
THE RISING SHARE OF NONMARITAL BIRTHS: FERTILITY CHOICE OR MARRIAGE BEHAVIOR?*

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Much of the sharp rise in the share of nonmarital births in the United States has been attributed to changes in the fertility choices of unmarried and married women—in response, it is often argued, to public policy. In contrast, we develop and test a model that attributes the rise to changes in marriage behavior, with no necessary changes in fertility. A variety of empirical tests strongly support this conclusion and invites focused attention to issues related to marriage behavior as well as to the interactions between marriage and fertility.

Over the past half-century, total birth rates in the United States have fluctuated widely, rising dramatically after World War II and then falling from the late 1950s to the mid-1970s, when they dropped below prewar rates. Since the mid-1970s, total birth rates have changed little, increasing only modestly for whites and barely at all for blacks. At the same time, however, birth rates for unmarried women have soared. Birth rates for married women have also increased substantially, but proportionately less than for unmarried women. Simultaneously, the share of unmarried women has risen sharply. Collectively, these trends yield a particularly striking increase in the ratio of unmarried births to total births, sometimes termed the illegitimacy ratio, but referred to here as the nonmarital fertility ratio (NFR).

In this article, we offer an explanation for the apparent paradox posed by several decades of rising marital and nonmarital birth rates on the one hand, and relatively constant total birth rates on the other. Our explanation relies on the role played by the share of unmarried women in determining marital and nonmarital birth rates and simultaneously explains the sharp increase in NFR over the same period.

Increases in NFR have for some time been a central focus of a vast literature, particularly with respect to the effects of public policy, such as the now displaced federal Aid to Families with Dependent Children (AFDC), the newer Temporary Assistance for Needy Families (TANF), and the Earned Income Tax Credit. In addition, the 1996 Federal Welfare Reform Act requires states to reduce out-of-wedlock pregnancies and to establish annual numerical goals for doing so. However, changes in NFR do not necessarily indicate changes in the underlying childbearing decisions that are often the real target of public policy.

NFR can be decomposed into three component factors: the nonmarital birth rate, the marital birth rate, and the fraction of women who are married. NFR will vary positively with the first factor and negatively with the two other factors. Some demographic studies of NFR (e.g., Smith, Morgan, and Koropecjy-Cox 1996) focused on measuring the contributions of these component factors and on exploring their separate determinants. Finding an effect of a policy change on marital or nonmarital birth rates, and therefore NFR, is taken in some studies as evidence that the policy change has affected childbearing behavior (e.g., Baughman and Dickert-Conlin 2003; Duncan and Hoffman 1990).

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The distinctive feature of the theory we develop is that marriage behavior is *causally* related to marital and nonmarital birth rates. A decline in the proportion of women who choose to marry increases both the marital birth rate *and* the nonmarital birth rate, with the nonmarital birth rate rising *relative* to the marital birth rate. Consequently, marriage behavior has equally strong direct and, via relative birth rates, indirect effects on NFR. Our empirical findings support the theory and suggest that the effects of marriage behavior on birth rates are quantitatively important. If so, the dramatic changes in NFR of the past several decades may be due less to changes in fertility behavior than to changes in marriage behavior.

Thus, we address one of the central questions posed by Smith et al. (1996:142): “Are the continued increases in the proportion of children in the population who are born out of wedlock now a function primarily of fertility changes among unmarried women?” Or, are they a function primarily of changes in marriage behavior, as argued even earlier by Smith and Cutright (1988)? Smith and Cutright foreshadowed this article’s central insight when they speculated that declines in marriage rates put upward pressure on nonmarital birth rates by adding to the unmarried population “...an aggregation of women who are differentially selected with respect to a crucial criterion for out-of-wedlock births...” (p. 244). While Smith and Cutright emphasized sexual activity as the selection criterion, desired family size serves that role in this article. We posit an exogenous distribution of desired fertility and a propensity to marry that is positively related to that desire. In the context of this framework, a decline in marriage rates changes the composition of the existing married and unmarried populations with respect to desired fertility, increasing the mean birth rate for both groups.

BACKGROUND AND OVERVIEW

Although we examine longer periods as well, this article is motivated by the trends in birth rates, marriage rates, and NFR evident for women in the prime childbearing years 20–39 over the period 1974–2000, a period of relatively flat total birth rates. The average (age-standardized) nonmarital birth rate for white women aged 20–39 more than tripled over the period, rising from 13 per 1,000 in 1974 to 46 per 1,000 in 2000.¹ The marital birth rate rose by nearly half, from 94 to 135. But the total birth rate for white women increased by less than a fifth, from 78 to 92. The patterns are similar for black women, though they begin and end at much higher rates than for white women. The nonmarital birth rate for black women increased from 55 to 74, roughly a third. The marital birth rate for black women also increased, from 108 to 132, a smaller increase of roughly a fifth. But the total birth rate for black women barely rose at all, from 86 to 87. For both blacks and whites, there was also a pronounced shift away from marriage between 1974 and 2000, with the share of unmarried white women rising from .25 to .45 and the share of unmarried black women rising from .52 to .73. Finally, as noted earlier, the trends in birth and marriage rates yielded even more dramatic increases in NFR, which rose from .03 to .21 for white women, a sevenfold increase, and from .30 to .57, almost double, for black women.

