TRANSITIONS FROM COHABITATION TO MARRIAGE

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For presentation at the Annual Meeting of the Population Association of America,
Philadelphia, PA, April 1, 2005.

U S C E N S U S B U R E A U

Abstract 100 words

Using the Survey of Income and Program Participation (SIPP) 2001 Panel, we use couple-month data and event history analysis to examine the association between characteristics of cohabiting couples and likelihood of marrying or dissolving. Characteristics include basic demographics, economic characteristics, and presence and type of children. This is the first prospective longitudinal analysis of cohabitation dynamics using monthly SIPP data and explicitly controlling for right censoring. We identify the importance of the presence and type of children, and of higher education in increasing the odds of marriage, and a potential curvilinear relationship between male partner's income and likelihood of marriage.

Introduction

By any measure, cohabitation has increased steadily since the late 1960s and early 1970s (Fields 2004; Casper and Cohen 2000). In 2003, approximately 4.6 million households (4.2 percent of all households) included a householder and his or her opposite-sex unmarried partner, up from 2.9 million in 1996 (Fields 2004). Although marriages following cohabitation are more likely to end in marital disruption, much of this relationship is explained by the endogenous relationship between those couples likely to cohabit and those likely to divorce (Lillard, Brien, and Waite 1995). Lillard et al.'s (1995) study described selectivity into cohabitation and the instability in subsequent marriages for a 1972 high school class cohort (National Longitudinal Survey 1972). Cohabitation has become more normative since then and the endogeneity of this relationship should be reexamined with contemporary cohorts.

There is little argument that cohabitation is a less stable family form than marriage (Bumpass, Sweet and Cherlin 1991; Schoen and Owens 1992; Schoen and Weinick 1993; Brown 2000), and that stability is one of the important factors related to children's well being (McLanahan and Sandefur 1994; Graefe and Lichter 1999; Bumpass, Raley and Sweet 1995, Bumpass and Lu 2000). Because of the relationship between family stability and child well being, the link between the presence of children in cohabiting couples and the transition from cohabitation to marriage needs to be explored.

Given the increase in cohabitation, the importance of cohabitation as a family context for children, and the potential selectivity of those cohabiting to enter marriages at greater risk of disruption, this paper explores which cohabiting couples marry and which couples dissolve. This research contributes to the literature on transitions out of cohabitation in several ways: it introduces a nationally representative longitudinal data source for all age groups; the coding of the presence of children distinguishes whether children present in the household are the biological or adopted child of both partners, or of just one partner; it explicitly controls for right censoring (truncated data as the panel progresses); and it identifies a potential curvilinear relationship between the male partner's income and the likelihood of marriage.

Background

Since the mid-1980s, numerous sociological and demographic articles about the trends in cohabitation, characteristics of cohabitors, and the effect of cohabitation on the probability and timing of marriage have been published (Smock 2000). Many studies have looked at the factors related to the transition from cohabitation to marriage (Bumpass and Sweet 1989; Bumpass, Sweet, and Cherlin 1991; Manning and Smock 1995; Smock and Manning 1997, Schoen and Owens 1992; Rindfuss and VandenHeuvel 1990; Lillard, Brien, and Waite 1995; Brown 2000; Sassler and McNally 2003; and Lichter, Qian and Mellot 2004).

The transition from cohabitation to marriage may be related to the current levels of cohabitation, delays in marriage, the divorce rate, the presence of children in unmarried-parent households, educational attainment, increasing time spent in the labor force before marriage, and growing labor force participation for women (Teachman, Tedrow, and Crowder 2000; Bennett, Bloom, and Craig 1992). The characteristics of cohabiting couples compared with married couples suggest couples who are at different stages in their life course (e.g., age, income, tenure, marital status) (Rindfuss and VandenHeuval 1990; Seltzer 2000; Fields 2004; Manning and Lichter 1996).

Differences in the likelihood of transitioning to marriage may be related to racial differences in historical patterns of family formation (Ruggles 1994; Morgan et al. 1993). Other studies also find, net of many other variables, that White cohabitors were more likely to marry than Blacks (Manning and Smock 1995; Brien, Lillard, and Waite 1999; Lichter, Qian, and Mellot 2004). Manning and Smock (1995) find that full time employment, primarily men's employment, deters separation and increases the likelihood of marriage versus separation, but only for Whites.

Findings about the relationship between the economic characteristics of the couple and the likelihood of transitioning from cohabitation into marriage are mixed. Smock and Manning (1997) use couple level data and test the hypothesis that increases in women's earnings and egalitarian attitudes over time have led to a greater influence of female earnings on the transition from cohabitation to marriage. They found, however, that despite trends toward more egalitarianism and economic independence for women, men's economic circumstances still carry more weight in predicting a transition from cohabitation to marriage. After repairing NSFH for biases created by missing data, Sassler and McNally (2003) found a negative relationship between the log of the male partner's income and the likelihood of marriage, controlling for other couple level characteristics. Both of these studies treat income as a continuous variable and so are unable to identify any possibility of a curvilinear relationship between income and the likelihood of marriage.

In recent cross sectional data, a greater proportion of cohabiting partners have similar employment, earnings, and education than spouses (Fields 2004). Sanchez, Manning and Smock (1998) found a higher likelihood of marriage for cohabiting couples that employed sex-specialized roles. Brines and Joyner (1999) found that cohabiting partners with more similar employment and earnings were less likely to separate, but that couples in which the woman earned more than her partner were more likely to disrupt.

Education is consistently identified as an important person-level predictor as to whether cohabiting couples will marry or dissolve their partnership. During the early period of rising rates of cohabitation, it was lower socioeconomic status couples who led the way into cohabitation, counter to many assumptions about cohabitation among college students at the time (Bumpass and Sweet 1989). Since this early period, increasing numbers of people at higher socioeconomic levels have also begun to cohabit, but the prevalence of cohabitation is still high among lower income and education groups. Higher education (bachelor's degree or more) is routinely associated with higher odds of marriage from cohabitation (Smock and Manning 1997; Thornton, Axinn, and Teachman 1995; Carlson, McLanahan, and England 2004; Lichter, Qian, and Mellot 2004).

School enrollment operates differently from educational attainment. Current school enrollment has been associated with a reduced likelihood of marriage for cohabitors, and has been shown to be a risk factor for disruption (Thornton, Axinn, and Teachman 1995; Manning and Smock 1995). However not all studies have shown this relationship for enrollment. Lichter et al. (2004) found no statistical relationship between enrollment and union transitions among cohabitors.

Although much of the available data on cohabitation are cross sectional, the dynamic relationships between cohabitation and family formation cannot adequately be described using cross-sectional snapshots. Since cohabitations are often relatively short-lived, looking at the group of couples who are cohabiting at any given time will necessarily miss many people who have cohabited in the past but have already married their partner, or are no longer living with him or her.

