Race Differences in Cohort Effects on Nonmarital Fertility
in the United States

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January 2008

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Abstract

We employ newly developed methods to disentangle age, period and cohort

effects on nonmarital fertility ratios (NFRs) from 1972 to 2002 for black and white

women aged 20-44 in the United States. We focus on three cohort factors: family

structure, school enrollment, and the sex ratio. For both blacks and whites, cohorts with

less traditional family structures have higher NFRs. Other results differ by race. The

impact of school enrollment on NFRs is significantly *negative* for whites, but

significantly *positive* for blacks. The impact of sex ratio is significantly negative for

blacks, but insignificant for whites. If black women and white women had cohort

characteristics typical of the other group, age-specific NFRs for black women would

decline markedly, while those for whites would increase markedly.

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One of the most striking demographic changes in recent decades has been a sharp

upturn in the United States and a number of other countries in nonmarital fertility ratios

(i.e., the percentage of births occurring to unmarried women, for simplicity NFRs). For

instance, in 1940, only 3.8% of all births in the United States were out of marriage

(Ventura and Backrach 2000); by 2005 this figure had risen to 36.9% (Martin, et al

2007). Researchers have offered a number of explanations for these changes, including

liberalization of social attitudes and controls, generous welfare policies, and changing

absolute and relative economic opportunities for women and men (e.g., Aassve 2003;

Bitler et al. 2004; Cherlin 1990, 1992; Duncan and Hoffman 1990; Fitzgerald and Ribar

2004; Harding and Jencks 2003; Moffitt 2001; Pagnini and Rindfuss 1993; Parnell et al.

1994; Seltzer 2000; Sweeney 2002; Teachman, et al. 2000; White and Rogers 2000;

Willis and Haaga 1996; Zavodny 1999).

As with many other demographic processes, however, these changes probably

reflect a combination of age, period, and cohort effects. Because younger women are

generally more likely than older women to have nonmarital births, changes in the age

composition of the population can result in a higher ratio of nonmarital births to total

births. At the same time, changes in societal norms that promote more accepting attitudes

toward nonmarital childbearing may produce higher NFRs for people of all age groups, a

history or period effect. Finally, increases in nonmarital fertility may occur because more

recent birth cohorts are especially prone to such births and, through a simple process of

cohort replacement, the share of these births increases (cf. Hoffman and Foster, 1997; South 1999).

Thus, one central purpose of this paper is to examine the extent to which variations in nonmarital fertility can be attributed to cohort effects, rather than age or period effects. We utilize cohort theory and the notion of cohort opportunity structures to examine the extent to which three factors associated with cohorts can explain variations in age-specific nonmarital fertility ratios: 1) the prevalence of nonmarital births during a cohort's birth years, indicative of family structure during childhood; 2) the school enrollment of the cohort in the late teens and early twenties, indicative of early investment in human capital formation; and 3) the sex ratio (males to females) in young adulthood, traditionally seen as one indicator of the availability of marriage partners.

The second major purpose of this paper is to examine the extent to which cohort effects on NFR vary between blacks and whites. While nonmarital fertility ratios have increased substantially for both blacks and whites in the United States in recent years, values for blacks have always been substantially higher than those for whites. For instance, in 2005 69.6 percent of all births to African American women were nonmarital, while the corresponding figure for white women was 25.3 percent (Martin et al., 2007, p. 61). Common explanations for these differences include variations in neighborhood environments and in marriage and economic opportunities (e.g., Brewster 1994a,b; Browning 2004; Browning and Olinger-Wilborn 2003; Crane 1991; Miller, et al. 2001; Roche et al. 2005; South and Baumer 2000; South and Crowder 1999). We suggest that the extensive patterns of racial segregation in the United States result in black and white women growing up in very different social environments. Incorporating recent analyses

of contextual influences on cohort effects (Stockard and O'Brien 2002a, 2006), we test the hypothesis that the different contexts in which black and white women grow up and live are related to different patterns of influence of cohort related variables on NFR.

The third major purpose of this paper is methodological in nature, using newly developed techniques to isolate variance components corresponding to age, period and cohort. It is impossible to obtain parametric estimates of the effects of age groups, periods and cohorts on demographic processes, since including these effects in a model results in perfect collinearity. But, a different issue is the estimation of the variance uniquely associated with age controlling for period and cohort effects, or cohort controlling for age and period effects, or period controlling for age and cohort effects. Until now, no studies have estimated the impacts of age, period and cohort (controlling for the other two) on macro-level variations in NFRs or, more importantly, empirical estimates of factors believed to account for the magnitude of these effects. This paper begins to fill this gap by employing newly developed analytic techniques based on the Age-Period-Cohort-Mixed Model (O'Brien et al. 2008) to identify variations uniquely associated with cohort effects. Results indicate that a greater portion of the variation in nonmarital fertility is uniquely associated with cohort factors for whites than for blacks and that there are substantial differences in the factors influencing cohort variations in NFRs for blacks and whites.

Background

While demographers have utilized the distinction between age, period, and cohort for more than a century (e.g. Lexis, 1880), it was the work of Ryder (1965) that emphasized the importance of considering cohort effects in understanding social change

to the general sociological community. Ryder described birth cohorts as moving in a two-dimensional space of time and age. For instance, a cohort born in the early 1950s has grown older (moving through age) as years have passed (moving through time). Cohort theory and research have been used to examine how this movement affects age and period variations in a variety of areas including criminal behavior (O'Brien 1989; O'Brien et al., 1999; O'Brien and Stockard 2002; Savolainen 2000), suicide (Stockard and O'Brien 2002a,b), anti-black prejudice (Firebaugh and Davis 1988), opinions on democracy and Nazism (Weil 1987), parental values (Alwin 1990), political orientation and voting (Firebaugh and Chen 1995; Alwin and Krosnick 1991), sex role attitudes (Mason and Lu 1988), and intellectual skills (Alwin 1991).

There are two major tenets of cohort theory: the "life stage principle" and the "lasting effects principle." The "life-stage principle" posits that the experiences of members of one cohort differ from those of another because they experience historical events at different ages or developmental periods. Infants experience historical events, such as the establishment of universal suffrage or the transition to a market economy, differently than those who are twenty-one, those of middle age, and those who have retired (Elder and Caspi 1990, Elder 1974, 1979; Firebaugh and Chen 1995; Elder, Modell, and Parke 1993). The "lasting effects principle" suggests that certain events can produce lasting changes in the attitudes and behaviors of cohort members. To be sure these are cohort effects, however, these changes must be analytically distinct from those associated with age and period.

