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Better Child Support Enforcement: Can It Reduce Teenage Premarital Childbearing?

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ABSTRACT

Stricter child support enforcement may reduce unwed childbearing by raising the costs of fatherhood. We investigate this hypothesis using a sample of young women from the National Longitudinal Survey of Youth, to which we add information on state child support enforcement. Models of the probability of a teenage premarital birth and of teenage premarital pregnancy and pregnancy resolution show that, during the early 1980s, teens living in states with higher rates of paternity establishment were less likely to become unwed mothers. This relationship is stronger for whites than for blacks. The findings suggest that policies that shift more costs of premarital childbearing to men are likely to reduce this behavior, at least among whites.

Key Words: Child support, paternity, premarital childbearing, teenage childbearing

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The proportion of American children born outside marriage has grown dramatically over the past three decades. In 1965, when Daniel Patrick Moynihan (1965) warned that marital instability and father absence were undermining the progress of black Americans, the nonmarital birth ratio — the percentage of births to unmarried mothers relative to all births — was approximately 23 percent for black Americans and 8 percent for the country as a whole. By 1996 the figures were 70 percent and 32 percent respectively (Ventura et al. 1998). Among teenagers, the nonmarital birth ratio is much higher – in 1996, 95 percent for blacks, 69 percent for whites and 76 percent for all teenage women (Ventura et al. 1998).

The premarital birth rate (number of births per 1000 unmarried women) has also grown steadily in recent decades. In 1965 it was 23.4 among all women 15-44 and 16.7 among teenagers 15-19 (7.9 for whites, 77.1 for non-whites) (U.S. Department of Health and Human Services 1995). By 1991 it had risen to 45.2 among all women 15-44 and 44.8 among teenagers. Since 1991 the rate has hovered around 45 among all women 15-44. Until 1995 it also remained around 45 for teenagers, but in 1996 it dropped to 42.9. The recent drop in the teen premarital birth rate was particularly large for blacks, but the racial difference in the teen rate remains substantial (34.5 for whites, 89.2 for blacks) (Ventura et al. 1998).

In response to these changes, policy makers and social scientists have become increasingly concerned about the causes and consequences of premarital childbearing and the extent to which government policies may have fostered its growth.[i] Despite empirical evidence to the contrary (Moffitt 1992 1998), many people believe that increases in welfare benefits — AFDC, food stamps, and Medicaid — are the major cause of increases in premarital childbearing

(Murray 1984, 1993). 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. State legislators have followed suit. These federal and state initiatives have included policies that lower AFDC benefits, limit eligibility via time limits and family caps, impose 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 proposals are driven by the assumption that women bear the primary responsibility for premarital childbearing and that "getting tough on mothers" is the best strategy for reversing the trend.

What is largely missing from recent debates is recognition of men's role in premarital childbearing and discussion of how government policies may affect men's behavior. There has been little notice of how the government's failure to establish paternity and to collect child support from nonresident fathers may have contributed to men's failure to take responsibility for contraception or to marry their sexual partners (Sorensen 1997). Although efforts to establish paternity and enforce child support have intensified during the past decade and were strengthened by the Personal Responsibility and Work Opportunity Reconciliation Act, they are generally viewed as ways of reducing welfare costs rather than as strategies for preventing premarital births.

The asymmetrical focus on women appears unreasonable. Decisions about sexual intercourse, contraception, and marriage involve two adults rather than one. The same often is true for abortion decisions. Moreover, policies designed to make single motherhood financially more difficult for women run the risk of harming children who are already born (Duncan &

Brooks-Gunn 1997). In contrast, policies aimed at making unmarried fatherhood more costly for men have the potential to both prevent premarital births and benefit existing children.

We use individual level data from the National Longitudinal Survey of Youth, combined with state level information on child support enforcement, to test the hypothesis that strong child support enforcement is associated with lower teenage premarital birthrates. We estimate two related models. The first simply analyzes the probability that a woman had a premarital birth as a teenager. The second recognizes that a teenager's route to premarital motherhood depends on several major outcomes or decisions: becoming pregnant before marriage and then, given a pregnancy, not aborting it and not marrying before the birth.[ii] We examine the factors, including measures of child support enforcement, that affect the probability of a premarital pregnancy, and its resolution. We report tentative evidence in support of the hypothesis.

