The Effect of Incremental Benefit Levels on Births to AFDC Recipients

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Abstract

We examine the relationship between fertility and incremental AFDC benefits using the 1990 Panel of the Survey of Income and Program Participation (SIPP). Estimating a logit equation for the probability of a higher-order birth among a sample of AFDC recipients, we find a positive coefficient (although statistically insignificant) on the incremental AFDC benefit level. However, we find a positive correlation between incremental benefits and fertility for several nonrecipient comparison groups which is larger than the positive correlation for AFDC recipients. This finding suggests that the previously estimated relationship between incremental benefits and fertility among AFDC recipients is largely the result of a spurious correlation. We find similar results among whites, blacks, and never-married women, but less consistent results among Hispanics and divorced or separated women. We infer from these results that family cap policies, which eliminate the incremental benefits entitled to AFDC recipients who have additional children, are not likely to result in a large reduction in the number of out-of-wedlock births to AFDC recipients.

INTRODUCTION

The family cap policy or child exclusion policy is one of the most hotly debated topics on the current welfare reform agenda. This policy eliminates the extra monetary benefits entitled to women who have additional children while receiving the public assistance program Aid to Families with Dependent Children (AFDC). Historically, all states provided incremental increases to AFDC benefits for additional children. Although not included as a component of the Personal Responsibility and Work Opportunity Reconciliation Act of 1996, 16 states currently have family cap policies as part of their waiver demonstrations.[1] In 1990, prior to the implementation of these policies, incremental benefits ranged from $24 to $134 depending on the state and family size.

[1] See Wiseman [1996] for a description of current state waiver provisions.

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Supporters of the family cap policy argue that the elimination of incremental benefits will reduce the incentive among AFDC recipients to have additional out-of-wedlock children. This argument is supported by the standard economic model of fertility which posits that the demand for children decreases when the net cost of having children increases.[2] Because the removal of incremental benefits increases the net cost of having children, it is suggested that family cap policies will reduce fertility. This theory, however, does not offer a prediction regarding the size of the effect on fertility. For example, if incremental benefit levels are low relative to the perceived costs of raising children, their effect on fertility may be small.[3]

The few previous empirical studies of the effect of incremental AFDC benefits on higher order births provide mixed results.[4] These studies use the variation across states, time, or both to estimate the size of this effect.[5] Using a sample of young women with one child from the National Longitudinal Survey of Youth (NLSY), Acs [1996] finds that the level of the incremental benefit does not have a statistically significant effect on the probability of a second birth for either blacks or nonblacks. In contrast, Argys and Rees [1996], also using NLSY data, find some evidence that incremental benefits have a positive effect on the fertility of AFDC recipients. Powers [1994] finds a small positive and statistically significant effect of the incremental benefit on the probability of a subsequent birth over a five-year interval among a sample of unmarried mothers from the National Longitudinal Survey of Women (NLSW). Robins and Fronstin [1996], using a sample of never-married women from the 1980 to 1988 Current Population Surveys (CPS), find that the incremental benefit level for a second child has a positive and statistically significant effect on fertility for black and Hispanic women, but not for white women. Finally, Grogger and Bronars [1996], using data on unwed mothers of twins and singletons from the 1980 Census, do not find evidence of a positive effect of incremental benefits on fertility. Evidently, the literature has not reached a consensus regarding the effect of incremental benefits on out-of-wedlock births among AFDC recipients.

The main goal of this article is to estimate the size of the effect of incremental benefits on the fertility of AFDC recipients. If we find a large positive effect of incremental benefits on fertility then we can infer that the implementation of family cap policies will likely result in a large reduction in the number of out-of-wedlock births to AFDC recipients. The finding of either a small positive effect or no effect suggests that family cap policies will likely fail in their goal of reducing out-of-wedlock births among AFDC recipients.

[2] See Becker [1981].

