

Childbearing, Marriage and Human Capital Investment*

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Abstract

This paper proposes and tests a simple joint explanation for i) increases in marital and nonmarital birth rates in the United States over recent decades, ii) the dramatic rise in the share of nonmarital births, and iii) the pronounced racial differences in the timing of childbearing. The explanation arises from differences across time and race in the attractiveness of marriage and opportunities for investment in human capital. For given preferences, a decline in the marriage rate *necessarily* causes both the marital and nonmarital birth rates to increase, with no change in the total birth rate. This model exhibits exceptional power in replicating salient features of childbearing behavior. Our results suggest that changes in marital and nonmarital birth rates, as well as in the share of nonmarital births, arose primarily from changes in marriage behavior, not from changes in fertility; and that racial differences in the timing of childbearing reflect early differences in human capital investment.

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I. Introduction

The extraordinarily rapid rise in nonmarital birth rates in the United States over the past several decades – and correspondingly in the share of births to unmarried women – has elicited calls of alarm from social observers, politicians, and researchers, alike, as well as a vast literature on the potential role of various public policies in fostering changes in childbearing behavior. A simultaneous, if proportionately smaller, rise in the birth rates of married women – against a backdrop of relatively flat total birth rates – has captured less attention, but poses an apparent paradox: How can the birth rates of married and unmarried women both rise, while the total birth rate (married and unmarried women combined) be relatively constant?

Racial differences in marriage and birth rates have been an additional source of concern in both academic and policy circles, especially the dramatically lower rates of marriage and higher rates of nonmarital births among blacks than whites. Less familiar, but nonetheless prominent in the data, is the contrast between higher *total* birth rates for blacks than whites in the teens and early twenties, and *lower* total birth rates for blacks than whites beginning in the mid-twenties. Clearly, black and white women time their childbearing differently. The timing difference has become more pronounced over the past decade despite increasingly similar total fertility rates among blacks and whites. Indeed, total fertility rates for the two groups have been virtually identical for the past several years, suggesting that desired family size may not differ significantly across the two races.¹

We propose and test a simple joint explanation for i) the increases in marital and nonmarital birth rates in recent decades, ii) the dramatic rise in the share of births to unmarried women, and iii) the pronounced racial differences in the timing of childbearing – one that does *not* appeal to differences in preferences regarding either the number or timing of births. We argue that the explanation arises, in substantial part, from differences across time and race in the attractiveness of marriage and opportunities

¹ The total fertility rate is published by the U.S. Department of Health and Human Services as an indicator of completed family size.

for investment in human capital. Our model produces a *causal* relationship between marriage behavior and measured birth rates that is independent of the preferences governing childbearing behavior. For given preferences regarding family size and the timing of births, a decline in the marriage rate necessarily causes both the marital and nonmarital birth rates to increase. The increases are the result of a selection effect typically neglected in studies of the determinants of nonmarital childbearing.² The steep decline in marriage rates overall, as well as persistent racial differences in marriage rates, make this selection effect central to the model's striking empirical power.

Human capital investment opportunities are important in explaining the timing of births in the model. If whites have better access to investment opportunities than blacks during the early childbearing years, then, other things equal, the total birth rates of young black women will exceed those of young white women. However, the opposite will be true at later ages. The birth rates for white women will exceed those of black women during the mid-to-late childbearing years. Hence, differences in investment opportunities are a viable explanation for the distinctive differences in the age profiles of black and white birth rates.

The model we present is stark, with sharply simplifying assumptions. Even so, the model exhibits exceptional power in replicating salient features of childbearing behavior for women in the United States. In doing so, it offers a common theoretical explanation for some of the most widely studied trends in fertility in the past several decades. The model provides closed-form solutions for marital and nonmarital birth rates, as well as the share of births to unmarried women (often termed the illegitimacy ratio, but referred to here as the nonmarital birth share). The solutions imply stringent restrictions on the relationships between these variables and the "single share" – the proportion of women who choose not to marry. The strength of the empirical support for the restrictions, given the stark simplicity of the theoretical model, suggests that the model identifies effects of central importance in

² A notable exception is Smith and Cutright (1988). The authors speculate (p. 244) 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...." While Smith and Cutright emphasize sexual activity as the selection criterion, desired family size serves that role in the present paper.

understanding the behavior of birth rates. The contribution of the paper is in isolating and illustrating these effects, and the highly stylized theoretical model serves this end well – a case, we argue, in which less is more, and “the proof of the pudding is in the taste.”

Overall, our findings suggest that key aspects of childbearing behavior in recent decades arise more from changes in marriage behavior than from changes in fertility behavior, *per se*. In addition, the timing of childbearing, including racial differences in timing, may be explained in substantial part by opportunities women have for investments in human capital. Our findings place much of the recent literature on the effects of public policies on fertility behavior in a different context, and suggest that future efforts could be productively directed toward understanding marriage behavior and the incentives for investments in human capital, especially for black women.

We begin in Section II with the theoretical model. Section III follows with the key empirical implications of the model, along with corresponding figures summarizing actual U.S. experience over the past several decades. Section IV presents formal empirical estimates of the model, specified in terms of predictions for birth rate ratios and the share of births to unmarried women. Section V discusses the implications of our findings and directions for future research.

II. Conceptual Framework

This section describes a simple theory of childbearing and marriage in the presence of opportunities for investments in human capital. The model offers two key insights central to understanding observed patterns in childbearing, including racial differences in those patterns. First, we suggest that increases in the share of unmarried women reflect changes in the marital status of women with a lower probability of giving birth than the average married woman, but a higher probability of giving birth than the average unmarried woman. Accordingly, when such women “leave” the pool of married women and enter the pool of unmarried women, average birth rates of *both* groups rise, even though total births (and the total birth rate) may not change. Thus, the apparent paradox posed by

increases in both marital and nonmarital births rates, absent a corresponding change in total birth rates, is not a paradox at all.³

The effect of marriage behavior on measured birth rates has corollary implications for the nonmarital birth share, hereafter denoted NBS. As demonstrated below, an increase in the single share produces an proportionate increase in NBS share for given marital and nonmarital birth rates. However, the single share exerts an additional, equally powerful, effect on NBS by raising the nonmarital birth rate relative to the marital birth rate. Consequently, we predict a *magnified* effect of increases in the single share on the nonmarital birth share. In view of declines in marriage rates in recent decades, the theory offers a compelling explanation for the pronounced increases in nonmarital birth shares over the period.

We illustrate the paper's second central insight by positing opportunities for investments in human capital pertinent to women primarily in their early childbearing years. The pursuit of these opportunities is assumed to produce delays in both childbearing and marriage. The consequences include not only a lower marriage rate among young women, but also a correspondingly lower birth rate among unmarried young women. To the extent that investment opportunities are disproportionately available to black and white women, the model predicts differences in the timing of births across races – but not necessarily in *total* lifetime birth rates – consistent with the patterns emerging in recent U.S. data. Finally, the model is remarkably successful in explaining differences in the nonmarital birth share across age groups and across races.

The key theoretical results follow directly from a few simple definitions and a small number of deliberately strong assumptions, as outlined below.

A. Definitions

Let

MB	be the number of births to married women,
UB	be the number of births to unmarried women,
M	be the number of married women,
U	be the number of unmarried women.

³ The birth rate behavior described here is an example of Simpson's Paradox (Simpson, 1951).

Define the total birth rate, the marital birth rate, the nonmarital birth rate, the nonmarital birth share, and the single share as follows:

TBR is $(MB+UB)/(M+U)$, the total birth rate,
 MBR is MB/M , the birth rate of married women, or the marital birth rate,
 UBR is UB/U , the birth rate of unmarried women, or the nonmarital birth rate,
 NBS is $UB/(MB+UB)$, the share of nonmarital births, or the nonmarital birth share,
 Su is $U/(M+U)$, the fraction of women not married, or the single share.

Note that the variable NBS – the share of births to unmarried women – can be written in terms of the single share (Su) and the *ratio* of the nonmarital to the total birth rate (UBR/TBR).

$$\begin{aligned}
 \text{NBS} &= [U/(M+U)] \bullet [UB/(MB+UB)] \bullet [(M+U)/U] \\
 &= [U/(M+U)] \bullet (UB/U) \bullet [(M+U)/(MB+UB)] \\
 &= \text{Su} \bullet (\text{UBR}/\text{TBR})
 \end{aligned}
 \tag{Eq. (1)}$$

Eq. (1) is a common basis for demographic decompositions of NBS (as in Smith *et al.*, 1996), and a focal point of our theoretical and empirical contributions.

B. Childbearing Behavior

Women vary in their preference for children, which is captured by a parameter, γ , that measures desired family size. In a fully general life-cycle model, marriage and births would be jointly and endogenously determined. Our more simple model is consistent with a “fixed-target” model of fertility (c.f. Heckman *et al.*, 1985). In the end, we rely on the strength of the empirical tests to justify the simplicity of the fixed-target approach. We abstract from the various factors that cause desired and realized births to differ, so that γ also measures total births to a woman during her childbearing years. Also, we focus on births during the prime adult childbearing years, which we assume begin at age 20 and extend through age 39 for all women.⁴ Since our data are available by 5-year age intervals, the theory is developed as a four-period model, with the four periods corresponding to the age intervals 20-24, 25-29, 30-34, and 35-39.⁵ Successive cohorts of women are indexed by the subscript τ , where τ denotes the

⁴ Elsewhere (Gray *et al.*, forthcoming b), we examine the marriage and fertility behavior of teenagers aged 15-19.

⁵ The data are described in detail in sections III and IV below.

chronological time at which a cohort is in the first period of its childbearing span. Choice of family size, γ , is exogenous with respect to the model (in particular, it is independent of marital status), though marital status is not independent of γ , as described below. Furthermore, γ is distributed uniformly across women on the interval $[0, P_\tau]$, where $0 \leq P_\tau \leq 1$ and P_τ is allowed to vary by cohort.⁶

Now consider a representative group of z_τ women drawn from cohort τ . If the women are indexed and ordered by their preference for children, then the γ associated with the n^{th} “ordered” woman in the group is $(n/z)P$, as shown in Figure 1 below. (Note that the cohort subscript τ has been suppressed to streamline notation.) Figure 1 orders women along the horizontal-axis and their corresponding preference for children, γ , along the vertical axis. The relationship depicted is linear, a consequence of assuming that γ is uniformly distributed across women in the group, and one that makes the calculation of lifetime births for partitions of the group particularly expedient.⁷

[Figure 1 here]

Propositions (P1) through (P3) below give the lifetime birth rates for the first n ordered women in the group, the remaining $(z-n)$ women, and the group as a whole. Later, when n is identified with the number of women in a group who choose not to marry at any age, these propositions produce, respectively, the nonmarital, marital, and total lifetime birth rates of the group.

- (P1) for the first n ordered women in the group, the average number of lifetime births per woman (the lifetime birth rate) is $(1/2)(n/z)P$,
- (P2) for the remainder of the group, the lifetime birth rate is $(1/2)[(n/z)+1]P$,
- (P3) the lifetime birth rate for the group as a whole is $(1/2)P$.

⁶ 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, p. 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 and recent work suggesting that fertility arises in substantial part from deep genetic influences (e.g. Kohler et al. 1999 and Rodgers et al. 2001).

⁷ As a practical matter, the uniform distribution will likely best approximate the actual distribution of lifetime births if the extreme value P is extrapolated from the mean observed in the data (i.e., twice the mean), rather than setting it equal to the extreme value one might observe with a “Hutterite” level of fertility.

In the simplest version of the model, we assume that the number of births to a woman of a particular age is a fraction θ_i of her lifetime birth rate, where $i = 1, 2, 3, 4$ indexes the four successive (five-year) periods of a woman's childbearing life. The θ_i may vary by cohort, but always sum to one – a woman achieves her desired number of lifetime births by the end of her childbearing years. Furthermore, $\theta_1 > \theta_2 > \theta_3 > \theta_4$ in this baseline model. That is, women spread childbearing out over time, but typically prefer to have more children early than late. Reasons could include biological/health factors, a positive discount rate, greater security during retirement, etc. It is against this baseline description of childbearing that we measure of the effects of investment opportunities available to women in their early adult years.

