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

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Abstract: Most studies of union formation focus on short-term probabilities 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 at least 8, 12, or even 24 years. We use data for female respondents in the 1979 National Longitudinal Survey of Youth to estimate choice models for multiple stages of the union-forming process. We then use the estimated parameters to simulate each woman's sequence of union transitions from ages 18-46, and use the simulated outcomes to predict probabilities that women with given characteristics follow a variety of long-term paths. We draw three broad conclusions. First, a representative woman has the same 28% chance of cohabiting or marrying *and* maintaining the union for 12+ years regardless of whether it is a first union formed by age 22, a first union formed by age 28, or a second union formed by age 34. Second, unions formed via cohabitation contribute significantly to the likelihood of experiencing a long-term union, and this contribution grows over the lifecycle. This finding reflects the fact that the high probability of *entering* a cohabiting union more than offsets the relatively low probability of *maintaining* it for the long-term. Third, the likelihood of forming a union and maintaining it for the long-term is highly sensitive to race, but is virtually invariant to factors that can be manipulated by public policy such as divorce laws, welfare benefits, and income tax laws.

Keywords: Marriage, cohabitation, divorce, long-term unions

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 characteristics? 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 46. These simulated outcomes allow us to predict the probability that a woman with a particular set of observed characteristics will marry or cohabit by a given age *and* remain with her partner for the long-term. We consider alternative definitions of "long-term" ranging from eight to 24 years, although 12 years is our primary definition because it is the longest duration we can consider for unions formed both early (by age 22) and later (by age 34).

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 58% chance of staying married for at least 12 years conditional on marrying

and, therefore, a 16% 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 50%, and her joint probability of forming a union *and* remaining with her partner increases to 22%. 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; Osborne, Manning and Smock 2007). 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 substantially increases the probability that a woman will have a relationship lasting 12 or more years. This finding holds when we use shorter or longer definitions of "long-term."

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 one prior union. Interestingly, we predict that all three "types" have roughly the same 22-25% chance of entering a union within four years and maintaining it for at least 12 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 (31% 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. We find the same qualitative patterns when we shorten the definition of "long-term" to eight years or extend it (for first unions) to as long as 24 years.

Our final 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, attitudes toward gender roles), 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 benefits). We find that predicted probabilities of entering and maintaining unions are often highly sensitive to demographic, background, and skill-related factors, but are largely

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 12 years is only 12% for a black, versus 22% for a representative woman and 25% for the same representative woman in a highly "pro-marriage" environment.

II. Background

Our analysis has three distinguishing characteristics: First, we estimate choice models that follow women through every transition between single, cohabiting, married, and separated or divorced over three decades. Second, we control for an unusually broad array of exogenous measures of family background, religious affiliation, earnings potential, marriage market characteristics, and policy factors. Third, we draw inferences by computing simulation-based predicted probabilities of forming unions and maintaining them for many years, rather than focusing on predicted probabilities of entering or exiting a single state. 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.

We begin by noting that our analysis, with its focus on *determinants* of long-term unions, is complementary to studies that are concerned with *benefits* of stable unions. Causal evidence that long-term unions are "better" than shorter unions is provided by Korenman and Neumark (1991) and Stratton (2002), who demonstrate that the wage premium accruing to married men increases with marital duration. Even more compelling is extensive evidence that the *absence* of union stability—*viz*., divorce—can have harmful effects on adults and especially their children (Amato 2000; Duncan and Hoffman 1985; Fomby and Cherlin 2007; Osborne and McLanahan 2007). In our view, the extensive literature on the benefits of union stability will be enhanced by additional evidence on the likelihood that individuals with given characteristics form and maintain long-term unions at different stages of the lifecycle.

We make a particularly important contribution to the strand of the literature that argues that cohabitation contributes to union instability. Evidence abounds that cohabiting unions tend to be short-lived relative to marriages (Bumpass and Lu 2000; Lichter *et al.* 2006; Osborne, Manning

and Smock 2007), and that cohabitation is linked to subsequent marital instability (Axinn and Thornton 1992; Dush, Cohan and Amato 2003; Brien *et al*. 1999). Our findings corroborate existing evidence that cohabitation is a common form of union entry, and that cohabiting unions tend not to last as long as marriages. More surprisingly, we also find that the estimated joint probability of forming a first union by age 22 and maintaining it for at least 12 years is 36% *higher* when we include unions entered via cohabitation than when we focus solely on unions entered via marriage—and this statistic increases to a staggering 187% when we consider second unions formed by age 34. While there is no question that cohabitation accounts for a substantial portion of *short-term* unions, our analysis is the first to demonstrate that it is sufficiently prevalent to make a significant contribution to *long-term* union formation as well.

Turning to the literature on the determinants of union formation, the majority of studies identify effects of covariates of interest 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; Lichter *et al.* (2006), Smock and Manning (1997), and Wu and Pollard (2000) examine the effects of employment and other factors on cohabitors' transitions into marriage; and Friedberg (1998) and Wolfers (2006) consider the effects of divorce laws on the probability of terminating a marriage.[1] 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."

To appreciate the limitations of existing evidence consider, for example, the finding that decreased welfare benefits raise the predicted probability that a single woman transitions to marriage, while lowering the predicted probability that she cohabits (Blackburn 2000; Grogger and Bronars 2001). This finding suggests that welfare policy can, in principle, be used to promote entry into marriage—but can it be used to promote long-term union formation? To answer that question, we must identify effects of welfare benefits on transitions *into* cohabitation and marriage, *between* cohabitation and marriage, and *out of* cohabitation and marriage, and then "add up" the predicted probabilities of marrying or cohabiting (not necessarily in the next year, but by a particular age of our choosing) and then maintaining the union for the long-term. The literature has focused almost single-mindedly on predicting probabilities of short-term transitions (*e.g.,* single-to-married). A key contribution of our analysis is that we demonstrate whether covariates that are known to have important effects on transitions into or out of unions also have

important effects on the likelihood of entering unions *and* maintaining them for the long-term.