The fact that both marital and nonmarital birth rates have risen substantially over the past quarter-century, despite little or no change in total birth rates, presents a particular challenge. Our explanation is simple. Increases in the population share of unmarried women are produced by changes in the marital status of women who have a low probability of giving birth when compared with the average married woman, but a high probability of giving birth when compared with the average single woman. Accordingly, when the proportion of women who are married declines and the share of unmarried women correspondingly increases,

1. Throughout the article, summary data for women aged 20–39 refer to weighted averages constructed from underlying data available by five-year age interval, where the weights are the average population shares of the women in each age group over the relevant sample period. See footnote 6 for a description of the data and sources.

the average birth rates of *both* groups rise.² Thus, changes in marital and nonmarital birth rates may arise, not from changes in childbearing behavior, but from changes in marriage behavior. The model further predicts that a rise in the share of unmarried women will increase the nonmarital birth rate proportionately more than the marital birth rate. That is, the ratio of the two birth rates will rise—a prediction that is consistent with observed trends in the data and the basis for our claim that changes in marriage behavior produce a *magnified* effect on NFR.

Our findings underscore the importance of studies that look simultaneously at fertility and marriage (e.g., Grogger and Bronars 2001; Upchurch, Lillard, and Panis 2002), as well as those that focus directly on the determinants of marriage and other forms of union formation (e.g., Bitler et al. 2004; Carlson, McLanahan, and England 2004; Fitzgerald and Ribar 2004; Moffitt 2000; Sigle-Rushton and McLanahan 2002). They also suggest caution in interpreting the results of studies that examine marital and nonmarital birth rates separately—and do not consider total birth rates—in order to assess the effects of public policy on fertility choices (e.g., Baughman and Dickert-Conlin 2003). The effects identified may be due to the influence of policy on marriage decisions, which then induce changes in measured birth rates, with no significant change in fertility choices or total fertility. Finally, our results suggest that at least some of the divergence in published results concerning the effects of public policy on fertility may be reconciled by accounting for differences in the treatment of marital status.

THE IDEA

This section describes a simple but joint theory of childbearing and marriage. The model is intentionally stylized because its purpose is to isolate and illustrate a particular effect of changes in marriage rates on nonmarital and marital birth rates and, hence, NFR. The effect we isolate occurs even in the absence of changes in individual childbearing behavior and can therefore reconcile the observation of simultaneous *increases* in nonmarital and marital birth rates with a *constant* total birth rate. It also predicts a disproportionately large effect of changes in marriage behavior on NFR. Indeed, we show in the next section that, when one accounts for the causal relationship between marriage behavior and birth rates, a remarkably high proportion of actual changes in NFR can be attributed to changes in marriage behavior alone.

Theoretical Framework

Our central theoretical findings follow directly from a few definitions and a small number of deliberately strong assumptions. The nonmarital fertility ratio, NFR, is defined as

$$NFR = UB / (MB + UB),$$

where *MB* is the number of births to married women, and *UB* is the number of births to unmarried women. Simple algebra allows us to express NFR in terms of the nonmarital birth rate, the total birth rate, and the fraction of women who are not married (hereafter termed the *unmarried share*):

$$NFR = Su \cdot (UBR / TBR), \quad (1)$$

where *U* is the number of unmarried women; *UBR*, the birth rate of unmarried women (or the nonmarital birth rate), is equal to UB / U ; *TBR*, the total birth rate, is equal to $(MB + UB) /$

2. The birth rate behavior described here is an example of Simpson's Paradox (Simpson 1951). As one referee pointed out, a better known example is the Yale student who transfers to Harvard, thereby (as told at Yale) raising the mean intelligence at both schools.

$(M + U)$; and Su , the unmarried share, is equal to $U / (M + U)$. As expressed in Eq. (1), NFR depends only on the *ratio* of the nonmarital to the total birth rate and on the unmarried share. It follows that NFR differs from Su only to the extent that the childbearing behavior of unmarried women deviates from that of the rest of the population. This observation is a common basis for demographic decompositions of NFR.

