The Impact of Social and Economic Policy on the Family Structure Experiences of Children in the United States

David Blau

Department of Economics and Carolina Population Center
University of North Carolina at Chapel Hill

Wilbert van der Klaauw

Federal Reserve Bank of New York

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Corresponding author: David Blau, Department of Economics, Gardner Hall, University of North Carolina, Chapel Hill, NC 27599-3305. (Tel) 919-966-3962. (Fax) (19-966-4986. David_blau@unc.edu

Abstract

We use panel data from the 1979 cohort of the National Longitudinal Survey of Youth (NLSY79) to estimate the effects of policy, labor market, and marriage market contextual variables on the fertility, union formation, union dissolution, type of union (cohabiting versus married), and father identity (biological versus step) choices of women born from 1957 to1964. We follow these women from the early 1970s as they enter adolescence through 2004, when they are in their 40s. We specify a model that can be used to trace through the consequences of these demographic behaviors for the family structure experiences of children. We allow the effects of several of the contextual variables to differ for whites, blacks, and Hispanics. The evidence presented suggests that the family structure experiences of the children of the NLSY79 have been influenced by the tax gains associated with childbearing, welfare reform, unilateral divorce laws, and the wage rates available to men and women in the labor market. The results suggest that other contextual variables such as child support enforcement efforts, the level of welfare benefits, the tax gain to marriage, the sex ratio, and the unemployment rate had little impact on the family structure experiences of these children.

1. Introduction

The most prevalent type of family in which children in the U.S. are raised today is the traditional two-biological-married-parent family. But in the past 30 to 40 years it has become increasingly common for American children to spend a significant portion of their childhood in alternative family structures. These alternatives include residing with a divorced, separated, or never-married mother without a man present; residing with one biological parent and a step-parent; residing with unmarried cohabiting parents; and residing with neither biological parent. The consequences for children of growing up in non-traditional family structures have received enormous attention from social scientists and policy makers. Children who grow up in a family in which both biological parents are present and married experience better education, employment, marriage, childbearing, and psychological outcomes than do their counterparts who spend substantial parts of their childhood living in alternative family structures. There is considerable evidence that at least part of the association between family structure and child outcomes is causal. There is a much still to be learned about the consequences of growing up in alternative family structures, but there is a consensus that family structure has important consequences for children.

In contrast, there is much less known about the causes of the rather dramatic changes in family structure observed in the U.S. in the past 30 to 40 years. In a recent authoritative review of this issue, Ellwood and Jencks (2004) argue that "The spread of single-parent families has been both an intellectual challenge and a source of persistent frustration for social scientists. ... there is still no consensus about why single parenthood spread, much less about why it spread faster in some populations than in others." (p. 25). They provide an extensive discussion of theoretical and empirical approaches based on the "standard economic model" derived from Becker (1981), which emphasizes economic incentives to marry. Based on their review of the literature, they find that the explanatory factors emphasized by this approach - male wages, female wages, public assistance, and sex ratios - have important effects on single parenthood, but that changes over time in these factors cannot account for most of the observed changes in family structure. Hence, the frustration noted above. However, Ellwood and Jencks offer several constructive recommendations for research in this field.

First, they emphasize that a major feature of change in recent years has been de-linking of marriage and childbearing decisions, particularly for certain population groups. Hence it is crucial to recognize that marriage and childbearing are in fact distinct decisions, and that treating "single parenthood" as one decision rather than the consequence of a number of distinct choices will often not be a productive analytical or empirical approach. Recent theoretical analyses by Akerlof, Yellen, and Katz (1996) and Willis (1999) recognize this point and offer alternative models to explain the rising prevalence of single parent families. Second, Ellwood and Jencks point out the importance of distinguishing between delay and permanent avoidance of marriage and childbearing. In some cases, the major changes have been in the timing of childbearing and marriage, while for others the most important aspect of change has been more radical, namely avoiding marriage and/or childbearing altogether. The standard economic model has little to say about timing. And most empirical analyses do not come to grips with this issue: they are either explicitly focused on outcomes at certain ages (e.g., marriage by age 24, or non-marital childbearing by age 19), or they look at marital and childbearing transitions over short periods of time. Ellwood and Jencks also make the case for including cohabitation in the analysis, avoiding restrictive functional forms in dynamic models (i.e. proportional hazards), and explaining differential family structure change by education and/or family background as well as race.

In this paper, we take up the challenge offered by Ellwood and Jencks, and propose a new approach to analyzing the determinants of family structure change. Our approach has four distinguishing features that in combination make it unique. First, we jointly model union formation, union dissolution, and childbearing decisions. Previous analyses have integrated some of these behaviors in a single model, but none have integrated the full range of behaviors needed for a thorough analysis of family structure change. Second, the analysis is dynamic, and distinguishes between the short run timing effects and the long run "avoidance" effects of key driving forces. Third, we consider all of the major proposed explanations, including changes in public assistance policy, child support enforcement, divorce laws, and tax laws; changes in labor market opportunities facing both men and women; and changes in marriage market conditions. By considering the main proposed driving forces jointly rather than focusing on one or two in isolation from the others, as in much of the literature, we provide a more robust accounting of the factors driving family structure changes. Fourth, and perhaps most important, we model the behavior of the adults who make union and childbearing decisions, but we derive from the model the consequences of these decisions for the family structure experienced by children. Thus we explicitly account for and model choices that are relevant for determining the *identity* of men who are in the mother's household, from the perspective of children: biological father or step father. This approach is unique in the literature on family structure changes.

In order to accomplish this rather ambitious agenda, we take an analytical approach that is quite different from previous research. Most previous studies analyze the behavior of a succession of birth cohorts in a time-series-of-cross-sections, or follow a few birth cohorts over a short period. Our approach is to follow the behavior of a small number of birth cohorts over a long period. Specifically, we use panel data from the 1979 cohort of the National Longitudinal Survey of Youth (NLSY79) to analyze the fertility, union formation, union dissolution, type of union (cohabiting versus married), and father identity (biological versus step) choices of women born from 1957 to1964. We follow these women from the early 1970s as they enter adolescence through 2004, when they are in their 40s. The rich event history data available in the NLSY79 makes it possible to provide an integrated analysis of all of the key behaviors that determine the family structure experiences of children. We account for dynamics in a rich and flexible way, and our analysis allows for unobserved heterogeneity across women. We analyze the effects of state-year-specific policy, labor market, and marriage market variables over a three decade period. There is substantial variation over time and across states in many of these variables, providing leverage for identifying the effects of these factors. We allow the effects of the contextual variables to differ for whites, blacks, and Hispanics.

There are some limitations of our data and analytic approach. For a given cohort, age and calendar time do not vary independently. Thus, for example, welfare reform occurred in the 1990s, when the NLSY79 cohort was well past the teenage years, so our approach cannot provide a credible estimate of the impact of welfare reform on the behavior of teenagers. But most of the other explanatory variables vary over time and across states throughout the period of our analysis. Thus we view this as a relatively minor disadvantage compared to the richness of the analysis that is possible using our analytical approach together with the NLSY79 data. A drawback of analyzing only a few birth cohorts is that we cannot easily use our results to provide an accounting of changes over a wide range of cohorts. But this is mainly a result of lack of comparable data for other cohorts. We do use our results to simulate the impact on the NLSY79 cohort of a series of counterfactual experiments involving alternative policy and labor market scenarios. A limitation of our model is that it contains a large number of parameters that must be jointly estimated. Even with a sample of about 4,500 women and the 8,000 children they have borne, it is not feasible to allow all of the parameters to vary freely. Hence, a number of restrictions must be imposed for statistical identification. Finally, a limitation of the NLSY79 data is that we have little information on children who do not live with the biological mother, so our analysis is limited to children who live with their biological mother. As discussed below, most children do live with their biological mother for most of their childhood, so this is not a significant drawback.

Our results indicate that some of the contextual variables have had fairly sizeable effects on the family structure experiences of children, and in many cases the effects differ substantially for children of white, black, and Hispanic women. For example, a higher female wage rate reduces the risk of the biological father entering the household and increases the risk of his exit for black and Hispanic women, but not for whites. A one standard deviation increase in the mean female wage offer, holding the male wage rate and other contextual variables constant, is estimated to reduce the cumulative (up to age 18) amount of time spent with the biological father present by over one year for children of black mothers, from an already low base of about four years. The main union transition rate changes that produce this decrease are an increase in the divorce rate and a higher rate of dissolution of cohabitations. The male wage rate also has some sizeable effects, almost always in the opposite direction of the effects of the female wage rate.

Welfare reform is estimated to have has small effects on family structure for children of white and Hispanic mothers, but a large impact for children of black mothers. Welfare reform reduced the rate of entry and substantially increased the exit rate of the biological father in black families, and also significantly increased the risk of entry of a stepfather. The net effect was a decrease of about 2.5 years in time spent with the biological father and an increase of about two years in time spent with a step father. Unilateral divorce laws and the tax gains associated with childbearing were estimated to have some important effects as well. The effects of child support enforcement are generally quite small across all outcomes. There are also small effects of the sex ratio and the level of welfare benefits.

We provide background and a selective literature review in Section 2. In section 3 we specify the model and econometric approach. Section 4 describes the data, section 5 presents the results, and section 6 concludes.

2. Background and Literature

The changes in family structure that are of interest here have been the result of a decline in marriage, increases in divorce and cohabitation, and an increased rate of childbearing outside of marriage. These changes are well-known and have been discussed extensively elsewhere (Bumpass and Lu, 2000; Bumpass, Sweet, and Cherlin, 1991; Cherlin, 1999; Fields and Casper, 2001; Martin et al., 2002). Here, we discuss their consequences for the family structure experiences of children, and then summarize findings from the literature on the causes of the changes.

Kreider and Fields (2005) summarize recent family structure patterns of children using data from the Survey of Income and Program Participation (SIPP). In 2001, 62% of children under the age of 18 were living with their married biological parents. Another 2.5% were living with their cohabiting biological parents. Seven percent of children were living with one biological parent and one step or adoptive parent (in 84% of these cases the biological parent was the mother). Twenty five percent were living with one parent only (in 88% of these cases, the parent was the mother). Finally, 4% were living with neither parent. For most of the 20th century up to 1970, the percentage of children living in a two parent family remained stable at 83-85%. The percentage living in a one parent family grew slightly from 10-11% in the first half of the century to 13% in 1970, and the percentage living with no parent fell from 5-6% to 3% in 1970 (Kreider and Fields, 2005). Between 1970 and 1990 the percentage in two parent families fell from 85 to 73% and the percentage in one parent families rose from 12 to 25%, with little further change since 1990. Thus as noted above, the traditional family structure remains the most common experience of children, but it is significantly less prevalent than in the past. The childbearing and union formation experiences of the NLSY79 cohort that we analyze began in the early-to-mid 1970s, coinciding with the onset of the major changes in these behaviors.

An important dimension of family structure change is variation by race and ethnicity. In 2001, 69% of non-Hispanic white children lived with both biological parents, compared to 32% of non-Hispanic black children, and 62% of Hispanic children (Kreider and Fields, 2005).[1] The stark contrast between blacks and whites has received significant attention in the literature. But Ellwood and Jencks (2004) note that there have also been large differences in family structure changes by various measures of family background, such as parental education and income.

Theories of union formation, union dissolution, and childbearing behavior emphasize several key observable explanatory variables, in addition to less easily observed factors such as preferences,

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[1]Non-Hispanic will be implicit henceforth when referring to whites and blacks.

attitudes, and norms. These include the wage rates available to men and women; the tax penalties or subsidies for marriage and childbearing; the generosity and terms of public assistance to low income families with children; the state of the marriage market; and the legal environment governing divorce and enforcement of child support obligations[2]. We briefly discuss findings from the literature on each of these explanatory factors.

Wage rates. Becker's (1981) theory of marriage implies that the difference in potential wage rates between men and women affects the gains from specialization within marriage. The higher a woman's potential wage rate, the greater is the opportunity cost of staying home and raising children. The higher a potential husband's wage rate relative to the woman's wage rate, the greater is the incentive to marry in order to reap the gains from specialization within marriage. A number of studies have found a negative effect of the average male wage rate and a positive effect of the average female wage rate on the prevalence of female headship. However, trends in wages do not contribute much to explaining the trend in single headship during the 1970s to 1990s.[3] It has been argued that expansion of the Earned Income Tax Credit (EITC) in the 1980s and 1990s caused an increase in the marriage tax penalty (Hotz and Scholz, 2003). However, empirical analyses have found little evidence that the EITC

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[2]Many studies have examined the impact of abortion legalization and the increased availability of oral contraceptives on demographic behavior. We do not focus on these factors in this study because both the legalization of abortion and the diffusion of easy access to oral contraceptives were completed by the early 1970s, before most of the women in our sample began childbearing and union formation.

[3]See, for example, Blau, Kahn, and Waldfogel (2000), Fitzgerald and Ribar (2004); Bitler et al. (2004), and Moffitt (2001).

has influenced marriage decisions.[4] The effect of male and female wage rates on fertility has also been studied; see Francesconi (2002) and references cited therein.

