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The Impact of Child Support Enforcement Policy on Nonmarital Childbearing

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Abstract

A simple model of fatherhood and marriage choice implies that stricter child support enforcement will tend to reduce nonmarital childbearing by raising the costs of fatherhood. We investigate this hypothesis by examining nonmarital childbearing during 1980–1993, a period when child support policy and enforcement underwent enormous changes. We use a sample of women from the Panel Study of Income Dynamics, to which we add information on state child support enforcement. We examine childbearing behavior between the ages of 15 and 44, both before marriage and during periods of nonmarriage following divorce or widowhood. Discrete-time hazard models of nonmarital childbearing provide evidence that women living in states with more effective child support collection were less likely to bear children when unmarried. The findings suggest that policies that shift more costs of nonmarital childbearing to men may reduce this behavior.

The Impact of Child Support Enforcement Policy on Nonmarital Childbearing

Births outside of marriage have grown dramatically over the past three decades in the United States. In 1966, the nonmarital birth ratio—the percentage of births to unmarried mothers relative to all births—was approximately 8 percent. Since 1994, the figure has been between 32 and 33 percent. Among teenagers the ratio is much higher—79 percent in 1999 (Ventura and Bachrach, 2000).

The nonmarital birth rate (number of births per 1,000 unmarried women) has also grown steadily. In 1965 it was 23.4 among all women aged 15–44 and 16.7 among teenagers aged 15–19 (U.S. Department of Health and Human Services, 1995). By 1994 it had risen to 46.9 among all women aged 15–44 and 46.4 among teenagers. The rate declined slightly, to 44.3, by 1998 among women aged 15–44 and to 41.5 among teenagers (Ventura and Bachrach, 2000).

In 1996 Congress enacted new legislation designed to reduce welfare eligibility and increase the costs of single motherhood as part of the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA). State legislation with similar motivation both preceded and followed the PRWORA. These federal and state initiatives have included policies that lower welfare benefits, limit eligibility, impose stronger work requirements on welfare recipients, and place restrictions on benefits to unwed parents under age 18 who do not live with their parents (or in another adult-supervised setting) and attend school. Most of these policies were driven, at least in part, by the assumption that the availability of welfare to unmarried women was a major cause of nonmarital childbearing (Murray, 1984, 1993), despite empirical evidence to the contrary (Moffitt, 1992, 1998).

The asymmetrical focus on women appears unreasonable. Decisions about sexual intercourse and marriage involve two persons rather than one. The same often is true for decisions about contraceptive use and abortion. Yet the research literature, as well as policy debates, has largely failed to recognize and critically analyze men's role in nonmarital childbearing and how government policies may influence men's behavior. The government's poor record of establishing paternity and enforcing payment of child support by nonresident fathers (Sorensen, 1997) may partly be responsible for men's failure to take responsibility for contraception or to marry their sexual partners. Although efforts to establish paternity and enforce child support have intensified during the past decade and were strengthened by the PRWORA, they have generally been viewed as ways of reducing the financial costs of public welfare rather than as strategies for preventing nonmarital births.

This paper presents a simple model of how the incentives of child support policy affect unmarried men's decisions about fatherhood and marriage. The model implies that stronger child support enforcement reduces the likelihood of nonmarital childbearing. We test this hypothesis using data from the Panel Study of Income Dynamics during 1980–1993, a period when child support policy and enforcement underwent enormous changes. We find evidence in support of the hypothesis.

U.S. CHILD SUPPORT POLICIES AND THEIR EXPECTED EFFECTS ON NONMARITAL CHILDBEARING

Until very recently, financial responsibility for children born outside marriage rested primarily with the mother and her family and with government. Mothers who met the income test, which included the vast majority of unwed mothers, were eligible for Aid to Families with Dependent Children (AFDC), Food Stamps, Medicaid, and in many cases housing subsidies. In contrast, unwed fathers were more or less free to shirk their parental obligations, and most did so (Garfinkel, 1992).

During the past quarter century, the federal government has taken a number of steps to prevent unmarried fathers from abandoning their children financially (Garfinkel, McLanahan, and Robins 1994; Garfinkel and McLanahan, 1986, 1994). In 1975, Congress created the Child Support Enforcement Program, which established local offices of child support enforcement and authorized federal matching funds for states to help locate absent parents, establish paternity, establish child support orders, and obtain child support payments (U.S. House of Representatives, 2000, Section 8). The 1984 Child Support Amendments extended this legislation by requiring states to withhold child support obligations from the paychecks of delinquent fathers and to develop legislative guidelines to be used in determining child support awards. In 1988, the Family Support Act mandated that states adopt presumptive guidelines for child support awards and initiate automatic withholding from fathers' paychecks, regardless of delinquency. The Act also included provisions aimed at strengthening paternity establishment for children born to unmarried parents. Reforms in the 1996 PRWORA sought to further improve the child support system's ability to establish paternity for children born outside of marriage, to locate nonresidential fathers, and to collect support payments.

The results of this new legislation have been striking with respect to children born outside marriage. The proportion of never-married mothers with a child support award grew from 12 percent in the early 1980s to over 20 percent in 1994 (Hanson, 1995). Paternity establishment ratios (the number of paternities established in a given year divided by the number of nonmarital births) increased from 20 percent to 46 percent over the same period, and to 64 percent by 1998 (Nichols-Casebolt and Garfinkel, 1991; U.S. House of Representatives 2000, Table 8-22).

A Theoretical Model of How Child Support Policy Affects Men's Fatherhood and Marriage Choices

Administration and enforcement of child support laws raise the likelihood that fathers who do not live with their children will nonetheless be required to make substantial financial contributions over many years to their support. Child support policy, therefore, increases the expected costs of fatherhood for absent fathers.[1] This disincentive would, other things equal, make men more reluctant to father children outside marriage and, if a nonmarital pregnancy occurs, make them more likely to marry before the birth.

For women, improved enforcement that leads to higher child support payments might appear to reduce the costs of children and create more incentive to have children outside marriage. However, given that a large proportion of both teenage and nonteenage women who give birth while unmarried are likely to go on welfare (Duncan and Hoffman, 1990; Foster, Jones, and Hoffman, 1998; Haveman and Wolfe, 1994) and given that welfare policy taxed child support payments during the time period we analyze

[1]For absent fathers who plan to provide support above the level required by child support policy, policy may not affect costs. But for no father does it lower the expected financial costs of fatherhood.