Most of the research cited above has moved toward employing more dynamic methods to consider the process of transitioning out of cohabitation. Many studies have examined cohort-based samples (Lichter, Qian, and Mellot 2004; Brine and Joyner 1999; Brien, Lillard, and Waite 1999; Rindfuss and VandenHeuval 1990; and Astone et. al. 1999) while other have used population based surveys such as the NSFH and the National Survey of Family Growth (NSFG). Cycle 5 of the NSFG has provided important additions to our understanding of cohabitation, although it is limited to women of reproductive age in 1995 (Bramlett and Mosher 2002). Both surveys benefit from the presence of retrospective cohabitation histories. Retrospective data has the advantage of capturing more cohabitation spells than cross sectional data, although recall can also be an issue. Many studies of cohabitation dynamics have used NSFH prospectively by combining it with its longitudinal follow-up, NSFH-2 (Bumpass and Sweet 1989; Bumpass, Sweet and Cherlin 1991; Schoen and Owens 1992; Schoen and Weinick 1993; Casper and Cohen 2000; Bumpass and Raley 1995; Manning and Smock 1995; and Smock and Manning 1997; Brown 2004; and Sassler and McNally 2003). Some of the dynamic methods used in these studies include multistate life table analyses, discrete-time hazards models and proportional hazards models.

This analysis contributes to the literature on transitions from cohabitation to marriage in several key ways. First, we introduce a previously untapped resource for the study of cohabitation dynamics—the 2001 panel of the Survey of Income and Program Participation. Second, these prospective data include a sample large enough to analyze Hispanics as a distinct subgroup; both members of cohabiting couples of all ages; a monthly direct measure of cohabitation; and indicators of the presence of both coresident parents as well as the type of relationship between child and parent (biological, step, adopted). These data also limit poor measurement due to respondent recall through frequent interviews (every 4 months). Third, these models explicitly control for censoring as a competing risk, and we identify characteristics of our sample that are associated with censoring. Using discrete-time hazards models, we explore the proximate characteristics of cohabiting couples that marry, dissolve, or continue to cohabit.

Data

This analysis is the first to use 2001 Survey of Income and Program Participation (SIPP) data to examine transitions from cohabitation to marriage and to dissolution. We use waves (interviews) 1 through 9 of the SIPP 2001 survey panel, which sampled about 35,000 households and interviewed them every four months from February 2001 through May 2004.

SIPP data are uniquely qualified to supplement the existing literature on cohabitation dynamics and the transitions from cohabitation to marriage. Baughman et al. (2002) examined several panels of SIPP data and compared direct and indirect estimates of cohabitation from SIPP with those from the Current Population Survey (CPS). They found the estimates from the SIPP, particularly the direct estimates, were better than those from the CPS, and that the SIPP is a rich and untapped source of data for cohabitation analyses. Bauman (1999) used the SIPP to examine the effect of including cohabitors in the measurement of poverty. The monthly data collected by SIPP and the short recall period (4 months) enhance the quality and usefulness of SIPP household data over other surveys with longer recall periods.

The prospective nature of SIPP reduces the recall bias for short spells of living arrangements, and other omissions common in retrospective history data. Unlike cohort studies like the NLSY and PSID, and surveys like the NSFG, the SIPP is nationally representative for *all* ages. SIPP includes both men and women and is large enough to examine differences in transitions for Hispanic respondents. In addition, the SIPP provides a wealth of information on other characteristics of respondents, including detailed income and public assistance indicators. SIPP also contains more detailed information about family relationships than most surveys. The type of relationship between children and their coresident parents can be identified (biological, adopted or step), even if neither of the child's parents is the householder.

The lack of a cohabitation history in SIPP, however, limits analyses because of substantial left censoring for spells already underway at the time of the survey. The number and timing of previous cohabitation transitions that may affect current living arrangements and transitions observed over the life of the panel cannot be identified. Sample attrition is always a problem for longitudinal data collection.[1] Because we measure cohabitation by current living arrangements – not recall--and the interval between interviews is short (4 months), this data source is of high quality and is useful to gain significant insight about cohabitation dynamics.

Methods

We examine the risk of transitions from cohabitation to marriage, dissolution, and survey censoring (cases lost to follow-up) using discrete-time event history models. This method avoids proportionality assumptions and allows the use of fixed and time-varying predictors (Allison 1984).

Our dependent variable is modeled using multinomial logistic regression to generate odds of (1) censoring versus continuing to cohabit; (2) dissolving versus continuing to cohabit; (3) marrying versus continuing to cohabit. For ease of discussion, an

[1] For details about sampling loss see the Source and Accuracy statement available on the Census Bureau website at: http://www.sipp.census.gov/sipp/sourceac/S&A01_w1tow6_cross_puf.pdf

additional model was run to generate the odds of marrying versus dissolving. Multinomial models are an appropriate analysis tool for looking at transitions out of cohabitation because each option is treated as a competing risk.

Spells of cohabitation are censored if a couple leaves the survey prior to the last interview in the panel. Cohabitation is defined as the presence of a person identified as the unmarried partner of the householder. About nine percent of all cohabiting couples in the SIPP do not include the householder.[2] Only one person in the household in a given month can be identified as the unmarried partner. Dissolution occurs when the householder and their partner no longer reside together at the same address. If an individual has a subsequent spell of cohabitation, he or she is included as part of a new set of couple-month observations; serial cohabitations are treated independently. Marriage is identified when a cohabiting couple changes marital status to "married spouse present" and are married to each other. In most cases this change occurs at the seam between waves in SIPP.[3] This would be an important consideration when

[2] In wave 2 of the 2001 SIPP panel a household relationship topical module is administered that allows the identification of cohabiting couples which do not include the householder. Since this analysis includes continuous information about cohabitations present at the beginning of the panel, in addition to those forming in wave 2 and later, the additional cohabitations identified by the household relationship topical module are not used in these analyses.

[3] Since SIPP data are collected in Waves occurring every 4 months, asking about the status and changes over the last 4 months (reference period), creates a disjunction or 'seam' between waves of data when the data are examined monthly. This seam represents the transition from monthly data collected in one wave and monthly data collected in the next wave. When a status is reported for a whole wave but not reported in the previous wave, the transition will occur on the seam, the first month of the new wave.

examining duration of spells. SIPP uses spouse "pointers" which identify the person to whom the record holder is married. Continuation couple-months, or months in which there is no change in cohabitation status for the couple, form the comparison group for the above transitions.

We chose to stratify the models by sex. Although each couple-month observation includes both the male partner and female partner's information, the specific characteristics of the individual are more easily understood when the models are run separately. High correlations between men's and women's characteristics in cohabiting couples also suggest that it is preferable to conceptualize the models separately by sex. We include household characteristics (common to both men and women) in each model, and include male personal income in the women's model rather than female personal income based on previous research establishing the link between male partner's earnings and transitions out of cohabitation (Sassler and McNally 2003; Smock and Manning 1997).

In addition, we show results of the full model separately by sex and age group (15-39 years and 40 years and above). SIPP data are among the only data sources that allow the characteristics for older cohabitors to be examined. These results are shown in Table 4.

Independent Variables

All characteristics except race and Hispanic origin are included as time-varying covariates and may change monthly. At the individual level, we include basic demographic characteristics: Age (15-29 years; 30-39 years; 40-49 years; 50 years and above); Race and origin (White non-Hispanic; Black non-Hispanic; other race non-Hispanic; Hispanic); and Marital status (ever married; never married).