Cohort Opportunity Structures

Understanding why cohort differences arise is a crucial theoretical issue. Elder

(1996: 5) suggests that one can view the historical periods in which cohorts are born and grow to maturity as "opportunity structures." These structures may result from historical events, such as wars and revolutions, economic depression, transformations of economic systems, or opportunities to participate in education or the electoral process. Demographic differences between cohorts such as increased immigration, changes in sex ratios, relative size, and variations in family structure also generate differences in opportunity structures. Whatever the genesis of these opportunity structures, they affect the resources of, and constraints on, the members of different birth cohorts.

We examine the extent to which variation in nonmarital fertility arises from variation across time and race in the opportunity structures faced by successive cohorts of women. We hypothesize three different demographic factors related to opportunity structures that may influence cohort variations in NFR: 1) childhood family structure, 2) access to and investments in education, and 3) the availability of marriage partners.

A growing literature has documented the relationship of early childhood family structure to early sexual activity and nonmarital fertility (e.g. Albrecht and Teachman 2003; DeLeire 2002; Kahn and Anderson 1992; Musick 2002; South 1999) as well as cohort variations in other risky behaviors such as lethal violence (O'Brien and Stockard 2002). We use as our measure of cohort family structure the proportion of births that were nonmarital at the time of the cohort's birth. O'Brien and Stockard (2002) note that this measure, not surprisingly, is very highly correlated with estimates of the percentage of cohort members growing up in single parent families from 1910 to 1990.[1]

This measure of family structure has the advantage of being directly associated with a cohort's time of birth and, in addition, is available for a large group of birth

cohorts. We hypothesize that it will be positively related to age-period-specific NFRs for the cohort in later life. In line with earlier work in this tradition, we view this influence as structural in nature, reflecting the extent to which members of birth cohorts with more nonmarital births tend to grow up in situations with less adult supervision and monitoring, greater influence of peers, and fewer monetary resources. In line with cohort theory, we suggest that these influences impact not just those in nontraditional families, but all members of a birth cohort.

Nonmarried fertility ratios (NFRs) may be influenced by women's access to higher education and other forms of human capital acquisition during the late teens and young adulthood. We measure access to higher education and skill levels by school enrollment in the late teens and early 20s and propose two potentially offsetting influences at play in the relationship of this variable to NFRs. First, increased education may enhance the effectiveness of social controls that discourage nonmarital childbearing, thus reducing nonmarital fertility ratios. Moreover, individual women seeking more education may choose to delay childbearing, especially nonmarital births. East (1998, p. 159), for example, concludes that girls' school and job aspirations are an important determinant of their attitudes toward nonmarital births, regardless of race or socioeconomic status.

A second possible influence, however, is the "independence effect," where women with more education (and, hence, likely greater potential earnings) are less reliant on a marriage partner in order to bear and raise children (Becker 1981; Bennett et al 1989; Espenshade 1985: 222-225). Thus, a cohort's higher enrollment in education during the late teens and early adult years could be positively, negatively or not

associated with NFRs throughout a cohort's primary childbearing years, contingent on

the relative magnitudes of the "discouragement/delay effect" versus the "independence

effect."[2]

Third, we include a simple measure of mate availability. A relatively large

literature documents the relationship of mate availability to marriage (Billy and Moore

1992; Blau, et al. 2000; Cready, et al 1997; Fossett and Kiecolt 1990; Guttentag and

Secord 1983; Lichter, et al. 1991, 1992; McLaughlin and Lichter 1997; Raley, 1996;

South 1996; South and Lloyd 1992; Willis 1999; Wilson 1987). Simply put, when there

are more available men, the chances of marriage, and thus of bearing children within

marriage, are greater. We expect that the cohort's ratio of men to women will be

negatively related to NFRs throughout the childbearing years, but especially so for black

women, where the ratio is less than one. Wilson (1987) and Willis (1999), among others,

have argued that the effects of mate availability are most critical when the ratio falls

below unity.

Finally, it is possible that the influence of cohort opportunity structures may vary

depending upon a cohort's age (cf. Kahn and Mason 1987; Pampel 1996). Particularly, it

is possible that a cohort's opportunity structures could affect the probability of nonmarital

births during the teen years in different ways than during the adult years. This could be

expected given both differences in the dynamics surrounding marriage decisions in the

teen and adult years as well as differences in impulse control and general decision

making patterns between teens and adults (see Gray et al 2006a,b). In addition, at least

for blacks, and increasingly for whites, the incidence of marriage after conception has

become quite low for teens. Thus, we focus our analysis on adults (women ages 20-44),

for whom, especially for blacks, there is much greater variability in nonmarital births.[3]

The Contextual Effect of Race

      Recent cross-cultural studies (e.g. Pampel, 1996, 2001; Stockard and O'Brien,

2002a, 2006) have added to understandings of cohort effects by demonstrating that the

impact of cohort opportunity structures may vary across different social contexts. In other

words, the extent to which a given factor may produce risk or opportunities for a cohort

may vary from one society to another, as a societal context either buffers or exacerbates a

cohort's risks or opportunities. We suggest that such contextual effects may be relevant in

understanding the ways in which opportunity structures have affected black and white

women's nonmarital fertility.

      A large literature has documented extensive segregation in American society,

both historically and in contemporary times, and especially between African Americans

and whites. The vast majority of blacks and whites have remarkably little meaningful or

repeated contact with someone from the other race group in their communities (e.g.

Massey and Denton, 1993), schools (Orfield et al, 1997), churches (Lincoln and Mamiya

1990; Roof and McKinney 1987), peer and friendship groups (Moody, 2002; Berry

2006), workplaces (Landry 1987), or families (Tucker and Mitchell-Kernan 1990). A

large literature has also examined the way in which the different neighborhoods in which

blacks and whites typically live contribute to racial differences in early sexual activity,

premarital childbearing and rates of marriage (e.g., Brewster 1994a,b; Browning 2004;

Browning and Olinger-Wilborn 2003; Crane 1991; Miller, et al. 2001; Roche et al. 2005;

South and Baumer 2000; South and Crowder 1999).

We extend this line of thinking by noting that *cohorts* of white and black women in the United States grow up within different social contexts – in neighborhoods, schools, churches, and families that are largely separate. Given large differences between white and black communities in demographic characteristics and economic well being, these differences result in different cohort opportunity structures. Theoretically, these different opportunity structures could affect African Americans and whites in two ways. First, to the extent that the cohorts within the two groups differ in average levels of cohort characteristics (e.g. different average family structures, educational levels, and/or sex ratios), they can contribute to different levels of nonmarital births throughout the life cycle.

In addition, however, it is possible that the impact of cohort related variables could vary from one social context to another. For instance, numerous researchers have noted that African American communities have developed strategies of coping with disadvantaged contexts that can compensate for lower levels of economic and social resources (e.g. Burton 1990, 1996; Stack 1975,Taylor et al 1988, Taylor et al, 1997). To the extent that such coping strategies compensate for losses in opportunities experienced by a cohort, it would be expected that cohort characteristics could have different effects from one race group to another.

