Reassessing the Consequence of Nonmarital Childbearing for First Marriage Formation

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EXTENDED ABSTRACT

Family change in the United States has called much attention to the formation of first marriages among unmarried mothers. Over the last half century, there have been dramatic increases in nonmarital childbearing. By the 1990s, about half of all children were raised at some point in single-parent families, and about 40 percent of them were born to unmarried mothers (Bumpass and Raley 1995; Ventura and Bachrach 2000). Witnessing such a profound change in American families, researchers and policy makers have sought to address the implications of nonmarital childbearing for the lives of unmarried couples and their children. Because nonmarital births occur disproportionately among disadvantaged women, how nonmarital childbearing is intertwined with family background and socioeconomic wellbeing has been a core subject of this line of research (Wu and Wolfe 2001). Marriage formation is of substantive interest, as marriage is found to be a positive correlate of economic viability, health, and child development (Amato 2007; Waite 1995). Indeed, marriage promotion programs targeted on unmarried couples—e.g., the Healthy Marriage Initiative—have been proposed and implemented as an effective welfare policy (Dion 2005).

In this study, we use data from the National Longitudinal Study of Adolescent Health (Add Health) and employ propensity score weighting models to examine the relationship between nonmarital childbearing and the formation of first marriages. While prior research and policy discussion have focused mostly on marriage effects, they have yet to provide insight into unmarried mothers' marital prospects. Understanding the relationship between nonmarital childbearing and marriage is critical because it informs clarification of potential marriage effects among unmarried mothers (Lichter, Graefe, and Brown 2003). If nonmarital childbearing is, *ceteris paribus*, a significant barrier to forming a marital union, reducing nonmarital birth may be a priority from a family policy perspective. In contrast, if a negative association between nonmarital childbearing and marriage formation is driven by common selection factors, improving socioeconomic conditions of disadvantaged women as a whole may be complementary to marriage policy.

Various economic and sociological theories of family-building behaviors have provided a guide for disentangling the relationship between nonmarital childbearing and first marriage formation (Becker 1991; Geronimus and Korenman 1992; Oppenheimer 1988). One view maintains that the presence of a child tends to decrease marital prospects for unmarried mothers not only because it imposes constraints and time limitations on those who engage in marriage markets but also because it presents untenable burdens to potential spouses. The other view holds that nonmarital birth *per se* is not likely a major cause of lower marriage rates among unmarried mothers given their disadvantaged background and the scarcity of "marriageable" men in lower SES marriage markets, implying nonmarital childbearing as an alternative family-building strategy to marriage. The effect of nonmarital childbearing on first marriage formation, therefore, still remains an empirical question.

In examining this issue, our study is concerned particularly with the roles of 1) selection on observed and unobserved characteristics and 2) other facets of family change associated with the prevalence of nonmarital childbearing. Several influential studies have already documented that women who have nonmarital births are less likely than single, childless women to enter marriage, and argue that this result is insensitive to selection bias (Bennett, Bloom, and Miller 1995; Graefe and Lichter 2002; Lichter and Graefe 2001). Models that past studies utilize to account for selection bias include within-family fixed-effects models and natural experiments. Within-family fixed-effects models compare sisters, one of whom had a nonmarital birth and the other did not, such that unobserved heterogeneity in family environment is effectively controlled for. Natural experiments compare women with a nonmarital birth with women who became pregnant outside of marriage but miscarried, under the assumption that because miscarriage likely occurs randomly, these two groups are similar in terms of observed and unobserved characteristics. While informative, the findings from these studies need to be crossvalidated with alternative modeling strategies to fully address the issue of selection bias.

In addition, the context of nonmarital fertility has changed in tandem with other changes in union and family formation. Although the majority of births to teenagers are nonmarital, the highest rates of nonmarital fertility are among women aged 20-24, followed by women aged 25-29 and women aged 18-19 (Martin et al. 2009); Cohabiting births have accounted for an increasing percentage of nonmarital fertility (over 40 percent); And racial and ethnic differences have complicated the relationship between nonmarital childbearing and marital formation (Graefe and Lichter 2002; Manning and Smock 1995; Wu, Bumpass, and Musick 2001). Without taking these aspects of family change into account, examinations of the effect of nonmarital childbearing on first marriage formation would be limited.