(1980–1993), this countervailing effect is likely to have been small. From 1980 to 1984, a mother on welfare retained none of the child support paid by the absent father. Rather, all payments were used to offset benefits paid from public funds. Between 1985 and 1993, a mother on welfare was allowed to keep only the first $50 of child support each month.[2] All payments above $50 went toward reducing public spending on welfare and did nothing to increase her children's standard of living. Indeed, from the mother's viewpoint stricter child support enforcement may actually have increased the cost of raising children if she had been getting informal support from the father and if that support ended because of stricter enforcement (Waller and Plotnick, 1999). One may reasonably conclude that child support policy during 1980–1993 did little to affect women's incentives regarding nonmarital childbearing.

A simple utility maximization model captures the essence of how the incentives set up by child support policy in the 1980s and early 1990s would affect men's behavior. We focus on male decision making because, as just argued, child support policy left women's incentives largely unaffected.

Consider an unmarried man romantically involved with an unmarried woman, but not cohabiting with her or sharing income to any significant extent. We assume he faces three fatherhood and marriage options. Under the "no-child" option, he avoids fatherhood and does not marry. Under the "marital-child" option a nonmarital pregnancy occurs and he marries (or cohabits with) his partner and becomes a father. He shares his income with both mother and child. Assuming no welfare case has opened, the child support agency does not become involved. In the "nonmarital-child" option, the man becomes a father but neither marries nor cohabits with the mother, while the mother and child go on welfare.[3]

Assume that the man's utility depends upon his personal consumption, consumption of the mother and child (whether living with him or not), and the nonfinancial benefits of parenthood and of

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[2]In response to the 1996 welfare reforms, most states have eliminated this $50 "pass-through."

[3]Another option is nonmarital pregnancy followed by abortion. We assume the man can unilaterally choose to try to avoid fatherhood by use of condoms or vasectomy. If contraceptive failure leads to an unplanned nonmarital pregnancy, he must choose between the other two options. Fatherhood is possible only if his partner agrees or an unplanned pregnancy occurs (and she does not seek an abortion). If she carries to term, we assume her own utility-maximizing calculus led her to expect she would be better off with a child whether he chooses the second or third option. If she refuses to bear children, the effect of child support incentives on the man's behavior is moot.

marriage (or cohabitation). Let C_p = personal consumption, C_k = consumption of the mother and child, Y = his own income, V = the implicit value to the man of being married to his current partner and of parenthood, V' = the sum of the increase in his earnings produced by marriage (Gray, 1997) plus savings from economies of scale from living as a family unit, and G = the welfare benefit received by the mother and child if they do not live with the father. Let K = 1 if he fathers a child, 0 if not; and let M = 1 if he marries, 0 if not. If a child is born, the woman is the primary parent and earns nothing. Ignore, for the moment, the existence of a child support program.

The man chooses among the combinations K = 0, M = 0; K = 1, M = 0 and K = 1, M = 1 to maximize his utility:

(1) $$\text{Max } U(C_p, K^*C_k, M^*V) \text{ s.t.}$$

$$Y + (1-M)^*K^*G + M^*V' = C_p + K^*C_k$$

Given his preferences, the man determines the maximum utility of each combination and makes a global utility-maximizing choice.[4] Without information on those preferences, the model does not predict which option he will choose.

A policy that requires child support payments is easily incorporated into the model. Suppose the father must pay S in child support to the mother and that S < G. Assume that paying S does not affect his own income and that welfare policy taxes child support payments 100 percent, so the mother's welfare benefit falls by S.[5]

The new budget constraint is

(2) $$Y + (1-M)^*K^*G + M^*V' - (1-M)^*K^*S = C_p + K^*C_k$$

Note that under the choices K = 0, M = 0 or K = 1, M = 1, the constraint is the same as in (1) and so is the maximum utility obtainable from both choices. If, however, he chooses K = 1, M = 0, the left-

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[4]By allowing C_k to enter the utility function only if there is a child, we assume he gets no utility from financing consumption of the woman if she is not the mother of his child.

[5]Child support policy does not appear to affect men's earnings (Klawitter 1994; Freeman and Waldfogel, 1998).

hand side of (2) falls by S. The maximum utility obtainable from this choice must fall, and it will fall monotonically in S. Hence, the likelihood that an unmarried man would choose the nonmarital-child option will decrease as S grows.

Figure 1 illustrates the budget constraints under each choice. The vertical axis measures the man's personal consumption; the horizontal one measures consumption of the mother and child. Under the no-child option, all income is devoted to personal consumption. The constraint collapses to a point on the vertical axis equal to his income, Y. (The line with slope –1 starting at Y is for reference.)

To analyze the marital-child option, we draw the constraint under the assumption that V + V' is positive.[6] To show a situation where the marital-child option does not necessarily provide more utility than the no-child option, the figure assumes the man's consumption must be less than Y. Then the marital-child option's constraint starts at Z and has slope of –1. Because mother and child cannot consume from V, the constraint does not extend to the horizontal axis.

Under the nonmarital-child option in the absence of a child support system, the mother and child receive a welfare benefit of G, while V = V' = 0. If the father contributes nothing to their support, he consumes Y. Recall that the mother earns nothing. The budget constraint therefore starts at point W. If the parents agree, the father may provide under-the-table support payments to the mother and his child. In that case, his personal consumption declines by one dollar for each dollar of support. Again the constraint has slope of –1. We draw the constraint under the assumption G > V + V' so that the marital-child option does not necessarily provide more utility than the nonmarital-child option.[7]

To incorporate child support payments into the figure, suppose the father must pay S in child support and that S < G. Under the assumption that S does not affect his own income, his maximum personal consumption shifts down to Y – S. Given that welfare policy taxes child support payments 100