Welfare. Many studies have analyzed the effect of welfare benefits on various aspects of family behavior, including union formation, union dissolution, and single motherhood. Moffitt (1998) reviewed this literature thoroughly and concludes that there is evidence of a positive association between welfare benefits and female headship, but the magnitude and precision of the estimated effect are rather sensitive to specification. Furthermore, the trend in real welfare benefits in the 1980s and 1990s was downward, which should have led to a decline in female headship rather than the increase that was observed. Some recent studies have found more consistent evidence of a positive association between welfare benefits and female headship among disadvantaged young women, for whom welfare is likely to be a relevant option (Rosenzweig 1999; Foster and Hoffman, 2001; Hoffman and Foster, 2000). Blau, Kahn, and Waldfogel (2004) find no evidence that welfare benefits affect the likelihood that a young woman is a single mother. Their results indicate that higher welfare benefits increase the likelihood that a single black mother is the head of her own household.

A recent literature examines the impact of welfare reform on family structure. This literature analyzes the impact of the introduction of time limits, family caps, work requirements, more generous earnings disregards, and other reforms allowed by state welfare waivers in the late 1980s and early to mid 1990s, followed by enactment of the federal Personal Responsibility and Work Opportunity

[4]Dickert-Conlin and Houser (1998); Ellwood (2000). There is no evidence on whether the EITC has influenced fertility. Other features of the tax code that result in marriage and childbearing penalties (or subsidies) have also been analyzed, with results generally suggesting small effects in the expected direction (see Alm and Whittington, 2003).

Reconciliation Act (PRWORA) in 1996. The majority of studies find that welfare reform caused an increase in marriage and a decrease in divorce (e.g. Acs and Nelson, 2004; Bitler et al., 2006; Gennetian and Miller, 2004) . However, evidence from social experiments undertaken as part of welfare reform shows no consistent impacts on union formation in the welfare population (Harknett and Gennetian, 2003), and there is evidence from a couple of studies that welfare reform actually caused a decrease in marriage (Bitler et al. 2004; Kaestner et al., 2003). Fitzgerald and Ribar (2004) find no significant impact of welfare reform on female headship, using SIPP data. Kearney (2004) finds no evidence that the family cap affects fertility. Lopoo and DeLeire (2006) report evidence that welfare reform caused the fertility rate of girls aged 15-17 to decline relative to the fertility rate of 18 year olds. This is attributed to the feature of welfare reform that required teenage mothers under age 18 to live with a parent or legal guardian in order to be eligible for welfare benefits. We include a measure of welfare reform in our analysis, but the women of the NLSY79 cohort were between the ages of 25 and33 when welfare reform began in the 1990s, so we cannot claim to be able to evaluate the impact of welfare reform at younger ages.

Divorce laws and child support enforcement. Many studies have analyzed the impact of unilateral divorce laws on the divorce rate and related outcomes. Peters (1986) finds no impact, but Friedberg (1998), Gruber (2004) and others do find a positive association. Wolfers (2004) reconciles these differences by showing that there is a positive short run impact of unilateral divorce but apparently no long run impact. This finding suggests the importance of dynamic considerations. Alesina and Giuliano (2006) find evidence that unilateral divorce reduces out of wedlock fertility, with no impact on marital fertility. They interpret this as indicating that when it is easier to escape marriage, women who plan to have a child are more willing to do so within marriage.

A large literature examines the impact of changes over time in state child support enforcement (CSE) laws on demographic outcomes, using a variety of different measures of the strictness of these laws and the means used to enforce them. Most findings indicate that tougher CSE reduces the rate of out of wedlock childbearing, the divorce rate, and the prevalence of single mother families.[5]

Marriage market. Blau, Kahn, and Waldfogel (2000) analyze the impact of marriage market conditions on the prevalence of marriage among young women (ages 16-24). They find that a greater net supply of women in a given education/race group reduces the likelihood of marriage for white women but not for blacks. Marriage market conditions at ages 25-34 affect marriage for both whites and blacks. Wood (1995) finds a positive effect of the supply of "marriageable" black men on the prevalence of marriage among black women. However, the decline in the supply of such men in the 1970s can explain only a tiny fraction of the decline in marriage among black women.

This brief summary of a large literature suggests some possible reasons for the lack of consensus about the causes of family structure change noted by Ellwood and Jencks (2004). Many studies analyze one explanatory factor at a time, and focus on only one or two outcomes. The time period and population analyzed vary widely across studies, as does the set of control variables. There has not yet been a study that examines all of the major proposed explanatory factors in a framework that allows the impact of these factors to be traced through from the demographic behavior of parents to the resulting family structure experiences of children. Our analysis fills this gap, and goes beyond

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[5]See Acs and Nelson (2004), Aizer and McLanahan (2006), Carlson et al. (2004), Case (1998), Garfinkel et al. (2004), Nixon (1997), and Plotnick et al. (2004).

previous studies by incorporating the distinction between biological and step fathers. The dynamic specification and longitudinal data allow us to model state and duration dependence as well contemporaneous and dynamic interdependence between fertility and union choices. Controlling for unobserved heterogeneity across women allows us to give a causal interpretation to the effects of previous demographic decisions on current choices. Controlling for state and year fixed effects allows us to separate the impact of the explanatory factors of interest from other unobserved aggregate trends and cross-state differences.

3. Model

Our goal is to understand the family structure experiences of children who reside with their biological mother. The family structures of interest are living with the biological mother and (1) the married biological father, (2) the cohabiting biological father, (3) a married step father, (4) a cohabiting step father, and (5) no man.[6] We assume that women become at risk of entering a union and conceiving a child at age 12. We use a discrete time framework in which the unit of time is a month. In a given month (t), woman i's situation is characterized by a set of fixed characteristics X_i such as her race, ethnicity, and year of birth; the outcomes of previous childbearing and union formation and dissolution decisions Y_{it}, such as the number of children born and their ages, current marital and cohabitation status, and marital and cohabitation history; and a set of policy, labor market, and marriage market variables Z_{ijt}, some of which may be choice-specific (j is the indicator for choices, defined below). We do not

[6] We do not distinguish living arrangements by the presence of grandparents or other non-parental adults (see Bitler et al., 2006, and DeLeire and Kalil, 2002).

model schooling and employment decisions, and we do not condition on the outcomes of previous education and employment decisions. We also do not model migration behavior, but we do condition on the state of residence.

Each period, a woman faces a set of childbearing and union options, from which she can choose one. Here and in what follows, a "union" refers to a co-residential romantic relationship, which may be a marriage or a cohabitation. For empirical tractability, we assume that at most one alternative can be selected from the choice set in a given month. The set of alternatives available to a woman in a given period depends on her previous choices. For example, if she is currently married, then the option of entering a marriage or cohabitation is not available. If she is currently pregnant, then conceiving a child is not an option. We assume that if she is in a cohabitation, then the only man whom she can marry in the current month is her partner. We also assume that if she is currently in a union, then the only man with whom she can conceive a child is her current spouse or partner. Let $A(Y_{it})$ denote the set of alternatives available to a woman given the values of her state variables in period t. The alternatives are specified below. The value to a woman of choosing alternative j is specified as

$$V_{ijt}{}^{*} = \beta_{1\mathrm{j}} X_i + \beta_{2\mathrm{j}} Y_{it} + \beta_{3\mathrm{j}} Z_{ijt} + \beta_{4\mathrm{j}} X_i Z_{ijt} + \beta_{5\mathrm{A}} \mu_{\mathrm{i}} + \epsilon_{ijt}, \qquad j \in A(Y_{it}) \tag{1}$$

where μ_{i} is a permanent woman-specific effect, and ϵ_{ijt} is an alternative-specific shock, assumed to be independently and identically distributed. Note that by including interactions among elements of X_i, and Z_{it}, we allow for the possibility that policy and labor and marriage market variables affect different groups differently. If a woman chooses the alternative with the highest value to her, and if ϵ_{ijt} follows the Type I Extreme Value Distribution, then the conditional (on μ) probability that she makes choice j, P_{ijt}, has the multinomial logit form:

$$P_{ijt} = \exp\{V_{ijt}\}/\sum_{k \in A(Y_{it})} \exp\{V_{ikt}\} \qquad (2)$$

where $V_{ijt} = V_{ijt}^* - \epsilon_{ijt}$. The conditional likelihood function contribution for woman i is formed as the product over the months for which she is observed of choice probabilities for her observed choices, conditional on μ. The unconditional likelihood contribution is the integral of the conditional likelihood over the distribution of μ. The latter is treated as a discrete random effect with a two-point distribution. The model is thus a discrete-time multi-state competing risks multinomial logit model of childbearing, union formation and dissolution, and "father identity." The model is estimated by maximum likelihood.

The full set of alternatives, not all of which are available in a given month, is

1. Conceive a child with the current man

2. Conceive a child with a new man

3. End the current union and become single

4. Enter a cohabiting union with the current man

5. Enter a cohabiting union with a new man

6. Marry the current man

7. Marry a new man

We consider only conceptions that lead to a live birth. Conception is treated as a choice but the birth is treated as a censoring event that ends the current pregnancy. Thus the duration of pregnancy and the decision to terminate a pregnancy are not treated as choices. Twin births are treated as an exogenous random event. A *new* man is one who is not the father of any of a woman's children and with whom she has never lived. The *current* man is her partner or spouse if she is currently in a union. If she is not in a union, the current man is the father of her most recent child conceived since the end of her last union, if any, or since she began conceiving children if she has never been in a union. If she is not in a union and has not given birth to any children since the end of the previous union (or ever, if she has never been in a union), then there is no current man, and alternatives 1, 4, and 6 are not available. If she is currently in a union or pregnant, then we assume that only the current man is relevant: she can conceive a child and enter a union only with the current man, so alternatives 2, 5, and 7 are not available.

Distinguishing between a new man and the current man is important because the choice between the two determines which of a woman's children will reside with, or be at risk of residing with, the biological father, and which with a step father. This important distinction has rarely been made in analyses of family formation behavior (see Graefe and Lichter, 1999, for an exception). We impose one key assumption in order to make it feasible to model the choice between a new man and the current man. If a woman ends a union with the current man or if she has a child with a new man, then she is not at risk of conceiving a child or entering a union again with the former current man. With this assumption, there is at most one current man. If she could go back to a previous man after ending a union or giving birth to a child fathered by a different man, there would be too many men to keep track of.

The model is quite rich and flexible. It allows for observed and unobserved heterogeneity, state dependence, duration dependence, and other forms of history dependence. The effects of policy and labor and marriage market conditions are allowed to vary by race and ethnicity.[7] In practice, the specification is restricted in various ways described below, in order to avoid an excessive number of

[7]Our specification, as is common in the literature, does not allow for anticipation effects, which would imply that leads in Z should be included in the model.

parameters. But even after imposing restrictions, the model is quite rich and allows a great deal of flexibility in the effects of interest. After estimating the model, we use it to simulate event histories for a set of artificial women under alternative policy, labor market, and marriage market regimes. The simulations provide a convenient and straightforward way to interpret the estimates. The set of behaviors that we model allows us to trace the dynamics of the five family structures described above for all of a woman's children.

4. Data

A. NLSY79

The NLSY79 began in 1979 with a sample of young men and women who were born between 1957 and 1964. They were interviewed annually from 1979 to1994 and biannually since 1994. We use data on female respondents through the 2004 interview, along with retrospective reports from the first interview about pre-1979 marriage and fertility behavior. The NLSY79 includes a representative cross-section sample and supplementary over-samples of blacks, Hispanics, low-income whites, and members of the armed forces. We exclude the low income white and armed forces supplementary samples, but the black and Hispanic supplementary samples are retained, and controls for race and ethnicity are included in the model. Here we briefly describe construction of the key variables; more details are available in Blau and van der Klaauw (2006).

In 1979, when the sample was between the ages of 14 and 22, 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 since the previous interview. We treat the date of separation as the date of the end of a marriage, since the issue of interest is the presence of a man in the mother's household. However, there are many temporary separations that are followed by reuniting. Modeling the process that determines whether a couple reunites after a separation would make an already rich analysis excessively complicated. Thus, we ignore temporary separations if the duration of the separation was less than or equal to two years. Cases in which a temporary separation lasted more than two years are censored at the date of separation.[8]

The survey has collected information on cohabitation in several different ways, including snapshots of cohabitations in progress at each interview date; the starting date of cohabitations that were in progress at the interview date, beginning with the 1990 interview; the starting date of cohabitations that turned into marriages that were in progress at the interview date, also beginning with the 1990 interview; and both the beginning *and* ending date of cohabitations that did not turn into marriages, as of the 2002 interview. We combined information from the various reports to form as complete a cohabitation history as possible. The cohabitation and marriage histories were combined to form a complete union history. We performed extensive consistency checks on the union history, and examined and corrected many cases by hand (the resulting code is available on request). However, we dropped 401 cases with unresolvable inconsistencies.

The month and year of birth is reported for each child, and beginning in 1984 women were asked the month in which each pregnancy began. We use this information to identify the month of

[8]There is one exception to this rule: if a woman never had any children prior to the end of a temporary separation that exceeded two years, her record is not censored, since there are no children affected by the separation.

conception. If the month of conception is missing, we assume the conception occurred 9 months prior to the birth. Beginning with the 1984 interview, the mother is asked whether the child's biological father is present in the household, for each of her co-resident biological children. Thus, when a woman lives with a man before or during the conception and birth, identifying fathers is straightforward. The more difficult cases are those in which a woman who has given birth to a child since the end of her previous union (or since she began bearing children, if she has never been in a union) conceives and bears another child while single. In such cases, we need to identify whether the father of the new child was the same man who fathered her previous child, but we can do this only if she subsequently enters a union (and is interviewed while the union is still in progress). If she never enters a union following the birth of a child, we cannot determine whether the father of that child was the current man or a new man. 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 modify the likelihood function to account for both of the possibilities, weighted by the probability (from eq. 2) that the father was the current man or a new man. Details are available on request.