C. Marriage Behavior

Our model of marriage allows the marriage behavior of women of a particular age to change over time – that is, to vary by cohort – in response to changes in the net benefits of marriage. The benefits to marriage are assumed to increase in the number of children a woman plans to have over her lifetime, but are not realized until she actually begins childbearing. Other factors relevant to the decision to marry – e.g. education levels, earnings, unemployment rates, etc. – are captured in a measure of net benefits that we denote “C”. These factors, which are assumed common to the women of a particular cohort, are the underlying source of the “exogenous” variation in S_u , by race and cohort (and therefore across time), that gives empirical content to our model. For a given value of C, there exists a critical value of γ , denoted γ^* , for which it is true that women with $\gamma > \gamma^*$ marry, and women with $\gamma < \gamma^*$ do not marry. Note that γ^* depends (positively) on C, and so may also vary by race and over time.

D. Investment Opportunities

The behavior described in the preceding sections is the benchmark against which we measure the effects of investment opportunities available to young women. To summarize, this baseline scenario is one in which lifetime births vary across women and may also vary across cohorts. Age-specific birth rates are a fraction of lifetime births, and decline monotonically with age. Women with sufficiently high birth rates marry, while those with low birth rates do not. A woman's marriage decision is made when

she is young (in the first childbearing period of life), and the proportion of women who remain single is constant over the life of a cohort.

We now introduce investment opportunities into the model. The opportunities, in the form of investments in human capital such as higher education, are limited to the youngest group of women, those in the first of the model's four childbearing periods. Of course, more generally, some investment opportunities might also extend into later childbearing periods, but here we assume that these are modest. If undertaken, investment indirectly raises utility later in life (e.g. through increased income), but requires delayed childbearing. We do not model the details of the cost-benefit calculation that determines whether a particular woman chooses to take advantage of investment opportunities when young. Rather, we assume that the presence of opportunities causes a fraction α of the youngest women to undertake investment and, consequently, to delay childbearing.

We allow the fraction of women who delay childbearing to vary by cohort. For simplicity, however, investment opportunities are assumed independent of the childbearing propensity, γ . Thus, the women who choose to delay are drawn randomly from each cohort and have lifetime childbearing propensities representative of their cohort. On average, these women plan the same number of lifetime births as women who do not delay, but distribute the births foregone in the first childbearing period over the remaining three age periods in proportion to the benchmark birth rates for each age.

E. Putting it Together

Finding age-specific, nonmarital birth rates for a cohort requires summing births over those married women who have chosen to delay marriage and childbearing in order to undertake investment and those who have not. The calculation of a cohort's age-specific marital and total birth rates involves similar aggregations for married women and for the cohort as whole. The algebra is straightforward, but tedious, and is provided in an appendix.

Because the model limits investment opportunities to the youngest group of women, the birth rates given in eqs. (12) through (14) below for the youngest age group (20-24) differ in important ways from the birth rates for older groups of women, which are given in eqs. (15) through (17).

Youngest Women

$$MBR_1 = \theta_1(1-2\alpha+Su_1)(P/2)/(1-\alpha) \quad \text{Eq. (12)}$$

$$UBR_1 = \theta_1(P/2)(Su_1 - \alpha)^2/(1-\alpha)Su_1 \quad \text{Eq. (13)}$$

$$TBR_1 = \theta_1(1-\alpha)(P/2) \quad \text{Eq. (14)}$$

Remaining Age Groups

$$MBR_i = \theta_i(P/2)(1+Su)[1+\alpha\theta_i/(1-\theta_i)] \quad \text{Eq. (15)}$$

$$UBR_i = \theta_i(P/2)Su[1+\alpha\theta_i/(1-\theta_i)] \quad \text{Eq. (16)}$$

$$TBR_i = \theta_i(P/2)[1+\alpha\theta_i/(1-\theta_i)], \text{ for } i = 2, 3, \text{ and } 4. \quad \text{Eq. (17)}$$

Recall that the subscript i represents age, by five-year interval, beginning with the first of four childbearing age intervals. The share of single women in the youngest age group, Su_1 , may differ from the single share common to the three older age groups, denoted simply Su . The difference is due to the assumption that young women who choose to delay childbearing until the second period of life also delay marriage until that time, increasing the single share for the youngest women relative to older age groups. The cohort subscript τ has been suppressed in the birth rate solutions, as elsewhere in the text.

We remind the reader that the solutions may vary across cohorts since the desired number of lifetime births, the baseline timing of births, investment opportunities that lead to delayed childbearing, and the attractiveness of marriage are all cohort specific. Furthermore, there is no *a priori* reason to believe that these factors are independent of race. Thus, the parameters that determine the birth rates given in eqs. (12) through (17) may vary by race, producing systematic differences in black and white birth rates. While blacks and whites are often presumed to differ in their preferences over the number and timing of births, we focus on the possibility that investment opportunities available to young adult women

differ for blacks and whites. Indeed, the data and statistical tests presented in subsequent sections strongly suggest that differences in marriage behavior and investment opportunities, rather than intrinsic differences in the desire for children, are the primary cause of racial differences in observed birth rates.

III. Interpretation and Empirical Implications

To facilitate interpretation and motivate our empirical tests, we begin by considering the benchmark case in which there are no investment opportunities available to young women. A focus of the paper's empirical applications is the hypothesis that this baseline scenario is more relevant to young black women than to young white women.⁸

A. Baseline: No Investment Opportunities

The absence of significant investment opportunities is captured by setting $\alpha=0$ in eqs. (12) through (17). This case produces birth rates that vary by age only to the extent that the baseline timing parameter, θ_i , differs by age:

$$MBR_i = \theta_i(P/2)(1+Su) \quad \text{Eq. (18)}$$

$$UBR_i = \theta_i(P/2)Su \quad \text{Eq. (19)}$$

$$TBR_i = \theta_i(P/2), \text{ for } i = 1, 2, 3, \text{ and } 4. \quad \text{Eq. (20)}$$

In eqs. (18) through (20) all three birth rates increase as the preference for children increases, as one would expect. Because we have assumed that the θ_i decline with age, all birth rates decline with age. Furthermore, given that women who prefer more children are more likely to marry in our model, the marital birth rate in eq. (18) exceeds the nonmarital birth rate in eq. (19). More interesting are the effects of changes in the single share, Su , on birth rates. Increases in Su cause *both* the marital and nonmarital birth rates to increase, even though the total birth rate is unaffected by Su . In addition, the nonmarital

⁸ Differences in the percentage of black and white women who complete college is one indicator of racial differences in human capital investment opportunities. In the late 1950s, the percentage of black women who had completed college was barely half that of white women. Even by 2004, the ratio had risen only to 65% (Digest of Education Statistics).

birth rate increases proportionately more than the marital birth rate. To see this, note that UBR_i increases linearly in S_u (for given values of θ_i and P), while MBR_i increases linearly in $(1+S_u)$. Thus, while the absolute size of the increases in UBR_i and MBR_i produced by an increase in S_u are identical, UBR_i (with a smaller initial value than MBR_i) increases *proportionately* more. It follows that UBR_i will also increase proportionately more than the total birth rate, TBR_i , when S_u increases.

Our model's predictions for the behavior of MBR , UBR , and TBR , as expressed in eqs. (18) through (20), are difficult to test directly because the parameters that appear in these solutions – the θ_i and P – are not observable. However, the model's predictions for the *ratios* of birth rates, in particular, MBR/TBR and UBR/TBR , can be tested directly, as can the model's implications for the nonmarital birth share, NBS . As eqs. (21) through (23) verify, these ratios depend only on the single share, S_u , which is observed in the data, but do not depend on age or the number or timing of births. Nor do they depend on the costs and benefits of marriage (captured in the parameter C), except as reflected indirectly in S_u .

$$MBRR_i = MBR_i/TBR_i = (1+S_u) \quad \text{Eq. (21)}$$

$$UBRR_i = UBR_i/TBR_i = S_u \quad \text{Eq. (22)}$$

$$NBS_i = S_u \bullet (UBR_i/TBR_i) = S_u^2 \quad \text{Eq. (23)}$$

Note that the birth rate ratios $MBRR_i$ and $UBRR_i$ both increase one-for-one with increases in S_u . Indeed, $UBRR_i$ is *equal* to S_u . Using this result in eq. (1), we see that an increase in S_u raises NBS_i both because S_u appears directly in the expression for NBS_i , and also because it raises $UBRR_i/TBR_i$. Thus, the single share has a magnified effect on the nonmarital birth share and, for the baseline case, NBS_i becomes simply S_u^2 , as indicated in eq. (23).

To summarize, when investment opportunities for young women are very limited, as hypothesized in the case of young black women, we should observe:

Implication 1: The sum $(1+S_u)$ closely tracks the ratio of the married birth rate to the total birth rate, $MBRR$.

Implication 2: The single share, S_u , closely tracks the ratio of the unmarried birth rate to the total birth rate, UBRR.

Implication 3: The squared value of the single share, S_u^2 , closely tracks the share of unmarried births, NBS.

These are strong implications, and stark in their simplicity. Furthermore, as emphasized above, they are independent of age and cohort.

Figures 2 through 4 below provide striking support for the hypothesized relationships in the case of black women, who may be argued to be disadvantaged in terms of investment opportunities. The plots cover the years 1969-2002, the period over which the necessary data are available separately for blacks.⁹ Figure 2 shows $(1+S_u)$ and the married birth rate ratio, MBRR, presented as time series. The data for women aged 20-24 appear in the first panel of figure 2. The data for women aged 25-29 appear in the second panel, and so on. The model predicts that $(1+S_u)$ and the married birth rate ratio will be equal, regardless of age or cohort. The overall consistency of figure 2 with this hypothesis is impressive – all the more so in view of the notoriously noisy data available separately on the numbers of married and unmarried black women. Figures 3 and 4 present strong evidence consistent with implications 2 and 3.

(Figures 2-4 here)

B. Investment Opportunities

Using the results of section A as a benchmark, we can now investigate the role of investment opportunities facilitated by delayed childbearing. As in the previous section, it will be useful to focus on the model's results in terms of implications for birth rate ratios, since this eliminates unobservable parameters. Because investment opportunities are hypothesized to be available only to the youngest group of women, it is important to separate effects for this age group from those over the remainder of the childbearing years.

⁹ Births by marital status are from National Vital Statistics Reports (2000, 48:16; 2002, 50:10; and 2003, 52:10). Total births are from Vital Statistics of the United States (www.cdc.gov/nchs/births.htm). The numbers of married women (defined as married, spouse present) and total women are from U.S. Bureau of Census, Current Population Reports, Series P-20, various dates. The number of unmarried women is calculated as the difference between total and married women.

Youngest Women

Because investment opportunities cause a representative group of the youngest women to delay childbearing, it reduces the number of both married and unmarried women in a cohort who give birth during the first childbearing period of life. This reduces the total number of children born to both married and single women when they are young. Those younger women who choose to delay childbearing also delay marriage, so the number of unmarried women in this age group rises relative to the benchmark case. With unmarried births falling and the number of unmarried women rising, the unmarried birth rate at this age (UBR_1) must be lower than in the benchmark case.

Like those who delay childbearing, the young women who do *not* delay childbearing are otherwise representative of their cohort, which means that non-delay women have children and marry at rates equal to the rates of the benchmark case. Given that none of the delaying women marry, the *only* married women in the youngest group are women who choose not to delay. Accordingly, the married birth rate of the cohort when it is young (MBR_1) is the same as in the benchmark case. The total birth rate (TBR_1), is a weighted average of UBR_1 and MBR_1 . Since UBR_1 falls relative to the benchmark value but MBR_1 is unchanged, TBR_1 falls – but proportionately less than UBR_1 . Thus, the presence of investment opportunities causes the ratio MBR_1/TBR_1 to rise, and the ratio UBR_1/TBR_1 to fall, relative to the baseline case. These effects are verified in eqs. (24) and (25), which follow directly from eqs. (12) through (14).

$$MBRR_1 = MBR_1/TBR_1 = (1+Su_1 - 2\alpha)/(1-\alpha)^2 \quad \begin{array}{l} = (1+Su_1) \text{ if } \alpha = 0 \\ > (1+Su_1) \text{ if } \alpha > 0 \end{array} \quad \text{Eq. (24)}$$

$$UBRR_1 = UBR_1/TBR_1 = [(Su_1 - \alpha)/(1-\alpha)]^2(1/Su_1) \quad \begin{array}{l} = Su_1 \text{ if } \alpha = 0 \\ < Su_1 \text{ if } \alpha > 0 \end{array} \quad \text{Eq. (25)}$$

Setting α equal to zero in eqs. (24) and (25) confirms that each birth rate ratio is equal to its benchmark value in the absence of investment opportunities. (Compare to eqs. (21) and (22).) In the presence of investment opportunities, however, the ratios deviate from their benchmark values in opposite

(and testable) ways, as discussed above. In particular, it can be shown that $MBRR_1$ exceeds $(1+Su)$, and $UBRR_1$ falls short of Su , by an amount that increases monotonically in α .