[3] Other factors potentially contributing to a small effect include the unplanned nature of many births, important nonpecuniary factors that affect the birth decision, and the short expected duration of welfare receipt for many participants.

[4] See Moffitt [1992] for a review of the studies focusing on the effect of AFDC benefits on first births. These studies also provide mixed results.

[5] In addition to these national-level studies, a few state-level studies exist. Keefe [1983] finds that the large increase in total and incremental AFDC benefits from 1970 to 1971 in California did not increase fertility among recipients in the state. Rank [1989] finds that AFDC recipients in Wisconsin have lower birthrates than women in the general population. Camasso [1995] reports preliminary results from an experimental evaluation of the family cap policy in New Jersey which show that there is no significant difference between the birthrate of an experimental group of AFDC recipients who do not receive incremental benefits and the birthrate of a control group who receives these benefits.

In this article, we contribute to the literature by using a special panel of the Survey of Income and Program Participation (SIPP) and a different estimation technique than has been used in previous studies. The SIPP does not restrict us to include only young or never-married women in our sample and allows us to identify AFDC receipt and marital status before and after childbirth. The latter characteristic of the SIPP is needed to implement the quasi-experimental design used in our analysis to estimate the effect of incremental benefits on fertility. In particular, we use several comparison groups, consisting of women who are unlikely to respond to variation in incremental benefits, to estimate the true effect of incremental benefits on the probability of higher-order births to AFDC recipients.[6]

Our analysis of the effect of AFDC incremental benefits on higher-order births provides several important findings. First, there does not appear to be a clear relationship between incremental benefit levels and higher-order birthrates among AFDC recipients in the raw data. Second, using a sample of AFDC recipients, we find a positive coefficient (although statistically insignificant) on the incremental AFDC benefit level in a logit equation determining the probability of a higher-order birth. Using a quasi-experimental design, however, we find evidence suggesting that this positive coefficient represents a spurious correlation between incremental benefits and the fertility of AFDC recipients. Finally, we find similar results among whites, blacks, and never-married women, but less consistent results among Hispanics and divorced or separated women. Overall, our results do not provide evidence of a large positive effect of incremental benefits on the fertility of AFDC recipients.

DATA

To examine the effect of incremental AFDC benefits on higher-order births, we use microdata from the 1990 Survey of Income and Program Participation (SIPP). The SIPP was created jointly by the Department of Health and Human Services and the Census Bureau to become a major source of information on the demographic and economic situation of the United States. One of the most important goals of the SIPP is to collect an extensive amount of data on the receipt of governmental programs, including AFDC.

The 1990 Panel contains approximately 21,500 households. For each adult resident in these households, it includes up to two and a half years of monthly data covering the period from late 1989 to early 1992. This panel of the SIPP contains an oversample of households headed by females, blacks, and Hispanics which is especially useful for our analysis. Using weights provided by the SIPP, our sample is representative of all civilian noninstitutionalized households in the United States.

We create our analysis sample by including all women between the ages of 15 and 44 who have at least one child. The length of time covered by the 1990 SIPP allows us to include up to two nonoverlapping one-year intervals for each

[6] Note that we examine only second or higher-order births in this analysis because family cap policies are intended to reduce subsequent births among AFDC recipients, not necessarily to avert first births among potential AFDC recipients. Women not currently receiving AFDC are actually eligible for the full guarantee which increases with family size. See Donovan [1995] for a complete description.

individual. This creates a sample of 13,512 person-years representing one to two observations for 7,474 different women.

We measure fertility by examining whether a mother has a child during a one-year interval. For our purposes, the birthrate is defined as the percent of mothers who have a child less than one year old at the end of the one-year interval.[7] Individual-level characteristics are measured in the month prior to this one-year interval. For example, we define AFDC recipients as those women who report receiving AFDC benefits in the month prior to the birth interval. To the SIPP, we append several measures of state-level demographic and economic characteristics as well as measures of state policies regarding the availability of abortion and contraceptive services. These state-level variables, which we describe in more detail later, are measured in 1990 or 1991 depending on the year of the observation.[8] All monetary variables are reported in constant 1990 dollars.

