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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 with 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 before marriage and during periods of non-marriage following divorce or widowhood. The estimates suggest that women living in states with more effective child support enforcement were less likely to bear children when unmarried. The findings suggest that policies that shift more costs of nonmarital childbearing to men may reduce nonmarital childbearing.

I3, J1
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Births outside of marriage have grown dramatically over the past three decades in the United States. In 1965, the nonmarital birth ratio — the percentage of births to unmarried mothers relative to all births — was 7.7 percent. Since 1994, the figure has been between 32 and 33 percent. Among teenagers the ratio is much higher – 82 percent in 2003 (Ventura & Bachrach 2000, Child Trends 2005).

The nonmarital birth rate (number of births per 1000 unmarried women) has also grown steadily. In 1965 it was 23.4 among all women 15-44 and 16.7 among teenagers 15-19. By 1994 it had risen to 46.9 among all women 15-44 and 46.4 among teenagers. In recent years the rate for teenagers has fallen under 40 while the overall rate has leveled off in the mid-40 range (Ventura & Bachrach 2000, Child Trends 2005).

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 is 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 Policy and Its Expected Effect 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 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 &

Robins 1994; Garfinkel & 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 states to adopt presumptive guidelines for child support awards and to 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. 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 & 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 non-teenage 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 (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 the man's utility depends upon his personal consumption, consumption of the mother and child (whether living with him or not), and the non-financial benefits of parenthood and of 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 (if any, Korenman & Neumark 1991, 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 gross 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

$$Y + (1\text{-}M)*K*G + M*V' - (1\text{-}M)*K*S = C_p + K*C_k \qquad (2)$$

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 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 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, the model's conclusions remains the same.

Related 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 & Charng 1989; Seltzer, McLanahan & Hanson 1998), fathers' remarriage (Bloom, Conrad & Miller 1998), and fathers' earnings (Klawitter 1994, Freeman & Waldfogel, 1998).[9]

Five studies are particularly relevant to the 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 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. (2003) also analyzes state level data and find that effective child support enforcement deters nonmarital births. The effect is robust across all models and specifications.

Huang (2002) and Plotnick et al. (2004) use micro-data to examine the effect of child support enforcement on nonmarital childbearing. Both use the National Longitudinal Survey of

Youth (NLSY) to analyze the likelihood that a woman's first birth is premarital. Focusing on the teenage years, Plotnick et al. (2004) finds that young women living in states with higher rates of paternity establishment are less likely to become unwed teenage mothers. Because of the nature of the NLSY and the focus on teenage behavior, the study examines behavior during 1979-1984. Huang (2002) examines 20 years of data and different indicators of support enforcement. He reports similar relationships when women are age 20 or older but, unlike Plotnick et al., not when they are teenagers.

These papers' findings are 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 non-resident 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. 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). Family planning policies, abortion policies and the availability of family planning and abortion services also are related to the likelihood of nonmarital childbearing (Lundberg and Plotnick 1995, Kirby 2001).

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. However, Heim (2003)'s analysis of panel data on state divorce rates finds that child support enforcement policy has no significant impact on divorce rates.

The analysis here extends this line of research in several ways. We examine outcomes during 1980-1993, a longer time period than analyzed in Plotnick et al. (2004) and one 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. We examine nonmarital childbearing over the full age range typically used in fertility studies, 15 to 44, whereas Plotnick et al.'s (2004) analysis of the teenage years omits most nonmarital births, which occur to women in their 20s. Instead of focusing on first nonmarital births like both micro-data studies, we consider all nonmarital childbearing, whether it occurs before a first marriage or during periods of nonmarriage following divorce or widowhood, and regardless of parity order. In addition, this study is the first to ground the empirical analysis in a formal utility maximization model.

Statistical Models, Data and Explanatory Variables

Because family law is ultimately a state responsibility, state success in child support enforcement varies widely. When the federal government began pushing child support enforcement reforms in the early 1980's, 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.[10] 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.[11] 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.[12] 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.[13] Then, if the woman is still unmarried, another risk period begins. . Each risk period is divided into risk years, the unit of analysis in the multivariate models. To obtain personal and family background variables, we merged risk period data with information from the PSID 1968-1993 family and individual files.