At each interview date we can determine from the household roster whether a given child is present in the mother's household. Modeling the processes that determine whether a child lives with the biological mother would be interesting, but is beyond the scope of this paper. These processes are thus treated as exogenous, and the number of children present in the mother's household is adjusted when a child moves in or out. Cases in which a child is away at school or living part-time with the mother are treated as if the child is living with the mother. The death of a child is treated as a censoring event, and children's records are censored at age 18.

After dropping cases with incomplete data or unresolved inconsistencies, we are left with a sample of 4,476 women out of 4,926 eligible for inclusion.[9] Descriptive statistics on the final sample of 4,476 women are displayed in Table 1, separately for whites, blacks, and Hispanics. The variables in the upper panel of the table are background characteristics of the women that are included as explanatory variables in some specifications in the model. These include the mother's parent's education, an indicator for whether the mother lived with both biological parents at age 14, an immigration indicator, and the woman's number of siblings, (as well as indicators for blacks and Hispanics). Black and Hispanic women had on average significantly more disadvantaged backgrounds than white women, based on these measures.

The middle panel of Table 1 summarizes a few of the outcomes of the demographic processes modeled here. The sample women were aged about 40 on average as of the last observation.[10] White women had given birth to an average of 1.71 children, and 21% had not given birth to any children as of the latest observation. Black and Hispanic women had about 0.2 to 0.3 more births on average than whites. Eighty nine percent of white women had ever been married, compared to 62% of black women and 82% of Hispanic women. Whites were also more likely to have ever cohabited.

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[9]The omitted cases include the 401 cases described above with inconsistent marriage and cohabitation histories, and another 32 cases with problematic data on children and fathers. Another 17 cases are lost as a result of missing or inadequate data on contextual variables, described further below.

[10]Women who attrited from the sample are included in the analysis, with attrition treated as an exogenous censoring event. Women who were interviewed in 2004 were between the ages of 39 and 47, but women who attrited before 2004 were younger at the time of their last observation, which explains why the average age at the last observation is only 40.

The lower panel of Table 1 summarizes some of the family structure outcomes experienced by the 8,027 children born to the sample women through 2004. The children were on average aged 13-14 on average at the time of the last observation (after truncating at age 18; without truncating, they were 14-16). Thirty one percent of children of white mothers had ever lived with the mother and no man, compared to 76% of the children of black mothers and 45% of the children of Hispanic mothers. Most children of white and Hispanic mothers lived with both biological parents at some point in their childhood (94% and 85%, respectively), compared to about half of the children of black mothers. Children of black mothers were more likely to live with a stepfather and a cohabiting man compared to children of white and Hispanic mothers, but the differences are smaller in these cases.

B. Contextual Data

The geo-coded version of the NLSY79 provides the state of residence at each survey date, and at the woman's birth and age 14. We collected data from a variety of sources on welfare benefits, welfare reform, child support enforcement, divorce laws, tax rates, and labor and marriage market conditions, and merged them with the NLSY79 by state and year. Here we briefly describe the key measures; the Appendix documents the data sources and describes how state of residence was assigned for non-survey years.

The real AFDC/TANF plus Food Stamp benefit for a family of four (single mother with three children under 18) with no other income is used as a measure of welfare generosity. The Personal Consumption Expenditure Deflator (PCED) was used to convert nominal dollar amounts into year 2000 real equivalents. Figure 1 shows the aggregate time trend of welfare benefits, with state-year observations weighted by their prevalence in our NLSY sample. Welfare benefits declined in real terms over much of the sample period, with a couple of episodes of relative stability. The month and year of implementation of major welfare waivers and the Temporary Assistance for Needy Families (TANF) program for each state are used to characterize welfare reform. The welfare reform variable indicates the presence of any major change in welfare rules authorized by a waiver or TANF.[11] The timing of welfare reform is shown in Figure 2.

The intensity of a state's child support enforcement efforts is measured by three variables commonly used in the literature: the amount of child support payments collected by the state enforcement agency per dollar of administrative expenditure; administrative expenditure per child support case; and the number of paternities established per out of wedlock birth. Aggregate trends in these variables are shown in Figures 3. The federal child support enforcement program began in 1976, and all of these variables are set equal to zero in earlier years. Child support enforcement has increased substantially over time, at somewhat different rates for each measure.

The month and year of passage of unilateral divorce laws were taken from Gruber (2004), which is an update of Friedberg's (1998) data. Figure 4 shows that most of the action in terms of passage of such laws is in the 1970s, but there were occasional cases in the 1980s and 1990s in which states passed a unilateral divorce law.

The TAXSIM program provided by the National Bureau of Economic Research (NBER) was used to compute the average tax rate for alternative filing statuses and numbers of children. The

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[11]TANF was implemented by all states, while not all states requested a welfare waiver. TANF incorporated many of the rule changes implemented by various states as part of their waivers. TANF was implemented by states between 1996 and 1998.

program accounts for all major features of the tax code, including the EITC and (beginning in 1977) state taxes. Rather than conditioning on the woman's observed income, we specify an arbitrary real income level that is used for all women in all years. This ensures that the only variation in the tax rate is due to tax code variation over time and across states. Two different income levels were used (in separate specifications): the real equivalent of the year 2000 poverty line for a family of three, and the real equivalent of 2000 median family income. The average tax rate characterizes the implications of alternative marriage and childbearing choices for take home income better than the marginal tax rate. The average tax rate is treated in our analysis as a choice-specific variable that depends on the marital status and number of children associated with each alternative a woman faces. Current marital status and number of children are lagged outcomes of the choice processes, but their potential endogeneity is accounted for in our estimation approach because of the inclusion of a permanent unobserved heterogeneity term. Thus, in our analysis the tax rate varies over time, across states, and by fertility and marital status. Figures 5 and 6 illustrate trends in tax rates for selected filing statuses and numbers of children. Figure 5 shows rapid growth in the tax subsidy to children for low-income women beginning in the 1980s. Much of this growth is a result of large expansions of the EITC, which provides benefits almost exclusively to low-income families with children (and is refundable; hence the possibility of a negative average tax rate). There is a marriage subsidy for childless low-income women, but not for women with one child. Figure 6 shows that child subsidies in the tax system are much smaller for median-income women, particularly if married, but there has been some growth over time in the tax subsidy to children. However, marriage subsidies are quite large at median income, both with and without children.

The last set of contextual variables characterize labor and marriage markets conditions in each state-year cell. The unemployment rate is used as measure of the state of the business cycle. The female wage rate is measured by the mean real full time average hourly earnings of women aged 16-45, separately for whites, blacks, and Hispanics. The state-year-specific mean wage rate is constructed separately for whites, blacks, and Hispanics using data from the Current Population Survey (CPS) by dividing weekly earnings in the survey week by hours of work per week. The age group 16-45 spans the (employment-eligible) age range of the NLSY sample in the years for which we have data. In order to avoid introducing composition effects into the wage trends, we regression-adjust wages for education and age. The wage measures used here are standardized to a constant level of education (high school graduate) and age (26-30). In order to smooth out spurious fluctuations due to small sample size in some cells, we use a three year moving average of wage rates, within state-sex-race/ethnicity groups. Wages differ significantly by race and ethnicity, and we exploit this in order to provide additional exogenous variation. The male wage rate is constructed in the same way. Note that the wage rate is *not* choice-specific: it is not conditioned on marital status or fertility. It is also not conditioned on the education or other human capital characteristics of the women in our sample. Figures 7a-c show that the male-female wage gap narrowed for all three groups through the mid 1990s, especially for Hispanics, but has been relatively constant more recently. In absolute terms, female wages have been growing relatively rapidly recently, and male wages began growing as well in recent years after a long period of stagnation. However, only for females are mean real wages higher today than in the 1970s.

The sex ratio (male/female) is constructed from the CPS, using year-state-age-race/ethnicity-specific counts. In this case, we limit the female CPS sample to the age range of the NLSY sample in a given year, and the male sample to an age range two years older than the female age range, to account for the average age gap between spouses. In order to avoid using excessively noisy data, we used a three year moving average, and then applied a minimum cell size criterion of 30 for both wage rates and the sex ratio. We lose 5.4% of the potential NLSY person-month observations as a result of this criterion, with the loss disproportionately larger for blacks and Hispanics. Figure 8 illustrates the trend in the sex ratio by race and ethnicity. There is no discernable trend for any of the groups, but there is a sharp drop in the late 1970s for all three groups. This may be due to the change from the May CPS as a data source in 1970-1978 to the Merged Outgoing Rotation Group files as a source for 1979-2004. There is no other obvious explanation for the sharp changes, and the reason for the apparent difference in measurement across the two sources is unclear (see the Appendix for documentation). The size of the drop is relatively uniform across states (not shown), so inclusion of period dummies in the model will ensure that this measurement issue does not have an undue influence on the results.

The figures shown here emphasize the time series variation in the contextual variables, but there is also a substantial amount of cross-sectional variation in many of them as well. To illustrate, Table 2 shows the R^2 from two regressions of each of the contextual variables: one on year and race/ethnicity fixed effects, and the second on year, race/ethnicity, and state fixed effects. Variation across states accounts for a major part of the overall variation in welfare benefits, child support enforcement, unilateral divorce, the unemployment rate, and wage rates. State-level variation is less important for welfare reform, the sex ratio, and tax rates. It is also worth noting that even after accounting for period, race/ethnicity, and state or region fixed effects, all of which are included in our analysis, there is a reasonable amount of remaining variation in most of the contextual variables, ranging from 5-15% of the total variation for welfare benefits, unilateral divorce, welfare reform, and several of the tax rates, to over 50% for the sex ratio.

5. Results

We estimated several different specifications of the model in order to examine sensitivity of the results to alternative controls for time trends and cross state differences. The model is nonlinear and has a large number of parameters, so it is not yet been possible to incorporate full sets of state fixed effects and calendar year fixed effects. The estimates reported here are for the richest specification we have been able to estimate. It includes a quadratic time trend as well as dummies for five year periods[12], census region dummy variables (nine regions), and dummies for the six largest states. The only individual-level exogenous variables included are dummies for the woman's race (black) and ethnicity (Hispanic), a linear term for her date of birth and a quadratic in her age. We estimated other specifications that included background characteristics of women, such as the family structure in which she resided at age 14, as well as her education and Armed Forces Qualification Test score. The estimated effects of the contextual variables were not sensitive to inclusion of these covariates.

The specification includes several outcomes of previous fertility, marriage, cohabitation, and choice-of-man decisions, including the number of children born to date, the number of children fathered by the current man, if any, the number of marriages and cohabitations, whether a single woman was in a

[12]For some choices, the effects of five year period dummies could not be estimated, so ten year periods are used instead for those cases.

cohabitation or a marriage in her previous spell, quadratics in the ages of her youngest and oldest children, and quadratics in the duration of her current marriage, cohabitation spell, union spell, single spell, and pregnancy. In the interests of empirical tractability we imposed a substantial number of exclusion restrictions in cases in which a given parameter consistently had small and statistically insignificant effects. However, the contextual variables are not excluded from any of the parameter vectors. We found that the effects of several of the contextual variables differed substantially by race and ethnicity. Thus, interactions between dummies for black and Hispanic and several of the contextual variables are included.

The parameter estimates and standard errors on the contextual variables are shown in Table 3.[13] The magnitudes of the parameters and in some cases even the signs are not directly informative about the size and signs of the effects of interest, so we do not discuss them in detail. It is worth noting that there are a total of 207 parameters on the contextual variables, of which 26 are estimated to be significantly different from zero at the 10% level. The majority of the significant parameters involve interactions with race or ethnicity.

We use the parameter estimates to simulate the life histories of 10,000 artificial women of each race/ethnicity who are subject to the risks characterized by the model. Each woman starts out single and with no children at age 12 . The estimated parameters are used to compute the probability of each of the three events that can occur to a single woman aged 12 with no children (enter a cohabitation, enter a marriage, conceive a child). A uniform random number generator determines which, if any,

[13]See Blau and van der Klaauw (2006) for discussion of results for the individual-level variables from a similar specification without the contextual variables.

event occurs, depending on the value of the random draw compared to the event probabilities. If the event is conceiving a child, a pregnancy duration is randomly assigned according to the observed distribution of pregnancy durations in the sample. The Y_{it} 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 45 of the woman.[14]

In the baseline simulation the contextual variables are set equal to their sample means, shown in Table 2.[15] A one standard deviation increase is simulated separately for each continuous variable (see Table 2 for the standard deviations), and the effect of changing the value from zero to one is simulated for the binary variables (welfare reform and unilateral divorce).[16] Table 4 shows simulation results for selected outcomes of women, separately for whites, blacks, and Hispanics. These outcomes are closely related to the family structure experiences of children, which are discussed below. And these are the outcomes examined most frequently in previous studies, so we can compare our findings for these outcomes to the literature. The first row of each panel shows the mean outcomes in the baseline simulation. The remaining rows show the mean changes in the outcomes resulting from each simulation.