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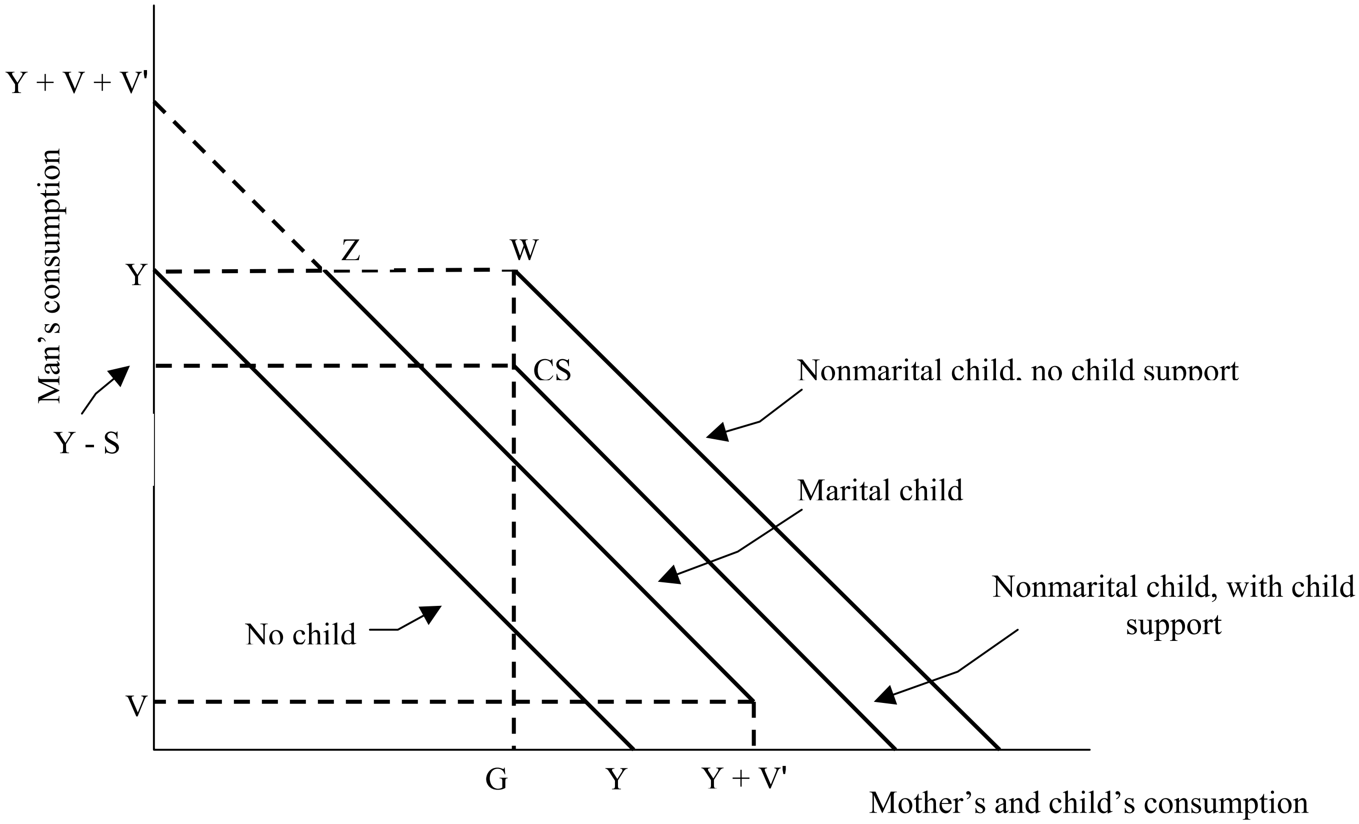
[6]If V + V' ≤ 0 the man would choose the marital-child option only if the value of parenthood (the utility contributed by C_k) at least compensated for the loss of utility due to V + V'.

[7]If G < V + V' the man would never choose the nonmarital-child option. This will be the case for some men. To illustrate where choice may be affected by child support policy, the figure assumes G > V + V'.

FIGURE 1
Budget Constraints for Different Fatherhood and Marriage Options



percent, the income of the mother and child remains constant. In the unlikely event that he so desires, the father can also provide under-the-table support payments. Hence, in a world with child support, the nonmarital-child constraint starts at CS and again has slope of –1.

Figure 1 makes clear the basis for hypothesizing that better child support enforcement will reduce nonmarital births. Forced payment of S reduces the maximum utility from choosing to locate on the nonmarital-child constraint. Higher values of S lead to higher reductions in maximum utility. But since child support policy does not change the no-child and marital-child constraints, the maximum utility available from these choices does not change. Other things equal, better enforcement—higher S—makes it more likely the man will choose to avoid unwed fatherhood.[8]

Besides raising S, stronger child support enforcement may take the form of raising the probability that an absent father will actually pay the award S as established by state guidelines and courts. If we replace S by its expected value, conclusions from the model remains the same.

Related Empirical Research

There is little research on the effects of child support enforcement on nonresident fathers' behavior. A few researchers have studied the association between child support and father-child contact (Seltzer, Schaeffer, and Charng, 1989; Seltzer, McLanahan, and Hanson, 1998), fathers' remarriage (Bloom, Conrad, and Miller, 1998), and fathers' earnings (Klawitter, 1994; Freeman and Waldfogel, 1998).

Four papers are particularly relevant to our argument that child support policy is likely to have empirically significant effects on nonmarital childbearing. Sonenstein, Pleck, and Ku (1994) find that a substantial proportion of adolescent males are aware of paternity establishment and may modify their sexual behavior and contraceptive use accordingly, especially if their peers are doing so. Case's (1998) analysis of state data reports that, net of economic and demographic conditions, states that adopted

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[8]Though incidental to the focus of this analysis, the figure also shows clearly that an increase in G will make the nonmarital-child option more attractive to the man.

presumptive guidelines for setting child support awards or allowed establishment of paternity up to age 18 had lower out-of-wedlock birth rates.

Garfinkel et al. (2002) also analyze state-level data and find that effective child support enforcement deters nonmarital births. The effect is relatively robust. Rather than getting weaker as more controls are added, the effect of child support emerges after the authors control for state and year fixed effects and become even stronger when state-specific time trends are included.

Plotnick et al. (forthcoming) is the only micro-data study that examines the effects of child support enforcement on nonmarital childbearing. It uses the National Longitudinal Survey of Youth (NLSY) to analyze the likelihood that a woman's first birth is both premarital and occurs while she is a teenager. It finds that women living in states with higher rates of paternity establishment are less likely to become unwed teenage mothers. Because of the structure of the NLSY and the focus on teenage behavior, the study only examines behavior during 1979–1984.

