The Effect of Child Support Enforcement Efforts on Nonmarital Fertility

Geoffrey L. Wallace
La Follette School of Public Affairs
University of Wisconsin–Madison

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INTRODUCTION

Recent research suggests that unmarried women respond to financial incentives in making fertility decisions.[1] This research and increased efforts at child support enforcement during the 1990s raise the question of whether fertility decisions are responsive to the strength of child support enforcement efforts. From the point of view of economic theory, the answer to this question is not clear. On one hand, increased efforts at child support enforcement may lead men to take measures that reduce the likelihood of their becoming fathers, and nonmarital births may be reduced. Men may also be more likely to marry if they intend to become fathers. On the other hand, increased child support enforcement efforts reduce the cost to single women of having children and may lead to increases in nonmarital fertility.

In this project I use individual-level data from the 2001 panel of the Survey of Income and Program Participation (SIPP) along with state-level data on child support collection rates, welfare rules, and unemployment rates to assess whether the strength of state child support enforcement efforts has an effect on fertility and marriage among single women. I find little evidence that child support enforcement efforts have any effect on nonmarital fertility or marriage. Indeed, I find little in the way of plausible policy effects.

BACKGROUND

In 1975 Congress enacted the Child Support Enforcement and Paternity Establishment Program, which authorized federal matching funds that states could use to assist in establishing paternity and child support orders and in collecting support from noncustodial parents to offset welfare payments or to increase the resources available to single-parent families. Since its inception the program has undergone a

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[1]See, for example, Huang (2002), Garfinkel, Huang, McLanahan, and Garylin (2003), and Acs and Nelson (2004).

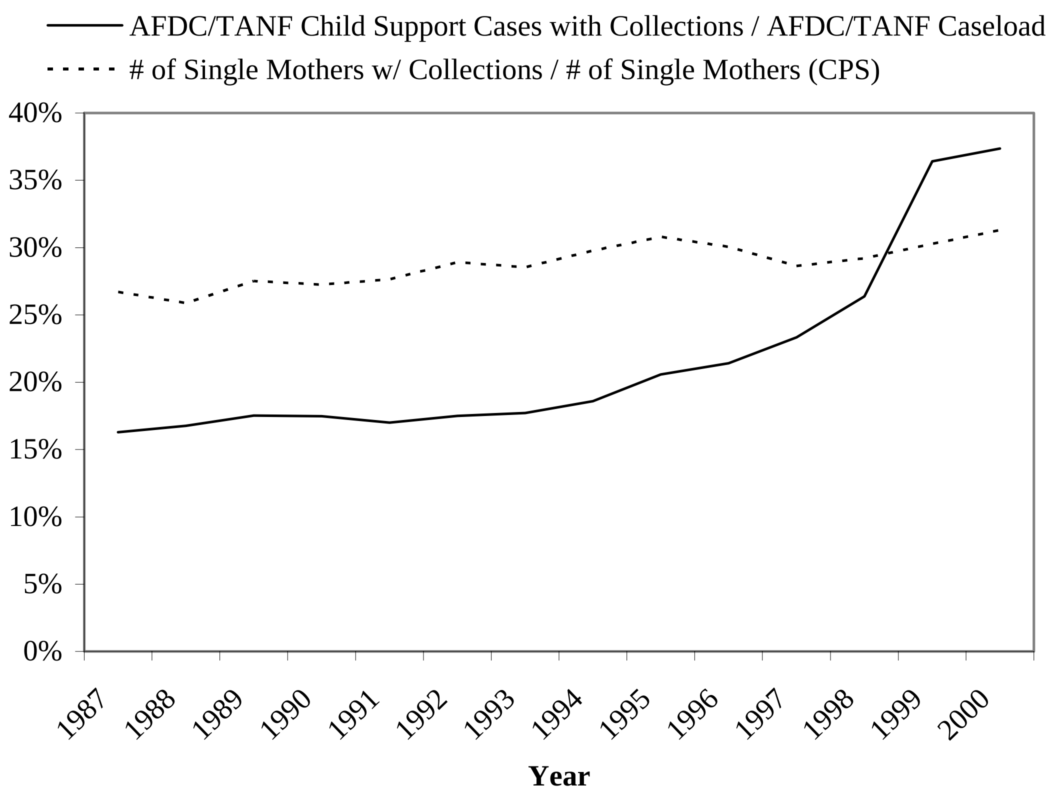
series of changes designed to aid the process of finding noncustodial parents, establishing paternity and support orders, and collecting child support.

