A Demographic Analysis

of the Family Structure Experiences of Children in the United States*

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Revised February 2008

In press, *Review of Economics of the Household*

*Financial support from NICHD grant HD45587 is gratefully acknowledged. Thanks to Karin Gleiter for expert programming. A previous version of this paper was presented at the 2005 Annual Meeting of the Population Association of America in Philadelphia, and in seminars at the Carolina Population Center, Cornell, Syracuse, NYU, and the 2005 NIH Workshop on Intergenerational Family Resource Allocation. We are grateful for comments by the editor, referees, and seminar and conference participants. The authors alone are responsible for the contents. The views expressed are those of the authors and do not necessarily reflect those of the Federal Reserve Bank of New York.

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Abstract

This paper analyzes the family structure experiences of children in the U.S. Childbearing and transitions among single, cohabiting, and married states are analyzed jointly. A novel contribution is to distinguish men by their relationship to children: biological father or stepfather. The analysis uses data from the NLSY79. A key finding is that children of black mothers spend on average only 33% of their childhood living with the biological father and mother, compared to 74% for children of white mothers. The two most important proximate demographic determinants of the large racial gap are the much higher propensity of black women to conceive children outside of a union, and the lower rate of "shotgun" unions for blacks compared to whites. Another notable finding is that cohabitation plays a negligible role in the family structure experiences of children of white mothers, and even for children of black mothers accounts for less than one sixth of time spent living with both biological parents.

JEL: J10

INTRODUCTION

Many children growing up in the United States in recent decades have spent a significant part of their childhood living in non-traditional family structures. Major increases in divorce, out-of-wedlock childbearing, and cohabitation have resulted in rapid growth in the prevalence of alternative family structures, such as living with the biological mother and the biological father in a cohabiting (unmarried) relationship, the biological mother and a step father, and the biological mother and no man. An important question of interest to social scientists, policy makers, and parents is how growing up in alternative family structures affects children, compared to the traditional experience of being raised by married biological parents. A large literature spanning several disciplines has analyzed the effects of children's family structure experiences on psychological, social, demographic, and economic outcomes, both during childhood and subsequently in adulthood.[1]

An important issue raised by such studies is what determines the family structure experienced by children during the course of their childhood. Factors that influence family structure may also have a direct impact on child outcomes, making it difficult to infer causality from correlations. But the family structure experiences of children are of independent interest as well. The proximate determinants of family structure are fundamental demographic behaviors: incidence and timing of childbearing, cohabitation, and marriage. A key point recognized in the literature is that the *interaction* between childbearing and marital behavior determines the family

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[1]Well-known examples include McLanahan and Sandefur (1994), Chase-Lansdale, Cherlin, and Kiernan (1995), and Hetherington and Stanley-Hagan (1999). Recent analyses include Aughinbaugh, Pierret, and Rothstein (2005), Gennetian (2005), Ginther and Pollak (2004), Hofferth (2006), Lang and Zagorsky (2001), and Sigle-Rushton, Hobcroft, and Kiernan (2005).

structure experienced by children. There are many studies of the demographic behaviors that influence family structure, but most of these studies do not draw implications from their findings for the family structure experienced by children. This is because the latter requires an integrated analysis. For example, divorce presumably affects children differently if they were alive at the time of the divorce compared to being born after the divorce. Similarly, the impact on children of a cohabiting versus a married relationship between a child's biological parents may depend on whether the child was born before or during the cohabitation. The impact on a child of being born out of wedlock is likely to depend on whether the mother and biological father subsequently marry or cohabit, and if so, how soon after the birth of the child.

A handful of studies analyze the implications of marital and childbearing behavior for the family structure experiences of children. But these studies have taken a limited perspective because they typically analyze only a subset of the relevant behaviors. For example, some studies focus only on divorce (Waite and Lillard 1991) or formation of formal unions (Upchurch, Lillard, and Panis 2001), while others focus only on children who were born out of wedlock (Brien, Lillard, and Waite 1999; Aquilino, 1996; Carlson, McLanahan, and England, 2004). An important issue in analyzing demographic behavior from the perspective of children is the identity of the man with whom a mother lives. A woman who has given birth to a child outside of a co-residential relationship is at risk of entering a co-residential relationship with the father of the child and with other men. When analyzing relationship formation from the mother's perspective, this distinction is rarely made (Graefe and Lichter 1999 is an exception). It could be important for the durability of the relationship, and thus important to analyze from the mother's perspective. But it is critical for understanding the family structure experiences of children born outside of a co-residential union.

In this paper we provide a more comprehensive demographic analysis of the family structure experiences of children than has been reported in previous studies. We jointly analyze transitions among co-residential union states defined by single, cohabiting, and married, together with childbearing and the identity of men from the perspective of children: biological father or stepfather. Modeling transitions into and out of cohabitation, marriage, and single status jointly with childbearing behavior provides a richer picture of family dynamics than does analyzing marital behavior in isolation. And modeling the identity of men from the perspective of children provides a unique perspective on transitions of children among living arrangements.

The analysis uses data from the 1979 cohort of the National Longitudinal Survey of Youth (NLSY79). These data provide detailed histories of the relevant demographic events, and unlike many other surveys, provide information on biological and non-biological fathers for each child.[2] These data have been used in many previous analyses of family structure, but we exploit the richness of the data more fully than in previous studies, particularly the information on the identity of men from the perspective of children.

The most striking finding of the analysis is the large difference between the children of black and white mothers in time spent living with the biological father. Children of non-Hispanic

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[2]See Andersson (2002) for a descriptive comparison of the family structure experiences of children in 15 countries, including the U.S. His findings show that "The USA stands out as one extreme case with its very high proportion of children born to a lone mother, with a higher probability of children who experience a union disruption of their parents than anywhere else, and with many children having the experience of living in a stepfamily." (Page 343). Heuveline et al. (2003) also report that the U.S. is an outlier in the sense that over half of all time spent by American children in single parent families is accounted for by children born to lone mothers. This is true of only three countries in their 17-country sample.

white mothers spend 74% of their childhood years living with their biological mother and father on average, compared to 33% for children of non-Hispanic black mothers. Children of Hispanic mothers are closer to whites at 65%.[3] It is also notable that the distribution of time spent living with the biological father is heavily skewed for children of black mothers: the median is only 0.3 years, and 49% of children of black mothers never live with the biological father. The distributions are much more symmetric for the children of white and Hispanic mothers. Surprisingly, estimates of the duration of childhood spent with both biological parents have not been presented previously in the literature.

An important source of the difference between blacks and whites is the role of the mother's status at the time of conception. For both groups, the mother's marital status at birth plays a dominant role in determining time spent with the biological father during childhood. However, if the mother was single at the time the child was *conceived*, the percent of childhood spent with the biological mother and father together is 46% for whites, and 14% for blacks. The black-white difference results from the fact that "shotgun" weddings and cohabitations are much more common for whites than for blacks (Manning, 2004; Manning, Smock, and Majumdar, 2004). Another important source of the difference is the higher propensity of black women to conceive children outside of a co-residential union. We find that controlling for the mother's family background and educational attainment does not reduce the black-white gap in the proportion of childhood spent with both biological parents by very much.

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[3]Here and throughout the paper, whites and blacks include only non-Hispanics. Hispanics can be either white or black (or another race). To save space, we focus mostly on children of black and white mothers. Results for Hispanics are almost always in between those for whites and blacks, and closer to whites. See the working paper version (Blau and van der Klaauw, 2007) for a more detailed discussion of results for Hispanics.

BACKGROUND

The major demographic trends that are the backdrop for our analysis are well known. The prevalence of marriage has declined in the last 30 years in the U.S., as a result of both later age at marriage and an increase in the proportion of adults who never marry (Fields and Casper, 2001). Many young unmarried adults live in cohabiting relationships (Bumpass, Sweet, and Cherlin, 1991). While marriage has declined, divorce has increased substantially (Kreider and Fields, 2002; Stevenson and Wolfers, 2007). Births to unmarried women increased from 18% of all births in 1980 to 33% in 2000 (Martin et al., 2002). However, children born to unmarried women do not necessarily live with only one biological parent. Thirty-nine percent of births to unmarried women from 1990 to 1994 were to women in cohabiting relationships (Bumpass and Lu, 2000). Thus while about half of children born recently are expected to ever spend time in a family with an unmarried parent (Cherlin, 1999), many of these children are in fact living with both biological parents during that time.

Our analysis is most closely related to two previous papers: Brien et al. (1999), and Graefe and Lichter (1999). Brien et al. present an integrated analysis of non-marital conception, entry to cohabitation, and entry to marriage, using 1986 data from the National Longitudinal Study of the Class of 1972. The analysis is integrated in the sense that the effect of one type of behavior on the others is modeled. For example, they find that a non-marital conception greatly increases the risk of marriage during the pregnancy, but reduces the risk of marriage for women who do not marry quickly in response to conceiving a child out of wedlock (see also Bennett, Bloom, and Miller 1995). Their analysis controls for unobserved heterogeneity using a random effects approach, uses rich monthly event histories, and a flexible specification of occurrence and duration dependence. Like Brien et al., we jointly analyze several key demographic behaviors, use rich multiple spell data, avoid restrictive specifications of duration and occurrence dependence, and control for unobserved heterogeneity. We go beyond Brien et al. by (1) modeling union dissolution behavior and marital conceptions, (2) using prospective data that are less subject to recall error, (3) using data through 2002, and most important, (4) explicitly incorporating the identity of men in the analysis, thereby making possible inferences about the family structure experiences of children from an analysis of the behavior of mothers.

Graefe and Lichter (1999) analyze marital transitions from the perspective of children. They use data from the NLSY79 covering children born from 1980-1992 to model transitions of children from a single-mother family to cohabitation and marriage, and transitions from cohabitation to single and married. In the latter model, they include as a covariate an indicator of whether the cohabiting man is the biological father of the child. This may be different for different children in the family, hence the use of the child as the unit of analysis. They find that biological children are at much greater risk of transition from cohabitation to marriage and at lower risk of transition to single, compared to stepchildren. Like Graefe and Lichter, we focus on the identity of men with whom mothers cohabit and marry, i.e. their relationship to the woman's children. We go significantly beyond Graefe and Lichter by (1) accounting for the identity of men in analyzing transitions from single to cohabitation and marriage; (2) analyzing the determinants of step versus biological father status by modeling conceptions jointly with marital behavior; and (3) recognizing that while child characteristics can influence marital transitions, the appropriate unit of analysis for marital transitions is the mother, not the child. Thus we analyze the behavior of women, but our modeling approach enables us to draw inferences about the implications of their behavior for the family structure experiences of children. We also use monthly instead of annual data, extend the analysis back to include all births and forward through 2002, and account for unobserved heterogeneity.