These papers' findings are all consistent with the argument that the deterrent effect on men will tend to be larger than any positive effect on fertility among unmarried women and with the hypothesis that increasing the costs of children for nonresident fathers will lower the incidence of nonmarital childbearing.

Indirect support for the argument that the incentives of child support policy may affect nonmarital childbearing comes from evidence that the incentives of other public policies affect this behavior. Lundberg and Plotnick (1995) find that family planning policies and availability affect the likelihood that women avoid premarital pregnancies, while abortion policies and availability of abortion services affect the likelihood that pregnant unmarried women carry their pregnancies to term. Welfare benefits largely limited to single mothers make lone motherhood more affordable and reduce the gains from marriage. Empirical evidence suggests that these incentives undermine marriage and promote nonmarital childbearing, though the importance of these effects remains uncertain (Blackburn, 2000; Moffitt, 1998).

A study by Nixon (1997) is also indirectly relevant. She finds that states with stricter child support enforcement regimes had lower rates of marital dissolution among families with children. By demonstrating that child support policies can affect a personal demographic choice such as marital status, this article suggests that such policies may also affect other personal demographic outcomes, such as premarital childbearing.

The empirical analysis we present extends Plotnick et al. (forthcoming) in several ways. It examines outcomes during 1980–1993, a longer time period with considerably more between- and within-state variation in support enforcement variables. During these years child support enforcement became much more prominent on the social welfare policy agenda. The sampling frame of the PSID allows us to examine nonmarital childbearing over the full age range often used in fertility studies, 15 to 44. Analyses of only the teen years omit most nonmarital births, which occur to women in their twenties. Instead of focusing on first premarital births, we analyze all nonmarital childbearing, whether it occurs before first marriage or during periods of nonmarriage following divorce or widowhood, and regardless of parity order.

STATISTICAL MODELS, DATA, AND EXPLANATORY VARIABLES

Because family law is ultimately a state responsibility, state programs for child support enforcement vary widely. When the federal government began pushing child support enforcement reforms in the early 1980s, some states were already relatively effective in paternity establishment, but most were not. Nearly all have improved their records—some dramatically, others not as much. This study exploits the varying vigor and commitment with which states have implemented their support enforcement programs to test the model's predictions.

We test the predictions by constructing spells of nonmarriage and estimate a discrete-time logit hazard model of whether a woman had a nonmarital birth before the spell ends, or not. We estimate the model using data on women in the PSID.[9] The sample initially consisted of 15,201 women whose marriage and childbirth histories were available in the 1985–1993 Marriage History File and the 1985–1993 Childbirth and Adoption History file.[10] We used the information in these two files to construct all periods when a woman faces the risk of having a nonmarital birth. For a typical woman, the first risk period starts, by assumption, at age 15.[11] It ends either with her first marriage or a nonmarital birth, or is censored if neither occurs before the last year in the data set or age 45 (whichever comes first). If it ends with marriage and the marriage dissolves before age 45, a second risk period begins. It may end via a second marriage, a nonmarital birth, or censoring. We proceed in parallel fashion for the period following the end of a second or higher order marriage. If a risk period ends with a nonmarital birth, we assume a postpartum infertility period of two months.[12] Then, if the woman is still unmarried, another risk period begins. To obtain personal and family background variables, we merged risk period data with information from the main PSID 1968–1993 family and individual files.

The child support policy variables are available for the 1980–1993 period. We divide each risk period into months and use information on annual state of residence to append values for these variables to each monthly risk period. To avoid left censoring, we restrict the sample to women who have one or

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[9]Testing the model with explicit data on men's behavior would be desirable. However, men consistently underreport the number or children they father. Underreporting is especially high among unmarried men. (See Garfinkel, McLanahan, and Hanson, 1998.) To avoid bias created by relying on men's reports, we use data on women.

[10]The Marriage History File reports a complete retrospective marriage history for a household head or wife of any age at the time of the 1985 interview. For a woman of any age who became a new head or new wife during 1986–1993, detail about only first and most recent marriage was reported. In all waves (1985–1993), detail about only first and most recent marriage was reported for other family unit members aged 12–44 at the time of the interview. Fortunately, we can obtain rather complete retrospective marriage histories for these women, since almost all the women in these two categories had at most two marriages. The 1985–1993 Childbirth and Adoption History File contains a complete retrospective birth history for a head or wife of any age and for other family unit members aged 12–44 at the time of the interview in all waves 1985–1993. The file includes records for women who have never had children.

[11]Births to girls younger than 15 are very uncommon.

[12]We assume two months, which is fairly short, because there are many cases in the PSID with a birth interval of only 11 or 12 months.

more risk periods starting no earlier than 1980. The resulting sample contains 5,195 women who have one or more risk periods starting no earlier than 1980. These women reported 1,220 nonmarital births.[13]

One can view S as the product of three factors: (1) the probability of having a child support obligation, (2) conditional on having an obligation, the probability that some or all of the obligation will be collected, and (3) the amount of the obligation that is paid. To capture the theoretical construct S in the empirical specification, we use three measures of state child support enforcement effectiveness that correspond to these factors and a fourth measure that encompasses them all.

Establishing paternity is a prerequisite to child support enforcement, for if paternity is not established there can be no legal child support obligation. The *paternity establishment rate* is measured as the total number of paternities established during a given year divided by the total number of nonmarital births during the same year.[14] This rate proxies the probability that an absent father of a nonmarital child will have a child support obligation.

Establishing paternity for children born outside marriage is a necessary but not sufficient condition for enforcing child support. To measure how well states collect child support obligations, we examine two additional measures. The *collection rate* is the fraction of child support cases that receive a child support payment; i.e., the probability that there is an obligation times the probability that the obligation is collected. *Collections per paying case* is the average amount paid by absent fathers who make at least partial payment of the obligation imposed by the formal child support system. Collections are deflated to 1987 dollars.

The theory of fatherhood and marriage choices implies that child support policy mainly affects behavior in situations where women are likely to go on welfare in the event of a nonmarital birth. This

[13]The sample includes cases from the poverty over sample in the PSID.