A one standard deviation increase in the monthly welfare benefit is estimated to cause age at first birth to increase by .108 for whites, and to decrease by .040 and .186 for blacks and Hispanics,

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[14]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. In the simulations, there are no deaths, no twin births, and no children who move in or out of the mother's household.

[15]The simulated women are assumed to have been born in January 1961 and to live in California. These assumptions determine the values of the period, region, and state effects, but do not affect the values of the contextual variables used in the simulations.

[16]The procedure is a bit different from the tax rate simulations: see the note to Table 4.

respectively. The effect on the number of children ever born is -.040 for whites, -.081 for blacks, and .066 for Hispanics. All of these effects are quite small, consistent with findings from previous research that welfare benefits have little impact on fertility (see the discussion in Section 2). Welfare benefits are also estimated to have small effects on the union formation and dissolution outcomes shown in Table 4. The largest effect is an increase of .049 in the probability of ever cohabiting for Hispanics, but the parameter estimates underlying this effect are not significantly different from zero.

Welfare reform, on the other hand, is estimated to have large effects on fertility, reducing age at first birth by more than one year for all three groups, and increasing the number of children ever born by .415 for whites and .058 for Hispanics, but decreasing children ever born by .147 for blacks. The effects of welfare reform on union formation and dissolution are also relatively large for all three groups: increases of .070 to .099 in the probability of every marrying (as of age 45). The probability of ever cohabiting declines for whites by .119 and for blacks by .031, and rises by .040 for Hispanics. The probability of ever divorcing (conditional on ever marrying) falls a bit for whites, and increases by .076 for blacks and .096 for Hispanics. The probability of ever remarrying, conditional on divorce, rises substantially for all three groups. The increase in marriage and decline in divorce are consistent with the findings of several previous studies of welfare reform, though as noted in Section 2 there is some dispute about such findings. We are not aware of any previous studies of the effect of welfare reform on fertility. The large effects of welfare reform are robust across all of the specifications estimated to date. However, richer controls for state and time effects are needed (dummy variables for every state and year) before we can be confident that these large effects are not caused by unobserved heterogeneity across states or by other unmeasured time trends. And as noted above, the women in the NLSY79 cohorts were in their 20s and 30s when welfare reform began in the early 1990s, so we do not claim that the effects estimated here can be applied to other cohorts.

The consequences of welfare reform for the family structure experiences of children cannot easily be directly inferred from these results. We explore these consequences in depth below, but a summary measure is included in the last column of Table 4: whether a woman ever had a child with no man present in the household. Welfare reform is estimated to have decreased the probability of this outcome for white women by .031, but increased it by .064 for black women and by .018 for Hispanic women.

A one standard deviation increase in the male wage rate is estimated to have some fairly large effects on white women, but none of the underlying coefficient estimates are significantly different from zero. Several of the black and Hispanic interactions are statistically significant, however, and the simulated effects are in some cases quite large. A higher male wage rate increases the number of children ever born to black women by .169, and reduces the age at first birth by .131. Cohabitation rises, ever-married remains unchanged, and the incidence of divorce falls. On balance, there is little change in the probability that a black woman ever has children with no man present. For Hispanic women, a higher male wage rate increases fertility, marriage, and remarriage, and reduces divorce. The net effect is a substantial decline of .139 in the probability that a Hispanic woman ever has children with no man present. Previous findings reported in the literature indicate a negative effect of the male wage rate on female headship, consistent with our results for Hispanics.

A higher female wage rate generally has effects that are of the opposite sign from those of the male wage rate. As with the male wage rate, the effects are not significantly different from zero for whites, but for blacks and Hispanics a higher female wage rate has negative effects on fertility that are significantly different from zero. Children ever born decline by about 0.1 for blacks and Hispanics. Marriage and remarriage decline as well, and divorce increases. All of these findings are consistent with results in previous studies. The net effect is a .044 increase in the probability that a black woman ever has children with no man present, and a .151 increase for Hispanic women.

Unilateral divorce laws are estimated to increase the incidence of divorce for whites and blacks, but to reduce it for Hispanics. All three of the underlying coefficient estimates are significantly different from zero. The increase in divorce is consistent with the findings of several previous studies, but as noted above, there is some disagreement about the size of the effect. Unilateral divorce increases fertility for whites and Hispanics, but reduces fertility by a small amount for blacks. The increase in fertility is consistent with some of Gruber's (2004) findings. Unilateral divorce increases the probability that a woman ever has children with no man present by .079 for whites and .018 for blacks, and decreases it by .114 for Hispanics.

The simulated effects of average tax rates are expressed in the form of a comparison between a hypothetical scenario with no tax gain from marriage (implemented by setting tax rates for single women equal to those for married women) and the mean observed tax gain from marriage, and similarly for the tax gains from additional children. The counterfactual in the latter simulation is all tax rates set to the zero-children level, for married and single women, respectively. The tax gain from marriage is estimated to have very small effects on the incidence of marriage for all three groups, consistent with findings of previous studies that focused specifically on the EITC. The tax gains from children are estimated to increase children ever born by a small amount, but to reduce the age at first birth by a fairly large amount in some cases. A higher tax gain to children conditional on marriage causes a decline in the probability that a white woman ever has children with no man present by .071, and a higher tax gain to children conditional on being single causes a decline in the probability by .176. These are rather large effects. The effects for blacks and Hispanics are much smaller.

The effects of child support enforcement (CSE) are generally quite small across all outcomes and measure of CSE. There are also small effects of the sex ratio and the unemployment rate. However, we have not yet explored whether the effects of these variables differ significantly by race or ethnicity.

Table 5 summarizes simulation results for several family structure outcomes of children. Welfare reform is estimated to have had a large impact on the family structure experiences of children of black mothers, decreasing the incidence of ever living with the biological father by .203, increasing the incidence of ever living with no father by .134, and increasing the incidence of ever living with a step father by .165. The probability that a child's mother was single at birth increases by about 0.2. The effects are much smaller and often of a different sign for whites and Hispanics. Unilateral divorce is estimated to have increased the incidence of children of Hispanic mothers living with the biological father by .055, and reduced the incidence of no father and step father by .118 and .087, respectively. A higher male wage rate has similar impacts for children of Hispanic mothers, with much smaller effects for children of white and black mothers. A higher female wage rate has the opposite effects, and the effects are again relatively large for Hispanics, and in this case also for blacks, but much smaller for whites. Finally, the tax gains from having a child while single *decreases* the incidence of children of white mothers ever living with no father. This somewhat surprising result may be a consequence of the fact that a higher tax gain to children while single substantially increases the incidence of cohabitation for children of white mothers.

A richer view of the effects of the contextual variables can be gained by examining family transition patterns generated by the simulations. Figure 9 shows how a one standard deviation increase in the female wage rate affects age-specific annual transition rates in living arrangements of children, along two dimensions: father status (biological, step) and union status (married, cohabiting, single).[17] For children of white mothers, all of the effects are small, as noted previously, and there is very little impact of the female wage rate on the cumulative amount of time spent with the biological father and a step father (shown in the last two panels of Figure 9a). The effects for children of black mothers are larger: a higher female wage rate reduces the risk of the biological father entering the household and increases the risk of his exit. The result is a decrease of over one year in cumulative time spent by a child up to age 18 with the biological father present, from an already low base of about four years. The main union transition rate changes that produce this decrease are an increase in the divorce rate and a higher rate of dissolution of cohabitations. Entry and exit rates of stepfathers both increase with a higher female wage rate, with little net effect on time spent with a stepfather. The effects of a higher female wage rate for Hispanics are quite similar to those for blacks, with the exception that cumulative time spent with a stepfather increases by about half a year.

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[17]Changes in father status and union status are of course not mutually exclusive events. For example, a divorce and exit of the biological father would both result from a divorce in which the husband was the biological father of at least one child in the mother's household. We present results separately for union and fatherhood transitions in order to avoid the excessively large number of graphs that would result from a full cross classification.

We analyzed the effects of several of the other contextual variables on transition patterns, but in order to avoid lengthening an already long paper we summarize the results here without presenting the underlying figures (the figures are available on request). The main channel through which a higher male wage rate affects family structure is a lower divorce rate (for blacks and Hispanics; there is a slight increase in divorce for whites). The net effects on cumulative time spent living with the biological father and step father are very small for children of white and black mothers. The effects are sizeable for children of Hispanic mothers: a one standard deviation increase in the male wage rate increases time spent living with the biological father by about two years, and reduces time spent living with step father by about one year. There is a small increase in the rate of entry of the biological father, but most of the effect occurs through a lower exit rate.

Welfare reform is estimated to have has small effects on family structure for children of white and Hispanic mothers, but a large impact for children of black mothers. Welfare reform reduced the rate of entry and substantially increased the exit rate of the biological father, and also significantly increase the risk of entry of a stepfather. These changes resulted from an increase in the marriage rate, but only for stepfathers, and an increase in marital dissolution. The net effect was a decrease of about 2.5 years in time spent with the biological father and an increase of about two years in time spent with a step father. These are large increases, particularly in percentage terms.

Unilateral divorce laws are estimated to have had small effects on family structure for children of white and black mothers, but a large impact for children of Hispanic mothers. For Hispanics, unilateral divorce laws increase the transition rate from single to cohabitation and from cohabitation to marriage, but not the transition rate from single to married. This is one of the few instances in which cohabitation plays an important role in family structure transitions. Unilateral divorce also substantially decreases the dissolution of both cohabiting and marital unions. The entry rate of the biological father increase, the exit rate decreases, and both the entry and exit rates of step father decline. The result is an increase of about two years in time spent with the biological father, and a decline of about one year in time spent with a step father.

An increase in the tax benefit of an additional child conditional on being married causes an increase in time spent with the biological father of about one year for children of white and black mothers, and about half a year fro children of Hispanic mothers. Time spent with a step father declines slightly for all three groups. The entry and exit rates of step fathers declines for all three groups, but the other channels differ. For whites, a decline in the exit rate of the biological father and a decline in divorce are significant contributors. For blacks, a large increase in the transition rate from cohabitation to marriage is the main channel, and for Hispanics this channel as well as an increase in the transition rate from single to married are bother important.

An increase in the tax benefit of an additional child conditional on being single causes a decline in time spent with the biological father by about two years for whites, .75 years for blacks, and one year for Hispanics. Time spent with a step father increases by about one year for whites but changes very little for blacks and Hispanics. There are some common causes across the three groups: increases in both the entry and exit rates of step fathers. But there are also some factors specific to each group: a large increase in the divorce rate and exit rate of the biological father for whites; and a large decrease in the transition rate from cohabitation to marriage for blacks and Hispanics.

6. Conclusions

The evidence presented here suggests that the family structure experiences of the children of women born from 1957 to 1964 have been influenced by the tax gain associated with childbearing, welfare reform, unilateral divorce laws, and the wage rates available to men and women in the labor market. The results suggest that other contextual variables such as child support enforcement efforts, the level of welfare benefits, the tax gain to marriage, the sex ratio, and the unemployment rate had little impact on the family structure experiences of these children. The methods we use to produce this evidence are rather new, and we view as encouraging the consistency between our findings and those of most previous studies, for the outcomes that can be compared. But we readily acknowledge several important limitations that make any strong conclusions or generalizations based on the results premature at best.

First, we have not yet been able to control for a full set of state fixed effects and a full set of calendar year fixed effects. Such controls are not a panacea for all forms of unobserved differences across states and over time that could be confounded with variation in the contextual variables. But these controls are common in the literature, and are especially important for inferences about the effects of welfare reform and unilateral divorce, for which implementation took place during a relatively short period of time. We plan to extend our controls for unobserved state and time fixed effects in the next draft of the paper.

Second, our results apply only to a narrow range of birth cohorts, and it is difficult to see how to generalize them. Ideally, we would do a cross cohort counterfactual analysis of how much of the difference in family structure outcomes across birth cohorts of women the model can account for by differences in the values of the contextual variables to which the cohorts were exposed. However, there is not enough variation in outcomes across the eight birth cohorts within the NLSY79 for such an analysis, and there are no other cohorts with sufficient information on the outcomes of interest. Thus, the best we might be able to do along these lines is to compare the predicted outcomes of the NLSY79 cohort under two alternative counterfactual scenarios: one in which the cohort was exposed to the early 1970s values of the contextual variables for their entire childbearing years, and another in which the cohort was exposed to the early 2000s values of the contextual variables for their entire childbearing years. We intend to explore such scenarios in the next draft.

Finally, while our model is rich, it omits some potentially important channels through which the contextual variables could affect family structure outcomes. We do not model time spent by children living without the biological mother, nor do we model the presence of other members other than fathers and siblings. Temporary separations are ignored as well. All of these channels are worth exploring in future work.

The motivation for our analysis was the challenge offered by Ellwood and Jencks (2004) to develop new approaches to analyzing the determinants of family structure change. We believe that our analysis has been successful in responding to several of their suggestions for new directions in this field. We treat marriage and childbearing as distinct decisions that interact in potentially complex ways to produce single parenthood; we distinguish between delay and permanent avoidance of marriage and childbearing; we include cohabitation in the analysis; we avoid restrictive functional forms in dynamic models; and we explain differential family structure change by race, though not yet by family background or education. But we have not yet been successful in using our modeling approach to explain changes across cohorts in family structure. This remains an important challenge for us and other researchers.