Summary and Hypotheses

To summarize: In this paper we examine variation in age-period-specific nonmarital fertility ratios (the percentage of live births to unmarried women) for whites and blacks over the latter third of the 20th century. We first examine the extent to which this variation represents age, period, and/or cohort effects and the extent to which

variations in opportunity structures for individual cohorts can account for these effects. We hypothesize i) that birth cohorts made up of higher percentages of members who were born out of wedlock will have higher nonmarital fertility ratios throughout their childbearing years; ii) that the influence of education in the early adult years on a cohort's nonmarital fertility may be positive, negative or non-existent, depending on the importance of education in delaying childbearing and enhancing social controls relative to its importance in generating potentially higher earnings and greater independence for women; and iii) that cohorts with higher ratios of males to females will have lower nonmarital fertility ratios, especially so if the ratio falls below unity. We speculate that each of these variables captures an array of opportunities that accrue to cohorts and that have varied across cohorts born throughout the 20th century.

Second, given the very different contexts in which black and white women grow to maturity in the United States, we hypothesize that these opportunity structures may have varying effects for the two race groups. Thus, we compare results between the race groups to consider the ways in which opportunity structures and their effects vary across the different social contexts associated with race in the United States.

Detecting Cohort Effects

To test these hypotheses we employ a newly developed methodology for disentangling age, period, and cohort effects. Given the recent nature of this technique we describe its characteristics in greater detail in the sections below. We begin, however, by describing the data used in our analysis.

Structure of the Cohort Data

Data on unmarried births were taken from *National Vital Statistics Reports* (2000, 48:16 and 2002, 50:10). Data on total births are from *Vital Statistics of the United States* (www.cdc.gov/nchs/births.htm). We limit our analysis to cohorts for which full data are available for whites and blacks, beginning in the year 1972.

The cohorts used in the analysis, their birth years, and the years at which they reached ages 20-24 and 40-44 are given in Table 1. As is typical in cohort analyses, 5-year birth cohorts are used, providing a range that is wide enough to provide reliable statistical estimates but narrow enough to ensure that members of a group have had relatively similar life experiences. The oldest cohort in the analysis was born between 1928-1932 and the youngest was born between 1978-1982. Tables 2 and 3 illustrate the structure of the data set. Table 2 includes the age-specific data for whites and blacks, while Table 3 includes the cohort-specific measures for both race groups. Table 4 presents summary descriptive statistics on each variable for both groups.

[Tables 1, 2, 3, and 4 about here]

In Table 2, data for each age group are portrayed in the rows, and data for each year are in the columns. The first two elements of each cell are the percentage of nonmarital births (NFRs) for a given age and period for whites and blacks, respectively. The third element is the number assigned to each birth cohort, corresponding to those listed in Table 1. For whites, values of NFRs vary from 2.7, for 25-29 year-olds in 1972, the first year for which we have data, to 44.6, for 20-24 year-olds in 2002. Values for blacks are consistently higher than the values for whites. As summarized in Table 4, the minimum NFRs for black women are 23.1, for 30-34 and 35-39 year-olds in 1972. The

maximum is 81.3, which occurred for the same age group and period as the maximum for whites: 20-24 year olds in 2002.

The influence of age on nonmarital fertility ratios can be seen by comparing data in the rows in Table 2. For whites, calculating the mean across all of the periods, the youngest age group (20-24) has the highest average NFR (23.4) while the oldest age group has the second highest (11.6). The middle age group (30-34) has the lowest NFR (7.4). The pattern is curvilinear with the 20-24 year olds having by far the highest NFR. Comparing the columns in Table 2 shows the extent to which NFRs have changed over historical periods. There has been a strong monotonic increase in the mean NFRs over the seven periods in our analysis. The mean was 4.1 in 1972 and 20.8 in 2002. Strong changes can be seen in each age group (across each row). For instance, the percentage of births to white unmarried women increased from 6% for 20-24 year-olds in 1972 to 45% in 2002. Changes in other age groups are also dramatic. For instance births to unmarried white women in their later 20s rose from only 3 percent in 1972 to over a fifth of all births by the turn of the century.

The diagonals can be compared to examine cohort effects. The oldest cohort for which full information is available is cohort 5, born in 1948-1952 and age 20-24 in 1972. Again, for whites, following the data for cohort 5, from the upper-left cell through the diagonal, it can be seen that its age-period-specific NFR was 5.7% in 1972, the lowest for all years for that age group. It fell somewhat when the cohort was in its late twenties and early thirties, but then rose to 8.9% in 1987, when the cohort was 35-39 years of age and rose again to 15.1% in 1992 when the cohort was 40-44 years of age. Cohort 6, in the next diagonal to the right, had an age-period-specific NFR of 8.3% in 1977, at the age of

20-24. The value was smaller over the next ten years, but then rose to a value of 11.4 for 35-39-year-olds in 1992 and 14.2% when the cohort was 40-44 in 1997.

Again, Table 3 reports the values of the measures of cohort characteristics for whites and blacks for each of the cohorts in the analysis. The first column is the cohort number, corresponding to Table 1. The second and third columns are the percentage of the birth cohort born outside of marriage, our measure of cohort variation in family structure, for whites and blacks respectively. Values for this measure are substantially lower for the older cohorts than for more recent cohorts, although the trend is not strictly linear, especially for whites. For this group, the lowest value occurs with cohort 5 (born 1948-1952) and the highest value with cohort 11. Values are substantially higher for blacks than for whites. In fact, as shown in Table 4, the range of values for white cohorts and black cohorts on this variable do not overlap: the maximum value for whites is 10.6 (for cohort 11), while the minimum value for blacks is 14.7 (for cohort 1)

The fourth and fifth columns in Table 3 give the sex ratios (based on national data) when the cohort was 20-24, which we calculated by simply dividing the number of men in this age group for a given cohort by the number of women and multiplying by 100. Data to calculate sex ratios also came from the U.S. Bureau of the Census: the web site at census.gov/popest/archives/pre-1980/PE-11.html for data before 1980 and selected editions of *Statistical Abstract of the United States* for later years. Data are based on the "resident" population, excluding those who are institutionalized. Values range from 100.9 to 107.0 for whites and 88.6 to 97.5 for blacks. The lowest sex ratios occur with the earlier cohorts and the larger sex ratios with the later cohorts, reflecting increased life expectancy for males in more recent periods. Note that the ranges of values for whites

and blacks do not overlap, due in large part to high incarceration rates and high death

rates among black men.