[14]The paternity establishment rate can be greater than one since the numerator includes children up to age 18 whereas the denominator includes children born during the past year. The results reported in the text, delete these outlier observations, but the results were virtually identical when the outliers were not deleted. Data on paternities established by each state in each year come from the federal Office of Child Support Enforcement's 1980–93 annual reports to Congress (U.S. Department of Health and Human Services, 1980–1993). Data on nonmarital births by state and year come from Report of Final Natality Statistics (National Center for Health Statistics, 1980–97).

insight suggests that the collection variables be measured with reference to AFDC recipients' child support cases rather than all cases. Accordingly, we present models in which the collection rate is specified as the fraction of AFDC child support cases that receive a child support payment, and the second indicator is collections per paying AFDC case. As a sensitivity check, we provide results for models in which both measures are computed using all cases as the base.

Data for the collection measures come from the annual reports to Congress of the federal Office of Child Support Enforcement (OCSE) for 1980–1993 (U.S. Department of Health and Human Services, 1980–1993). These indicators are probably measured accurately, since they come from administrative records. The all-cases measures are likely to be dominated by collections from formerly married fathers, because divorce cases are easier to process and have a higher payoff for the state than never-married cases. Because formerly married mothers account for between 25 and 30 percent of nonmarital births (Bumpass and Sweet, 1989), we expect the all-case indicators to have some effect on nonmarital childbearing. The AFDC-case measures are more likely than the all-case measures to be influenced by collections from never-married fathers. But, even the AFDC-case measures are heavily influenced by payments from formerly married fathers.

We also examine an alternative indicator of child support policy success. The *effectiveness ratio* is defined as the total amount of child support a never-married mother receives divided by the amount she would be expected to receive under an "ideal" child support system. Garfinkel, Miller, McLanahan, and Hanson (1998) developed the effectiveness ratio using Current Population Survey (CPS) data.[15] The effectiveness ratio is superior to the collection rate insofar as it is based on never-married mothers. It is

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[15]The March CPS does not identify children owed support. Garfinkel et al. assume that all of the mother's children are eligible. The numerator is based on a mother's report of income from child support. To construct the denominator, Garfinkel et al. first estimate the nonresident father's income, using the mother's characteristics, then compute the expected child support payment for each mother by applying the Wisconsin guidelines to the father's estimated income. The Wisconsin guidelines set child support orders equal to 17 percent, 25 percent, 29 percent, 31 percent, and 33 percent of nonresident father's income for one, two, three, four, and five or more children respectively. The ratio is calculated for each never-married mother in the state and aggregated across all never-married mothers to obtain the state-level variable. The guidelines are used as a way of standardizing child support obligations across states. Using an alternative standard, such as "income shares," yields similar rankings for the states.

also superior to all the other measures because it encompasses all of them. However, it is likely to be measured with considerable error because both the numerator and denominator come from the March CPS and CPS sample sizes of never-married mothers in most states are very small.

Higher values of each measure indicate better support enforcement. The theory therefore predicts each measure will be negatively related to the likelihood of a nonmarital birth.

Tables 1 and 2 and Figure 2 provide descriptive information for the enforcement measures. Table 2 shows that correlations among the enforcement measures are modest. Figure 2 arrays states based on their mean values of the collection rate for AFDC cases during 1980–1993. The middle line plots the means. The upper line plots each state's maximum value on the measure during this period; the lower line plots each minimum. The figure shows considerable variation both between and within states. The substantial within-state variation is essential for identifying policy effects in state fixed-effects specifications. Figures for the other enforcement measures (available upon request) similarly exhibit extensive variation between and within states.

In addition to child support enforcement, a number of other factors may affect decisions about nonmarital childbearing. Welfare benefits are an important incentive created by state policy that may affect fertility and marriage behavior. We measure the welfare guarantee as the AFDC cash benefit (in 1987 dollars) provided to a four-person family with no other income.

Whatever the relationship between child support incentives and women's fertility and marriage decisions, it is well known that women's personal and family background variables do affect the likelihood of nonmarital childbearing (Kirby, 2001). These variables partly capture differences in family resources and in nonmonetary benefits and costs associated with this outcome. We include a limited number of such exogenous variables in the models. There are three race/ethnicity dummy variables—non-Hispanic black, Hispanic, and "other race." We include age at the start of each risk period. Religion is indicated with three dummy variables—Baptist, other Protestant, and Catholic. The omitted category is "other religion or none." Mother's marital status when the woman was born and mother's and father's

