Economic Incentives and Family Formation

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March 2005; revised February 2009

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This research was funded by a grant to Light from the National Science Foundation (grant SES-0415427) and a grant to Omori from the Japan Society for the Promotion of Science (Grant-in-Aid for Scientific Research (B)16530169); we thank both agencies for their generous support. We also thank Taehyun Ahn and Ranajoy Ray-Chaudhuri for excellent research assistance, Daniel Feenberg and the National Bureau of Economic Research for the use of TAXSIM to calculate income taxes, and Robert Moffitt for data on state welfare benefits.

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Abstract: This study identifies the effects of economic factors that can be directly manipulated by public policy on women's union-forming decisions. We jointly model transitions made by never-married women to cohabitation or marriage, cohabiting women to marriage or separation, and married women to divorce. We control for expected income tax burdens, maximum allowed state AFDC or TANF benefits, average state Medicaid expenditures, and parameters of state laws governing divorce and the division of property, along with a wide array of family background, personal, and environmental characteristics. We compare the estimated effects of alternative policy interventions to each other, and to the estimated effects of nonpolicy factors. In addition to focusing on the predicted effect of each factor on each individual transition (single to married, *etc*.), we compute their effects on the predicted probability of long-term marriage and long-term unions of any type (marriage or cohabitation). We find that each policy variable except the income tax "marriage penalty" is a potentially important determinant of long-term union formation. However, several factors that are outside the control of policy makers, such as religion, childhood household composition and the presence of children also have very large, potentially offsetting effects.

I. Introduction

Family formation in the U.S. changed dramatically over the last three decades: marriage rates declined, divorce rates rose sharply, and cohabitation among unmarried couples became increasingly common. These patterns have been carefully documented (Bumpass and Lu 2000; Bumpass, Sweet and Cherlin 1991; Cherlin 1992), but researchers continue to seek a better understanding of the behavioral mechanisms by which individuals choose their marital status. To what extent does family background affect these choices? To what extent are they financial decisions? What prompts individuals to revise past decisions and dissolve relationships or marry their cohabiting partners? In the face of rapid social change, we must continually refine our answers to these questions—particularly if incentives to choose one marital status over another are to be used as policy tools for improving the welfare of families. From elimination of the income tax "marriage penalty" to no-fault divorce laws to welfare reform, policy initiatives that address the link between marriage and well-being require a solid understanding of the forces that drive union-forming decisions.

In the current study, we use data from the 1979 National Longitudinal Survey of Youth (NLSY79) to identify the effects of a broad set of economic and non-economic factors on women's decisions to cohabit, marry, and divorce. Studies of union formation often consider such economic factors as schooling attainment, employment status, and wages, which are viewed as proxies for the gains resulting from the division of labor and income pooling within households (Moffitt 2000; Oppenheimer 2000; Xie *et al.* 2003). We focus instead on economic costs and benefits *conferred by law* as a function of marital status. In particular, we control for expected state income tax burdens, maximum allowed state AFDC or TANF benefits, average state Medicaid expenditures, and parameters of state laws governing divorce and the division of property. A primary goal of our analysis is to learn whether economic incentives that can be directly manipulated by public policy have important effects on women's union-forming decisions.

Elements of our analytical strategy have been seen in earlier research, but we take an unusually comprehensive approach by incorporating the following features. First, we control for a number of economic factors simultaneously, along with family background factors (*e.g.*, religion, living arrangements at age 14), individual factors (*e.g.,* ability test scores, race/ethnicity) and environmental factors (county unemployment rates and racial composition).

Other researchers have looked at the effects of tax policy (Alm and Whittington 1999; Whittington and Alm 1997) *or* welfare policy (Bitler *et al.* 2004; Ellwood and Bane 1985; Grogger and Bronars 2001) *or* divorce laws (Friedberg 1998; Nakonezny, Shull and Rodgers 1995; Wolfers 2006) on marriage-related transitions. We learn not only whether each of these factors "matters," but how their estimated effects on women's marital transitions compare to each other *and* to the estimated effects of factors that policy-makers cannot directly manipulate. We also allow the estimated effects of select factors to vary across racial groups.

Second, we focus on economic factors that are exogenous to marriage-related decisions. For example, we control for *expected* state income tax burdens rather than *actual* tax burdens, and for the generosity of states' welfare benefit rather than individuals' past or current welfare recipiency. By eliminating endogenous variation in employment, earnings, and other factors that can depend on marital status, we are able to isolate the "true" effects of economic incentives on union formation. Identification is aided by our reliance on state-level policies which, unlike federal laws, generate within-year variation in the data.

Third, we acknowledge the sequential nature of union-forming decisions by estimating a model in which women decide on an annual basis whether to be single, cohabit, or marry. While most researchers examine the transitions undertaken by individuals who are currently single (*e.g.,* Lundberg and Rose 2003; Xie *et al.* 2003) *or* cohabiting (*e.g.,* Lichter, Qian and Mellot 2006; Smock and Manning 1997) *or* married (*e.g.,* Whittington and Alm 1997; Wolfers 2006), we estimate all three stages jointly and allow the unobserved factors influencing these decisions to be correlated across alternatives and over time. This approach enables us to assess the effect of *each* observed factor on *each* type of transition. For example, we consider raising the income tax "marriage penalty" from the sample mean to the 90th percentile (holding other factors constant), and compute the effect of this intervention on the predicted probabilities that a single woman cohabits or marries, a cohabiting woman marries or becomes single, and a married woman divorces.

Because we consider the sequential process of forming and dissolving unions, we are also able to compute the effect of each observed factor on the predicted probability of getting married and *staying* married or, more generally, forming and maintaining *any* union. Our examination of year-to-year transitions may reveal, for example, that increased AFDC or TANF benefits are expected to *decrease* the likelihood of single-to-marriage transitions, but *increase* the likelihood of single-to-cohabiting and cohabiting-to-marriage transitions. Does this intervention lead to more marriage and, in particular, more long-term marriage? To answer these questions, we use our underlying estimates to compute the predicted probabilities of two alternative paths: marry by age 25 and remain married for at least nine years and, alternatively, form any union (marry or cohabit) by age 25 and maintain that union for at least nine years. By moving the focus from year-to-year transitions to these hypothetical paths, we are able to determine which policy (and non-policy) interventions are effective in promoting the formation of long-term unions.

II. BACKGROUND

Like many contributors to the modern, socioeconomic study of family formation, we begin with the premise that people engage in optimizing behavior. As individuals choose among alternative marital states (single, cohabiting, married, divorced), they are assumed to select the best option. In this study, we consider the role of several factors—welfare benefits, income tax obligations, and divorce laws—that can be manipulated by public policy to affect the economic costs or benefits associated with each option. We also consider numerous factors—religion, race and ethnicity, ability test scores, local labor market conditions, *etc.*—that affect the value of each option by altering preferences or the potential gains to union formation. In this section, we explain how policy-related economic factors are expected to influence union-forming decisions, and we briefly summarize existing empirical evidence on the importance of these key factors.

A. Welfare Benefits

The now-defunct Aid to Families with Dependent Children (AFDC) program provided cash benefits to low-income, single mothers in the U.S. and, in some instances, to two-parent families where one parent was not biologically related to the children. Other two-parent families received no benefits unless a parent's unemployment status made them eligible for AFDC-UP. Thus, the program imposed a "marriage tax" insofar as nonmarital fertility would typically increase cash benefits and marital fertility would not. For many years, this marriage penalty was reinforced by the fact that Medicaid eligibility was tied directly to AFDC eligibility. However, reforms introduced in the late 1980s and 1990s increased Medicaid income eligibility beyond the limit set by AFDC, while also eliminating the requirement that children live with a single or cohabiting parent to be eligible for Medicaid. The Personal Responsibility and Work Opportunity Reconciliation Act of 1996 replaced AFDC with Temporary Assistance for Needy Families (TANF). Under this program, states increased eligibility for two-parent families, reduced the generosity of benefits (in part by imposing time limits), and operated welfare-to-work programs; each component of TANF is predicted to make marriage a more attractive option than it was under AFDC programs, although welfare-to-work programs could also discourage marriage by making women more economically independent (Oppenheimer 2000).[1]

Regardless of which regime we consider, welfare programs in the U.S. provide cash benefits that are tied to varying degrees to the recipient's marital status. Moreover, benefit levels—and, therefore, the gaps in expected benefits between married and unmarried women—vary dramatically across states. Table 1 shows the maximum AFDC payment available for a family of four in 1995 in several states. The least generous states (Mississippi, Alabama and Tennessee) pay between $144 and $226 per month, while the median state (Maryland) pays $450 and the most generous state (Alaska) pays $1,025. Similar cross-state variation is seen for average Medicaid expenditures for a family of four. This cross-state variation in the "cost" of marriage, as well as the additional variation caused by policy changes over time, is one avenue by which we can assess the effect of economic factors on union formation.