We conclude this section by noting that our multi-stage approach is not without precedent in the union formation literature. Bramlett and Mosher (2002) estimate transitions into and out of multiple stages (single, cohabiting, married, divorced). 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) and Steel *et al.* (2005) jointly model fertility and transitions into and out of cohabitation and marriage. Van der Klaauw (1996) estimates transitions into and out of marriage (ignoring cohabitation) jointly with labor force participation decisions, while Keane and Wolpin (2010) also estimate a 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, 2012) provide such estimates. In other respects, Blau and van der Klaauw (2010), Keane and Wolpin (2010), Steel *et al.* (2005) 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).

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, as discussed in section IV, is initialized 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:

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

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*; note that alternative *j*=*c* (cohabitation) is unavailable when the current stage is *g*=3 or *g*=5 (marriage). The model allows observed factors to vary over time (within and between stages) and across alternatives for each woman, 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 (so each stage is estimated independently of the others), 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 over a 28-year interval, from age 18 to

age 46; we use age 46 as the endpoint because it is the median age among women who are seen over the entire history of the NLSY79 (see section IV). For our baseline simulations, we assign each woman her actual, time-constant values of factors Z and her actual values for time-varying covariates X and Y that prevail at t=1 (age 18). 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 rely on actual covariate values observed at $t=1$, we can simulate each woman's outcomes over the entire 28-year panel regardless of whether she participates in the NLSY79 for that long.

To supplement the 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 or to have a child born by age 18 (both of which are elements of Z), 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, each woman's history from age 18 to 46 is simulated for each of 150 random draws from the estimated distribution of the parameter estimates; we use age 46 as the endpoint because it is the median age among women who are seen over the entire history of the NLSY79 (see section IV). The means and standard deviations of the simulated outcomes constitute our predicted probabilities of following various long-term paths.[2] We consider three types of paths: (*a*) early first unions, which require a transition from single (with no prior unions) to a first cohabiting union or a first marriage between ages 18 and 22; (*b*) later first unions, which require a similar transition between ages 24 and 28; and (*c*) second unions, which require a transition from separated or divorced to a cohabiting union or marriage between ages 30 and 34, conditional on terminating a first union of either type at age 30. For all three types of paths, we compute probabilities of *entering* a union (marriage, cohabitation, or either) within the given four-year window, *conditional* probabilities of maintaining the union for the long-term, and *joint* probabilities of entering the union within the four-year interval and maintaining it for the long-term.

Our primary definition of a "long-term" union is 12 or more years. Given that our simulations follow women to age 46, 12 years is the longest duration we can consider for second unions

formed by age 34. For comparison, we use eight years as an alternative definition of "long-term" for all three types of paths. We also consider 18 years (the longest duration we can consider for unions formed by age 28) for both early first unions and later first unions, and 24 years (the longest possible duration for unions formed by age 22) for early first unions.

IV. Data

A. Sample Selection

The data used to estimate the multi-stage choice models described in section III must satisfy three criteria. First, the data must come from a panel survey that follows a large sample of women from their teenage years into their late 40s or beyond. These features allow us to observe long-term unions that were formed not only at early ages, but also by individuals in their mid-30s. Second, we require detailed information on transitions between single, cohabiting, and marriage, including partner identifiers that enable us to track the beginning and end of each unique union. Third, we require a rich array of background factors along with geographic information that can be merged with external data sources to construct exogenous, "rules-based" policy variables (*e.g.*, potential welfare benefits and income tax burdens) and exogenous measures of local marriage markets.

The survey that best satisfies these criteria is the 1979 National Longitudinal Survey of Youth (NLSY79), which 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 (the last interview year for which we have data) due, in large 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 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, which is 2008 (age 44-48) for those who participate throughout the duration (thus far) of the survey. 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.[3] To identify current union status and transition dates, we do not rely solely on marital status at each interview date. Instead, we use "clean" start and end dates for each marriage provided by the survey, all available information on start and end dates for cohabitation spells, and identifiers for cohabiting partners and spouses. This information allows us to determine, for example, (*a*) that a single woman transitions into cohabitation within the next 12 months even if she reappears as single at the next interview; (*b*) 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; and (*c*) whether a woman who appears as cohabiting followed by married in successive interviews has married her cohabiting partner or formed a new union.[4]

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. 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 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, school enrollment, the presence of children, and other factors that are determined jointly with union transitions, we use time-invariant demographic, family background, and skill measures identified at (approximately) age 18. Rather than control for actual welfare benefits or income tax obligations (or predicted amounts based on *actual* income) that are endogenous to union transitions due to their dependence on economic status, we rely on cross-state and cross-year variation in welfare policy and income tax laws. Summary statistics for our covariates (for the stage 1 and stage 3 samples) are summarized in table 1.

Our union-related history variables include current spell duration in years and its square to account for duration dependence. 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 Hispanic or non-Hispanic black, with non-Hispanic/nonblack the omitted group. We also

control for whether the woman is foreign born, her 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, exogenous determinants of union transitions (*e.g.,* Bennett *et al.* 1989; Manning and Smock 1995; Phillips and Sweeney 2005).