TABLE 1
Descriptive Statistics

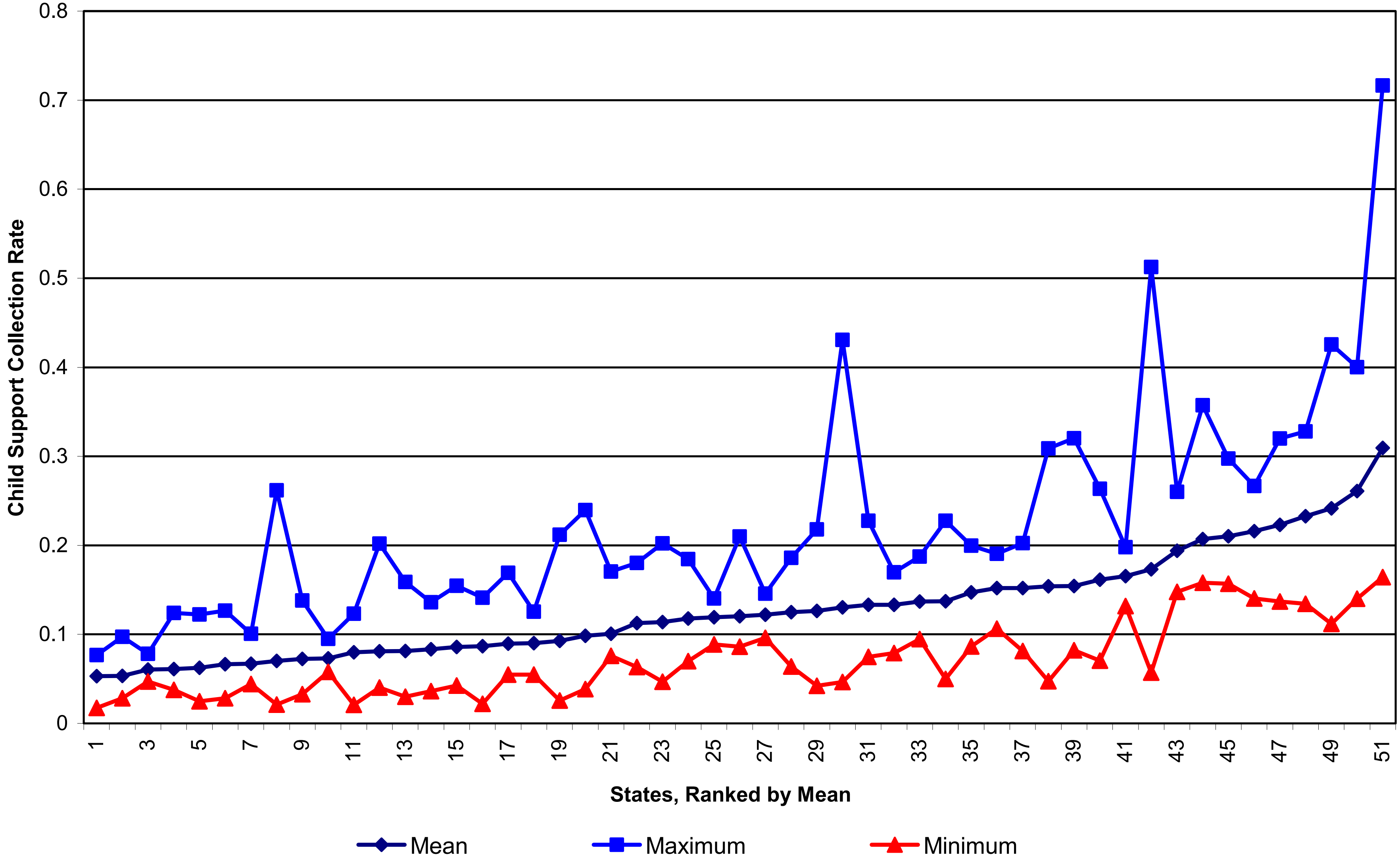
	Mean	Standard Deviation	Minimum	Maximum
Paternity establishment rate	.368	.195	.011	.997
Fraction of cases with collection	.159	.068	.021	.460
Collections per case	477	259	13	3,020
Collections per paying case	2,945	938	115	13,942
Fraction of AFDC cases with collection	.115	.055	.02	.43
Collections per AFDC case	253	123	5	1,737
Collections per paying AFDC case	2,289	660	45	6,203
Effectiveness ratio	17.8	4.1	6.7	36.2
AFDC guarantee	503	221	154	1,077
(All monetary measures in 1987 $s)				
State missing dummy	.038	.190	0	1
Age of woman for the risk year	23.04	7.46	15	45
Black	.534	.499	0	1
Hispanic	.076	.265	0	1
Other race	.01	.099	0	1
Baptist	.418	.493	0	1
Other Protestant	.277	.448	0	1
Catholic	.177	.382	0	1
Woman’s mother never married	.198	.399	0	1
Woman’s mother widowed	.01	.097	0	1
Woman’s mother divorced	.051	.220	0	1
Missing dummy for woman’s mother’s marital status	.254	.435	0	1
Woman’s mother is high school graduate	.357	.479	0	1
Woman’s mother has some college	.138	.344	0	1
Woman’s mother is college graduate	.095	.293	0	1
Missing dummy for woman’s mother’s educational attainment	.022	.146	0	1
Woman’s father is high school graduate	.303	.460	0	1
Woman’s father has some college	.124	.329	0	1
Woman’s father is college graduate	.119	.324	0	1
Missing dummy for woman’s father’s educational attainment	.152	.359	0	1
N = 22,941				

TABLE 2
Correlations among Child Support Enforcement Measures

	Fraction of AFDC CS cases with collection	Collection per paying AFDC CS case	Effectiveness Ratio
Paternity establishment rate	.20	-.02	.18
Fraction of AFDC child support cases with collection		-.26	.24
Collection per paying AFDC child support case			-.18

	Fraction of cases with collection	Collection per paying case	Effectiveness ratio
Paternity establishment rate	.39	.20	.18
Fraction of cases with collection		.13	.35
Collection per paying case			.26

FIGURE 2
Mean, Maximum, and Minimum Child Support Collection Rates for AFDC Cases, by State, 1980-1993



education are also included as control variables. We do not include the woman's own educational attainment because it is endogenous to fertility and marriage behavior.

Our preferred models include state fixed effects and state fixed effects interacted with a linear time trend. Models with state fixed effects and year effects gave similar results.[16]

RESULTS

Table 3 reports the results from three different regressions from models that use all risk periods for each woman in the sample. All models include the policy, personal, and family background variables, and include state fixed effects and state fixed effects interacted with a linear time trend. Column 1 presents the results for the regressions which includes the three measures of enforcement based on OCSE data for AFDC cases, column 2 presents results for the three measures of enforcement based on OCSE data for all cases, and column 3 presents the results for the single comprehensive measure of effectiveness based on CPS data.

All three columns show that more stringent child support is associated with less nonmarital childbearing. The coefficient on the collection rate is negative and significant, though weakly so in column one. The coefficient on the effectiveness ratio is negative and significant. Neither the paternity establishment rate nor collections per paying case is significant.

Though they are not the focus of this study, findings on welfare and the personal background variables are of interest per se and because they affect confidence in the child support results. Consistent with theory and most other recent research, the welfare variable is positive, but weak in all three specifications. The coefficients in Table 3 range from .001 to .002 and are significantly different from zero at only the .10 level in two regressions, and not even that in the third.

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[16]To be able to estimate state effects, we exclude women in states where no one reported a nonmarital birth. Standard errors are corrected to account for the inclusion of multiple observations on the same woman.