CONCEPTUAL ISSUES AND EMPIRICAL MODEL

We focus on the marital and childbearing behavior of women. The behavior of men is obviously relevant, but the complexity of modeling male and female behavior jointly, as well as the limitations of the available data, necessitate a focus on the woman's perspective. We assume that women become at risk of entering a co-residential union and conceiving a child at age 12. We use a monthly discrete time framework.[4] For empirical tractability, it is assumed that at most one demographic event can occur in a given month, where events include conceiving a child, giving birth, beginning a co-residential relationship, and ending a co-residential relationship. With periods of one month, it is rare for two events to occur in the same month. The advantage of this assumption is that it allows us to use a competing risks empirical model in which there is no chance that two or more events can occur in the same period.[5]

We consider three co-residential romantic relationship states, which for brevity we refer

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[4]Lillard's (1993) continuous time model of simultaneous hazards is in principle an attractive framework for the analysis. However, it is very difficult to deal with missing data in Lillard's model. We discuss below how missing data are handled.

[5]The assumption that at most one event occurs per month implies that a woman cannot end a co-residential relationship with one man and begin a co-residential relationship with another man in the same month. A new co-residential relationship can be formed only if the woman is single at the beginning of the month. And a newly formed co-residential relationship cannot be dissolved in the period in which it was formed. Hence, we assume that all partnerships last at least one month, and all single spells last at least one month.

to henceforth as *union* states: single, married, and cohabiting. If a woman was single in the previous period, her current period union options are to remain single, enter a cohabitation, and marry. If she was cohabiting in the previous period, then her union choices are to end the relationship and become single, continue the cohabitation, and marry her partner. If the woman was married in the previous period, then her union choices are to end the marriage and become single, or continue the marriage.

We consider only conceptions that lead to a live birth. The focus of the analysis is the family structure experiences of children, so pregnancies that do not result in a live birth are not of direct interest.[6] If a conception leading to a live birth occurs, the duration of the pregnancy is taken as given. Thus, while conception is a risk in the model, conditional on conceiving a child birth is not a risk. A mother is at risk of changes in her *union* status while pregnant, but birth itself is treated as an exogenous censoring event that ends a pregnant spell. Thus, the timing of births, conditional on the date of conception, is taken as given. The occurrence of twin births is also taken as given.

The most novel aspect of our analysis is the focus on the identity of men. The relevant issues are (1) whether a newly formed union is with a man who has previously fathered any of the woman's children, and (2) whether a newly conceived child was fathered by the same man who fathered the woman's youngest child, if any. These factors determine whether a given child lives with - or is at risk of living with - his biological father or a stepfather. Note that stepfather

[6]Ignoring pregnancies that do not lead to a live birth results in some women being treated as if they were at risk of conception in some months in which they actually are pregnant. And the occurrence of a miscarriage or abortion could affect the demographic processes of interest. Accounting for pregnancies that do not result in a live birth would require modeling at least two additional events - miscarriage and abortion - in an analysis that is already quite rich.

here means a co-residing man who is not the child's biological father, regardless of whether the union is a cohabitation or a marriage. One key assumption is invoked in order to make the analysis tractable: if a woman ends a union with a given man, she is not at risk of conceiving a child or entering a union again with that man.[7] Without this assumption, it would be necessary to keep track of all men with whom a woman ever lived or who fathered any of her children, because she would be at risk of entering a relationship or conceiving a child with all such men.

With this assumption, we can limit attention to at most two men in any given period. One is the *current* man, defined as the partner or spouse if in a union, or, if single, the father of the woman's most recent child born since the end of the last union (or since she began bearing children if she has never been in a union; this will be implicit henceforth). The other man is a *new* man, defined as a man who has not fathered any of her children and with whom she has never lived. For a woman in a union, only the current man is relevant: if she lives with a man in month t-1, that man is by assumption the only man with whom she can live and conceive a child in month t. For a single woman with no children born since the end of the last union, only the new man is relevant: there is no current man. If she enters a union or conceives a child, it must be with a new man. For a single woman who has given birth to at least one child since the end of her most recent union, both the current man and a new man are relevant. If she enters a union, it matters whether it is with the current man, because in this case some of her children will have their biological father present. Alternatively, if the relationship is with a new man, then all of her children will have a stepfather present. And if she conceives a child with the current man, the

[7]Note that this assumption does not rule out a woman having multiple children with a given man with whom she does not live. Such behavior is ruled out only if the woman bears a child with a different man between pregnancies with the man in question.

number of children at risk of living with the biological father in the event of a future union with this man increases by one. If the conception is with a new man, then children fathered by any previous men are (by assumption) no longer at risk of living with the biological father in the event of a future union.[8]

In order to describe the model more formally, let B_t denote the number of children born to a single woman since the end of the previous union. Consider a single pregnant woman with B_t>0. Define N_t=1 if the father of the unborn child is a new man, and N_t=0 if the father is the current man. In any given month, a woman is at risk of one or more of the following events: (1) conceiving a child with the current man; (2) conceiving a child with a new man; (3) ending a union and becoming single; (4) entering a cohabitation with the current man; (5) entering a cohabitation with a new man; (6) marrying the current man; (7) marrying a new man. A *state* is defined by the unique *set* of risks to which a woman is subject. The set of risks she faces depends on her pregnancy status, marital status, B_t, and N_t. Table 1 identifies eight states defined by the unique set of risks to which a woman is subject. In states 1-4, she is not pregnant, and is therefore at risk of conceiving a child. In states 5-8 she is pregnant and is not at risk of conceiving. In states 1, 2, 5, and 6, she is single and is therefore at risk of entering a cohabitation or marriage. In state 1 there is no current man (B_t=0), so the relationship or conception can only be with a new man. In state 2 there is a current man (B_t>0): the father of the most recent of the

[8]Note that the new man this period could be different from the new man in previous periods. There is no way to determine this empirically, because we know nothing about non-co-residential romantic relationships unless and until they become co-residential. For example, we do not know how long a woman may have been dating a given man. This could be relevant in determining the risk that he fathers a child or enters a union with the woman. See Carlson et al. (2004) for an analysis of this issue using data from the Fragile Families study.

children born since the end of the previous union, so she is at risk of entering a relationship or conceiving a child with *both* the current man and a new man. In state 5, B_t>0 and the father of the child with whom she is pregnant is the current man. Thus she is at risk of entering a relationship with the current man only. In state 6, she may or may not have given birth to any children since the end of the previous union, but the father of the child with whom she is pregnant is a new man, so she is at risk of entering a relationship only with this new man (who becomes the current man upon conception of the child). In states 3 and 7 she is cohabiting, and is at risk of marrying the current man and ending the relationship. In states 4 and 8 she is married and is at risk of ending the relationship.

We specify an index function $h_{js}(t)$ for the occurrence of the j^{th} event (j=1,...,7) in month t while in state s (s=1,...,8). Note that not all risks are relevant in each state; Table 1 shows the set of events of which the woman is at risk when in state s, J_s. We assume that

$$h_{js}(t) = Z(t)\alpha_{js} + X\beta_{js} + \varepsilon_{jst}, \qquad j \in J_s, \tag{1}$$

where $Z(t)$ is a vector of polynomials in duration and age (duration of the current union, duration of pregnancy, duration since previous birth, mother's age, etc.), X is a set of variables that are constant within a spell in a given state (e.g., number of children, number of children with biological father present, marital status, race), α_{js} and β_{js} are coefficient vectors, and ε_{jst} is a disturbance. Within a spell, all durations and ages are perfectly collinear, but the α_{js} coefficients are identified by variation across spells in the calendar month of the begin date of the spell and by the existence of multiple spells in a given state. Additive separability of $Z(t)$ and X is not crucial, and is selectively relaxed in the empirical analysis. This expression for $h_{js}(t)$ can be interpreted as an approximation to the combination of decision rules and stochastic processes that determines the risk of occurrence of the j^{th} event while occupying state *s*. The value of alternative marital and fertility choices will depend on the woman's marital and fertility history if these variables affect the flow of utility she receives from the alternative choices or if they affect her expectations about future outcomes, including her expectation about the effects on her children of alternative marital choices. Hence, *X* includes measures of her marital and fertility history. The value of the alternative choices may depend on other observable factors such as her age and race, for similar reasons. The value of alternative choices will also depend on unobserved factors, such as the characteristics of her current partner, and the perceived state of the marriage market. These are captured by ε_{jst}. Note that this specification is very flexible: the effects of all state variables can vary freely with the state occupied, that is with the combination of union status, pregnancy status, and partner status.

The disturbance ε_{jst} is specified as $\varepsilon_{jst} = \rho_{js}\mu + \eta_{jst}$, where μ is a permanent error component, ρ_{js} is an event-and-state-specific factor loading, and η_{jst} is an independently and identically distributed error component that is assumed to follow a Type I extreme value distribution. Ignoring persistent unobserved heterogeneity would result in the *Z(t)* and *X* variables picking it up, and would lead to invalid causal interpretations of correlations between current and previous marital and fertility behavior. For example, women who have children at a young age may be quite different in unobserved ways from women who delay childbearing, and μ will help control for such differences. The extreme value assumption is convenient, because the resulting event probabilities have the multinomial logit form, conditional on μ. We assume that μ has a discrete step function distribution with two mass points, resulting in a mixture model. The model is estimated by maximum likelihood, integrating μ out of the likelihood function.

DATA

The NLSY79 cohort contains individuals born from 1957 to 1964. They were interviewed annually from 1979-1994 and have been interviewed biennially since 1994. We use data through the 2002 interview. The representative cross-section sample and supplementary over-samples of blacks and Hispanics are used in the analysis. There are 4,926 women in these groups.

The NLSY79 data provide many advantages for the analysis, compared to other possible data sources, including information allowing the identity of men to be determined from the perspective of children and large enough samples to analyze whites, blacks, and Hispanics separately. The data have one main disadvantage compared to other commonly used data sources in the family demography literature. The NLSY79 follows a birth cohort of women and the children born to these women. The children themselves do not form a well-defined cohort: their birth dates range from 1970 through 2002 (and beyond), and their only common link is that they were all born to women who are themselves part of a well-defined birth cohort. Thus the analysis includes children born over a long period of time, in which there were major changes in some of the demographic behaviors of interest. It is clearly inappropriate to use these data to draw inferences about birth cohort trends, but this is not our goal.

In 1979, the survey collected information on the beginning and ending dates (to the nearest month) of up to two marriages. In subsequent waves, information has been collected on up to three changes in marital status that occurred since the previous interview. Changes include marriage, separation, re-uniting after a separation, divorce, death of a spouse, and re-marriage.[9]

The survey has collected information on cohabitation in several different ways: (1) At each interview date, the respondent is queried about her relationship to other members of her household. "Partner" is one of the relationship codes included in the resulting household roster. (2) Beginning in 1990, respondents were asked to report the date on which the cohabitation began for cohabitations that were in progress at the interview date. And if a respondent is married at the interview date, she is asked whether she lived with her spouse before the marriage began, and if so when the cohabitation began. These questions elicit a more precise date for the beginning of a cohabitation, but the information is based on recall, so we use it only if it does not conflict with the household roster. (3) The cohabitation questions were completely redesigned in the 2002 interview. For the first time, both the beginning *and* ending date of cohabitations that did not turn into marriages were ascertained. And cohabitations that lasted less than three months are ignored.