DESCRIPTIVE RESULTS

Birthrates

Birthrates vary tremendously across demographic groups in the United States. The results reported in Table 1 provide evidence of this variation. We report birthrates for various subgroups of our sample of women (ages 15 to 44) who currently have at least one child. We divide the sample into three groups that we utilize in the estimation section of the article: (a) single mothers who receive AFDC; (b) single mothers who do not receive AFDC; and (c) married mothers who do not receive AFDC.[9] Among these three groups, unmarried AFDC recipients have the highest probability of giving birth during a one-year interval (0.087), whereas single mothers who do not receive AFDC have the lowest probability (0.034). This evidence might indicate that AFDC receipt among single women is associated with a higher birthrate. The reported birthrates, however, do not control for other factors that are potentially correlated with AFDC receipt.[10] In addition, never-married women have much higher birthrates than divorced or separated women.

Among both groups of single women, birthrates vary substantially by race; however, there is much less variation by race among married women.[11] For example, Hispanic single mothers have birthrates that are 2.6 to 2.9 times larger than those of white single mothers. In contrast, the birthrates of white, black, and Hispanic married mothers do not differ by more than 30 percent.

[7] We count twins and other multiple births as a single birth because of their unanticipated nature.

[8] There are a few exceptions. The restriction on teen abortions and the Medicaid funding for abortion variables are measured in 1990, and the number of abortion providers variable is measured in 1991.

[9] We exclude married women who receive AFDC-UP (Unemployed Parent) benefits from our sample. Families participating in this program comprise only 7.3 percent of the total families receiving AFDC [U.S. Congress, House Committee on Ways and Means, 1994, Table 10-24].

[10] See Table A.1 in the Appendix for evidence of the large differences between these groups in the mean values of many individual-level characteristics.

[11] We define our racial groups as follows. White is defined as those who report a white race, but do not report a Hispanic ancestry. Black is defined as those who report a black race, and Hispanic is defined as those who report a Hispanic ancestry, but do not report a black race.

Table 1. Birthrates among women with children (ages 15–44), 1990 Survey of Income and Program Participation (SIPP).

	Single AFDC recipients		Single nonrecipients		Married nonrecipients	
	Birthrate	Sample size	Birthrate	Sample size	Birthrate	Sample size
Total	0.087	1120	0.034	2734	0.068	9658
Marital status						
Never married	0.110	658	0.055	807	—	—
Widowed	0.066	16	0.052	131	—	—
Divorced	0.053	250	0.021	1308	—	—
Separated	0.052	196	0.029	488	—	—
Race						
White	0.050	383	0.025	1600	0.067	7354
Black	0.101	520	0.035	791	0.063	818
Hispanic	0.144	170	0.067	287	0.082	1060
Asian, Native American, or other race	0.059	47	0.147	56	0.069	426
Age group						
Ages 15–24	0.156	336	0.087	377	0.164	696
Ages 25–34	0.072	532	0.043	1113	0.100	4418
Ages 35–44	0.018	252	0.057	1244	0.020	4544
Number of children						
1 child present	0.103	410	0.044	1462	0.121	3177
2 children present	0.063	358	0.022	899	0.044	4198
3 children present	0.092	217	0.024	277	0.030	1653
4 or more children present	0.087	135	0.025	96	0.048	630

Notes: The birthrate is the percent of women who have a child during the one-year interval. All birthrates are calculated using weights provided by SIPP.

For each of our three main groups, birthrates decline with the age of the mother. In addition, birthrates are the highest for women who only have one child and change relatively little across larger numbers of children.