Union formation and stability are intrinsically tied to childbearing, yet jointly estimating our five-stage model along with a model of fertility, and simulating outcomes, is beyond the scope of this study.[5] Therefore, we control for a dimension of fertility that is arguably exogenous to the union-forming process: whether a child was born to the woman prior to her 18th birthday. In principle, women could plan to marry or cohabit in response to this childbirth, but we find that none of the early child-bearers in our sample enters a union within the first year of the observation period.

To explore the effects of religious affiliation and family 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 family attitudes, 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 gender 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 attitudes 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 a dummy variable indicating whether the woman completes grade 12. The AFQT score is derived from scores on the Armed Services Vocational Aptitude Batter (ASVAB) which was administered to NLSY79 respondents in 1980, and the Rosenberg score is derived from a 10-item scale administered during the 1980 interview. Neither the ASVAB nor the Rosenberg scale were administered prior to 1980 (when our sample members were 16-20), so we necessarily deviate from our strategy of measuring covariates prior to the age 18 starting date. 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 also use data reported through 1980 to determine whether each woman completed grade 12, given that students who delay school entry, repeat a grade, or are born late in the calendar year do not necessarily complete high school by age 18. 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 rarely 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. We experimented with the inclusion of scores for the Rotter Locus of Control scale, but found that this additional measure had no affect on our estimates.

Our final set of controls is intended to capture characteristics of marriage markets and exogenous, policy-influenced 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 (Hispanic, black, or neither), 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.

Turning to policy factors, we measure the ease with which divorce can be obtained in two dimensions: whether divorce is bilateral or unilateral, and whether property settlements are

determined on a fault or no fault basis. These measures are specific to the state of residence and calendar year corresponding to the person-year observation; see the appendix for details on the construction of these and other policy variables. Unilateral divorce means the consent of both parties is not needed, and speaks to the right to divorce; no fault means marital misconduct need not be established, and speaks to the cost of divorce. While many analysts control for one dimension or the other (Friedberg 1998; Peters 1986; Wolfers 2006), we believe it is optimal to control for both dimensions, as argued by Iyvarakul *et al.* (2011). We distinguish between fault and no fault with respect to property settlement rather than divorce *per se* because all states but New York granted no fault divorce throughout our observation period. We sub-classify the "unilateral" states into those that only grant unilateral divorce following a mandatory separation versus those without a separation requirement, given that separation requirements (ranging from six months to two years for most states) substantively alter the right to divorce. As summarized in table 1, this yields a four-way classification of states that offer unilateral divorce; the omitted category consists of all states that require bilateral consent for a divorce to be issued. Alternative specifications that, for example, control for the duration of the mandatory separation and/or whether common law marriage is recognized proved not to alter our findings.

Our remaining policy variables are the maximum, monthly AFDC or TANF benefit available 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. As with the divorce variables, 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; see the appendix for details.[6] 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, 2012) 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.[7] 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 access to reading materials at age 14=0. Our next "type" is a woman who gave birth prior to her 18th birthday. Turning to the religion/attitudes variables, we next assign all women weekly church attendance, Baptist=1 (given that the Baptist indicator generally has a bigger estimated effect on union transitions than other religion categories), and a traditional attitudes score of four, which corresponds to the 90th percentile for the full sample of 2,761 women. Next, we consider a high-skill type by assigning all women the completion of 12th grade and 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 pro-marriage environment by assigning a county sex ratio corresponding to the 90th percentile for the full sample of 2,761 women, 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, we briefly assess the estimated probabilities of transitions in the "next period" to highlight the fact that conventional, short-term estimates leave us with little insight into how covariates of interest will affect probabilities of entering and forming long-term unions. Table 3 shows estimated short-term 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.

Most estimated marginal effects in table 3 have the expected sign. Single black women with no prior unions are 4.5 and 5.2 percentage points less likely to enter cohabitation and marriage, respectively, than are their observationally equivalent nonblack counterparts; if these women *do* cohabit, they are 7.6 percentage points less likely than nonblacks to marry their partner and an imprecisely estimated 1.4 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 gender 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 1.3 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 raises the likelihood of cohabiting in stage 1 by 2.7 percentage points, and subsequently raises the stage 2 likelihood of dissolving a cohabiting union by 1.2 percentage points. Laws that allow unilateral divorce and no fault property settlements are associated with an increased probability of union dissolution, increased AFDC/TANF benefits raise (lower) the predicted probability of SC (SM) transitions; and an increase in the cost of marriage via a larger income tax marriage penalty is predicted (albeit imprecisely) to deter entry into marriage among both single and cohabiting women.

While the estimates summarized in table 3 contain few surprises, it is noteworthy that only one covariate—"lived with single mother"—proves to be consistently union detracting or enhancing. That is, "lived with single mother" has a negative estimated effect on all transitions *into* unions and a positive estimated effect on all transitions *out of* unions. All other covariates are found to have inconsistent estimated effects on union formation. For example, a 10-point increase in AFQT scores is predicted to lower the probability of entering cohabiting and marital unions in stage 1, but also to lower the probability of exiting both types of unions in stages 2-3. Clearly, these estimates leave us unable to predict qualitatively whether high-skill, single women—or any "type" of woman other than those raised by a single mother—are more or less likely than their counterparts to form a union and maintain it for the long-term. Our simulations are designed to facilitate this type of long-term inference.

Even with clear-cut determinants, 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 18 percentage points (0.045x4) 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 the long-term. 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? Twelve years? Twenty-four 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 unions by age 22 and maintaining those unions for at least 12 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 by age 28, and second unions formed by age 34. We described the simulations and alternative long-term paths in section III.B, but it is worth reiterating that estimates in all three tables (*a*) are based on the same simulated outcomes from age 18 to 46 for a uniform sample of 2,761 women; (b) present predicted entry probabilities over a four-year window (ages 18-22, 24-28, or 30-34 for early first, later first, and second unions, respectively); and (*c*) use 12 or more years as the definition of long-term. We focus on 12 year-long unions because this is the longest duration that we can apply uniformly to early first, later first, and second unions. We conclude this section by demonstrating how predicted probabilities change when we use alternative definitions of "long-term" ranging from eight to 24 years (table 7).