The nonmarital birth share for this age group is found by substituting eqs. (13) and (14) into (1):

$$NBS_1 = [(Su_1 - \alpha)/(1-\alpha)]^2 \quad \begin{array}{l} = (Su_1)^2 \text{ if } \alpha = 0 \\ < (Su_1)^2 \text{ if } \alpha > 0 \end{array} \quad \text{Eq. (26)}$$

NBS_1 reduces to its benchmark value, $(Su_1)^2$, when α is zero, but is otherwise less than $(Su_1)^2$. Thus, the model predicts that the presence of investment opportunities – e.g. access to higher education or on-the-job training – drives a wedge between the values of $(Su_1)^2$ and NBS_1 . Note that NBS_1 itself is invariant with respect to α , since the direct negative effect of α in eq. (26) is exactly offset by the indirect positive effect that α exerts through Su_1 . Intuitively, the only births among the youngest group of women are due to the women in the non-delay group, and this group has marital and nonmarital births at the same rates as in the benchmark case.

Remaining Age Groups

In the remaining childbearing periods, women who did not delay childbearing in order to undertake investment continue having children at the benchmark rates. Those who delayed now begin having children and they do so at higher rates than their non-delay counterparts – a catch-up effect. However, the nonmarital, marital, and total birth rates of the delay group all rise in the same proportion, so the birth rate *ratios* of these women equal those of their non-delay counterparts. Accordingly, the birth rate ratios for the cohort as a whole are the benchmark values given by eqs. (21) and (22). Similarly, the nonmarital birth share is equal to its benchmark value, given by eq. (23), for women in the three later periods of childbearing.

To summarize, if investment opportunities are significantly available to young white women only, we expect a divergence from the baseline case for this group only. Accordingly, we should observe:

Implication 1: $MBRR$ exceeds $(1+Su)$ in early childbearing years, but equals $(1+Su)$ in later childbearing years .

Implication 2: UBRR is less than S_u in early childbearing years, but equals S_u in later childbearing years.

Implication 3: NBS exceeds the squared value of the single share, S_u^2 , in early childbearing years, but equals S_u^2 in later childbearing years.

Figures 5 through 7 below provide considerable support for the hypothesized relationships in the case of white women, who we presume have access to significant investment opportunities during their early childbearing years. The plots cover the years 1957-2002, with the exception of women aged 35-39, for whom data by five-year interval begin in 1968. In Figure 5, the married birth rate ratio substantially exceeds $(1+S_u)$ for the youngest group of women, while the two series match very closely for older age groups, consistent with the first of the three implications above.

(Figures 5-7 here)

Figure 6 provides evidence on the second implication. In the first panel of figure 6, the unmarried birth rate ratio, UBRR, lies substantially below S_u , as predicted by the theory. The difference between the two series is much smaller in the remaining panels, which is also consistent with the theory. However, the second and last panels diverge, at least on average, from the strict prediction that UBRR and S_u are *exactly* equal. In the second panel, for women aged 25-29, UBRR falls somewhat short of S_u , which is quite plausibly explained by generalizing the theory to permit investment opportunities in the second childbearing period, as well as the first. In the final panel, for ages 35-39, S_u now tends to fall somewhat short of UBRR, at least on average, a divergence not explained by the same generalization. Instead, it may reflect the more complex nature of childbearing and marriage status at the end of the childbearing years. In any event, we examine these implications more formally in Section IV.

In Figure 7, which examines the third and most comprehensive implication of the theory, we once again see striking consistency between the data and the theory. The nonmarital birth share, NBS, substantially exceeds its benchmark value, S_u^2 , for the youngest age group, while the two variables take very similar values for the three older age groups.

C. Racial Differences

The model implications illustrated thus far are general, in that they are independent of the unobservable parameters that capture the desire to have children (P) and the timing of births (the θ_i). That is, the implications hold even if blacks and whites have very different “tastes” regarding family size and the timing of births. Here we explore a more restrictive form of the model, one in which blacks and whites are assumed to share the same preferences with respect to the number and timing of children. The focus is on the effect of investment opportunities on age-specific total birth rates. If preferences regarding childbearing are the same, then eqs. (14) and (17) imply that racial differences in total birth rates are due solely to racial differences in the investment opportunities available to young women.

We begin by examining the assumption that blacks and whites are similar in the number of children they wish to have. Figure 8 suggests that this is a good description of the facts for the age groups and time periods covered in this study. The figure presents estimates by race of the number of births occurring between the ages of 20 and 39 for married and unmarried women combined. The estimates are generated by (i) multiplying the annualized birth rates per thousand women by five to obtain 5-year birth rates, (ii) dividing by 1,000 to obtain births per woman, and (iii) adding together the resulting birth rates for women aged 20-24 at time t , women aged 25-29 at time $t+5$, women aged 30-34 at time $t+10$, and women aged 34-39 at time $t+15$. The sum is recorded on the vertical axis of figure 8 against the year in which the women were aged 20-24, which identifies cohort, on the horizontal axis. As the figure shows, these age-aggregated birth rates are very similar for blacks and whites throughout the period, becoming virtually indistinguishable for more recent cohorts.

(Figure 8 here)

Births occurring over the ages 20 to 39 capture most, but certainly not all, lifetime births. The wider is the age range covered, however, the fewer and older are the cohorts for which this methodology can be used to estimate lifetime birth rates. Basing estimates of lifetime birth rates on five-year birth rates for women aged 15-44, for example, means that the latest cohort for which estimates can be constructed was aged 15-19 in 1977. Still, even these estimates suggest considerable – and increasing – similarity in

desired family size for blacks and whites. Black women in this cohort had roughly 2.3 lifetime births as compared to 2.0 births for whites. Furthermore, given the behavior of total fertility rates since the late 1970s, the difference seems likely to have narrowed further for more recent cohorts. Indeed, to the extent that the total fertility rates for women aged 15-44 are an indicator of lifetime birth rates – which would be the case in the absence of changes over time in childbearing preferences and the age composition of the population – the evidence suggests that lifetime birth rates for very recent cohorts of black and white women are virtually identical.¹⁰

Absent investment opportunities and differences in childbearing preferences, our baseline model predicts that total birth rates will decline with age, but will be the same for blacks and whites of a given age. Incorporating investment opportunities into the model lowers birth rates for the youngest group of women and raises them for older groups (compare eqs. (14) and (17) for positive values of α). If, however, investment opportunities have been significantly available for whites but much less so for blacks, then the ratio of the first-period birth rate to the second period birth rate should be lower for whites than for blacks. That is, one should observe that black women “front-load” their childbearing relative to white women. Indeed, as figure 9 demonstrates, this is the case. The steady decline in the ratio for whites further suggests expanding opportunities for young white women over the period.

(Figure 9 here)

The theory also predicts that while the youngest group of black women should have higher birth rates than the youngest group of white women, older groups of blacks should have *lower* birth rates than older groups of whites. Alternatively stated, the ratio of black to white births rates should exceed unity for the youngest women, but should be less than one for older groups, as confirmed in figure 10.

(Figure 10 here)

¹⁰ See “Births: Final Data for 2002”, *National Vital Statistics Reports*, Vol. 52, No. 10, December 17, 2003, Centers for Disease Control and Prevention, Table 4. The total fertility rate for a particular year is the expected number of births to a hypothetical cohort of 1,000 women over the ages 15-44 under the assumption that they experience the same age-specific birth rates observed in that year. It is an estimate of total lifetime births under the assumption that current fertility patterns are maintained.

The empirical implications formulated in this section, along with the informal evidence presented, are provocative. The implications are qualitatively unambiguous, in that effects are always signed, as well as quantitatively specific. For example, in the absence of investment opportunities at a particular age, the model predicts that the ratio of the married birth rate to the total birth rate will equal one plus the single share. That the data appear to so closely match such stark predictions is strong motivation for the more formal statistical tests to which we turn next.

IV. Empirical Tests.

Our theoretical model yields distinctive parameterizations for the marital birth rate ratio (MBRR), the nonmarital birth rate ratio (UBRR), and the nonmarital birth share (NBS). The parameterizations differ in precise ways depending upon whether or not there are opportunities for human capital investment. In this section, we estimate and formally test the model using the data introduced in section III. The pooled data (two races and four age groups) yield a total of 309 observations. For each of the model's key predictions, we present estimates for three cases of interest. In the first, no opportunities for investment in human capital are present. In the second, investment opportunities are hypothetically available only to young white women aged 20-24. In the third, investment opportunities may also be available to young black women aged 20-24. While the data for white and black women are pooled in the estimation, we test for the likely differences between the two groups based upon the model predictions. Similar results are obtained when specifications for white and black women are estimated separately.

Empirical models for MBRR, UBRR, and NBS follow directly from eqs. (24) – (26). In each case, the model predictions below are nested in a single equation incorporating the binary variable W_{20-24} , which takes the value one for white women aged 20-24 and zero otherwise:¹¹

$$MBRR_i = b_0 + b_1 Su_i + b_2 W_{20-24} + b_3 Su_i W_{20-24}, \quad i = 1, 2, 3, 4, \quad \text{Eq. (27)}$$

$$\text{where } b_0 = 1, b_1 = 1, b_2 = -\alpha^2/(1-\alpha)^2 < 0, b_3 = [1/(1-\alpha)^2] - 1 > 0.$$

¹¹ Error terms are omitted for simplicity. Properties are discussed below, and robust standard errors are reported.

$$UBRR_i = c_0 + c_1 Su_i + c_2 W_{20-24} + c_3 Su_i W_{20-24} + c_4 (1/Su_i) W_{20-24}, \quad i = 1, 2, 3, 4, \quad \text{Eq. (28)}$$

$$\text{where } c_0 = 0, c_1 = 1, c_2 = -2\alpha/(1-\alpha)^2 < 0, c_3 = [1/(1-\alpha)^2] - 1 > 0, c_4 = \alpha^2/(1-\alpha)^2 > 0.$$

$$NBS_i = d_0 + d_1 Su_i^2 + d_2 W_{20-24} + d_3 Su_i^2 W_{20-24} + d_4 Su_i W_{20-24}, \quad i = 1, 2, 3, 4, \quad \text{Eq. (29)}$$

$$\text{where } d_0 = 0, d_1 = 1, d_2 = \alpha^2/(1-\alpha)^2 > 0, d_3 = [1/(1-\alpha)^2] - 1 > 0, d_4 = -2\alpha/(1-\alpha)^2 < 0.$$

A. Married Birth Rate Ratio

For each variable, we begin by estimating the model corresponding to the special case of no investment opportunities. Setting $\alpha=0$ in eq. (27) produces the coefficient values $b_0=b_1=1$ and $b_2=b_3=0$, yielding the prediction that, in the absence of investment opportunities, MBRR is identically equal to $(1+Su)$ for every age group. Accordingly, only a constant and Su appear in the corresponding empirical specification, and the predictions tested are that both the constant and the coefficient on Su equal one. Indeed, estimates of this model reported in column (1) of table 1 yield a constant (0.9732) that is not significantly different from one at the five percent level but is significantly different from zero. Similarly, the coefficient on Su (1.1379) is not significantly different from one, but is significantly different from zero. Inferences are based on robust standard errors, reported in parentheses. The latter correct for serial correlation, heteroskedasticity, and contemporaneous correlations among the cross-sectional age groups.¹²

(Table 1 here)

While the data do not reject the baseline model reported in the first column of table 1, they nonetheless support expanding the model to account for human capital investment by young white women. A positive value of α in eq. (27) produces model predictions for white women aged 20-24 that contrast sharply with the no-investment case – an intercept strictly *less* than one and a coefficient on Su^2 strictly *greater* than one. For all other race/age groups, however, the model predictions are the same as in the no-investment case; both the constant and the coefficient on Su should equal one.