Numerous researchers have already exploited cross-state and cross-year variation in benefit levels to assess the empirical effects of AFDC, TANF, and Medicaid on union formation. As expected, studies of the AFDC-marriage link generally find that increased benefit levels decrease the likelihood that single women marry (Blackburn 2000; Grogger and Bronars 2001; Hoynes 1997; Lichter, McLaughlin and Ribar 2002; Winkler 1994) and increase the likelihood of divorce (Ellwood and Bane 1984; Hoffman and Duncan 1995). Moffit, Reville and Winkler (1998) demonstrate that a surprisingly high percent of AFDC recipients cohabit, presumably because they are not penalized for monetary contributions made by unmarried partners. Bitler *et al.* (2004) find that the transition from AFDC to TANF caused fewer marriages but also less divorce, while Kaestner and Kaushal (2005) find little effect of TANF on marriage rates. These conflicting results are consistent with the fact that TANF programs simultaneously encourage marriage by increasing eligibility for married women and discourage marriage by promoting female employment. Yelowitz (1998) observes that the theoretical effect of Medicaid expansion on marriage rates is similarly ambiguous, but provides empirical evidence that the net effect of

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[1]See Bitler *et al*. (2004), Blank (2002), Hoynes (1997), Moffitt (1990, 1992) and Yelowitz (1998) for additional details on each program's characteristics and predicted effects on union formation.

increased eligibility on entry into marriage is positive. While a great deal has been learned about the effects of welfare benefits on union formation, Medicaid has received much less attention than AFDC/TANF and cohabitors' transitions have been largely overlooked in this literature.[2]

B. Income Taxes

A husband and wife in the U.S. who earn similar levels of taxable income often face a higher federal tax burden than they would face if unmarried. A marriage penalty arises when the standard deduction for a married couple is less than twice the standard deduction for a single filer (*e.g.,* \$6,550 versus \$3,900 in 1995). Similarly, a married couple in which one partner earns all or most of the taxable income generally receives a marriage bonus by using the larger standard deduction. The Economic Growth and Tax Relief Reconciliation Act of 2001 alleviated the "marriage penalty" for many low-income couples and is expected to eliminate the penalty for all couples by 2010. However, state income tax burdens will continue to vary with marital status because many states impose a tax penalty or bonus due to differential standard deductions. Moreover, because states vary dramatically in their income tax rates and allowable deductions, the difference between the tax owed by a given couple if single (or cohabiting) and the tax owed if married can vary across states even in the absence of a marriage penalty or bonus.[3]

To illustrate how tax burdens vary with marital status, in table 2 we present the taxes owed by a hypothetical couple in 1995, assuming one partner earns \$35,000 and the other earns \$20,000. We also assume this couple has no other taxable income, no itemized deductions, and no dependents. Under these assumptions, their joint federal income tax liability is \$8,077 if they are single (regardless of whether they cohabit), with each partner paying tax on the portion of his/her income that exceeds the standard deduction of \$3,900. The same couple pays \$8,503 if they are married and file jointly; in this scenario, a tax is imposed on their total income net of a standard deduction of only \$6,550. If they live in Minnesota, they owe an additional \$3,152 in state taxes if they are single and \$3,419 if they are married; in other words, the "marriage penalty" increases \$267 above and beyond the \$426 penalty imposed by federal law. Although Minnesota tax law does not allow personal exemptions and imposes an 8% marginal tax rate for

[2]Lichter *et al.* (2006), and Manning and Smock (1995) identify the effects of *actual* welfare receipt, rather than exogenous potential benefits, on cohabitors' transitions.

[3]Additional details on the relevant tax laws and theoretical effects of taxes on marriage decisions can be found in Alm, Dickert-Conlin and Whittington (1999), Chade and Ventura (2005), and Feenberg and Rosen (1995).

all incomes considered in this example, the additional penalty arises because federal taxable income (adjusted for federal standard deductions) is used as Minnesota taxable income. Because Texas has no state income tax, this couple will owe the same amount (zero) regardless of marital status. In California, they will pay $145 more in state income tax if they are single than if they are married ($2,132 versus $1,987). A "marriage bonus" arises because California taxes federal adjusted gross income (which does not reflect the differential federal standard deduction), while using a lower tax rate for married couples filing jointly than for singles.[4]

In a seminal study of the marriage-tax link, Alm and Whittington (1999) control for individuals' federal income tax burdens in modeling transitions into marriage and find that the "marriage tax" is associated with a slight decrease in the probability that women marry. While most studies in this vein examine the effect of taxes on transitions into marriage (Alm and Whittington 1995a, 1995b; Brien, Lillard and Stern 2006; Lopez-Laborda and Zarate-Marco 2004), Whittington and Alm (1997) find that tax policy that penalizes married couples also has a small, positive effect on the probability of divorce. We are unaware of a study that examines the effects of tax policy on transitions from cohabitation to marriage. Moreover, most existing studies rely on federal tax laws, which only vary over time, and compute income tax burdens on the basis of *actual* (endogenous) earnings rather than predicted earnings. Our analysis fills these gaps in the literature.

C. Divorce and Property Division

When the decision is made to dissolve a union, married couples are governed by state laws regarding grounds for divorce, the determination of alimony, and the division of property. Most states grant "no fault" divorces, but about half the states consider fault when determining alimony, and states differ significantly in their laws governing property division. No state has explicit laws stating how unmarried, cohabiting couples should divide their property upon dissolving their union. Although 11 states and the District of Columbia recognize common law marriage, couples are required to have cohabitated for many years and to have marital intent. In some states, courts are willing to grant "palimony" to unmarried partners if the couple has a written or implied agreement concerning property settlements. In short, legal protection varies significantly across states for both married and unmarried couples.

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[4]California's personal exemption credit of $66 for singles and $132 for married couples is marriage neutral.

In table 3, we summarize laws governing divorce and the division of property that prove to play an important role in our analysis. Using 1985 for illustration, we see that only two states (South Dakota and Utah) use "fault" as the sole grounds for divorce, meaning the spouse initiating the lawsuit must prove that his/her partner was at fault in order to be granted a divorce. While the remaining 48 states and the District of Columbia grant "no-fault" divorces, 15 of them (those with a "yes" in column B) also consider "fault" as grounds for divorce. Column C indicates that in 1985, 20 states (including the District of Columbia) consider "fault" when determining alimony and the division of property; the remaining 31 states had switched to "no-fault" decisions. From column D, we see that 23 states do not require a mandatory separation before a divorce is granted, while the remaining states impose waiting periods that range from 6 to 36 months.

There is considerable disagreement in the literature over whether the trend toward "no fault" divorce laws affects divorce decisions. Peters (1986, 1992) argues that couples offset the direct effects of no fault laws via contracting. Allen (1992) questions her assumption that utility is perfectly transferable among spouses. Mechoulan (2006) observes that couples who marry prior to the adoption of no fault laws are unable to arrange the appropriate contingent contract unless they have perfect foresight about impending changes in the law. From an empirical standpoint, it is difficult to identify whether divorce laws have a causal effect on divorce decisions. Bougheas and Georgellis (1999), Ellman and Lohr (1998), Friedberg (1998), and Nakonezny *et al.* (1995) differ in the time periods considered, whether short-term or long-term effects are estimated, and how divorce outcomes are measured. A recent study by Wolfers (2006) suggests that the introduction of "no fault" divorce had a small, short-lived effect on divorce rates. Although voices in the policy arena are currently calling for a return to "fault" divorce as a means of lowering divorce rates in the U.S., the literature has yet to establish whether such a switch would significantly affect divorce decisions.