The column A estimates in table 4 are based on simulations in which each woman is assigned her actual, first period covariate values; marital history variables (spell duration, number of prior unions, *etc.*) 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).[8] (As the note to table 4 indicates, these and all other probabilities are precisely estimated.) The conditional probabilities in rows a′-c′ indicate that among women who make a single-to-married transition by age 22, 58% are predicted to remain married to the same spouse for at least 12 years. Among women making a single-to-cohabitation transition, 35% are expected to remain with the same partner for 12+ years; this includes women who convert their cohabiting union to marriage.[9] All told, 50% of women who form a union by age 22 can be expected to maintain the union for at least 12 years. Finally, the joint probabilities in rows a″-c″ reveal that 22% (0.45·0.50) of single, 18 year-old women are expected to form a union by age 22 *and* maintain that union for at least 12 years; these women have a 16% chance of marrying without prior cohabitation and remaining with their spouse for at least 12 years, but only a 6% chance of

cohabiting (and perhaps subsequently marrying) 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 58% chance of lasting 12 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 almost six percentage points, or by 36% relative to the "entry via marriage" probability of 0.16. In short, cohabitation is less likely than marriage to lead to a long-term union, yet is a sufficiently common form of entry that it substantially increases the overall chance of experiencing a long-term union.

Turning to the estimates for different "types" of women in columns B-F 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-c), to maintain those unions for 12 or more years (rows a'-c') or to enter *and maintain* those unions (rows a″-c″).[10] For example, we predict that disadvantaged blacks (column B') have a 22.9% chance of forming a union by age 22 and only a 9.8% chance of forming the union and maintaining it for the long-term; both predictions are roughly half as large as the corresponding estimates in column A. More noteworthy is the finding that cohabitation plays a smaller role in long-term union formation among blacks than among other women. Using disadvantaged blacks for illustration, "entry by cohabitation" raises the predicted probability of union entry by 49% (the percent increase from 0.154 in row a to 0.229 in row c), lowers the predicted conditional probability by 40% (0.224 in row b' relative to 0.529 in row a') and raises the predicted joint probability by 21% (the percent increase from 0.081 in row a″ to 0.098 in row c″); the comparable numbers in column A are 59%, 60% and 36%. Women who give birth as teenagers (column C) look similar to blacks in that their predicted entry, conditional, and joint probabilities are also lower than what is seen in column A. The most striking finding for teen mothers is that their conditional probability of maintaining unions formed via marriage (row a') is only 0.447, which is dramatically lower than what we see for blacks or any other type. These women are six percentage points less likely than their column A counterparts to experience a long-term union (0.157 versus 0.221) largely because they are *far* less likely to maintain their marriages.

At the other extreme are high skill women (column E), for whom the conditional probabilities of remaining with one's partner are the highest of any type: 0.715 and 0.438 for unions entered via marriage and cohabitation, respectively. Moreover, cohabitation contributes more to long-term unions for this group than for any other. The predicted probability of an early single-to-cohabitation transition is 0.16, which raises the predicted likelihood of an early first union (relative to the "entry by marriage" option in row a) by 68%. Similarly, the predicted joint probability of 0.070 (row b") accounts for a 42% increase in the predicted probability of experiencing a long-term early union.

Perhaps the most surprising finding in table 4 is that the remaining types—women who attend church regularly and report traditional family attitudes (column D), and women in a pro-marriage environment (column F)—are only four percentage points more likely than their column A counterparts to experience a long-term union entered via marriage (row a"), and only three percentage points more likely to experience a long-term union of any kind (row c"). Relative to the column A sample, both types show a slight tendency to substitute from cohabitation towards marriage when entering early first unions (rows a-b) and a slight increase in the predicted probability of maintaining unions entered via marriage (row a'). However, neither shift is sufficiently pronounced to increase substantially the likelihood of experiencing a long-term union entered via marriage, let alone to increase the "bottom line" likelihood of experiencing a long-term union of any type.[11]

To summarize, the findings in table 4 corroborate well-established evidence that blacks have lower probabilities of forming unions than nonblacks, and that highly-skill women tend to delay marriage. However, the "new" findings in table 4 are two-fold. First, while blacks and teen mothers are less likely than others to experience (that is, to form and maintain) a long-term first union, high-skill women, religious women, and women in pro-marriage environments are no more likely than others to do so. Second, looking across columns we find that the predicted probabilities in row c" are 2-7 percentage points (20-40%) higher than the estimates in row a". This indicates that for all "types" of women, 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 2-7 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 all types. For example, in column A of table 5 the predicted probabilities of marrying or forming any union are 0.23 and 0.41, 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 12 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 (column F) have a 25% chance of experiencing an early or 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 3-9 percentage points (33-60%) 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 12 years (row b') when formed late versus early.

Next, 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 12 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 than 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 9% chance of remarrying and remaining remarried for at least 12 years; the comparable estimates are 28% and 16% in table 4, and 23% and 15% 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 32% chance of cohabiting within four years (row b), a 48% chance of remaining with that partner for 12 years (row b') and a 15% chance of entering and maintaining a union formed via cohabitation (row b"); the comparable estimates in table 5 are 16%, 41% and 7%.