Variation across women in fertility is captured by a parameter, γ , which measures the probability that a particular woman will give birth to a child during the observation period (e.g., a year). We assume that γ is exogenous with respect to the model (in particular, it is independent of marital status) and is drawn from a continuous uniform distribution defined on the interval $[0, P]$, where $0 \leq P \leq 1$.³ Furthermore, the net benefits to marriage, C , are increasing in γ . Other factors that affect the net benefits of marriage, which are later allowed to vary by race and over time in our empirical work, are assumed for now to be the same for all women in the population under consideration.⁴ Thus, the women who choose to marry are those for whom the net benefits of marriage are positive—those with relatively high γ s. Conversely, women with relatively low γ s and negative net benefits do not marry. For a large population of women, these assumptions produce the following expressions for nonmarital, marital, and total birth rates:⁵

$$UBR = (1/2)(Su)P, \quad \text{where } Su = [U / (M + U)]. \quad (2)$$

$$MBR = (1/2)(Su + 1)P, \quad (3)$$

$$TBR = (1/2)P. \quad (4)$$

It follows that the ratio of the nonmarital birth rate to the total birth rate is equal to Su , so that the *ratio* of the two depends *solely* on marriage behavior:

$$(UBR / TBR) = Su. \quad (5)$$

Substituting Eq. (5) into Eq. (1) yields the theoretical result that is the cornerstone of our empirical contribution—NFR is simply the square of Su :

$$NFR = Su^2. \quad (6)$$

Finally, Eq. (6) implies that the change in NFR induced by a change in Su depends on the level of Su .

$$dNFR / dSu = 2Su. \quad (7)$$

3. We are far from alone in assuming that family formation and marriage are driven by innate unmeasured propensities that vary across women (e.g., Upchurch et al. 2002:313). The assumption that childbearing behavior is determined by a single characteristic drawn from a uniform distribution is, however, exceptionally strong. In this regard, our work more closely parallels Udry's (1994, 2000) model of within-sex differences. A more flexible distributional assumption (e.g., the gamma), while perhaps more realistic, produces birth rate ratios that are functions of the unobservable factors represented by P and C in our model. That these factors are eliminated from birth rate ratios in the case of the uniform distribution produces our sharp empirical implications.

4. Factors widely studied by demographers, sociologists, and economists include educational levels, earnings, unemployment rates, and search costs. These factors are the underlying source of the exogenous variation in Su , by race and across time, that gives empirical content to our model.

5. If women are indexed and ordered by γ , then the γ associated with the n th woman in an arbitrarily large population of z women is $(n / z)P$; the expected birth rate of the first n ordered women in the population is $1/2(n / z)P$; the expected birth rate of the remainder of the population is $1/2[(n / z) + 1]P$; and the expected birth rate of the population as a whole (the total birth rate) is $1/2P$. Our assumptions imply that the first U ordered women in the population are unmarried, and the remaining M women are married. Thus, we set $n = U$ and $z = U + M$ to produce Eqs. (2) through (4).

The relationships described in Eqs. (2) through (7) produce the our key predictions:

Prediction 1 (P1). Increases in Su cause increases in both UBR and MBR —with a proportionately larger increase in UBR —and no necessary change in TBR .

Prediction 2 (P2). Increases in Su produce a magnified increase in NFR ($NFR = Su^2$) because the direct effect of a shift away from marriage is magnified by its effect on the ratio of UBR to TBR .

Prediction 3 (P3). $dNFR / dSu$ is linear in Su , with changes in NFR exceeding changes in Su when Su is large (greater than 0.5) and falling short of changes in Su when Su is small (less than 0.5).

Prediction 4 (P4). NFR , unlike its component birth rates, does not depend on P , the upper limit of the uniform distribution describing γ . Thus, NFR is independent of factors that might plausibly be expected to influence childbearing behavior.

Discussion

We describe a simple model in which increases in the population share of unmarried women are associated with changes in the marital status of women who have a lower probability of giving birth than average married women, but a higher probability of giving birth than the unmarried women they join. Thus, the average birth rates of both groups rise with Su , as noted in (P1). Yet, because the propensity to bear children is independent of marriage behavior, the number of children born and the total birth rate do not necessarily change. Less intuitive, perhaps, is the implication that as Su rises, UBR increases by proportionately more than MBR . The assumption that γ is uniformly distributed means that as women shift from married to unmarried, the absolute increases in the birth rates of the two groups are the same. However, because UBR is always less than MBR , the same absolute increase in the two birth rates produces a larger percentage increase in UBR .