U.S. Child Support Policies and Their Expected Effects on Premarital 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; Beller & Graham 1993).

During the past decade and a half, 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 (CSE) 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

1998: 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 a number of provisions aimed at strengthening paternity establishment for children born to unmarried parents. Reforms in the 1996 Personal Responsibility and Work Opportunity Reconciliation Act place are aimed at further improving the child support's 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. Paternity establishment ratios (the number of paternities established in a given year divided by the number of nonmarital births) have increased from 20 percent in the early 1980s to 46 percent in 1994 (Nichols-Casebolt & Garfinkel 1991, U.S. House of Representatives 1998, table 8-22). The proportion of never married mothers with a child support award has grown from 12 percent to over 20 percent over this same period (Beller & Graham 1993; Hanson 1995).

Why child support policy may affect premarital childbearing

Sociological and psychological theories, as well as substantial empirical research, conclude that becoming a premarital parent is a product of many influences (Miller & Moore 1990; Moore et al. 1995; Federal Interagency Forum on Child and Family Statistics 1998). These include biological factors, individual social and personality characteristics, family influences, and peer, community and other contextual influences. Economists and some sociologists argue in addition that the likelihood of premarital parenthood is partly influenced by economic incentives embodied in the labor market, such as women's earnings opportunities or availability of "marriageable males," and in some public policies (Duncan & Hoffman 1990, Lundberg & Plotnick 1995, South 1996; Willis & Haaga 1996; Wilson 1987).

Improved 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. This change in the policy regime increases the expected costs of fatherhood for absent fathers. According to economic theory (Becker 1991; Weiss & Willis 1985; Willis 1994; Willis & Haaga 1996), this disincentive would, other things equal, make men more reluctant to father children outside marriage (and to divorce if there are children from the marriage).

If child support enforcement only affected men's incentives, the theoretical link from improved enforcement to less premarital childbearing would be clear. For unmarried women, though, improved enforcement that leads to higher child support payments might appear to reduce the costs of children and, hence, create more of an incentive to have children outside marriage.

Given the existence of welfare and its rules about the treatment of child support during the time period we analyze, the size of this countervailing effect on women was likely to have been small. During that period, a mother on welfare was allowed to keep only the first $50 of child support each month.[iii] If the father paid more than $50, all the additional money went toward reducing the state's 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. In contrast, unmarried men were supposed to pay the full child support order regardless of how much (or how little) it helped their children. Thus, the effect of the disincentive on unmarried men's motivation to avoid nonmarital parenthood was likely to have been stronger than the reverse incentive effect for unmarried women. This reasoning leads us to the hypothesis that, on net, better child support enforcement will be associated with lower premarital birthrates.

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 & Charng 1989; Seltzer, McLanahan & Hanson 1998), fathers' remarriage (Bloom, Conrad & Miller 1996), and fathers' earnings (Klawitter 1994, Freeman and Waldfogel, 1998).

Three papers are particularly relevant to our argument that child support policy is likely to have empirically significant effects on premarital childbearing. Sonenstein, Pleck and Ku (1994) found 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. Besley and Case's (1997) analysis of state data reported that, net of economic and demographic conditions, states which adopted presumptive guidelines for setting child support awards or allowed establishment of paternity up to age 18 had lower out-of-wedlock birth rates. Gaylin et al. (1996) also analyzed state level data and found that effective child support enforcement deters nonmarital births. 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 women and with our hypothesis that increasing the costs of children for non-resident fathers will lower the incidence of premarital childbearing.

Indirect support for the argument that the incentives of child support policy may affect premarital childbearing comes from evidence that the incentives of other public policies affect this behavior. Lundberg and Plotnick (1995) found 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 limited to single mothers make women more able to afford to be a lone mother and reduce the gains from marriage. The empirical evidence has suggested that these incentives undermine marriage and promote out-of-wedlock childbearing, though the importance of these effects remains uncertain (Moffitt 1998). Nixon (1997) is also indirectly relevant. It found 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 highly personal demographic choice such as marital status, this article suggested such policies may also affect other highly personal demographic outcomes such as premarital childbearing.