III. Modeling Transitions Into and Out of First Unions

A. Choice Model

We consider three stages in the union-forming process. In stage 1, single women with no prior marriage or cohabiting experience decide whether to stay single, cohabit, or marry; each woman in our sample begins the decision-making process in stage 1. Women who cohabit advance to stage 2, in which first-time cohabiters decide whether to continue cohabiting, dissolve their union, or marry. Anyone who marries in stage 1 or stage 2 advances to stage 3, where they decide whether to remain married or divorce. We focus strictly on *first* unions in this study, so individuals leave the sample upon dissolving their first cohabiting union or marriage. If neither of these events occurs, they are followed until the date of their last NLSY79 interview.

We assume that in each period—defined as a one-year interval—individuals choose the stage-specific alternative that maximizes their expected utility. We assume the expected utility of alternative *j* for individual *i* in stage *g* at time *t* can be expressed as a linear function of various observed and unobserved factors. That is,

$$V_{igt}^{j} = \beta_{g}^{j} X_{igt}^{j} + \varepsilon_{igt}^{j} \quad \text{for } j = s,c,m \text{ and } g = 1,2,3 \tag{1}$$

where X_{igt}^{j} represents observed factors (including current and past spell durations) and ε_{igt}^{j} represents unobserved factors affecting the value of alternative *j* for individual *i* in stage *g* at time *t*. The model allows both observed and unobserved factors to vary across individuals, over time (within and between stages), and across alternatives. In addition, the parameters (β_{g}^{j}) describing the effect of X_{igt}^{j} on expected utility are allowed to vary across stages because the presence of children, relative income tax burdens, divorce laws, and many other factors are likely to have a different effect on the value of being single than on the value of cohabitation or marriage.

B. Estimation issues

In estimating the three-stage model described above, we drop the assumption made throughout the literature that unobserved factors affecting the value of each alternative are uncorrelated across stages. To see why this assumption is undesirable, consider a specific example: suppose single individuals with an unobservable factor such as liberal views toward gender roles tend to cohabit rather than marry. Hence, the stage 2 sample of cohabiters is a select subset of the stage 1 population insofar as individuals tend to possess this particular unobserved characteristic. Suppose further that an observed factor such as ability (measured via test scores), while initially uncorrelated with liberal views, also tends to increase the attractiveness of cohabitation relative to marriage among singles. Clearly, we now have a correlation between the observed and unobserved factors among the stage 2 sample. If we assume transitions made by cohabitors are independent of the prior decision to cohabit, we are unable to separate the *causal* effect of ability from the confounding effect of liberal views. More generally, estimates of *all* stage 2 and stage

3 decisions and *all* stage 1 decisions beyond the first period will be biased unless we explicitly account for the sequential nature of the decision-making process.

Following Cameron and Heckman (1998, 2001) and Hotz *et al.* (2002), we assume the unobserved factors affecting the value of each alternative in each stage can be characterized by a one-factor loading structure:

$$\varepsilon_{igt}^{j} = \alpha_{g}^{j}\phi_{i} + \nu_{igt}^{j} \tag{2}$$

where α_g^j is a factor loading (to be estimated) specific to alternative *j* and state *g*, ϕ_i represents time-invariant, individual-specific unobservables, and ν_{igt}^j represents other unobservables that vary across individuals, across alternatives, and over time. We assume ν_{igt}^j is independent of ϕ_i and independent across individuals, alternatives, and time periods. In other words, neither within-stage nor cross-stage correlations in the unobservables arise from this component of the residual. We also assume ν_{igt}^j is drawn from an extreme value distribution, which means the transition probabilities that form the likelihood function have a logistic structure.

We assume the other unobserved factor, ϕ_i, is independent across individuals. The error structure given by equation (2) implies that ϕ_i is the source of any intertemporal (or cross-alternative) dependence in the choices made by individual *i,* conditional on her observables. To eliminate the dynamic selection bias illustrated above, we need only integrate out ϕ_i after making an assumption about its distribution. We assume ϕ_i is drawn from the standard normal distribution. In estimating the model, we restrict the coefficients and loading factor associated with one alternative to be zero for identification, and we use 10-point Gauss-Hermite integration to evaluate the log-likelihood function.[5]

IV. DATA

A. Sample Selection

Our primary data source is the 1979 National Longitudinal Survey of Youth (NLSY79). The survey began in 1979 with a sample of 12,686 individuals born in 1957-1964. The original

[5]We experimented with allowing the factor loadings to vary across age intervals, and with assuming ϕ_i is drawn from a discrete distribution with as many as five support points. Neither innovation substantively changed our findings.

sample is 60% nonblack, non-Hispanic ("white"), 25% black, and 15% Hispanic, and roughly 50% male. Although some attrition has occurred, most respondents were surveyed annually from 1979 to 1994 and biennially thereafter. We use data for survey years 1979 through 2004.[6]

In selecting a sample for our analysis, we begin by eliminating the 6,403 NLSY79 respondents who are male. We confine our attention to women because they (and their children) tend to be the focus of marriage-related public policy. Next, we eliminate 1,424 female respondents whose 20th birthdays occur more than six months before the 1979 interview date and an additional 1,360 women who marry or cohabit prior to their 20th birthday. These selection rules produce a sample of 3,499 women who are single (never married, never cohabited) when first observed. We use the age 20 cutoff in order to initialize stage 1, to the extent possible, on the basis of an exogenous factor (age) rather than self-determined events such as observed first unions. Ideally, we would initialize stage 1 when each woman starts making union-forming decisions, but this occurs prior to the start of the survey for many women, given that individuals can cohabit long before they can legally marry.[7]

Among the 3,499 women who are observed from their 20th birthdays onward and who are single at that starting date, we eliminate another 71 individuals because key variables are missing. The remaining samples consists of 3,428 women, 1,056 (31%) of whom are black and 517 (15%) of whom are Hispanic; we refer to the remaining 1,855 women as white. We obtain an annual observation for each of these individuals, starting in 1979 and terminating when their first union ends or they are last interviewed. This strategy produces a sample of 46,675 person-year observations.[8]

B. Transitions between Single, Cohabiting, and Married

For each person-year observation, we identify the respondent's state as single, cohabiting, or married, and we determine which new state, if any, she transits into during the succeeding year. If a woman is single at the time of the 1990 and 1991 interviews, for example, we must ensure

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[6]Confidential geocode data for 2006, which we require to learn each respondent's state of residence, will not be available until fall 2008.

[7]We could use an earlier initialization age such as 17, but more than 3,000 respondents are past their 17th birthday when interviewed in 1979. We would either have to drop these respondents or estimate key, time-varying information such as state of residence for the pre-survey years.

[8]We include observations for noninterview years (1995, 1997, *etc.*) by identifying marital status, number of children, and residential location from the surrounding, non-missing years.

that no marriage or cohabitation takes place between interviews in order to identify the observation as “single-to-single” (SS). Similarly, if she reports herself as cohabiting in 1990 and 1991, we must make sure she is with the same partner to classify the observation as “cohabiting-to-cohabiting” (CC) rather than “cohabiting-to-single” (CS).

To associate each person-year observation with a transition type, we use all available information on marriage, cohabitation, and divorce. NLSY79 respondents report their marital status at each interview, and they also provide a complete event history of the dates when each marriage begins and ends. From 1990 onward, dates for cohabitation spells are reported as well. Although start and end dates for cohabitation spells are not reported prior to 1990, we know whether the respondent is cohabitating at the time of each interview, and we also have identifier codes for each cohabiting partner. If a respondent is cohabiting in two successive interviews but with different partners, we will correctly identify the CS transition. Similarly, when a respondent divorces and remarries between interviews there is no danger of treating him as continuously married; the transition will be correctly identified as “married-to-single” (MS).