¹² All of the key variables exhibit non-stationarity, reflected by a failure to reject the null of a unit root in each case. In previous work (Gray *et al.*, forthcoming a), we find significant co-integrating vectors for the corresponding variables. Hence, here we employ least-squares estimates of the coefficients, along with robust standard errors.

Thus, if there is substantial human capital investment by white women aged 20-24, but not by women in other race/age groups, we again expect to find that the estimated values of b_0 and b_1 are not significantly different from one. However, we also expect a *negative* intercept shift for the youngest group of white women, identified by the binary variable W_{20-24} in eq. (27), and a *positive* coefficient on the interaction between S_u and W_{20-24} . That is, we expect $b_2 < 0$ and $b_3 > 0$. All these predictions are born out in the second column of table 1. We note that the interaction terms implied by the presence of investment opportunities for young white women are not only highly significant, but also collectively add significantly to the power of the equation, evidence of substantial investment by this group.¹³

Eq. (27) provides the theoretical values of the coefficients b_2 and b_3 , each of which is uniquely related to the delay parameter, α . These restrictions can be solved individually for α and, in conjunction with the estimates of b_2 and b_3 reported in the column (2) of Table 1, imply values of the delay parameter. Although produced by separate calculations, the implied values of α are identical to two decimal places -- 0.30 in each case.¹⁴ Thus, estimates of the theory's implications for MBRR suggest that 30 percent of white women aged 20-24 delay childbirth and marriage.

The third column of table 1 reports the results of expanding the empirical model to account for the possibility of significant investment opportunities for young black women. We expect the coefficients related to the binary variable for young black women (i.e., B_{20-24} and $S_u B_{20-24}$) to be insignificant in the absence of substantial investment opportunities for this group. That prediction is born out in column 3 of table 1. Neither coefficient is significantly different from zero – indeed, neither even exceeds the corresponding standard error.

B. Unmarried Birth Rate Ratio

The results for UBRR are reported in table 2, also along with robust standard errors in parentheses. The model predictions for the no-investment case, produced by setting $\alpha=0$ in equation (28),

¹³ Expanding the empirical model to account for the possibility of human capital investment by white women aged 25-29 yields insignificant results for this older age group.

¹⁴ For a value of α equal to 0.30, the data fail to reject the model restrictions given in eq. (27), either individually or collectively, at the five percent level.

are $c_1=1$ and $c_0=c_2=c_3=c_4=0$. That is, UBRR is identically equal to the single share, S_u , for every age group. Accordingly, the empirical model for this case includes, as for MBRR above, only a constant and S_u . Here, however, the predictions tested are a constant of *zero* along with a coefficient on the single share (S_u) equal to one. The estimates reported in column (1) of table 2 are consistent with these predictions – a constant (0.0066) that is not significantly different from zero at the five percent level and a coefficient on S_u (0.9369) that is not significantly different from unity.

(Table 2 here)

Again, however, the data support expanding the model to account for human capital investment by young white women. A positive value of α in eq. (28) implies a negative intercept shift for the youngest group of white women, identified by W_{20-24} , a positive coefficient on the interaction between S_u and W_{20-24} , and a positive coefficient on a similar interaction of W_{20-24} with $1/S_u$. That is, we expect $c_2 < 0$, $c_3 > 0$, and $c_4 > 0$. The predicted values of the intercept and the coefficient on S_u remain zero and one, respectively – that is, $c_0 = 0$ and $c_1 = 1$. These predictions are born out in column (2) of table 2, which reports highly significant coefficients on the interaction terms associated with investment by young white women, and a substantial increase in the power of the equation.¹⁵

The coefficients associated with the variable W_{20-24} and its interactions are, again, uniquely related to the delay parameter, α . As in the case of MBRR, these restrictions can be solved individually for α and, in conjunction with the estimates of c_2 , c_3 , and c_4 reported in the column (2) of Table 2, imply values of the delay parameter, α . The implied values are again nearly identical; all lie in the range .38 to .39.¹⁶ Thus, estimates of our model's implications for UBRR suggest that just under 40 percent of white women aged 20-24 delay childbirth and marriage.

¹⁵ Expanding the empirical model to account for the possibility of human capital investment by white women aged 25-29 yields much smaller, but statistically significant, results for those coefficients, consistent with a model in which smaller investments continue into the late twenties.

¹⁶ The coefficient restrictions given in eq. (28) cannot be rejected at the five percent level, either individually or collectively, for choices of a common α in the range .38 to .39.

The third column of table 2 reports the results of expanding the empirical model to account for the possibility of significant investment opportunities for young black women. The additional terms (i.e., B_{20-24} , SuB_{20-24} , and $(1/Su)B_{20-24}$) are statistically insignificant, suggesting, as for MBRR in table 1, the absence of substantial investment opportunities for this group.

C. Nonmarital Birth share

The results for NBS are presented in table 3. For the no-investment case, produced by setting $\alpha=0$ in equation (29), the theory predicts $d_1=1$ and $d_0=d_2=d_3=d_4=0$. That is, NBS is identically equal to Su^2 for every age group. Accordingly, only Su^2 and a constant appear in the corresponding empirical specification, and the predictions tested are that the constant is zero and the coefficient on the Su^2 is one. Indeed, the estimates of this model reported in column (1) of table 1 yield a constant (-0.0065) not significantly different from zero at the five percent level and a coefficient on Su^2 (0.9733) not significantly different from unity.

(Table 3 here)

Once again, the data support expanding the model to account for human capital investment by young white women. A positive value of α in eq. (29) implies a *positive* intercept shift for the youngest group of white women, identified by W_{20-24} ; a *positive* coefficient on the interaction of Su^2 with W_{20-24} ; and a *negative* coefficient on the interaction of W_{20-24} with Su . Thus, we expect $d_2>0$, $d_3>0$, and $d_4<0$. For all other groups, the model predicts a zero intercept and a coefficient on Su^2 equal to one – $d_0=0$ and $d_1=1$. All these predictions are born out in column (2) of Table 3. As in the cases of MBRR and UBRR, highly significant coefficients on the interaction terms associated with investment by young white women and a substantial increase in the power of the equation provide evidence of investment by young white women.¹⁷

¹⁷ Permitting human capital investments to extend into the second age group (25-29) yields insignificant results for the associated coefficients.

The coefficients associated with the variable W_{20-24} and its interactions are, once again, uniquely related to the delay parameter, α . These restrictions can be solved individually for α and, in conjunction with the estimates of d_2 , d_3 , and d_4 reported in the column (2) of Table 3, imply values of the delay parameter, α . In this case, the implied values all lie in the range .40 to .42.¹⁸ Thus, estimates of our model's implications for NBR suggest that slightly more than 40 percent of white women aged 20-24 delay childbirth and marriage.

Finally, column 3 of table 3 reports insignificant results for additional terms that capture the possibility of investment opportunities for young black women (i.e., B_{20-24} , Su^2B_{20-24} , and SuB_{20-24}). As in earlier cases, none of the individual coefficients for these variables exceeds the corresponding standard error, once again suggesting the absence of substantial opportunities for this group.

Overall, the results reported in this section provide considerable support for three key predictions of our theoretical model: 1) changes in marital and nonmarital birth rate ratios (MBRR and UBRR) and the nonmarital fertility share (NBS) all appear to have been driven primarily by changes in marriage behavior, rather than by changes in fertility, *per se*; 2) opportunities for investments in human capital play a distinctive role in the timing of marriage and fertility over the adult childbearing years; and 3) racial differences in these investments explain key differences in the marital and fertility patterns for black and white women. The coefficient estimates are inconsistent with significant delays in childbearing and marriage by young black women in response to human capital investment opportunities. By contrast, our empirical results are strikingly consistent in supporting the presence of such delays among young white women, with implied values for the proportion of young white women who delay childbearing and marriage in order to make investments in human capital ranging from .30 to .42.

¹⁸ The coefficient restrictions given in eq. (29) cannot be rejected individually at the five percent level for choices of a common α in the range .40 to .42; however, the full set of five restrictions are collectively rejected.

V. Discussion and Concluding Remarks

Our theoretical model and empirical estimates attribute observed increases in nonmarital and marital birth rates in recent decades primarily to a decline in the marriage rate, not to changes in fertility behavior, *per se*. Our explanation is consistent with increases in both nonmarital and marital birth rates, even in the absence of changes in the total birth rates. As the proportion of women who marry declines, the birth rates of unmarried and married *necessarily* rise, all else the same, with the increase in the unmarried birth rate necessarily rising proportionately more than either the married or total birth rates. Accordingly, the nonmarital birth share – the share of births to unmarried women – also rises. It follows that the observation of higher marital and nonmarital birth rates, along with a higher nonmarital birth share, for black women than for white women may reflect differences in marriage behavior, not desired family size. Indeed, the marriage rate is substantially lower for blacks than whites, while *total* birth rates for black and white women are nearly identical in recent years.

The theoretical model further predicts that opportunities for investment in human capital early in the childbearing years will introduce differences between black and white women in the timing of childbearing and marriage decisions. Empirical estimates suggest that a substantial portion of the youngest group of white women (ages 20-24) delay marriage and childbearing to pursue investments in human capital, while the youngest group of black women do not – at least not significantly so. These results, along with the apparent similarity in total lifetime birth rates for blacks and whites, imply that black women have more children early in their childbearing years than white women, while the reverse is true later in life, predictions confirmed by the data.

Our findings are particularly relevant in the context of studies that take marriage behavior as given in evaluating the effects of public policy on fertility, especially on differences between black and white women. (See, for example, Baughman and Dickert-Conklin, 2003.) The direct role of policy on changes in fertility will likely be overstated in these studies relative to the role of changes in marriage behavior. Presumably the distinction is important, both for the construction of public policy and for the

evaluation of the effects of public policy. Hence, future research might productively be focused more directly on changes in marriage behavior, a direction taken in recent papers by Bitler et al (2004), Fitzgerald and Ribar (2004), Grogger and Bronars (2001), and Moffitt (2000).

Finally, our findings suggest that much of the difference between birth rates for black and white women arises from differences in decisions about investments in human capital early in the childbearing years. It will be important to explore more directly the link to investments in human capital, which we have done only indirectly here. Nonetheless, the paper's results suggest that increasing the incentives and opportunities for human capital investment by young black women may be an especially useful direction for public policy aimed at reducing nonmarital births among young women.

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Table 1 Married Birth Rate Ratio (MBRR)

(robust standard errors)

Variable	Predicted	No	Investment	Investment
	Coeff	Investment	Whites	Whites & Blacks
		(1)	(2)	(3)
C	1	0.9732** (0.0988)	0.9890** (0.0375)	1.0400** (0.0201)
Su	1	1.1379** (0.2067)	0.9863** -0.0789	0.8206** (0.0521)
W(20-24)	-		-0.1808* (0.0787)	-0.2318* (0.0624)
Su*W(20-24)	+		1.0156** (0.1954)	1.1813** (0.1461)
B(20-24)	0			0.1001 (0.1043)
Su*B(20-24)	0			0.1042 (0.1593)
R2		0.6812	0.8954	0.9187
nobs		309	309	309

*(**) significantly different from zero at the 5 (1) percent level.

Notes: Dependent variable is the married birth rate ratio, the ratio of the marital birth rate to the total birth rate (MBRR). Su is the single share, the ratio of unmarried to total women. Data are age group (20-24, 25-29, 30-34, 35-39) by race (white, black), 1957-2002 for white women (1968-2002 for 35-39), and 1969-2002 for black women. See text for further details. Standard errors are robust standard errors, corrected for serial correlation, time-varying heteroskedasticity, and contemporaneous correlations.