Incremental Benefits in the AFDC Program

AFDC is the primary cash assistance program in the United States for single women with children. Historically, the program has been funded jointly by the states and federal government, but has been administered solely by the states. In the recently enacted welfare reform legislation, states have even more flexibility in their AFDC programs, which is funded through federal block grants. Because benefit levels are determined by each state, there currently exists substantial variation in benefit levels across states and family sizes. In this article, we use this variation in benefit levels to identify the effect of incremental AFDC benefits on higher-order births.

In Table 2, we report the maximum AFDC benefit level for a family of two and the incremental AFDC benefit levels for several family sizes in each state

Table 2. Maximum total and incremental AFDC benefit levels by state and family size (1990).

State	Maximum guarantee ($) (1 child)	Incremental benefit ($)				Incremental benefit/ maximum guarantee (%) (1 child)
		1 Child	2 Children	3 Children	4 Children	
Alabama	88	30	29	30	29	34.1
Alaska	752	94	94	94	94	12.5
Arizona	233	60	60	59	60	25.8
Arkansas	162	42	43	39	45	25.9
California	560	134	130	116	117	23.9
Colorado	280	76	76	80	78	27.1
Connecticut	524	125	107	97	101	23.9
Delaware	265	68	69	68	68	25.7
District of Columbia	321	88	90	76	101	27.4
Florida	225	69	52	54	53	30.7
Georgia	229	44	49	47	31	19.2
Hawaii	480	122	123	121	123	25.4
Idaho	254	63	40	42	34	24.8
Illinois	268	99	47	71	60	36.9
Indiana	229	59	58	59	58	25.8
Iowa	347	63	66	51	60	18.2
Kansas	338	71	61	55	55	21.0
Kentucky	196	32	57	48	43	16.3
Louisiana	138	52	44	43	39	37.7
Maine	337	116	116	116	116	34.4
Maryland	309	87	81	75	56	28.2
Massachusetts	446	93	89	92	94	20.9
Michigan	438	93	107	99	139	21.2
Minnesota	437	95	89	76	76	21.7
Mississippi	96	24	24	24	24	25.0
Missouri	232	57	49	46	43	24.6
Montana	286	73	74	74	73	25.5
Nebraska	293	71	71	71	71	24.2
Nevada	270	60	60	60	60	22.2
New Hampshire	442	64	57	55	75	14.5
New Jersey	322	102	64	64	64	31.7
New Mexico	210	54	53	54	53	25.7
New York	522	118	116	119	86	22.6
North Carolina	236	36	25	27	25	15.3
North Dakota	313	73	86	65	55	23.3
Ohio	274	60	79	70	55	21.9
Oklahoma	252	73	78	69	67	29.0
Oregon	369	63	94	90	87	17.1
Pennsylvania	330	91	93	93	80	27.6
Rhode Island	440	103	77	76	88	23.4
South Carolina	165	41	42	42	42	24.8
South Dakota	333	44	44	44	44	13.2
Tennessee	141	43	40	38	41	30.5
Texas	158	26	37	25	38	16.5
Utah	310	77	65	63	52	24.8
Vermont	556	106	80	92	58	19.1
Virginia	294	60	56	78	30	20.4
Washington	404	97	88	90	92	24.0
West Virginia	201	48	63	48	53	23.9
Wisconsin	440	77	100	91	58	17.5
Wyoming	320	40	30	60	60	12.5
State average	315	72	69	67	65	22.8

Note: All benefit levels are taken from the 1990 *Green Book,* U.S. Congress, House of Representatives, Committee on Ways and Means.

in 1990.[12] The total monthly AFDC benefit for a women with one child ranges from a low of $88 in Alabama to a high of $752 in Alaska. Incremental AFDC benefits also vary markedly across states. In 1990, AFDC participants with one child received from $24 in Mississippi to $134 in California for having a second child. These incremental benefit levels are partially related to the guarantee level in the state, but there is some variation. The incremental benefit measured as a percent of the maximum AFDC benefit for a family of two ranges from 12.5 percent in Alaska and Wyoming to 37.7 percent in Louisiana.