In the top panel of table 4, we show the distribution of year-to-year transitions for our person-year sample, disaggregated by race/ethnicity. Because most respondents spend the bulk of the observation period either single or married, SS and MM transitions account for almost 90% of all observations. However, SS accounts for 62% of observations contributed by black women, and only 35-40% of the observations for whites and Hispanics; this is consistent with black women spending smaller portions of their lives in marriage relative to nonblacks (Bennett, Bloom and Craig 1989; Cherlin 1992; Raley 1996). Cohabitation spells are relatively short-lived, on average, so CC transitions account for only about 2% of each subsample’s observations. While respondents can contribute multiple SS, CC, and MM transitions, they can contribute only one SC or SM transition, one CS or CM transition, and one MS transition. As a result, these transition types account for, at most, 4.5% of the person-year observations for each subsample.

In the bottom panel of table 4, we characterize the sequence of transitions for each subsample of women. We see that 12.5% of whites, 29% of blacks and 15% of Hispanics are single (S) for the entire observation period, while 42% of whites, 23% of blacks and 37% of Hispanics transition from single to married (SM) and another 17-27% transition from single to married to divorced (SMS). Together, these three patterns account for roughly three-quarters of each subsample. The remaining women are seen cohabiting at some point during the observation period. Between 12% and 15% of women separate from their first cohabiting partner (SCS), while 15% of whites and 6-7% of blacks and Hispanics marry their cohabiting partner and, in 35-50% of these cases, eventually divorce. Unsurprisingly, the divorce rate among individuals whose prior patterns are SCM is generally higher than the divorce rate seen among individuals who transition directly from single to married. This is consistent with the evidence in Axinn and Thornton (1992), Bumpass and Lu (2000), Lillard, Brien and Waite (1995), and Sweet and Bumpass (1992).

C. Covariates

To control for potential welfare benefits, we include measures of the maximum, monthly AFDC or TANF benefit available for a family of four, divided by the implicit price deflator for gross domestic product (GDP), and the average Medicaid expenditure for a family of four, divided by the consumer price index (CPI) for medical care. Both measures are specific to the state of residence and calendar year corresponding to the person-year observation.[9] Summary statistics for these and other variables used in our analysis appear in table 5.

To measure the expected income tax penalty (or bonus) associated with marriage, we begin by predicting the income of both the sample member and her spouse or partner for every person-year observation; if the woman is currently single we use an "expected" partner who is the same race as the woman, but two years older. We then assign each woman and her (real or expected) partner the average income earned by individuals of the same sex, race and age. By combining these expected income values with information on the couple's state of residence, we compute the couple's combined state income tax liability, first under the assumption that they are cohabiting or single and filing separately, and again under the assumption that they are married and filing jointly. The income tax variable that we include in our model is the difference between these predicted tax burdens.

Our predicted tax variables are highly correlated with tax obligations based on actual income, but they depend entirely on state of residence, calendar year, and exogenous determinants of income such as age and sex. In addition, they vary considerably both within and

[9]These variables are from the welfare benefit database ben_dat.txt that is available, along with documentation, at http://www.econ.jhu.edu/People/Moffitt/datasets.html and from the Urban Institute's welfare rules database available at http://www.urban.org/toolkit/databases/index.cfm.

across years because we rely on state income tax policies rather than federal laws. In table 5, we see that the mean of this tax variable (deflated by the implicit price deflator for GDP) is $53, with a large standard error of $180.

Our covariates also include measures of state laws governing divorce, the division of property, and alimony. As discussed in section II, there is considerable variation across states in whether "fault" *must* be established as grounds for divorce, whether "fault" *can* be used as ground for divorce, or whether the state is strictly "no fault." We use two dummy variables indicating whether the state uses "fault" or "no fault" in the given calendar year as the sole grounds for divorce; the omitted category identifies states that grant "no fault" divorces while also allowing "fault" to be established. We also include a dummy variable indicating whether the state uses "no fault" in determining alimony and the division of property. In addition, we control for the length of time that a married couple is required to separate before a no-fault divorce is granted. This variable equals zero if the state imposes no separation requirement.

Table 5 also summarizes the remaining covariates included in our models. To control for family influences that might affect union formation, we include an array of dummy variables indicating the woman's reported religion, if any ("none" is the omitted category) and another set of dummies indicating whether the woman lived with both biological parents or her mother only at age 14. We also control for the highest grade completed by the respondent's mother.

To control for race and ethnicity, we include dummy variables indicating whether the woman is black or Hispanic (with white the omitted group), and we also interact these race/ethnicity dummies with selected covariates identified in table A-1. We determine which interactions to include by estimating models in which every covariate is allowed to vary by race/ethnicity and eliminating interactions that prove to be statistically insignificant *and* have no important effect on the estimates reported in tables 6-7. This "intermediate" strategy allows us to identify important racial/ethnic differences in the estimated slopes (rather than assume that race operates solely through the intercept), while reducing the number of parameters that we estimate. Other personal characteristics include age-adjusted scores on the Armed Forces Qualifying Test (AFQT), which we view as an exogenous measure of the woman's earnings potential.[10] We also

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[10]AFQT scores are derived from scores on the Armed Services Vocational Aptitude Battery, which was administered to NLSY79 respondents in 1980. We regress percentile AFQT scores on a set of birth-year dummies and use the residual as the age-adjusted score.

control for the number of children in the household, and whether an adult other than the respondent and her spouse/partner lives in the household.

We construct two environmental variables by merging county-level data from the City and County Data Book (collected by the U.S. Census Bureau) with the NLSY79 data, using county of residence indicators from the NLSY79 Geocode file.[11] From these Census data, we compute the county unemployment rate and the percent of the county population that is the same race/ethnicity as the respondent. We also constructed measures of per capita personal income in the county and the percent of the county population that is married, but these variables proved to have little explanatory power so we do not include them in our models. Variables such as these have been proposed by Lichter, LeClere and McLaughlin (1991) and Gould and Paserman (2003) to control for economic opportunity and the characteristics of local marriage markets.

The probability of a union-related transition has been shown to change dramatically with both current spell duration and the duration of past spells (Bennet, Blanc and Bloom 1988; Lichter *et al.* 2006; Lillard *et al.* 1995). We allow for duration dependence in a flexible manner by including dummy variables indicating whether the spell is in its first year, second year, and so forth. We also include the completed duration of the woman's "single" spell (which is equivalent to controlling for the age at which she formed her first union) among the stage 2 and stage 3 covariates, and we control for the completed cohabitation duration (which is zero if the woman transitioned directly from single to marriage) in stage 3.

V. FINDINGS

A. Predicted Probabilities of Year-to-Year Transitions

Table A-1 contains estimated parameters for our choice model. Because these estimates cannot be readily interpreted we immediately turn our attention to table 6, which shows the predicted probability that a representative woman of each race/ethnicity makes each annual transition (excluding SS and MM), and the effects on these predicted probabilities of various hypothetical interventions. Our representative woman is assumed to be one year into her given spell and to have characteristics equal to the mean or mode for her stage-specific sample. Specifically, her state's welfare benefits, the "marriage penalty" she pays in state income taxes, and the

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[11]The City and County Data Books do not provide annual observations for each variable of interest, so we use the closest available observation for each person-year observation in our sample.

mandatory separation she has to complete to be granted a divorce are all equal to the sample mean. This woman lives in a state that uses both "fault" and "no-fault" as grounds for divorce and "fault" as grounds for property division and alimony, and she is a Protestant who lived with both parents at age 14, has no children, lives with no adults aside from a spouse or partner, and has all other variables (AFQT score, county unemployment rate, *etc.)* equal to the sample mean.[12]

To assess the estimated effects of welfare benefits, we consider increasing potential AFDC or TANF benefits from the sample mean to an amount that would place the state's generosity in the 90th percentile of the distribution. Our computations suggest that this change in policy causes single women (*i.e*., women in stage 1) to favor cohabitation over marriage: the predicted probability of a SC transition increases by 50% for whites, 83% for blacks, and 33% for Hispanics, while the predicted probability of SM falls by about 25% for all three groups. These patterns, which mimic the findings of the AFDC-based studies discussed in section II, reflect the fact that AFDC programs severely limit the benefits available to married women. Benefit levels prove to have small and statistically insignificant effects on the transitions undertaken by cohabiting women (stage 2), but the marriage disincentive reappears among married black women in stage 3, whose predicted probability of divorce increases by 34% when benefits are raised. For whites and Hispanics, however, increased benefits are predicted to lower divorce probabilities by about 12%. Given that TANF programs are in place by the time many women in our sample enter stage 3, these results suggest that the marriage incentives built into TANF are effective for select segments of the population.