Changes in the child support program in 1984 mandated that administrative systems be set up by the states to expedite the process of obtaining and enforcing child support orders and gave state child support enforcement agencies access to the IRS for the purpose of locating and verifying the income of noncustodial parents. The Family Support Act of 1988 contained a mandate that states attempt to establish paternity for all children under 18. To meet this mandate states were encouraged to set up administrative procedures for establishing paternity by genetic testing in cases where paternity was contested. As part of the 1996 welfare reform legislation, states were required to establish a database of new employees, and employers were required to provide the name and Social Security number of all new employees to the states and, by proxy, to a national new employee database that can be used to locate noncustodial parents for the establishment of paternity or enforcement of a support order. When a noncustodial parent who is delinquent in child support payments is located using this system, employers are immediately instructed to begin withholding child support from the parent's wages.[2]

The effect of these and other changes in child support enforcement practices can be seen in Figure 1, which plots two measures of the child support collection rate over time. The first measure is the number of AFDC/TANF child support enforcement cases with collections in a year, divided by the average monthly AFDC/TANF caseload. The second measure is the number of single mothers with collections in the March Current Population Survey (CPS), divided by the total number of single mothers in the CPS. Although both collection rates increase over time, the AFDC/TANF collection rate increased more, particularly after 1998. From the early 1990s to 2000 that rate more than doubled. Although limited in coverage to AFDC/TANF child support cases, this rate is more accurate than alternative administrative

[2]There are exceptions to this immediate withholding. If the noncustodial parent can show good cause, withholding may be delayed indefinitely. If the custodial and noncustodial parent can reach a suitable arrangement, immediate holding may be also be delayed.

FIGURE 1
Child Support Enforcement Measures by Year



collection rates, and is thought to be broadly consistent with overall collection rates. Both measures are comprehensive in that they reflect the evolution in combined efforts to locate noncustodial parents, to establish paternity and support orders, and to increase collections.

In attempting to assess the effect of increased child support enforcement efforts on nonmarital fertility, it is important to note that changes in the child support enforcement system did not occur in isolation. During the early and mid-1990s many states implemented experimental welfare reform programs under waivers from the Department of Health and Human Services, and after 1996 all states implemented TANF programs, which set a maximum limit of five years on welfare receipt and mandated work requirements. Either as part of waiver programs or as part of their TANF plans, a number of states adopted family caps, under which welfare mothers who give birth to an additional child do not receive an increase in assistance. Also, a series of expansions of the Earned Income Tax Credit (EITC) began in 1986. All of these policies, either directly or indirectly, could influence nonmarital fertility. In examining the effect of child support enforcement efforts, it is important to control for these other policy changes.

DATA

To construct the data used for this paper I merged individual-level data from the 2001 SIPP with state-level policy and economic variables. The SIPP is a longitudinal data set published by the U.S. Bureau of the Census. Once every four months SIPP participants are asked for information about their income, earnings, and program participation over the previous four months. In addition to the standard questionnaire that is administered every four months, one topical module questionnaire is administered. These topical modules obtain retrospective or more detailed information about a particular area of interest to researchers.

In this paper I make use of the second wave core data file (W2CDF) and the second wave topical module file (W2TMF). The W2CDF provides information on educational attainment, race and Hispanic origin, year and month of birth for each person in the sample, and residential location (state of residence and metro status). The W2TMF provides retrospective information on marital and fertility histories.

Using information in these files I first identified all women who were potentially eligible[3] for a first or second nonmarital birth between January 1987 and December 2000 and had all of their children living with them at time the topical module questionnaire was administered.[4] From this initial group of women I selected a sample of white, black, and Hispanic women who resided in states that are uniquely identified and well represented in the initial SIPP sample.[5] Women included in the sample were followed until they experienced a nonmarital birth, married, reached the end of the sample period, or reached the age of 45.[6]

Including only women who meet the criteria outlined above leaves a sample of 5,848 women spanning nearly 400,000 person months. Because of the nature of the sample design I do not observe first or second nonmarital birth intervals[7] for all of these women. Rather, I observe first birth intervals for 5,422 women who turned 16 between January 1987 and December 2000 and second birth intervals for 1,332 women during the same period. There are 801 women for whom I observe the time till the first nonmarital birth and at least part of the time till the second birth.

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[3]A woman is assumed to be eligible for first birth at age 16 and a second birth 9 months after giving birth to her first child. There were a number of cases where a woman gave birth to her first child prior to turning 16. These cases were treated as outliers and dropped from the sample.

[4]Because the topical module file only contains the year of birth of each female respondent's first and last child, it is impossible to determine the precise month in which her children were born. In pinpointing the timing of first births, I rely on the information contained in the core data file. In cases where the topical module file indicates that all children are living with their mother, the data file should contain a record for each of the mother's children indicating a year and month of birth. These birth months and years were then checked against the information on birth year of the first and last children contained in the module file to eliminate cases where births were before January 1987 or after December 2000, cases with twins, and cases where there are apparent inconsistencies between the two files.