We combined information from the interview date, the retrospective reports, and the

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[9]Because we are interested in the implications of marital behavior for the family structure experiences of children, the end of a marriage is defined to occur at the time of separation rather than divorce. However, some separations are temporary, and modeling the process that determines whether a separated couple reunites would make the analysis overly complex. Therefore, we ignore separations that result in reuniting rather than divorce, if the temporary separation lasts less than two years. If a temporary separation lasts more than two years, we censor the observation at the date of separation. The only exception is if the woman had not conceived any children before the end of the separation. We also allow for transitory separations involving cohabitation, again if the separation lasts no more than two years. The date of separation was not ascertained for marriages that ended in divorce before the 1979 interview. We use the date of divorce as the ending date of the marriage in these cases (2.4% of first marriages). In the sample of 4,926 women, there were 1,676 separations, of which 19% were temporary. The median duration of a temporary separation is 17 months, and 60% were shorter than two years and therefore ignored. The other 40% were right-censored.

2002 interview to form as complete a cohabitation history as possible. The beginning and ending dates of many cohabitations could be identified only to the nearest survey date. Rather than discard such cases, we include them in the analysis and modify the likelihood function to account for the lack of precise beginning and ending dates. If we know that a cohabitation began sometime between the 1986 and 1987 interview dates, for example, the likelihood function includes terms for each of the possible beginning months, weighted by the probability that the cohabitation began in the given month. This uses the available information as efficiently as possible. However, cohabitations that began and ended before the 1979 interview or between interviews are missed completely (except for those that began and ended between the 2000 and 2002 interview and lasted at least three months), so the cohabitation histories are incomplete. The details of how the likelihood function accounts for uncertain cohabitation begin and end dates, as well as for other incomplete or missing data of the type discussed later in this section**,** are in the working paper version of this article (Blau and van der Klaauw, 2007).[10] The cohabitation and marriage histories were combined to form a complete union history.

The month and year of birth of each child is reported by the respondent. Beginning in 1984, women were asked when each pregnancy began. We use this information to identify the

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[10]Sixty percent of freestanding cohabitations had a beginning date that was not known to the nearest month, and 95% had an uncertain end date. Forty percent of cohabitations that turned into marriages had an uncertain begin date (the end date in this case is the date of marriage, which is always known to the nearest month). Another consequence of the fact that some cohabitation dates are not known to the exact month is that there are cases in which the *sequence* of events cannot be determined. For example, if a cohabitation began sometime between the 1986 and 1987 interview dates and a birth also occurred in this interval, then we do not know whether the woman was single or cohabiting at the time of the birth. Ambiguity about the sequence of events occurs at least once for 9% of women in the sample. Rather than discarding such cases, we modify the likelihood function to account for all the possible sequences in which events could have occurred, weighted by the probability of each sequence.

month of conception. If the information is missing, we assume the conception occurred 9 months prior to the birth.[11]

Identifying fathers is one of the critical tasks. Beginning with the 1984 interview, the mother is asked for every biological child present in her household whether the biological father of the child is present. Thus, when a woman lives with a man before or during the conception or birth of a child, identifying fathers is straightforward. The more difficult cases are those in which a woman conceives and bears a child while single. In such cases, we can identify whether the father of the child was the current man or a new man only if she subsequently moves in with a man. Similarly, following the birth of a child while single, if she moves in with a man and the union ends before the 1984 interview, then we cannot determine whether the father of that child was the current man or a new man. And if a man moves in and out between interviews, we cannot determine the father of the child. Of the 1,086 cases in which a child was conceived and born to a single woman who had given birth to a child since the end of her previous union, we are able to identify whether the father is the current man or a new man in 35% of the cases. Rather than discard the remaining cases, we again modify the likelihood function to account for both possibilities, weighted by the probability that the father was the current man or a new man. The details of this approach are available in Blau and van der Klaauw (2007).

Finally, at each interview date we can determine from the household roster whether a given child is present in the mother's household. We do not model the processes that determine

[11]Sixty percent of live births had an observed conception date. For these cases, the birth occurred 9 months after the conception in 55% of cases, 8 months after in 30% of cases, and 10 months after in 7% of cases. Child deaths are treated as exogenous censoring events, and the spell in progress at the time a child died is right-censored. A new spell then begins with one less child present.

whether children move in and out of the mother's household. If a child moves out of the mother's household, the spell in progress is right-censored at the date of the last interview in which the child was known to be present, and a new spell begins with one less child present. If the child subsequently moves back in, the spell in progress is right-censored at the last date at which a child was known to be living outside the mother's household, and a new spell begins. We treat cases in which a child is away at school or living part-time with the mother as if the child is living with the mother. Eighteen percent of children ever move out before age 18, and 60% of those children ever move back in before age 18.

After dropping cases with inconsistent union histories, event histories that violate the assumptions of our model, and other cases with problematic data, we are left with a sample of 4,480 women out of 4,926.[12] These women bore 7,970 children as of their latest interview or as of the date at which their event history was censored. The average age at the last date of observation is 38.7, and for 74% of the sample this date corresponds to the 2002 interview. Table 2 summarizes the union and childbearing behavior of the sample, separately for white and black

[12]There are 415 cases in which there were important inconsistencies in the union history that could not be resolved. Cases that violate the assumptions of the model include women who (1) end a marriage or cohabitation with a man, and subsequently form a new co-residential union with the same man (except cases of temporary separation, as discussed above); (2) women who bear a child with one man, then with a second man, and then with the first man; and (3) cases in which two or more demographic events occur in the same month (e.g. marriage and birth; conception and exit from cohabitation). There were 114 cases in which a woman apparently violated assumption (1) and otherwise would have been kept in the sample; 68 cases in which a woman apparently violated assumption (2); and 65 cases in which assumption (3) was apparently violated. We say "apparently" because it is likely that some of these cases are a result of errors in identifying men, but there is no way to determine this. A total of 446 cases with inconsistencies in the union history or violations of model assumptions were dropped**.**

women.[13] There are large differences between blacks and whites in union and childbearing behavior. 89% of white women and 61% of black women had ever married as of the last date of observation. Of women who ever married, 37% of whites and 52% of blacks ever ended a marriage (divorced or permanently separated). 25% of the white sample ever experienced the end of a marriage when children were present, compared to 42% of blacks. 60% of white women who experienced the end of a marriage ever remarried, compared to 30% of blacks. 42% of whites and 35% of blacks ever cohabited. 74% of cohabitations of whites ended with a transition to marriage, compared to 59% for blacks.

22% of whites ever conceived a child while single, compared to 64% of blacks. 11% of whites ever gave birth to a child while single, compared to 58% of blacks. These rates are somewhat lower than the period rates for 2000 (Martin et al., 2002). Only 7% of women ever conceived a child while cohabiting, while 4% of whites and 8% of blacks ever gave birth while cohabiting.[14] These figures illustrate the importance of using monthly data in order to capture union formation and dissolution events that occur during pregnancy. "Shotgun" union formation appears to be quite important for whites: while 21.7% of white women ever conceived a child while single, only 10.7% ever gave birth while single. The comparable figures for blacks are 63.9% and 58.4% , indicating that shotgun unions are a less significant phenomenon for blacks.

[13]As discussed above, cases in which the sequence of events is uncertain are included in the analysis. They contribute terms in the likelihood function for all possible sequences, weighted by the estimated probability from the model of each alternative sequence. The descriptive statistics presented in Tables 2 and 3 weight each alternative possible sequence equally.

[14]Data from the early 1980s show that 4 and 13% of births were to cohabiting women, for whites and blacks, respectively (Bumpass and Lu, 2000). The comparable figures for the early 1990s were 9 and 16, respectively.

The data also show that a sizeable fraction of cohabitations convert to marriage following conception. 36% of white mothers ever experienced an episode of being single with children, compared to 82% of black mothers. 24% of white mothers and 36% of black mothers ever had an episode in which a child had a stepfather present.

Table 3 summarizes family structure experiences from the perspective of children, based on data up to the last age observed or the month in which the child turned age 18, whichever comes first. 16% of children of white mothers were conceived while the mother was single, compared to 64% of children of black mothers. 52% of children of white mothers who were conceived while single were born in a union, most in a marriage, compared to 10% of children of black mothers. 30% of children of white mothers ever experienced an episode in which they lived with the mother without any man, compared to 76% of children of black mothers. 17% of children of white mothers and 27% of children of black mothers ever lived with a step father. 94% of children of white mothers and 52% of children of black mothers ever lived with their biological father. Of children whose biological father was not present at birth, the biological father moved into the child's household in 25% and 15% of cases for children of white and black mothers, respectively. Of children whose biological father was ever present in the child's home, the father moved out of the child's household in 24% and 44% of cases for children of white and black mothers, respectively. Of children who ever lived without a man present, a step father moved into the household in 53% and 36% of cases for children of white and black mothers, respectively. 36% of children of white mothers who ever have a stepfather present experienced the exit of the stepfather, compared to 49% of children of black mothers.

RESULTS

The estimated coefficients of our discrete-time competing risks transition model, which for each origin state corresponds to a panel data multinomial logit model**,** are presented in the Appendix in Table A-1. The coefficients are not particularly informative, so we do not discuss them in detail. It is worth noting, however, that race effects are important, as expected: only 3 out of 21 intercept shifts are *not* significantly different from zero at the 5% level for blacks. Differences between Hispanics and whites are less often significant. More generally, the majority of coefficient estimates (57%) are significantly different from zero at the 5% level. The transitions with relatively imprecise estimates are those for which the samples at risk are relatively small.

In order to illustrate the implications of the estimates, we used them to simulate the life histories of 50,000 artificial women who are subject to the risks characterized by the model. Each woman starts at age 12 in state 1 (single with no children). The estimated parameters for state 1 are used to compute the probability of each of the three events that can occur to a woman in state 1. A random number generator determines which, if any, event occurs. If the woman experiences an event, she changes states accordingly. If the event is conceiving a child, a pregnancy duration is randomly assigned according to the observed distribution of pregnancy durations in the sample. If no event occurs, she remains in state 1. The state variables are updated according to which event, if any, occurred, and the process is repeated for the next month. If pregnant, the birth occurs at the assigned duration. The process continues to age 39 of the woman.[15] Simulations are computed separately for whites, black, and Hispanics, in each case integrating over the estimated heterogeneity distribution.