Appendix

Wage rates. The mean full time hourly wage rate was computed for men and women aged 16-45 by year, state, and race/ethnicity from the Merged Outgoing Rotation Group (MORG) files of the Current Population Survey (CPS) for 1979-2004, and from the May CPS files for 1970-1978. The wage rate is computed by dividing weekly earnings by hours of work per week. Cases were included in the computation only if weekly earnings were at least $150 (in year 2000 dollars), hours of work were at least 30, and the resulting hourly wage rate was between $2.00 and $200.00. Weekly wages were topcoded at $999 from 1970-1988, $1,923 from 1989-1997, and $2,884 from 1998 on. Wages were deflated using the Personal Consumption Expenditure Deflator (PCED, base year 2000). The sampling weight provided in the CPS files was used in the calculations. Before 1977, some states are not separately identified, so for those years the mean wage for the group of states (by sex, year, and race/ethnicity) is assigned to each state in the group.

Weekly earnings are given in categorical form before 1973 in the May CPS files. The midpoint of each category is used in this case, with $600 assigned for the highest category in 1972 (when the lower limit is $500), and $300 assigned in 1970-1971 (when the lower limit of the highest category is $200). In the 1973-1978 May CPS files, the continuous weekly earnings variable is missing for some cases, but the categorical version of earnings is also on the file for those years. If the categorical variable is not missing, it is used to compute the wage when the continuous measure is missing (the categories are the same in 1973-1978 as in 1972). Hispanic ethnicity is not identified in the May CPS in 1970-1972. The real 1973 means by state were used for 1970-72 for Hispanics.

The wage rate is regressed on education dummies (four groups), age dummies (six groups), and state of residence, separately by year, sex, and race/ethnicity. A wage rate is predicted for each employed individual, holding education constant at high school graduate and age constant at 26-30. Wages are then averaged within state-year-sex-race/ethnicity cells. In order to smooth out spurious fluctuations due to small sample size in some cells, we use a three year moving average of wage rates, within state-sex-race/ethnicity groups. Cells with fewer than 30 cases (after averaging) are omitted.

Sex Ratio. The sex ratio (males/females) is computed by state, year, and race/ethnicity using the CPS MORG files for 1979-2004 and the March CPS supplements for 1970-1978. The ratio is computed using females who are in the same age range at the interview date as the NLSY cohort in a given year, and males who are two years older than the females. Sample weights are used. Hispanic ethnicity is not identified in the March CPS in 1970, so Hispanics are assigned the 1971 sex ratio for 1970. The MORG files contain individuals aged 16 and above, so in 1979 the sex ratio excludes females aged 14 and 15 and their male counterparts (aged 16-17), and in 1980 the sex ratio excludes females aged 15 and males aged 17. We use a three year moving average of the sex ratio, within state-sex-race/ethnicity groups. Cells with fewer than 30 cases (after averaging) are omitted.

Child Support Enforcement. Data on administrative expenditure, caseloads, paternity establishments, and child support collections by state agencies were collected from annual reports to Congress by the Office of Child Support Enforcement, and the Green Book. Data on out of wedlock births for some years were collected from the National Center for Health Statistics Vital Statistics data. Data on out of wedlock births for 2004 are not available. Out of wedlock births for 2004 are estimated by applying the state-specific fraction of births that were out of wedlock in 2003 to the state total number of births in 2004. In some of the early years of the Child Support Enforcement program (1976-

1980), the paternity establishment rate is missing for some states. The missing years were assigned either the overall state mean, or the predicted value from a state-specific regression with a time trend. Caseload data are not available for 1976 and 1977, the first two years of the program. The 1978 values are used for 1976 and 1977. All child support variables are set equal to zero for 1970-1975.

Welfare Benefit. Data for the AFDC/TANF cash benefit for a family of four with no income are from Robert Moffitt's welfare benefits file for the years 1970-1998. Data for 1999-2004 are from the 2004 Green Book, the Congressional Research Service (2005), and the Urban Institute's Assessing the New Federalism web site. In some recent years data are only available for a family of three. The benefit for a family of four was estimated by applying the state-specific ratio of benefits for households of size and three and four, which are both available for 1996-1998 and 2003-2004. The Food Stamp guarantee for a family of four is from Moffitt's data base for 1970-1998, updated with data from the web site of the Food and Nutrition Service.

Welfare Reform. The timing of implementation of welfare waivers and TANF are from the web site of the Office of the Assistant Secretary for Planning and Evaluation, Department of Health and Human Services.

Divorce Laws. The year of passage of a unilateral divorce law is from Gruber (2004), Table 1.

Tax Rates. Tax rates are computed using the NBER's TAXSIM program. Tax rates were computed for two income levels: the poverty line for a family of three (one adult and two related children) in 2000 ($13,874), and for median family income in 2000 ($50,372), both adjusted for inflation in other years. All income was assumed to be from earnings. Child care expenditure for a poor family was assumed to be 23% of income, and for a median-income family 6% of income (Johnson, 2005). All children were assumed to be under 17 for purposes of the child tax credit. Taxes were computed for alternative numbers of children (0-9) and filing statuses (single, head of household, and married filing jointly). State taxes are included from 1977-2004, but are not included for 1970-1976. In married families, 60% of earnings were allocated to the husband and 40% to the wife.

Other Aggregate Variables. The state-year-specific unemployment rate is taken from the Moffitt data base, updated after 1998 with data from the BLS web site.

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Table 1: Descriptive Statistics on Individual Variables

	White	Black	Hispanic
Background Characteristics of women			
Father's years of education	12.2	10.1	8.2
Mother' years of education	11.9	10.7	7.9
Woman lived with both biological parents at 14	0.79	0.50	0.68
Number of siblings	3.1	4.7	4.5
Immigrant	0.031	0.025	0.261
Other characteristics of women			
Completed education			
Armed Forces Qualification Test score (percentile)			
Demographic outcomes of women			
Number of children	1.71	1.89	1.99
No children	0.21	0.18	0.17
Ever married	0.89	0.62	0.82
Ever cohabited	0.43	0.36	0.39
Age at last observation	40.5	39.6	40.1
Number of women	2,292	1,338	846
Family structure outcomes of children			
Ever lived with no father	31.0	76.1	45.0
Ever lived with stepfather	17.8	28.1	23.6
Ever lived with biological father	94.0	52.0	84.9
Ever lived with cohabiting father	13.8	26.8	23.1
Age of child at last observation	12.9	14.0	13.3
Number of children	3,864	2,496	1,667

Notes: Father's and mother's education are missing for some cases. The statistics reported are for non-missing cases only. The child outcomes are censored at age 18.

Table 2: Summary Statistics on Contextual Variables

	Mean	SD	R^2 in regression on year and race/ethnicity fixed effects	R^2 in regression on year, race/ethnicity, and state fixed effects
Welfare benefit	1013	265	0.19	0.93
Paternity establishment rate per out of wedlock birth	.348	.279	0.40	0.55
Child support collections per dollar of administrative expenditure	3.02	2.15	0.33	0.76
Administrative expenditure per child support enforcement case	.130	.103	0.69	0.81
Unilateral divorce law in effect	.552		0.06	0.97
Welfare reform in effect	.210		0.88	0.89
Unemployment rate	6.51	2.16	0.52	0.71
Sex ratio (male to female)	.952	.112	0.43	0.50
Male wage rate	12.12	1.80	0.46	0.86
Female wage rate	9.70	1.30	0.16	0.80
Average tax rate (ATR), married no children	.170	.018	0.49	0.56
ATR, married one child	.061	.062	0.91	0.93
ATR, single no children	.228	.023	0.72	0.89
ATR, single one child	.077	.065	0.85	0.88

Note: Unit of observation is a state-year-race/ethnicity cell. Observations are weighted by the cell sample size in the NLSY sample. Average tax rates facing married women with 2-9 children are .030, .028, .027, .027, .026, .026, .025, and .025, respectively. Average tax rates facing single women with 2-9 children are .035, .031, .030, .028, .028, .027, .027, and .026, respectively.

Table 3: Selected Coefficient and Standard Error Estimates

	Conceive		Become	Enter Cohabitation		Marry	
	Current Man	New Man	Single	Current Man	New Man	Current Man	New Man
	1	2	3	4	5	6	7
Welfare Ben.	-0.012	0.010	-0.032	0.135	0.003	0.005	**-0.055**
	(0.019)	(0.034)	(0.033)	(0.118)	(0.034)	(0.032)	(0.029)
*Black	-0.034	-0.006	0.025	0.143	-0.061	0.042	0.073
	(0.030)	(0.043)	(0.047)	(0.136)	(0.059)	(0.050)	(0.050)
*Hispanic	0.026	0.021	0.046	0.004	0.049	-0.071	0.052
	(0.025)	(0.050)	(0.051)	(0.152)	(0.056)	(0.050)	(0.041)
Unil. Divorce	0.110	-0.058	**0.390**	-0.337	0.218	0.039	-0.073
	(0.082)	(0.151)	(0.138)	(0.442)	(0.148)	(0.150)	(0.127)
*Black	**-0.173**	0.024	**-0.237**	0.404	-0.185	-0.022	**0.273**
	(0.089)	(0.141)	(0.135)	(0.428)	(0.156)	(0.143)	(0.149)
*Hispanic	0.007	0.155	**-0.725**	0.009	0.065	**0.535**	-0.064
	(0.119)	(0.247)	(0.218)	(0.603)	(0.249)	(0.266)	(0.221)
Welfare Ref.	0.191	0.452	-0.062	**1.888**	-0.365	0.365	**0.942**
	(0.207)	(0.451)	(0.334)	(0.822)	(0.485)	(0.377)	(0.418)
*Black	**-1.351**	0.000	0.125	-2.597	0.610	-0.202	-0.273
	(0.797)	(0.000)	(0.516)	(3.869)	(0.902)	(0.860)	(1.159)
*Hispanic	-0.375	0.000	0.175	-1.062	**1.206**	0.055	0.394
	(0.314)	(0.000)	(0.464)	(1.332)	(0.601)	(0.644)	(0.660)
Male wage	-0.093	-0.001	0.035	0.350	0.043	-0.030	0.029
	(0.054)	(0.098)	(0.096)	(0.270)	(0.096)	(0.093)	(0.080)
*Black	**0.179**	0.033	-0.092	-0.071	0.038	-0.160	0.028
	(0.064)	(0.099)	(0.108)	(0.278)	(0.114)	(0.112)	(0.112)
*Hispanic	**0.164**	-0.158	**-0.281**	-0.084	**0.257**	0.077	0.066
	(0.059)	(0.125)	(0.129)	(0.332)	(0.130)	(0.121)	(0.094)
Female wage	0.115	-0.046	0.119	-0.573	-0.027	-0.096	0.009
	(0.070)	(0.151)	(0.127)	(0.379)	(0.130)	(0.120)	(0.107)
*Black	-0.204	-0.001	0.132	0.115	-0.104	0.144	-0.197
	(0.091)	(0.157)	(0.151)	(0.424)	(0.169)	(0.155)	(0.157)
*Hispanic	**-0.232**	0.229	0.257	0.124	-0.355	0.269	-0.179
	(0.108)	(0.205)	(0.230)	(0.577)	(0.243)	(0.219)	(0.169)
Tax rate	**-2.232**	0.873	**3.763**	0.000	0.000	-1.322	-1.381
	(1.078)	(1.775)	(1.439)	(0.000)	(0.000)	(1.631)	(1.590)
*Black, Hisp.	**2.589**	-1.692	**-3.605**	0.000	0.000	**4.655**	2.296
	(1.341)	(1.850)	(1.946)	(0.000)	(0.000)	(1.859)	(2.217)
Sex ratio	0.218	-0.229	-0.174	1.289	0.437	0.441	0.222
	(0.215)	(0.309)	(0.355)	(0.935)	(0.413)	(0.394)	(0.363)
Unem. rate	-0.019	-0.006	-0.012	-0.029	0.013	-0.030	-0.004
	(0.010)	(0.016)	(0.018)	(0.047)	(0.018)	(0.018)	(0.014)
Paternity	-0.134	0.125	0.175	-0.086	0.044	0.010	-0.052
	(0.124)	(0.186)	(0.217)	(0.682)	(0.217)	(0.217)	(0.194)
Coll/Admin	0.008	0.004	-0.028	0.001	0.016	**-0.047**	-0.006
	(0.014)	(0.020)	(0.025)	(0.070)	(0.023)	(0.024)	(0.022)
Admin/Case	-0.385	0.817	-0.972	-0.174	**1.326**	0.685	0.367
	(0.447)	(0.763)	(0.822)	(2.800)	(0.722)	(0.697)	(0.569)

Notes: Welfare benefit is in units of 100 dollars per month. Paternity is the paternity establishment rate per out of wedlock birth. Coll/Admin is the ratio of child support dollars collected to administrative expenditure. Admin/case is administrative expenditure per child support case. There is a single interaction term for the tax rate, a dummy equal to one if the woman is Black or Hispanic. Coefficients estimates in bold are significantly different from zero at the 10% level.