In stage 1, the estimated effects of Medicaid benefits are similar to what we saw for AFDC/TANF. That is, single women in each race/ethnic group are expected to substitute away from marriage toward cohabitation in response to an increase in the generosity of their states' Medicaid program. In contrast to what we found for AFDC/TANF benefits, this marriage disincentive now carries over to stage 2: increased Medicaid benefits lower the predicted probability of CM transitions by around 30% for all three groups although, interestingly, these women tend to substitute toward dissolution rather than continued cohabitation. In stage 3, we now predict that all married women (including blacks) are 42% less likely to divorce when Medicaid benefits become more generous—a finding that is consistent with the marriage-

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[12] We use the delta method for computing standard errors for the conditional probabilities.

enhancing effects of Medicaid expansions estimated by Yelowitz (1998).

We find that marriage penalties arising from state income tax laws have small, statistically insignificant effects on the estimated transition probabilities of white and Hispanic women. However, we predict that cohabiting black women are 16% less likely to marry (and 14% more likely to continue cohabiting) when the "marriage tax" increases, while married black woman are 11% more likely to divorce. These estimated effects are statistically significant and consistent with expectations: when tax policy increases the relative cost of marriage, individuals substitute cohabitation or being single for marriage. Surprisingly, the estimated effects among single black women are also statistically significant but "wrong-signed." One explanation is that for single black women, the value of a partner with similar earnings potential (which is what generates a tax penalty rather than a tax bonus) more than offsets the additional tax burden. Given that we compute tax burdens for single women based on the income of a hypothetical partner (see appendix), another explanation is that blacks are less likely than whites and Hispanics to choose a mate with our assumed age, race, and schooling attainment.

To introduce a hypothetical change in divorce law, we assume the woman's state requires that "fault" be established before granting a divorce, and that a mandatory separation of 36 months (which falls in the 90th percentile of the separation duration distribution) is imposed; the baseline assumption is that both "fault" and "no fault" divorces are granted and that a 10 month separation is mandated. For whites and Hispanics, this change in policy is expected to increase the likelihood of marriage, both by increasing the probability of SM transitions by about 36% and lowering the probability of MS transitions by 61%. This intervention has an even larger effect on the predicted transition probabilities of black women: the predicted probability of a SM transition increases by 63% while, contrary to conventional wisdom, the predicted probability of MS transitions also increases by a staggering 245%.[13] These estimated effects are larger than what is typically found in the literature (*e.g.,* Friedberg 1998; Wolfers 2006), presumably because we cannot separate the effects of divorce law from the effects of state-specific attitudes toward divorce that are correlated with policy.

Turning to factors that are less readily manipulated by policy, both childhood household

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[13]Although the estimated MS effect for blacks and all estimated stage 2 effects are very large in magnitude and statistically significant, we observe only a handful of cohabiting unions and marriages among blacks in "fault" states and virtually all those unions end in dissolution.

composition and the presence of children prove to have similar effects on union formation. For all race/ethnic groups, the "intervention" of having lived with a single mother at age 14 is predicted to lower the probability of SM transitions by as much as 40%, lower the probability of CM transitions by about 11% (although the CM effects are imprecisely estimated), and raise the probability of CS (MS) transitions by about 60% (31%). Similarly, having two children is predicted to lower the probability of SM (CM) transitions by about 40% (19-27%). While the predicted probability of divorce for blacks and Hispanics is 22-28% higher when children are present than when they are not, children are predicted to lower the probability of MS transitions by 20% among whites; this racial difference may be due to the fact that white women are more likely than nonwhites to be married to the biological father of their children. Aside from the negative effect of children on white women's predicted divorce probabilities, both interventions serve to deter entry into marriage (both SM and CM) and promote union dissolution (both CS and MS).

Raising a woman's AFQT score—which is an exogenous predictor of her earnings potential—from the sample mean to the 90th percentile has an imprecisely estimated, negative effect on SC and SM transitions. However, this "intervention" is found to encourage marriage once she enters stages 2 and 3: the predicted probability of CM transitions increases by 20-33%, while the predicted probability of MS transitions declines by around 35% for all three race/ethnic groups. It appears that the marriage-deterring effect of economic independence dominates among single women, while the increased gains to marriage associated with higher earnings dominates among women with partners.

The estimated effects of the remaining factors—religion, county unemployment rates, and county race composition—vary widely across race/ethnic groups. We predict that single, white women who are Baptist are 53% more likely than their Protestant counterparts to marry. Among whites who opt for cohabitation, the predicted probability of separation is 64% higher for Baptists than for Protestants. Among blacks, both of these predicted effects are opposite in sign. While higher unemployment rates are predicted to raise the probability of CS transitions for all three race/ethnic groups, black cohabitors are the only subsample to substitute away from marriage in the face of weak labor market conditions. Living in a county with an unusually high proportion of residents of the same race increases the predicted probability that single black women marry *and* that married black women divorce. Among whites and Hispanics, both estimates are much smaller in magnitude and of the opposite sign.

To summarize, a number of distinct patterns emerge from the estimates seen in table 6. First, large changes in AFDC/TANF benefits, Medicaid benefits, and divorce laws (but not income tax penalties) prove to have substantial effects on the predicted probabilities of select union transitions. Second, several factors that are outside the control of public policy, including religion, childhood living arrangements, and the presence of children have equally large predicted effects. Third, the estimated effects of many factors differ radically across our three race/ethnic groups, with black women often behaving differently than whites and Hispanics. Fourth, many interventions that we consider are predicted to promote marriage in one stage but discourage it in another. As a result, our current focus on annual transition probabilities leaves us without a clear understanding of which factors encourage long-term marriages or long-term unions in general. We address this latter issue in the next subsection.

B. Predicted Probabilities of Long-Term First Unions

To assess the effect of each intervention on long-term union formation, we first compute the predicted probability that a representative woman in each race/ethnic group follows two hypothetical paths: marry by age 25 and remain married for at least nine years, and form any union (marriage or cohabitation) by age 25 and maintain that partnership for at least nine years. These predicted probabilities are shown in the top panel of table 7. The predicted probabilities of long-term marriage and long-term unions for a representative white woman are 0.375 and 0.411, respectively; as shown in the right-most column, the inclusion of cohabitation increases the likelihood of a long-term union by almost 10%. Table 7 also shows that black and Hispanic women are less likely than whites to have long-term unions, and that cohabitation accounts for a larger proportion of such unions among blacks than among other women.

The remainder of table 7 shows how our assumed "interventions" change the predicted probabilities of following each hypothetical path. For example, an increase in AFDC/TANF benefits lowers a representative white woman's predicted probability of entering a long-term marriage by 14.9%, and lowers her predicted probability of "any union" by 8.3%. Because the formation of "any union" is now 18.2% more likely than marriage (versus 9.6% without this intervention), it is apparent that generous welfare benefits cause a shift toward cohabitation among women with long-term relationships. Although an assessment of the benefits to marriage and cohabitation is beyond the scope of our analysis, there is little reason to believe that cohabitation is less preferred than marriage in this context, given that we focus on long-term unions. Most evidence on the relative merits of marriage is based on the fact that cohabiting unions tend to be quite short (Bumpass *et al.* 1991; Lichter *et al.* 2006); when long-term unions are analyzed, marriage is not necessarily more beneficial than cohabitation (Willets 2006).