The bottom row of table 6 reveals 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 0.11 for poor blacks to 0.22 for teen mothers to 0.25-0.29 for every other type, while the comparable estimates in table 4 range from 0.11 to 0.25. 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 12+ years (row c") is two to three times higher than their predicted probability of entering and maintaining a marriage. Moreover, cohabitation makes the *largest* relative contribution to the probability of experiencing a long-term union blacks and teen mothers, which are the groups with the *lowest* chance of experiencing a long-term union.

The last step of our investigation is to determine whether our findings are sensitive to the definition of "long-term." In table 7 we report conditional and joint probabilities of maintaining each type of union (early first, later first, and second) for durations ranging from eight to 24 years; entry probabilities are excluded from table 7 because they are invariant to the definition of long-term. The simulations used for table 7 are based on actual covariate values, so these estimates should be compared to the column A estimates in tables 4-6.

As expected, the estimated conditional probability of maintaining each type of union declines as we increase union duration from 8 to 12 to 18 to 24 years. For example, the predicted

probability of maintaining an early first union entered via marriage falls from 0.716 to 0.583 to 0.417 to 0.292 across this range of union durations. The predicted joint probabilities necessarily decrease as well, given that they are the product of the conditional and entry probabilities, the latter of which are invariant to duration. Despite these quantitative differences, table 7 reveals remarkable stability across union duration in the qualitative patterns highlighted earlier. In particular, the relative contribution of cohabitation to long-term unions—measured by the percent by which row c" exceeds row a"—is 35-39% for early first unions, 51-58% for later first unions, and 184-187% for second unions. These percentages are slightly higher when we consider unions lasting 18 or 24 years than when we consider 8- or 12-year durations. Nonetheless, our finding that "entry by cohabitation" makes an increasingly important contribution to long-term union formation over the life-cycle proves not to depend on our definition of long-term.

VI. Conclusions

Most analysts who study 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 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 factors that cause individuals to enter a union and maintain it for the long-term. Our contribution is to estimate a series of stage-specific transition models, use the estimates to simulate women's sequences of union-related outcomes from age 18 to 46, and then use these simulated outcomes to predict probabilities of experiencing 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 12 years (or, alternatively, 8, 18, or 24 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 a focus on 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 high-skill women are less likely to dissolve unions, and that cohabiting unions are less likely than marriages to endure. Even here, however, we provide new

insights by *quantifying* the odds of experiencing a long-term union. 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 25% chance of entering a union (via cohabitation or marriage) within four years, but only a 12% chance of forming a union and maintaining it for at least 12 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 16% chance of marrying by age 22 *and* remaining with her husband for at least 12 years, and a 22% chance of forming any union (via cohabitation or marriage) and maintaining it for at least 12 years. That is, the "entry by cohabitation" option raises the predicted probability of entering and maintaining an early union by six percentage points, or 36%. The contribution of "entry by cohabitation" increases to 52% for women who form a first union at ages 24-28, and to an astounding 187% for women who form a second union at ages 30-34. We find similar patterns when we substitute both longer and shorter durations as alternative definitions of "long-term." 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, as are women who give birth as teenagers. However, none of the policy factors that we considered (divorce law, expected income tax obligations, expected welfare 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.

[1]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.

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

[3]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 in some years is only known at the time of each interview, so we assume residential changes take place half way between successive interview dates.

[4]A potential shortcoming of the NLSY79 is that from 1979 to 1989 (when respondents in our sample were ages 25-29), only cohabitations that spanned the interview date were identified. This caused the shortest cohabitation spells to be under-sampled, and could affect our inferences if "true" first spells are systematically and substantially shorter than first spells that span interview dates. To assess this problem, we use the 1997 National Longitudinal Survey of Youth—which tracks all cohabitation spells in an event history format—to identify cohabitation spells from age 18 to 25-29; *i.e.,* we follow respondents, all of whom were born in 1980-4, from their 18th birthday to the 2009 interview date. Of the 8,426 cohabitation spells observed for 8,115 respondents, only 14% would be "lost" if we required them to be in-progress at an interview date. Among respondents with more than one spell, only 4.4% have a "lost" spell that is shorter than the second spell. Extrapolating to the NLSY79, it appears unlikely that we significantly over-predict the duration of cohabitation spells simply because a small number of spells that do no span interview are unobserved.

[5]Brien *et al*. (1999), Lillard and Waite (1993), and Steele *et al.* (2005) are examples of studies that jointly estimate union formation/dissolution and childbearing, although they consider fewer stages than we do and do not simulate outcomes across multiple stages. Simply adding covariates for childbearing to our model would be inappropriate because these measures are endogenous to union-related decisions.

[6]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 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.

[7]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.

[8]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*.

[9]The majority of women who remain with their cohabiting partner for at least 12 years convert the cohabitation to marriage. Fewer than 5% of the simulated paths underlying the row b' prediction reveal 12 consecutive years of cohabitation.

[10]The sole exception is that black women lag the column A sample by only 1.6 percentage points in their conditional probability of maintaining a union formed via marriage (row a').

[11]One concern is that we estimate trivial effects of "pro-marriage" factors because divorce laws and income tax bonuses only affect high income women, while AFDC/TANF payments only affect low income women. To explore this, we allowed the coefficients for each variable in this group to differ for high and low income women defined on the basis of race, mother's schooling, own schooling, and AFQT scores. We experimented with multiple definitions of high and low (expected) income, and we allowed each policy variable to have varying effects in isolation and in combination with other interactions. For each alternative specification, we failed to reject the null hypothesis that the coefficients are equal for alternative income groups. Our findings were also unchanged when we experimented with alternative measures of divorce laws and income tax marriage penalties/bonuses.

[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.