[5]Main, Vermont, South Dakota, North Dakota, and Wyoming are not uniquely identified by the SIPP. Women residing in Alaska, Delaware, the District of Columbia, Hawaii, Idaho, Montana, Nevada, New Mexico, Rhode Island, and Utah were excluded because these states are not well represented in the SIPP.

[6]Exact marriage dates cannot be determined from the W2TMF. For reasons of privacy protection the public use W2TMF only includes the year of first marriage. In constructing the data used in this analysis I assigned women who were married in a particular year the earliest possible marriage data. So, for example, if the year a respondent was first married is recorded as 1990, I assign her a marriage date of January 1990.

[7]A birth interval is defined as the length of time starting when a women becomes eligible for a birth (at age 16 in the case of a first birth interval, or 9 months after her first birth in the case of a second birth interval) and ending with her next birth. This is not the standard definition of birth intervals used by demographers. In the demography literature a birth interval is typically defined as the length of time between two births.

Sample means and standard deviations of the variables used in this analysis are shown in Table 1 for the entire sample and by race (white or nonwhite). Time-varying variables are shown with a *t* subscript and reflect values for the first at-risk month. Most of the variables in Table 1 are self-explanatory, but some require additional explanation (shown on the table). “Family cap” indicates whether a state had a family cap in place, “other reform” indicates the presence of statewide welfare reform waivers other than a family cap, and TANF indicates whether a TANF program had been implemented.[8]

Both the AFDC/TANF collection rate and the CPS collection rate are comprehensive measures of the strength of child support enforcement efforts at the state level. They were described earlier, in connection with Figure 1. I use AFDC/TANF cases with collections as the basis for the first measure, because reliable information on the total number of cases with collections is not available over time.[9] Although the CPS collection rate has the advantage of reflecting collections across all single mothers, it is an estimate, and thus reflects a degree of measurement error.[10] Issues of coverage and measurement error aside, both of these rates should reflect efforts to increase the establishment of paternity and support orders as well as efforts to enforce support orders.

Figure 2 provides an estimate of the cumulative probability of a nonmarital birth. Because there is a one-to-one relationship between age and first birth, age is on the horizontal axis in Panel A. In Panel B, which describes second-birth patterns, the number of years between first and second births are on the horizontal axis. In Panel A we see striking differences between the first-birth patterns of never-married whites and nonwhites. At every age nonwhites are approximately twice as likely to have experienced a

[8]Information on welfare reform measures was pieced together from a variety of sources, most notably the Department of Health and Human Services, Office of the Assistant Secretary for Planning and Evaluation Web site, the State Policy Demonstration website, and the Urban Institute’s Welfare Rules Database.

[9]Historically, states have not kept accurate records on nonwelfare child support collections. See Huang et al. (2001) for a detailed discussion of this issue.

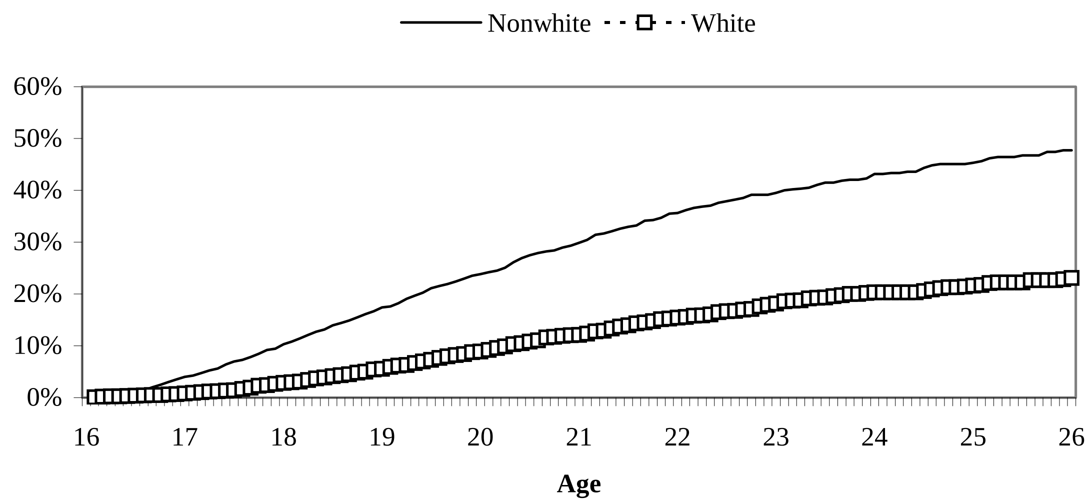
[10]In 2000, 95 percent confidence bands for the CPS enforcement rate ranged from ± 0.06 percentage points in California to ± 1.39 percentage points in Maine.