The final two columns in Table 3 give the percentage of women enrolled in

school at ages 18-21, when the cohort was 18-22 years of age. Values for cohorts 3 and

later were obtained directly from the U.S. Bureau of the Census

(http://www.census.gov/population/www/socdemo/school.html). Values for earlier

cohorts were predicted using regression techniques with median years of schooling at age

25-29 as a predictor (R squared = .93) and adjusting for differences between blacks and

whites using historical trends. Values range from 16.6 to 56.2 for whites and 14.3 to 51.4

for blacks, with the lowest values for the earliest cohorts. For each cohort, whites have

higher school enrollment in young adulthood than blacks, although the values are much

closer than for the other cohort characteristics. For whites the percentage enrolled in

school increases monotonically for the cohorts in the analysis. For blacks, however, the

increase is not monotonic. The value for cohort 8 (32.5) is lower than that of the two

adjacent cohorts (36.0 for cohort 7 and 38.0 for cohort 9).

Identifying Age, Period and Cohort Effects

While we could continue to follow the experiences of each birth cohort by tracing

descriptive data through the diagonals of the table, such comparisons do not allow us to

parsimoniously and statistically separate the impact of age patterns and historical trends

from cohort effects. Traditionally, the most difficult problem facing analysts who wish to

understand the presence of cohort effects is the linear dependence of age, period and

cohort. If one knows the age and period, one knows the cohort associated with the age-

period-specific NFRs. This linear dependence has impeded the development of reliable

and accurate estimates of the existence of cohort effects, independent of age and period, using OLS techniques.

One way around this impediment is to substitute one or more characteristics of the cohorts that are theorized to affect the dependent variable rather than to use dummy variables for cohorts. For example, we might measure the percentage of the birth cohort born to unwed mothers (the cohort measure of family structure used here). This breaks the linear dependency between the age-group dummy variables, period dummy variables, and the effects of cohorts. It also may provide an explanation for the variation of cohorts on the dependent variable (this variation may be significantly associated with the variation in the cohort characteristic). A disadvantage is that the cohort characteristics may not account for all of the variation associated with cohorts that is not associated with the age and period dummy variables.

The approach that we use allows us to estimate the amount of variation in the age-period-specific NFRs associated with cohorts after controlling for age and period dummy variables and then, using cohort characteristics, assess how much of that variance is accounted for by the cohort characteristics. The method was recently introduced (O'Brien et al., 2007) and involves a mixed model in which age and period dummy variables are treated as fixed effects and cohorts are treated as random effects. The dummy variables for age and period provide strong controls for the effects of maturation and history, while the remaining variation associated with cohorts (after this strong control) is estimated by the variance associated with the random effect of cohorts.

Since this method is new, we outline its rationale (for details see O'Brien et al. 2007). In an Age-Period-Cohort model, we cannot estimate all of the dummy variables

for age groups, periods, and cohorts simultaneously (even when omitting a reference category for each set of dummy variables). Using a mixed model, however, we can control for the age groups and periods using dummy variables and in the same analysis (using dummy variables for cohorts) examine whether there is a pattern for the residuals that is consistent with cohort effects. If there are cohort effects, and we have taken into consideration the age group and period dummy variables, we would expect that the residuals (observed minus predicted NFRs) would form a pattern along the cohort diagonals. If some cohorts have particularly high NFRs, their residuals (after controlling for age and period) should be positive. If other cohorts have particularly low NFRs, their residuals should be negative. It is the variance of these cohort diagonals that the random effects term for cohorts estimates. [4]

In our analyses we focus on cohorts to see if there is a unique effect of cohorts after controlling for age and period. The question is whether cohort membership is associated with NFRs after controlling for age and period. We can do the same thing for periods and age-groups by treating one of them as the random variable and the other two sets of dummy variables as fixed. We can evaluate if, for example, periods are associated with NFRs after controlling for age groups and cohorts. The method does not tell us how to attribute the variance associated jointly with two or more of these sets of dummy variables (a common problem in regression analysis), and this problem is exacerbated, because any linear effect of time of cohort birth and NFRs is fully accounted for by the age group and period dummy variables (O'Brien 2000). The method does assess whether there is a *unique* association between cohorts and the age-period-specific NFRs.

Specifically, we use the following procedure to estimate the unique variance associated with cohorts. First, to determine the amount of variance associate with NFRs, we run a mixed model with only the intercept as an independent variable. The residual variance from this model is the total variance of the NFRs that we seek to explain [var(tot)]. Second, we then run a mixed model with just the fixed effects for the age and period dummy variables. The residual variance tells us how much of the variance in NFRs is not associated with the age and period dummy variables [var(age,period)]. Third, we run a mixed model that adds the cohort dummy variables as random effects. The residual variance tells us how much variance in the NFRs is not associated with age, period, and cohort [var(age, period, cohort)].

We are then able to estimate the amount of variance uniquely associated with cohorts by calculating the difference, var(age,period) - var(age, period, cohort). The proportion of the total variance associated uniquely with cohorts is calculated as: [var(age,period) - var(age, period, cohort)] / var (tot), and the proportion of the total variance associated with age, period and cohort is calculated as: [var(tot) - var(age, period, cohort)] / var(tot). Later, using these variances, we report the proportion of the total variation in NFRs associated with age, period, and cohort and the proportion of the total variation uniquely associated with each of these factors.[5]

Empirical Specifications

We implement this approach using the Proc Mixed procedure in SAS (2004, version 9.1). In addition to allowing us to model fixed and random effects, it also allows us to estimate the autocorrelation of "adjacent" observations within cohorts (they are treated as repeated measures). For example, from Table 2 we see that cohort 2 (those

born between 1933 and 1937) has several age-period-specific NFRs as does cohort 3

(those born between 1938 and 1942), etc. We estimate an AR(1) model that takes into

consideration the correlation between the residuals for these observations within the

cohorts.

We label this approach the Age-Period-Cohort Mixed Model. When we add

cohort characteristics to specifically model the cohort effects, we label this the Age-

Period-Cohort Characteristic Mixed Model. The second approach provides estimates of

the variation uniquely associated with cohorts, with age groups, periods, and the cohort

characteristics controlled. Models with cohort characteristics can also provide improved

estimates of age and period effects.   In both models we can estimate the autocorrelation

within cohorts and the addition of the random cohort effects should improve the estimates

of standard errors of the parameters.