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 CSE 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 hardly at all. Variability in state programs offers an opportunity to explore the impact of public policy. This study exploits the varying vigor and commitment with which states have implemented their CSE programs to assess the impact of alternative CSE policies on premarital teenage childbearing.

To analyze this relationship we estimate two related models using data on individual young women. The first model's outcome is whether a woman had a premarital birth as a teenager, or not. The second model jointly considers four major possible outcomes of premarital teenage sexual behavior: no premarital pregnancy, a premarital pregnancy terminated by abortion, a premarital pregnancy followed by marriage before the birth, and a premarital pregnancy ending in a premarital birth. (We do not consider the choice between raising the child as an unmarried mother and placing the infant for adoption because the placement option is so rarely exercised.) Because the outcomes in both models are categorical ("yes-no") and we wish to estimate determinants of the probability that a woman experienced each outcome, standard linear regression methods are not appropriate. Instead, we estimate the first model using the simple logit estimation method and the second using the multinomial logit method. (Note that research based on aggregate level data, which lack information on the resolution of premarital pregnancies, can not examine the more detailed behavioral relationships in our multinomial logit model.)

The individual level data come from the National Longitudinal Survey of Youth. The sample initially consisted of 2,153 young women in the NLSY who were age 14 to 16 in 1979. We follow each woman's fertility and marital history until she reaches age 20 (i.e. depending on age in 1979, until 1983, 1984 or 1985) and examine only first premarital pregnancies. (We do not examine out-of-wedlock births of young women who married and divorced, and then became pregnant, all before age 20.) We drop women with insufficient information in their fertility and marital records to indicate whether and when they first became premaritally pregnant, whether they aborted such a pregnancy or carried it to term, and whether they married between a premarital conception and delivery. After we drop cases with missing independent variables, the sample for the logit models has 1865 cases.

We estimate models for the entire sample. Because racial differences in teenage fertility and marriage behavior are substantial, of great interest, and likely reflect different causal processes, we also analyze separately non-Hispanic whites and non-Hispanic blacks (hereafter simply referred to as whites and blacks). There are too few Hispanic cases (99) to estimate separate models for that group.

Table 1 presents basic statistics on the fertility and marital behavior of the sample, adjusting for sampling weights. The pregnancy rate among the entire sample was 28 percent. Of these, 11 percent ended by miscarriage or still birth. Of the pregnancies that continued, 30 percent were aborted. Of the live births, 63 percent were to unmarried women. The others married between conception and birth. The net result was a premarital birth rate of 10.9 percent. The corresponding figures for whites were 24, 11, 37, and 48 percent, and a net premarital birth rate of 6.6 percent. For blacks, they were 48, 12, 12, 94 and a net premarital birth rate of 34.8 percent.[iv]

[Table 1 about here]

We analyze three key measures of state child support enforcement effectiveness and efficiency: number of paternities established as a percentage of nonmarital births, average annual support collections per AFDC case, and AFDC child support dollars collected per administrative dollar spent. The AFDC specific variables are used because most premarital births occur among women who become eligible for this program. Hence, these indicators are likely to show stronger associations with premarital births. The number of paternities established and the other two enforcement variables are from Office of Child Support Enforcement (1982). The number of nonmarital births in the state is from vital statistics data for the corresponding years.[v]

Higher values of each measure indicate better child support enforcement. The theory discussed in section 2 leads us to expect each measure to be negatively related to the probability of a premarital birth in both the simple and multinomial logit models. The theory does not, however, lead to unambiguous predictions about the effects of each measure on the probability of a premarital pregnancy terminated by abortion and the probability of a premarital pregnancy followed by marriage before the birth. Suppose better enforcement leads to less sexual activity, better contraception, and, so, fewer premarital pregnancies. Then better enforcement would tend to be associated with fewer abortions and marital births. It is possible, though, that the likelihood of abortion could rise while the likelihood of a marital birth could fall, or vice versa. All one can say for certain is, given a reduction in the likelihood of a premarital pregnancy, the combined likelihood of abortion, marital birth and premarital birth must fall.

On the other hand, suppose better enforcement has little effect on the chances of premarital pregnancy, but encourages men to avoid becoming unmarried fathers. Then, holding the probability of premarital pregnancy constant, better enforcement would be associated with more abortions and more legitimated births.