Table 2 Unmarried Birth Rate Ratio (UBRR)

(robust standard errors)

Variable	Predicted Coeff	No Investment (1)	Investment Whites (2)	Investment Whites & Blacks (3)
C	0	0.0066 (0.0869)	0.0235 (0.0413)	0.0053 (0.0418)
Su	1	0.9369** (0.1600)	0.9928** -0.0813	1.0538** (0.0942)
W(20-24)	-		-1.9704** (0.7156)	-1.9523** (0.6280)
Su*W(20-24)	+		1.6832** (0.6180)	1.6222** (0.5305)
1/Su*W(20-24)	+		0.3883* (0.1780)	0.3883* (0.1574)
B(20-24)	0			-0.8260 (2.1143)
Su*B(20-24)	0			0.5534 (1.3729)
1/Su*B(20-24)	0			0.2454 (0.7797)
R2		0.7265	0.9154	0.9216
nobs		309	309	309

*(**) significantly different from zero at the 5 (1) percent level.

Notes: Dependent variable is the unmarried birth rate ratio, the ratio of the nonmarital birth rate to the total birth rate (UBRR). Su is the single share, the ratio of unmarried to total women. Data are age group (20-24, 25-29, 30-34, 35-39) by race (white, black), 1957-2002 for white women (1968-2002 for 35-39), and 1969-2002 for black women. See text for further details. Standard errors are robust standard errors, corrected for serial correlation, time-varying heteroskedasticity, and contemporaneous correlations.

Table 3 Nonmarital Birth Share (NBS)

(robust standard errors)

Variable	Predicted Coeff	No Investment (1)	Investment Whites (2)	Investment Whites & Blacks (3)
C	0	-0.0065 (0.0287)	0.0071 (0.0108)	-0.0017 (0.0085)
Su2	1	0.9733** (0.0829)	1.0066** (0.0331)	1.0736** (0.0359)
W(20-24)	+		0.4564** (0.1397)	0.46522** (0.0981)
Su2*W(20-24)	+		1.9537** (0.5189)	1.8867** (0.3385)
Su*W(20-24)	-		-2.2503** (0.5776)	-2.2503** (0.3929)
B(20-24)	0			0.1936 (0.3316)
Su2*B(20-24)	0			0.4368 (0.5665)
Su*B(20-24)	0			-0.6752 (0.8826)
R2		0.9156	0.9799	0.9842
nobs		309	309	309

*(**) significantly different from zero at the 5 (1) percent level.

Notes: Dependent variable is the nonmarital birth share, the ratio of the nonmarital birth rate to the total birth rate (NBS). Su is the single share, the ratio of unmarried to total women. Data are age group (20-24, 25-29, 30-34, 35-39) by race (white, black), 1957-2002 for white women (1968-2002 for 35-39), and 1969-2002 for black women. See text for further details. Standard errors are robust standard errors, corrected for serial correlation, time-varying heteroskedasticity, and contemporaneous correlations.

Figure 1. Women Ordered by Preference for Children (γ)