Four of our assumed "interventions"—increased AFDC/TANF and Medicaid benefits, living without a father at age 14, and the presence of children—prove to discourage long-term union formation for all women and, in most cases, cause a shift from marriage toward cohabitation among women who form long-term unions. These cumulative effects cannot be discerned from table 6, given that a factor might be predicted to lower the probability that a single woman marries while simultaneously lowering the probability that a married woman divorces. By "adding up" the entire sequence of underlying transition probabilities, we now see that all four interventions are expected to produce fewer long-term unions. As we established in section II, it is widely understood that welfare programs can be (and have been) tailored to affect marriage and cohabitation decisions. Our estimates confirm that reduced benefit levels can be expected to raise the likelihood of long-term unions—but we also find that a woman's childhood household composition and her decision to have children have large estimated effects that can potentially offset such policies.

The remaining economic factors that we consider—an increased "marriage tax" and a switch to "fault" divorce"—require careful interpretation to avoid putting undue weight on the occasional anomalous estimate. An increase in the "marriage tax" lowers the predicted probability that white and Hispanic women form long-term unions, but these effects are small in magnitude and imprecisely estimated. For blacks, the same intervention is predicted to *increase* the probability of long-term unions because, as noted in the preceding subsection, we appear not to be capturing the true cost of marriage facing single black women. We believe the "non effect" of tax policy estimated for whites and Hispanics is the more reliable finding. Similarly, when assessing the estimated effects of divorce policy, we discount the results for blacks because small sample sizes lead us to find (table 6) that a switch to "fault" *increases* the predicted probability of divorce. Focusing instead on the findings for white and Hispanic women, we conclude that "conservative" divorce laws (lengthy, mandatory separation requirements coupled with a requirement that fault be established) increase the predicted probability of long-term unions.

Religion is a factor that operates differently for blacks and nonblacks. We predict that white and Hispanic Baptists are significantly more likely to form long-term unions than are observationally equivalent Protestants (although very few Baptist Hispanics appear in our sample), while the opposite effect is seen among blacks. While higher AFQT scores are associated with small increases in the predicted probability of long-term unions for all three race/ethnic groups, increased racial representation in one's county of residence lowers the predicted probability that black women form long-term unions (by raising the likelihood of divorce more than the likelihood of marriage) while having virtually no effect for whites and Hispanics. In general, the characteristics of local marriage markets and labor markets prove to have trivial estimated effects on union formation.

VI. CONCLUDING COMMENTS

Current U.S. public policy can be characterized as pro-marriage. Examples include the Economic Growth and Tax Relief Reconciliation Act of 2001, which changed federal tax law to provide "marriage penalty" relief; the 1996 Healthy Marriage Initiative, in which numerous federal programs provide services to help couples sustain their marriages; and Covenant Marriage laws passed in Arkansas, Oklahoma, and Louisiana that allow married couples to limit the grounds by which they can be divorced. Many social scientists have taken a pro-marriage stance in their research by arguing that marriage enhances a range of important outcomes. Prominent examples include Waite (1995) and Waite and Gallagher (2000).

In our view, discussion about union formation will benefit from additional information on two issues. First, social scientists should continue to learn whether marriage (or cohabitation) *causes* various outcomes rather than simply being correlated with them. Progress has been made in assessing the causal effects of marriage and cohabitation on wages (Cornwell and Rupert 1997; Gray 1997; Korenman and Neumark 1991; Loh 1996; Stratton 2002), family income (Light 2004), and selected child outcomes (Levine and Painter 2000), but much "pro-marriage" evidence continues to be based on cross-sectional correlations. Second, if the promotion of marriage is judged to be desirable, we should learn more about how marriage decisions can be influenced—that is, we should identify factors that can be *manipulated* by public policy and that *causally* increase the probability that individuals will get married and stay married. Without taking a stand on whether such policy is desirable, we contribute evidence on this second issue.

We have found that income tax policy that raises the relative cost of marriage generally has very small, imprecisely estimated effects on women's union forming decisions. However, the other economic factors that we consider—AFDC/TANF benefits, Medicaid benefits, and divorce law—prove to be important determinants of union-related transitions. While the effects of these factors on union formation have been studied before, we are able to compare their estimated effects side by side, and use our sequential choice model to determine which policy interventions are expected to promote long-term unions. This proves to be important, given that a particular intervention may be predicted to lower the probabilities of both entry into and exit from marriage, for example, or lower the probability of entry into marriage while raising the probability of entry into cohabitation. We find that increased AFDC/TANF and Medicaid benefits substantially lower the predicted probability of long-term marriages and long-term unions of any type (marriage or cohabitation), while eliminating the "no fault" option from divorce law significantly raises the predicted probability of long-term unions. At the same time, we find that factors that are not easily controlled by public policy, such as religion, childhood household composition, and the presence of children, are often predicted to be equally important determinants of long-term unions. Our estimates suggest that policy interventions may have to be quite severe to offset the effects of non-policy related factors.

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Table 1: AFDC and Medicaid Payments Available in Select States
(Calendar Year 1995)

	Maximum AFDC Payment		Average Medicaid Expenditure	
Rank	State	Amount	State	Amount
1	Mississippi	144	Arizona	61
2	Alabama	194	Hawaii	87
3	Tennessee	226	Washington	231
26	Maryland	450	S. Carolina	397
49	Vermont	731	Maine	641
50	Hawaii	859	Alaska	642
51	Alaska	1,025	New York	655

Note: Payments shown are nominal, monthly dollar amounts for a family of four.

Source: Robert Moffitt's welfare benefit database; see text for details.

Table 2: Federal and State Income Tax Obligations in Select States, by Filing Status
(Tax Year 1995)

	Earned Income	Filing Status	Federal		Minnesota[a]		Texas		California[b]	
			Tax	Rate[c]	Tax	Rate[c]	Tax	Rate	Tax	Rate[c]
a)	$20,000	single	$2,411	0.15	$980	.080	$0	0	$495	0.14
b)	$35,000	single	$5,666	0.28	$2,172	.080	$0	0	$1,637	0.06
c)	$55,000	single	$8,077		$3,152		$0		$2,132	
d)	$55,000	joint	$8,503	0.28	$3,419	.080	$0	0	$1,987	0.08
d)-c)			$426		$267		$0		-$145	

[a]Minnesota uses federal taxable income (earned income in this example) as taxable income.
[b]California uses federal adjusted gross income as taxable income.
[c]Marginal income tax rate.
Note: Each individual/couple is assumed to have no taxable income other than earned income, no itemized deductions, and no dependents. Row c) shows the (nominal) total tax bill for a cohabiting couple with the assumed income levels, while row d) shows the tax bill if they marry and file jointly.
Source: Tax forms available at http://www.irs.gov/formspubs, http://www.taxes.state.mn.us and /taxes and http://www.ftb.ca.gov/forms.

Table 3: Summary of State Laws Governing Divorce and Division of Property

State	A.	B.	C.	D.	State	A.	B.	C.	D.
AK			2005+	24	MT		Yes	1975	6
AK		Yes	1974	24	NE		Yes	1972	0
AZ		Yes	1973	0	NV			1973	12
AR			1979	18	NH			2005+	24
CA		Yes	1969	0	NJ			1980	18
CO		Yes	1971	0	NM			1976	0
CT			2005+	18	NY			2005+	12
DE			1974	6	NC			2005+	12
DC		Yes	2005+	12	ND			2005+	0
FL		Yes	1986	0	OH			2005+	12
GA			2005+	0	OK			1975	0
HI			1960	24	OR		Yes	1971	0
ID			1990	0	PA	1980		2005+	24
IL	1983		1977	24	RI			2005+	36
IN			1973	0	SC			2005+	12
IA		Yes	1972	0	SD	1989		2005+	0
KS			1990	0	TN			2005+	24
KY		Yes	2005+	2	TX			2005+	36
LA			2005+	6	UT	1987		1987	36
ME			1985	0	VT			2005+	6
MD			2005+	12	VA			2005+	12
MA			2005+	0	WA		Yes	1973	0
MI		Yes	2005+	0	WV			2005+	12
MN		Yes	1974	0	WI		Yes	1977	0
MS			2005+	0	WY			2005+	0
MO			2005+	12					

A. Fault must be established for a divorce to be granted until this year; if no year is given, fault was dropped as the sole grounds for divorce prior to 1979.