TABLE 1
Sample Means and Standard Deviations

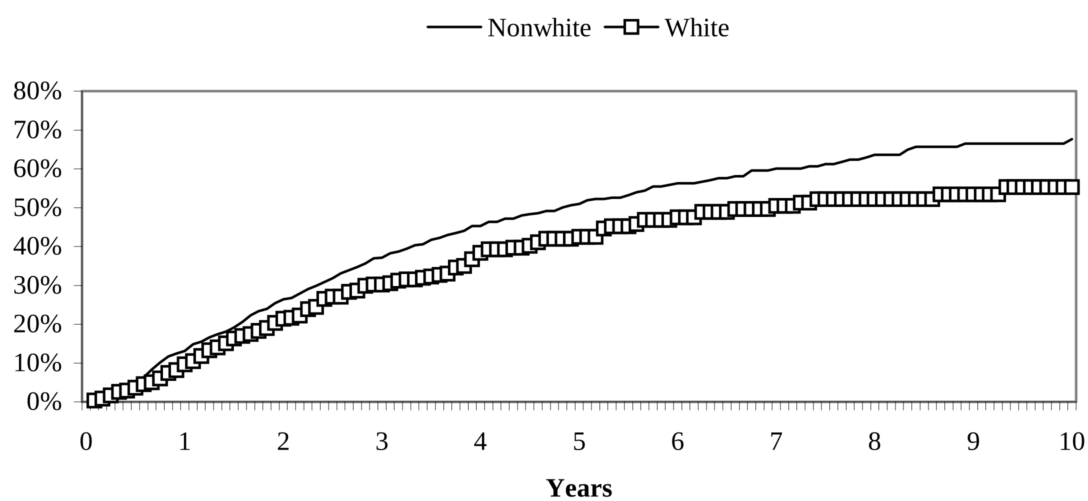
Variable	First Interval (*N=5,422)*	Second Interval (*N=1,332)*
AGE	16.000	22.5625
		(4.7331)
HIGH SCHOOL DROPOUT$_t$ (=1)	0.2353	0.2087
	(0.4242)	(0.4065)
HIGH SCHOOL GRADUATE$_t$ (=1)	0.2722	0.3806
	(0.4451)	(0.4857)
SOME COLLEGE$_t$ (=1)	0.4924	0.4107
	(0.5000)	(0.4921)
WHITE (=1)	0.6741	0.4602
	(0.4687)	(0.4986)
BLACK (=1)	0.1619	0.3491
	(0.3684)	(0.4768)
HISPANIC (=1)	0.1640	0.1907
	(0.3703)	(0.3930)
METRO RESIDENT (=1)	0.7942	0.7940
	(0.4043)	(0.4065)
FAMCAP$_t$	0.1627	0.1628
(=1 if the state in which a respondent resides has a family cap)	(0.3691)	(0.3826)
OTHER REFORMS$_t$	0.1121	0.1224
(=1 if the state in which a respondent resides has a statewide waiver other than a family cap)	(0.3156)	(0.3278)
TANF$_t$	0.3054	0.3093
(=1 if the state in which a respondent resides implemented a TANF program)	(0.4606)	(0.4624)
URATE	5.5646	5.6575
(State monthly unemployment rate)	(1.6196)	(1.6011)
AFDC/TANF collection rate	0.2215	0.2260
(Fraction of AFDC/TANF child support cases with collections)	(0.1359)	(0.1409)
CPS collection rate	0.2908	0.2877
(Fraction of single mothers with child support collections from the CPS)	(0.0762)	(0.0732)

FIGURE 2
Cumulative Probability of a Nonmarital Birth by Race

A. First Birth



B. Time till Second Birth



nonmarital birth. By age 26 an estimated 48 percent of nonwhites will have given birth nonmaritally, compared with only 23 percent of whites. The differences in cumulative second nonmarital birth probabilities are not as striking, but are still apparent. The second-birth probabilities are high: when a woman has experienced one nonmarital birth, the chance that she will experience another in the next ten years is greater than 50 percent.

METHODS

To evaluate the effect of child support enforcement efforts on nonmarital fertility I make use of an event-history model in which the length of a woman's birth interval is determined as a function of her monthly probabilities of having a birth and getting married. For both first and second birth intervals, monthly birth rates are specified as a function of variables thought to influence the likelihood of a birth in any given month. Included in these variables are demographic controls and policy variables, including measures of child support enforcement efforts. For first birth intervals, monthly marriage rates are also specified as a function of these same variables. Monthly marriage rates are not estimated for women in their second birth intervals. The reason for the difference in modeling treatment between first and second birth intervals is that most nonmarital first birth intervals end in marriage rather than in a nonmarital birth, whereas only a small number of second birth intervals end in marriage. Indeed, there are so few second birth intervals that end in marriage that it is not feasible to estimate monthly marriage rates for women who already have one nonmarital birth: there are simply not enough data available.