We use the natural log of the age-period-specific NFRs as the dependent variable

in all of our analyses. The use of the natural log can incorporate likely nonlinearities in

the relationships, which is helpful since we anticipate specific cohort variables (esp.,

family structure and school enrollment) to have initially larger and then smaller effects,

which is consistent with the logarithmic specification. Put differently, we expect cohort

related variables to affect age-period-specific nonmarital fertility ratios proportionately

whether those rates are for the youngest or oldest age groups or the earliest or latest

periods in our data. Following O'Brien and Stockard (2003), we also log the measures of

the cohort characteristics. The double logarithmic transformation ensures that

proportionate shifts in the percentage of each cohort variable are associated with

proportionate changes in NFRs, thereby facilitating the interpretation of coefficients ("b")

associated with the cohort characteristics: a one percent change in the independent

variable will be associated with a "b" percent change in the dependent variable.[6]

    We first examine the effects of age and period (as fixed effects) with cohorts as

random effects (Model 1) and then add the three measures of cohort opportunity structure

(Model 2): family structure at birth (the prevalence of nonmarital births during the birth

years of a cohort), investments in human capital in young adulthood (school enrollment

from ages 18-21), and the availability of marriage partners (the sex ratio of males to

females in young adulthood). In Model 3 we add an autocorrelation term to better

estimate the standard errors of the parameter estimates. We compute the results

separately for whites and blacks and then examine more closely the implications of the

race differences that we find. As noted above, the values of the cohort characteristics, as

well as the level of NFRs, differ markedly between whites and blacks.

    Through simple substitution in the regression equations, we examine ways in

which black and white NFRs would theoretically alter if each group had the cohort

characteristics of the other, while retaining the underlying pattern of age, period, and

cohort effects. A predicted score is calculated for each age-group and year combination

within each race group, the results are then converted from the logged values to the actual

values, and then averages are computed across age and period.

Results

Comparing Cohort, Age, and Period Effects

    Employing the techniques described above, we calculated the percentage of the

total sum of squares for the nonmarital fertility ratio that is associated with age, period,

and cohort in each of our analyses, both in total and uniquely. For both whites and blacks

age, period, and cohort together account for 99% or more of the total sum of squares. In other words, these three factors account for the vast majority of the variation in the age-period-specific NFRs. For whites slightly more than one-fourth of the variation (28 percent) is associated uniquely with one of these factors: 6 percent with cohorts, 15 percent with age groups, and 7 percent with periods. For blacks, the explained variation is more confounded, with only 16% of the total variation uniquely explained. One percent is uniquely associated with cohorts, 8 percent uniquely associated with age, and 7 percent uniquely associated with period.

Results reported in Model 1 in Table 5 indicate that the variance due to cohorts is statistically significant for both whites and blacks. Raudenbush and Bryk (2002) caution that the best test for assessing the significance of such random effects variance estimates is the likelihood ratio chi-square test, which requires that we compare minus 2 times the log likelihood ratio for the model that contains the random effects for cohorts with the same statistic for a model that does not contain the random effect for cohorts. Minus two times the log likelihood for the model that does not contain the random cohort effect (not shown in Table 5) is -25.7. When we subtract minus two times the log likelihood in Model 1 (Table 5) for whites from this, the result is a likelihood ratio chi-square statistic with a chi-square of 47.7 [= –25.7 – (–73.4)] with one degree of freedom for the variance estimate. This is statistically significant (p < .001). The comparable calculation for the black data results in a chi-square of 11.11 with one degree of freedom (p < .001).

[Table 5 About Here]

Model 2 adds the three measures of cohort opportunity structure, and Model 3 adds the autocorrelation term to adjust for multiple observations within a cohort. The

results in Models 2 and 3 are very similar. We focus on those in Model 3 because they

represent the most articulated specification. [7]

In Model 3 the likelihood ratio chi-square test supports the results reported for the

test of the significance of the variance estimates for the random effects of cohorts. While

the residual cohort variance declines substantially for whites when explicit cohort

characteristics are included, it remains significant (p<.05). For blacks, the residual cohort

variance also declines substantially, to the point where it is no longer significant at the

.05 level. Thus, the three explicit cohort variables account for most of the influence of

cohort-specific variance, especially for black women.

We can estimate the extent to which the cohort characteristics capture the random

cohort effects by assessing the reduction in the random cohort variance after we add the

cohort characteristics to the model. For example, the cohort variance for whites in Model

1 is .0448 and in Model 3 is reduced to .0020. That represents a reduction in the variance

associated with the cohort diagonals of 96% [= ((.0448 - .0020)/.0448) × 100]. For

blacks the reduction from Model 1 to Model 3 is 89%. For both race groups most of the

unique variance associated with cohorts is associated with the three characteristics of

cohorts that are central to our analysis.

Cohort Effects and the Context of Race

The coefficients associated with the three cohort related variables may be

interpreted as regression coefficients and indicate that the cohort variables are related to

age-period specific nonfertility fertility ratios, but that the social context of race very

much influences the effects of two of these characteristics. We had expected that the

influence of cohort family structure would be positive, and our hypothesis receives strong

support for both whites and blacks. For both race groups, cohorts born in environments with more nonmarital births are more likely to have nonmarital births throughout adulthood. As noted above, because we have logged both the independent and dependent variables, the coefficients may be easily interpreted in (approximate) percentage terms. For white cohorts, an increase of 1 percent in a cohort's nontraditional family structure is related to a .50 percent increase in nonmarital births throughout the cohort's lifespan as covered in our analysis; for black cohorts, a similar increase in childhood family structure is related to a .42 percent increase in adult nonmarital births, with other variables in the model held constant.

In contrast, the influence of education works in opposite directions for whites and blacks. For white cohorts, education enrollment in young adulthood significantly *reduces* nonmarital fertility throughout the adult years, perhaps reflecting the hypothesized "discouragement" or "delay" effect described earlier. A one percent increase in enrollment from when the cohort was 18-21 is related to a .90 percent decline in nonmarital fertility throughout adult ages. For black cohorts, education enrollment in young adulthood is significantly related to *higher* nonmarital fertility ratios, perhaps due to the dominance of the "independence" effect. A one percent increase in school enrollment is related to a .26 percent increase in the nonmarital birth ratio.

The predicted negative influence of the sex ratio is supported only for blacks, where, as expected, cohorts with more males relative to females are significantly less likely to have nonmarital births throughout adulthood. The magnitude of the coefficient indicates that a one percent increase in the ratio of males to females in young adulthood is related to a 2.4 percent decrease in NFRs for black adult cohorts. For whites, there is a

positive, but insignificant, relationship. These findings are consistent with arguments by

Wilson (1987) and Willis (1999) that mate availability is particularly crucial to black

women because the sex ratio for blacks falls below unity.

Age and Period Effects Controlling for Cohort Characteristics

Although not central to our hypotheses, the results related to age and period

effects should be noted, for they confirm the importance of including cohort effects in an

analysis of variations over time or across age groups. Because age, period, and cohort

effects are correlated and can be confounded, it is best to interpret the age and period

effects in models that contain explicit cohort characteristics that are strongly related

to the dependent variable. However, even in this case, age and period may receive credit

for the remaining effects of cohorts that are linearly associated with age and period.