B. "Yes" means both fault and no-fault divorces are granted; otherwise, only no-fault divorces are granted.

C. Fault is considered in deciding property division and alimony until this year; 2005+ means fault remained in effect beyond 2004.

D. Months of required separation before a no-fault divorce is granted.

Source: Ellman and Lohr (1998), Mechoulan (2001), American Bar Association (2004)

Table 4: Distribution of Year-to-Year Marital Status Transitions
and Distribution of "Overall" Transition Patterns, by Race/Ethnicity

	Whites		Blacks		Hispanics	
Year-to-Year Transition	Number	Percent	Number	Percent	Number	Percent
Single to:						
Single (SS)	8,750	35.5	9,115	61.6	2,929	40.4
Cohabiting (SC)	523	2.1	224	1.5	113	1.6
Married (SM)	1,100	4.5	528	3.6	327	4.5
Cohabiting to:						
Cohabiting (CC)	442	1.8	281	1.9	173	2.4
Single (CS)	213	0.9	156	1.1	78	1.1
Married (CM)	285	1.2	66	0.5	34	0.5
Married to:						
Married (MM)	12,879	52.3	4,120	27.8	3,443	47.5
Divorced (MS)	431	1.8	317	2.1	148	2.0
All person-year obsns.	24,623	100.0	14,807	100.0	7,245	100.0
Overall transition pattern	Number	Percent	Number	Percent	Number	Percent
S	232	12.5	304	28.8	77	14.9
SC	25	1.4	2	0.2	1	0.2
SCS	213	11.5	156	14.8	78	15.1
SCM	174	9.4	33	3.1	22	4.3
SCMS	111	6.0	33	3.1	12	2.3
SM	780	42.1	244	23.1	191	36.9
SMS	320	17.3	284	26.9	136	26.3
All individuals	1,855	100.0	1,056	100.0	517	100.0

Table 5: Summary Statistics for Covariates Used in Choice Model

Variable	Description	Mean	S.D.
	Legal factors		
AFDC	Maximum monthly AFDC/TANF benefit for family of four[a]	475.44	177.39
Medicaid	Average monthly Medicaid expenditure for family of four[a]	416.75	220.13
Marriage tax	State income tax if married minus state income tax if unmarried[ab]	53.05	179.76
Fault	1 if "fault" only grounds for divorce	0.01	
No fault	1 if "no fault" only grounds for divorce	0.32	
Property	1 if "no fault" used for property division and alimony	0.38	
Separation	Separation (in months) required before no fault divorce can be granted	10.80	11.26
	Family background		
Baptist	1 if Baptist	.28	
Protestant	1 if Protestant	.21	
Catholic	1 if Catholic	.36	
Other relig	1 if other religion	.11	
Live both	1 if lived with mother/father at age 14	.71	
Live mom	1 if lived with mother only at age 14	.17	
Mom HGC	Mother's highest grade completed	11.12	3.14
	Personal characteristics		
Black	1 if black	.32	
Hispanic	1 if Hispanic	.16	
AFQT	Age-adjusted AFQT score	1.19	28.12
Children	Number of children in household	0.83	1.14
Adults	1 if living with other adults (not a partner)	0.38	
	Environmental factors (county-specific)		
Unemp rate	Unemployment rate	7.27	3.24
Same race	Percent of population of same race/ethnicity	57.57	32.05
	Number of person-year observations	46,675	

[a]AFDC and Income tax are divided by the implicit price deflator for GDP; Medicaid is divided by the CPI for medical care (base year=2000).

[b]Based on predicted income. See text for details.

Note: The model also includes dummy variables indicating current spell duration, measures of previous spell durations (time spent single and cohabiting), and interactions between Black and Hispanic and selected variables. See table A-1.

Table 6: Percent Change in Predicted Probability of Year-to-Year Marital Status Transition Due to Changes in Observed Characteristics, by Race/Ethnicity

		Stage 1		Stage 2			Stage 3
	Race	SC	SM	CS	CC	CM	MS
Baseline predicted probability of year-to-year transition[a]	White	.018*	.108*	.022*	.639*	.339*	.061*
	Black	.012*	.065*	.079*	.461*	.460*	.083*
	Hispanic	.006	.022	.058*	.128	.142	.017
Percent change in prediction if							
Maximum AFDC or TANF payment raised from mean to 90th percentile	White	50.0*	-23.2*	-4.6	-0.2	0.3	-13.1*
	Black	83.3*	-27.7*	-3.8	0.2	0.4	33.7*
	Hispanic	33.3*	-25.7*	0.0	0.0	0.6	-11.6*
Average Medicaid expenditure raised from mean to 90th percentile	White	50.0*	-66.7*	186.4*	8.1	-27.1*	-42.6*
	Black	95.8*	-14.4*	168.4*	2.4	-31.3*	-42.2*
	Hispanic	33.3	-68.6*	180.0*	4.5	-29.8*	-41.9*
Marriage tax raised from mean to 90th percentile	White	5.6	-3.7	-4.6	-0.8	1.8	-1.6
	Black	-33.3*	12.3	12.7	13.9*	-15.9*	10.8*
	Hispanic	0.0	-5.7	0.0	-0.4	2.2	-2.3
State divorce laws changed to "fault" and separation raised from mean to 90th percentile	White	-33.3*	36.1*	999.9*	-99.9*	-99.8*	-60.7*
	Black	-58.3*	63.1*	999.9*	-99.9*	-99.9*	245*
	Hispanic	-33.3	37.1*	999.9*	-99.9*	-99.7*	-60.5*
Religion changed from Protestant to Baptist	White	-11.1	52.8*	63.6*	-6.4*	8.0	3.3
	Black	0.0	-35.4*	-58.2*	55.5*	-45.7*	3.6
	Hispanic	133.3*	228.6*	999.9*	4.3	-99.7*	4.7
Living arrangements at age 14 changed from both parents to mother only	White	122.2*	-39.8*	59.1*	3.4	-10.3	31.2*
	Black	-16.7	-13.9*	55.7*	2.0	-11.7	31.3*
	Hispanic	100.0*	-40.0*	60.0*	2.0	-11.6	34.9*
AFQT score raised from mean to 90th percentile	White	-11.1	-7.4	-13.6	-10.2*	20.1*	-36.1*
	Black	-16.7	-7.7	-63.3	-22.6*	33.3*	-34.9*
	Hispanic	-16.7	-8.6	-10.0	-5.6	25.9	-34.9*
Number of children raised from zero to two	White	16.7*	-38.0*	9.1	11.9*	-22.7*	-19.7*
	Black	16.7	-38.5*	11.4	17.1*	-19.1*	21.7*
	Hispanic	66.7	-40.0*	0.0	6.1	-26.5	27.9*
County unemployment rate raised from mean to 90th percentile	White	5.6	-0.9	40.9*	-9.4*	15.0*	-1.6*
	Black	0.0	0.0	58.2*	16.1*	-25.9*	-1.2
	Hispanic	0.0	-2.9	50.0	-5.1	20.4	0.0
Percent of county population that is same race raised from mean to 90th percentile	White	-16.7*	-3.7	4.6	-3.4	6.2	-11.5*
	Black	0.0	38.5*	1.3	-5.2	4.4	47.0*
	Hispanic	-16.7	-5.7	10.0	-1.7	8.3	-11.6*

See notes on next page.

Table 6 (notes)

[a]All covariates equal the stage-specific sample mean if continuous or mode if discrete; current spell duration is assumed to be one year.

*Reject null hypothesis that change in predicted probability (or level of predicted probability in top panel) is different than zero using 10% significance level.

Note: Predicted percent changes are top-coded at 999.9%; these very large estimated effects arise in stage 2 when, due to small sample sizes, virtually all women with a given characteristic make the same transition. Computations are based on the estimates shown in table A1.