Figure 1a shows the simulated mean cumulative duration of time spent with the biological father from birth through the 18th birthday, separately for children of white and black mothers.[16] Children of white mothers can expect to spend about 11 years with the biological father by age 18, compared to 3 years for children of black mothers. Accounting for the fact that not all simulated children are observed through age 18, 74% of observed childhood is spent living with the biological father by children of white mothers, compared to 33% for children of black mothers (see Table 4). This large difference between blacks and whites is implied by well-documented differences in racial patterns of union formation, dissolution, and childbearing, but to our knowledge has never previously been directly illustrated**.** Even more striking than differences in the means are differences in the skewness of the distributions. Table 4 summarizes the distribution of lifetime years spent with the biological father as of the last age observed for each child. 49% of children of black mothers are estimated to never live with their biological father during childhood, compared to 6% of children of white mothers. These simulated

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[15]The sample includes women up to age 45, but the average age at the last observation is about 39. The simulations are truncated at age 39 in order to allow a comparison to the data. See the working paper version (Blau and van der Klaauw, 2007) for a comparison of the simulations to the data. As in the data, some children are not observed for their entire childhood in the simulations. The simulated data for children are truncated at age 18, as were the real data. In the simulations, there are no deaths or children who move in or out of the mother's household. The only exogenous explanatory variable other than race/ethnicity is the woman's date of birth, which is set to the sample mean for all women in the simulations.

[16]It is implicit here and throughout the remaining discussion that the child lives with the biological mother. Hence, "time spent with the biological father" means "time spent with the biological father and mother together."

percentages are very close to the actual percentages of 6% and 48% implied by the descriptive statistics shown in Table 3.

The remaining panels of Figure 1 illustrate the impact of the mother's marital status at birth on time spent with the biological father. Marital status at birth is clearly the major determinant of time spent with the biological father, but it is not the only determinant. If the mother was single at the time the child was born, the mean percent of childhood spent with the biological father is 21% for whites and 9% for blacks.[17]

The two most important proximate demographic determinants of the large racial gap in time spent living with both biological parents are the much higher propensity of black women to conceive children outside of a union, and the lower rate of "shotgun" unions for blacks compared to whites. If black women conceived children outside of a union at the same rate as white women, and all other black-white differences in behavior remained the same, the black-white difference in the percent of childhood spent with the biological father would be 35% smaller. If black women entered shotgun unions while single and pregnant at the same rate as whites, other things equal, the gap would be 38% smaller.[18] Other black-white differences in demographic behavior have much smaller effects.

Figure 2 summarizes the underlying transition patterns for entry and exit of the biological father from the child's home, aggregated to the annual level. The annual rate of entry of the

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[17]Marital status at the time of conception plays a dominant role in family structure experiences for children of black mothers, but much less so for children of white mothers. If the mother was single when the child was *conceived*, the mean percent of childhood spent with the biological father is 46% for whites and 14% for blacks (not shown).

[18]See Blau and van der Klaauw (2007) for details on these calculations.

biological father to the child's household (for children at risk of this event) is around 8% for children of white mothers during the first year of life, versus 5.5% for blacks. The rate drops rapidly with age and converges to 2% during the third year and less than 1% during the fifth year for both groups. Thus if a biological father was absent from the child's home at birth he is likely to enter the child's household very early in the child's life or not at all. The rate at which biological fathers move out is about 5% in the first year of life for children of black mothers, versus about 3% for children of white mothers. The exit rate drops with age, more rapidly for blacks, but convergence is quite slow. The monotonic decline in the rate of exit of the biological father as a child ages contrasts with the typical pattern of the hazard of divorce, which *increases* with marriage duration in the first year of marriage before declining (see Waite and Lillard, 1991, for example). Note that the simulations on which the figures are based do not hold constant marital duration, the mother's age, the number of siblings, and other factors that may affect the risk of exit of the biological father.

Figure 3a shows the cumulative duration of time spent with a stepfather. Racial differences are negligible in this case: by age 18, the expected duration of time spent with a stepfather is about three years for blacks and whites. Figure 3b shows the patterns conditional on the mother being single at the time the child was born. A more noticeable racial gap appears in this case: children of white mothers who were single at birth accumulate about 6 years with a stepfather by age 18, compared to about 3 years for children of black mothers. The underlying annualized transition rates shown in Figure 4 indicate that the annual rate of gaining a stepfather rises with age until the teenage years, and is persistently higher for children of black mothers. The exit rate of stepfathers rises with age until about age 5, and is higher for blacks than for whites, suggesting that step families are less stable for blacks than for whites. The greater instability for blacks is masked in the cumulative averages in Figure 3 because blacks experience both greater entry and greater exit of stepfathers than whites. Of all time spent living with the biological mother and a father figure, the simulations indicate that stepfathers account for 13% and 35% for children of white and black mothers, respectively (not shown). The large share for blacks is not a result of a large absolute share of childhood spent with a stepfather: on average 10% of childhood is spent with a stepfather by children of both white and black mothers (not shown).[19] Rather, only 33% of childhood is spent with the biological father by children of black mothers (see Table 4), so stepfathers loom relatively large by comparison.

Figure 5 summarizes annual transition rates of children among union states. Many studies have presented such estimates for adults, but they have rarely been presented for children, and never for the full set of transitions shown here. The annual entry rate to cohabitation is 3-4% for children of black mothers during the preschool years, compared to 6-7% for children of white mothers. The cohabitation entry rate rises until age 8 for whites, and the racial gap falls after age 8. The transition rate from single to married is highest for children of white mothers in the first year of life at 8%, declines to 4% by age 5, and then continues to decline gradually. The pattern for children of black mothers is similar but at a lower level. The rate at which cohabitations are converted to marriage (shown in Figure 5c) is relatively constant during the preschool ages at 18-20% per year for whites and 12-13% for blacks (with Hispanics at 12-13%, not shown). The conversion rate remains roughly constant until the teen years, when it begins to fall for whites.

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[19]These numbers are derived from calculations for all simulated children, regardless of the last age at which they are observed.

The rate at which cohabitations dissolve is 9-10% per year for blacks through age 10, and declining thereafter. The rate for whites is around 8% at age 1 and declines slowly with age. Finally, the annual risk of experiencing a divorce is 5% for blacks through age 7, then declining slowly to about 4% by age 18. The divorce rate for whites is about 2.5% in the preschool years, then declining slowly.

The cohabitation experiences of children have received a lot of attention recently, as a result of the large increase in the incidence of cohabitation. Brown (2004) finds that children living in biological-cohabiting-parent families experience worse outcomes than children residing with married biological parents. Raley and Wildsmith (2004) argue that ignoring cohabitation biases estimates of family instability experienced by children. Our estimates imply that cohabitation accounts for 2.7% of the time spent living with both biological parents during childhood for whites, 14.3% for blacks, and 7.3% for Hispanics (not shown). Thus, despite the relatively high incidence of cohabitation (see Table 2), it accounts for a negligible share of time spent with both biological parents, except for the children of black mothers.

Another issue of considerable interest is the extent of instability in family structure, particularly by the age at which a child experiences a transition. Table 5 shows the mean number of simulated changes in family structure of various types experienced by children in three different age groups: 0-5, 6-11, and 12-17.[20] These age groupings correspond roughly to early childhood, middle childhood, and adolescence. These are hypothesized by developmental psychologists to be distinct stages in the developmental life course, with qualitatively different

[20]The figures in Table 5 may underestimate the true number of family structure changes, as our analysis ignores temporary separations and misses cohabitations that begin and end between interviews before 2000, and cohabitations that ended before the first interview.

effects of family structure (e.g. Hill et al., 2001; Moore et al. 2001). The first row of panel A shows that children of white mothers experienced on average .106 instances of a man entering the household in their first six years, .097 in the second six years, and .046 in the third six years, for a total of .249 from ages 0-17 (a man who lives with the child at the time of the birth is not at risk of moving in). The corresponding figures for children of black mothers are roughly twice as large, and the mean number of “father figure” entries per child is .447 from age 0-17. The next four rows break out entries by the relationship of the man to the child (biological or step father) and by union status at the time of entry (cohabitation versus marriage). The most commonly experienced entry at all ages is a cohabiting step father, with a married step father also common for children of black mothers, and somewhat less frequent for children of white mothers. Entry of the biological father is relatively rare for whites (about 12% of total entries), but not as rare for blacks (about 25% of total entries).

Panel B shows comparable figures for exits of men from the household. The total number of exits per child is much more similar across blacks and whites than in the case of entry, at .318 for whites and .376 for blacks. The composition of exits is quite different, however, with 72% of exits for children of white mothers accounted for by the married biological father, compared to 47% for children of black mothers.

Panel C shows the mean total number of transitions per child, and the frequency distribution of the number of transitions. 69% of children of white mothers experience no family structure transitions during childhood, compared to 52% of children of black mothers. Transitions are about equally likely to occur in the first and second parts of childhood, and less likely to occur in the last third of childhood (although many simulated children are not observed for their entire childhood). About 6% of children of white mothers experience two or more transitions during the preschool years, compared to 7% of children of black mothers.

DISCUSSION

An important question is what explains the black-white differences in demographic behavior that lead to the large gap in time spent living with the biological father. Our analysis is descriptive, so we cannot provide any definitive answers here. We can, however, examine whether the differences can be accounted for by differences in the family background and other characteristics of black and white women. Family background characteristics were not included in the specification because of the large number of additional parameters that would be required (21 per additional variable; see Table A-1). However, with a simpler estimation approach, it is possible to include family background variables. In this approach we estimate eight multinomial logit models independently, one for each state, without accounting for unobserved heterogeneity or integrating over alternative possible sequences of events and spell beginning and ending dates. The simulated share of childhood spent with the biological father based on this approach is very similar to the share based on the more complex estimation approach. This gives us some confidence that this simpler approach will provide a reasonable indication of the importance of family background variables. Using this approach, we added to all of the logit models a set of categorical indicators for the family structure the woman herself experienced at age 14[21]; a dummy variable for immigrants; education of the woman's mother and father (with dummies for

[21]The categories are lived with (1) biological mother and biological father; (2) biological mother and another man; (3) biological mother and no man; (4) another woman and biological father; (5) no woman and biological father; and (6) any other living arrangement.

missing values and the missing cases set to zero); and the woman's number of siblings. There are large differences on average between whites and blacks in many of these variables. For example, 79% of white women lived with both biological parents at age 14, compared to 50% of black women. Parental education is higher for whites by two years for fathers and one year for mothers, and black women have 4.7 siblings compared to 3.1 for whites. Based on these estimates, we find that if black and white women had the same mean background characteristics, the black-white difference in the percent of childhood spent with both biological parents would be 9-19% smaller. Thus, we conclude that differences in family background matter, but they are not a major factor in accounting for the large black-white gap in time spent with the biological father.

