and maintain them for the long-term.

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Appendix: Construction of Environmental Covariates

Divorce variables: Using the state of residence and calendar year corresponding to each person-year observation, we characterize the prevailing divorce law using a five-way classification scheme. "Unilateral divorce and no fault for property" means the woman can obtain a unilateral divorce *and* property settlement without having to establish "fault" (marital misconduct) and without a mandatory separation requirement. "Unilateral divorce and fault for property" refers to a more restrictive environment in which unilateral, no fault, no mandatory-separation divorce is granted, but fault must be established for the court to make a property settlement. "Unilateral, mandatory separation ≤ 1 year" means no-fault divorce is granted only after a mandatory separation of one year or less, while "unilateral, mandatory separation > 1 year" means a mandatory separation of more than one year is required before a unilateral, no-fault divorce is granted. The omitted group identifies state-year observations where fault must be established for a divorce to be granted. This five-way taxonomy is ordered in the sense that "unilateral, no fault for property" laws are considered to impose the lowest costs on divorce, while the lack of unilateral divorce (the omitted group) is considered to raise the cost of divorce. Individual women may encounter exceptions depending on whether they invoke community property laws, seek alimony or child custody agreements, and/or have a prenuptial agreement. We do not define our covariates on the basis of these individual characteristics because they may be endogenous to union-related outcomes.

We construct our divorce variables using information in Friedberg (1998) supplemented by data available at abanet.org.

AFDC/TANF benefits: We assign the state- and year-specific maximum, monthly AFDC or TANF benefit available to a family of four, divided by the implicit price deflator for gross domestic product. These variables are independent of each woman's income, family status, and other determinants of her AFDC/TANF eligibility status, all of which are likely to be endogenous to union formation. These values are taken from the welfare benefit database ben_dat.txt available at http://www.econ.jhu.edu/People/Moffitt/datasets.html.

Medicaid expenditure: We assign the state- and year-specific average Medicaid expenditure for a family of four, divided by the consumer price index for medical care. As with our AFDC/TANF measures, this "expected" benefit is independent of individual (and potentially endogenous) factors such as income and family status. The data are from the Urban Institute's welfare rules database available at http://www.urban.org/toolkit/databases/index.cfm

State income tax marriage penalty: To construct this variable, our first step is to use 1979-2006 NLSY79 data for all male and female respondents who are age 18 or older to estimate earnings models for eight separate samples defined by marital status (single, cohabiting, married, or divorced) and sex. We use individuals' total earnings for the prior year as the dependent variable. Regressors are the year-specific implicit price deflator for personal consumption expenditure; county- and year-specific per capita income; the county- and year-specific unemployment rate; a quartic in age; age-adjusted AFQT scores; dummy variables indicating the current highest grade completed is 0-11, 12, 13-15, or 16+; and indicators for whether the respondent is black or Hispanic. Our second step is to use the eight sets of estimated parameters to compute year-specific, race/ethnicity-specific, sex-specific, marital status-specific predicted incomes, which we use to identify the *median* predicted income for each sample. In step 3, we associate each person-year observation in our stage 1-5 samples with the median predicted incomes for both men and women in the same stage (single, cohabiting, married, or divorced), in the same year, and with the same race/ethnicity as the respondent. Our final step is to use Taxsim (available at http://www.nber.org/~taxsim/) to compute the state income tax liability for the median man and median woman, first assuming they are married and filing jointly, and then assuming they are single or cohabiting and filing separately.

The difference between the state income tax liability if married and the tax liability if single or cohabiting is the variable used in our state-specific choice models. This variable identifies the expected income tax penalty (or bonus) associated with marriage for a median woman who shares the sample member's race/ethnicity, marital status, and state of residence, and who has a (potential) partner with the same race/ethnicity and marital status. Our measure is correlated with tax obligations based on actual income, but within-stage variation is entirely dependent on cross-year and cross-state variation in income tax laws. We rely on state income tax laws rather than federal income tax laws because the latter only varies across years, and is difficult to separate from aging effects.