A detailed description of the estimation procedure is provided in Appendix A. We turn now to the analytic results.

RESULTS

The monthly probabilities of a nonmarital birth and marriage (for women in their first birth intervals) are modeled as a function of age, metro status, educational level,[11] the AFDC/TANF collection rate or the CPS collection rate, the state monthly unemployment rate, and indicators of a family cap implementation, implementation of a welfare reform waiver other than a family cap, and TANF implementation. Separate estimates of monthly birth rate specifications are obtained for whites and nonwhites (blacks and Hispanics).[12] In addition to the variables listed above, each of the estimated specifications also includes month, year, and state effects (dummy variables). Month effects are included to control for seasonal effects on birth and marriage.[13] Year effects are included in the birth and marriage rate specifications to control for changes in policy or exogenous behavioral changes that occurred at the national level over the course of the sample period that may affect nonmarital fertility or marriage.[14] State effects are included in the specifications as control for differences in unmeasured state-level factors that may affect fertility and/or marriage and are potentially correlated with the policy variables. These unmeasured state-level factors may include, but are not limited to, the availability of abortion services, differences in levels of child care subsidization and/or availability, and location-specific differences in preferences affecting demand for children. Because nominal AFDC/TANF benefit levels did not change

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[11]The education variables are omitted from first birth and marriage rate specifications. The reason for this omission has to do with the retrospective nature of the information used to construct the sample. Women are followed from the time they turn 16; women who turned 16 between January 1997 and December 2000 are included in the sample. The problem is that educational levels are only observed in late 2000 or early 2001. These levels may be a cause of particular fertility and marital history or may reflect a particular marital and fertility history. Because the educational levels may be both a cause and effect of marital and fertility histories, they are not included in the set of independent variables in the first birth and marriage rate. Because most women will have completed all of their schooling by the beginning of their second nonmarital birth interval, educational levels will not be affected by the timing of second nonmarital births. Thus, education levels are included in the second birth rate specifications.

[12]Sample sizes were sufficiently small to preclude estimating specifications for blacks and Hispanics separately.

[13]Month dummy variables are excluded from the marriage rate equations because all marriages are presumed to occur in January. See footnote 6 above.

[14]One policy change with implications for subsequent nonmarital fertility that occurred on the national level was a change to the structure of the Earned Income Tax Credit (EITC) that allowed for more generous subsidies for families with two or more children as compared with one child beginning in 1991.

much over the course of the late 1980s and 1990s, the state fixed effects also serve as a proxy for the baseline level of welfare generosity across states.

The estimation results for whites are shown in Table 2. The first two columns contain estimates of the determinants of the nonmarital birth rate during the first birth interval. Column (1) uses the AFDC/TANF collection rate as a measure of child support enforcement efforts, and column (2) uses the CPS collection rate. Columns (3) and (4) mirror columns (1) and (2) except that they contain the estimates of the determinants of the marriage rate during the first birth interval The column (1) specification was estimated simultaneously with the column (3) specification, and the column (2) specification was estimated simultaneously with the column (4) specification. Columns (5) and (6) contain estimates of the birth rate during the second birth interval.

Examining the estimates in columns (1) and (2) of Table 2, the only variable that has a statistically significant impact on the nonmarital birth rate among never-married white women is metro status: white women residing in metro areas are much less likely to have a nonmarital birth. None of the policy variables have statistically significant effects. The column 1 estimate indicates that the AFDC/TANF collection rate has a positive effect on first-birth rates among never-married whites. The coefficients shown in Table 2 can be interpreted as marginal fractional changes for small deviations in the independent variable. Thus, the column 1 estimate implies that a 0.10 increase in the AFDC/TANF collection rate would increase the first-birth rate for never-married whites by 3.8 percent (0.10 x 0.38 x 100). This is a very small impact. The column (2) estimate of the impact of child support is nearly as small in absolute terms but goes in the opposite direction. A 0.10 increase in the CPS collection rate would reduce the nonmarital birth rate by approximately 2.9 percent.

Moving to the marriage rate specifications in columns (3) and (4), we see that never-married white women residing in metro areas are also much less likely to get married in a given month. When the AFDC/TANF collection rate is used as the measure of child support enforcement efforts, increases in child support enforcement are associated with an increase in marriage. An increase in the AFDC/TANF

TABLE 2
Estimates of the Determinants of the Nonmarital Birth Rate by Birth Interval (Never-Married Whites)