In addition to paternity establishment and child support enforcement, a number of other factors may affect out-of-wedlock birthrates. We include welfare, abortion and family planning variables to measure other incentives created by state public policies that are likely to affect teen fertility and marriage behavior. The welfare guarantee, generally available only if the young woman bears a child out-of-wedlock, measures an important long run consequence of adolescent fertility and marriage decisions. We measure the guarantee as the AFDC cash benefit provided to a family with no other income plus the amount of food stamps it would receive (measured in 1983 dollars). Other measures of the guarantee are likely to produce similar results, as shown in Lundberg and Plotnick (1995.)

As proxies for the short run cost of obtaining an abortion, we include three measures of the availability of abortion and of state funding policies for abortion. These measures vary by state and, where data are available, over time.

The first measure is a series of three dummy variables to represent policy on state funding of abortions for needy women. The first dummy variable equals one if a state funded all or all medically necessary abortions. The second equals one if the state pays for abortions for medical and/or emotional reasons; and the third equals one if the state pays only if full-term pregnancy may endanger the life of the woman or if the pregnancy resulted from rape or incest. The omitted category is a policy of not paying for abortions for any reason. The second measure is an index of the restrictiveness of state abortion law. The third, the percentage of counties within a state with large providers of abortions, indicates the availability of abortion. We also include a dummy for the presence of state laws, regulations and policies which restrict the advertisement, sale, or licensing of contraceptives as a proxy for the short run cost of obtaining contraceptive services and supplies.[vi]

One issue arises in assigning values of the CSE, welfare, abortion and family planning variables. If a girl becomes premaritally pregnant, we could assign values of her state's variables for the year she became pregnant. But what year's values should we assign to girls who avoid pregnancy? To be consistent, for each girl in the sample we assign values for her state of residence when she was 16. We do not have CSE information for 1979. For sample members who were 16 in 1979, we use 1980 CSE values for their 1979 state of residence.

Personal and family background variables also affect the overall likelihood of premarital childbearing and each of the outcomes in the multinomial logit model (Moore et al 1995). These variables partly capture differences in family resources and in non-monetary benefits and costs associated with the outcomes. There are two family structure dummy variables: lived in a mother-only family, or in one of a variety of other family structures. The omitted category is "lived with both natural parents." There is a dummy variable for whether an adult female, usually the mother, worked in the paid labor force when the sample member was age 14. Variables for mother's education, number of siblings and whether a foreign language was spoken at home are included. Religiosity is gauged by a dummy set to one if the young woman reports she never or rarely attends religious services. The model also includes age of menarche because teenagers who reach physical maturity earlier may be more likely to begin sexual activity earlier, thereby running a higher risk of becoming pregnant. Last, there are dummy variables for being black and Hispanic.

Labor market conditions may also affect behaviors related to premarital childbearing. For example, better employment conditions for women may reduce marriage by increasing women's independence, but by encouraging women to avoid pregnancy to take advantage of earnings opportunities, they could reduce birth rates more than marriage rates and thereby decrease out-of-wedlock childbearing. Previous literature finds that poor employment opportunities for men (Wilson 1987; Lichter et al. 1992; Mare & Winship 1992) and good employment opportunities for women (South & Lloyd 1993; Schultz 1994; McLanahan & Casper 1995) lead to higher out-of-wedlock birthrates. We test the sensitivity of the findings on the child support policy variables by adding indicators of state labor market conditions to the model.

Results

Full Sample: Table 2 contains estimates from the simple logit model of the relationship between child support enforcement policies and the likelihood of teenage premarital childbearing. The model in column one includes the personal, family background, welfare benefit, abortion, and family planning variables described earlier, and all three measures of child support enforcement. Because the three measures are strongly intercorrelated (simple correlations range between .35 and .60), column two shows results when each is the only CSE variable entered in the regression.