In addition to the variation across states, incremental AFDC benefits vary across family sizes. For many states, the amount of the incremental benefit can increase or decrease notably from one family size to the next. A few states have substantially different incremental benefit levels across family sizes. For example, AFDC recipients in Illinois receive $99, $47, and $71 for their second, third, and fourth child, respectively, and AFDC recipients in Oregon receive $63 and $94 for their second and third child, respectively.

Birthrates and AFDC Benefits

As a first pass at determining the potential effectiveness of family cap policies, we examine the simple relationship between birthrates and incremental benefits in our data. At the top of Figure 1, we plot the birthrate for various categories of incremental benefit levels in our sample of single mothers receiving AFDC. The results indicate that the birthrate among AFDC participants falls within a fairly narrow range across these incremental benefit level categories. The birthrate varies from a low of 7.8 percent for women who receive $120 to $139 if they have another child to a high of 10.2 percent for women who receive $40 to $59.[13] There does not appear to be a clear pattern of birthrates across incremental benefit levels among AFDC recipients. In the bottom two panels of Figure 1, we plot birthrates for single mothers who do not receive AFDC and married mothers who do not receive AFDC. Among single nonrecipients, birthrates appear to rise, fall, and rise again as the incremental benefit increases. Birthrates clearly rise with incremental benefits for married nonrecipients. Overall, we do not find evidence of a strong positive relationship between incremental benefits and fertility among AFDC recipients in the raw data.

ESTIMATION

The Determinants of Births among AFDC Recipients

The results presented in the previous section, although suggestive, do not rule out the possibility that incremental AFDC benefits affect fertility. After controlling for demographic and policy-related factors that differ across states and family sizes, we may discover an important relationship between incremental AFDC benefits and higher-order birthrates. We address this concern by estimating a model of the birth decision for AFDC participants. Theoretically, women decide to have an additional child if the expected utility associated

[12] Although not reported, we find similar differences across states and family sizes for incremental benefits in 1991. There is some additional variation in incremental AFDC benefits over time, but this is small relative to the variation in each year.

[13] The difference between the estimates is not statistically significant at the $\alpha = .05$ level.