We also examined the role of two other key variables, the woman's own education and her cognitive achievement, the latter measured by the percentile score on the Armed Forces Qualification Test (AFQT). We do not model the processes that determine these two variables, and we do not claim to estimate the causal impact of education and cognitive ability on demographic behavior. This exercise should be viewed simply as an attempt to determine how much of the black-white gap in residence with the biological father can be accounted for by these two important observable characteristics of women. As in the case of family background, the black-white difference in the mean AFQT score is substantial: 53.0 for whites versus 22.8 for blacks. However, blacks and whites are much closer on mean completed years of education: 13.7 for whites versus 13.1 for blacks. The estimates indicate that education is a highly significant determinant of all of the demographic behaviors, but the black-white education gap accounts for only 4-5% of the biological father co-residence gap, conditional on equalized family background. Conditional on equalized family background and education, the AFQT gap accounts for 19-49% of the remaining gap, depending on which group's means are used. This leaves some ambiguity about how much of the gap can be explained by cognitive ability, but it suggests that cognitive achievement may be an important source of the gap. These results should be interpreted carefully. The analysis has not identified the main causal factors that determine the gap. For example, we do not know the importance of poor economic prospects of potential mates, cultural attitudes, psychological factors, or other factors.

CONCLUSIONS

Our demographic analysis indicates that children experience family structure changes during childhood at rates that differ considerably by race. While 30% of children of white mothers experience either an exit or entry of a man into the household during childhood, 48% of children of black mothers do so. 6% of children of white mothers never live with their biological father, compared to 49% of children of black mothers. Overall, children of white mothers spend on average 74% of their childhood with the biological father, while children of black mothers spend on average 33% of childhood living with the biological father. Including stepfathers, the figures are 82% and 43% of childhood spent with any father figure present. Despite a high incidence of cohabitation, only 3% (14%) of the total time spent with the biological father of children of white (black) mothers occurs during cohabitation.

We conclude by noting some important limitations of the analysis, in addition to those already discussed. The results are restricted to one cohort of women, and this cohort has not yet completed its childbearing, union formation, and union dissolution behavior. We cannot extrapolate from the results to infer how the family structure experiences of children born to mothers at relatively old ages will be influenced by union formation and dissolution behavior beyond age 45. The choice of whether a child lives with the biological mother is not modeled. All of the simulation results are conditional on children living with their biological mother for their entire childhood. In fact, most children do spend most or all of their childhood with their biological mother, but accounting for time spent by children away from the biological mother would provide a more complete picture of family structure experiences. The same is true for other aspects of family structure, such as the presence of grandparents, step siblings and half siblings in the child's household. Finally, characteristics of men and children (other than number and ages) are not considered in the analysis. We did investigate whether the sex composition of children affected any of the demographic behaviors, but there was no evidence of such effects. It would be quite interesting to extend the analysis to incorporate the choice among "types" of men; for example low versus high education, low versus high income. It would also be useful to explicitly model temporary separations and reconciliations, since these could affect children.

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Table 1: Definition of states and events

State	Preg-nant	Marital status	B_t	N_t	Events for which the woman is at risk
1	no	single	0	NA	2. Conceive with new man. 5. Cohabit with new man. 7. Marry new man
2	no	single	>0	NA	1. Conceive with current man. 2. Conceive with new man. 4. Cohabit with current man. 5. Cohabit with new man. 6. Marry current man 7. Marry new man
3	no	cohabiting	NA	NA	1. Conceive with current man. 3. Become single. 6. Marry current man
4	no	married	NA	NA	1. Conceive with current man. 3. Become single.
5	yes	single	>0	0 (current)	4. Cohabit with current man. 6. Marry current man
6	yes	single	0 or >0	1 (new)	5. Cohabit with new man. 7. Marry new man
7	yes	cohabiting	NA	NA	3. Become single. 6. Marry current man
8	yes	married	NA	NA	3. Become single

Notes: NA indicates Not Applicable.
B_t is the number of children born since the end of the previous union.
N_t=0 if the current pregnancy, if any, is with the current man, and N_t=1 if it is with a new man.
The numbering of events corresponds with the order in which they are listed in the text. In state 6, the "new" man is the man with whom she conceived the current pregnancy. This man became the current man at the time of the conception, but we refer to him as the new man to avoid confusion.

Table 2: Characteristics of Sample Women

	White	Black
Ever married	88.6	61.2
Ever ended marriage conditional on ever married	37.2	52.3
Ever ended marriage conditional on kids present and ever married	25.4	41.7
Ever married more than once conditional on experiencing end of first marriage	60.5	30.0
Ever cohabited	42.2	34.6
Ever move from cohabitation to marriage	34.7	21.2
Percent of cohabitation spells that end in marriage	74.5	59.1
Percent of cohabitation spells that end in marriage, conditional on a child being born during cohabitation	60.0	56.2
Number of children ever born = 0	21.7	18.2
Mean number of children ever born	1.69	1.87
Ever conceive a child while single	21.7	63.9
Ever give birth while single	10.7	58.4
Ever conceive a child while cohabiting	6.8	7.3
Ever give birth while cohabiting	4.0	7.6
Ever conceive a child while married	67.9	34.5
Ever give birth while married	71.5	38.0
Ever had children living with no man (if ever had kids)	36.5	82.3
Ever had children without biol father present (if ever had kids)	38.4	82.5
Ever had children with a stepfather present (if ever had kids)	24.2	35.6
Mean age at first birth	25.1	21.7
Marital status at first birth:		
single	11.4	66.6
cohabiting	3.2	4.0
Married	85.3	29.4
Mean age at last observation	39.0	38.4
Sample size	2,286	1,340

Notes: All entries are sample means (multiplied by 100 for binary variables). Observations are weighted by the inverse of the number of distinct event histories per woman. See the text for description of how multiple event histories are generated. Only conceptions resulting in a live birth are included here. In cases in which the date of an event is uncertain, the earliest possible date is used to compute the statistics. However, this has very little impact on the statistics; we recomputed the statistics using the latest possible date, and they were always within two percentage points of the statistics shown in the table, and usually were identical.

Table 3: Family Structure Experiences of Sample Children in the First 18 years

	White	Black
Mother' marital status at conception:		
Single	15.7	63.9
Cohabiting	4.6	5.2
Married	79.8	31.0
Mother's marital status at birth:		
Single	8.2	58.8
Cohabiting	2.8	5.3
Married	88.9	36.0
Conceived while single: percent born in cohabitation	5.7	2.4
Conceived while single: percent born in marriage	46.5	7.8
Conceived in cohabitation: percent born in marriage	54.4	21.3
Ever lived with no father	30.1	75.7
Percent of time with no father if >0	38.1	65.0
Percent of time with no biological father if >0	58.9	75.7
Ever live with stepfather	17.1	26.8
Percent of time with stepfather if >0	38.9	31.6
Ever live with biological father	94.0	51.6
Percent of time with biological father if >0	82.9	66.8
Ever lived with cohabiting father	13.5	25.9
Percent of time lived with cohabiting father (if >0)	19.0	19.1
Biological father ever moved in (if not present at birth)	25.4	15.3
Biological father ever moved out (if ever present)	24.3	43.8
Step father ever moved in (if biological father ever not present)	53.0	36.2
Step father ever moved out (if ever present)	36.4	49.4
Mean age at last observation	12.7	14.7
Sample size	3,818	2,479

Note: Unit of analysis is a child. See notes to Table 2.

Table 4: Simulated Distribution of Observed Childhood Years Spent with Biological Father

Percentile	White	Black
10	0.8	0
25	4.0	0
50	8.8	0.3
75	13.3	7.4
90	16.8	13.3
Mean	8.8 (73.8%)	4.0 (33.3%)
Percent zero	6.3	48.8

Note: The simulation runs through age 18 or the last observed age, whichever is less. The figures in parentheses next to the means are the mean percent of all observed simulated childhood years spent with the biological father and mother, accounting for the fact that many children are observed for less than 18 years.

Table 5: Family structure changes experienced by children, by age range

	White				Black			
Age of child in years:	0-5	6-11	12-17	Total	0-5	6-11	12-17	Total
A. Man enters household	.106	.097	.046	.249	.212	.148	.087	.447
Step father, cohabitation	.049	.063	.030	.143	.071	.088	.055	.215
Step father, marriage	.028	.032	.016	.076	.043	.049	.030	.121
Biological father, cohabitation	.010	.001	.000	.011	.044	.006	.001	.050
Biological father, marriage	.018	.001	.000	.019	.055	.005	.001	.060
B. Man exits household	.159	.110	.050	.318	.165	.133	.077	.376
Cohabiting step father	.007	.015	.009	.031	.014	.029	.022	.065
Married step father	.006	.019	.018	.043	.014	.046	.037	.091
Cohabiting biological father	.013	.002	.000	.015	.033	.009	.002	.043
Married biological father	.134	.073	.022	.229	.112	.055	.017	.177
C. Man enters or exits	.265	.206	.096	.567	.378	.282	.164	0.823
Number of transitions								
0	.803	.843	.923	.695	.703	.778	.867	.524
1	.139	.116	.061	.140	.228	.172	.106	.255
2	.049	.035	.014	.103	.060	.043	.024	.135
3+	.009	.007	.002	.062	.010	.007	.004	.086

Notes: All entries except in the last four rows are the mean number of transitions of the indicated type experienced by a child during the ages shown in the column headers. The last four rows show the distribution of the total number of transitions per child. The calculations are not conditioned on father presence or marital status at birth.