The child support and other policy variables are available starting in 1980. We use information on annual state of residence to append values for these variables to each risk year. To avoid left censoring, we restrict the sample to women who have one or more risk years starting no earlier than 1980. The resulting sample contains 5,195 women. They reported 1,220 nonmarital births.[14]

We examine five alternative indicators of S. Because higher values of each indicator reflect better support enforcement, the theory predicts each will have a negative relationship with the likelihood of a nonmarital birth.

The first is an index that measures the extent to which a state has adopted child support enforcement legislation. This legislative index covers each step of the enforcement process: establishing paternity, obtaining an award, and collecting payments. It includes eight forms of child support legislation: genetic tests, paternity establishment, numerical guidelines, presumptive guidelines, wage withholding under delinquency, immediate wage withholding for new cases, universal wage withholding, and state income tax intercept. Genetic testing permits the father's genetic test results to be used to establish paternity. Paternity establishment allows for the chance to establish the paternity throughout the child's minority until age 18. Numerical guidelines provide a nonbinding set of guidelines to advise judges in their enforcing child support laws. Presumptive guidelines require that these guidelines be used unless the judge can cite "good reason" to deviate. Wage withholding under delinquency indicates that the state has a system similar to income tax withholding that allows deductions of child support obligations and any arrearages from the obligor's paycheck. Federal law now requires all new or modified support orders for welfare recipients to include immediate withholding of support, but this was not so in the 1980s and early 1990s. Last, a state income tax intercept gives the state authority to garnish state income tax refunds up to the amount of overdue child support. The index ranges from zero for states with no law to eight for states with all eight laws. We assume a 1-year lag between legislative enactment and implementation. [15]

The second indicator of the vigor of child support enforcement is child support enforcement expenditures for each state. Expenditures reported by the Office of Child Support

Enforcement are divided by the number of single-mother families in that state. The latter is obtained from the March Current Population Survey. Other studies of behavioral effects of child support policy (Huang et al. 2002, Huang 2002, Freeman & Waldfogel 2001) have used these indicators.

We also examine two more proximate indicators of S. The "effectiveness ratio" is defined as the amount of child support collected on behalf of never-married mothers as a percentage of the total amount of child support owed for all never married mothers assuming the state used the Wisconsin percent-of-income guidelines. The guidelines are a way of standardizing child support obligations across states.[16] The amount owed depends upon the number of children owed support and the income of the non-resident father. Non-resident father income is estimated as a function of the demographic characteristics of the mother and state median earnings and unemployment rates. The ratio is calculated for each never married mother in the state and aggregated across all such mothers to obtain the state-level indicator.

The fourth indicator uses CPS data to compute the total amount of child support received by single mothers divided by the number of single mothers in the state (including those who received none). This ratio reflects not only the strength of child support enforcement, but also differences across states in fathers' ability to pay child support and other factors, such as labor market conditions rates and welfare generosity, that affect the difficulty of collecting child support. For example, low-income fathers have less ability to pay support, and high unemployment further reduces a father's ability to pay support. Race/ethnicity may be important because black fathers are more likely to have children on welfare, who will gain little or nothing from his child support contributions. We, therefore, adjust this measure to purge it of differences across states in the difficulty of enforcement. This is done by first regressing the child support payment on demographic characteristics of the mother – years of schooling, race and ethnicity, whether an immigrant, age, number of children, whether any child is less than age six, and state characteristics – the median earnings of males age 18 to 65 in the state, the maximum AFDC plus food stamp benefit and the state unemployment rate. We use the regression estimates to predict the average amount of child support received in each state. We then take the ratio of the actual to the predicted amount to obtain an adjusted child support payment ratio, and standardize it. This measure is similar, though not identical, to the effectiveness ratio.

Because child support laws and expenditures on child support enforcement help determine the effectiveness ratio and the adjusted child support payment ratio, one may prefer reduced form estimates based on the first two indicators. On the other hand, the latter two reflect aggregate conditions over which an individual has no control and, hence, are likely to be exogenous to any woman's behavior. Both directly capture three key aspects of the child support enforcement process in one parameter: 1. the probability of having a child support obligation, 2. conditional on having an obligation, the probability that some or all it will be collected, and 3. the amount of the obligation that is paid. Thus, both may be better indicators of S than measures of laws and expenditures.