Also at the person level, we include the following socioeconomic indicators: Educational attainment (less than high school; high school degree; some college; bachelor's degree or higher); School enrollment (enrolled full time or part time; not enrolled); Employment (in labor force; not in labor force); personal income (less than $1,500 per month; $1,500 to $2,999 monthly; $3,000 to $4,499 monthly; $4,500 per month or more).

Since the federal family poverty level is established based on a family definition which does not include cohabiting couples, and using the household poverty level may include other adults' income beyond that of the couple, we chose not to include poverty as a covariate in the final models. We examined household poverty in earlier models, but decided to replace it with the receipt of cash benefits by anyone in the household in combination with the male partner's personal income. Male partner's personal income has been shown to have a persistent and stronger relationship with cohabitation transitions than couple income, female partner's personal income, or income differentials (Sassler and McNally 2003; Smock and Manning 1997). The receipt of cash public assistance carries additional meaning beyond that of a simple poverty indicator. Cash benefit receipt also indicates an interaction with state and federal welfare systems, a situation with may be independently related to cohabitation and the likelihood of marriage.

Lastly, at the household level, SIPP data allow the monthly identification of household composition, including the presence of a mother and father and type (biological, adopted, or step) of relationship between the child and each parent. This allows a detailed coding of the <u>presence of children</u>. Joint biological or adopted children of both partners can be differentiated from the presence of other children (biological or adopted children of only one partner, or other children in the household), as well as from households with no children under age 18. This important distinction is related to how and when the cohabitation forms – whether following prior childbearing. We expect that the effect of joint biological/adopted children on the transition to marriage will differ by the age of the youngest child. We expect that couples who have recently had a birth (those with a joint child under 1) will be more likely to marry, and those with joint children of any age will be more likely to marry than those without children.

Descriptive Characteristics

Table 1 presents the characteristics of the sample of couples cohabiting at wave 1 of the 2001 SIPP. The estimates shown in Table 1 are weighted to account for sample design and represent a national estimate of cohabiting couples in thousands.[4] In the spring of 2001, SIPP identified 4.6 million households with a householder and his or her opposite-sex unmarried partner. Most cohabiting men and women were under age 40 (68.9 percent of women and 65.1 percent of men). Half of the couples were within three years of age of each other, 50.9 percent. As expected, when the age difference exceeded three years, women were more often younger than their male partners (35.4 percent of couples). However, 13.7 percent of couples included a woman more than three years older than her partner.

This analysis is one of the first to extend the study of cohabitation transitions beyond Whites and Blacks in a nationally representative survey. Most respondents are White non-Hispanic (69.3 percent of male partners and 71.1 percent of female partners), slightly lower than their distribution in the total population over age 15 (estimates from wave 1 of the 2001 SIPP indicate that 73.0 percent of men and 72.5 percent of women 15 and over were White non-Hispanic). Black non-Hispanic males were 12.1 percent of cohabitors and 10.1 percent of the total population over 15, while Black non-Hispanic

[4] The multivariate results in Tables 2 through 4 are unweighted. The units of analyses in these models are couple-months of exposure to the risk of a transition out of cohabitation. There is no appropriate weight for this unit of analysis.

females were 10.6 percent of cohabitors and 11.8 percent of the total population of adult females. Hispanic cohabitors are also over represented among cohabitors compared with the total population of adults (14.7 percent of male cohabitors and 13.7 percent of female cohabitors are Hispanic compared with 12.2 percent and 10.9 percent of adult men and women in the population in 2001).

Partners differed in race from each other 6.2 percent of the time, and one partner was Hispanic and the other was non-Hispanic 6.9 percent of the time. Census 2000 data show 12 percent of opposite-sex partners differed in race from each other, while 6 percent of opposite-sex cohabiting couples involved one partner who was Hispanic and one who was non-Hispanic (Simmons and O'Connell 2003). The definition of race which was used in Census 2000 had more categories than that used in SIPP 2001, which would increase the percentage of interracial couples, all else being equal.

Most cohabitors are never married (56.2 percent and 55.3 percent for men and women respectively). Approximately a third of both men and women cohabitors are divorced (36.6 percent of men and 34.5 percent of women). Over half (55.2 percent) of cohabiting couples had no children in the household, 24.2 percent had a joint child in the household, 20.5 percent did not have a joint child, but lived with a child who was not the child of both partners. In comparison, Census 2000 data also showed that 57 percent of opposite-sex unmarried partner households had no children under 18 present in the household (Simmons and O'Connell 2003).

Education is an important predictor of cohabitation transitions. In our sample, 7.5 percent of men and 11.2 percent of women were enrolled either full or part time in school. Thirty-five percent or more of the partners were high school graduates (38.8 percent and 34.5 percent for men and women respectively), but 20.5 percent of male partners and 18.3 percent of female partners were not high school graduates. The percentages of male and female partners who had earned at least a bachelor's degree were similar--14.6 percent for men and 15.7 percent for women.

We code employment status as either in the labor force or not. Most cohabitors were in the labor force (88.7 percent of the men and 79.0 percent of the women). Almost all of those in the labor force were actively employed. Just 5.3 percent of male and 3.9 percent of female cohabitors were in the labor force but unemployed.

We experimented with several operationalizations of income and welfare receipt. Income is shown rather than earnings because the inclusion of older cohabitors increases the likelihood that unearned income (e.g., social security, retirement) will be relevant to understanding the economic well being of the household. We ran models including household poverty instead of income, but decided not to include this measure in the final model set. Poverty status includes the income of any additional adults present, which is likely unrelated to the choice for a cohabiting couple to marry. In general the models including poverty status showed that those in households below 200 percent of poverty were less likely to marry and more likely to dissolve. Overall, 11.5 percent of couples were in households below poverty, and 33.7 percent were in households falling below 200 percent of poverty.

We chose to use the male partner's personal monthly income and the household receipt of cash benefits rather than poverty. The median monthly male income was $2,000. For 30.0 percent of the couples the male monthly income was less than $1,500. Another 36.0 percent of couples fell into the category that included the median ($1,500 to $2,999 per month), and 26.4 percent had male monthly incomes of $3,000 or more. Cash public assistance was received in 9.7 percent of the cohabiting couple's households.

Multivariate Models

Men

Table 2 presents the first set of models for men. The basic relationships shown in Model 1 include age, race/origin, and marital status. From this model, a distinct relationship is shown by age. Older cohabitors are less likely to marry or dissolve versus continuing, and less likely to marry than dissolve. Censorship is negatively related to age, a finding that is common in studies about who is likely to leave the survey by attrition. The only significant effect for younger men is a higher odds of dissolution versus continuing (51 percent more likely).

Compared with non-Hispanic White men, all race and origin groups were more likely to censor versus continue. They were also less likely to marry versus continue than White non-Hispanics. Black non-Hispanic men were less likely to marry versus dissolve than White non-Hispanic men. Ever-married men were more likely to either dissolve or marry than to continue. Characteristics of a respondent's life course, such as a prior marriage or children, are indicators that could be related to having a perspective that encourages evaluation of the current situation. As evidenced from these results, this type of evaluation could be one of the factors relating prior marriage to a more rapid dissolution or marriage versus continuation of the current situation.