TABLE 3
Effect of Child Support Enforcement Policies on the Probability of Nonmarital Childbearing, Full Sample, Alternative Specifications

Independent Variable	AFDC Case Measures	All Case Measures	Comprehensive Measure
Paternity establishment rate	-.3876 (.3502)	-.3411 (.3489)	NA
Fraction of cases with collection	-1.866* (1.075)	-2.492** (1.081)	NA
Collections per paying case	.00003 (.00006)	-.0000096 (.00004)	NA
Effectiveness ratio	NA	NA	-.0366*** (.0127)
Welfare guarantee (in $100s)	.0020* (.0012)	.0021* (.0012)	.0013 (.0012)

Note: Models include state dummies and state-time interactions.

*Significant at .10 level.
**Significant at .05 level.
***Significant at .01 level.

Table 4 presents coefficients for the personal background variables from the model in Table 3, column 1. The corresponding coefficients for the other two models were virtually identical. Consistent with other research, the estimates show that black and Hispanic women are more likely to have a nonmarital birth relative to non-Hispanic white women. The dummy variable for the small "other race" category is not significant. We observe a strongly significant age effect. The older a woman at the start of a risk period, the less likely she is to have a nonmarital birth. Religion is related to nonmarital childbearing, too. Protestants other than Baptists are less likely to have nonmarital births. Catholics are also less likely, but the coefficient is just weakly significant. Being Baptist is not related to this behavior. Being born to a mother who was unmarried at the time of the birth is weakly associated with a greater likelihood of the daughter becoming an unwed mother herself.[17] Daughters of mothers who completed college are less likely to have a nonmarital birth. Father's education tends to be negatively associated with nonmarital childbearing as well.

Age and Marital Status Interactions

We were interested in examining whether child support and welfare differentially affected nonmarital childbearing for younger versus older women and for women before and after a marriage. One hypothesis is that policy would more strongly affect the behavior of older women and women who had been married because, being older and more experienced, their behavior would be more rational and therefore more likely to be affected by economic incentives. Our conceptual analysis, however, suggests that the asymmetric incentives of child support enforcement are most important for the male partners of low-income women who are most likely to be welfare recipients for a long time. Thus we expect to find that child support enforcement will have bigger deterrent effects for younger and never-married mothers.

Tables 5 and 6 present only the child support and welfare coefficients interacted with age and marital status from models that include the same personal and family background variables, state fixed

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[17]We use mother's marital status at the birth because we do not have adequate data on family structure during a sample member's youth.

TABLE 4
Effect of Family Background Characteristics on the Probability of Nonmarital Childbearing, Full Sample

Black	1.130***
	(.106)
Hispanic	.486***
	(.168)
Other race	.144
	(.390)
Age at start of risk period	-.032***
	(.005)
Protestant, non-Baptist	-.425***
	(.114)
Catholic	-.237*
	(.136)
Baptist	-.139
	(.094)
Mother never married	.158*
	(.091)
Mother widowed	.438
	(.277)
Mother divorced	.192
	(.135)
Mother is high school graduate	-.108
	(.078)
Mother has some college	-.170
	(.118)
Mother is college graduate	-.315**
	(.129)
Father is high school graduate	-.272***
	(.084)
Father has some college	-.524***
	(.148)
Father is college graduate	-.220
	(.157)
Constant	-2.924***
	(.498)

Notes: Omitted categories are non-Hispanic white, other or no religion, married/separated, and mother [father] did not complete high school.
Coefficients are for the model in Table 3, column 1. Model also includes state dummies and state-time interactions.
*Significant at .10 level.
**Significant at .05 level.
***Significant at .01 level.

TABLE 5
Effect of Child Support Enforcement Policies on the Probability of Nonmarital Childbearing, by Age

Independent variable	AFDC Case Measures	All Case Measures	Comprehensive Measure
Risk Periods through Age 25			
Paternity establishment rate	-.3765 (.3805)	-.3170 (.3795)	NA
Fraction of cases with collection	-2.519** (1.211)	-3.290*** (1.247)	NA
Collections per paying case	-.0000053 (.00007)	-.00005 (.00005)	NA
Effectiveness ratio	NA	NA	-.0374** (.0157)
Welfare guarantee	.0020 (.0013)	.0024* (.0013)	.0014 (.0013)
Risk Periods, Age 26 and Higher			
Paternity establishment rate	-.0650 (.7392)	-.0130 (.7409)	NA
Fraction of cases with collection	.2603 (2.50)	.689 (2.23)	NA
Collections per paying case	.0001 (.0001)	.00004 (.00009)	NA
Effectiveness ratio	NA	NA	-.012 (.023)
Welfare guarantee	.0018 (.00327)	.0022 (.0026)	.0019 (.0028)

Note: Models include state dummies and state-time interactions.

*Significant at .10 level.
**Significant at .05 level.
***Significant at .01 level.

TABLE 6
Effect of Child Support Enforcement Policies on the Probability of Nonmarital Childbearing, by Marital Status

Independent Variable	AFDC Case Measures	All Case Measures	Comprehensive Measure
Risk Periods for Never-Married Women			
Paternity establishment rate	-.3328 (.3634)	-.2897 (.3622)	NA
Fraction of cases with collection	-2.000* (1.14)	-2.852** (1.16)	NA
Collections per paying case	.00001 (.00007)	-.00002 (.00004)	NA
Effectiveness ratio	NA	NA	-.0400*** (.013)
Welfare guarantee	.0021* (.0012)	.0023* (.0012)	.0012 (.001)
Risk Periods for Formerly Married Women			
Paternity establishment rate	-1.645 (1.746)	-1.473 (1.743)	NA
Fraction of cases with collection	-.756 (4.41)	.6583 (4.410)	NA
Collections per paying case	.0002 (.0003)	-.0000076 (.00017)	NA
Effectiveness ratio	NA	NA	.0486 (.0703)
Welfare guarantee	.0073 (.0052)	.0082 (.0053)	.0103** (.0051)

Note: Models include state dummies and state-time interactions.

*Significant at .10 level.
**Significant at .05 level.
***Significant at .01 level.

effects and fixed effects interacted with a time trend as those in Table 3. As in Table 3, column 1 presents the results for regressions which includes the three measures of enforcement based on OCSE data for AFDC cases, column 2 presents results for the measures of enforcement based on OCSE data for all cases, and column 3 presents the results for the single comprehensive measure of effectiveness based on CPS data. We define premarital risk periods to begin at age 15 and continue until a woman experiences a premarital birth, marries, leaves the risk pool for some other reason, or until the period is censored. Most risk periods occur before the first marriage (81 percent), as do nearly all nonmarital births (90 percent). Thus, we expect the pattern of findings in Tables 5 and 6 to be similar, and for the most part they are.