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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, no separation, no fault property settlement" means the woman can obtain a unilateral divorce without a mandatory separation period, *and* she can obtain a property settlement (as well as the divorce) without having to establish fault (marital misconduct). "Unilateral divorce, no separation, fault for property settles" 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, separation, no fault property settlement" and "unilateral, separation, fault for property settlement" are defined similarly, but in both categories unilateral, no fault divorce is granted only after a mandatory separation requirement is met. The omitted group identifies state-year observations where divorce is granted only with bilateral consent. This five-way taxonomy is ordered in the sense that bilateral divorce (the omitted group) limits the right to divorce relative to unilateral divorce with a separation requirement which, in turn, is more restrictive than unilateral divorce without a separation requirement. Similarly, the need to establish fault raises the cost of divorce relative to no fault property settlements. Individual women may encounter exceptions to these general restrictions 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 and 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 1979-2008 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 separated/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

separated/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	
Fertility						
1 if child born by age 18	.04		.06		.03	
Religion and attitudes						
1 if raised 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, 1979	.21		.23		.22	
1+ times/week, 1979	.45		.45		.47	
Traditional attitudes score, 1979	1.83	1.66	1.83		1.80	1.65
Skill levels						
AFQT percentile score, 1980	42.27	27.83	41.02	28.98	46.42	27.63
Rosenberg self esteem score (10-34), 1980	18.17	4.01	18.09	4.06	17.96	3.93
1 if highest grade completed ≥ 12, age 18	.82		.81		.85	
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 and						
no separation, no fault property settlement	.22		.20		.24	
no separation, fault for property settlement	.29		.30		.29	
separation, no fault property settlement	.06		.05		.06	
separation, fault for property settlement	.24		.23		.23	
Max. monthly AFDC/TANF benefit, $100s	6.12	2.53	4.92	2.88	3.28	3.09
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.

†Time-varying variables (includes all environmental factors).

Table 2: Multinomial Logit Estimates for Stages 1-5

	Stage 1: 1st single spell				Stage 2: All cohab. spells			
	S to C		S to M		C to M		C to S	
Covariate	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Constant	-5.071	1.547	-2.015	1.355	.850	1.980	-2.191	1.744
Spell duration	.095	.022	.049	.018	-.107	.037	-.273	.031
Spell duration squared/10	-.061	.013	-.063	.011	-.030	.041	.072	.023
Age began spell					-.026	.012	-.068	.010
Number of prior cohabitation spells					-.197	.125	-.240	.111
Number of prior marriages					.000	.086	-.037	.082
Family background								
1 if black	-1.349	.184	-1.058	.144	-.715	.228	-.014	.186
1 if Hispanic	-.763	.186	-.227	.151	-.388	.225	-.052	.193
1 if foreign born	-.240	.158	.038	.126	-.128	.215	-.099	.180
Mother's highest grade completed	.005	.014	-.030	.011	.038	.019	.048	.016
1 if lived with single mother, age 14	-.333	.150	-.168	.142	-.336	.176	.168	.140
mother and stepfather, age 14	-.075	.178	.069	.170	.009	.197	.175	.167
mother and father, age 14	-.615	.140	.024	.131	-.003	.153	-.074	.137
1 if access to reading materials, age 14	-.032	.130	.010	.111	-.116	.186	-.106	.120
Fertility								
1 if child born by age 18	-.241	.190	-.150	.174	-.217	.183	-.072	.131
Religion and attitudes								
1 if raised Baptist	-.271	.192	.273	.186	.245	.216	-.046	.203
Catholic	-.284	.192	-.008	.186	-.028	.209	-.148	.197
other Christian	-.334	.190	.112	.180	.187	.210	-.046	.198
other religion	-.354	.208	.256	.193	.136	.243	.172	.220
1 if attends church 1+ times/month	.038	.094	.049	.085	.080	.123	-.037	.109
1+ times/week	-.278	.087	.116	.071	.241	.109	.003	.091
Traditional attitudes score	-.043	.024	.005	.019	-.110	.031	-.003	.026
Skill levels								
AFQT score	-.005	.002	-.006	.001	.005	.002	-.002	.002
Self esteem score	.008	.009	-.022	.008	-.027	.011	-.002	.010
1 if highest grade completed ≥12, age 18	.050	.091	.128	.075	-.090	.115	-.116	.114
Environmental factors								
Percent same race in county	-.007	.002	-.001	.002	-.002	.002	-.002	.002
Percent men in county	.075	.030	.019	.027	-.016	.035	.076	.035
Population density in county	-.013	.008	.009	.007	.002	.008	.006	.006
1 if unilateral, no separation, no fault prop.	-.100	.138	.039	.110	.325	.161	.057	.141
if unilateral, no separation, fault property	.024	.113	.005	.091	.098	.146	.116	.122
if unilateral, separation, no fault property	.112	.163	.210	.131	.306	.195	.353	.178
if unilateral, separation, fault property	-.034	.119	.146	.091	-.149	.153	-.065	.131
Maximum monthly AFDC/TANF benefit	.059	.020	-.062	.017	-.034	.026	-.010	.022
State income tax marriage penalty	.035	.029	-.024	.027	-.011	.041	.011	.036
Log likelihood	-8,552.70				-3,894.27			
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	-.364	1.659	-4.356	1.806	-4.897	2.563	3.785	2.620
Spell duration	.029	.021	-.151	.030	-.143	.036	-.005	.039
Spell duration squared/10	-.037	.010	.014	.020	-.005	.021	-.028	.022
Age began spell	-.053	.012	-.066	.010	-.106	.014	-.168	.018
Number of prior cohabitation spells	.297	.091	.080	.046	-.197	.075	.285	.087
Number of prior marriages			-.128	.068	-.070	.092	.272	.209
1 if cohabited with spouse before marriage	-.267	.127					-.509	.155
Family background								
1 if black	-.055	.159	-.561	.177	-.843	.238	.533	.299
1 if Hispanic	-.158	.171	.159	.182	-.539	.250	-.182	.322
1 if foreign born	-.197	.163	-.401	.217	-.078	.216	-.554	.354
Mother's highest grade completed	.022	.015	.009	.017	.007	.021	-.016	.027
1 if lived with single mother, age 14	.335	.156	-.015	.156	-.055	.213	-.130	.261
mother and stepfather, age 14	.051	.181	.103	.180	.268	.246	-.265	.269
mother and father, age 14	-.019	.148	.073	.145	.188	.196	-.140	.234
1 if access to reading materials, age 14	.141	.120	.051	.132	.412	.204	-.551	.183
Fertility								
1 if child born by age 18	.448	.181	.108	.129	-.092	.179	-.427	.231
Religion and attitudes								
1 if raised Baptist	.021	.195	.054	.187	-.045	.338	-.129	.346
Catholic	-.111	.199	-.052	.192	-.025	.336	.031	.355
other Christian	.038	.200	.053	.187	-.102	.342	-.360	.340
other religion	.052	.207	-.107	.202	-.002	.351	-.228	.364
1 if attends church 1+ times/month	-.174	.098	.032	.106	.160	.147	-.168	.179
1+ times/week	-.194	.082	-.139	.091	.271	.126	-.140	.144
Traditional attitudes score	-.033	.027	.037	.026	.044	.034	.014	.036
Skill levels								
AFQT score	-.011	.002	.001	.002	.001	.003	.002	.003
Self esteem score	-.002	.010	.001	.010	-.006	.014	-.025	.017
1 if highest grade completed ≥ 12	.006	.092	-.235	.103	-.062	.144	-.330	.155
Environmental factors								
Percent same race in county	-.001	.002	.003	.002	-.006	.003	-.005	.004
Percent men in county	-.037	.033	.090	.035	.132	.050	.006	.050
Population density in county	-.001	.007	.005	.007	.019	.014	-.005	.015
1 if unilateral, no separation, no fault prop.	.350	.136	-.135	.145	-.388	.186	-.221	.230
if unilateral, no separation, fault property	.314	.115	-.103	.127	-.259	.158	.039	.206
if unilateral, separation, no fault property	.086	.193	.014	.191	-.745	.294	.450	.314
if unilateral, separation, fault property	.147	.120	-.051	.137	-.185	.177	.208	.214
Maximum monthly AFDC/TANF benefit	-.043	.022	.019	.022	-.041	.032	-.047	.039
State income tax marriage penalty	-.005	.032	-.032	.041	-.094	.068	-.051	.065
Log likelihood	-3,738.33		-4,516.28				-1,396.64	
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