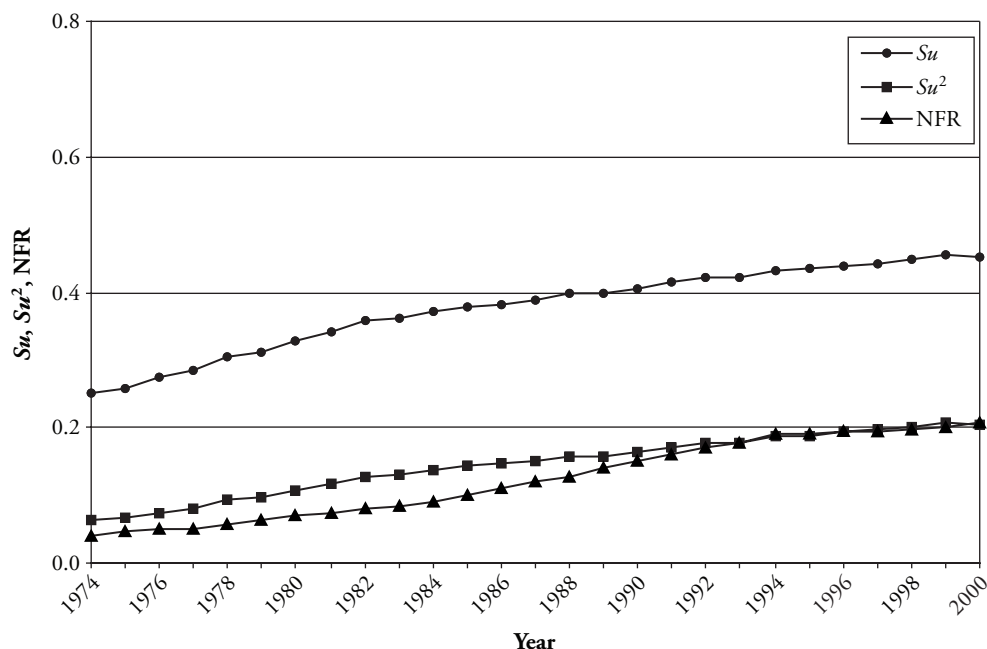
The magnified effect of marriage behavior on NFR described in (P2) arises because Su has both direct and indirect effects on NFR . The direct effect is captured in the first term on the right side of Eq. (1). For a given value of (UBR / TBR) , NFR increases proportionately with Su . However, this effect is magnified by Su 's equally strong effect on the ratio (UBR / TBR) , the second term on the right side of Eq. (1). The corollary prediction (P3) follows from Eq. (7). The size of the change in NFR caused by a change in Su depends linearly on the initial value of Su . Changes in NFR exceed (fall short of) changes in Su as Su exceeds (falls short of) 0.5.

The prediction that NFR is independent of P , the upper limit of the distribution from which γ is drawn (P4), is important because it allows us to set aside in our study of NFR many factors that might be expected to influence childbearing behavior. Since each birth rate is proportional to P , the ratio of any two birth rates is independent of P . Changes in P have the same proportional effect on childbearing for all women and so do not affect birth rate *ratios*. Thus, NFR , equal to the product of Su and the birth rate ratio (UBR / TBR) , is independent of P .

EMPIRICAL EVIDENCE

To answer the question of how well our model performs in explaining NFR , we examine several forms of empirical evidence. The annual U.S. data employed in these exercises are constructed from information available by five-year age interval for married and unmarried women over various post–World War II periods.⁶ Black women and white women are

6. Data for births by marital status are from National Vital Statistics Reports (National Center for Health Statistics 2000, 2002). Total births are from Vital Statistics of the United States (National Center for Health Statistics

Figure 1. Su^2 as a Predictor of NFR for White Women Aged 20–39

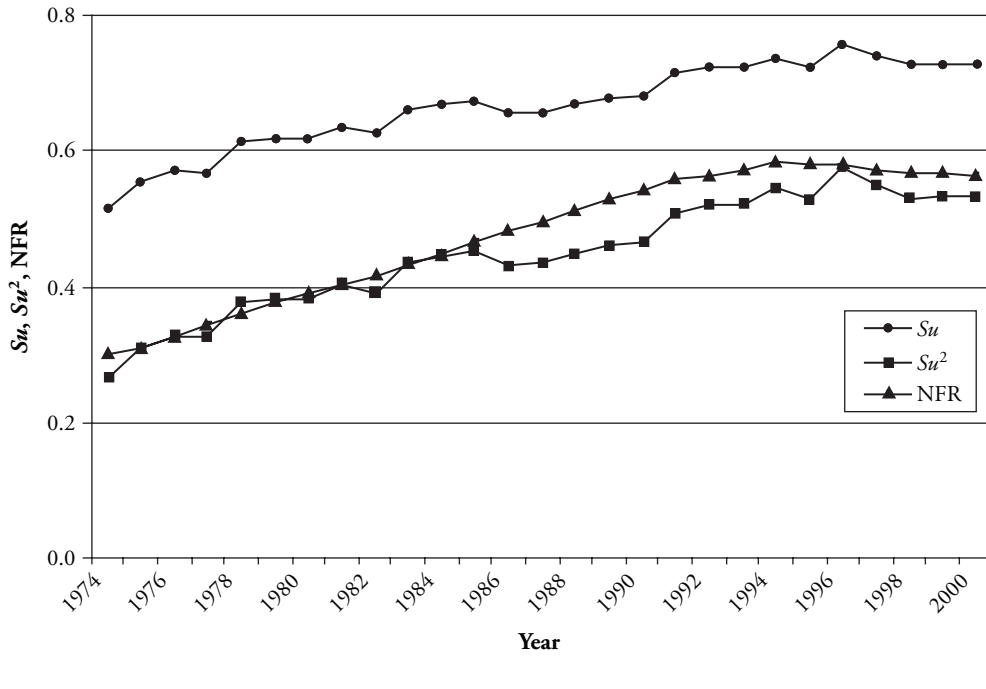
examined separately, although we find little difference in the ability of the model to explain behavior across the two races. Because biological, legal, and social factors may limit the relevance of our theoretical model to teenagers and older women, our empirical work focuses on the prime adult childbearing years of 20–39. However, we briefly explore the issue of teenage fertility and marriage behavior at the end of the section.