	First Interval				Second Interval	
Variable	Birth Rate		Marriage Rate		Birth Rate	
	(1)	(2)	(3)	(4)	(5)	(6)
Age	—				-0.0965	-0.0940
					(0.1406)	(0.1406)
Age*age/100	—	—	—	—	0.1153	0.1112
					(0.2491)	(0.2493)
Metro	-0.5856**	-0.5860**	-0.5554**	-0.5454**	-0.4631*	-0.4601*
	(0.1210)	(0.1209)	(0.0844)	(0.0850)	(0.2519)	(0.2525)
High school graduate	—	—	—	—	-0.1643	-0.1629
					(0.2618)	(0.2621)
Some college	—	—	—	—	-0.6373**	-0.6362**
					(0.2696)	(0.2689)
AFDC/TANF collection rate	0.3811	—	0.4699*		0.2450	—
	(0.4139)		(0.2837)		(0.3521)	
CPS collection rate	—	-0.2964	—	-0.1076	—	1.3201
		(0.9590)		(0.7188)		(1.8858)
Family cap	-0.0149	-0.0436	-0.0639	-0.0952	0.6654*	0.6733*
	(0.1732)	(0.1746)	(0.1489)	(0.1502)	(0.3521)	(0.3368)
Other reforms	0.1951	0.1870	-0.5223**	-0.5354	-0.2375	-0.2823
	(0.1984)	(0.2003)	(0.1714)	(0.1735)	(0.3651)	(0.3615)
TANF	0.2419	0.2369	-1.4142**	-1.4155**	-1.2006**	-1.2345**
	(0.2634)	(0.2648)	(0.2026)	(0.2029)	(0.5630)	(0.5616)
State unemployment rate	-0.0512	-0.0503	1.4340	1.4156	0.0927	0.0929
	(0.0774)	(0.0775)	(0.0535)**	(0.0531)	(0.0996)	(0.0983)
State effects	Included	Included	Included	Included	Included	Included
Year effects	Included	Included	Included	Included	Included	Included
Month effects	Included	Included	Omitted	Omitted	Included	Included
Duration controls	Included	Included	Included	Included	Included	Included

** Statistically significant at a 0.05 significance level.
* Statistically significant at a 0.10 significance level.

collection rate of 0.10 would increase the marriage rate by nearly 5 percent. This effect is statistically significant at a significance level of 0.10. Interestingly, the coefficient on the CPS collection rate has the opposite sign and is not statistically different from zero. The only other policy variable in either the column (3) or (4) specification that has a statistically significant effect is the TANF indicator. Although the estimates suggest that marriage rates are substantially lower under TANF than they were pre-TANF, the coefficient is identified primarily on the basis of a pre-1998 and post-1998 comparison. Because the model also contains year dummy variables which are highly correlated with the TANF interpretation of the coefficient, it is difficult to determine whether there is any information in the TANF coefficient estimate.

The estimates of the determinants of second nonmarital births rates for whites, shown in columns (5) and (6) of Table 2, indicate that older women, women residing in metropolitan areas, and women who completed more schooling have lower second nonmarital birth rates. As was the case with first births, many of the state-level variables are imprecisely estimated. There is no evidence that either measure of child support enforcement has an effect on second-birth rates among never-married white women. The CPS collection rate does have a very large impact, but its standard error is also very large. The only state-level variables that are statistically significant are the indicators of family cap and TANF implementation. The coefficient estimate associated with the family cap variables implies that family cap implementation leads to an increase, rather than reduction, in second nonmarital births. This is the opposite of the intended effect of the policy. The coefficient on the TANF variable is consistent with conventional wisdom concerning the effects of welfare reform on nonmarital fertility, but as alluded to above, caution must be used in interpreting the TANF coefficients because of the collinearity of the TANF variable with the year effects.

Estimates for nonwhites (blacks and Hispanics) are shown in Table 3. All else equal, Hispanic women have lower nonmarital first-birth rates than black women. Residing in a metro area is also associated with lower nonmarital first-birth rates. The AFDC/TANF collection rate is estimated to have a

TABLE 3
Estimates of the Determinants of the Nonmarital Birth Rate by Birth Interval (Never-Married Nonwhites)