Table 4: Simulation Results for Outcomes of Women

A. Whites

	Age at first birth	Ever have a child with no man present	No. of children born	Ever cohab it	Ever marry	Ever divorce (if ever married)	Ever remarry (if ever divorced)
Baseline value:	24.4	.327	2.07	.362	.926	.329	.523
Change due to:							
Welfare benefit	.108	-.013	-.040	.017	-.007	-.020	-.028
Welfare reform	-1.806	-.031	.415	-.119	.070	-.014	.205
Unilateral divorce	.035	.079	.089	.087	.001	.090	.046
Unemployment rate	.143	-.001	-.040	.009	.003	-.004	-.007
Sex ratio	-.020	-.010	.052	.002	.009	-.007	.014
Male wage rate	.270	.019	-.155	.027	.007	.023	-.001
Female wage rate	-.088	.036	.140	-.000	-.003	.026	.004
Child support enforcement							
Paternity establish. rate	.027	.018	-.026	.013	.001	.015	.003
Collection rate/admin exp.	.012	-.008	.025	.002	.001	-.018	-.014
Admin. expenditure/case	-.089	-.023	-.005	.006	.013	-.023	.019
Average tax rate							
Gain from marriage	-.001	.002	.005	.001	.006	.020	.024
Gain from child (married)	-.360	.071	.085	.027	.000	.074	.046
Gain from child (single)	.168	-.176	.008	-.075	-.006	-.176	-.049

B. Blacks

	Age at first birth	Ever have a child with no man present	No. of children born	Ever cohab it	Ever marry	Ever divorce (if ever married)	Ever remarry (if ever divorced)
Baseline value:	21.3	.854	2.28	.545	.839	.426	.439
Change due to:							
Welfare benefit	-.040	-.016	-.081	-.010	.005	.011	.003
Welfare reform	-1.057	.064	-.147	-.031	.088	.076	.118
Unilateral divorce	-.004	.018	-.065	.004	.007	.065	.033
Unemployment rate	-.008	.003	-.008	.005	-.007	-.014	-.017
Sex ratio	.042	-.009	-.003	.002	.008	-.002	.009
Male wage rate	-.131	-.009	.169	.037	.001	-.054	-.019

Female wage rate	.019	.044	-.090	.013	-.033	.071	-.013
Child support enforcement							
Paternity establish. rate	-.035	.003	.015	.006	.001	.004	-.001
Collection rate/admin exp.	.039	-.002	.039	.006	-.009	-.024	-.017
Admin. expenditure/case	-.109	-.014	.055	.013	.017	-.021	.014
Average tax rate							
Gain from marriage	-.063	.008	.004	.008	-.003	-.005	-.005
Gain from child (married)	-.011	.023	.008	.032	-.020	-.019	-.040
Gain from child (single)	-.359	-.003	.013	-.022	.033	.030	.074

C. Hispanics

	Age at first birth	Ever have a child with no man present	No. of children born	Ever cohab it	Ever marry	Ever divorce (if ever married)	Ever remarry (if ever divorced)
Baseline value:	23.4	.489	2.18	.392	.928	.388	.554
Change due to:							
Welfare benefit	-.186	.028	.066	.049	-.009	.007	-.011
Welfare reform	-1.379	.018	.058	.040	.099	.096	.313
Unilateral divorce	-.205	-.114	.192	-.028	.023	-.104	.054
Unemployment rate	.072	.007	-.044	.004	-.003	.003	-.013
Sex ratio	.027	-.007	.023	.005	.004	.004	.007
Male wage rate	.272	-.139	.123	-.007	.025	-.107	.043
Female wage rate	-.437	.151	-.110	.010	-.019	.162	-.006
Child support enforcement							
Paternity establish. rate	.016	.022	-.036	.009	.000	.020	-.002
Collection rate/admin exp.	-.009	.002	.022	.002	-.005	-.009	-.018
Admin. expenditure/case	-.170	-.007	.004	.001	.007	-.013	.017
Average tax rate							
Gain from marriage	-.040	.001	-.005	.010	-.004	-.011	-.009
Gain from child (married)	.033	.019	-.023	.024	-.008	-.013	-.030
Gain from child (single)	-.226	-.031	.037	-.021	.018	-.002	.061

Notes: The baseline values of the explanatory variables are the sample means shown in Table 3. Each row shows the simulated change in the outcomes of interest caused by a one standard deviation increase in the explanatory variable (see Table 3 for the standard deviations), or the effect of changing the variable from 0 to 1 for welfare reform and unilateral divorce. The simulation for the average tax rate gain from marriage replaces the tax rates on single women with a given number of children with the corresponding tax rates for married women, thus eliminating the tax gain from marriage. The results shown in the table are the effect of moving from no tax gain from marriage to the baseline (average) observed tax gain. The simulations for the tax gain from having a child set all tax rates for a given marital status equal. The results shown in the table are the effect of moving from no tax gain from children to the baseline (average) observed tax gain. See Table 3 for the baseline tax rates.

Table 5: Simulation Results for Outcomes of Children

A Whites

	Ever live with biological father	Ever live with no father	Ever live with step-father	Ever live with cohabiting parents	Ever live with married parents	Mother single at conception	Mother single at birth
Baseline outcome value:	.964	.260	.178	.152	.992	.139	.044
Change due to:							
Welfare benefit	.006	-.014	-.014	-.004	.000	.007	-.003
Welfare reform	.023	-.030	-.004	-.026	.008	-.019	-.027
Unilateral divorce	-.003	.065	.050	.050	-.001	-.007	-.001
Unemployment rate	.000	-.001	-.002	.002	-.001	-.002	.001
Sex ratio	.008	-.009	-.007	-.004	.002	-.013	-.008
Male wage rate	.001	.021	.009	.013	-.002	.007	-.001
Female wage rate	-.003	.028	.024	.020	-.001	-.013	.004
Child support enforcement							
Paternity establish. rate	-.002	.016	.011	.010	.000	.007	.002
Collection rate/admin exp.	-.002	-.009	-.007	-.004	-.004	-.001	.004
Admin. expenditure/case	.003	-.019	-.012	-.011	.001	.004	-.003
Average tax rate							
Gain from marriage	.002	.001	.001	.002	.001	-.007	-.004
Gain from child (married)	.005	.065	.048	.038	.001	-.006	-.009
Gain from child (single)	-.001	-.171	-.123	-.100	-.001	-.017	.003

B Blacks

	Ever live with biological father	Ever live with no father	Ever live with step-father	Ever live with cohabiting parents	Ever live with married parents	Mother single at conception	Mother single at birth
Baseline outcome value:	.433	.790	.653	.508	.918	.707	.636
Change due to:							
Welfare benefit	.043	-.014	-.028	-.011	.001	.000	-.028
Welfare reform	-.203	.134	.165	-.005	.011	.194	.189
Unilateral divorce	.009	.015	.010	-.001	-.002	-.004	-.007
Unemployment rate	-.019	.007	.009	.009	-.003	.011	.015
Sex ratio	.024	-.014	-.013	.008	.007	-.017	-.020

Male wage rate	.007	-.014	.001	.022	-.005	-.018	-.010
Female wage rate	-.062	.066	.049	.051	-.015	.057	.060
Child support enforcement							
Paternity establish. rate	-.015	.011	.011	.010	-.002	.012	.013
Collection rate/admin exp.	-.013		-.001	.000	-.008	.003	.008
Admin. expenditure/case	-.012	-.005	.007	.005	.006	.007	.007
Average tax rate							
Gain from marriage	-.011	.005	.005	.010	-.002	.011	.012
Gain from child (married)	-.051	.030	.032	.043	-.019	.032	.047
Gain from child (single)	.041	-.021	-.020	-.041	.026	-.015	-.033

C. Hispanics

	Ever live with biological father	Ever live with no father	Ever live with step-father	Ever live with cohabiting parents	Ever live with married parents	Mother single at conception	Mother single at birth
Baseline outcome value:	.859	.411	.315	.277	.971	.269	.167
Change due to:							
Welfare benefit	-.007	.020	.022	.041	-.008	.015	.009
Welfare reform	.045	.027	.060	.027	.022	-.047	-.051
Unilateral divorce	.055	-.118	-.087	-.072	.022	-.041	-.061
Unemployment rate	-.010	.008	.005	.008	-.004	.011	.010
Sex ratio	.018	-.012	-.011	.000	.002	-.016	-.019
Male wage rate	.073	-.138	-.103	-.070	.010	-.109	-.082
Female wage rate	-.066	.164	.114	.073	-.013	.125	.075
Child support enforcement							
Paternity establish. rate	-.008	.019	.013	.010	-.001	.016	.009
Collection rate/admin exp.	-.004	-.003	-.001	.005	-.003	.002	.004
Admin. expenditure/case	-.001	-.010	-.002	-.003	.002	.009	.001
Average tax rate							
Gain from marriage	-.008	.002	.002	.002	-.001	.010	.010
Gain from child (married)	-.035	.024	.023	.028	-.009	.020	.038
Gain from child (single)	.033	-.035	-.027	-.039	.020	-.011	-.034

Notes: See Table 4.

Figure 1: Trend in real monthly welfare benefit for a family of 4

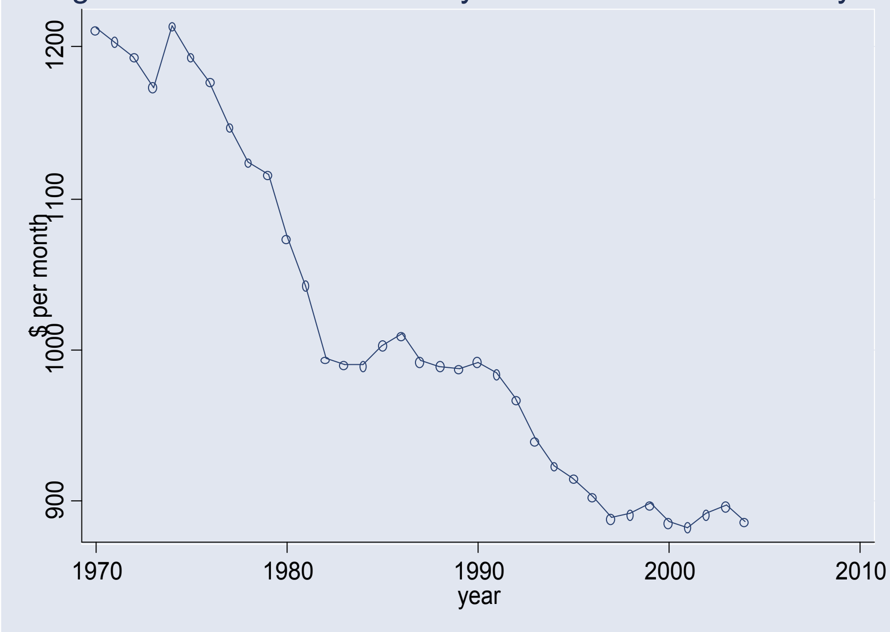


Figure 2: Welfare Reform (Waiver or TANF in place)