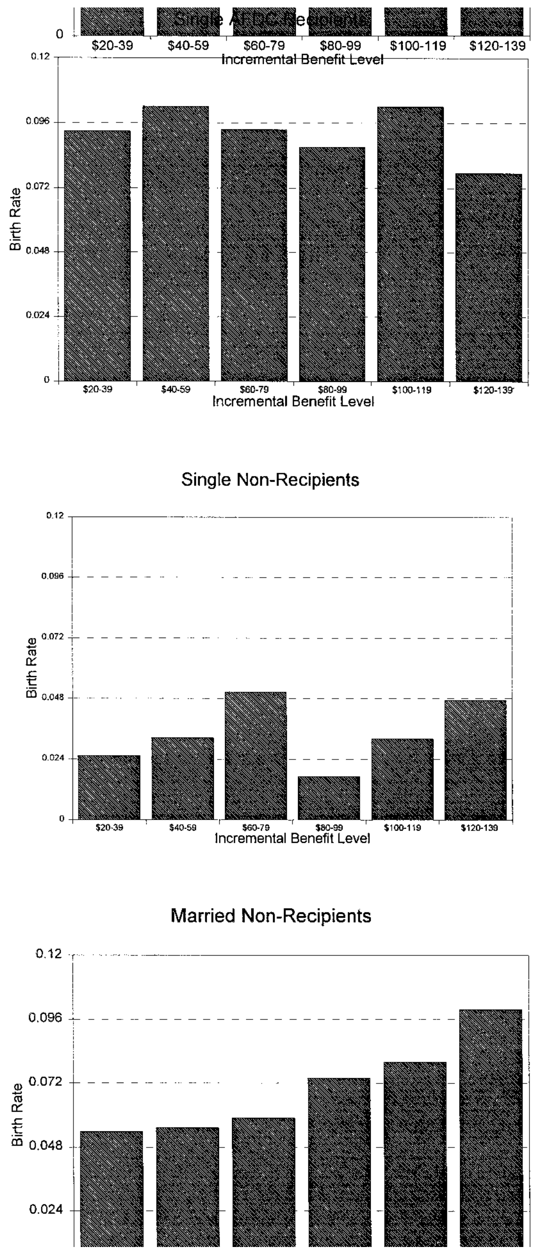

Figure 1. Higher-order birthrates by incremental AFDC benefit levels.

with having this child is larger than the expected utility from not having this child. Although it is impossible to identify all of the monetary and nonmonetary factors that contribute to mothers' expected utility, we can include proxies for a number of important factors. We also do not attempt to estimate a structural model of the decision. Instead, we estimate a reduced form of the solution to the utility maximization problem which is represented by the following logit regression:

$$Prob(B_{it+1} = 1) = \frac{e^{\beta' X_{it} + \gamma I_{it}}}{1 + e^{\beta' X_{it} + \gamma I_{it}}} \tag{1}$$

where B_{it+1} represents the birth decision for individual i between time t and $t + 1$, X_{it} is a vector of individual-level and state-level characteristics, and I_{it} is the potential incremental AFDC benefit level. In all of our logit regressions, the dependent variable is equal to 1 if the woman has a child during a one-year interval and 0 otherwise. We are primarily interested in the sign and magnitude of γ because it represents the effect of the incremental benefit on the probability of a higher-order birth after controlling for other factors.

We estimate equation (1) using a sample of single women who receive AFDC and have at least one child. In Table 3, we report results for the total AFDC sample and separate results for samples of white, black, and Hispanic women.[14] Although not reported, each of the specifications includes an intercept and an indicator variable denoting the year of the observation. The included independent variables are measured in the month prior to the one-year interval in which a birth may occur. Mean values of each independent variable for the full sample are reported in Table A.1 of the Appendix. For the full sample, the average birthrate is 0.0935 and the average incremental AFDC benefit a woman would receive by having another child is $78.

Specification 1 of Table 3 reports the results for the full sample of AFDC recipients. The estimates indicate that being black or Hispanic increases the probability of having a child relative to being white. In addition, the birth probability increases with age for younger mothers and decreases with age for older mothers. Graduating from high school, having a job, the level of other family income, and living with parents all decrease the probability of having a child.[15] The number of previous children and age of the youngest child also affect the birth probability.

In addition to these individual-level characteristics, we include controls for the availability of abortion and contraceptive services at the state level.[16] We include these state policy variables to control for their effect on fertility and their correlation with incremental AFDC benefit levels. If states with generous AFDC benefits also provide more abortion and contraceptive services, then we would understate the effect of AFDC benefits on fertility by excluding these controls. Their inclusion provides us with unbiased estimates of the effect of incremental benefits on fertility. All else equal, we expect that the availability

[14] The standard errors reported in Table 3 are corrected for the possibility of including two observations per individual.

[15] Other family income includes family income not earned by the individual or received as a means-tested cash transfer.

[16] See Lundberg and Plotnick [1995] and Powers [1994] for recent studies which include similar measures.

Table 3. Logit regressions for probability of birth for single women with children who receive AFDC (ages 15–44), 1990 Survey of Income and Program Participation (SIPP).