Table 1: Summary Statistics for Stage 1 and Stage 3 Samples

	Stage 1: 1st single spell				Stage 3: 1st marriages	
	First observation		All observations			
Covariates	Mean	S.D.	Mean	S.D.	Mean	S.D.
Spell duration, years	0		6.82	6.79	8.41	6.14
Age began spell	18	0.00	18	0.00	23.63	4.35
Number of prior cohabitation spells	0		0		.36	.62
Number of prior marriages	0		0		0	0
1 if cohabited with spouse before marriage	0		0		.26	
Family background						
1 if black	.29		.42		.23	
1 if Hispanic	.17		.16		.19	
1 if foreign born	.07		.06		.07	
Mother's highest grade completed	10.86	3.04	10.94	3.16	10.90	3.10
1 if lived with mother, age 14	.20		.23		.15	
mother and stepfather, age 14	.08		.07		.08	
mother and father, age 14	.65		.63		.71	
1 if access to reading materials, age 14	.89		.89		.90	
Religion and values						
1 if Baptist	.29		.33		.26	
Catholic	.35		.32		.38	
other Christian	.20		.20		.21	
other religion	.12		.11		.12	
1 if attends church 1+ times/month	.21		.23		.22	
1+ times/week	.45		.45		.47	
Traditional values score	1.83	1.66	1.83	1.81	1.80	1.65
Skill levels						
AFQT percentile score	42.27	27.83	41.02	28.98	46.42	27.63
Rosenberg self esteem score (10-34)	18.17	4.01	18.09	4.06	17.96	3.93
1 if final highest grade completed < 12	.08		.07		.05	
= 13-15	.26		.27		.27	
≥ 16	.23		.28		.27	
Environmental factors						
Percent same race in county	59.50	33.58	52.27	32.43	60.03	31.31
Percent men in county	48.44	1.28	48.44	1.28	48.84	1.24
Population density in county/1000	2.26	6.30	2.93	7.47	1.69	5.33
1 if unilateral divorce + no fault for property	.22		.20		.24	
+ fault for property	.29		.30		.29	
+mandatory separation ≤ 1 year	.15		.14		.15	
+mandatory separation > 1 year	.15		.15		.14	
Max. monthly AFDC/TANF benefit, $100s	6.12	2.53	4.92	2.88	3.28	3.09
Average Medicaid expenditure, $100s	4.13	1.08	3.50	1.70	2.42	2.09
State income tax marriage penalty, $100s	.25	.79	.46	1.17	.28	1.17
Number of observations	2,761		20,810		25,566	
Number of women	2,761		2,761		2,178	

Note: Covariates also includes indicators that mother's highest grade completed, AFQT score, and self esteem score are missing; stage-specific sample means are used to replace missing values. Summary statistics for stages 2, 4 and 5 are available upon request.

Table 2: Multinomial Logit Estimates for Stages 1-5

Covariate	Stage 1: 1st single spell				Stage 2: All cohab. spells			
	S to C		S to M		C to M		C to S	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Constant	-5.277	1.590	-2.039	1.362	.795	1.971	-1.911	1.732
Spell duration	.107	.022	.058	.018	-.107	.037	-.273	.031
Spell duration squared/10	-.067	.013	-.068	.011	-.029	.041	.072	.023
Age began spell					-.027	.012	-.067	.010
Number of prior cohabitation spells					-.166	.127	-.221	.112
Number of prior marriages					.017	.088	-.036	.082
Family background								
1 if black	-1.256	.189	-1.003	.147	-.778	.228	-.032	.185
1 if Hispanic	-.683	.187	-.169	.155	-.439	.227	-.036	.194
1 if foreign born	-.179	.163	.119	.128	-.143	.220	-.092	.181
Mother's highest grade completed	.027	.015	-.020	.012	.029	.020	.045	.017
1 if lived with single mother, age 14	-.348	.151	-.185	.145	-.315	.176	.144	.139
mother and stepfather, age 14	-.063	.179	.042	.172	.028	.198	.165	.164
mother and father, age 14	-.565	.141	.016	.133	.015	.153	-.091	.135
1 if access to reading materials, age 14	.001	.133	.009	.114	-.083	.188	-.099	.122
Religion and values								
1 if Baptist	-.330	.186	.242	.180	.272	.218	-.052	.208
Catholic	-.297	.187	-.029	.180	.002	.210	-.160	.202
other Christian	-.327	.185	.104	.180	.229	.212	-.026	.202
other religion	-.367	.202	.250	.188	.186	.241	.182	.223
1 if attends church 1+ times/month	.058	.095	.054	.086	.086	.123	-.014	.108
1+ times/week	-.223	.088	.144	.072	.221	.109	.004	.092
Traditional values score	-.050	.024	.004	.019	-.098	.031	-.001	.026
Skill levels								
AFQT score	.002	.002	-.002	.002	.003	.002	-.002	.002
Self esteem score	.002	.010	-.025	.008	-.022	.011	.001	.010
1 if highest grade completed <12	.449	.136	-.344	.137	-.159	.148	-.223	.117
13-15	-.200	.095	-.251	.081	.208	.115	-.183	.010
16+	-.665	.110	-.557	.090	.316	.149	-.035	.138
Environmental factors								
Percent same race in county	-.007	.002	-.001	.002	-.002	.003	-.001	.002
Percent men in county	.079	.031	.023	.027	-.019	.039	.069	.035
Population density in county	-.013	.008	.009	.007	.001	.009	.001	.007
1 if unilateral divorce + no fault for property	-.121	.139	.015	.110	.296	.160	.018	.151
+ fault for property	.042	.114	.013	.091	.078	.147	.100	.124
+mandatory separation ≤ 1 year	.057	.132	.078	.107	.034	.159	.017	.143
+mandatory separation > 1 year	-.114	.133	.200	.100	-.093	.177	.062	.143
Maximum monthly AFDC/TANF benefit	.053	.020	-.061	.017	-.025	.025	.002	.021
Average Medicaid expenditure	-.062	.045	-.009	.029	.109	.036	-.063	.035
State income tax marriage penalty	.037	.030	-.025	.028	-.013	.042	.006	.036
Log likelihood	-8,505.38				-3,892.51			
Number of observations	20,810				4,721			
Number of women	2,761				1,292			