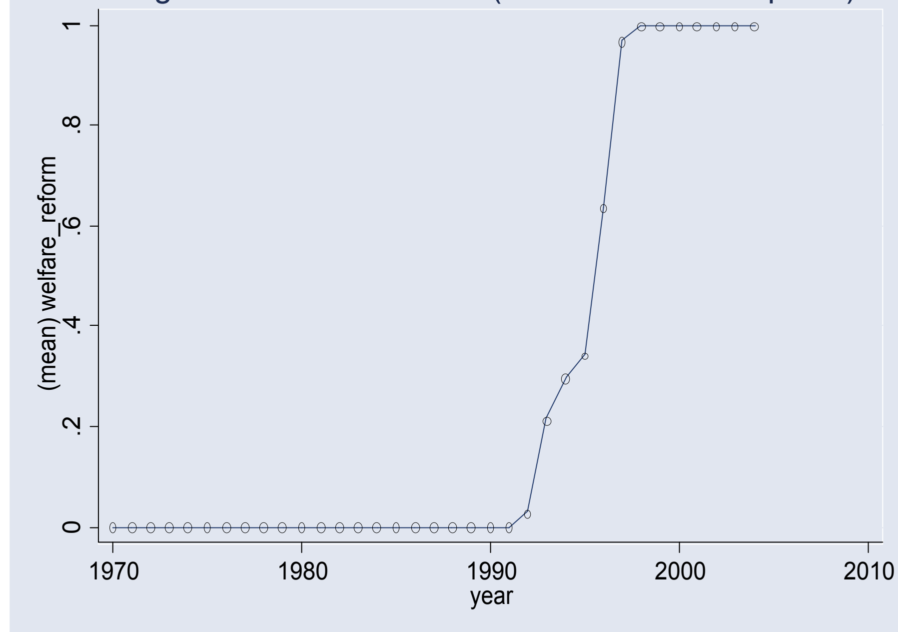


Figure 3: Trends in Child Support Enforcement

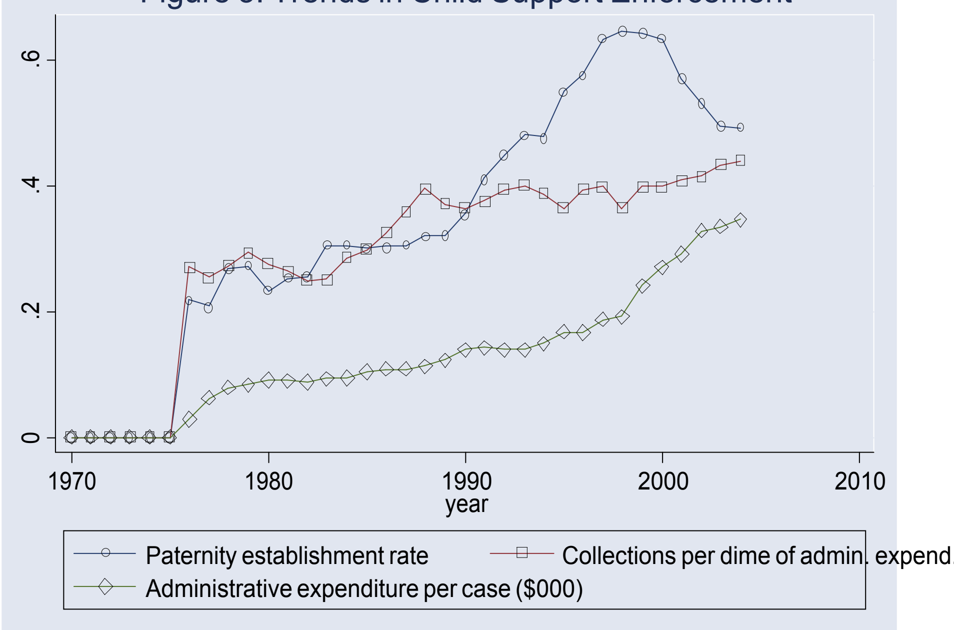


Figure 4: Unilateral divorce law

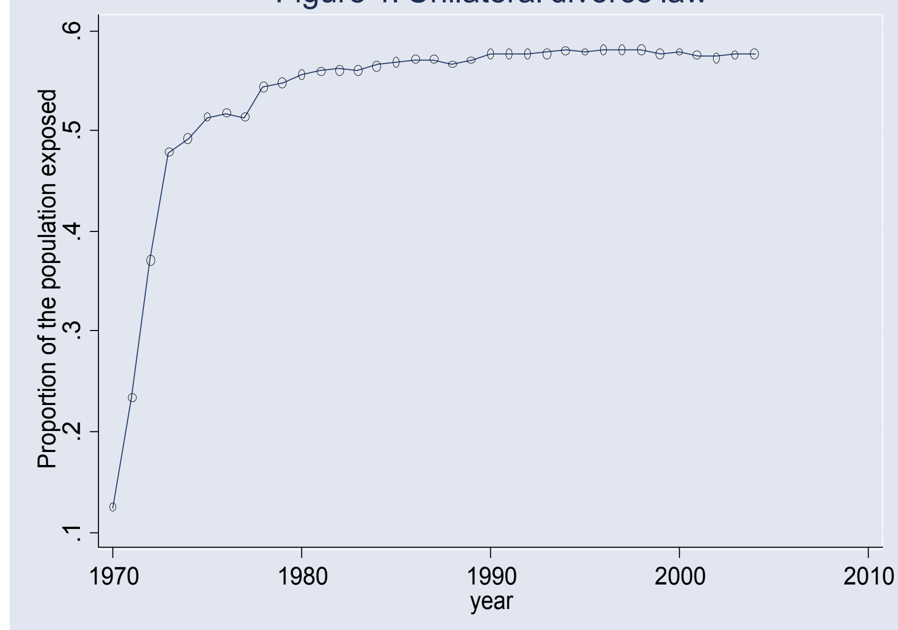


Figure 5: Trends in selected average tax rates at low income

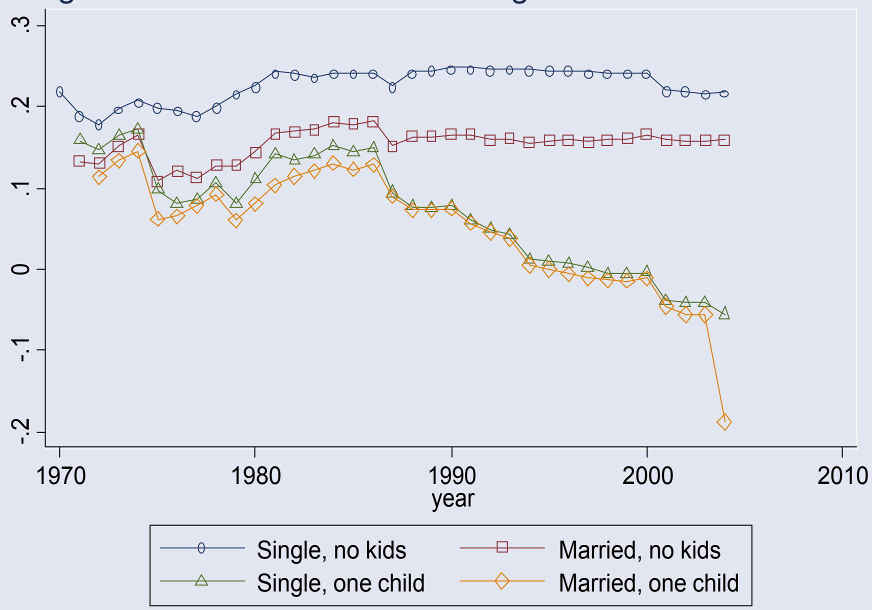


Figure 6: Trends in selected average tax rates at median income

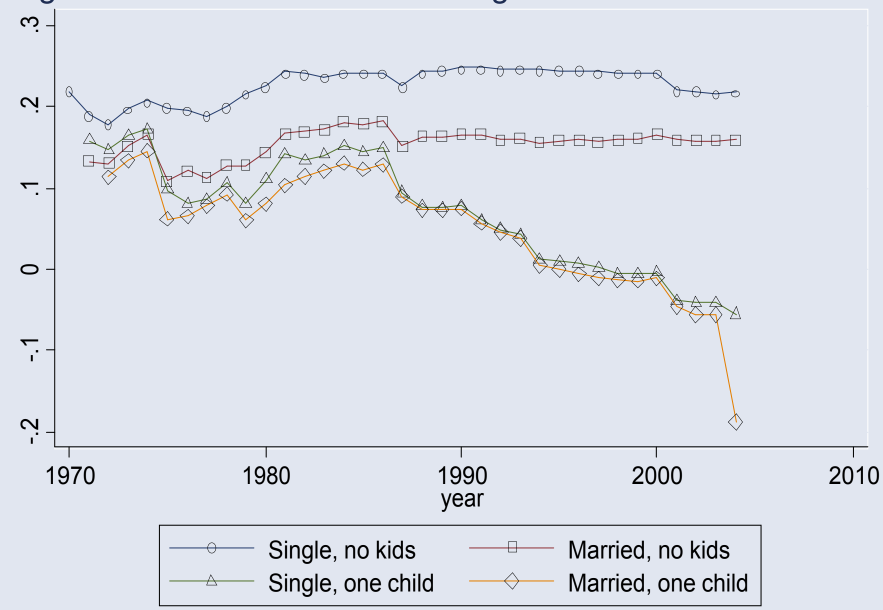


Figure 7a: Trends in male and female wage rates for whites

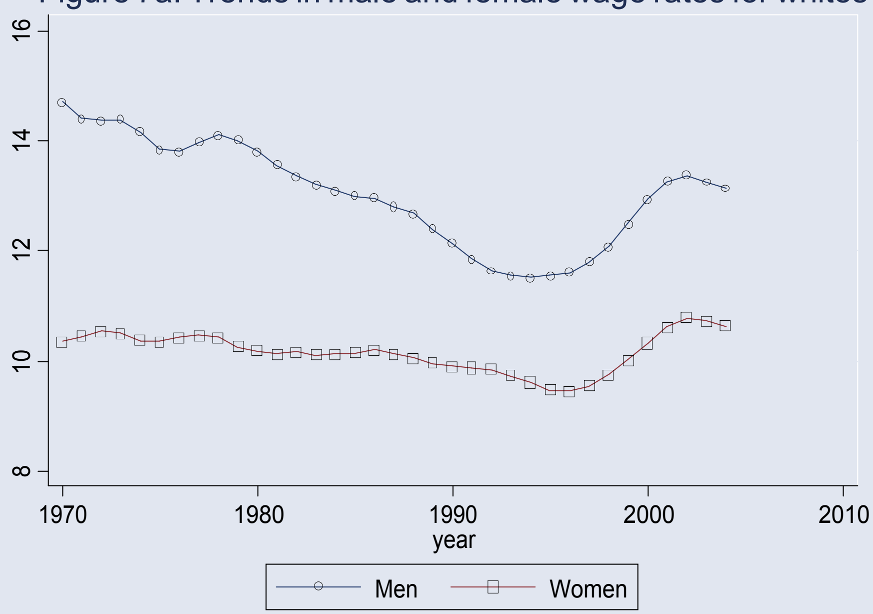


Figure 7b: Trends in male and female wage rates for blacks

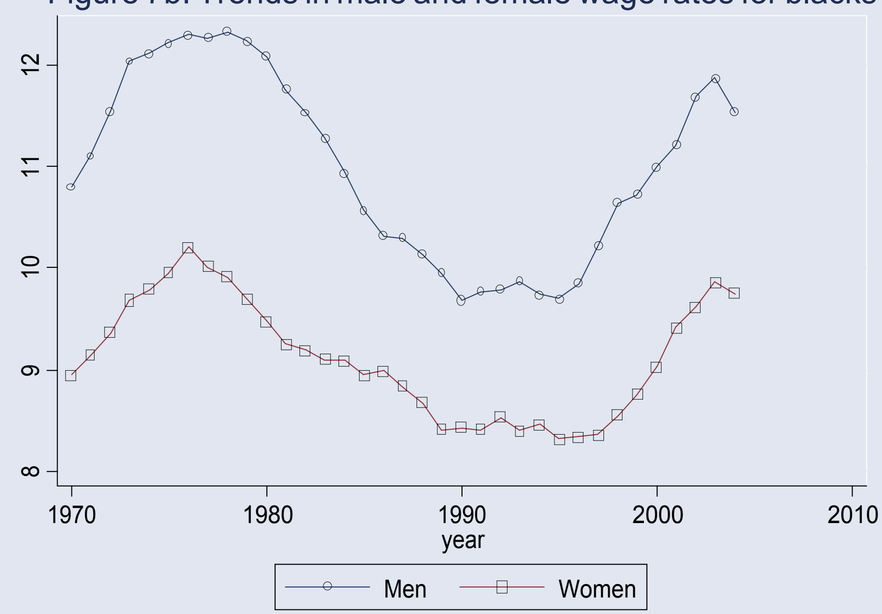


Figure 7c: Trends in male and female wage rates for Hispanics

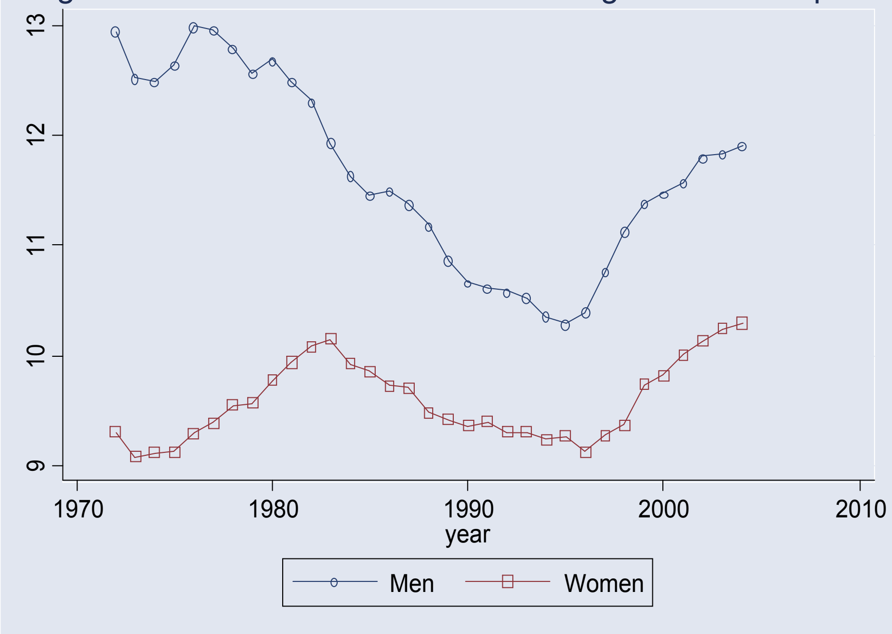
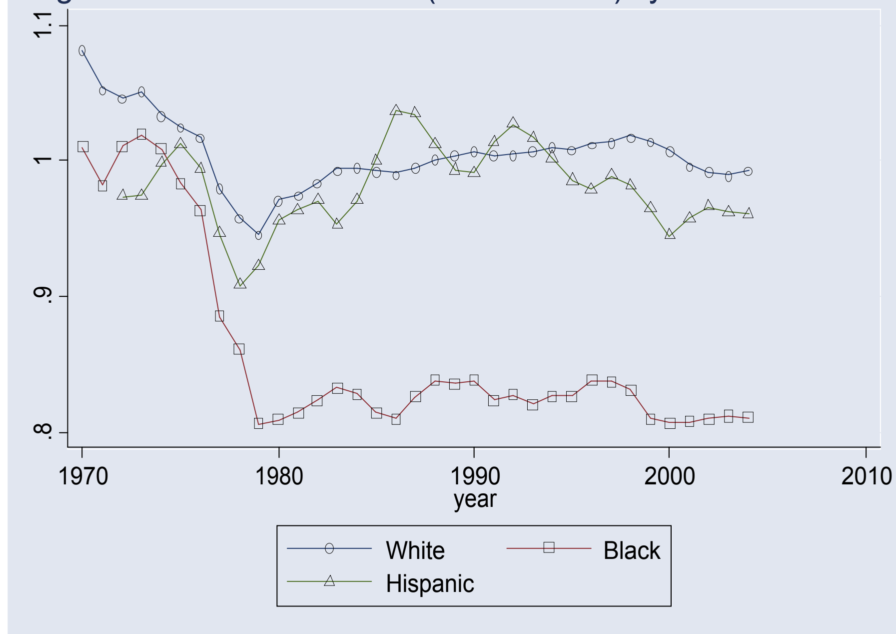