These characteristics are fixed and specific to each cohort and their effects are controlled

for in Models 2 and 3 when we examine the age and period dummy variables. That is not

the case for Model 1 that includes cohorts as a random variable but does not contain

cohort characteristics. The random effects represent the residual variance associated with

the cohort diagonals.

The coefficients associated with age in Model 3 indicate that, for both whites and

blacks, once cohort characteristics are controlled and relative to ages 40-44, NFRs are

significantly higher for ages 20-24 and lower for ages 30-39. This pattern matches that

observed in Table 2 and described above. The coefficients associated for period differ for

whites and blacks and also from the pattern seen in Table 2, where a monotonic increase

over time appeared. For whites, NFRs, once cohort characteristics are controlled, are

significantly lower from 1972 through 1987 than in 2002, but NFRs in 1992 and later are

not significantly different from those in 2002. For blacks, once cohort characteristics are

controlled, NFRs are significantly lower in 1972 and 1977 than in 2002, but significantly

higher from 1987 through 1997. In other words, with explicit, and theoretically and

substantively relevant, cohort characteristics included, there is not a monotonic increase

in NFRs over the years in our analysis. This curvilinear pattern is especially strong for

blacks.

What if White Cohorts Were Black, and Vice Versa?

Table 6 reports the values of hypothetical NFRs calculated using the equations in

Model 3 in Table 5 for each race group, but substituting the average values of the cohort

characteristics for the other group.[8] These calculations assume that the magnitude of age,

period, and cohort influences on NFRs (and the intercept) remains the same for whites

and for blacks, and that only the value of the cohort characteristics is altered.

[Table 6 About Here]

This theoretical manipulation results in dramatic changes in the values of the

dependent variable. For whites, if the pattern of age, period, and cohort influences stayed

the same but the magnitude of the cohort characteristics matched that of blacks, their age-

period-specific nonmarital fertility ratios would rise substantially, more than doubling in

size in all instances. For blacks, the opposite occurs. If their pattern of age, period and

cohort influences stayed the same and their cohort characteristics matched those of

whites, their age-period specific NFRs would decline markedly, to far less than half of

the actual values. Clearly, differences in the cohort factors and their effects play a

persistent role in the racial differences in nonmarital fertility ratios. Inspection of the

calculations indicates that the largest changes in this theoretical manipulation result from the different levels of the measure of family structure.

Discussion and Conclusion

This paper examines age, period, and cohort variations in the nonmarital fertility ratios of black and white women in the United States, focusing on the adult childbearing years (i.e. women aged 20-44). Our objective is to explore the role of cohort opportunity structures in discouraging (or encouraging) nonmarital births throughout life (a cohort effect) and the ways in which these effects vary from one race group to another (a contextual effect). We focus on three variables theoretically related to cohorts' opportunity structures: childhood family structure, investments in education in the late teens and young adulthood, and the ratio of males to females in young adulthood. A critical distinguishing feature of this paper is the recent methodological innovations that we exploit to appropriately disentangle the hypothesized cohort effects from related age and period effects. Our results support our hypotheses regarding both the influence of cohorts' opportunity structures and contextual effects associated with race, with effects that differ sharply by race for two of the three cohort related variables.

Childhood family structure, measured as the proportion of the cohort born to unwed mothers, has the expected positive association for both groups, although the influence is slightly stronger for whites than for blacks. We suggest that all children in cohorts with more nonmarital births would be expected to experience less adult supervision and monitoring and a greater influence of peers, and our results indicate that, for both whites and blacks, this experience is associated with a cohort's probability of having more nonmarital births throughout adulthood. To the extent that this measure

captures the effects of supervision, monitoring, and socialization processes, the effects appear to operate in relatively similar manners across racial contexts.

Our second measure of cohort opportunity structures, education, exhibits strong contextual influences. The effect is negative and statistically significant for whites and positive and statistically significant for blacks. Thus, for white cohorts, there is some indication of support for our hypothesis that education in young adulthood represents an investment in human capital that results in delaying childbearing and possibly marriage, as well. It may also be related to whites' greater probability of marriage over the life span and after giving birth without being married (Bennett, et al, 1989; Harknett and McLanahan 2004). For black cohorts, the dynamics appear different, with greater investment in education perhaps resulting in a greater probability to be financially independent and thus making childbearing outside of marriage more economically feasible. Further research is, of course, needed to examine why the discourage/delay influence appears stronger than the independence effect for white women, while the reverse is true for black women.[9]

The third measure of cohort opportunity structure – mate availability, captured by the sex ratio – also produced sharply different results for blacks and whites. Consistent with Wilson (1987) and Willis (1999), results are significant in the predicted negative direction only for blacks, where the ratio is below unity for all cohorts.[10] For whites, for all cohorts, there was an excess of men when the cohorts were in their early 20s. These results reflect both the low ratio of marriageable men for blacks, which arises in large part from higher incarceration rates and higher death rates among black men (O'Brien and Stockard 2002, Pettit and Western 2004). Additionally, we note that black women

are far less likely than white women to marry interracially (Tucker and Mitchell-Kernan 1990), which compounds this problem.

As shown in Table 6, the differences in NFRs between blacks and whites are large for all age groups and periods. The average age-period-specific rate for blacks is over three times as large as that for whites. Our results suggest that these differences may result from two different pathways, both involving the different contexts in which the cohorts have been raised. First, black cohorts have higher levels of risk factors included in our models: women more often grow up in single parent families and the pool of marriageable men is smaller. The calculations summarized in Table 5 illustrate the impact of these differences. If white cohorts had the characteristics of black cohorts, but the same underlying pattern of age, period, and cohort effects, their nonmarital fertility ratios (NFRs) would be much higher; and if black cohorts had the characteristics of whites, their nonmarital fertility ratios (NFRs) measures would be substantially lower.

Second, cohort effects have different levels of unique influence and cohort opportunity structures are related in different ways to NFRs for the two groups. The different relationship of greater investment in education is striking. For whites, cohort opportunity structures related to higher levels of education are related to lower nonmarital fertility ratios, while for blacks such structures actually appear to promote higher ratios, net of other variables.[11] Similarly, higher sex ratios (more males relative to females) have a significant negative influence on the nonmarital fertility ratios of black cohorts, but no significant relationship to the NFRs of white cohorts.

The differing influences of age, period, and cohort bolster the conclusion regarding the important contextual effects of race. Using the Age-Period-Cohort Mixed

Model approach, we determined that cohort factors have a much more important unique influence for whites than for blacks. For blacks, cohort factors explain a very small proportion of the total variation in NFRs and the influences of age, period and cohort are much more confounded.