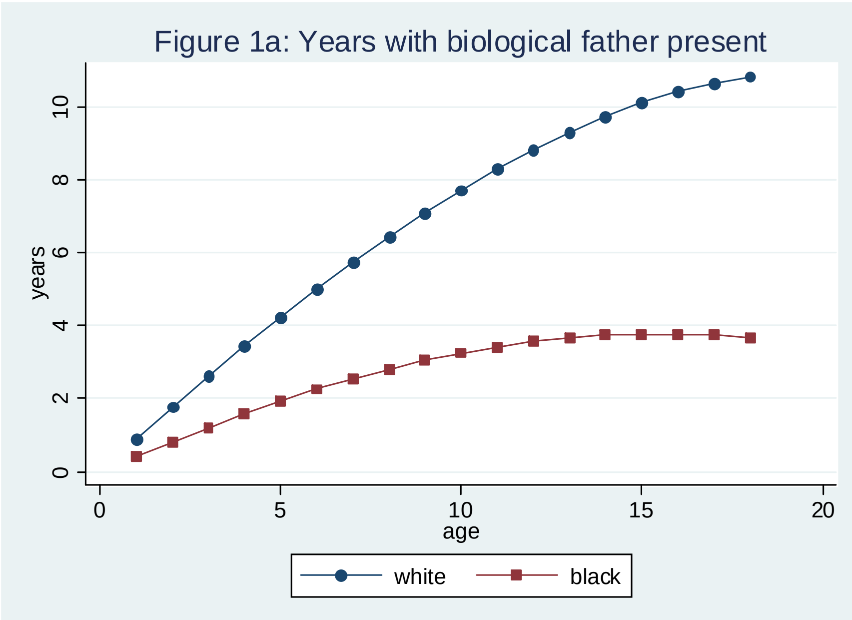


Figure 1a: Years with biological father present

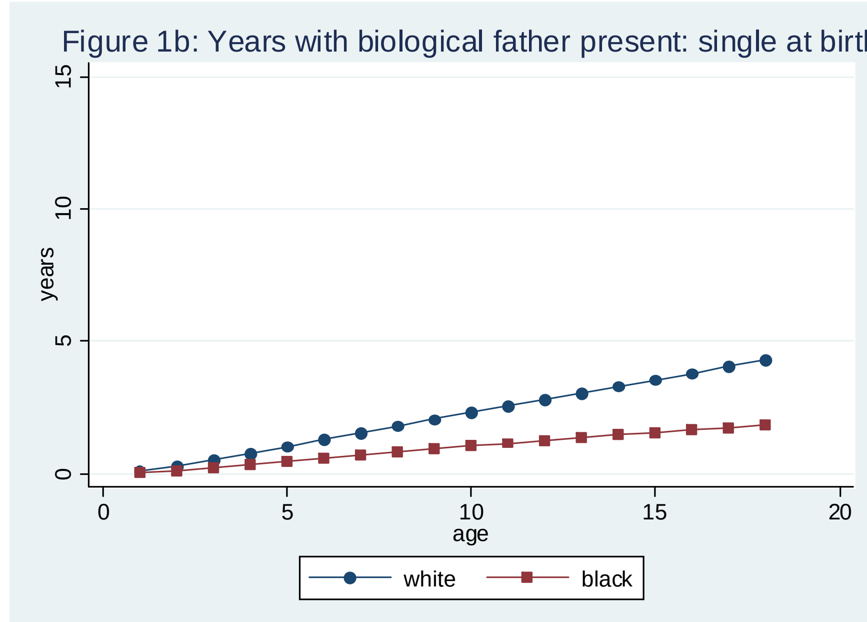


Figure 1b: Years with biological father present: single at birth

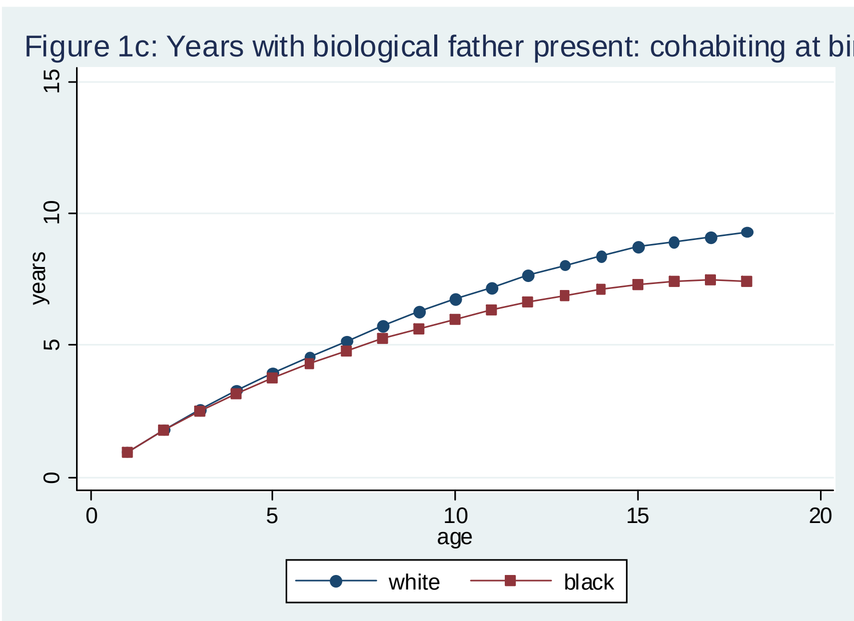


Figure 1c: Years with biological father present: cohabiting at birth

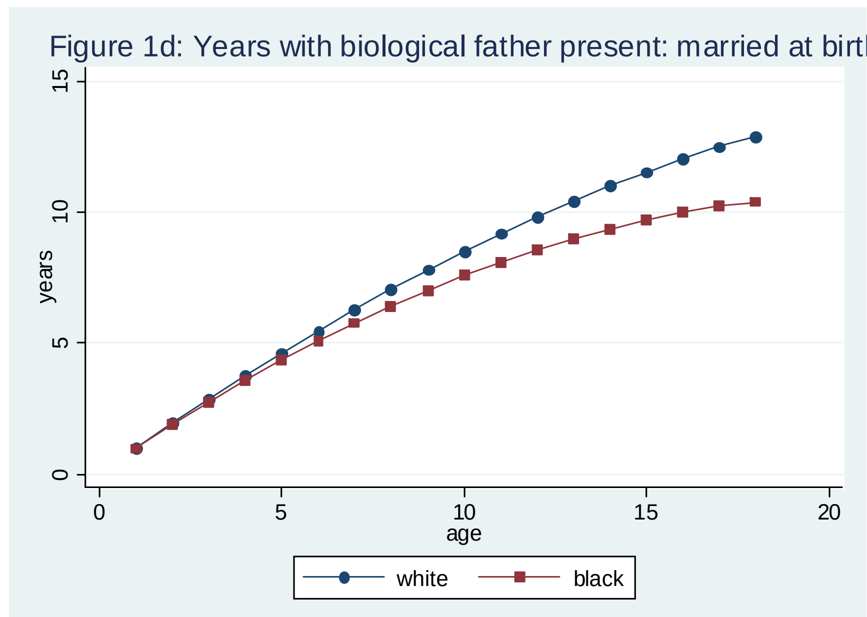


Figure 1d: Years with biological father present: married at birth

Figure 2a: Annual rate of entry of biological father [if at risk]

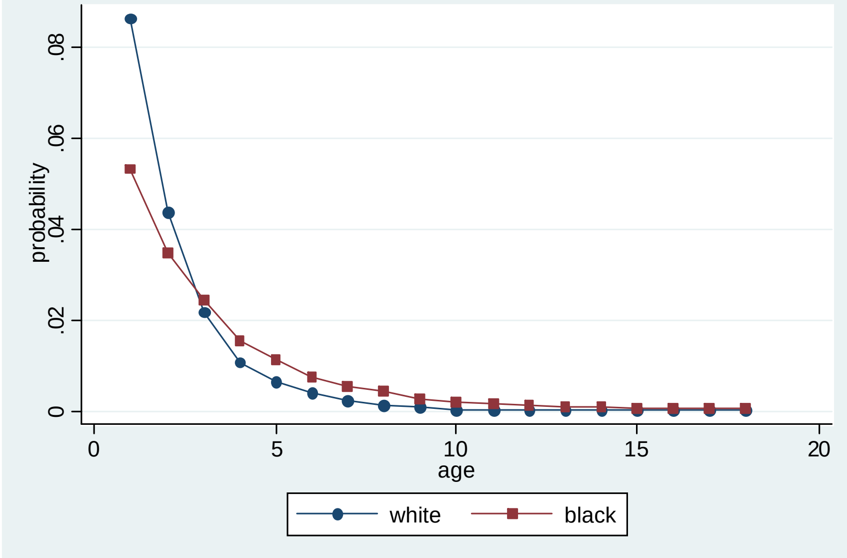


Figure 2b: Annual rate of exit of biological father [if at risk]