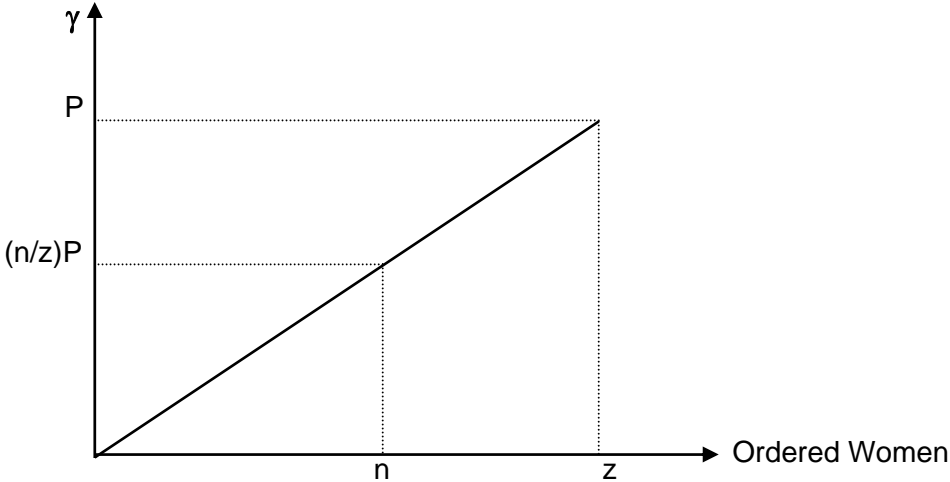


Figure 2. MBRR and (1+Su) by Age, Black Women 20-39

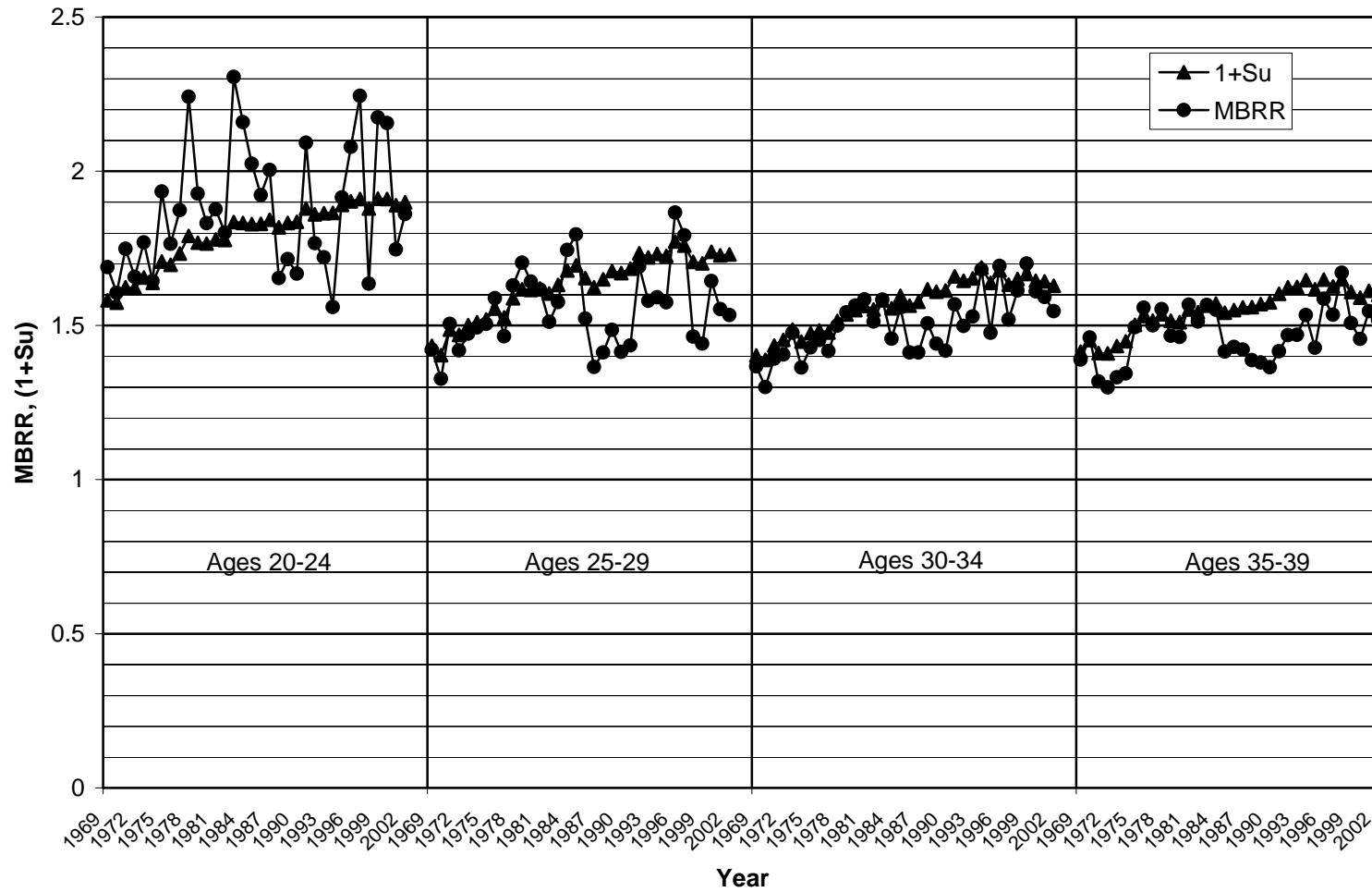


Figure 3. UBRR and Su by Age, Black Women 20-39

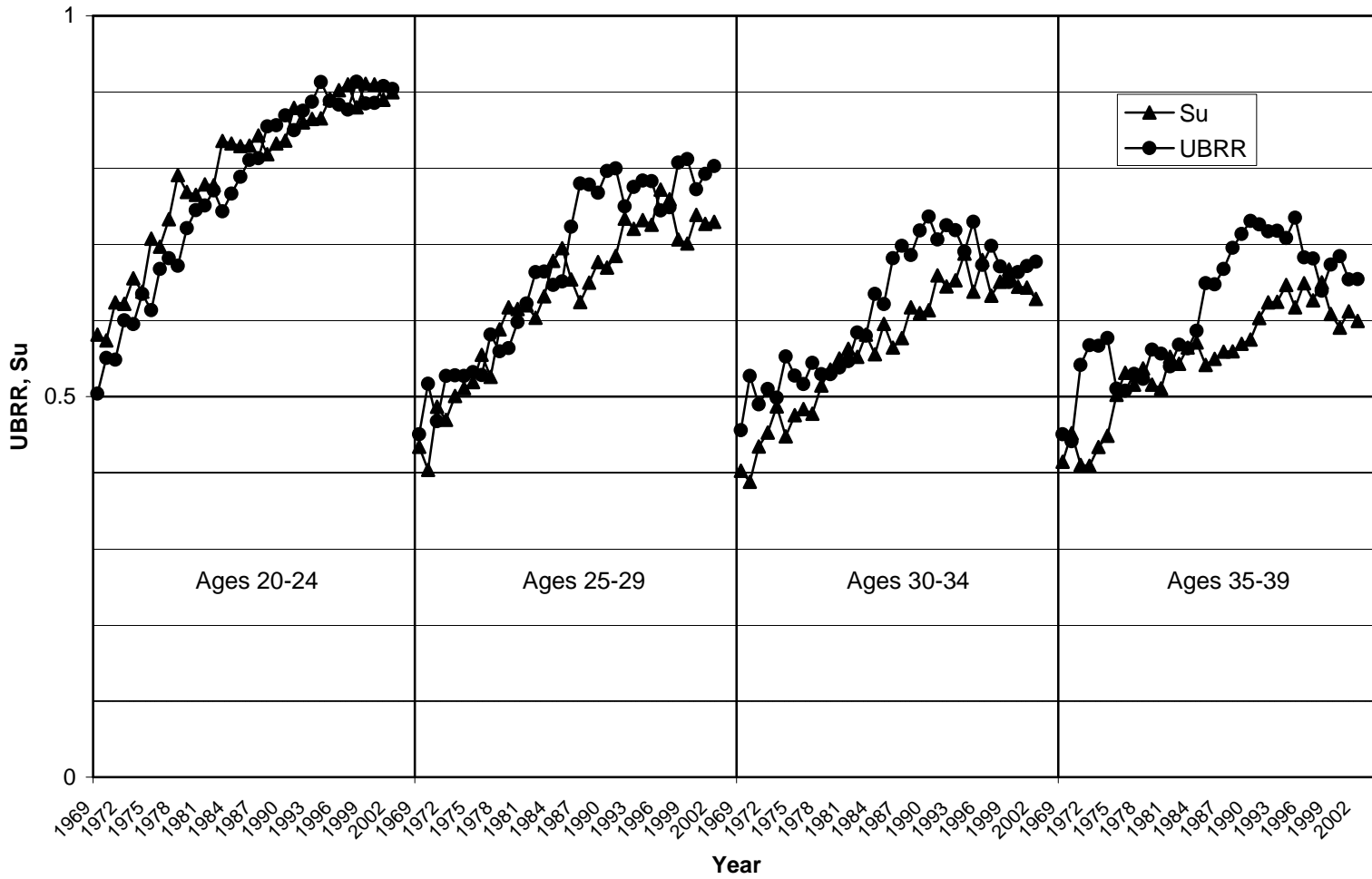


Figure 4. NBS and Su2 by Age, Black Women 20-39

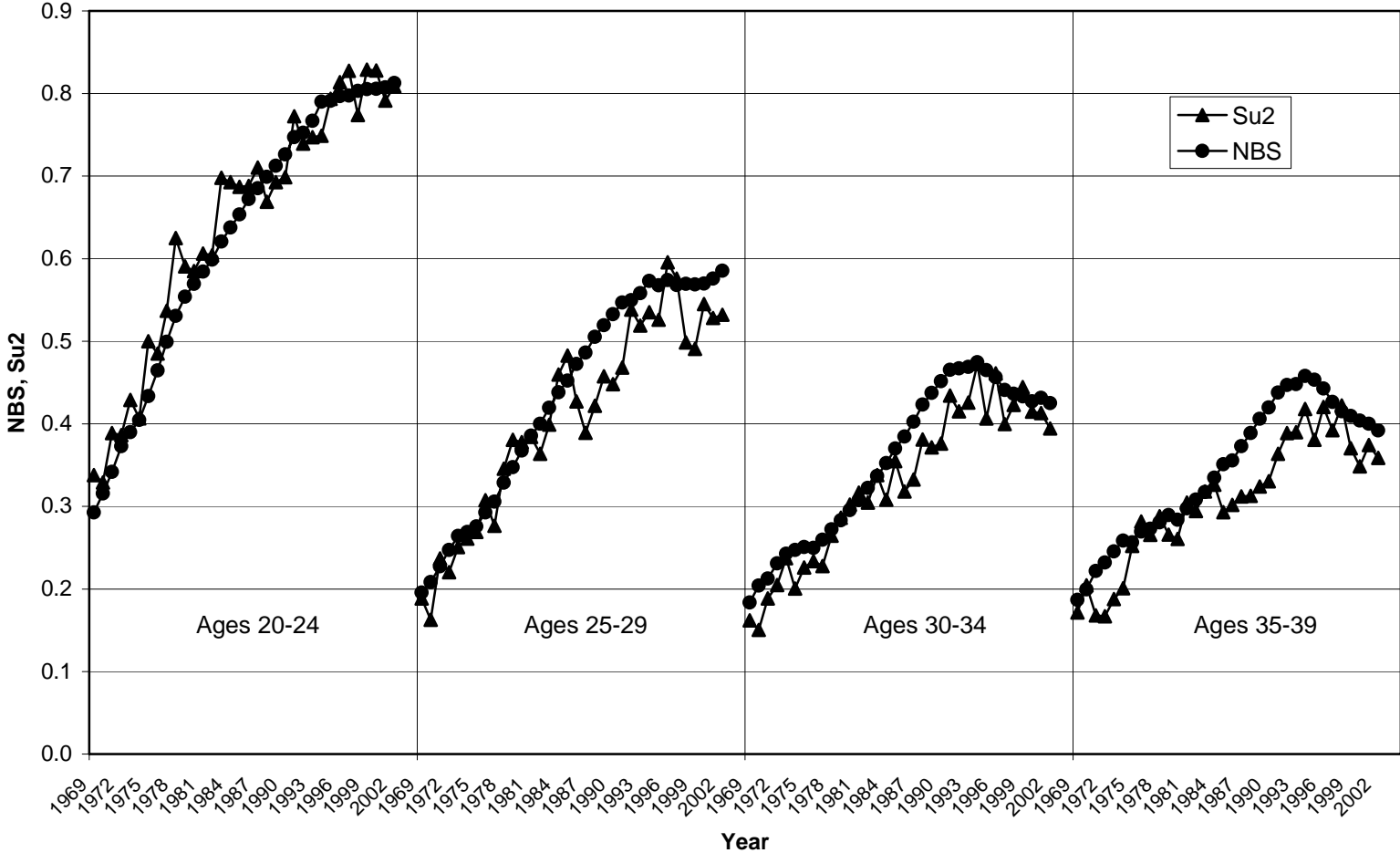


Figure 5. MBRR and (1+Su) by Age, White Women 20-39

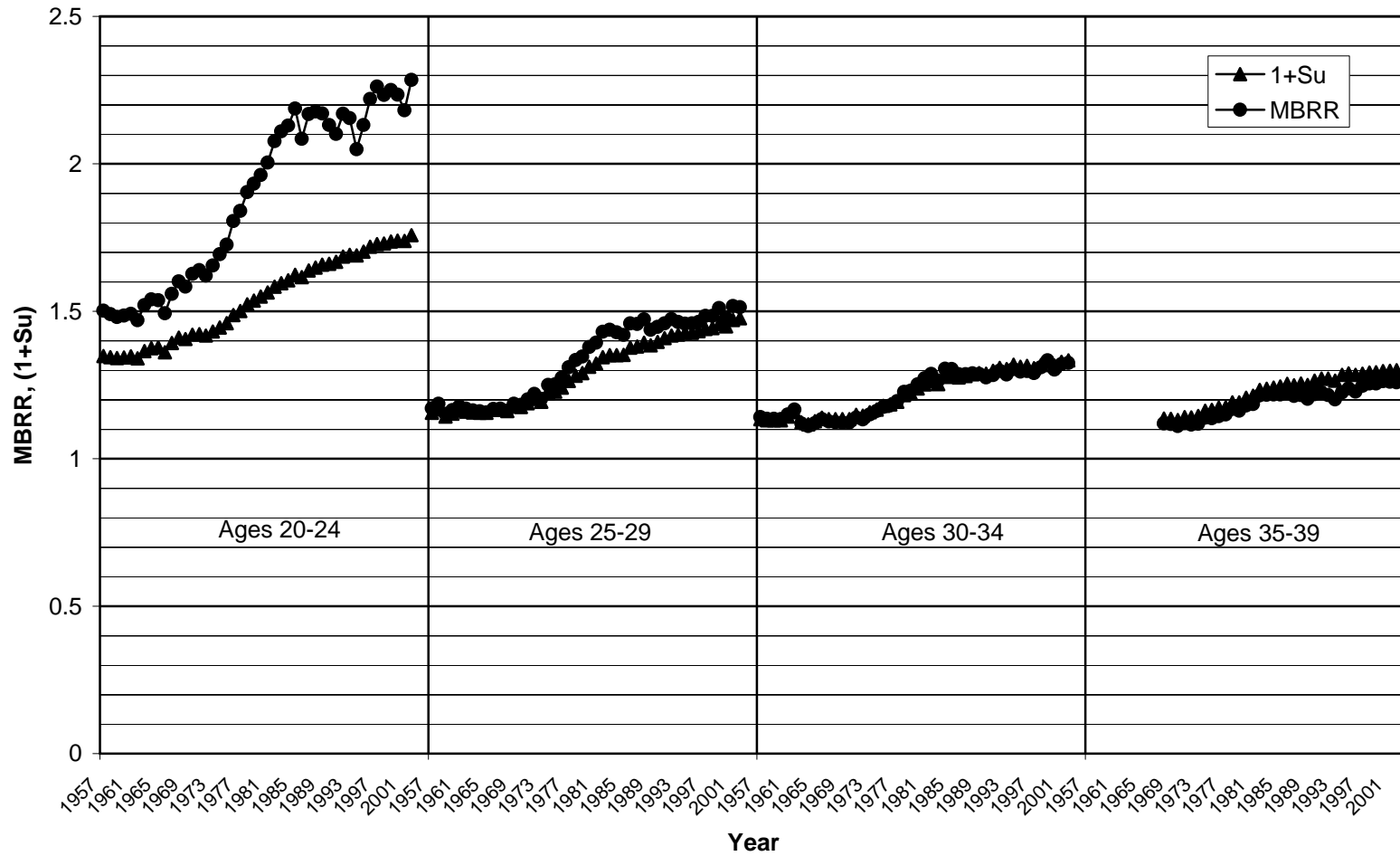