Figure 8: Trends in sex ratio (men/women) by race and ethnicity



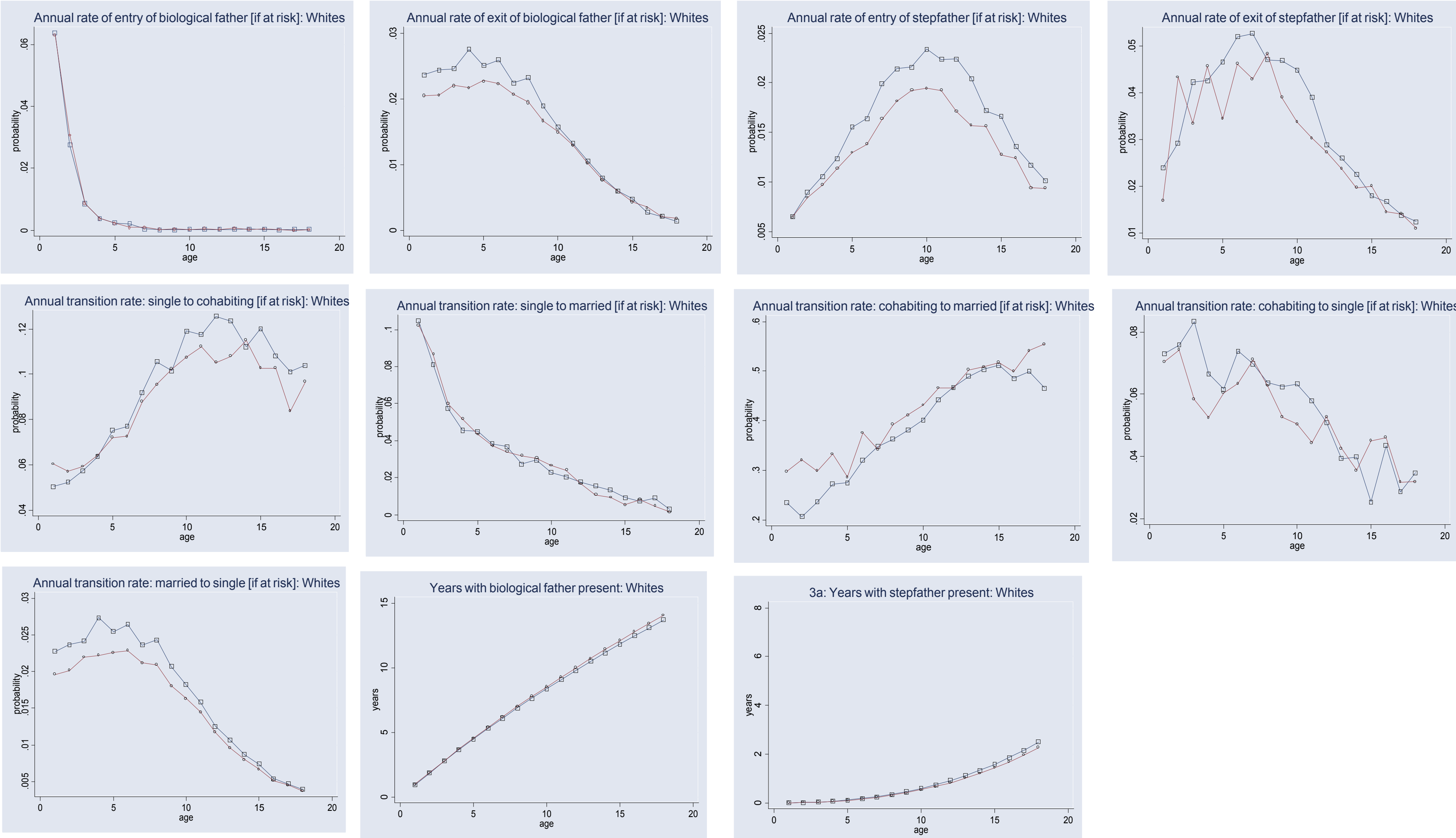


Figure9a:EffectofanIncreaseintheFemaleWageRateonSelectedFamilyStructureTransitionsofChildrenofWhiteMothers

Notes:Circles(inred)arethebaseline,squares(inblue)showtheeffectofaonestandarddeviationincreaseinthefemalewagerate

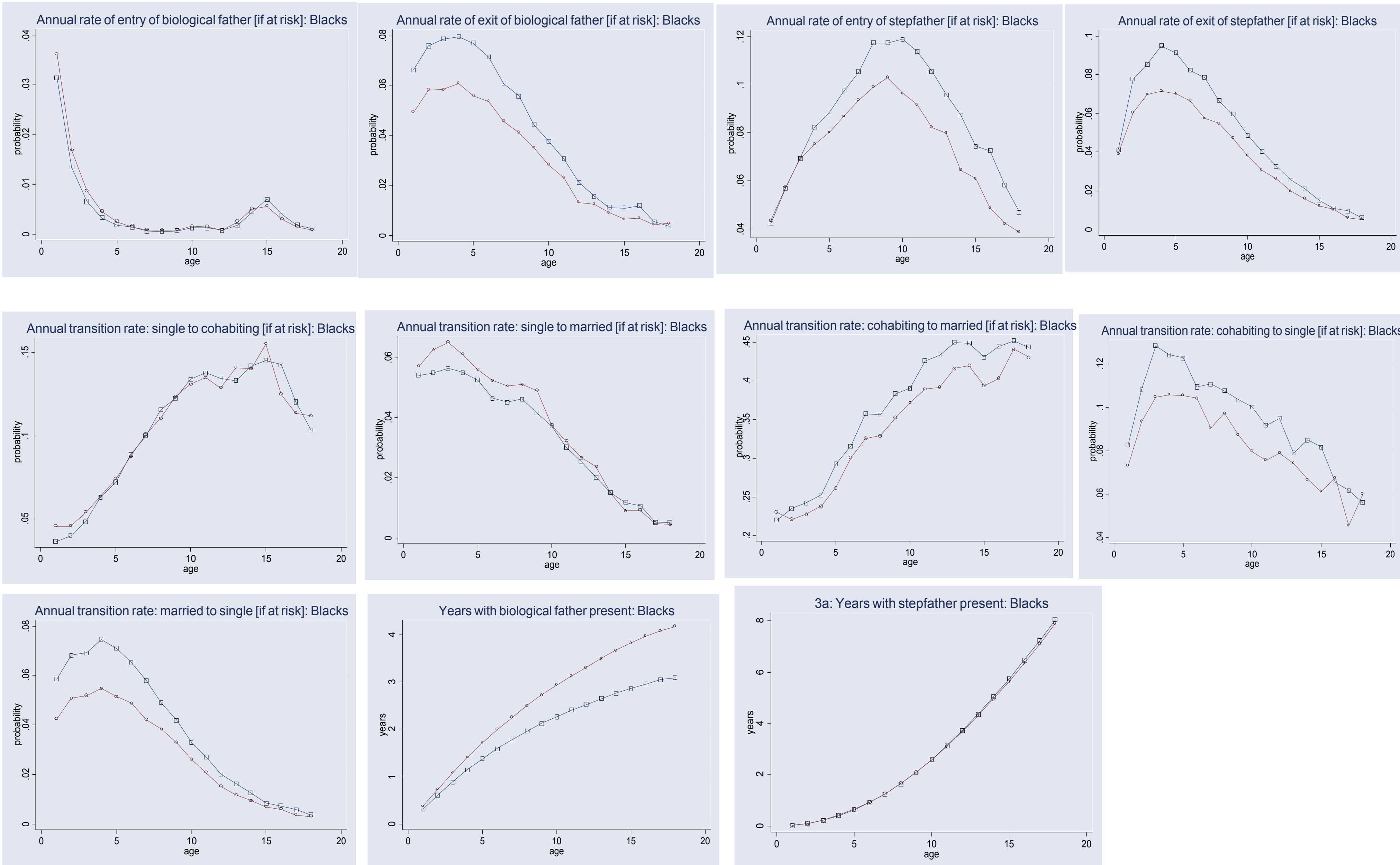


Figure9b:EffectofanIncreaseintheFemaleWageRateonSelectedFamilyStructureTransitionsofChildrenofBlackMothers

Notes:Circlesarethebaseline,squaresshowtheeffectofaonestandarddeviationincreaseinthefemalewagerate

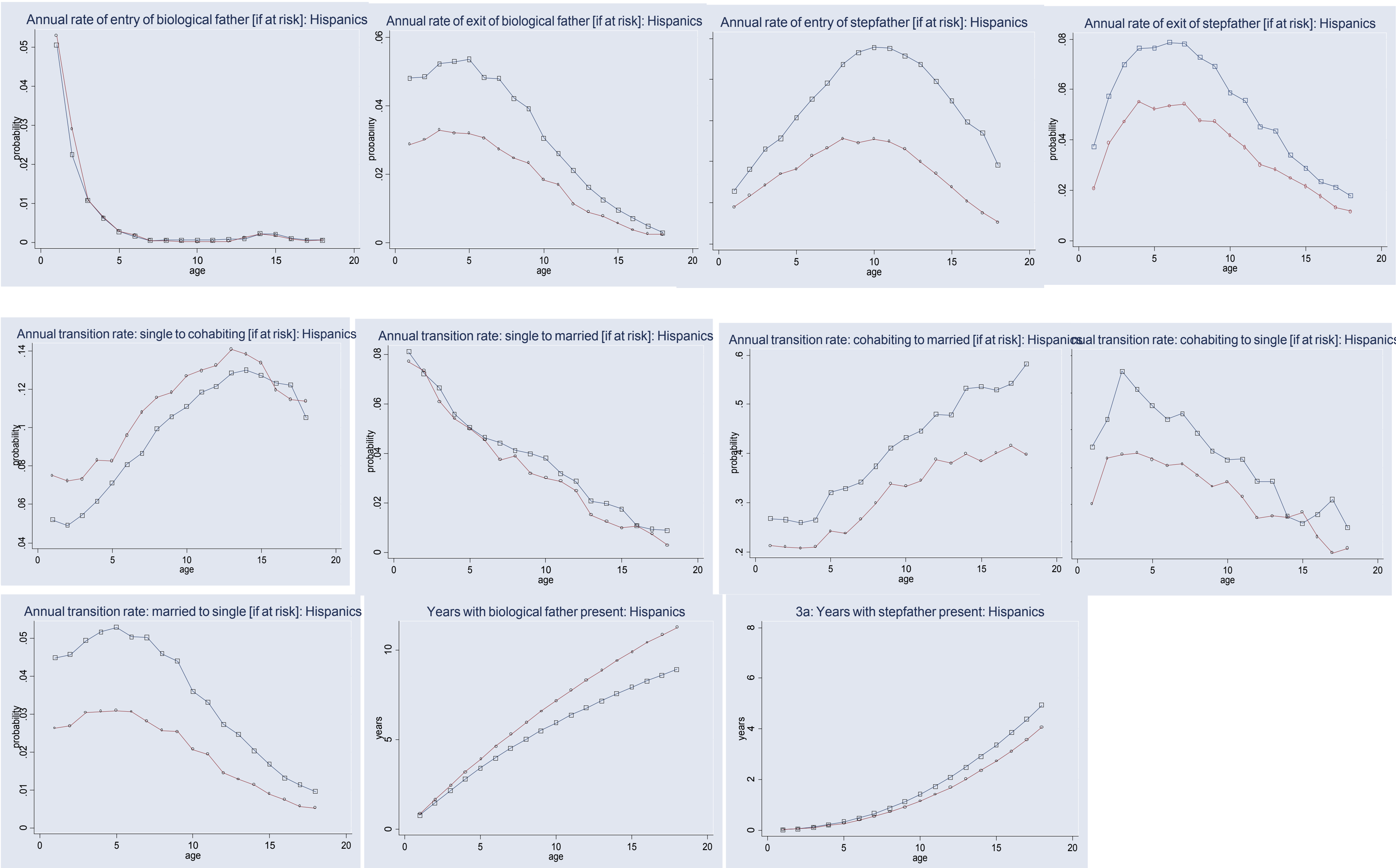


Figure9c:EffectofanIncreaseintheFemaleWageRateonSelectedFamilyStructureTransitionsofChildrenofBlackMothers

Notes:Circlesarethebaseline,squaresshowtheeffectofaonestandarddeviationincreaseinthefemalewagerate.