	Specification			
Sample	1 (Total)	2 (White)	3 (Black)	4 (Hispanic)
Incremental AFDC benefit	0.0083	0.0071	0.0024	−0.0233
	(0.0066)	(0.0173)	(0.0093)	(0.0258)
Black	0.7628	—	—	—
	(0.3251)			
Hispanic	1.2536	—	—	—
	(0.3335)			
Asian, Native American, or other race	0.3368	—	—	—
	(0.5949)			
Age	0.2426	0.3276	0.3675	0.0541
	(0.1912)	(0.4261)	(0.2499)	(0.2743)
Age squared/100	−0.6574	−0.7753	−0.9468	−0.2416
	(0.3848)	(0.7806)	(0.4860)	(0.5068)
High school graduate	−0.4502	−0.8980	−0.1122	−1.3254
	(0.2325)	(0.5523)	(0.3040)	(0.7628)
With a job	−0.2107	−1.2354	−0.1174	1.0623
	(0.3840)	(1.2822)	(0.5088)	(1.1112)
Never married	0.0141	0.8630	−0.7224	−0.0236
	(0.3506)	(0.6166)	(0.4875)	(0.6647)
Live with parents	−0.1680	—	0.1388	−0.1019
	(0.3255)		(0.4493)	(0.8080)
Log other family income	−0.0990	−0.2662	−0.1303	−0.0405
	(0.0412)	(0.1178)	(0.0561)	(0.1154)
Number of children (ages ≤18)	−0.2375	2.9095	−0.3589	−0.6867
	(0.3836)	(1.9020)	(0.5417)	(1.3790)
Number of children (ages ≤18) squared	−0.0295	−0.9746	−0.0089	0.0466
	(0.0717)	(0.5496)	(0.0908)	(0.2332)
Number of children (ages ≤5)	0.0591	0.3407	0.7321	−4.5471
	(0.4361)	(1.2473)	(0.6161)	(1.2556)
Number of children (ages ≤5) squared	0.1011	0.2108	−0.0055	1.2943
	(0.1068)	(0.4312)	(0.1332)	(0.3496)
Age of youngest child	0.0856	0.3197	0.1576	−0.3856
	(0.1080)	(0.2682)	(0.1708)	(0.2914)
Age of youngest child squared	−0.0078	−0.0307	−0.0053	−0.0125
	(0.0077)	(0.0363)	(0.0106)	(0.0287)
Restrictions on teen abortions	0.1724	0.1913	−0.0501	7.0092
	(0.3020)	(0.7783)	(0.4276)	(2.9361)
No medicaid funding for abortions	−0.1604	−0.1328	−0.0168	−0.5571
	(0.3905)	(0.9107)	(0.5854)	(1.4230)
Abortion providers/women (ages 15–44)	−9.2362	−12.3692	−6.4448	45.7623
	(9.2776)	(24.8105)	(12.6955)	(44.3829)
Log expenditures on contraceptive services	0.3131	−0.8510	0.3455	5.5901
	(0.3262)	(0.6417)	(0.5560)	(1.9363)
Unemployment rate	−0.0499	−0.8490	−0.0026	1.7740
	(0.1341)	(0.5328)	(0.1882)	(0.8393)
Female labor force participation rate	0.0372	−0.1453	0.0463	0.3988
	(0.0439)	(0.1685)	(0.0609)	(0.2512)
Log median household income	−1.2906	−3.2537	0.1941	10.4804
	(1.3436)	(4.3477)	(1.7288)	(7.1609)
Male-to-female ratio (ages 15–44)	−0.4090	8.3707	−6.8943	−5.5467
	(4.0200)	(9.6174)	(6.2312)	(16.5577)
Sample birthrate	0.0935	0.0574	0.1058	0.1412
Sample size	1120	383	520	170
Log likelihood	−295.26	−59.59	−150.66	−47.88

Notes: The dependent variable is equal to 1 if the woman has a child during the one-year interval. Standard errors are in parentheses below the coefficient estimates and are corrected for the possibility of including two observations per individual. All equations include an intercept and an indicator variable denoting the year of the observation.