Given these concerns, this study extends previous literature in three important ways. First, we adopt a propensity score weighting framework for event history analysis to address selection bias (Hernán, Brumback, and Robins 2000; Robins 1999). As detailed below, the basic idea of this framework is to make the treated group (women who had a nonmarital birth at time *t*) and the control group (women who did not at time *t*) similar in terms of observed covariates through inverse-probability-of-treatment weighting (IPTW) estimators. To do so, we employ a logistic model to calculate the conditional probabilities of giving a nonmarital birth at time *t* as propensity scores, *ps*, and weight each woman by the inverse of the propensity score. Women in the treated group at time *t* are given a weight of 1/*ps*, thereby assigning those with the higher propensity scores less weight and those with the lower propensity scores more weight. Meanwhile, women in the control group at time t are given a weight of $1/(1 - ps)$, thereby assigning those with the higher propensity scores more weight and those with the lower propensity scores less weight. In this way, the propensity score weighting method attempts to achieve an approximate randomization of treatment assignment. The weights can be understood as the number of copies of each observation that creates a pseudo-population in which both time-constant and time-varying covariates do not predict the probability of giving a nonmarital birth (Hernán, Brumback, and Robins 2000). Any difference in first marriage formation, as a result, can be attributable to the effect of nonmarital childbearing.

This framework has a clear advantage over other models commonly used for event history analysis with respect to accounting for confounding by time-varying covariates. Previous research tends either to include time-varying covariates at time t–1 or to simply exclude them in the models. This strategy is problematic because it ignores the fact that some of the time-varying covariates both affect and are affected by treatment assignment, and obscures dynamic relationships between treatment and these time-varying covariates. In the literature, educational attainment and cohabitation status have been identified as such time-varying covariates (Aassve 2003; Brien, Lillard, and Waite 1999). We therefore introduce these two time-varying factors in our weighting models. However, the propensity score weighting framework holds the same assumption as conventional regression models that treatment assignment is independent of the outcome of interest conditional on observed covariates. Because the concern about selection bias in the association between nonmarital childbearing and first marriage formation is often focused on unobserved heterogeneity, we consider a sensitivity analysis for propensity score weighting models developed by Robins and colleagues in a case where we find a significant association (Robins 1999; Brumback et al. 2004).

Second, this study pays close attention to the roles that age, relationship status, and race/ethnicity play in estimating the effect of nonmarital childbearing and first marriage formation. We reestimate our propensity score weighting models in the following way: 1) to examine whether unmarried teen mothers differ from unmarried non-teen mothers by entry into first marriage, we compare the full sample with a subsample that excludes unmarried teen mothers; 2) to investigate whether unmarried cohabiting mothers differ from unmarried single mothers by forming a first marriage, we compare the full sample with a subsample that excludes unmarried single mothers; and 3) to address racial/ethnic differences, we fit our weighting models to race/ethnicity-specific samples. Given data limitations, we focus on non-Hispanic whites, non-Hispanic African Americans, and Hispanics.

Third, we use the Add Health data to examine whether and how nonmarital first birth exerts influence on the formation of first marriages among a current cohort of young women. While most of the data used in previous studies are valuable for documenting historical changes in the relationship between nonmarital childbearing and marriage, those data do not contain a representative sample of women whose childbearing and union formation have occurred at the turn of the 21st century. This limitation is unfortunate, given that marriage promotion programs have been extensively implemented over the last decade. We fill this gap by utilizing the recently released Add Health Wave IV survey.

Add Health is a longitudinal study of a nationally representative sample of adolescents in grades 7-12 in the United States during the 1994-95 school year (Harris et al. 2009). The Add Health cohort has been followed into young adulthood with four in-home interviews, the most recent in 2008, when the sample was aged 24-32. The Wave I data produced a total sample size of 20,745 adolescents, 10,480 of whom are female. Their parents also were interviewed in Wave I. In 2008, approximately 15,700 Wave I respondents, 8,352 of whom are female, were re-interviewed in

Wave IV to investigate developmental and health trajectories across the life course of adolescence into young adulthood. The Wave IV survey contains detailed life history data on fertility, union formation, and educational attainment. The Wave I survey provides rich sets of variables that are measured at the individual-, family-, school-, and neighborhood-levels, many of which are unobservable in the previous literature but known to affect both nonmarital childbearing and entry into marriage. For example, Add Health allows us to measure individuals' socioemotional traits, personality, risk behaviors, parent-child relationships, peer networks, and school climate, alongside an array of standard sociodemographic characteristics.

