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A tale of two shares: The relationship between the "illegitimacy" ratio and the marriage share

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Abstract

Our model implies a magnified effect of marriage rates on the illegitimacy ratio. For U.S. data, plots and regression estimates support the prediction that the share of unmarried births is driven by the *square* of the share of unmarried women.

Keywords: Fertility; Marriage; Non-marital births

JEL classification: J10; I38

1. Introduction

In recent decades both the proportion of women who are unmarried and the share of births to unmarried women have risen sharply. (While the latter is often referred to as the illegitimacy ratio, we refer to it below as the unmarried fertility share, or UFS.) Proportional increases in the unmarried fertility share (UFS) are an expected consequence of declines in the population share of unmarried women. However, recent increases in UFS have far exceeded increases in the share of unmarried women and have been the central focus of a vast literature, particularly regarding the effects on child-bearing behavior of public policies such as the now displaced Aid to Families with Dependent Children and the continuing Earned Income Tax Credit (e.g. Baughman and Dickert-Conlin, 2003).

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Demographic studies of UFS (e.g., Smith et al., 1996) typically focus on three factors — the share of unmarried women, the unmarried birth rate, and the married birth rate. These studies implicitly assume that the three factors independently influence changes in UFS, and so suggest a proportional relationship between UFS and the share of unmarried women, holding constant child-bearing behavior. By contrast, we propose in Section 2 a simple model in which marriage behavior causes changes in unmarried and married birth rates without necessarily causing changes in child-bearing behavior. The model implies that UFS is driven by the *square* of the population share of unmarried women.

Section 3 explores the ability of the model to explain observed changes in UFS, i.e., the extent to which recent disproportionate changes in UFS may be due to the effects of changes in marriage behavior on measured birth rates rather than changes in child-bearing behavior. Our findings support the model and underscore the importance of studies that look simultaneously at fertility and marriage (e.g., Grogger and Bronars, 2001) and find direct effects of policies on marital status.

2. The idea

The theory illustrates a simple idea: increases in the share of unmarried women are produced by 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 these women leave the pool of married women and enter the pool of unmarried women, average birth rates of *both* groups rise, even though the total number of births (and the total birth rate) may not change. Furthermore, the *proportional* increase in the unmarried birth rate exceeds that of the married birth rate, producing a magnified effect of changes in the share of unmarried women on UFS.

2.1. Definitions

The unmarried fertility share is defined as

$$\mathrm{UFS} = \mathrm{UB}/(\mathrm{MB} + \mathrm{UB}),$$

where

MB = number of births to married women, and
UB = number of births to unmarried women.

UFS can be rewritten as

$$\mathrm{UFS} = [\mathrm{UB}/(\mathrm{MB} + \mathrm{UB})][(\mathrm{M} + \mathrm{U})/\mathrm{U}][\mathrm{U}/(\mathrm{M} + \mathrm{U})] = (\mathrm{UBR}/\mathrm{TBR})\mathrm{Su} \qquad (1)$$

where

M = number of married women,
U = number of unmarried women,
UBR = UB/U = the unmarried birth rate,
TBR = (MB+UB)/(M+U) = the total birth rate,
Su = U/(M+U) = the fraction of women not married.

UFS is proportional to Su, provided the birth rate ratio UBR/TBR is unchanged — the standard demographic approach.[1]

2.2. Assumptions and implications

Women vary in their preference for children, captured by a parameter, γ, that measures the probability of a particular woman giving birth during the observation period (e.g. a year). We assume that γ is independent of marital status and distributed uniformly on the interval $[0, P]$, where $0 \leq P \leq 1$. Women are ordered and indexed by γ. Implications:

- the average birth rate across all women is $(1/2)P$, or TBR=$(1/2)P$.
- the γ associated with the nth ordered woman in a total population of z women is $(n/z)P$.
- the average birth rate of the first n ordered women is $(1/2)(n/z)P$.