In Model 2 we add measures of the presence of children, educational attainment and school enrollment. Age effects remain largely unchanged with the exception that the youngest group of men (15-29 years old) has a 30 percent higher likelihood of marrying versus continuing, and the relationship between age and the odds of dissolving versus continuing for the oldest two groups is now non-significant. Significant effects remain indicating higher odds of censoring and lower odds of marrying among other non-Hispanics compared with White non-Hispanic men.

Ever-married men are more likely than never-married men to marry versus continuing to cohabit. Those men who have older children with their partner are less likely to marry than to either continue to cohabit or to dissolve. Those men in households without joint children, but with other children present are twice as likely to dissolve versus continue to cohabit compared with men in couples with no children present. Men in couples with other children present are also nearly half as likely to marry as to dissolve than men in couples without children.

Educational attainment has a strong relationship with the likelihood of marriage. Men with some college or more are more likely to marry than to continue to cohabit or to dissolve compared with men who have a high school degree. Men with less than a high school diploma are significantly less likely to marry than to continue to cohabit. School enrollment has no significant relationship on transitions out of cohabitation for men.

In Model 3, employment, income and cash benefits are added to the model. The relationships between age and the transitions out of cohabitation are similar to those described for Model 2. Other non-Hispanic men continue to be more likely to censor, and less likely to marry versus continuing to cohabit compared with White non-Hispanic men.

The effects of marital status, the presence of children and school enrollment were unchanged from model 2. Model 3 shows that men with some college or more were more likely to marry than to continue; they are also more likely to marry than dissolve compared with men with a high school diploma.

Employment status did not have a significant effect on transitions out of cohabitation for men. Men whose income was $3,000 to $4,499 per month were 35 percent more likely to marry than to continue to cohabit, while men whose incomes were less than $1,499 per month were more likely to dissolve versus continue compared with men whose income was between $1,500 and $2,999. Household receipt of cash public assistance was associated with a higher likelihood of dissolving versus continuing, and a lower likelihood of marrying versus continuing or dissolving.

Women

When the same sets of models are run using the female cohabitor's characteristics, similar patterns of association among the predictors and transitions out of cohabitation are found. In Model 1, which included only demographic characteristics, the pattern is similar to that found for men, with young women (15 to 29 years) being more likely to dissolve versus continue to cohabit, and older women (50 years and over) less likely to censor, dissolve or marry than continue to cohabit. As was also true for men, Black non-Hispanic women were less likely to marry versus continue to cohabit, and less likely to marry versus dissolve than White non-Hispanic women. Hispanic women were more likely to censor versus continue than White non-Hispanic women, and less likely to either dissolve or marry than continue. Ever married women were more likely to dissolve versus continue than never married women.

In the second model for women, which adds predictors for the presence of children, educational attainment, and school enrollment, the relationship between age and transitions out of cohabitation remains the same in general. The only coefficient for race and origin which is still significant is the one for Hispanic women, which shows they are more likely to censor versus continue than White non-Hispanic women, and that Hispanic women are less likely to dissolve than continue. The coefficients for ever married women are no longer significant in this model.

As we found for men, couples with older joint children have a lower likelihood of marrying versus continuing to cohabit, and a lower likelihood of marrying versus dissolving than couples without children. Couples with other children present are more than twice as likely to dissolve versus continue, and nearly half as likely to marry versus dissolve than couples without children. The main association between educational attainment and transitions out of cohabitation is for women with at least a bachelor’s degree. These women are more than twice as likely to marry versus either continuing to cohabit or dissolving. While school enrollment was not related to transitions out of cohabitation for men, women enrolled full or part time were more likely to dissolve versus continue to cohabit and less likely to marry versus dissolve than women who were not currently enrolled.

In the full model for women (Model 3), the relationship between predictors that appeared in previous models and transitions out of cohabitation remain basically unchanged. The full model also includes employment, male partner's personal income and household receipt of cash benefits. Women's employment was not related to transitions out of cohabitation. Couples in which the male partner had a monthly income of $3,000 to $4,499 were more likely to marry versus continue than couples in which the male partner made $1,500 to $2,999 monthly. Couples in households which received cash benefits were nearly half as likely to marry versus dissolve than couples in households that did not receive cash assistance. The odds of marrying and dissolving did not differ for couples in which the male partner had a monthly income of $4,500 and those in which the male partner had a monthly income of $1,500 to $2,999.

In Table 4, we show the final models for men and women stratified by age (15-39 years old and 40 years old and over). Based on the persistent findings related to age for men and women in the models shown in Tables 2 and 3, we felt it necessary to explore whether patterns of transitions out of cohabitation differ substantially for younger and older adults. The emergence of distinct patterns is expected based on a life course perspective, and is a component of cohabitation that will only become more important to understand as the baby-boom cohort continues to age.

Models 1 and 3 in Table 4 show the likelihood of marriage and dissolution for younger men and women respectively. Young Hispanic men and women are less likely to dissolve than to continue to cohabit. This is an important finding, consistent with research on cohabitation among Hispanics both in the U.S. and elsewhere (Martin 2002; Landale and Forste 1991). Additionally, other non-Hispanic young men are less likely to marry versus continuing to cohabit when compared with White non-Hispanic young men. Both models show an increased likelihood for censoring among other non-Hispanic men and women, with younger Hispanic men also being more likely to censor versus continuing to cohabit.

Being ever married is not related to transitions out of cohabitation for young men and women. For younger adults, having older children together (joint children 1 year old or older) is associated with decreased odds for marrying versus continuing to cohabit or versus dissolution. This effect is observed for both young men and women. There is no effect of having a joint child under age 1. Also, for both men and women, the presence of a child of just one partner increases the odds of dissolution versus continuing, and decreases the likelihood of marriage versus dissolution.

Young men and women with at least a bachelor's degree have higher odds of both marrying versus continuing to cohabit and of marrying versus dissolution. For young women, current school enrollment increases the likelihood of dissolution versus continuing, as well as reducing the odds for marrying versus dissolving.

For young men, being out of the labor force is marginally related ($p<0.10$) to a reduced risk of marrying versus continuing and versus dissolution. For young women, an increased likelihood of censoring was the only significant effect related to being out of the labor force. The only significant relationship observed for male partner's personal income is a higher odds of marrying versus continuing to cohabit for couples in which male monthly personal income was between $3,000 and $4,499 compared with those with monthly incomes from $1,500 to $2,999. This effect is observed for both men and women.

Public assistance receipt significantly decreases the likelihood of marriage versus continuing for men. Women show the same effect. Both men and women living in households receiving cash benefits show a decrease in the likelihood of marriage versus dissolution. For younger women, living in a household that received cash public assistance reduced the likelihood of censoring.

Among older men and women (models 2 and 4 in Table 4) race and origin fail to attain significance. Previous marriage increases the odds of dissolving versus continuing for women 40 years and over. For men 40 and over, previous marriage was only significantly related to a lower risk of censoring.