[Table 2 about here]

Column one shows a weak relationship between CSE and premarital childbearing. Two coefficients have the anticipated sign. The coefficient on the paternity establishment percentage is significant, but only marginally at the 10 percent level. However, when one CSE measure is entered at a time, one is statistically significant at the 5 percent level and one marginally at the 10 percent level, both with the anticipated negative coefficient. [vii]

Table 3 contains estimates of the multinomial logit model with its more disaggregated outcomes. The set of explanatory variables is the same as in the model in column 1 of table 2. The omitted outcome is "no premarital pregnancy." The results are similar. A higher paternity establishment percentage is negatively associated with the likelihood of a premarital birth at the 5 percent level of significance. Its coefficient is negative for the other two outcomes – abortion and marital birth – but not statistically significant. That the coefficient is negative in all three columns suggests, however, that better paternity establishment mainly leads to fewer premarital births by discouraging premarital pregnancies. As in table 1, the other two CSE variables show no relationship with the outcomes. When we estimate the same multinomial model, but include just one CSE variable at a time, the paternity establishment percentage and AFDC collections per case are each significant with the expected negative sign for the premarital birth outcome (not shown in table 3).

[Table 3 about here]

The nested logit model can also estimate the determinants of these outcomes (Lundberg & Plotnick 1995). The nested logit model relaxes the multinomial logit model's assumption of "independent irrelevance of alternatives" but requires imposing a specific nesting sequence. We estimated a nested logit model in which the abortion decision is conditional on the occurrence of a premarital pregnancy, and the marriage decision is conditional on continuing the pregnancy. Results are consistent with those in tables 2 and 3.

<u>Racial Subsamples</u>: Estimating the models on a sample restricted to young white women yields the results in the middle two columns of table 2 and the middle panel of table 3. The patterns of coefficients generally resemble those for the full sample. There are fewer significant effects, perhaps because of smaller sample size. The importance of the paternity establishment percentage again stands out.

The last two columns of table 2 and the third panel of table 3 provide corresponding estimates for the subsample of young black women. There is little evidence that child support enforcement policies are related to the likelihood of premarital childbearing among black teenagers. No CSE variable is significant in any of the simple logit models. (Lundberg and Plotnick (1995) similarly report much weaker relationships between the welfare, abortion and family planning variables and premarital childbearing among blacks than among whites.) When the outcomes are disaggregated, one coefficient on the AFDC collections per case measure is marginally significant. The consistent significant relationship seen earlier between better paternity establishment and less premarital childbearing vanishes. On the other hand, all point estimates for the paternity establishment and AFDC collections measures are negative, which suggests there may be a weak relationship.

Sensitivity of Findings to Labor Market Variables: To test the sensitivity of the findings on the child support policy variables, we extend the model to include several indicators of state labor market conditions: median female manufacturing earnings, median manufacturing earnings, the unemployment rate, and the male and female labor force participation rates. We include all five indicators in the models, then enter them singly. As with the other state level variables, we assign values to each young woman based on her state of residence when she was 16.

Findings on the CSE variables are robust to these six variations. In the simple logit model, the coefficient on the paternity establishment rate is negative in all six equations and statistically significant at the 10 percent in three, and at the 5 percent level in two. The negative coefficient on AFDC collections per case appears in all 6 equations and is significant only at the 10 percent level in three of them. The coefficient on the AFDC collections per administrative variable is again positive and nonsignificant in all six variations.

Similarly, in the multinomial logit model, the coefficient on the paternity establishment rate for the premarital birth outcome is negative in all six variations. In four of the six models it is again significant at the 5 percent level, and in one at the 10 percent level. The other 8 coefficients in table 3 are not statistically significant. Of the corresponding 48 coefficients in the six variations, 46 have the same sign and are also not significant. One flips sign while remaining nonsignificant. And one (for the paternity establishment rate in the marital birth outcome) is again negative but barely significant at the 10 percent level.

Discussion and Conclusion

The results of this study, which are the only ones based on individual-level data, provide encouraging but tentative evidence of an inverse relationship between teenage premarital childbearing and child support enforcement vigor. The simple logit estimates of this relationship for the overall and whites samples generally show that teenage women who lived in states that did a better job of establishing paternity during the early 1980s had a lower probability of becoming unwed mothers. This relationship was much weaker for young black women. The multinomial logit analysis provides suggestive, though not conclusive evidence that better paternity establishment mainly led to fewer premarital births by discouraging premarital pregnancies. The two other indicators of state child support enforcement effectiveness and efficiency that we analyze - average annual support collections per AFDC case, and AFDC child support dollars collected per administrative dollar spent - show little relationship with the outcomes.