	First Interval				Second Interval	
Variable	Birth Rate		Marriage Rate		Birth Rate	
	(1)	(2)	(3)	(4)	(5)	(6)
Age	—	—	—	—	0.0361	0.0362
					(0.1344)	(0.1341)
Age*age/100	—	—	—	—	-0.2058	-0.2060
					(0.2484)	(0.2475)
Hispanic (vs. black)	-0.6132**	-0.6175**	1.0240**	1.0239**	0.1218	0.1177
	(0.1132)	(0.1138)	(0.1806)	(0.1803)	(0.1825)	(0.1824)
Metro	-0.2862*	-0.2842*	-0.4187*	-0.4113*	-0.0055	-0.0023
	(0.1624)	(0.1628)	(0.2235)	(0.2233)	(0.2355)	(0.2343)
High school graduate	—	—	—	—	-0.5789**	-0.5821**
					(0.1587)	(0.1581)
Some college	—	—	—	—	-0.6742**	-0.6740**
					(0.1870)	(0.1873)
AFDC/TANF collection rate	0.9457**	—	-1.0191	—	0.3882	—
	(0.4076)		(0.7730)		(0.8721)	
CPS collection rate	—	0.3451	—	0.7056	—	0.3076
		(1.0968)		(1.7754)		(1.5855)
Family cap	0.2233	0.2020	-0.3025	-0.2439	-0.2178	-0.2029
	(0.1725)	(0.1742)	(0.2821)	(0.2853)	(0.2488)	(0.2519)
Other reforms	-0.1699	-0.1988	-2.0272**	-1.9588**	0.0192	-0.0022
	(0.1817)	(0.1818)	(0.3208)	(0.3123)	(0.2703)	(0.2738)
TANF	0.2941	0.2766	-1.4671	-1.4144	0.5687	0.5526
	(0.2745)	(0.2745)	(0.4273)	(0.4175)	(0.4904)	(0.4903)
State unemployment rate	0.0867	0.0918	1.3681**	1.3709**	0.0835	0.0874
	(0.0728)	(0.0723)	(0.1137)	(0.1154)	(0.0918)	(0.0919)
State effects	Included	Included	Included	Included	Included	Included
Year effects	Included	Included	Included	Included	Included	Included
Month effects	Included	Included	Omitted	Omitted	Included	Included
Duration controls	Included	Included	Included	Included	Included	Included

** Statistically significant at a 0.05 significance level.
* Statistically significant at a 0.05 significance level.

very large positive, and statistically significant, effect on the first-birth rate among nonwhites. The magnitude of the coefficient implies that a 0.10 increase in the AFDC/TANF collection rate would increase the nonmarital first-birth rate by nearly 10 percent. The effect of the CPS collection rate in column (2) is positive, but it is also small and imprecisely estimated. None of the other policy variables in the nonmarital first-birth rate specifications are statistically different from zero.

The marriage rate equation estimates for nonwhites are shown in columns (3) and (4) of Table 3. These estimates must be interpreted carefully because the estimates are not based on a large number of marriages. There were 861 white women who were married prior to having a first birth, whereas only 288 nonwhite women were married prior to having their first birth. The estimates suggest that Hispanic women are far more likely to marry than black women. They also suggest that women in metro areas are less likely to marry. There is no strong evidence to suggest that child support collection affects marriage rates among never-married nonwhites. The coefficient estimates on the child support variables in the nonwhite marriage rate equations are large, but they are also of opposite signs and imprecisely estimated. The only policy variable that is statistically significant is the indicator of a pre-1996 waiver other than a family cap. This coefficient suggests that these reforms led to decreases in marriage among never-married black women without children. As these women are only indirectly or tangentially affected by such reforms, this effect must be treated with suspicion. One other variable that has a strong estimated effect on marriage rates among black and Hispanic women is the state monthly unemployment rate. Increases in state monthly unemployment rates are associated with increases in marriage rates among never-married nonwhites with no children.

The estimates of the determinants of a second nonmarital birth among nonwhites appear in columns (5) and (6) of Table 3. Only the educational level variables are statistically significant in these specifications. Graduation from high school or completing some post-high school training dramatically reduces the probability of a second nonmarital birth among black and Hispanic women. Both child support variables are associated with an increase in second nonmarital birth rates, but the standard errors associated with the coefficient estimates are large and hypotheses that the true coefficient values are zero cannot be rejected.

CONCLUSION

In this paper I have attempted to assess the impact of child support enforcement efforts on nonmarital fertility. From the point of view of economic theory, the effect of child support enforcement on fertility is ambiguous. On one hand, in states that vigorously pursue child support enforcement, there is an increased likelihood of an additional stream of income for female-headed families. All else equal, these enforcement efforts may have the unintended consequence of reducing the cost of nonmarital childbearing by providing an income stream to single mothers, thereby reducing the cost of nonmarital childbearing. On the other hand, in states that are aggressive in establishing and collecting child support orders, there is an increased incentive for men to forgo nonmarital parenthood. This incentive might lead to reduced birth rates among women. Aggressive enforcement efforts may also have a positive effect on marriage formation, as parents believe it is less costly to provide support for one household than two.

I examined the impact of state-level child support enforcement efforts using a sample constructed from the 2001 SIPP. The empirical strategy involved estimating first and second nonmarital birth rates and marriage rates by race. Birth rates and marriage rates were modeled as a function of individual-level demographic variables and state-level policy and economic variables, including two measures of child support collection rates. I find little evidence that child support enforcement efforts have any impact on nonmarital fertility. There is a large and statistically significant positive effect of child support collection rates on first-birth rates among nonwhite women, but most of the estimated child support enforcement effects were small and/or imprecisely estimated.