	Stage 1		Stage 2		Stage 3
Covariate	SC	SM	CM	CS	MS
Black (0 to 1)	-.045	-.052	-.076	.014	-.002
	(4.74)	(7.19)	(3.71)	(0.51)	(0.35)
1-year increase in mother's highest grade completed	.000	-.002	.003	.006	.001
	(0.50)	(2.72)	(1.46)	(2.62)	(1.49)
Lived with single mother, age 14 (0 to 1)	-.011	-.008	-.042	.034	.011
	(2.34)	(1.17)	(2.34)	(1.57)	(1.92)
Access to reading materials, age 14 (0 to 1)	-.001	.001	-.012	-.013	.004
	(0.25)	(0.10)	(0.51)	(0.75)	(1.23)
Child born by age 18 (0 to 1)	-.008	-.007	-.023	-.005	.016
	(1.35)	(0.86)	(1.22)	(0.30)	(2.05)
Baptist (0 to 1)	-.010	.016	.032	-.013	.001
	(1.57)	(1.46)	(1.16)	(0.46)	(0.21)
Attend church 1+ times/week (0 to 1)	-.010	.007	.030	-.006	-.006
	(3.34)	(1.79)	(2.25)	(0.46)	(2.37)
1-point increase in traditional attitudes	-.002	.000	-.013	.002	-.001
	(1.88)	(0.36)	(3.71)	(0.62)	(1.74)
High school graduate (0 to 1)	.002	.007	-.008	-.014	.000
	(0.45)	(1.62)	(0.61)	(0.93)	(0.60)
10-point increase in AFQT score	-.002	-.003	.007	-.004	-.003
	(2.68)	(4.03)	(2.99)	(1.59)	(6.81)
1-point increase in self esteem score	.000	-.001	-.003	.001	-.000
	(0.94)	(2.97)	(2.38)	(0.22)	(0.63)
10-point increase in percent men in county	.027	.009	-.039	.012	-.011
	(2.47)	(0.74)	(0.84)	(2.34)	(1.13)
Unilateral divorce, no separation, no fault property settlement (0 to 1)	-.004	.002	-.040	.000	.011
	(0.21)	(0.39)	(1.91)	(0.44)	(2.36)
100-dollar increase in maximum AFDC/TANF benefit	.001	-.003	-.004	-.005	-.013
	(2.99)	(3.86)	(1.31)	(0.17)	(1.97)
100-dollar increase in state income tax marriage penalty	.001	-.001	-.002	.019	-.015
	(1.27)	(0.92)	(0.32)	(0.38)	(0.17)
Unconditional transition probability	.038	.058	.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.