The stronger estimated effect for the paternity establishment variable makes sense. Information on how effectively a state establishes paternity is probably fairly easy for a man in his teens or twenties to learn by observing how often it happens to other unwed fathers in his circle of friends and relatives and in his neighborhood. Given this information, some knowledge of the long run consequences of legal paternity, and a desire to minimize the costs of unwed fatherhood, men may well change their sexual and other behavior to reduce these costs. In contrast, information about how effectively a state collects child support on behalf of AFDC mothers could reflect how well the state does in collecting child support from divorced rather than unwed fathers and, in any case, is much harder for such men to know and, hence, act upon.

We use the results in the first panel of table 3 to simulate the potential effect of better paternity establishment. For each woman in the sample, we first use observed values for her personal characteristics and the 3 CSE variables in the state where she lives to compute the probability she will become an unwed mother. The mean simulated probability is .1000. We then compute the probabilities if the paternity establishment rose 20 percent in every state. The simulated mean probability falls to .0946, or by 5 percent. Given that some states are doing a good job of paternity establishment and child support enforcement, in a second simulation we consider a world in which poorer performing states improve their paternity establishment performance to the median state (before the simulated improvement). In this scenario the simulated mean probability falls to .0922, or by 8 percent. If every state reached the establishment percentage currently attained by the best states (around 75 percent), the simulated mean probability drops to .0501, or 50 percent. At the same time, the mean probability of avoiding a teenage premarital pregnancy is .7554 at the simulated baseline and rises to, respectively, .7652, .7673, and .8446.

For the white sample, parallel simulations yield a base level probability of a premarital birth of .0598 and declines to .0536 (10 percent), .0517 (14 percent), and .0159 (73 percent). The likelihood of avoiding a teenage premarital pregnancy is .7875 at the simulated baseline and rises to, respectively, .7982, .7991, and .8708.

Successful programs to prevent teen pregnancy and childbearing have proven remarkably difficult to create and sustain (Maynard, 1995). Our results suggest that greater success at establishing paternity and enforcing child support, thereby shifting more of the cost of childbearing from teenage women to their partners, may help reduce premarital childbearing. If future research confirms this study’s findings, improved child support enforcement may inadvertently turn out to be one of the more potent interventions for addressing this social problem.

The data and analysis have several limitations. The data are from the late 1970s and early 1980s. We do not know whether the relationships we report would hold in the current regime of child support enforcement. Because of limitations in our sample, the models do not include state dummy variables for the states (sometimes described as state “fixed effects” models) and, thus, the estimates of policy effects may be biased. (We note, however, that results from studies of the child support – nonmarital birth relationship based on aggregate data were robust to the inclusion of state fixed effects.) Also, we are not able to directly measure men’s behavior with the data set. The results reported here, then, are best viewed as a weak test of the relationship between child support policy and nonmarital childbearing and need confirmation with other data.

States are increasingly addressing their child support enforcement responsibilities, and the results, both direct and indirect, have the potential to create large benefits for single parents and their custodial children, as well as citizens and taxpayers in general. In contrast, cutting welfare benefits to single mothers and implementing other policies designed to make single motherhood financially more difficult will have, at best, a modest negative effect on the rate of premarital childbearing, but will run the risk of harming the short and long run prospects of children who are already born.

References

Becker, G. S. (1991). A treatise on the family. (Enlarged ed.). Cambridge, MA: Harvard University Press.

Beller, A. H., & Graham, J. W. (1993). Small change: The economics of child support. New Haven, CT: Yale University Press.

Besley, T., & Case, A. (1997). Women in politics: Finding instruments for incidence analysis. Unpublished manuscript, Princeton University.

Bloom, D., Conrad, C., & Miller, C. (1996). Child support and fathers' remarriage and fertility. Cambridge MA: National Bureau of Economic Research Working Paper.

Duncan, G., & Brooks-Gunn, J. (Eds.). (1997). Consequences of growing up poor. New York: Russell Sage Foundation.