Due to the absence of any women 40 and over with infants, "couple has infant together" and "couple has older children together" were collapsed to reflect couples with any joint children in model 4. For both men and women 40 and over, the presence of a child who was not a joint child increased the odds of both marriage and dissolution over continuing to cohabit. This reflects the fact that these couples are at a different point in their life course than younger couples, for whom the relationship between the presence of other children and the likelihood of marriage was not significant. The presence of children from a previous partnership/marriage is likely one of the more significant catalysts for making decisions about the current cohabitation and transitioning either to marriage or dissolution.

Among older men, the attainment of a bachelor's degree or higher is only significant for reducing the likelihood of censoring. For older men, having some college education but no bachelor's degree decreases the odds of dissolution versus continuing, and increases the odds of marriage versus dissolution. Also, among older men with less than a high school degree a lower likelihood of marriage versus continuing is marginally significant.

For women age 40 and over, the only significant relationship for educational attainment is an increased odds of marriage versus continuing among those women with a bachelor's degree or more.

Men 40 years and over with personal incomes of less than $1,500 per month are more likely to dissolve versus continuing to cohabit. Men with incomes over $4,500 per month are more likely to censor than to continue to cohabit. Women 40 years and over show no significant relationship between their male partner's personal monthly income and cohabitation transitions. In the models for the older adults, household receipt of public assistance is only significant for males and increases the likelihood of dissolution versus continuing to cohabit.

Discussion & Conclusions

This analysis contributes to the literature on the dynamics of transitions from cohabitation to marriage by using monthly prospective data that show the contemporaneous characteristics of the couple making the transition.

Overall we show a strong and persistent relationship between high educational attainment, presence of children who are not the joint children of both partners, female partner's school enrollment, male partner's personal income, age, and household receipt of cash assistance with transitions out of cohabitation.

Having at least a bachelor's degree is related to a higher likelihood of marriage, and is one of the strongest predictors in the models. This confirms the findings of prior research (Smock and Manning 1997; Thornton, Axinn, and Teachman 1995; Carlson, McLanahan, and England 2004; Lichter, Qian, and Mellot 2004). Since education is a good indicator of socioeconomic status, these results suggest that those who can afford to marry, or have a higher income potential, are more likely to do so. These results also suggest that a bachelor's degree may represent a threshold level of education with respect to the transition from cohabitation to marriage. For young adults, those who have earned a bachelor's degree may be more likely to consider their education sufficiently completed to enter a marriage. School enrollment is only important for younger women as a predictor of dissolution versus marrying or continuing to cohabit. These results confirm those found in prior research (Thornton, Axinn and Teachman 1995; Manning and Smock 1995). This result follows the finding for educational attainment, also suggesting that cohabitors who have not yet completed their schooling are less likely to marry and more likely to dissolve.

Prior research has connected child well being with transitions in family structure and nonmarital parenting (McLanahan and Sandefur 1994; Graefe and Lichter 1999;

Bumpass, Raley and Sweet 1995, Bumpass and Lu 2000). Forty-one percent of cohabiting couples (Fields 2004) have children under 18 living in the household, based on CPS data. In the sample used in this paper, 45 percent of the couples cohabiting at Wave 1 have children present; 24 percent of the Wave 1 couples have children who are the biological or adopted child of both partners. The detailed coding of the type of relationship between children and parents employed in this analysis has been shown to be of paramount importance for understanding transitions out of cohabitation.

Contrary to our initial expectation that a joint child would be associated with a higher likelihood of marriage, in no model was the presence of a joint child under 1 year associated with transitions out of cohabitation. It seems likely that couples who would marry to legitimate their child's birth had already done so before the birth of the child, or would perhaps report themselves as married. The presence of older joint children was associated with a lower likelihood of marriage versus continuing or dissolving the union. This effect is consistent with a couple that has cohabited for a longer period of time and may have less pressure to marry or a lower intention to marry.

In all models, couples with children of just one partner were more than twice as likely than couples with no children to dissolve versus continue to cohabit and were about half as likely to marry versus dissolve. Only in models for older adults did the presence of children of just one partner have a positive relationship with the likelihood of marriage versus continuing to cohabit. This may indicate couples for which the presence of the children of only one partner acts as a catalyst for a decision to either marry or dissolve. Except for older couples, the presence of children of just one partner is associated with increased instability of the cohabitation. Using SIPP data, which can be used to determine whether the household contains joint children of the couple or children of just one partner, we show that the presence of children of previous partnerships is an extremely important predictor of dissolution distinct from the presence of joint children of the couple. The only predictor in the models that is roughly similar in magnitude is that for men or women who have at least a bachelor's degree.

As in previous research, we have included the male partner's income in these analyses. However, unlike other studies, we did not code income as a continuous variable, instead categorizing income with median income included in the reference group. This categorization enables us to see an unusual pattern of association between income and transitions out of cohabitation. Those couples in which the male partner's income is \$3,000 to \$4,499 per month are more likely to marry versus continue to cohabit than couples whose income falls into the group which includes the median (\$1,500 to \$2,999 monthly). Our expectation that income in the highest category would also be positively associated with marriage was not met. While this could be related to an insufficient sample in the highest income group, it may suggest a curvilinear relationship such that those couples at the highest income levels may be content to continue cohabiting. The lack of a positive association between the highest income group and marriage may also indicate a potential interaction between income, age, and the odds of marriage, since older couples may have retirement or other estate assets which make marriage a more complicated option.

Since the association between poverty and transitions out of cohabitation (in models not shown) was similar to results shown for cash public assistance, and cash assistance may be a better indicator of interaction with public state and federal welfare systems, we chose to use cash public assistance in the models over poverty status. In pre-welfare reform assistance systems, this interaction could be expected to reduce the likelihood of marriage due to eligibility requirements of the AFDC program. We find that receipt of cash benefits is associated with a lower likelihood of marriage versus both dissolving and continuing to cohabit. Although the full model controls for male partner’s personal income, it is impossible to distinguish the differential effect of poverty from welfare system interaction, both of which may be measured in the cash benefits variable.

Since all models showed a strong and persistent relationship between age and transitions out of cohabitation, we stratified the final models, running them separately for younger (15-39 years) and older (40 and over) adults. The results of these analyses point to different processes at work for these two groups, and suggest that the characteristics related to transitions out of cohabitations occurring at different points in the life course may differ. In sum, our results for older adults suggest a pattern in which continuing to cohabit is as likely as a transition out of cohabitation, except in cases where children are present in the household. As the baby-boom cohort continues to age, further research exploring cohabitation dynamics for older adults will be necessary to clarify the ways in which the dynamics of cohabitation differ by age.

Unlike previous research, this study explicitly controls for the competing risk of censoring among cohabiting couples. Being 50 years and over is consistently associated with a lower risk of censoring versus continuing to cohabit. Men with at least a bachelor's degree also have a lower risk of censoring. Groups with a higher risk of censoring than continuing include: young Hispanic and young other non-Hispanic men, older women with children of just one partner present. While these findings point to some potential differences between those who leave the sample by attrition and those who do not, on the whole the prospective design and short recall period of the SIPP seem to limit bias associated with censoring of cohabitation spells. Cohabitation spell length and prior spells of cohabitation are unmeasured in this analysis. Left censoring and information about previous cohabitations are significant issues that may affect the association between couple's characteristics and transitions out of cohabitation, which could be overcome by the addition of a cohabitation history to the SIPP.