In addition to child support enforcement, 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 these benefits as the AFDC cash benefit (in 1987dollars) provided to a four person family with no other income – i.e. the welfare guarantee. We control for the restrictiveness of state abortion policy and the availability of abortion services, both of which may affect whether a woman carries a nonmarital pregnancy to term (Lundberg & Plotnick, 1995) with three indicators. One is a dummy variable indicating a state law requiring parental consent for a minor to obtain an abortion. A second dummy indicates whether a state restricts use of Medicaid funds for abortions. Both variables are from Levine (2004, table 2.3.). Abortion availability is indicated by the fraction of counties in a state with no abortion provider of any size.[17] One would expect the welfare guarantee and all three abortion variables to be positively related to the likelihood of a nonmarital birth.

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 are related to the likelihood of nonmarital childbearing (Kirby, 2001). These variables partly capture differences in family resources and in non-monetary benefits and costs associated with nonmarital childbearing. We include a limited number of such exogenous variables in the models. We include age at the start of each risk period. There are two race/ethnicity dummy variables – non-Hispanic black and Hispanic. 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 education are also control variables. We do not include the woman's own educational attainment because it is endogenous to fertility and marriage behavior.

We estimate three alternative specifications: one with state fixed effects, a second with state and year effects, and a third with state effects and state effects interacted with a linear time trend. To be able to estimate state effects, we exclude women in states where no one reported a nonmarital birth.

Results

Table 1 reports the results from 12 models that use all risk periods in the sample. Four models include one enforcement measure; the fifth includes both the laws and expenditures measures. All also include welfare, abortion, personal, and family background variables, and one of three fixed effect specifications.

The laws index and expenditure measure are weakly associated with nonmarital childbearing. All three coefficients on the laws index are negative, but only one is significant. Only one coefficient on expenditures is negative; it is significant. When these two indicators are entered jointly (not shown), results for each are similar. The effectiveness ratio is consistently associated with less nonmarital childbearing. The coefficient is negative and significant at the one percent level in columns 1 and 3, and at the ten percent level in column two. All three coefficients on the collections measure are negative and two are significant at the one percent level.

Though they are not the focus of this study, findings on the welfare, abortion and personal background variables are of interest per se and because they affect confidence in the results for the child support policy measures. Table 2 presents coefficients for these variables from the model in table 1, row 1, column 3. Corresponding coefficients from the other models in table 1 were similar.

There is no relationship between the welfare guarantee and nonmarital childbearing, a result consistent with much other research. Living in a state with restrictions on Medicaid funding of abortion has a significant and substantively large association with nonmarital childbearing. Living in such a state is associated with a 0.0243 increase in the probability of having a nonmarital birth each year, relative to a mean of 0.0483. Black and Hispanic women are much more likely to have a nonmarital birth relative to non-Hispanic white women. Being black or Hispanic is associated with a 0.0558 or 0.0388 increase in the probability. We observe a strongly significant, negative age effect. A five-year age difference is associated with a difference in the probability of nonmarital childbearing of 0.0083. Religion is related to nonmarital childbearing. Relative to women raised with no religion or a non-Christian one, Protestants other than Baptists are less likely to have nonmarital births. The effect is large in substantive terms, but weaker than the race and ethnicity effects. Catholics are also less likely, but the coefficient is not quite significant (p = 0.12). Being Baptist is not related to this behavior. Being born to a mother who had never married at the time of the birth is associated with a greater likelihood of the daughter becoming an unwed mother herself.[18] Mother's education has a monotonically negative relationship with a daughter's likelihood of having a nonmarital birth. Father's education tends to be negatively associated with non-marital childbearing as well, though the coefficient on graduated from college is significant only at the 0.12 level. The results for the welfare, abortion, and demographic variables are consistent with other research (Kirby 2001).

Age and marital status interactions

We were interested in examining whether child support differentially affected non-marital 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 3 and 4 present the child support policy coefficients estimated on subsamples based on age and marital status and for models that include the same control variables as in table 1. We define premarital risk periods to begin at age 15 and continue until a woman has a premarital birth, marries, or leaves the risk pool for some other reason, or until the period is censored. Most risk periods occur before the first marriage (81%), as do nearly all nonmarital births (90%). Thus, we expect the pattern of findings in tables 3 and 4 to be similar and for the most part they are.