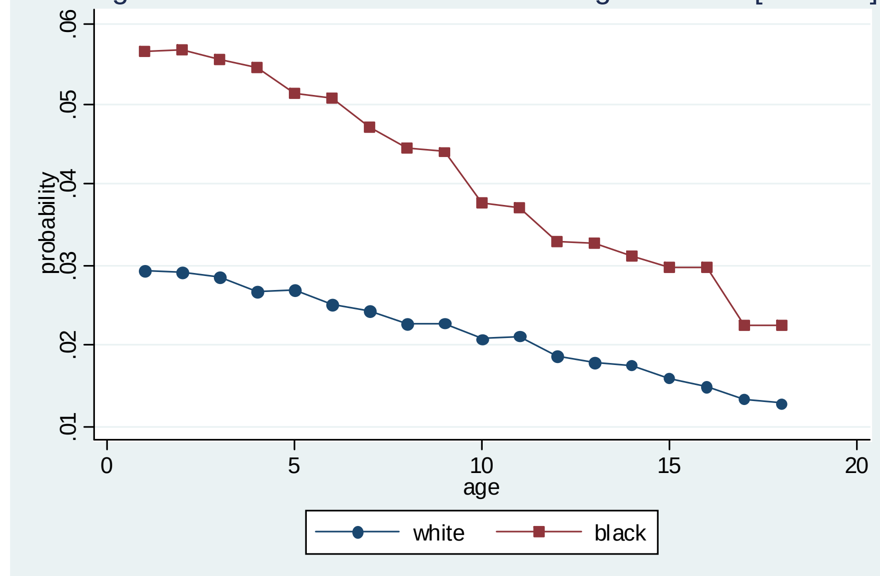


Figure 3a: Years with stepfather present

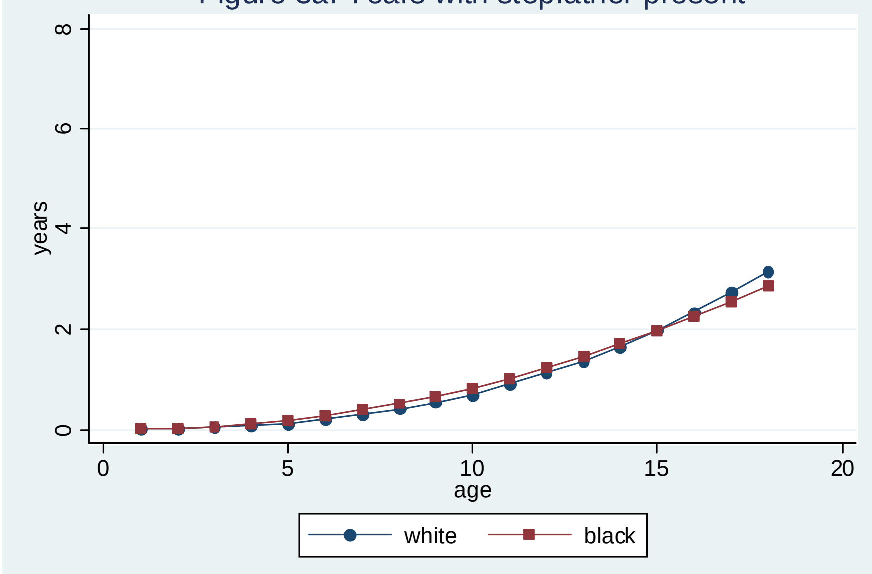


Figure 3b: Years with stepfather present: single at birth

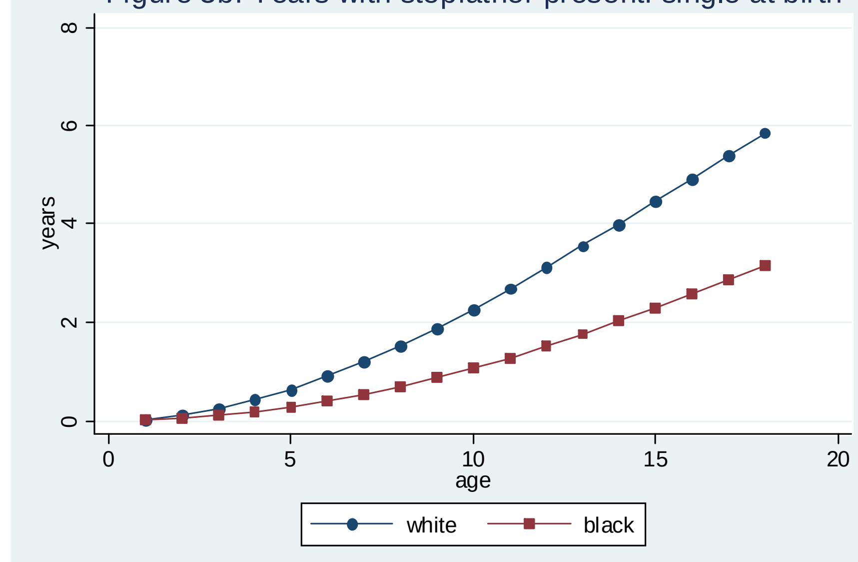


Figure 4a: Annual rate of entry of stepfather [if at risk]

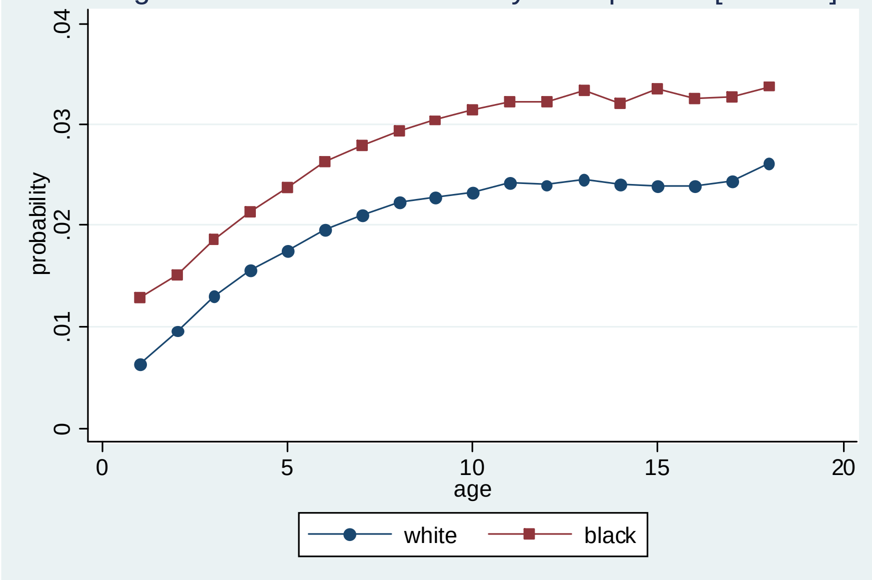


Figure 4b: Annual rate of exit of stepfather [if at risk]

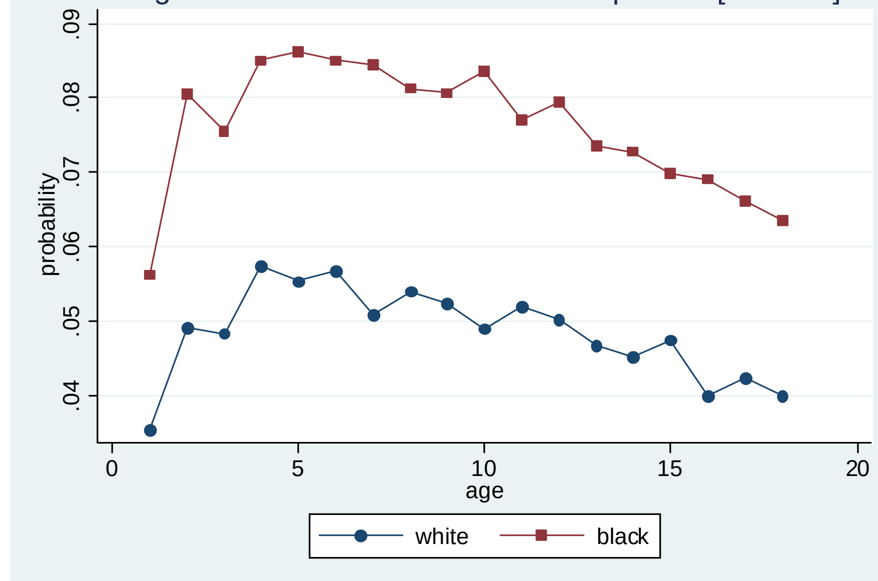


Figure 5a: Annual transition rate: single to cohabiting [if at risk]

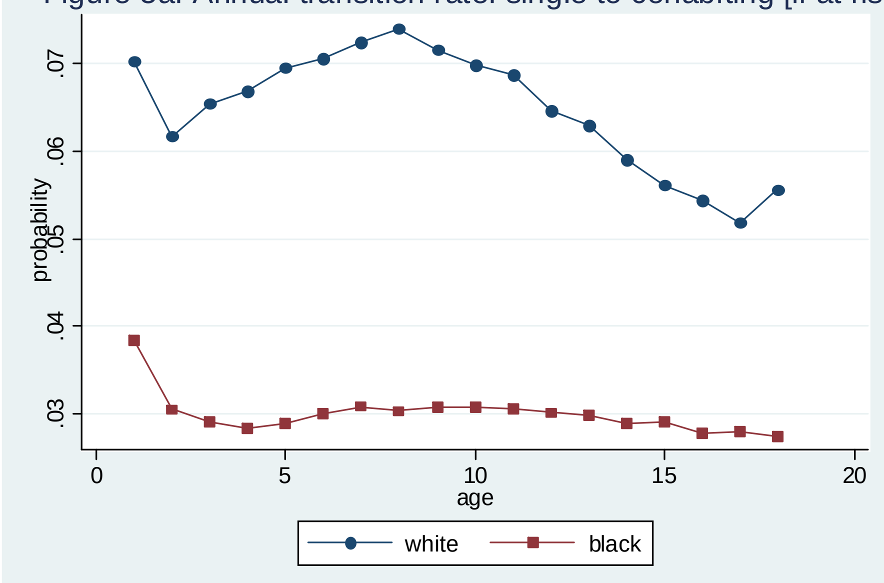
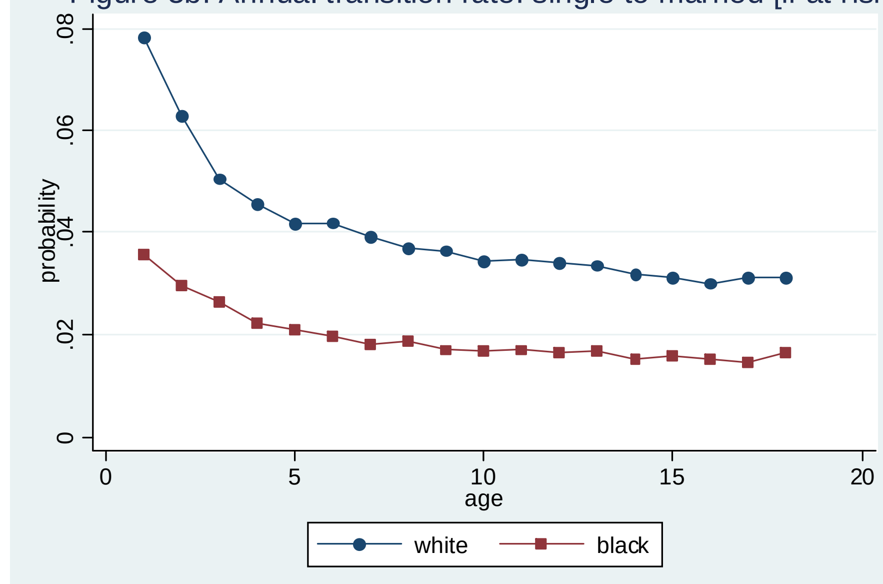


Figure 5b: Annual transition rate: single to married [if at risk]



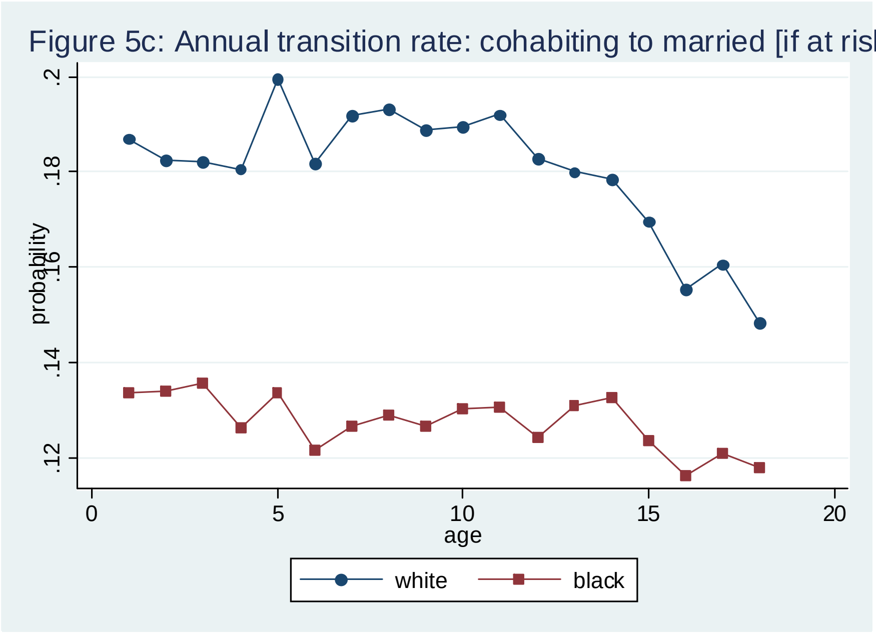


Figure 5c: Annual transition rate: cohabiting to married [if at risk]
probability
.12
.14
.16
.18
.2
0
5
10
15
20
age
white
black