This research introduces a rich new data source for the study of cohabitation dynamics. We identify the importance of the presence and type of children in the household, the continued importance of higher education in increasing the odds of marriage, and the potential for the curvilinear relationship between income and the likelihood of marriage. Future research in this area should further explore these relationships and the differences in the dynamics of cohabitation for older versus younger cohabitors.

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Table 1. Characteristics of Men, Women, and Couples Cohabiting at Wave 1

(Numbers in thousands.)	Men		Women		Couple	
	Number	Percent	Number	Percent	Number	Percent
Total	4,550	100.0	4,550	100.0	4,550	100.0
Outcome						
Censor	X	X	X	X	1,413	31.1
Dissolve	X	X	X	X	760	16.7
Marry	X	X	X	X	888	19.5
Continue	X	X	X	X	1,489	32.7
Age						
15 to 29 years old	1,605	35.3	1,965	43.2	X	X
30 to 39 years old	1,354	29.8	1,170	25.7	X	X
40 to 49 years old	852	18.7	837	18.4	X	X
50 years and over	738	16.2	578	12.7	X	X
Age gap						
Woman is more than 3 years younger	X	X	X	X	1,608	35.4
Within 3 years	X	X	X	X	2,317	50.9
Man is more than 3 years younger	X	X	X	X	625	13.7
Race and origin						
White non-Hispanic	3,154	69.3	3,236	71.1	X	X
Black non-Hispanic	550	12.1	482	10.6	X	X
Other non-Hispanic	175	3.9	210	4.6	X	X
Hispanic (of any race)	671	14.7	622	13.7	X	X
Couple is mixed race	X	X	X	X	282	6.2
One partner Hispanic, other non-Hispanic	X	X	X	X	313	6.9
Marital status						
Ever married	1,995	43.9	2,035	44.7	X	X
Divorced	1,664	36.6	1,568	34.5	X	X
Never married	2,555	56.2	2,515	55.3	X	X
Educational attainment						
Less than high school graduate	930	20.5	833	18.3	X	X
High school graduate	1,766	38.8	1,569	34.5	X	X
Some college	1,188	26.1	1,434	31.5	X	X
Bachelor's degree or more	666	14.6	713	15.7	X	X
Enrollment						
Is enrolled in school full or part time	341	7.5	509	11.2	X	X
Is not enrolled	4,210	92.5	4,041	88.8	X	X
Presence of children						
Couple has infant together	X	X	X	X	259	5.7
Couple has older child together	X	X	X	X	843	18.5
Children are present	X	X	X	X	935	20.5
No children present	X	X	X	X	2,513	55.2
Employment						
In labor force	4,038	88.7	3,593	79.0	X	X
Employed	3,799	83.5	3,414	75.0	X	X
Unemployed	239	5.3	179	3.9	X	X
Not in labor force	512	11.3	957	21.0	X	X
Personal Income						
$1 to $1499/month	1,367	30.0	1,984	43.6	X	X
$1500 to 2999/month	1,639	36.0	1,306	28.7	X	X
$3000 to 4499/month	747	16.4	427	9.4	X	X
$4500 plus/month	457	10.0	264	5.8	X	X
Median monthly personal income	$2,000	X	$1,500	X	X	X
Public Assistance						
Household receives cash benefits	X	X	X	X	442	9.7
Poverty						
Below poverty level	X	X	X	X	524	11.5
100 to 199 percent of poverty	X	X	X	X	965	21.2
200 percent or more of poverty	X	X	X	X	3,061	67.3
Median monthly household income	X	X	X	X	$3,430	X

X - Not applicable.

Source: U.S. Census Bureau, Survey of Income and Program Participation, 2001 panel, Wave 1

Table 2. Multinomial Logistic Regression Models of Marriage, Dissolution and Censoring Among Cohabiting Couples: Men

	Model 1								Model 2								Model 3							
	Censor vs. continue		Dissolve vs. continue		Marry vs. continue		Marry vs. dissolve		Censor vs. continue		Dissolve vs. continue		Marry vs. continue		Marry vs. dissolve		Censor vs. continue		Dissolve vs. continue		Marry vs. continue		Marry vs. dissolve	
Characteristic	Odds Ratio		Odds Ratio		Odds Ratio		Odds Ratio		Odds Ratio		Odds Ratio		Odds Ratio		Odds Ratio		Odds Ratio		Odds Ratio		Odds Ratio		Odds Ratio	
Age																								
15 to 29 years old	0.99	n.s.	1.51	***	1.21	n.s.	0.79	n.s.	0.95	n.s.	1.46	**	1.30	*	0.89	n.s.	0.95	n.s.	1.48	***	1.30	*	0.88	n.s.
30 to 39 years old (R)																								
40 to 49 years old	0.79	*	0.78	+	0.51	***	0.66	*	0.79	*	0.83	n.s.	0.54	***	0.66	*	0.79	*	0.83	n.s.	0.54	***	0.66	*
50 years and over	0.49	***	0.63	**	0.36	***	0.56	*	0.51	***	0.80	n.s.	0.36	***	0.44	***	0.53	***	0.78	n.s.	0.39	***	0.50	**
Race and origin																								
White non-Hispanic (R)																								
Black non-Hispanic	1.19	+	1.16	n.s.	0.65	*	0.56	**	1.17	n.s.	1.08	n.s.	0.74	n.s.	0.69	+	1.19	n.s.	1.03	n.s.	0.79	n.s.	0.76	n.s.
Other non-Hispanic	1.37	*	0.84	n.s.	0.48	*	0.57	n.s.	1.41	*	0.84	n.s.	0.45	*	0.54	n.s.	1.42	*	0.82	n.s.	0.46	*	0.56	n.s.
Hispanic (of any race)	1.18	+	0.83	n.s.	0.67	*	0.81	n.s.	1.16	n.s.	0.79	n.s.	0.91	n.s.	1.15	n.s.	1.16	n.s.	0.78	+	0.93	n.s.	1.19	n.s.
Marital status																								
Ever married	1.00	n.s.	1.32	*	1.29	*	0.98	n.s.	0.94	n.s.	1.14	n.s.	1.33	*	1.17	n.s.	0.94	n.s.	1.14	n.s.	1.31	*	1.14	n.s.
Never married (R)																								
Presence of Children																								
Couple has infant together									1.05	n.s.	1.22	n.s.	1.11	n.s.	0.91	n.s.	1.06	n.s.	1.20	n.s.	1.15	n.s.	0.96	n.s.
Couple has older child together									0.92	n.s.	1.21	n.s.	0.66	*	0.55	**	0.92	n.s.	1.21	n.s.	0.66	*	0.55	**
Other children are present									1.16	+	2.19	***	1.16	n.s.	0.53	***	1.16	+	2.17	***	1.17	n.s.	0.54	***
No children present (R)																								
Educational attainment																								
Less than high school graduate									1.02	n.s.	0.91	n.s.	0.66	*	0.72	n.s.	1.03	n.s.	0.86	n.s.	0.70	*	0.82	n.s.
High school graduate (R)																								
Some college									0.97	n.s.	0.86	n.s.	1.39	**	1.61	**	0.96	n.s.	0.88	n.s.	1.32	*	1.51	*
Bachelor's degree or more									0.77	*	0.95	n.s.	1.89	***	2.00	***	0.75	*	0.97	n.s.	1.70	***	1.75	**
Enrollment																								
Is enrolled in school full or part time									1.11	n.s.	1.28	n.s.	0.88	n.s.	0.69	n.s.	1.11	n.s.	1.22	n.s.	0.95	n.s.	0.77	n.s.
Is not enrolled (R)																								
Employment																								
In labor force (R)																								
Not in labor force																	0.90	n.s.	1.00	n.s.	0.82	n.s.	0.82	n.s.
Personal Income																								
Less than $1499/month																	1.09	n.s.	1.23	+	0.96	n.s.	0.78	n.s.
$1500 to 2999/month (R)																								
$3000 to 4499/month																	1.13	n.s.	1.00	n.s.	1.35	*	1.34	n.s.
$4500 plus/month																	1.09	n.s.	1.14	n.s.	1.19	n.s.	1.04	n.s.
Public Assistance																								
Household receives cash benefits																	1.09	n.s.	1.30	+	0.67	+	0.52	*
Intercept	-3.76		-4.54		-5.64		-1.10		-3.77		-4.18		-5.46		-1.28		-3.74		-4.08		-5.47		-1.39	
Catmod Likelihood ratio (chi-square)	63.10								1321.53								4172.06							
Pr >Chi-square	0.7638								0.9972								1.0000							
N=	39023								39023								39023							