Child support enforcement has a stronger relationship with non-marital births among younger women. Eleven of the 12 coefficients in the top panel of table 3 are negative and six of these are statistically significant. For older women, eight coefficients are negative and only two are significant. The contrast between never married and formerly married women is similar. Eleven of 12 coefficients in the top panel of table 4 are negative, with seven significant. For formerly married women, only one of the six negative coefficients is significant and there are two anomalous significant positive coefficients.

Interacting enforcement measures with race and ethnicity

Racial and ethnic differences in nonmarital fertility are substantial, of great interest, and likely reflect different causal processes. To examine this issue, we extend the models by interacting the child support policy variables with dummy variables for Hispanic and white. Panel one of table 5 presents the child support policy coefficients for blacks – the reference group. The second and third panels show the coefficients on the interactions terms. The models include the same control variables as the models in table 1.

The results suggest important differences across race and ethnic groups in the relationship of child support enforcement policy to nonmarital childbearing. Black nonmarital childbearing is closely associated with enforcement policies.[19] Ten point estimates are negative. Seven of these are significant at the one or five percent level; one is significant at the ten percent level.

For Hispanics, the four coefficients on the laws index and effectiveness ratio that correspond to the significant estimates for blacks are negative but not significant. These relationships for Hispanics are statistically indistinguishable from those for blacks. The trivial positive point estimate for the expenditure measure in column 3 implies that this relationship for Hispanics is also statistically indistinguishable from the corresponding estimate in panel 1. However, for the adjusted child support payment ratio, the point estimates for the net effect on Hispanics, obtained by summing the corresponding coefficients in panels one and two, suggest that this variable not significantly related to Hispanic behavior.

For whites, the two coefficients on the laws index and expenditure measure that correspond to the significant estimates for blacks are negative and either insignificant or marginally significant. So, similar to Hispanics, these two relationships for whites are basically the same as those for blacks. In contrast, the three point estimates for the effectiveness ratio are positive and differ significantly from the estimates for blacks. Summing each of these estimates with the corresponding one for blacks indicates that the net effect of this explanatory variable on whites is essentially zero. The three point estimates on the adjusted child support payment ratio are positive, which suggests the relationship with nonmarital childbearing is weaker for whites.

The pattern of coefficients in table 5 indicates that blacks show the most responsiveness to the incentives of the child support system, followed by Hispanics. The associations between enforcement policy and nonmarital childbearing are notably weaker for whites.

<u>The magnitude of the deterrence effect</u>

We use the results in row 3, column 3 of table 1 to simulate the potential effect of better enforcement. For each risk year in the sample, we first use observed values for all the explanatory variables to compute the probability it will end with a nonmarital birth. The mean simulated probability is 0.0483. We then compute the probabilities with the effectiveness ratio set 20 percent higher in every state. The simulated mean probability falls to 0.0424, or by 12.2 percent. In a second simulation we consider a world in which poorer performing states improve their effectiveness to the median state value during 1980-1993 (17.65). In this scenario the simulated mean probability falls to 0.0413, or by 14.5 percent. If every state's effectiveness ratio is set to no less than that attained by the state at the 90th percentile during this period (22.30), the simulated mean probability drops to 0.0385, or 20.3 percent below the baseline.

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 try to avoid unwed fatherhood. Though better enforcement might provide women more incentive to bear 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 non-marriage among women.

Discrete time hazard models for the full sample, and for samples restricted to risk periods before age 26 and before a first marriage are consistent with the theory's prediction. They show an inverse relationship between nonmarital childbearing and four measures of child support enforcement vigor. During 1980-1993, women who lived in states that did a better job of enforcing child support had a lower probability of having a nonmarital birth. Specifications that allow the coefficients on the enforcement variables to differ by race and ethnicity show that the relationships are strongest for blacks. Single black mothers are the most likely to receive welfare, so finding that their nonmarital childbearing appears most sensitive to the incentives of child support policy is consistent with theoretical expectations.