Continued.

Table 2: Continued

Covariate	Stage 3: 1st marriages		Stage 4: Non-1st single spells				Stage 5: Non-1st marriages	
	M to S		S to C		S to M		M to S	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Constant	-.565	1.670	-4.258	1.794	-4.853	2.581	3.017	2.584
Spell duration	.031	.021	-.147	.030	-.145	.036	-.006	.039
Spell duration squared/10	-.037	.010	.011	.020	-.005	.021	-.028	.022
Age began spell	-.045	.012	-.064	.010	-.108	.014	-.165	.019
Number of prior cohabitation spells	.271	.090	.062	.048	-.198	.076	.283	.086
Number of prior marriages			-.140	.069	-.060	.092	.227	.211
1 if cohabited with spouse before marriage	-.270	.126					-.489	.156
Family background								
1 if black	.005	.160	-.530	.176	-.851	.238	.594	.299
1 if Hispanic	-.133	.172	.158	.184	-.509	.249	-.100	.321
1 if foreign born	-.200	.165	-.376	.216	-.068	.215	-.472	.352
Mother's highest grade completed	.028	.015	.012	.018	.007	.022	-.010	.027
1 if lived with single mother, age 14	.332	.156	-.019	.157	-.054	.216	-.062	.263
mother and stepfather, age 14	.019	.181	.112	.181	.290	.247	-.189	.272
mother and father, age 14	-.035	.148	.076	.146	.180	.198	-.106	.240
1 if access to reading materials, age 14	.123	.120	.058	.132	.422	.206	-.504	.192
Religion and values								
1 if Baptist	.040	.196	.040	.190	-.063	.339	-.131	.354
Catholic	-.096	.200	-.060	.195	-.056	.337	.067	.364
other Christian	.075	.201	.056	.189	-.145	.343	-.344	.353
other religion	.068	.208	-.116	.205	-.026	.350	-.233	.375
1 if attends church 1+ times/month	-.181	.098	.042	.105	.160	.147	-.128	.177
1+ times/week	-.176	.082	-.131	.091	.283	.126	-.098	.145
Traditional values score	-.033	.023	.032	.026	.040	.034	.023	.036
Skill levels								
AFQT score	-.009	.002	.003	.002	-.001	.003	.001	.003
Self esteem score	-.006	.010	.001	.010	-.006	.014	-.027	.016
1 if highest grade completed <12	.228	.133	.119	.132	-.010	.181	.094	.237
13-15	-.071	.088	-.202	.103	-.067	.132	.150	.147
16+	-.359	.116	-.307	.141	-.148	.168	-.215	.217
Environmental factors								
Percent same race in county	-.001	.002	.003	.002	-.006	.003	-.003	.004
Percent men in county	-.035	.033	.088	.035	.132	.050	.012	.050
Population density in county	-.001	.007	.005	.007	-.017	.014	-.001	.014
1 if unilateral divorce + no fault for property	.341	.136	-.148	.145	-.361	.187	-.216	.225
+ fault for property	.311	.115	-.102	.128	-.263	.159	.034	.201
+mandatory separation ≤ 1 year	.088	.133	-.214	.148	-.279	.196	.251	.248
+mandatory separation > 1 year	.151	.134	.148	.152	-.301	.203	.212	.222
Maximum monthly AFDC/TANF benefit	-.043	.022	.022	.021	-.053	.031	-.047	.039
Average Medicaid expenditure	.006	.034	-.039	.032	-.076	.049	-.048	.058
State income tax marriage penalty	-.007	.032	-.030	.039	-.091	.068	-.047	.066
Log likelihood	-3,734.76		-4,514.04				-1,400.10	
Number of observations	25,566		10,850				6,092	
Number of women	2,178		1,492				724	

Note: Each stage includes indicators that mother's highest grade completed, AFQT score, and self esteem score are missing; stage-specific sample means are used to replace missing values. Standard errors account for nonindependence of residuals across observations for the same person.