Duncan, G., & Hoffman, S. (1990). Welfare benefits, economic opportunities and out-of-wedlock births among black teenage girls. Demography 27(4), 519-536.

Federal Interagency Forum on Child and Family Statistics. (1998). Nurturing fatherhood: Improving data and research on male fertility, family formation and fatherhood. Proceedings of the conference on fathering and male fertility. Washington DC.

Freeman, R. and Waldfogel, J. (1998). Does Child Support Enforcement Affect Male Labor Supply? in I. Garfinkel, S. McLanahan, D. Meyer, & J. Seltzer, (Eds.), Fathers under Fire: The revolution in child support enforcement, New York: Russell Sage.

Garfinkel, I. (1992). Assuring child support. New York: Russell Sage Foundation.

Garfinkel, I., & McLanahan, S. S. (1986). Single mothers and their children: A new American dilemma. Washington, DC: The Urban Institute Press.

Garfinkel, I., & McLanahan, S. S. (1994). Single-mother families, economic insecurity, and government policy. In S. Danziger, G. Sandefur, & D. Weinberg (Eds.), Confronting poverty: Prescriptions for change. Cambridge, MA: Harvard University Press.

Garfinkel, I., McLanahan, S.S. & Robins, P. (1994). The relationship between child support enforcement tools and child support outcomes. In I. Garfinkel, S. McLanahan, & P. Robins (Eds.), Child support and child well-being (pp. 133-170). Washington, DC: The Urban Institute Press.

Gaylin, D., Garfinkel, I., McLanahan, S. S., & Huang, C. C. (1996, April). Will child support enforcement reduce nonmarital childbearing? Paper presented at the annual meeting of the Population Association of America.

Hanson, T. (1995). Unpublished tables, Princeton University.

Klawitter, M. (1994). Child support awards and the earnings of divorced noncustodial fathers. Social Service Review, 68(3), 351-369.

Lichter, D., McLaughlin, D., Kephart, G. & Landry, D. (1992). Race and the retreat from marriage: A shortage of marriageable men. American Sociological Review, 57(6), 781-799.

Lundberg, S. J., & Plotnick, R. D. (1995). Adolescent premarital childbearing: Do economic incentives matter? Journal of Labor Economics, 13(2), 177-200.

Mare, R. D. and Winship, C. (1991). Socioeconomic change and the decline of marriage for blacks and whites. in C. Jencks & P. Peterson, (Eds.), The urban underclass. Washington, DC: The Brookings Institution.

Maynard, R. (1995). Teenage childbearing and welfare reform: Lessons from a decade of demonstration and evaluation research. Children and Youth Services Review, 17(1/2), 309-332.

McLanahan, S. and Casper, L. (1995). Growing diversity and inequality in the American family.

in R. Farley (Ed.), State of the union: America in the 1990s (Volume 2). New York: Russell Sage Foundation.

Miller, B. C. & Moore, K. A. (1990). Adolescent sexual behavior, pregnancy and parenting: Research through the 1980s. . Journal of Marriage and the Family, 52(4), 1025-1044.

Moffitt, R. (1992). Incentive effects of the U. S. welfare system: A review. Journal of Economic Literature, 30(1), 1-61.

Moffitt, R. (1998). The effect of welfare on marriage and fertility. In R. Moffitt (Ed.), Welfare, the family and reproductive behavior (pp. 50-97). Washington, DC: National Academy Press.

Moore, K., Miller, B. C., Glei, D., & Morrison, D. R. (1995). Adolescent sex, contraception, and childbearing: A review of recent research. Washington, DC: Child Trends, Inc.

Moynihan, D. P. (1965). The Negro family: The case for national action. Washington, DC: Office of Planning and Research, U. S. Department of Labor.

Murray, C. (1984). Losing ground: American social policy, 1950-1980. New York: Basic Books.

Murray, C. (1993). Welfare and the family: The U.S. experience. Journal of Labor Economics, 11(1), S226-S261.

Nichols-Casebolt, A., & Garfinkel, I. (1991). Trends in paternity adjudications and child support awards. Social Science Quarterly, 72(1), 83-97.

Nixon, L. A. (1997). The effect of child support enforcement on marital dissolution. Journal of Human Resources, 32(1), 159-181.