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 reported here and in other studies of this relationship consistently 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 significantly. If every state's effectiveness ratio had been no less than the value at the 90th percentile, the findings predict that nonmarital childbearing would have been 20 percent lower. A conventional service intervention that reduced nonmarital childbearing by 20 percent would be viewed as a major success. Note that even the 90th percentile effectiveness ratio during the 1980s and early 1990s leaves ample room for improved enforcement.

This paper is but one of a handful suggesting that child support policy affects 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 reducing this behavior. It would be a refreshing change to find an unintended consequence of social policy intervention that has large, positive impacts.

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Table 1

Effect of Child Support Enforcement on the Logarithm of the Odds of Nonmarital Childbearing, Full Sample, Alternative Specifications

	Fixed effect specification		
Independent variable	State effects	State and year effects	State effects and state-time interaction
Laws index	-.0021 (.0170)	-.0304 (.0359)	-.0994** (.0340)
Expenditures on enforcement	.0006 (.0007)	.0016# (.0010)	-.0021* (.0010)
Effectiveness ratio	-.0402** (.0121)	-.0236# (.0132)	-.0411** (.0128)
Adjusted child support payment ratio	-.2239** (.0751)	-.1372 (.0868)	-.2803** (.0952)

N = 22,107 in all models.
\# Significant at .10 level
* Significant at .05 level
** Significant at .01 level

All models include the Medicaid funding dummy variable. Results are similar when other indicators of abortion policy and access are used instead.

Table 2

Effect of Family Background Characteristics and Welfare Benefits on the Logarithm of the Odds of Nonmarital Childbearing, Full Sample

Welfare guarantee (in $100s)	-.00013 (.0010) [.00006]
Restricted Medicaid funding of abortion	.5283** .1968 [.0243]
Black	1.2137** (.1071) [.0558]
Hispanic	.8438** (.1633) [.0388]
Age at start of risk period	-.0360** (.0047) [-.00165]
Protestant, non-Baptist	-.4386** (.1170) [-.0205]
Catholic	-.2084 (.1345) [-.0096]
Baptist	-.1332 (.0984) [-.0061]
Mother never married	.1616 # (.0931) [.0074]
Mother widowed	.3469 (.2811) [.0160]
Mother divorced	.2223 (.1397) [.0102]

Mother graduated high school	-.1469# (.0797) [-.0068]
Mother had some college	-.2091# (.1210) [-.0096]
Mother graduated college	-.3310* (.1313) [-.0152]
Father graduated high school	-.3079** (.0855) [-.1142]
Father had some college	-.5185** (.1512) [-.0238]
Father graduated college	-.2486 (.1587) [-.0114]
Constant	-2.1311** (.4487)

Standard errors are in parentheses. The marginal effect of a one unit change on the probability of nonmarital birth, relative to the mean probability of 0.0483, is in brackets.

N = 22,107. 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 1, row 1, column 3.

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

Table 3

Effect of Child Support Enforcement on the Logarithm of the Odds of Nonmarital Childbearing, by Age

Risk Periods Through Age 25

	Fixed effect specification		
Independent variable	State effects	State and year effects	State effects and state-time interaction
Laws index	-.0077 (.0194)	-.0076 (.0398)	-.0707# (.0393)
Expenditures on enforcement	-.0001 (.0008)	.0010 (.0011)	-.0022# (.0012)
Effectiveness ratio	-.0458** (.0139)	-.0326* (.0151)	-.0537** (.0168)
Adjusted child support payment ratio	-.1412 (.0898)	-.0744 (.1011)	-.2476* (.1143)

N = 15,429

Risk Periods at Age 26 and Higher

	Fixed effect specification		
Independent variable	State effects	State and year effects	State effects and state-time interaction
Laws index	.0130 (.0369)	-.0365 (.0765)	-.0863 (.0618)
Expenditures on enforcement	.0007 (.0013)	.0010 (.0018)	-.0028 (.0019)
Effectiveness ratio	-.0140 (.0223)	.0077 (.0244)	-.0217 (.0227)
Adjusted child support payment ratio	-.3635* (.1574)	-.2721 (.1799)	-.4309* (.1964)

N = 6,440

Significant at .10 level

* Significant at .05 level

** Significant at .01 level

All models include the Medicaid funding dummy variable. Results are similar when other indicators of abortion policy and access are used instead.