Table 3: Estimated Marginal Effects of Select Covariates on One-Year Transitions

Covariate	Stage 1 SC	Stage 1 SM	Stage 2 CM	Stage 2 CS	Stage 3 MS
Black (0 to 1)	-.041 (6.44)	-.049 (6.69)	-.082 (4.05)	.012 (0.46)	.000 (0.03)
Foreign born (0 to 1)	-.006 (1.24)	.007 (0.94)	-.014 (0.60)	-.010 (0.40)	-.006 (1.32)
1-year increase in mother's highest grade completed	.001 (1.90)	-.001 (1.84)	.002 (1.01)	.006 (2.52)	.001 (1.83)
Lived with single mother, age 14 (0 to 1)	-.011 (2.44)	-.009 (1.25)	-.039 (2.16)	.029 (1.38)	.011 (1.92)
Access to reading materials, age 14 (0 to 1)	.000 (0.12)	.000 (0.08)	-.008 (0.34)	-.013 (0.73)	.003 (1.08)
Baptist (0 to 1)	-.012 (1.95)	.014 (1.36)	.036 (1.29)	-.015 (0.51)	.001 (0.20)
Attend church 1+ times/week (0 to 1)	-.008 (2.67)	.008 (2.12)	.027 (2.06)	-.005 (0.40)	-.005 (2.16)
1-point increase in traditional values	-.002 (2.08)	.000 (0.33)	-.012 (3.27)	.002 (0.63)	-.001 (1.47)
Highest grade completed ≥ 16 (0 to 1)	-.020 (6.44)	-.026 (6.44)	.043 (2.07)	-.014 (0.73)	-.010 (3.36)
10-point increase in AFQT score	.001 (1.04)	-.001 (1.11)	.004 (1.46)	-.007 (1.38)	-.003 (4.81)
1-point increase in self esteem score	.000 (0.34)	-.001 (3.29)	-.003 (2.08)	.001 (0.54)	-.000 (0.61)
10-point increase in percent men in county	.028 (2.53)	.010 (0.74)	-.040 (0.87)	.011 (2.16)	-.010 (1.07)
Unilateral divorce + no fault for property (0 to 1)	-.004 (0.99)	.012 (1.92)	-.012 (0.63)	.012 (0.54)	.005 (1.07)
100-dollar increase in state income tax marriage penalty	.001 (1.27)	-.001 (0.95)	-.018 (0.36)	.001 (0.25)	-.020 (0.21)
Unconditional transition probability	.037	.057	.141	.179	.031

Note: Figures represent the estimated marginal effect of the indicated change in the given covariate, holding all other covariates at the sample mean. Robust z-statistics are in parentheses. Shaded cells identify union-detracting marginal effects; *i.e.,* effects that decrease (increase) the likelihood of entering (exiting) unions.

Table 4: Predicted Probabilities of Entering and Maintaining "Early" First Unions for 18 Year-Old Women with no Prior Unions

Simulated pattern	Covariates used for simulations[a] Actual values (A)	Black (B)	Poor black (B')	Church-goer (C)	Trad'l values (C')	College (D)	High skill (D')	High % male (E)	Pro-mar. setting (E')
Entry into early (age 18-22) first union									
a. Marry by age 22; no prior cohabitation	.281	.172	.156	.301	.338	.226	.192	.284	.275
b. Cohabit by age 22; no prior marriage	.167	.093	.073	.146	.125	.115	.127	.181	.186
c. Marry or cohabit by age 22	.449	.265	.229	.446	.463	.341	.318	.465	.460
Conditional probability									
a'. Stay married for 8+ years, given a.	.717	.696	.666	.729	.738	.781	.849	.727	.769
b'. Stay with partner for 8+ years, given b.	.420	.348	.286	.443	.429	.470	.510	.404	.412
c'. Stay with partner for 8+ years, given c.	.606	.573	.544	.636	.654	.676	.714	.601	.625
Joint probability									
a''. Stay married for 8+ years, and a.	.202	.119	.103	.219	.249	.177	.163	.207	.211
b''. Stay with partner for 8+ years, and b.	.070	.032	.021	.065	.054	.054	.065	.073	.077
c''. Stay with partner for 8+ years, and c.	.272	.152	.125	.284	.303	.231	.227	.280	.288

Note: Predictions are the mean simulated outcomes for a sample of 2,761 women seen at age 18 with no prior unions. Each woman's history from age 18 to 42 is simulated 150 times, using a random draw from the estimated parameter distributions summarized in table 2. The standard error of each mean is no greater than 0.003.