Future research in this area is of course important, for our results provide only a very macro, broad ranging picture of the last third of the twentieth century. Three areas of potentially important research can be noted. First, even though the patterns of race-ethnic segregation in the United States are strongest for African-Americans, it would be important to examine, as further data become available, the extent to which cohort effects vary among other race-ethnic groups. Second, segregation occurs not just by race-ethnicity, but also by social class (see Jargowsky, 1997); and there is extensive class variation within race-ethnic groups. An important line of research could explore the joint impact of contexts of class and race. Third, our analysis is of course macro in nature. Future research could potentially examine the linkages between the macro processes that we have examined here and the interactions and relationships on the micro level that underlie these data as well as the parallels between micro level processes and the macro level processes that we have hypothesized underlie our findings.

Overall, our findings indicate that cohort opportunity structures have powerful effects protracted across the childbearing lifespan for women, even with explicit age and period effects held constant, but that the pattern of these effects varies across the context of race. Indeed, if black women and white women had cohort characteristics typical of the other group, age-specific NFRs for black women would decline markedly, while those for whites would increase markedly. Hence, cohort related variables in conjunction

with the segregated contexts in which cohorts grow to maturity appear to contribute

substantially to black-white differences in NFRs in adulthood in the United States.

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Table 1: Cohorts Used in the Analysis and Birth Years

Cohort #	Age 20-24 in	Age 40-44 in	Birth Year
1	1952	1972	1928-1932
2	1957	1977	1933-1937
3	1962	1982	1938-1942
4	1967	1987	1943-1947
5	1972	1992	1948-1952
6	1977	1997	1953-1957
7	1982	2002	1958-1962
8	1987		1963-1967
9	1992		1968-1972
10	1997		1973-1977
11	2002		1978-1982

Table 2: Age-Period-Specific Nonmarried Fertility Ratios (NFR) by Age, Period, and Race, 1972-2002

Age		Period						
		1972	1977	1982	1987	1992	1997	2002
20-24	NFR-Whites	5.65	8.26	13.42	21.58	31.71	38.41	44.6
	NFR-Blacks	37.28	49.93	59.88	68.52	75.24	79.76	81.3
	Cohort	5	6	7	8	9	10	11
25-29	NFR-Whites	2.71	3.57	6.2	9.72	14.27	16.9	20.7
	NFR-Blacks	24.73	30.57	40.01	48.64	54.98	56.8	58.5
	Cohort	4	5	6	7	8	9	10
30-34	NFR-Whites	3.23	3.56	5.13	7.42	10.15	10.54	11.8
	NFR-Blacks	23.08	25.97	32.23	40.25	46.7	44.11	42.5
	Cohort	3	4	5	6	7	8	9
35-39	NFR-Whites	3.99	5.11	6.86	8.91	11.4	11.23	11.8
	NFR-Blacks	23.19	27.29	30.82	37.28	44.7	42.65	39.2
	Cohort	2	3	4	5	6	7	8
40-44	NFR-Whites	4.96	7.79	10.75	12.86	15.12	14.19	15.3
	NFR-Blacks	24.26	28.38	33.23	37.98	43.34	44.75	41.6
	Cohort	1	2	3	4	5	6	7

Table 3: Cohort Characteristics by Race

Cohort Number	Childhood Nonmarital birth ratio (birth years)		Sex Ratio, 20-24		School enrollment, 18-21	
	White	Black	White	Black	White	Black
1	1.9	14.7	100.9	88.6	16.6	14.3
2	2.1	16.1	101.0	90.6	20.4	17.5
3	1.9	17.1	101.3	91.3	21.6	21.3
4	2.0	16.9	102.1	92.4	29.6	20.3
5	1.7	17.6	102.4	93.0	33.0	28.9
6	1.8	20.1	102.8	94.0	35.5	35.3
7	2.4	21.9	101.9	94.4	37.7	36.0
8	3.9	26.3	101.3	93.4	44.0	32.5
9	5.6	35.2	105.0	96.8	49.0	38.0
10	7.2	44.0	104.5	97.4	52.6	47.2
11	10.6	48.4	107.0	97.5	56.2	51.4

Table 4: Descriptive Statistics, All Variables, by Race

	Whites	Blacks
Age-Period-Specific Nonmarital Births (Nonmarried Fertility Ratio)		
Mean	12.3	43.4
Minimum	2.7	23.1
Maximum	44.6	81.3
Standard Deviation	9.5	15.6
Cohort Nonmarital Birth Ratio		
Mean	3.1	23.2
Minimum	1.7	14.7
Maximum	10.6	48.4
Standard Deviation	2.1	8.8
Cohort Sex-Ratio		
Mean	102.5	93.6
Minimum	100.9	88.6
Maximum	107.0	97.5
Standard Deviation	1.4	2.1
Cohort School Enrollment (18-21)		
Mean	35.9	31.0
Minimum	16.6	14.3
Maximum	56.2	51.4
Standard Deviation	10.0	8.9
Number of cases	35	35

Table 5: Results for Blacks and Whites

	Models for Whites						Models for Blacks					
Fixed Effects	Model 1		Model 2		Mode 3		Model 1		Model 2		Model 3	
	Coeff	t	Coeff	t	Coeff	t	Coeff	t	Coeff	t	Coeff	t
Intercept	2.9402	41.08***	-7.5354	-0.95	-7.3375	-0.97	3.7637	177.55***	13.0068	3.31*	12.1937	3.31*
2024	0.5449	6.47***	0.6244	3.44**	0.5972	3.46**	0.5678	25.07***	0.3198	6.38***	0.3255	7.09***
2529	-0.1821	-2.78*	-0.1188	-0.86	-0.1502	-1.13	0.2087	10.44***	0.0253	0.67	0.0269	0.77
3034	-0.4060	-8.58***	-0.3659	-3.88***	-0.3896	-4.40***	0.0177	1.01	-0.1050	-3.97**	-0.1064	-4.48***
3539	-0.2687	-8.49***	-0.2407	-4.60***	-0.2348	-4.44***	-0.0230	-1.46	-0.0839	-4.69***	-0.0819	-4.19***
4144	0.0000		0.0000		0.0000		0.0000		0.0000		0.0000	
1972	-1.5326	-12.31***	-1.6502	-6.10***	-1.6002	-6.19***	-0.6539	-20.94***	-0.2871	-3.89**	-0.2919	-4.31***
1977	-1.1524	-10.92***	-1.2357	-5.49***	-1.1822	-5.51***	-0.4537	-15.88***	-0.1464	-2.35*	-0.1501	-2.60*
1982	-0.6521	-7.54***	-0.7214	-3.96***	-0.6845	-3.94**	-0.2546	-9.96***	-0.0057	-0.11	-0.0023	-0.05
1987	-0.2479	-3.66**	-0.3003	-2.19*	-0.2546	-1.95[a]	-0.0705	-3.09**	0.1204	2.80*	0.1290	3.33**
1992	0.0248	0.49	-0.0130	-0.14	0.0142	0.16	0.0559	2.73*	0.1833	5.80*	0.1912	6.90**
1997	-0.0053	-0.15	-0.0190	-0.35	-0.0091	-0.17	0.0392	2.09[a]	0.1044	4.90***	0.1058	4.61***
2002	0.0000		0.0000	.	0.0000	.	0.0000		0.0000		0.0000	
Cohort Specific Effects												
Log NFR			0.4739	6.32***	0.5017	7.05***			0.4077	6.49***	0.4165	7.26***
Log Sex Ratio			2.8653	1.65	2.7819	1.66			-2.5499	-2.82*	-2.3547	-2.77*
Log School Enrollment			-0.9409	-3.06**	-0.8995	-3.06*			0.2933	4.19**	0.2629	4.04**
Random Effects												
Cohort	0.0448	2.3*	0.0018	1.64[a]	0.0020	1.97*	0.0020	2.03*	0.0001	0.87	0.0002	1.59[a]
Residual	0.0019	3.46***	0.0020	3.49***	0.0020	3.19***	0.0008	3.47***	0.0007	3.54***	0.0007	3.26***
AR(1)					-0.5785	-3.23**					-0.5656	-3.14**
Model Fit Statistics												
- 2 x LL	-73.4		-103.3		-105.5		-128.4		-149.7		-152.4	
BIC	-42.2		-64.9		-64.7		-97.3		-111.3		-111.7	