Data Overview

We begin with the prediction that NFR is equal to the squared value of Su . Figures 1 and 2 compare the model's prediction of NFR with actual experience over the years 1974–2000, the period emphasized in earlier sections. The figures plot the values of NFR, Su , and Su^2 separately for white women and black women aged 20–39. Overall, the match between historically observed values of NFR and the model's prediction, Su^2 , is strikingly close. In Figure 1, for whites, both the level and rate of increase of NFR over time correspond closely to the level and rate of increase of Su^2 . Although the measure of Su is noisier for black women, the same general correspondence is evident in Figure 2. Thus, the much higher levels of Su^2 for blacks than for whites produce correspondingly higher levels of NFR. The values of NFR range from a low of .04 in Figure 1 to a high of .58 in Figure 2,

n.d.). The numbers of married and unmarried women are from Current Population Reports (U.S. Census Bureau, various years). Married women are those categorized by the Census Bureau as "married, spouse present." Birth and population data are available for teenagers and for the five-year age groups 20–24, 25–29, and 30–34 beginning in 1957 for whites and 1969 for blacks. Data for women aged 35–39 begin in 1968 for whites and 1969 for blacks. Teenagers are those aged 14–19 prior to 1980 and those aged 15–19 in 1980 and subsequent years.

Figure 2. Su^2 as a Predictor of NFR for Black Women Aged 20–39



over half the total possible variation in NFR. That Su^2 appears to explain NFR well over such a wide range of values suggests a robust relationship.

To assess the sensitivity of these observations to age-related differences, we also examine data for individual five-year age groups.⁷ Table 1 presents the results of simple numerical calculations for how well changes in Su^2 explain changes in NFR between 1974 and 2000. For each age group and race, Table 1 shows that changes in Su^2 correspond quite closely to the change in NFR. The ratio of the predicted NFR to the actual NFR deviates from 1 by just 6% on average, and in only one case—for black women aged 30–34—does the ratio fall more than 7 percentage points away from unity. Overall, the calculations presented in Table 1 are consistent with the hypothesis that observed changes in NFR have been caused primarily by changes in marriage behavior rather than by changes in fertility behavior per se.⁸

While Table 1 focuses on the model prediction (P2) that NFR is equal to the square of Su , Figure 3 examines evidence for the prediction (P3), that $dNFR / dSu$ increases linearly in Su , with changes in Su exceeding changes in NFR ($dNFR / dSu < 1$) when Su is less than 0.5, and falling short of changes in NFR ($dNFR / dSu > 1$) when Su is greater than 0.5. Figure 3 plots the ratio $dNFR / dSu$ against Su for blacks and whites in each of the four age groups in our study. The ratio $dNFR / dSu$ is calculated by dividing the change in NFR

7. This exercise is similar in spirit, though not in detail, to the demographic decompositions reported in Smith et al. (1996:147, table 3) for the period 1982–1992. Whereas our calculations assign to Su both its direct effect and its indirect effects (through birth rates) on NFR, Smith et al. assign to Su only its direct effect on NFR.

8. For both Table 1 and Figure 3, results over longer periods for which relevant data are available are similar and are available on request.

Table 1. Ratio of Changes in Su^2 to Changes in NFR From 1974 to 2000^a

Age Group	White Women			Black Women		
	NFR Change (1)	Su^2 Change (2)	Ratio of (2) to (1) (3)	NFR Change (4)	Su^2 Change (5)	Ratio of (5) to (4) (6)
20–24	0.357	0.348	0.976	0.401	0.421	1.050
25–29	0.158	0.152	0.965	0.301	0.284	0.945
30–34	0.077	0.072	0.931	0.180	0.215	1.192
35–39	0.066	0.062	0.937	0.145	0.148	1.016

^aThe model predicts that the ratios in columns 3 and 6 will be 1.000.

between 1974 and 2000 by the change in Su over the same period. For the level of Su , which appears on the horizontal axis, we use the average of its values in 1974 and 2000. The reference line, which plots Eq. (7), has an intercept of 0 and a slope of 2.