+=p<.10 *=p<.05 **=p<.01 ***=p<.001
(R) Reference (comparison) group omitted from regression.
Source: U.S. Census Bureau Survey of Income and Program Participation, 2001 panel

Table 3. Multinomial Logistic Regression Models of Marriage, Dissolution and Censoring Among Cohabiting Couples: Women

	Model 1				Model 2				Model 3			
	Censor vs. continue	Dissolve vs. continue	Marry vs. continue	Marry vs. dissolve	Censor vs. continue	Dissolve vs. continue	Marry vs. continue	Marry vs. dissolve	Censor vs. continue	Dissolve vs. continue	Marry vs. continue	Marry vs. dissolve
Characteristic	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio
Age												
15 to 29 years old	1.09 n.s.	1.44 **	1.25 +	0.87 n.s.	1.08 n.s.	1.41 **	1.33 *	0.95 n.s.	1.07 n.s.	1.42 **	1.32 *	0.93 n.s.
30 to 39 years old (R)												
40 to 49 years old	0.91 n.s.	0.77 +	0.59 **	0.76 n.s.	0.91 n.s.	0.90 n.s.	0.63 **	0.70 n.s.	0.91 n.s.	0.89 n.s.	0.64 **	0.72 n.s.
50 years and over	0.49 ***	0.55 ***	0.38 ***	0.69 n.s.	0.50 ***	0.78 n.s.	0.40 ***	0.51 *	0.48 ***	0.75 n.s.	0.41 ***	0.55 *
Race and origin												
White non-Hispanic (R)												
Black non-Hispanic	1.15 n.s.	1.15 n.s.	0.66 *	0.57 *	1.14 n.s.	1.05 n.s.	0.79 n.s.	0.75 n.s.	1.15 n.s.	1.01 n.s.	0.84 n.s.	0.83 n.s.
Other non-Hispanic	1.15 n.s.	0.80 n.s.	0.71 n.s.	0.89 n.s.	1.18 n.s.	0.80 n.s.	0.67 n.s.	0.85 n.s.	1.20 n.s.	0.77 n.s.	0.69 n.s.	0.89 n.s.
Hispanic (of any race)	1.27 *	0.71 *	0.59 **	0.83 n.s.	1.22 *	0.67 **	0.79 n.s.	1.19 n.s.	1.22 +	0.66 **	0.80 n.s.	1.21 n.s.
Marital status												
Ever married	1.07 n.s.	1.28 *	1.17 n.s.	0.92 n.s.	1.02 n.s.	1.05 n.s.	1.22 n.s.	1.16 n.s.	1.01 n.s.	1.06 n.s.	1.20 n.s.	1.13 n.s.
Never married (R)												
Presence of Children												
Couple has infant together					1.06 n.s.	1.21 n.s.	1.11 n.s.	0.92 n.s.	1.04 n.s.	1.16 n.s.	1.20 n.s.	1.03 n.s.
Couple has older child together					0.91 n.s.	1.17 n.s.	0.67 *	0.57 **	0.90 n.s.	1.17 n.s.	0.69 *	0.59 *
Other children are present					1.13 n.s.	2.17 ***	1.17 n.s.	0.54 ***	1.13 n.s.	2.15 ***	1.19 n.s.	0.56 ***
No children present (R)												
Educational attainment												
Less than high school graduate					1.12 n.s.	1.11 n.s.	0.70 +	0.63 *	1.12 n.s.	1.03 n.s.	0.78 n.s.	0.76 n.s.
High school graduate (R)												
Some college					0.97 n.s.	0.96 n.s.	1.25 +	1.30 n.s.	0.98 n.s.	0.97 n.s.	1.22 n.s.	1.25 n.s.
Bachelor's degree or more					0.91 n.s.	0.99 n.s.	2.11 ***	2.14 ***	0.92 n.s.	1.03 n.s.	1.94 ***	1.89 **
Enrollment												
Is enrolled in school full or part time					0.94 n.s.	1.37 *	0.85 n.s.	0.62 *	0.93 n.s.	1.33 *	0.87 n.s.	0.65 *
Is not enrolled (R)												
Employment												
In labor force (R)												
Not in labor force									1.14 n.s.	1.09 n.s.	0.89 n.s.	0.82 n.s.
Personal Income												
Less than $1499/month									1.07 n.s.	1.23 +	0.94 n.s.	0.76 +
$1500 to 2999/month (R)												
$3000 to 4499/month									1.14 n.s.	0.99 n.s.	1.35 *	1.37 n.s.
$4500 plus/month									1.04 n.s.	1.11 n.s.	1.18 n.s.	1.06 n.s.
Public Assistance												
Household receives cash benefits									1.04 n.s.	1.24 n.s.	0.67 +	0.54 *
Intercept	-3.79	-4.74	-5.50	-0.76	-3.80	-4.21	-5.29	-1.07	-3.78	-4.15	-5.23	-1.09
Catmod Likelihood ratio (chi-square)	99.96				1306.76				3803.35			
Pr >Chi-square	0.0088				0.8802				1.0000			
N=	39023				39023				39023			

+=p<.10 *=p<.05 **=p<.01 ***=p<.001
(R) Reference (comparison) group omitted from regression.
Source: U.S. Census Bureau Survey of Income and Program Participation, 2001 panel

Table 4. Multinomial Logistic Regression Models of Marriage, Dissolution and Censoring Among Cohabiting Couples by Age