Office of Child Support Enforcement. (1982). Child support enforcement: Seventh annual report to Congress. Washington, DC: The Department of Health and Human Services Administration for Children and Families.

Schultz, T. P. (1994). Marital status and fertility in the United States: Welfare and labor market effects. Journal of Human Resources, 29(2), 637-669.

Seltzer, J. A., Schaeffer, N. C., & Charng, H. W. (1989). Family ties after divorce: The relationship between visiting and paying child support. Journal of Marriage and the Family, 51(4), 1013-1031.

Seltzer, J. A., McLanahan, S. S., & Hanson, T. (1998). Will child support enforcement increase father-child contact and parental conflict after separation? in I. Garfinkel, S. McLanahan, D. Meyer, & J. Seltzer, (Eds.), Fathers under Fire: The revolution in child support enforcement, New York: Russell Sage.

Sonenstein, F. L., Pleck, J., & Ku, L. (1994, May). Child support obligations and young men's contraceptive behavior: What do young men know? Does it matter? Paper presented at the annual meeting of the Population Association of America, Miami, FL.

Sorensen, E. (1997). A national profile of nonresidential fathers and their ability to pay child support. Journal of Marriage and the Family, 59, 785-797.

South, S. J. (1996). Mate availability and the transition to unwed motherhood: A paradox of population structure. Journal of Marriage and the Family, 58(2), 265-279.

South, S.J. and K.M. Lloyd. (1992). Marriage opportunities and family formation: Further implications of imbalanced sex ratios. Journal of Marriage and the Family 54(3), 440-451.

U. S. Department of Health and Human Services. (1995). Report to Congress on out-of-wedlock childbearing . Hyattsville, MD: U. S. Public Health Services.

U. S. House of Representatives. (1998). 1998 Green book: Background material and data on programs within the jurisdiction of the Committee on Ways and Means. Washington, DC:

U. S. Government Printing Office.

Ventura, S. J., Martin, J. A., Curtin, S. C., & Mathews, T. J. (1998). Report of final natality statistics, 1996 (Monthly Vital Statistics Report, vol. 46, no.11, supplement). Washington DC: National Center for Health Statistics.

Weiss, Y., & Willis, R. J. (1985). Children as collective goods in divorce settlements. Journal of Labor Economics, 3(3), 268-292.

Willis, R. J. (1994, September). A theory of out-of-wedlock childbearing. Paper presented at the Workshop on Expanding Frameworks for Fertility Research, National Research Council, Woods Hole, MA.

Willis, R. J., & Haaga, J. G. (1996). Economic approaches to understanding nonmarital fertility. In R. L. Casterline & K. Foote (Eds.), Fertility in the United States: New patterns, new theories. (A supplement to Population and Development Review, vol. 22) (pp. 67-86). New York: The Population Council.

Wilson, W. J. (1987). The truly disadvantaged. Chicago: University of Chicago Press.

Endnotes

[i] In the rest of the text, we use “premarital” in place of the more comprehensive “nonmarital” because the empirical work estimates models of the first teenage premarital birth. We do not examine out-of-wedlock births of women who divorced and then became pregnant.

[ii] We do not consider the choice between raising the child as an unmarried mother and placing the infant for adoption because the placement option is so rarely exercised.

[iii] In response to the 1996 welfare reforms, some states have relaxed this rule and allowed more child support to “pass through” to the mother, while others have reduced the pass through. During the time period we analyze, however, the pass-through was fixed at $50 by federal law.

[iv] Lundberg & Plotnick (1995) discuss underreporting of abortions in this sample.

[v] The appendix provides descriptive statistics on all explanatory variables.

[vi] An appendix available on request provides details on the source and nature of these variables.

[vii] Though they are not the focus of this study, findings on the policy and family background variables may be of interest. There are statistically significant relationships ($p < .05$) between living in a state with higher welfare benefits and a less liberal policy on public funding of abortions, and the likelihood of having a premarital birth. Background variables with a statistically significant association with having a premarital birth are living at age 14 in a mother-only family (+), living at age 14 with a mother who works outside the home (-), mother’s schooling (-), and being black (+). Complete regression outcomes are available upon request to the first author.