As Figure 3 shows, the eight observations provided by the two races and four age groups in our sample deviate very little from the linear relationship predicted by the theory, with the observation for black women aged 30–34 lying furthest from the reference line included in the figure. Furthermore, with only one exception (again, blacks aged 30–34), changes in Su exceed (fall short of) changes in NFR when Su is less than (greater than) 0.5. In the three cases in which Su is less than 0.5, the corresponding values of $dNFR / dSu$

Figure 3. $dNFR / dSu$ and Su , 1974–2000

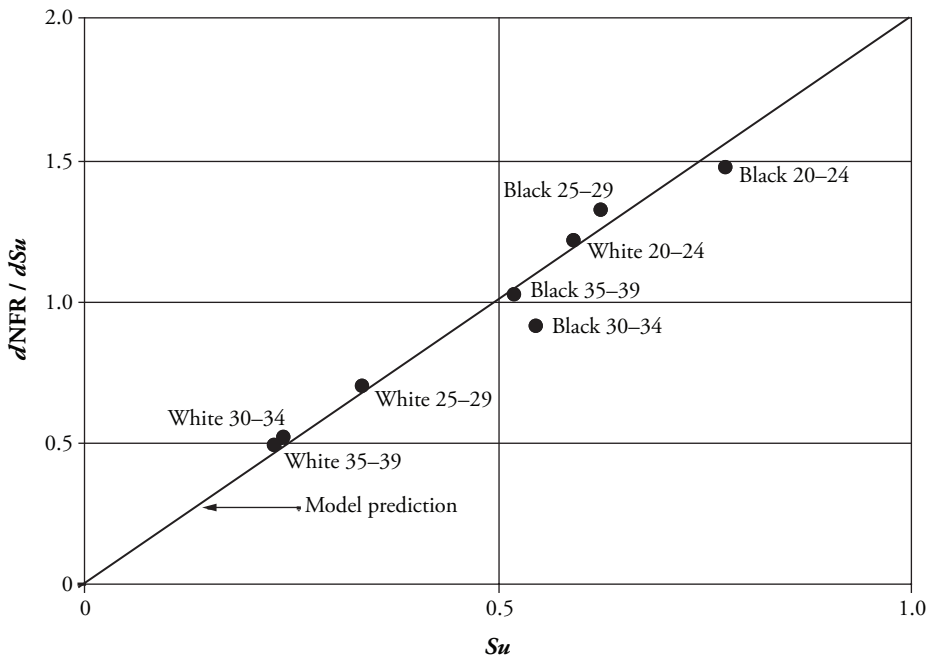


Table 2. Regression Estimates of the Link Between NFR and Su^2

Specification	(1)	(2)	(3)
Constant	-0.005* (0.002)	0.013* (0.003)	0.006 (0.009)
Su^2	0.985* (0.008)	0.973* (0.014)	0.965* (0.022)
[5% confidence interval]	[0.968–1.001]	[0.945–1.000]	[0.923–1.008]
AR(1) _i ^a	yes	yes	yes
Fixed Effects ^b			
Age group	no	yes	yes
Race	no	yes	yes
Race × Age group	no	no	yes
R^2 (transformed data)	0.984	0.986	0.987
Adjusted R^2 (transformed data)	0.980	0.986	0.986
<i>df</i>	283	279	276

Notes: Standard errors are in parentheses. Data are pooled across five-year groups over the years 1957–2000 for white women aged 20–34, 1968–2000 for white women aged 35–39, and 1969–2000 for black women aged 20–39.

^aEstimates are corrected for first-order (AR(1)) autocorrelation specific to age group by race.

^bFixed effects are introduced, respectively, for age group, race, and the interaction between age group and race. The latter interactions do not improve the adjusted R^2 .

* $p < .05$

are less than 1, and in all but one of the five cases in which Su is greater than 0.5, $dNFR / dSu$ is greater than 2. The one exception is black women aged 30–34, for whom Su is close to the critical value of 0.5. Overall, Figure 3 illustrates a remarkable consistency between actual experience over the period 1974–2000 and the predictions of our simple model.