Overall the results indicate a very limited role for policy, be it child support enforcement policy or welfare policy, to influence fertility outcomes. This lack of measurable policy effects is both good and bad news for policy makers. Although increasing child support collections may increase the resources available to single-parent families and/or offset the cost of state TANF programs, such actions will not lead to increases in marriage or reductions in nonmarital fertility. On the other hand, states that aggressively pursue child support collections need not be concerned about providing an incentive for nonmarital births.

Appendix A

Estimation Procedure

The primary approach toward estimation in this analysis is taken from Prentice and Gloeckler (1978). They provide an extension to the Cox proportional hazard model, which allows estimation of grouped or discrete duration data in a manner that does not require making functional form assumptions concerning the parametric form of the baseline hazard. A primary advantage of this model is that the coefficients can be readily interpreted as marginal percentage changes.

In the analysis, a nonmarital birth interval can end in two ways: in the birth of a second child, or in marriage. I am interested in estimating the likelihood of a second nonmarital birth and a marriage for women who have no prior fertility history. For women with one prior nonmarital birth I am only interested in estimating the likelihood of a second nonmarital birth. The model with two risks (the risk of a nonmarital birth and the risk of marriage) is significantly more complex than the model with only one risk, but both models use the same general framework. In the interest of providing a flavor for the estimation without being overly complex, I outline the less complex model below.

Consider the chance that an individual's birth falls in the interval of time $[t, t+h)$, conditional on the individual having not giving birth until time t. The limit of this conditional probability, $h \to \infty$, is the birth rate. More formally,

$$\lim_{t \to \infty} \frac{\Pr(t <= \tau < t + h \mid \tau > t)}{h} = \lambda(t)$$

The proportional hazard assumption amounts to a functional form restriction on $\lambda(t)$. More specifically, the proportional hazard model assumes that

$$\lambda(t) = \lambda_0(t) \cdot \exp\left(x(t) \cdot \beta\right)$$

where the function $\lambda_0(t)$ is the baseline hazard, $x(t)$ is a vector of covariates that are allowed to depend on time, and β is a vector of parameters. Next, consider the probability that an individual has a birth interval of length t or greater. This probability is the survivor function. In the proportional hazard model the survivor function takes the form

$$S(t) = \exp\left(-\int_0^t \lambda(u)du \right).$$

Assuming the vector of covariates is constant over the interval $[j, j+1)$ for all j, $S(t)$ can be rewritten as

$$S(t) = \exp\left(-\sum_{j=0}^{t} \int_j^{j+1} \lambda(u)du \right) = \exp\left(-\sum_{j=0}^{t} \exp\left[\gamma(j) + x(j)\cdot\beta\right] \right)$$

where

$$\gamma(j) = \ln\left[\int_j^{j+1} \lambda_0(u)du \right].$$

Once the survivor function is known, the likelihood function follows straightforwardly. Consider a sample consisting of n birth intervals. Let T_i denote the length of the ith individual's birth interval. These birth intervals can be either uncensored or censored. For censored birth intervals I adopt the convention that the person is known not to have given birth until the end of the $T_i th$ period. In addition to the length of birth intervals a sequence of explanatory variables, $x(t)$, $(t = 1,2..,T)$, are observed for each individual in the sample. Letting $\delta_i = 0$ for all right-censored observations and $\partial_i = 1$ for all observations with completed birth intervals, the likelihood function is

$$L(\gamma, \beta) = \prod_{i=1}^{n} \left[1 - \delta_i \exp\{-\exp[\gamma(T_i) + x_i(T_i)\cdot\beta]\}\right]$$

$$\times \prod_{t=1}^{T_i - \delta_i} \left[\exp\left\{ -\sum \exp[\gamma(t) + x_i(t)\cdot\beta] \right\} \right]. \qquad (1)$$

Note that the first part of equation (1) is the probability of giving birth in period T_i, assuming the birth interval was not censored, while the second part of equation (1) is the probability that a birth interval lasts for at least $T_i - \delta_i$ months. Given (1), the log likelihood function is provided by

$$l(\gamma, \beta) = \sum_{i=1}^{n} \left\{ \ln\left[1 - \delta_i \exp\{\gamma(T_i) + x(T_i) \cdot \beta\}\right] - \sum_{t=1}^{T_i - \delta_i} \exp\{\gamma(t) + x_i(t) \cdot \beta\} \right\}. \qquad (2)$$

For the purpose of this analysis, several restrictions on $\gamma(\cdot)$ will be imposed. The approach taken in this analysis is to let $\gamma(t)$ take the form of a step function where there is a potential for a step every 12 months. I also artificially censor individuals with histories greater than 8 years. This censoring eliminates the need to estimate the γ 's associated with large numbers of periods where there are not many births and results in a loss of only a minimum of information.

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