Table 7: Percent Change in Predicted Probability of "Long Term" Unions
Due to Changes in Observed Characteristics, by Race/Ethnicity

	Race	(A) Marriage	(B) Any Union	(A to B) % change [a]
Baseline predicted probability of marrying or forming any union by age 25 and staying in union for at least 9 years[b]	White	.375	.411	9.6
	Black	.242	.274	13.2
	Hispanic	.171	.188	9.9
Percent change in prediction if [c]				
Maximum AFDC or TANF payment raised from mean to 90th percentile	White	-14.9	-8.3	18.2
	Black	-37.2	-24.8	15.5
	Hispanic	-19.3	-13.3	18.1
Average Medicaid expenditure raised from mean to 90th percentile	White	-36.3	-16.3	43.9
	Black	-72.3	-41.6	138.8
	Hispanic	-42.1	-24.5	43.4
Marriage tax raised from mean to 90th percentile	White	-3.7	-2.4	11.1
	Black	11.6	4.4	5.9
	Hispanic	-5.3	-3.7	11.7
State divorce laws changed to "fault" and separation raised from mean to 90th percentile	White	52.3	38.9	16.5
	Black	-49.4	-44.6	14.2
	Hispanic	22.2	14.9	11.9
Religion changed from Protestant to Baptist	White	26.1	20.7	4.9
	Black	-9.5	-11.3	11.0
	Hispanic	144.4	124.5	1.0
Living arrangements at age 14 changed from both parents to mother only	White	-42.1	-29.9	32.7
	Black	-23.1	-24.5	11.3
	Hispanic	-43.9	-30.9	35.4
AFQT score raised from mean to 90th percentile	White	9.9	9.7	9.5
	Black	13.2	14.2	14.2
	Hispanic	3.5	2.1	8.5
Number of children raised from zero to two	White	-25.1	-20.4	16.4
	Black	-40.5	-35.4	22.9
	Hispanic	-40.4	-29.8	29.4
County unemployment rate raised from mean to 90th percentile	White	0.3	1.0	10.4
	Black	0.4	-1.5	11.1
	Hispanic	0.0	0.0	9.9
Percent of county population that is same race raised from mean to 90th percentile	White	3.2	2.2	8.5
	Black	-5.8	-8.0	10.5
	Hispanic	0.6	-1.1	8.1

See notes on next page.

Table 7 (notes)

[a]This column shows the percent change in the predicted probability of “any union” relative to marriage for the given row.

[b]All covariates except duration-related variables equal the sample mean if continuous or mode if discrete.

[c]The “marriage” and “any union” columns show the percent change in the predicted probability relative to the baseline predictions shown in the first rows.

Note: Computations are based on the estimates shown in table A1.

Table A1: Maximum Likelihood Estimates for Dependent Sequential Choice Model

	SC		SM		CS		CM		MS	
Variable	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat.	Coeff.	t-stat	Coeff.	t-stat
Constant	-5.20	9.75	-1.80	3.67	-4.30	4.55	-2.56	3.46	-3.01	6.79
Black	-.79	1.43	-1.10	2.65	-.23	.25	2.10	2.29	-.70	1.38
Hispanic	-1.48	1.29	1.23	1.76	.15	.09	-22.05	.01	-.27	1.16
Duration=1 year	1.48	4.36	.04	.07	-1.35	1.88	.39	.97	1.07	8.32
2 years	1.54	5.31	.27	.65	-1.18	2.15	.25	.69	.91	6.80
3 years	2.02	8.39	.46	1.45	-.83	1.91	.18	.49	.62	4.34
4 years[a]	1.99	9.18	.60	2.36	-.46	1.26	.33	.89	.80	5.82
5 years	1.95	9.44	.70	3.37					.42	2.61
6 years	2.11	10.93	.66	3.71					.27	1.63
7 years	1.87	9.25	.55	3.36					.17	.98
8 years	1.88	9.09	.11	.64					.17	.95
9 years[a]	1.79	8.33	.30	1.81					.14	.75
Single duration					.12	2.82	-.01	.36	-.05	3.65
Cohab. duration									.11	2.23
Cohab.*Black									-.23	2.34
Cohab.*Hispanic									-.29	1.85
AFDC/1000	1.21	3.32	-.91	3.13	-.12	.26	.01	.03	-.50	1.59
AFDC/1000*Black	.73	1.10	-.18	.35					1.56	2.95
Medicaid/1000	.32	.43	-1.13	1.63	1.43	1.24	-.59	.52	-.63	.82
Medi./1000*Black	-5.49	.11	.11	2.30						
Income tax/1000	.49	1.20	-.35	1.17	-.39	.79	.17	.39	-.06	.30
Inc tax/1000*Black	-3.45	3.52	1.39	2.26	.29	.25	-2.50	1.81	.55	1.23
Fault	-.63	1.42	.50	2.12	1.74	.01	-3.23	.01	-.97	1.09
Fault*Black									2.06	1.62
No fault	.46	2.84	.02	.21	.87	3.27	-.07	.34	.25	1.83
No fault*Black	-.95	2.90	.09	.40					.25	1.04
Property	.06	.60	-.14	1.89	.15	.77	.05	.29	.04	.44
Property*Hispanic	-.06	.22	.49	2.45	-1.14	2.08	-.19	.34		
Separation/100	1.16	1.78	-.63	1.37	1.73	1.92	-1.98	2.21	-.38	.06
Separ/100*Black	-.02	1.55	.72	.98					1.73	2.01
Separ/100*Hispanic					-1.52	.73	5.56	2.64		
Baptist	-.59	2.18	.97	3.45	.30	.81	.13	.40	-.18	.95
Baptist*Black	.18	.73	-.56	2.59						
Baptist*Hispanic	1.30	1.04	-1.65	2.01	.97	.59	-4.32	.04		

Note: Continued on next page.

Table A1: Continued

Variable	SC		SM		CS		CM		MS	
	Coeff.	t-stat	Coeff.	t-stat	Coeff.	t-stat.	Coeff.	t-stat	Coeff.	t-stat
Protestant	-.35	1.66	.53	.24	-.26	.67	-.01	.04	-.23	1.17
Protestant*Black	.18	.73			1.87	3.65	1.19	2.41		
Protestnt*Hispanic	.23	.17	-2.46	2.71	-1.16	.58	21.19	.00		
Catholic	-.25	1.17	.55	2.50	.06	.17	.31	1.05	-.37	1.91
Catholic*Hispanic	.38	.34	-1.60	2.35	.80	.53	21.09	.01		
Other religion	-.62	2.51	.52	2.18	.13	.32	-.08	.23	-.29	1.36
Other rel.*Hispanic	.89	.72	-2.17	2.56	.58	.34	21.03	.01		
Live both	-.65	4.45	.14	1.24	.00		.19	1.01	-.26	2.51
Live mom	.11	.61	-.38	2.17	.43		.05	.21	.04	.32
Live mom*Black	-1.01	3.69	.36	1.79						
Mom HGC	.04	2.37	-.04	3.26	.12		.08	2.60	.04	2.22
Mom HGC*Black					.01		-.13	2.02	-.06	2.25
AFQT/100	-.36	1.95	-.19	1.42	-.14		.72	2.33	-1.19	6.48
AFQT/100*Black					-1.69		.63	.83		
Children	.06	.94	-.26	4.31	-.02		-.18	1.95	-.11	1.99
Children*Black									.22	3.02
Children*Hispanic	.20	1.37	.01	.09					.23	2.54
Adults	-.41	5.06	-.31	5.19	-.01		.02	.09	.57	5.62
Adults*Black					.73		.33	.68		
Unemp rate/100	.63	.55	-.08	.10	9.72		5.19	2.23	-.43	.34
Unemp/100*Black					-2.97		-14.98	2.40		
Same race/100	-.47	1.43	-.10	.46	.16		.23	.54	-.38	1.31
Same /100*Black	.45	.70	.91	2.05					1.66	3.28
α_s^c	1.53 (5.18)									
α_s^m	-1.23 (3.07)									
α_c^s	1.44 (2.76)									
α_c^m	0.28 (0.79)									
α_m^s	0.02 (0.21)									
Function value	-15,147.81									

[a]For cohabitation (single and married) spells, the duration variable equals one if the spell is in its fourth (ninth) year or greater.

Note: Columns labeled "t-stat" give absolute values of asymptotic t-statistics. These statistics appear in parentheses for the "alpha" variables, which are the factor loadings.