[a] significant at .10 level; * significant at .05 leve; ** significant at .01 level; *** significant at .001 level. All tests are two-tailed.

Table 6: Predicted Nonmarital Fertility Ratios by Age and Period of Whites and Blacks If They Were to Have the Cohort Characteristics of the Other Group

| | Whites | | Blacks | |
	Actual Values	Predicted with Black Cohort Characteristics	Actual Values	Predicted with White Cohort Characteristics
Total	12.3	31.2	43.4	14.6
Age				
20-24	23.4	55.1	64.6	19.3
25-29	10.6	26.1	44.9	14.3
30-34	7.4	20.5	36.4	12.5
35-39	8.5	24.0	35.0	12.8
40-44	11.6	30.3	36.2	13.9
Period				
1972	4.1	9.2	26.5	10.8
1977	5.7	14.0	32.4	12.4
1982	8.5	23.0	39.2	14.4
1987	12.1	35.3	46.5	16.4
1992	16.5	46.2	53.0	17.5
1997	18.3	45.1	53.6	16.1
2002	20.8	45.5	52.6	14.4

Note: The values in this table represent the NFRs that each race group would have if they had the cohort characteristics of the other race group (as reported in Table 3), but the influence of cohort characteristics on NFRs remained (as reported in Table 5). Values were calculated by substituting mean values of logged cohort characteristics in Table 3 in the equations for Model 2 for each race group, as given in Table 4. Predicted logged values for each age group and year were calculated and average scores by age and period were then calculated by summing across the cases in each age or period and averaging across the 5 age groups or 7 periods. Antilogs were then taken to convert the values to the NFRs shown above.

Endnotes

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[1] The correlation between the percentage of cohort members in single-parent families between ages 5 and 9 and the percentage of births that were nonmarital at the time of the cohort's birth is .98, and the correlation between first differences of these two measures is .90 (O'Brien and Stockard 2002; footnote 5.).

[2] The youngest age group in our analysis is 20-24 years of age, and our measure of school enrollment is based on ages 18-21. We chose this measure because it has much more variability than a measure based on earlier ages, even though it taps investment in human capital after the point at which some women may have already experienced nonmarital child bearing. It is then possible that nonmarital fertility in the teen years may have influenced decisions regarding educational investment. Our strong controls for age help to control for this dynamic.

[3] Analyses that include teens are available on request from the authors. They do not negate the general conclusions of this paper.

[4] Since there are many options that a researcher might use we provide the SAS program that we used to estimate our most complicated model: Model 3:

```
proc mixed method=ml noclprint scoring = 15 covtest;
   class age year cohort;
   model lnnfr = age year lnnmbb lnsexratio lnschool /solution ddfm=bw;
   random cohort / solution;
   repeated/ type=ar(1) sub=cohort;
 run;
```

Where age represents the dummy variables for age-groups, year the dummy variables for periods, and cohorts the dummy variables for cohorts (they are transformed in to dummy variables using the class statement). Lnnfr is the logged rate NFRs, lnnmb is the logged percent of nonmarital births for the birth cohorts, lnsexratio is the logged sex ratio for the cohorts; and lnschool is the log of the rate of school enrollment.

[5] The variance associated with age, period, and cohort "dummy variables" is essentially the same no matter which of two of these factors are fixed and which is treated as random. For example, in the six estimates we make of the variance associated with age, period, and cohort in our analyses (age as random, period as random, cohort as random for both blacks and whites) the difference in the estimates of the proportion of variance associated with age, period and cohort is never more than .16%. The details of this procedure appear in O'Brien et al. (2008).

[6] Results with unlogged variables are substantively similar to those reported here and are available upon request.

[7] We note, however, that the AR(1) term does not improve the fit of the model according to BIC (which takes into consideration the number of variables and the number of cases in the analysis as well as the log likelihood for the model) for whites but does for blacks. Neither for whites or blacks does the addition of the AR(1) term significantly improve the fit of the model according to the likelihood ratio chi-square test [chi-square = 2.2 (= −103.3 − (−105.5)) with one degree of freedom (p > .10) for whites and chi-square = 2.7 (=−149.7 − (−152.4)) with one degree of freedom (p = .10) for blacks].

[8] Means for the logged cohort characteristics were as follows: ln cohort nonmarital births: black mean = 3.0866, white mean = .966899; ln sex ratio: black mean = 4.539, white mean = 4.6297; ln school enrollment, black mean = 3.389, white mean = 3.5379.

[9] This result is at odds, at least implicitly, with evidence in Bennett *et al.* (1989) that increased education is associated with *higher* rates of eventual marriage among black women.

[10] Incarceration rates and other age- and time-varying factors are captured by the age and period effects. In addition, as discussed above, our measure is based on the "resident" population, excluding institutionalized people. Hence, we use the sex ratio as a measure of mate availability intrinsic to a cohort.

[11] The differential impact of education on NFRs affect, of course, the projections in Table 5. Whites have, on average, higher levels of school enrollment in young adulthood than blacks (see Table 3). While for whites higher levels of enrollment result in lower levels of NFR, for blacks higher levels, which result from the substitutions used in the calculations for Table 5, result in higher levels of NFR.