Table 4

Effect of Child Support Enforcement on the Logarithm of the Odds of Nonmarital Childbearing, by Marital Status

Never Married Women

	Fixed effect specification		
Independent variable	State effects	State and year effects	State effects and state-time interaction
Laws index	-.0225 (.0182)	-.0362 (.0374)	-.0835* (.0356)
Expenditures on enforcement	-.0002 (.0007)	.0013 (.0010)	-.0022* (.0011)
Effectiveness ratio	-.0426** (.0122)	-.0256# (.0137)	-.0500** (.0129)
Adjusted child support payment ratio	-.2013** (.0784)	-.1092 (.0902)	-.2799** (.0999)

N = 18,034

Formerly Married Women

	Fixed effect specification		
Independent variable	State effects	State and year effects	State effects and state-time interaction
Laws index	.1146* (.0534)	-.0089 (.1255)	-.1965# (.1143)
Expenditures on enforcement	.0050** (.0021)	.0040 (.0031)	.0003 (.0040)
Effectiveness ratio	-.0176 (.0547)	.0032 (.0532)	.0152 (.0663)
Adjusted child support payment ratio	-.3666 (.2662)	-.2723 (.3198)	-.5445 (.3480)

N = 3,568
# Significant at .10 level
* Significant at .05 level
** Significant at .01 level

All models include the Medicaid funding dummy variable. Results are similar when other indicators of abortion policy and access are used instead.

Table 5

Effect of Child Support Enforcement on the Logarithm of the Odds of Nonmarital Childbearing, with Racial and Ethnic Interactions

Coefficient for blacks

	Fixed effect specification		
Independent variable	State effects	State and year effects	State effects and state-time interaction
Laws index	-.0147 (.0183)	-.0494 (.0372)	-.1146** (.0347)
Expenditures on enforcement	.0006 (.0007)	.0015 (.0010)	-.0021* (.0010)
Effectiveness ratio	-.0497** (.0132)	-.0316* (.0142)	-.0515** (.0138)
Adjusted child support payment ratio	-.2499** (.0792)	-.1607# (.0903)	-.3048** (.0980)

Hispanic interaction term

	Fixed effect specification		
Independent variable	State effects	State and year effects	State effects and state-time interaction
Laws index	.0397 (.0500)	.0526 (.0548)	-.0509 (.0574)
Expenditures on enforcement	.0007 (.0015)	.0009 (.0015)	.0001 (.0015)
Effectiveness ratio	-.0011 (.0456)	-.0019 (.0464)	-.0055 (.0470)
Adjusted child support payment ratio	.1324 (.1605)	.1196 (.1618)	.1367 (.1761)

Table 5 (continued)