Figure 6. UBRR and Su by Age, White Women 20-39

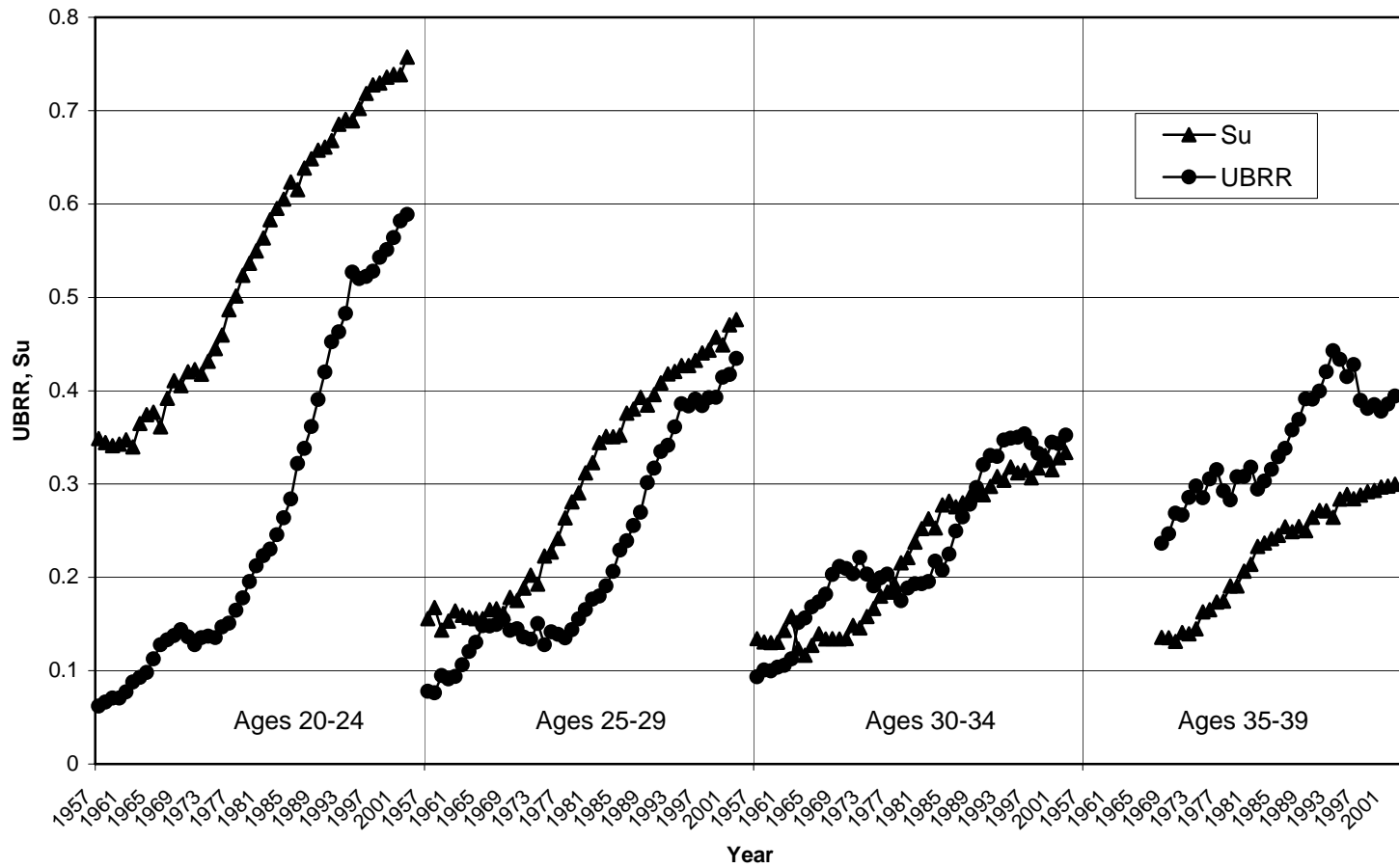


Figure 7. NBS and Su2 by Age, White Women 20-39

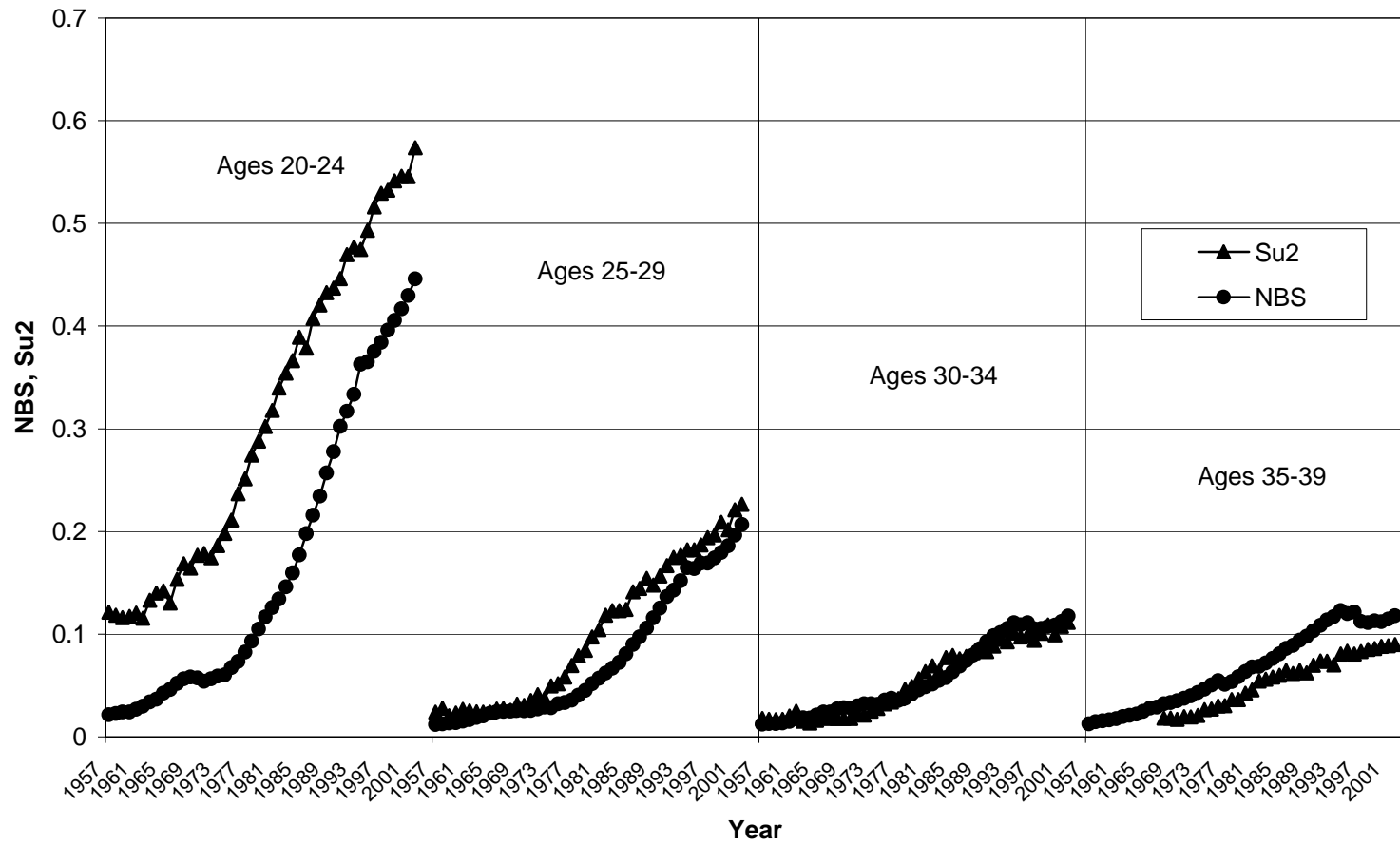


Figure 8. Cumulative Births, Ages 20-39

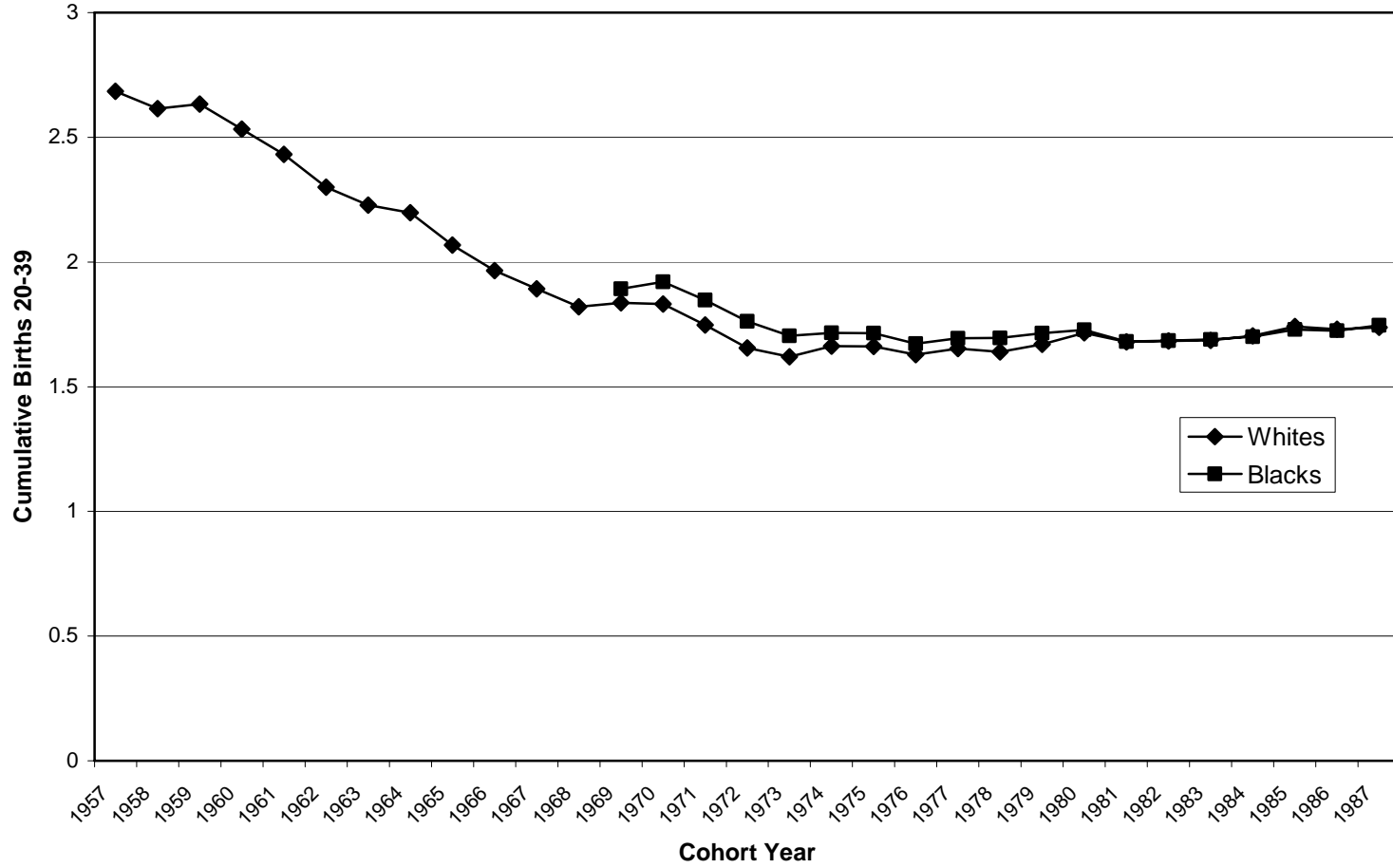


Figure 9. Ratio of TBR for Ages 20-24 to TBR for Ages 25-29

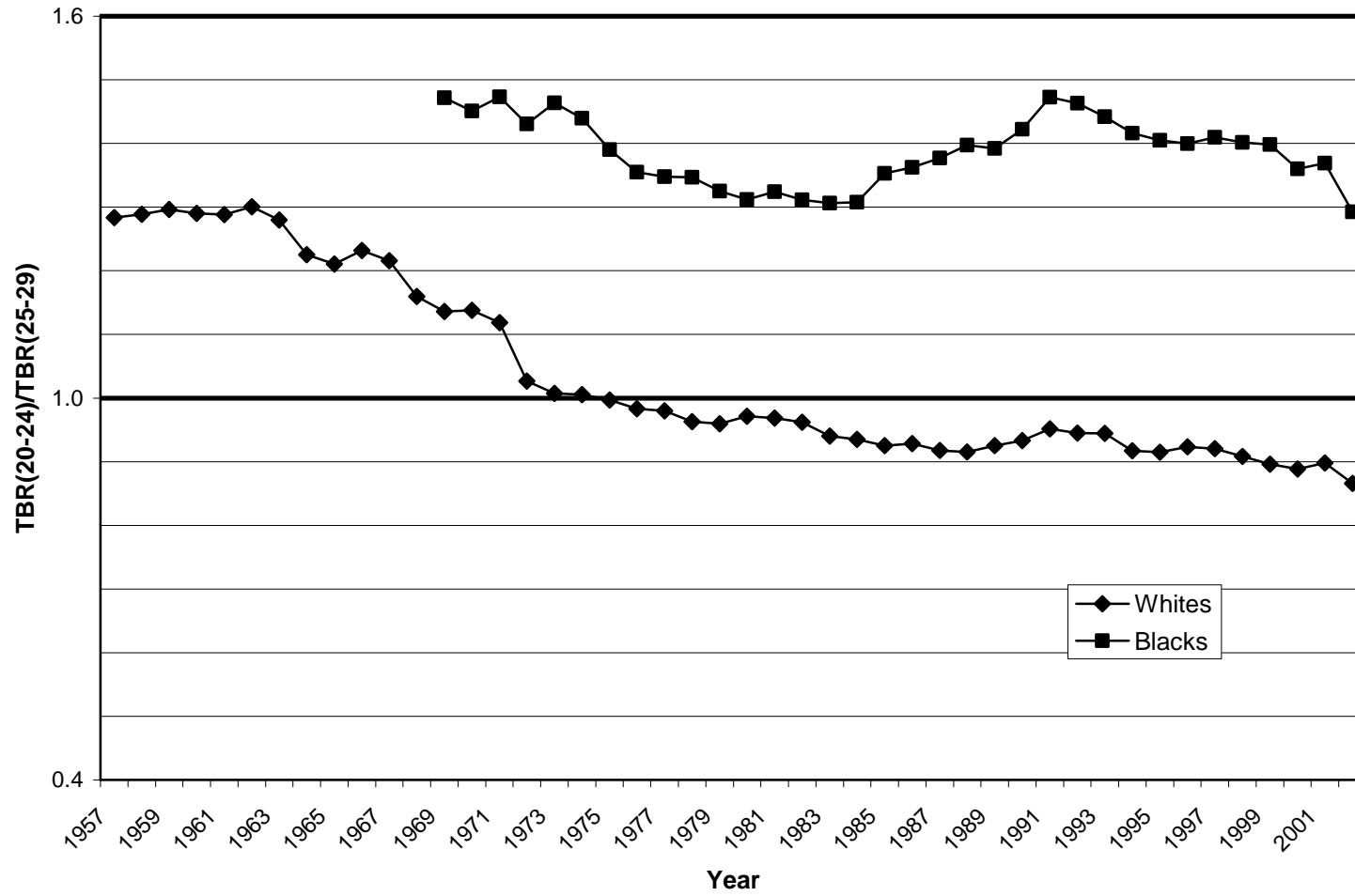
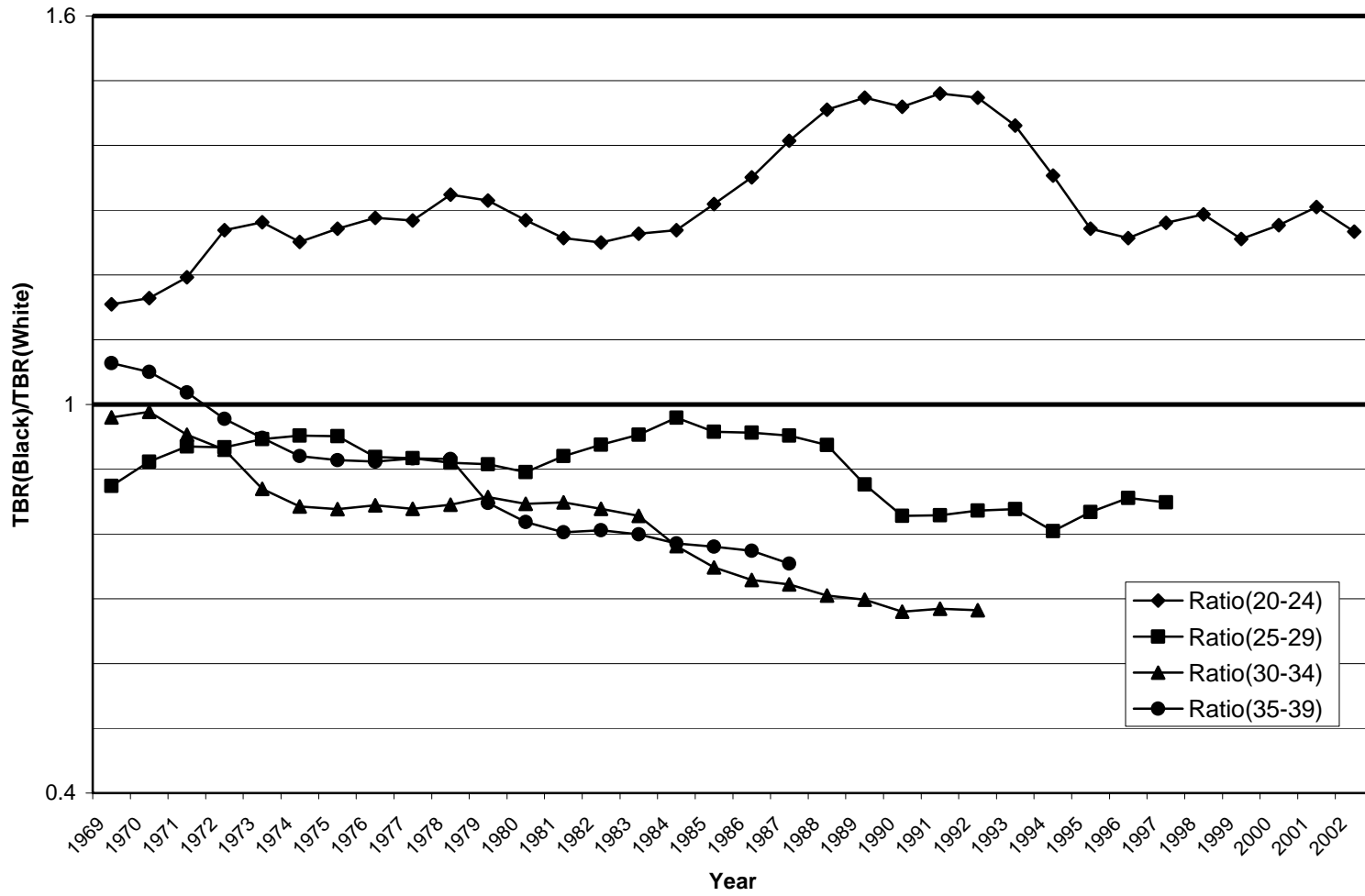