Table 4: Predicted Probabilities of Entering and Maintaining "Early" First Unions for 12 or More Years
(for 18 year-old women with no prior unions)

Simulated pattern	Covariates used for simulations†						
	Actual values (A)	Black (B)	Poor black (B')	Teen birth (C)	Trad'l values (D)	High skill (E)	Pro-mar. setting (F)
Entry into early first union							
a. Marry by age 22; no prior cohabitation	.280	.166	.154	.258	.336	.235	.328
b. Cohabit by age 22; no prior marriage	.165	.088	.075	.141	.127	.160	.150
c. Marry or cohabit by age 22	.446	.254	.229	.400	.463	.394	.478
Conditional probability							
a'. Stay married for 12+ years, given a.	.583	.567	.529	.447	.613	.715	.617
b'. Stay with partner for 12+ years, given b.	.351	.277	.224	.289	.360	.438	.323
c'. Stay with partner for 12+ years, given c.	.497	.467	.429	.391	.543	.603	.525
Joint probability							
a''. Stay married for 12+ years, and a.	.163	.094	.081	.116	.206	.168	.203
b''. Stay with partner for 12+ years, and b.	.058	.024	.017	.041	.046	.070	.048
c''. Stay with partner for 12+ years, and c.	.221	.119	.098	.157	.252	.238	.251

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 46 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 0.018, and for all means are statistically distinguishable from zero at a significance level of 0.05.

†All women are assigned their actual covariate values in column A. In subsequent columns, all women are assigned: black (B); black plus nonforeign born, mother completed grade 9, lived with single mother, no reading access (B'); child born by age 18 (C); regular church attendance, Baptist, traditional attitudes score at the 90th percentile (D); high school graduate plus AFQT and self esteem scores at the 90th percentile (E); state without unilateral divorce plus percent men at the 90th percentile, AFDC/TANF and income tax marriage penalty at the 10th percentile (F). See text for details.

Table 5: Predicted Probabilities of Entering and Maintaining "Later" First Unions for 12 or More Years
(for 24 year-old women with no prior unions)

Simulated pattern	Covariates used for simulations†						
	Actual values (A)	Black (B)	Poor black (B')	Teen birth (C)	Trad'l values (D)	High skill (E)	Pro-mar. setting (F)
Entry into later first union							
a. Marry by age 28; no prior cohabitation	.228	.141	.130	.208	.269	.188	.265
b. Cohabit by age 28; no prior marriage	.177	.100	.084	.151	.138	.168	.162
c. Marry or cohabit by age 28	.405	.241	.214	.359	.406	.356	.426
Conditional probability							
a'. Stay married for 12+ years, given a.	.665	.658	.624	.545	.689	.777	.696
b'. Stay with partner for 12+ years, given b.	.441	.386	.327	.394	.444	.526	.406
c'. Stay with partner for 12+ years, given c.	.567	.546	.507	.482	.606	.658	.586
Joint probability							
a''. Stay married for 12+ years, and a.	.151	.093	.081	.114	.185	.146	.184
b''. Stay with partner for 12+ years, and b.	.078	.039	.028	.060	.061	.088	.066
c''. Stay with partner for 12+ years, and c.	.229	.132	.109	.173	.246	.234	.250

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.

†See table 4.

Table 6: Predicted Probabilities of Entering and Maintaining Second Unions for 12 or More Years
(for women who end their first union at age 30)

Simulated pattern	Covariates used for simulations†						
	Actual values (A)	Black (B)	Poor black (B')	Teen birth (C)	Trad'l values (D)	High skill (E)	Pro-mar. setting (F)
Entry into second union							
a. Marry by age 34; prior union	.131	.095	.065	.122	.166	.135	.221
b. Cohabit by age 34; prior union	.306	.228	.218	.343	.304	.284	.319
c. Marry or cohabit by age 34	.438	.323	.283	.464	.471	.419	.540
Conditional probability							
a'. Stay married for 12+ years, given a.	.659	.534	.448	.540	.658	.733	.618
b'. Stay with partner for 12+ years, given b.	.525	.440	.365	.449	.539	.578	.480
c'. Stay with partner for 12+ years, given c.	.565	.468	.384	.474	.581	.628	.536
Joint probability							
a''. Stay married for 12+ years, and a.	.086	.050	.029	.066	.109	.099	.137
b''. Stay with partner for 12+ years, and b.	.161	.100	.080	.154	.164	.164	.153
c''. Stay with partner for 12+ years, and c.	.247	.151	.109	.220	.274	.263	.289

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.

†See table 4.

Table 7: Predicted Probabilities of Maintaining Early First, Later First, and Second Unions for Alternative Durations (X)

	Early 1st unions			Late 1st unions		2nd unions
Simulated Pattern	X=8	X=18	X=24	X=8	X=18	X=8
Conditional probability						
a'. Stay married for X+ years, given a†	.716	.417	.292	.778	.515	.747
b'. Stay with partner for X+ years, given b†	.423	.263	.194	.511	.383	.588
c'. Stay with partner for X+ years, given c†	.608	.360	.256	.661	.446	.636
Joint probability						
a''. Stay married for X+ years, and a†	.201	.117	.082	.177	.117	.098
b''. Stay with partner for X+ years, and b†	.070	.043	.032	.090	.068	.180
c''. Stay with partner for X+ years, and c†	.271	.160	.114	.267	.180	.278

Note: All women are assigned their actual covariate values. Estimates should be compared to the column A estimates in tables 4-6. Outcomes are simulated to age 46, so the maximum duration (X) that can be used is 24 years for first unions entered by 22, 18 years for second unions entered by age 28, and 12 years for second unions entered by age 34.

†X refers to the duration under consideration (8, 18, or 24 years); a, b and c refer to the entry probabilities shown in column A of tables 4-6, which are invariant to the chosen duration (X).