We further assume that the net benefits to marriage are increasing in a woman's preference for children, captured by γ, and decreasing in a fixed cost, C, common to all women. Implications: there exists a critical value of γ, γ_c, such that

– all women with a $\gamma > \gamma_c$ marry.
– all women with a $\gamma < \gamma_c$ do not marry.

- The critical value γ_c is increasing in C.

Since women are ordered by γ, the first U ordered women will be unmarried and the remaining M women will be married. It follows that:

- UBR, the average birth rate of unmarried women, is $(1/2)(\mathrm{Su})P$.

After substitution for UBR and TBR, Eq. (1) yields the key prediction of the model:

$$\mathrm{UFS} = \mathrm{Su}^2, \text{ or } \log\mathrm{UFS} = 2 \log\mathrm{Su}. \tag{2}$$

That is, UFS is simply the squared value of Su, implying a log–linear relationship with a coefficient of 2. Importantly, the "taste" for children, captured by the parameter P, does not appear in Eq. (2). While UBR and TBR both increase with P, it is their *ratio*, which does *not* depend on P, that determines UFS.

2.3. Discussion

A woman's marital status depends on the net benefits to marriage, determined in part by her child-bearing propensity, γ. Married women occupy the upper end of the uniform distribution that describes γ, and unmarried women the lower end. If Su rises due to, for example, changes in labor market

[1] While the model could be adapted to deal with co-habitation and other variations, we presume here an unambiguous and unchanging definition of marriage.

opportunities, married women with the lowest child-bearing propensities shift from married to unmarried.[2] Accordingly, both married and unmarried birth rates rise, even if P, and therefore the total birth rate, does not. In addition, UBR (with a smaller initial value than MBR) rises proportionately more than MBR, producing a rise in UBR relative to both MBR and TBR. (Indeed, UBR/TBR is *equal* to Su.).

Referring to Eq. (1), it is evident that an increase in Su raises UFS both because of its direct effect on UFS and because it raises UBR/TBR. Thus, the impact of an increase in Su on the unmarried fertility share is larger than a typical demographic calculation would suggest: UFS is a power function (the square) of Su. Furthermore, because USF is independent of P, many factors that one might expect to influence both marriage and child-bearing behavior exert their influence on UFS *only* through Su.

3. Data and evidence

Our age-specific data for births to and numbers of married and unmarried women are available separately for women in the United States annually beginning in 1957 for whites and 1969 for blacks, with data for both extending through 2000.[3] We focus on the prime adult child-bearing years of 20–44 and construct composite measures of Su and UFS by weighting each age cohort in the data by its average share of the total over the period.[4] This fixed-weight composite avoids shifts induced solely by changes in the age distribution of the population of women.

Figs. 1 and 2 plot log UFS against log Su for white and black women, respectively. The data track our model's prediction of a log–linear relationship with a slope of 2 quite closely. Furthermore, they are clearly inconsistent with the alternative hypothesis of a slope of 1. While the data appear to suggest non-linearities early in the period for white women, when values of UFS and Su are quite low (i.e., when the log values are most negative), higher order terms do not enter significantly in the regressions below. Overall, the model's prediction that log UFS equals 2 times log Su holds surprisingly well.

Table 1 presents estimates of the log–linear form of Eq. (2) separately for white and black women (with a predicted coefficient of 2). Ordinary least squares (OLS) estimates evidence high measures of overall fit. As one might expect, the OLS estimates exhibit substantial first-order autocorrelation in the residuals, 0.88 for whites and 0.46 for blacks. Accordingly, we also present AR1 estimates adjusted for the estimated autocorrelation.[5]

[2] Cready et al. (1997), Lichter et al. (1991), Moffitt (2000), and others, find evidence that increases in the female wage relative to the male wage lead to lower marriage rates, as implied by models of marriage in the tradition of Becker (1973), which emphasize the gains from specialization. Moffitt, in particular, finds that women's rising relative wages *over time* are important in explaining increases in the population share of unmarried women *over time*. The female relative wage is, of course, just one of many factors thought to affect marriage decisions. Other widely studied factors include welfare eligibility and benefits, male wages, the unemployment rates of both genders, the sex ratio, access to birth control and abortion, and changing social norms. Informative surveys include Moffitt (1992, 1998) and Fitzgerald and Ribar (2004).