White interaction term

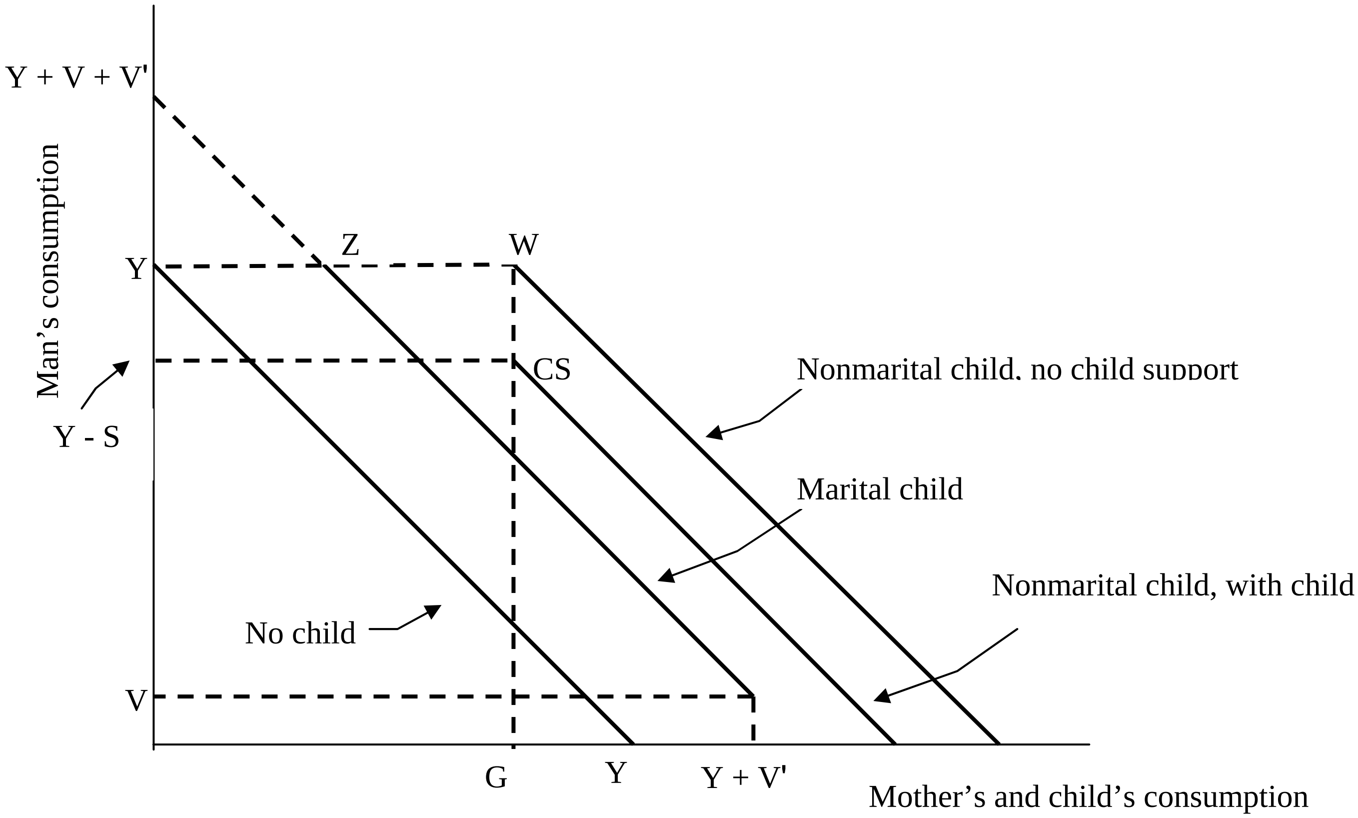
	Fixed effect specification		
Independent variable	State effects	State and year effects	State effects and state-time interaction
Laws index	.0516 (.0324)	.0645# (.0354)	-.0708# (.0371)
Expenditures on enforcement	.0000 (.0010)	-.0000 (.0010)	-.0000 (.0010)
Effectiveness ratio	.0543* (.0239)	.0450# (.0235)	.0687** (.0252)
Adjusted child support payment ratio	.0918 (.1130)	.0803 (.1117)	.0980 (.1190)

N = 22,107 in all models
\# Significant at .10 level
* Significant at .05 level
** Significant at .01 level

All models include the Medicaid funding dummy variable. Results are similar when other indicators of abortion policy and access are used instead.

Figure 1

Budget Constraints for Different Fatherhood and Marriage Options



Endnotes

[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.

[2] In response to the 1996 welfare reforms, most states have eliminated the 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.

[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 & Waldfogel, 1998).

[6] If $V + V' \leq 0$ 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 $G > V + V'$.

[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.

[9] See Lerman and Sorenson (2003) for a survey of the effects of child support incentives on behavior.

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

[11] 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.

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

[13] 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.

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

[15] Information on legislation was collected mainly from various years of the State Legislative Summary from the National Conference of State Legislatures and the OCSE Legislative Tracking System Report from the U.S. Department of Health and Human Services. Inconsistencies between the two sources were resolved by examining each state's existing laws.

[16] Garfinkel, Miller, McLanahan, and Hanson (1998) provide further details on the construction of the effectiveness ratio. We thank Theresa Heintze for sharing the effectiveness ratio data used in this analysis and Lenna Nepomnyaschy for sharing the regression based measures. Alternative measures are discussed in Nepomnyaschy (2003).

[17] We also examined two alternative indicators – the fraction of counties without a non-hospital provider of 400 or more abortions, and the fraction of women in a state living in counties without a non-hospital provider of 400 or more abortions. Results for the child support enforcement variables were virtually the same for all three indicators and for the two abortion policy dummy variables.

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

[19] Huang (2002) also reports stronger effects of enforcement on blacks than whites. He does not report separate estimates for Hispanics.