Statistical Tests

The close correspondence between the (squared) share of unmarried women and the nonmarital fertility ratio apparent in both Figures 1 and 2, as well as between *changes* in the two variables, is also supported by formal statistical tests. For these tests, we turn to standard regression techniques that exploit both the cross-sectional and time-series dimensions of our data over the longer period for which data on both NFR and Su are available. Data are pooled across the available five-year groups over the years 1957–2000 for white women aged 20–34, 1968–2000 for white women aged 35–39, and 1969–2000 for black women aged 20–39. Table 2 presents estimates of the relationship between NFR and Su^2 for the full, unbalanced panel, pooled by race and age group. The estimates are corrected for first-order autocorrelation specific to race and age group. For white women, the autocorrelation coefficients range from 0.70 to 0.90, while for black women they range from 0.20 to 0.55.

The estimated effect of Su^2 on NFR reported in column 1 of Table 2 is 0.985, significantly different from zero at the 5% level, but not significantly different from the predicted value of 1.00. The fit of the equation—as measured by the adjusted R^2 for the transformed data—is quite high at 0.980, as one might expect given the time series component of the data.⁹ Column 2 reports the results of including controls (fixed effects) for the five-year age groups and for race. While the addition of these controls raises the adjusted R^2 slightly,

9. Including a time trend in the model has little effect on the results reported here. In the preferred specification in column 2, for example, the estimated coefficient on Su^2 declines only slightly, from .973 to .951.

the estimate of the coefficient on Su^2 is virtually unchanged, at 0.973. Additional controls for the age group and race interactions, reported in column 3, do not increase the adjusted R^2 , and the coefficient on Su^2 is again virtually unchanged at 0.965. Hence, the overall explanatory power of the model is remarkably strong, and its prediction that the coefficient on Su^2 be 1.00 is supported in the data, regardless of whether age group and race-specific controls are included.

Teenagers

To this point, we have concentrated solely on women aged 20–39. To what extent should our model yield accurate predictions for teenagers? One might argue that teenage pregnancies do not typically arise from a deliberate assessment of desired family size, as assumed in our model. We know from the field of neuroscience, for example, that the prefrontal cortex in teenagers is still relatively underdeveloped, limiting in varying degrees their ability to resist impulses and to make deliberate decisions about future consequences of current actions (e.g., Goldberg 2001; Kelley, Schochet, and Landry 2004).

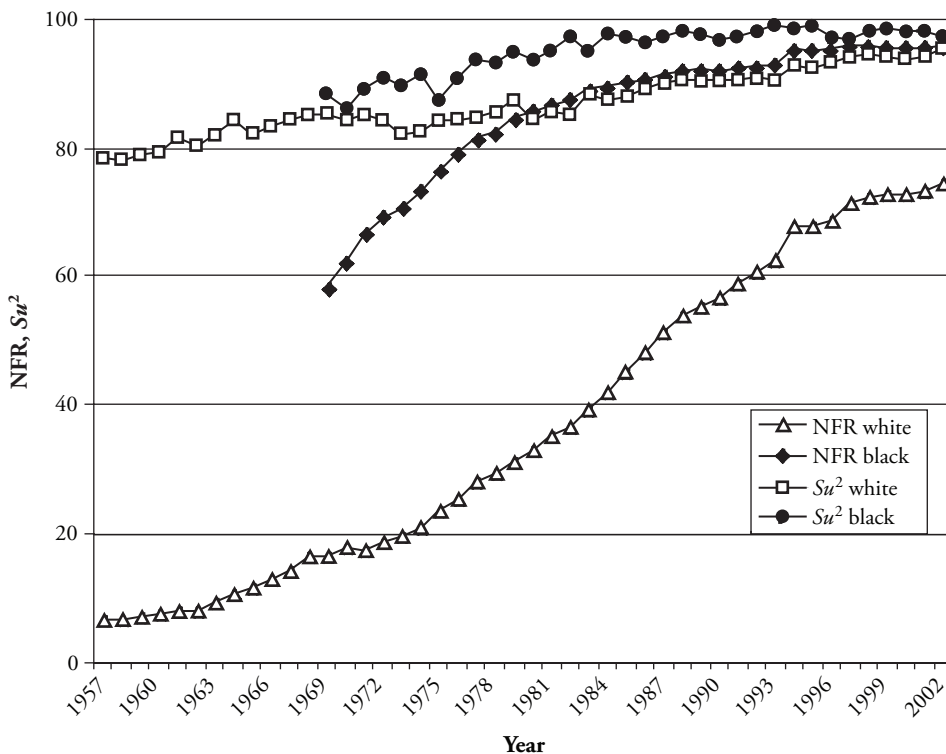
Furthermore, teenage marriage is not an unconstrained outcome. A variety of legal and cultural factors impinge on teenage decisions regarding marriage. All states restrict the age of marriage, with marriage under the age of 18 commonly permitted only with parental consent or, in some cases, pregnancy. Whether a marriage occurs, especially in the ages requiring parental consent, is likely to be heavily influenced by parents and by social conventions that may vary by time, place, and race (Delaire and Kalil 2002; Kahn and Anderson 1992). So-called shotgun weddings, for example, were common in the 1950s and 1960s, but less so today.