Child support enforcement has a much stronger effect on deterring nonmarital births among young and unmarried women. Indeed, enforcement has no deterrent effect on women older than age 25 and on previously married women. Similarly, welfare has a stronger effect on nonmarital births among younger than among older women. But, surprisingly, it has a stronger effect on nonmarital births among women previously married as compared to those never married.

Racial and Ethnic Interactions

Racial differences in nonmarital fertility are substantial, are of great interest, and likely reflect different causal processes. To examine this issue, we extend the models by interacting the four child support policy variables with the dummy variables for non-Hispanic black and for Hispanic.

Table 7 presents only the race/ethnicity-specific child support and welfare coefficients from models that include the same control variables as the models in Table 3.[18] The race/ethnic coefficients are reported in three panels.

The general pattern of results is consistent with that in Table 3, though not surprisingly, given smaller subsample sizes, a smaller proportion of the coefficients are statistically significant. The

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[18]We show the sum of the main policy effect coefficient and the coefficient on the race interaction, with the appropriate standard error, rather than the actual main and interaction coefficients and standard errors.

TABLE 7
Effect of Child Support Enforcement Policies on the Probability of Nonmarital Childbearing, Alternative Specifications with Racial Interaction

Independent Variable	AFDC Case Measures	All Case Measures	Comprehensive Measure
White			
Paternity establishment rate	-.733 (.498)	-0.9007* (.5070)	NA
Fraction of cases with collection	-1.793 (1.254)	-1.9780 (1.2459)	NA
Collections per paying case	-.0002 (.0001)	-0.0001 (.0001)	NA
Effectiveness ratio	NA	NA	-.025 (.023)
Welfare guarantee	.0016 (.0013)	0.0016 (.0013)	.0006 (.0013)
Black			
Paternity establishment rate	-.266 (.358)	-.179 (.362)	NA
Fraction of cases with collection	-2.043* (1.093)	-2.764** (1.131)	NA
Collections per paying case	.00004 (.0001)	-.000001 (.000004)	NA
Effectiveness ratio	NA	NA	-.040*** (.013)
Welfare guarantee	.0019 (.0012)	.0020* (.0012)	.0014 (.0012)
Hispanic			
Paternity establishment rate	-2.204* (1.271)	-1.840 (1.299)	NA
Fraction of cases with collection	.624 (1.864)	-.703 (1.725)	NA
Collections per paying case	.0003 (.0002)	.00005 (.0001)	NA
Effectiveness ratio	NA	NA	-.056** (.027)
Welfare guarantee	.0018 (.0013)	.0023* (.0013)	.0019 (.0013)

Note: Models include state dummies and state-time interactions.
*Significant at .10 level.
**Significant at .05 level.
***Significant at .01 level.

effectiveness ratio is negative for all groups, but not significantly different from zero for whites. The collection rate variable is negative in most cases, but only significant for blacks. The paternity establishment coefficients all are negative and, interestingly, become weakly significant for whites and Hispanics in some specifications. As in Table 3, the measure of collections per paying case is consistently insignificant. The welfare coefficients are all positive, but none are statistically significant at the .05 level.

Black nonmarital childbearing is most strongly associated with child support enforcement policies. White behavior is least associated with such policies. Hispanic behavior is in between. This accords with the extent to which women from these groups were involved with the welfare system during the years covered by our data.

The Magnitude of the Deterrence Effect

We use the results in column 2 of Table 3 to simulate the potential effect of better collection rates. For each risk period in the sample, we first use observed values for all the explanatory variables to compute the probability it will end with a nonmarital birth. Since the time unit is one month, we convert the result to show the probability that a nonmarital birth will occur within 12 months of the start of the risk period. The mean simulated probability is .0466. We then compute the probabilities with the AFDC case collection rate set 20 percent higher in every state. The simulated mean probability falls to .0450, or by 3.6 percent. In a second simulation we consider a world in which poorer performing states improve their collection rates to the median state value during 1980–1993 (.116). In this scenario the simulated mean probability falls to .0447, or by 4.1 percent. If every state's collection rate is set to no less than that attained by the state at the 90th percentile in 1993 (.240), the simulated mean probability drops to .0372, or 20.2 percent.

DISCUSSION AND CONCLUSION

A simple economic model of men's fatherhood and marriage choices demonstrates that better enforcement of child support obligations makes it more likely that men will choose to avoid unwed fatherhood. Though better enforcement might provide women more incentive to have children outside marriage, during the period analyzed in the empirical work, child support policy left women's childbearing and marriage incentives largely unaffected. To test the model, we use data from the PSID to analyze spells of nonmarriage among women.

Discrete-time hazard models for the full sample, and for samples restricted to risk periods before a first marriage and before age 26, are consistent with the theory's prediction. They show an inverse relationship between nonmarital childbearing and several measures of child support enforcement vigor. During the 1980–1993 period, women who lived in states that did a better job of collecting child support, as indicated by both the collection rate and the effectiveness ratio, had a lower probability of having a nonmarital birth. Specifications that provide policy effects by race and ethnic group show that the effects are strongest for blacks. The other indicators of enforcement effectiveness—paternity establishment and collections per paying case—show little relationship with nonmarital childbearing.

Successful programs to prevent teen pregnancy and childbearing have proven difficult to create and sustain (Maynard, 1995; Kirby, 2001). We do not know whether the relationships reported here would hold in the current regime of child support enforcement and welfare policy following the 1996 reforms. But the results suggest that greater success in enforcing child support, thereby shifting more of the cost of childbearing from unmarried women to their partners, may help reduce nonmarital childbearing by a significant amount. A conventional service intervention that reduced nonmarital childbearing by 25 percent would be viewed as a major success. Note that even the 90th percentile collection rate as of 1993 leaves ample room for improved enforcement.

This paper is but one of a handful suggesting that child support policy can affect nonmarital childbearing behavior. If future research confirms the tenor of these findings, improved child support enforcement may inadvertently turn out to be one of the more potent interventions for addressing this problem. It would be a refreshing change to find an unintended consequence of social policy intervention with large, positive impacts.

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