This study employs discrete-time event history modeling in a propensity score weighting framework to investigate the relationship between nonmarital childbearing and the formation of first marriages. A key feature of our analytic strategy is to fit the models to a person-year data file rearranged from the Add Health data using an IPTW estimator, such that the treated (women who had a first nonmarital birth) and control (women who did not) groups at year t are balanced on observed covariates. We consider age 11, or the earliest age at which a respondent was interviewed for the Wave I survey, as start of follow-up time for our analysis. We construct two types of weights, one based on treatment assignment and the other on censoring. First, let B_t be 1 if a respondent gave a nonmarital first birth at time t and 0 otherwise, and $\boldsymbol{X_0}$ be a vector of time-constant covariates measured at Wave I. For each of time-varying covariates, we use overbars to denote the history of that covariate up to time t. For example, $\overline{\boldsymbol{X}}_{it} = \{\boldsymbol{X}_{i0}, \boldsymbol{X}_{i1}, \dots, \boldsymbol{X}_{it}\}$ is the covariate process for respondent i up to time t. Then the weight for treatment assignment is given by

$$w_{it}^{B} = \prod_{k=0}^{t} \frac{1}{Pr(B_k | \overline{B}_{ik-1}, \boldsymbol{X}_{i0}, \overline{\boldsymbol{X}}_{ik-1})} \ ,$$

where the denominator states the probability that the respondent received her own observed treatment at time t, given her prior treatment and covariate history. Using this weight, however, is known to lead estimators to large variance because a small number of observations with the extreme weights are likely to dominate the estimation process (Hernán, Brumback, and Robins 2002). To increase efficiency, we use the "stabilized" weight

$$sw_{it}^{B} = \prod_{k=0}^{t} \frac{Pr(B_{ik}|\bar{B}_{ik-1}, \boldsymbol{X}_{i0})}{Pr(B_{k}|\bar{B}_{ik-1}, \boldsymbol{X}_{i0}, \bar{\boldsymbol{X}}_{ik-1})} \ ,$$

where the numerator states the probability that the respondent received her own observed treatment at time *t*, given her prior treatment history and time-constant covariates, but not further adjusting for her time-varying covariate history.

Second, let C_t be 1 if a respondent was right-censored by time *t* and 0 otherwise. Assuming that the censoring process is noninformative after adjusting for observed covariates, the stabilized weight for censoring is calculated analogously to sw_{it}^{B}:

$$sw_{it}^{C} = \prod_{k=0}^{t} \frac{Pr(C_{ik} = 0|\bar{C}_{ik-1} = 0, \bar{B}_{ik-1}, \boldsymbol{X}_{i0})}{Pr(C_{ik} = 0|\bar{C}_{ik-1} = 0, \bar{B}_{ik-1}, \boldsymbol{X}_{i0}, \bar{\boldsymbol{X}}_{ik-1})} \ ,$$

which is the ratio of the respondent's probability of remaining uncensored by time *t*, given her treatment history and time-constant covariates, to her conditional probability of remaining uncensored by time *t*, further adjusting for her time-varying covariate history.

Third, the final stabilized weight is given by

$$sw_{it} = sw_{it}^{B} \times sw_{it}^{C} \ ,$$

which states the inverse of the conditional probability that a respondent had her observed treatment and censoring history by time *t*. Because both sw_{it}^{B} and sw_{it}^{C} are unknown, we estimate them from pooled logistic models that treat each person-year as a unit of observation and include all measured covariates.

Finally, we fit the following weighted discrete-time event history model to examine the effect of nonmarital childbearing on the formation of first marriages:

$$\log\left[\frac{P_{it}}{1-P_{it}}\right]_{sw} = \alpha_t + B_{it}\beta + \boldsymbol{X}_{i0}\gamma \ ,$$

where P_{it} denotes the conditional probability of entering a first marriage for woman i at year t and α_t is a set of time dummy variables. Note that, as suggested earlier, the effect of nonmarital childbearing may be confounded by the presence of time-varying covariates—e.g., educational attainment and cohabitation status—if they affect subsequent nonmarital childbearing status and first marriage formation and are also affected by past nonmarital childbearing status. To account for this confounding, we do not include the time-varying covariates as explanatory variables in the model above but use them to calculate the weights. The model is estimated with robust standard errors to correct for within-individual correlation (Hernán, Brumback, and Robins 2000).

Our analysis will present the following tables and figures:

Table 1. Descriptive Statistics

Table 2. Logit Model Predicting Nonmarital Childbearing Status

Table 3. Stabilized Weights for Treatment and Censoring

Figure 1. Unadjusted Hazard Rates of First Marriage by Nonmarital Childbearing Status

Table 4. IPTW Estimates of the Effect of Nonmarital Childbearing on First Marriage Formation

Table 5. Sensitivity Analysis

Table 6. IPTW Estimates of the Effect of Nonmarital Childbearing on First Marriage Formation, Teen vs. Non-Teen Unmarried Mothers

Table 7. IPTW Estimates of the Effect of Nonmarital Childbearing on First Marriage Formation, by Cohabitation Status

Table 8. IPTW Estimates of the Effect of Nonmarital Childbearing on First Marriage Formation, by Race/Ethnicity

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