To the extent that these factors, rather than latent preferences for children, drive teenage pregnancy and marriage, observed patterns of teenage marriage and fertility will not correspond to those predicted by our model for adults. Indeed, Figure 4, which plots NFR and Su^2 over the period 1957–2000 for white teenagers and 1969–2000 for black teenagers, shows that NFR falls far below the model prediction early in the sample period for each race. However, the figure does appear to be broadly consistent with trends in legal and social factors over the period. If teenage pregnancies are largely unplanned events that affect a relatively small number of teens and frequently lead to (shotgun) marriage, then NFR will be small and Su^2 will be large, consistent with the early years in Figure 4. NFR is small, since most births, as opposed to conceptions, will be marital. In the extreme, if all pregnancies lead to marriage, NFR is 0. The single share and Su^2 , on the other hand, are high because teenage pregnancy and marriage are relatively uncommon. As occurrences approach 0, Su^2 approaches its upper bound of 1.

Many of the idiosyncratic legal and social factors influencing teenage births and marriage appear to have been more important earlier in the postwar period than later, and stronger for whites than for blacks. If so, then we would expect to see the gap between NFR and Su^2 for both races close over the period, with the gap larger for whites than for blacks throughout. This is the pattern we see in Figure 4. Indeed, by the end of 2000, the prediction that NFR equals Su^2 increasingly holds for whites and is satisfied for blacks.

Of course, the fact that neither NFR nor Su^2 can exceed 1 is a reason for the gap between them to close as NFR and Su both rise, but that still leaves unexplained the systematic difference between the two races. To account for legal and social factors that may vary by race and time in the context of our model, we estimate the relationship between NFR and Su^2 separately for teenagers, as reported for adult women in the second column of Table 2, but with a time trend that varies by race. The coefficient on Su^2 is 0.9149, significantly different from 0 but not from unity—consistent with our model and estimates for adults in Table 2. Nonetheless, the substantial divergence between the data and our model prediction early in the period reminds us that our strong assumptions can limit the model's applicability in important dimensions.

Figure 4. NFR and Su^2 , Teenagers



CONCLUSION

We have argued that the rising share of births to unmarried women in the United States over the past several decades may be due less to changes in the underlying fertility of individual women than to changes in their marital status. The argument is motivated by the fact that in recent decades in the United States, birth rates for unmarried women have soared, birth rates for married women have also tended to increase, and yet total birth rates have either remained flat or increased only slightly for specific age groups. If the primary origins of these trends were fundamental changes in underlying fertility, then why has the total birth rate risen so much less, proportionately, than either the nonmarital birth rate or the marital birth rate?

Our explanation relies on the effects that marriage behavior has on the composition of married and unmarried women. To illustrate the effects, we develop a model in which childbearing behavior at the level of the individual is characterized by a single preference parameter that varies across women. This simple framework allows us to isolate the effects of changes in marriage behavior on nonmarital birth rates and shares, holding constant underlying fertility decisions. The model predicts outcomes consistent with U.S. experience in recent decades: a decline in marriage will cause increases in the nonmarital birth rate, the marital birth rate, and the nonmarital birth rate *relative* to the marital birth rate, even

though the total birth rate does not change. Thus, in addition to the direct effect of marriage behavior on the nonmarital fertility ratio standard to demographic models, our model suggests an indirect effect that operates through relative birth rates. Indeed, the nonmarital fertility ratio reduces to a power function—the square—of the share of unmarried women.

U.S. data from the post–World War II period are remarkably consistent with the various predictions of the model, as are demographic projections performed for black women and white women aged 20–39. Regression estimates fail to reject the model, regardless of whether controls for specific age groups and race (or their interactions) are included in the regressions. Our findings lend support to the view that the soaring ratio of nonmarital to total births has arisen primarily from changes in marriage behavior, rather than from changes in underlying fertility choices, and provide insights into the dual relationship between fertility and marriage behaviors. Not surprisingly, the behavior of teenagers diverges from the model's predictions early in the period, for reasons we explore, but increasingly match those of the model by the end of the period.

A number of recent studies have sought to identify factors, including public policies, that affect the fertility decisions of unmarried women. As this article shows, factors that *appear* to affect fertility decisions—because they affect birth rates or shares—may actually exert no influence at all on fertility behavior. Instead, their influence may be on marriage decisions, which in turn affect measured birth rates and shares, with no necessary change in fertility choices. Thus, the article provides new arguments and evidence that changes in the nonmarital fertility ratio have been driven primarily by changes in marriage behavior over the past several decades. Of course, the model is stark, and the evidence is limited in some important dimensions. That said, we believe the article identifies a relatively unexplored and potentially powerful effect of marital choice on our most common measures of fertility behavior.

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