Figure 10. Ratio of Black to White Age-Specific Birth Rate



Appendix

Women who do not delay

Let T be total number of women in cohort t . Since a fraction α of these women choose to delay childbearing, the number of women who do *not* delay (and therefore begin giving birth in the first period of life) is $(1-\alpha)T$. The childbearing and marriage behavior of non-delay women conform to the baseline behavior described in sections II.B. and II.C. of the text. We identify by U^N the number of non-delay women who choose to be single, and by M^N the number who choose to marry in the first period of life. Since it is the first U^N ordered women in the set of $(1-\alpha)T$ non-delay women who are single, and the remainder who marry, setting $n=U^N$ and $z=(1-\alpha)T$ in propositions (P1) and (P2) produces the non-marital and marital life-time births rates of these women. Age-specific birth rates are produced by multiplying the resulting lifetime birth rates by the θ_i , which determine the timing of births in the baseline case.

Recognizing that $(1-\alpha)T$ is equal to (U^N+M^N) then produces

$$UBR_i^N = \theta_i(1/2)(Su^N)P, \quad \text{where } Su^N = [U^N/(M^N+U^N)], \quad (A1)$$

$$MBR_i^N = \theta_i(1/2)(1+Su^N)P, \quad (A2)$$

$$TBR_i^N = \theta_i(1/2)P, \quad \text{for } i = 1, 2, 3, 4. \quad (A3)$$

Note that Su^N is the share of single women in the non-delay group, which will differ from the single share for the cohort as a whole during the first period.

Women who delay

For the αT women who decide to delay childbearing, the number of births in the first period of life is zero. Because the benefits of marriage are presumed to depend on the *presence* of children, we assume that none of these women marry in the first period of life. Accordingly, the non-marital and total birth rates of this group are zero in the first period of life. Since there are no married women among this group in the first period, the married birth rate is undefined.

As described earlier, the women who delay childbearing do not reduce lifetime births, but redistribute childbearing over the remaining three childbearing periods, resulting in age-specific birth

rates equal to lifetime birth rates multiplied by $\theta_i/(1-\theta_i)$ for $i=2, 3$, and 4 . The marriage criterion of the delaying women in a cohort are no different from those who do not delay – the marriages just take place a period later. Denote by U^D the number in this group who do not marry in the second period (or later) and by M^D the number who do marry, where $U^D+M^D=\alpha T$. For the delaying group of women, then,

$$UBR_i^D = [\theta_i/(1-\theta_i)](1/2)(Su^D)P, \quad \text{where } Su^D = [U^D/(M^D+U^D)], \quad (A4)$$

$$MBR_i^D = [\theta_i/(1-\theta_i)](1/2)(1+Su^D)P, \quad (A5)$$

$$TBR_i^D = [\theta_i/(1-\theta_i)](1/2)P, \quad \text{for } i = 2, 3, 4. \quad (A6)$$

Putting it Together

1. First Period Analysis

Recall that there are no first period births or marriages among women who delay childbearing. It follows that the number of births to unmarried women in cohort τ during the first period of life can be expressed as the product of two terms. The first is the number of women who choose not to delay childbearing in the first period and also choose to be single, or U^N . The second term is the first-period birth rate of this sub-population, UBR_1^N . The number of non-delay women in cohort τ is $(1-\alpha)T$ and so the number of those women who do not marry is equal to $Su^N(1-\alpha)T$. The first-period birth rate of this group is given by equation (A1). So the total number of births in the first period (all of which occur among the non-delay population) is

$$UB_1^N = U^N UBR_1^N = \theta_1 Su^N (1-\alpha)T(1/2)(Su^N)P = \theta_1 (Su^N)^2 (1-\alpha)T(P/2).$$

Similarly, the number of marital births to the cohort during the first period is given by

$$MB_1^N = M^N MBR_1^N = \theta_1 (1-Su^N)(1-\alpha)T(1/2)(1+Su^N)P = \theta_1 [1-(Su^N)^2](1-\alpha)T(P/2).$$

Total births, the sum of UB_1^N and MB_1^N , are

$$TB_1^N = \theta_1 (1-\alpha)T(P/2).$$

To obtain the first-period unmarried birth rate for the cohort as a whole, UBR_1 (no superscript), divide the number of first-period unmarried births in the cohort by the total number of women in the cohort who are unmarried in the first period. First-period unmarried births, due entirely to the non-delay

group of women, are UB_1^N , given above. The number of single women in the cohort is the sum of the number of non-delay women who do not marry, $Su^N(1-\alpha)T$, and the number of delaying women (none of whom marry), αT . Accordingly, the first-period unmarried birth rate for cohort τ can be expressed as

$$\begin{aligned} UBR_1 &= \theta_1(Su^N)^2(1-\alpha)T(P/2)/[Su^N(1-\alpha)T + \alpha T] \\ &= \theta_1(Su^N)^2(1-\alpha)(P/2)/[Su^N(1-\alpha) + \alpha]. \end{aligned} \quad (A7)$$

Similarly, the first-period married and total birth rates are

$$\begin{aligned} MBR_1 &= \theta_1[1-(Su^N)^2](1-\alpha)(P/2)/(1-Su^N)(1-\alpha) \\ &= \theta_1[1+(Su^N)](P/2) \end{aligned} \quad (A8)$$

$$TBR_1 = \theta_1(1-\alpha)(P/2) \quad (A9)$$

Finally, denote the first-period single share of the combined (delay and non-delay) population by Su_1 (different from Su^N). The share Su_1 , which is the share observed in the data, is equal to the total number of women in the cohort who are unmarried in the first period (derived earlier) divided by the size of the cohort, or

$$Su_1 = [Su^N(1-\alpha)T + \alpha T]/T = Su^N(1-\alpha) + \alpha$$

Solving for Su^N in terms of Su_1 gives

$$Su^N = (Su_1 - \alpha)/(1-\alpha) \quad (A10)$$

Substituting (A10) into (A7) through (A9) allows us to write the first-period birth rates in terms of the *observed* first-period share of unmarried women in a cohort.

$$UBR_1 = \theta_1(P/2)(Su_1 - \alpha)^2/(1-\alpha)Su_1 \quad (A11)$$

$$MBR_1 = \theta_1(1-2\alpha+Su_1)(P/2)/(1-\alpha) \quad (A12)$$

$$TBR_1 = \theta_1(1-\alpha)(P/2) \quad (A13)$$

2. Analysis for the Remaining Periods

For the remaining three childbearing periods, the birth rates for non-delay women are provided in (A1) through (A3), and those for the delay group in (A4) through (A6). The unmarried share of the non-delay women is unchanged at its first-period value, Su^N . However, some of the delay group, none of whom married in the first period of life, will marry in the second period of life when they begin to have

their children. Since the benefits of marriage depend on lifetime births, and these women exhibit the same distribution of γ as those who did not delay, they will marry in the same proportion as the non-delay women – that is, $Su^D = Su^N$. Accordingly, the share of unmarried women for the cohort as a whole, Su_i , and the unmarried shares of the non-delay and delaying portions of the cohort, are the same for $i=2, 3$ and 4. We denote the common value by Su .

From the birth rates and the numbers of unmarried, married, and total women in the two groups of women – delay and non-delay – the unmarried, married, and total births of each group for the last three childbearing periods can be calculated:

$$UB_i^N = U^N UBR_i^N = Su(1-\alpha)T\theta_i(1/2)SuP = \theta_i Su^2(1-\alpha)T(P/2),$$

$$MB_i^N = M^N MBR_i^N = (1-Su)(1-\alpha)T\theta_i(1/2)(1+Su)P = \theta_i(1-Su^2)(1-\alpha)T(P/2),$$

$$TB_i^N = \theta_i(1-\alpha)T(P/2),$$

$$UB_i^D = U^D UBR_i^D = Su\alpha T[\theta_i/(1-\theta_i)](1/2)SuP \\ = [\theta_i/(1-\theta_i)]Su^2\alpha T(P/2),$$

$$MB_i^D = M^D MBR_i^D = (1-Su)\alpha T[\theta_i/(1-\theta_i)](1/2)(1+Su)P \\ = (1-Su^2)\alpha T(P/2),$$

$$TB_i^D = [\theta_i/(1-\theta_i)]\alpha T(P/2), \text{ for } i = 2, 3, \text{ and } 4.$$

Summing UB_i^N and UB_i^D and dividing by the sum of U^N and U^D gives the per-period unmarried birth rates for the cohort as a whole, UBR_i , as given by equation (15) of the text, repeated here for convenience.

Similar calculations produce MBR_i and TBR_i , as given in equations (16) and (17) of the text:

$$UBR_i = \theta_i(P/2)Su[1+\alpha\theta_i/(1-\theta_i)] \tag{A15}$$

$$MBR_i = \theta_i(P/2)(1+Su)[1+\alpha\theta_i/(1-\theta_i)] \tag{A16}$$

$$TBR_i = \theta_i(P/2)[1+\alpha\theta_i/(1-\theta_i)], \text{ for } i = 2, 3, \text{ and } 4. \tag{A17}$$