	MEN								WOMEN							
	Age 15 to 39				Age 40 and over				Age 15 to 39				Age 40 and over			
	Model 1				Model 2				Model 3				Model 4			
	Censor vs. continue	Dissolve vs. continue	Marry vs. continue	Marry vs. dissolve	Censor vs. continue	Dissolve vs. continue	Marry vs. continue	Marry vs. dissolve	Censor vs. continue	Dissolve vs. continue	Marry vs. continue	Marry vs. dissolve	Censor vs. continue	Dissolve vs. continue	Marry vs. continue	Marry vs. dissolve
Characteristic	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio	Odds Ratio
Race and Origin																
White non-Hispanic (R)																
Black non-Hispanic	1.24 n.s.	1.04 n.s.	0.80 n.s.	0.77 n.s.	1.03 n.s.	0.93 n.s.	0.70 n.s.	0.75 n.s.	1.13 n.s.	1.06 n.s.	0.93 n.s.	0.87 n.s.	1.20 n.s.	0.81 n.s.	0.54 n.s.	0.66 n.s.
Other non-Hispanic	1.52 *	0.87 n.s.	0.45 *	0.52 n.s.	1.20 n.s.	0.68 n.s.	0.54 n.s.	0.80 n.s.	1.44 *	0.65 n.s.	0.59 n.s.	0.90 n.s.	0.75 n.s.	1.03 n.s.	0.87 n.s.	0.84 n.s.
Hispanic (of any race)	1.25 *	0.71 *	0.93 n.s.	1.32 n.s.	0.88 n.s.	0.92 n.s.	0.89 n.s.	0.97 n.s.	1.19 n.s.	0.54 **	0.75 n.s.	1.39 n.s.	1.32 n.s.	1.31 n.s.	1.05 n.s.	0.80 n.s.
Marital Status																
Ever married	1.02 n.s.	0.90 n.s.	1.15 n.s.	1.27 n.s.	0.72 *	1.42 n.s.	1.42 n.s.	1.00 n.s.	0.96 n.s.	0.85 n.s.	1.13 n.s.	1.33 n.s.	1.09 n.s.	1.60 +	0.93 n.s.	0.58 n.s.
Never married (R)																
Presence of Children																
Couple has infant together[1]	1.06 n.s.	1.26 n.s.	1.10 n.s.	0.87 n.s.	1.04 n.s.	1.42 n.s.	1.99 n.s.	1.40 n.s.	1.05 n.s.	1.28 n.s.	1.20 n.s.	0.94 n.s.				
Couple has older child together	0.95 n.s.	1.28 n.s.	0.61 **	0.47 **	0.92 n.s.	0.96 n.s.	0.96 n.s.	1.00 n.s.	0.95 n.s.	1.22 n.s.	0.63 *	0.51 **	0.88 n.s.	0.82 n.s.	0.93 n.s.	1.13 n.s.
Other children are present	1.18 n.s.	2.07 ***	1.04 n.s.	0.50 ***	1.24 n.s.	2.58 ***	1.89 **	0.73 n.s.	1.12 n.s.	2.14 ***	1.06 n.s.	0.50 ***	1.43 *	2.48 ***	1.84 *	0.74 n.s.
No children present (R)																
Educational Attainment																
Less than high school graduate	1.04 n.s.	0.92 n.s.	0.79 n.s.	0.86 n.s.	1.00 n.s.	0.76 n.s.	0.49 +	0.65 n.s.	1.17 n.s.	1.12 n.s.	0.78 n.s.	0.70 n.s.	1.02 n.s.	0.78 n.s.	0.76 n.s.	0.97 n.s.
High school graduate (R)																
Some college	0.94 n.s.	0.93 n.s.	1.28 +	1.38 n.s.	1.02 n.s.	0.68 +	1.33 n.s.	1.96 *	0.97 n.s.	1.01 n.s.	1.19 n.s.	1.18 n.s.	0.94 n.s.	0.84 n.s.	1.25 n.s.	1.49 n.s.
Bachelor's degree or more	0.80 n.s.	0.88 n.s.	1.63 **	1.85 *	0.63 *	1.02 n.s.	1.52 n.s.	1.49 n.s.	0.95 n.s.	0.99 n.s.	1.82 ***	1.85 **	0.77 n.s.	1.06 n.s.	2.09 *	1.96 n.s.
Enrollment																
Is enrolled in school full or part time	1.18 n.s.	1.34 n.s.	1.00 n.s.	0.75 n.s.	0.66 n.s.	0.82 n.s.	0.98 n.s.	1.20 n.s.	0.95 n.s.	1.39 *	0.88 n.s.	0.63 *	0.82 n.s.	1.53 n.s.	1.31 n.s.	0.86 n.s.
Is not enrolled (R)																
Employment																
In labor force (R)																
Not in labor force	0.97 n.s.	1.06 n.s.	0.57 +	0.54 +	0.73 +	0.88 n.s.	1.10 n.s.	1.26 n.s.	1.23 *	1.06 n.s.	0.92 n.s.	0.86 n.s.	0.83 n.s.	1.26 n.s.	0.87 n.s.	0.69 n.s.
Personal Income																
Less than $1499/month	1.04 n.s.	1.16 n.s.	1.00 n.s.	0.86 n.s.	1.23 n.s.	1.53 *	0.94 n.s.	0.61 n.s.	1.06 n.s.	1.16 n.s.	0.92 n.s.	0.79 n.s.	1.08 n.s.	1.36 n.s.	0.99 n.s.	0.73 n.s.
$1500 to 2999/month (R)																
$3000 to 4499/month	1.13 n.s.	1.00 n.s.	1.30 +	1.30 n.s.	1.16 n.s.	1.01 n.s.	1.45 n.s.	1.44 n.s.	1.14 n.s.	1.05 n.s.	1.28 +	1.22 n.s.	1.10 n.s.	0.79 n.s.	1.48 n.s.	1.88 n.s.
$4500 plus/month	0.82 n.s.	1.09 n.s.	1.14 n.s.	1.05 n.s.	1.63 **	1.18 n.s.	1.24 n.s.	1.06 n.s.	0.91 n.s.	1.13 n.s.	1.24 n.s.	1.10 n.s.	1.35 n.s.	1.06 n.s.	0.93 n.s.	0.88 n.s.
Public Assistance																
Household receives cash benefits	0.92 n.s.	1.01 n.s.	0.52 *	0.51 +	0.94 n.s.	1.61 *	0.89 n.s.	0.56 n.s.	0.68 *	1.12 n.s.	0.58 +	0.52 +	1.24 n.s.	1.31 n.s.	0.74 n.s.	0.57 n.s.
Intercept	-3.26	-3.93	-5.04	-1.11	-4.06	-4.60	-4.99	-0.39	-3.38	-4.01	-4.73	-0.72	-3.90	-4.41	-4.86	-0.45
Catmod Likelihood ratio (chi-square)	1850.24				1061.43				2279.22				1040.30			
Pr >Chi-square	1.0000				1.0000				1.0000				1.0000			
N=	23242				15781				25292				13731			

1 For Model 4 for women age 40 and over, the categories "couple has infant together" and " couple has older children together" are combined.

+=p<.10 *=p<.05 **=p<.01 ***=p<.001

(R) Reference (comparison) group omitted from regression.

Source: U.S. Census Bureau Survey of Income and Program Participation, 2001 panel