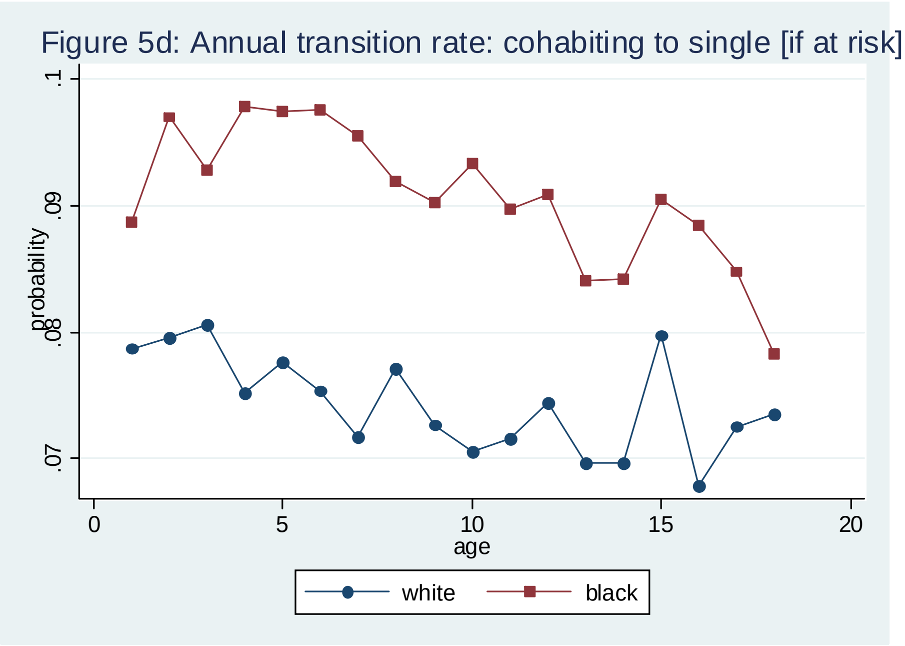


Figure 5d: Annual transition rate: cohabiting to single [if at risk]
probability
.07
.08
.09
.1
0
5
10
15
20
age
white
black

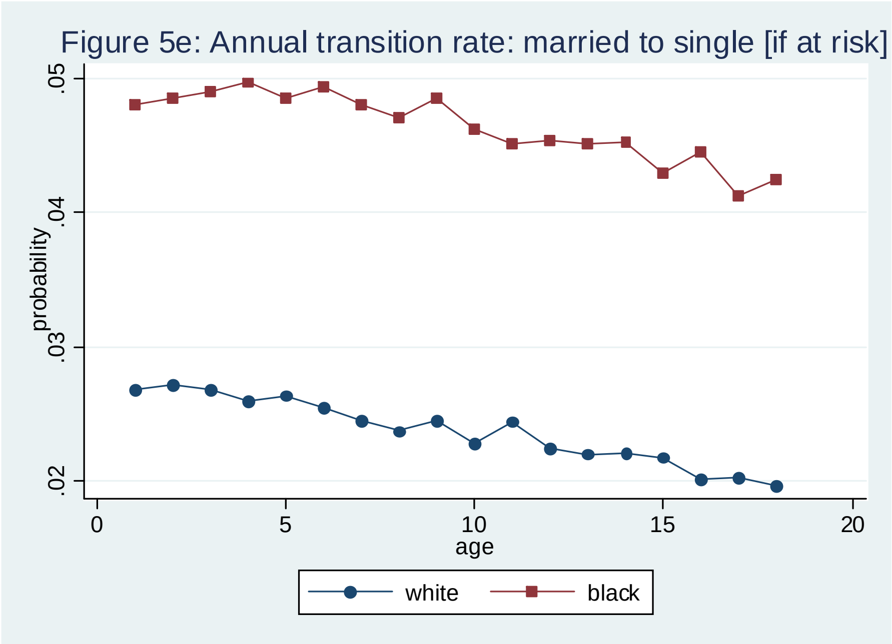


Figure 5e: Annual transition rate: married to single [if at risk]
probability
.02
.03
.04
.05
0
5
10
15
20
age
white
black

Table A-1: Coefficient Estimates

	state 1 Single, not pregnant, B_t=0			state 2 Single, not pregnant, B_t>0						state 3 Cohabiting, not pregnant		
event:	2 Conceive, new man	5 Cohabit, New man	7 Marry, New man	1 Conceive, cur. man	2 Conceive, new man	4 Cohab, new man	5 Cohab., cur. man	6 Marry, new man	7 Marry, cur. man	1 Conceive, Cur. man	3 Become single	6 Marry cur. man
Intercept	**-14.745**	**-20.637**	**-18.500**	**-4.046**	**-8.113**	**-10.150**	**-14.295**	**-5.483**	**-10.255**	-1.967	**-2.266**	**-6.670**
Numfath										0.168	**-0.201**	-0.177
Num marr.	-0.173	**-0.283**	-0.149									
Num cohab.	0.016	-0.146	**-0.582**							-0.203	**0.309**	-0.139
Prev marr.	0.281	**0.532**	**-0.733**									
Prev. cohab.	0.226	**0.723**	**-0.687**									
Age youngest				0.322	1.106	**-4.308**	**1.035**	-0.625	**2.983**	0.246	0.302	-0.292
Age oldest	-0.360	**-1.043**	**-1.406**	0.991	-0.008	-0.885	-0.841	**-1.713**	**-1.758**			
Age mother	**4.224**	**8.183**	**7.023**	-0.220	1.576	**2.975**	**4.821**	1.095	2.475	-1.236	**-1.241**	**2.264**
Dur. of cohab.										-0.737	**2.378**	**-2.331**
Dur. Single	**1.591**	**1.133**	**1.710**									
Agey sq.				**-3.031**	-0.347	**1.369**	**-0.485**	-1.454	**-1.035**	-0.479	-0.090	0.152
Ageo sq.	**-0.339**	**0.299**	**0.398**	-0.259	-0.232	0.349	0.303	**0.834**	**0.622**			
Age mom sq.	**-0.659**	**-1.115**	**-0.925**	-0.120	-0.300	**-0.384**	**-0.682**	-0.267	-0.449	0.101	0.110	**-0.324**
Dur coh. sq.										-0.151	**-1.732**	**0.633**
Dur sing. sq.	**-0.535**	**-0.388**	**-0.724**									
Black	**1.383**	**-0.789**	**-0.637**	**0.616**	0.208	**-0.766**	**-0.815**	**-0.952**	**-0.741**	**0.367**	**0.318**	**-0.287**
Hispanic	**0.672**	**-0.437**	**0.125**	**0.714**	0.028	-0.112	-0.129	**-0.717**	-0.586	**0.297**	0.006	**-0.341**
Hisp*Prevmarr	0.192	0.195	-0.256									
Hisp*prevcoh	-0.014	0.213	-0.130									
Factor Load	**2.378**	**1.510**	**1.730**	-0.287	0.372	0.157	**0.780**	-0.018	0.908	**0.513**	**-0.232**	-0.075
No. of events	1,844	1,629	2,197	429	392	155	231	174	145	403	582	1,156
No. of monthly observations		709,514					116,287				55,372	

	state 4 Married, not pregnant		state 5 Single, pregnant, N_t=0		state 6 Single, pregnant, N_t=0		state 7 Cohabiting, pregnant		state 8 Married, pregnant
event:	1 Conceive, cur man	3 Become Single	4 Cohabit, Cur. man	6 Marry cur. man	5 Cohabit, New man	7 Marry, New man	3 Become single	6 Marry, Cur. man	3 Become single
Intercept	**-3.014**	**-5.813**	**-5.669**	0.541	**-16.802**	-1.041	**-5.306**	**-2.360**	**-5.962**
Numfath	**-0.330**	**-0.245**					0.349	**-0.767**	
Mother's DoB					0.450	**-0.567**			
Num. kids									**0.297**
Num marr	-0.094	-0.129							
Numc					**-1.126**	**-1.410**			

Prev marr					-0.465	**-1.441**			
Prev cohab					-0.368	**-2.165**			
Age youngest	**2.036**	**0.661**							
Age mother	0.243	0.280	0.197	**-1.146**	**4.965**	0.634			**-0.692**
Dur. marriage	0.289	**2.538**							
Dur. single					0.287	1.279			
Dur. union	-0.203	**-2.379**							
Dur. pregnancy					**0.586**	**0.716**			
Age young. sq.	**-1.780**	**-0.219**							
Agemom sq.	**-0.128**	-0.080			**-0.626**	-0.014			
Dur marr sq.	-0.080	**-1.244**							
Dur single sq.					-0.165	**-0.683**			
Dur union sq.	-0.329	**1.039**							
Dur preg sq.					**-0.075**	**-0.091**			
Dur preg = 1				0.172					
Black	**-0.115**	**0.596**	-0.461	**-1.636**	**-1.520**	**-2.028**	-0.024	**-0.992**	**1.263**
Black*agey	**0.532**	-0.017							
Black*agey sq	**-0.756**	-0.123							
Hispanic	**0.120**	0.088		-0.813	-0.209	**-0.536**	-0.772	**-0.784**	0.519
Hisp*prevmarr					-0.101	-0.860			
Hisp*prevcoh					-0.064	0.442			
Factor Load	**-0.490**	**0.609**		-0.926	0.149			0.120	**1.264**
Prob. weight	**0.509**								
No. of events	4,945	1,632	18	41	131	453	26	190	84
No. of monthly observations	503,928		3,783		19,390		3,875		53,223

Notes: See Table 1 for complete definitions of the states and events. Coefficient estimates in bold are significantly different from zero at the 5% level.

B_t is the number of children born since the end of the previous union.
$N_t=0$ if the current pregnancy is with the current man, and $N_t=1$ if it is with a new man.

DoB = Date of Birth.
Numfath = number of children fathered by the current man (if currently in a union).
Numc = number of children fathered by the current man (if not currently in a union).
Dur union = duration of the current union, including both the cohabitation and marriage (if currently married and cohabited before the marriage began).
Prev marr = 1 if the union status before the current single spell was married.
Prev cohab = 1 if the union status before the current single spell was cohabitation.
Dur preg = 1 is a dummy if currently in the first month of a pregnancy.

All ages and durations are measured in months. All except the duration of pregnancy are divided by 100. All quadratics in age and duration (except pregnancy) are divided by 10000.

The mother's date of birth is measured in months since January 1900, divided by 100.
The random effect was omitted in three of the 21 transitions, as no support was found in the data for its addition to these transitions (the inverse of the outer product of the likelihood derivatives became ill behaved, so that the model would not converge). The two mass points for the discrete random effect were set equal to 0 and 1.
The number of events observed in each state and the number of monthly observations are weighted by the inverse of the number of distinct event histories per woman.