[a]All women are assigned their actual covariate values with the following exceptions: All women are assigned black (column B), and black plus nonforeign born, mother completed grade 9, lived with single mother, no reading access (B'). All women are assigned regular church attendance (column C), and regular church attendance plus Baptist and a traditional values score at the 90th percentile (C'). All women are assigned highest grade≥16 (column D), and highest grade≥16 plus AFQT and self esteem scores at the 90th percentile (D'). All women are assigned county percent male at the 90th percentile (column E), plus a state without unilateral divorce and a state income tax marriage penalty at the 10th percentile (E'). See text for details.

Table 5: Predicted Probabilities of Entering and Maintaining "Later" First Unions for 24 Year-Old Women with no Prior Unions

Simulated pattern	Covariates used for simulations[a] Actual values (A)	Black (B)	Poor black (B')	Church-goer (C)	Trad'l values (C')	College (D)	High skill (D')	High % male (E)	Pro-mar. setting (E')
Entry into later (age 24-28) first union									
a. Marry by age 28; no prior cohabitation	.228	.147	.134	.242	.271	.191	.161	.233	.222
b. Cohabit by age 28; no prior marriage	.176	.106	.084	.156	.136	.128	.141	.190	.194
c. Marry or cohabit by age 28	.404	.253	.218	.398	.407	.319	.303	.422	.417
Conditional probability									
a'. Stay married for 8+ years, given a.	.770	.759	.738	.779	.789	.821	.879	.786	.813
b'. Stay with partner for 8+ years, given b.	.510	.447	.399	.529	.511	.546	.580	.484	.493
c'. Stay with partner for 8+ years, given c.	.667	.629	.608	.681	.696	.711	.739	.650	.664
Joint probability									
a''. Stay married for 8+ years, and a.	.175	.112	.099	.189	.214	.157	.142	.183	.181
b''. Stay with partner for 8+ years, and b.	.090	.048	.033	.082	.070	.070	.082	.092	.096
c''. Stay with partner for 8+ years, and c.	.265	.159	.132	.271	.284	.227	.224	.275	.277

Note: Predictions are the mean simulated outcomes for the same sample described in the note to table 4. We condition on women who remain single (with no prior cohabitation spells or marriages) at age 24.

[a]See table 4.

Table 6: Predicted Probabilities of Entering and Maintaining Second Unions for Women who End Their First Union at Age 30

Simulated pattern	Covariates used for simulations[a]								
	Actual values (A)	Black (B)	Poor black (B')	Church-goer (C)	Trad'l values (C')	College (D)	High skill (D')	High % male (E)	Pro-mar. setting (E')
Entry into second union at age 30-34									
a. Marry by age 34; prior union	.131	.091	.065	.148	.159	.156	.140	.148	.191
b. Cohabit by age 34; prior union	.317	.230	.231	.284	.305	.274	.303	.335	.337
c. Marry or cohabit by age 34	.448	.321	.296	.432	.463	.431	.443	.483	.528
Conditional probability									
a'. Stay married for 8+ years, given a.	.766	.683	.572	.752	.775	.793	.841	.742	.760
b'. Stay with partner for 8+ years, given b.	.569	.483	.427	.582	.576	.576	.638	.556	.546
c'. Stay with partner for 8+ years, given c.	.626	.540	.459	.640	.644	.655	.702	.613	.623
Joint probability									
a''. Stay married for 8+ years, and a.	.100	.062	.037	.111	.123	.124	.118	.110	.145
b''. Stay with partner for 8+ years, and b.	.180	.111	.099	.165	.176	.158	.193	.186	.184
c''. Stay with partner for 8+ years, and c.	.280	.173	.136	.276	.298	.282	.311	.296	.329

Note: Predictions are the mean simulated outcomes for the same sample described in the note to table 4. We condition on women who form a first union (cohabitation, marriage, or cohabitation that converts to marriage with the same partner) and then terminate that union at age 30.

[a]See table 4.