[3] Data on unmarried births are from National Vital Statistics Reports (2000, 48:16 and 2002, 50:10). Data on total births are from Vital Statistics of the United States (www.cdc.gov/nchs/births.htm). Data for the numbers of married and unmarried women are from U.S. Bureau of Census, Current Population Reports, Series P-20, various dates.

[4] Age cohorts are 20–24, 25–29, 30–34, 35–44.

[5] In formal time-series tests, a unit root cannot be rejected for either series for whites or blacks. For whites, a significant co-integrating equation yields a coefficient of 1.85, which is significantly different from one but not two. For blacks, there is no significant co-integrating equation, likely due to low power. We note, however, that the most significant equation yields a coefficient of 2.18.

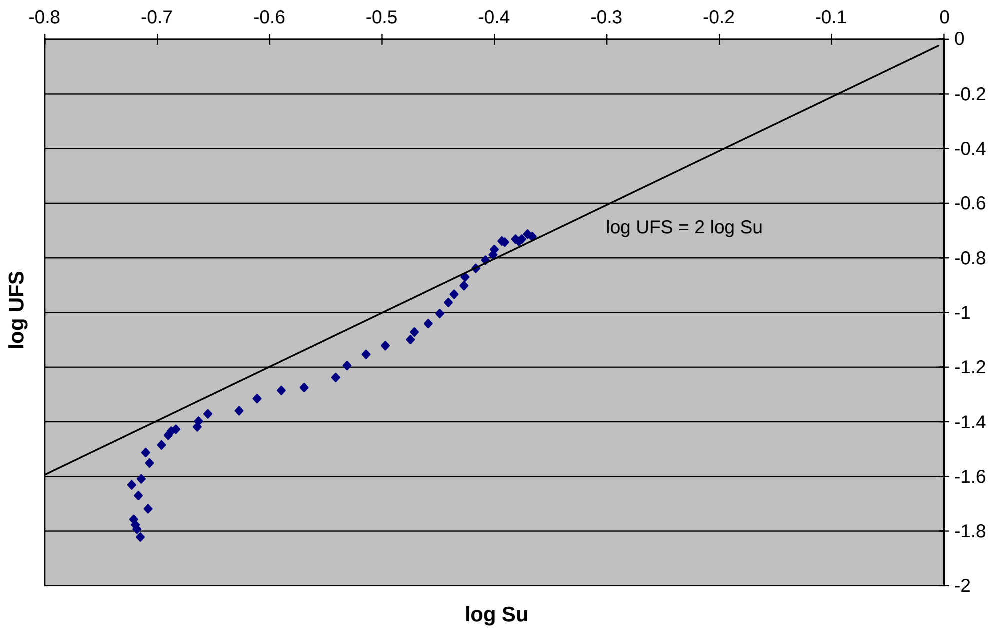

Fig. 1. White women 20–44, 1957–2000.

For white women, the slope estimates are 2.59 and 1.94, respectively, for the OLS and AR1 specifications. In the relevant AR1 specification, the point estimate is significantly different from 1, the slope predicted by the traditional demographic model (Eq. (1)), yet insignificantly different from 2, the slope predicted by our

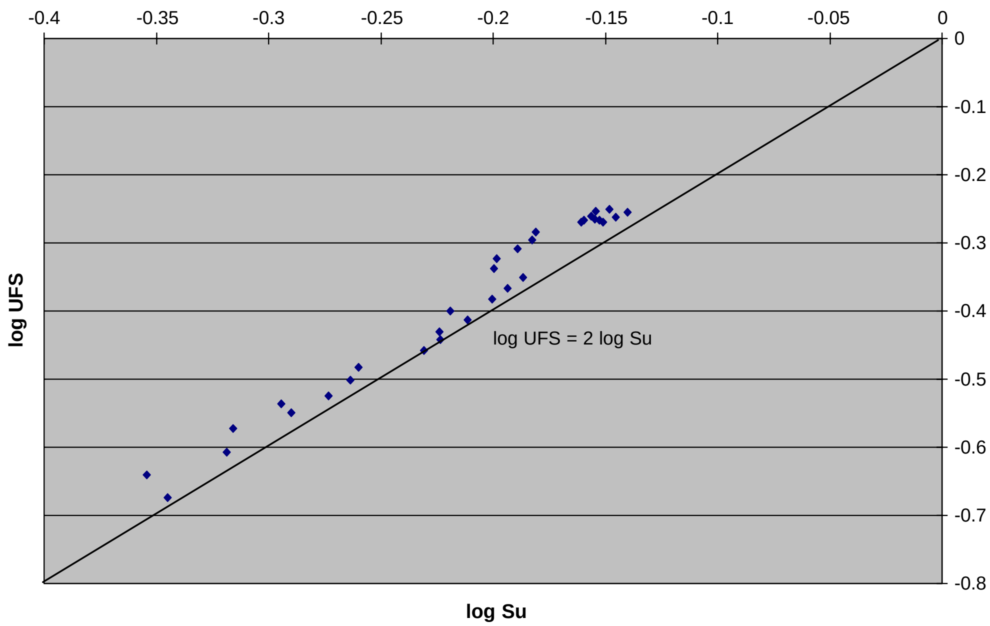

Fig. 2. Black women 20–44, 1969–2000.

Table 1
Regression estimates for log UFS[+]

	White women 20–44, 1957–2000		Black women 20–44, 1969–2000	
	OLS	AR1	OLS	AR1
Constant	0.223**	−0.007	0.047**	0.020**
	(0.050)	(0.011)	(0.014)	(0.008)
log Su	2.590**	1.939**	2.038**	1.938**
	(0.088)	(0.174)	(0.060)	(0.066)
t-statistic for null of:				
Slope = 1	18.068**	5.397**	17.300**	14.212**
Slope = 2	6.705**	−0.351	0.633	−0.939
R-squared	0.954	0.866[++]	0.987	0.966[++]
Rho	–	0.880	–	0.460

[+] Standard errors in parenthesis.
[++] AR1 *R*-squares are for transformed data.
** Significant at 5% level.

model. For black women, the estimates are 2.04 and 1.94, respectively. In the AR1 specification, the point estimate is significantly different from 1, but not from the predicted value of 2. Furthermore, despite the stark simplicity of our model, the intercepts are either insignificant or relatively small.

One might speculate on the possibility of a variety of omitted variables. Recall, however, that factors determining the preferences for children are already subsumed, if simplistically, within our model. Given Su, for example, UFS is invariant to changes in P, the maximum value of γ. Given the striking correspondence between the predicted and actual relationships between log UFS and log Su evident in Figs. 1 and 2 and the nearly perfect match of the estimated coefficients in Table 1 to their predicted values, the model appears to have substantial empirical power.

4. Concluding remarks

Our simple, but joint, model of fertility and marriage yields a sharp result — the share of births to unmarried women is driven by the square of the share of unmarried women, whether or not individual fertility behavior is changing. Plots and regression estimates yield strong support for this prediction of the model. Our findings suggest that some of the emphasis on changes in fertility behavior in explaining the rising share of births to unmarried women might be productively redirected toward exploring the role and determinants of changes in marriage behavior. Moreover, previous studies of fertility behavior, to the extent that marital status is taken as given, may confound fertility and marriage behavior. Of course, our model is simple; our data and tests perhaps more suggestive than definitive. It would be surprising, though, if more elaborated theoretical models or more detailed data entirely upset the strong evidence found here.

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