of contraceptive and abortion services in a state reduces the birth probability by decreasing the likelihood that a woman becomes pregnant and carries the baby to term. The variables that serve as proxies for the availability of abortion services in the state are: (a) whether a state imposes parental notification or consent requirements on teen abortions; (b) whether a state disallows Medicaid funding for abortions; and (c) the number of abortion providers per woman (ages 15 to 44) in the state.[17] Our proxy for access to contraceptive services is the log of total expenditures on contraceptive services per woman (ages 15 to 44) in the state.[18]

Consistent with our expectations, we find that women in states with parental consent restrictions on teen abortions have a higher birth probability and that more abortion providers per woman in the state decreases the birth probability. Unexpectedly, the amount spent on contraceptive services per woman in a state has a positive effect on the probability of having a child and the presence of state restrictions on Medicaid funded abortions decreases the birth probability. The coefficients on all of these variables, however, are imprecisely measured.

We also include other state-level variables in our regression: the unemployment rate; female labor force participation rate; log median household income; and male-to-female ratio (ages 15 to 44). These variables are added to the regression to control for observable economic and demographic differences across states and provide mixed results.

Our main goal is to determine the size of the effect of incremental AFDC benefits on fertility. We use the maximum incremental AFDC benefit available to the individual based on her state of residence, year of response, and number of previous children as our measure of incremental benefits. The coefficient estimate for this variable is large and positive, although statistically insignificant at the $\alpha = 0.05$ level (p-value = 0.210). The lack of statistical significance for this coefficient estimate suggests that incremental benefits may not have a positive effect on fertility; however, it is just as likely that the true effect is much larger than that implied by our coefficient estimate. With this in mind, we find that the magnitude of the estimated coefficient implies a large effect of incremental benefits on fertility. The estimates predict that the elimination of incremental benefits (a 100 percent decrease) would reduce the birthrate among AFDC recipients by nearly 60 percent.[19]

Although not reported, we try a number of alternative measures of the incremental AFDC benefit level. Our results do not change when we use the log incremental AFDC benefit, a quadratic specification for the incremental benefit, or the incremental AFDC plus food stamp benefits. We calculate the last measure taking into account the full interaction between the benefit levels for the two programs. We also estimate a logit regression that includes the total AFDC guarantee in addition to the incremental benefit. The coefficient on the incremental benefit is negative and larger in absolute value, but the correlation

[17] These data are reported and described in Blank, George, and London [1996].

[18] Expenditures on contraceptive services by state are reported in Gold and Daley [1991] and Daley and Gold [1993].

[19] The average elasticity of the birth probability with respect to the incremental benefit is 0.583. Because of the nonlinearity of the logit model, the average elasticity is equal to the sample average of $I_{it}\gamma/(1 + EXP[Z_{it}\pi])$, where $Z_{it} = (X_{it}, I_{it})$ and $\pi = (\beta, \gamma)$.

between the guarantee and the incremental benefit is equal to 0.84, potentially indicating a problem with collinearity between these two variables.

In Specifications 2, 3, and 4, we report estimates by race. Specification 2 reports estimates using a sample of white AFDC recipients.[20] The coefficient on the incremental AFDC benefit is slightly smaller than the one for the full sample with a larger standard error. The coefficient for blacks (Specification 3) is positive, but much smaller than the coefficient for the full sample. The coefficient in the Hispanic regression (Specification 4) is negative and imprecisely measured. The large standard error on this coefficient is likely due to both the regional concentration of Hispanics and the small sample size for this group. These results provide some evidence of a positive relationship between fertility and incremental benefits for whites and blacks, but not for Hispanics.

The results presented in Table 3 suggest that incremental benefits may have a large positive effect on fertility among AFDC recipients. The evidence of this effect, however, is provided by several imprecisely measured coefficients. We now turn to using a different approach for estimating the effect of incremental benefits on fertility.

Using Comparison Groups to Estimate the Effect of Incremental Benefits

In the previous analysis, we cannot rule out the possibility that the positive coefficient estimate on the incremental AFDC benefit is due to the omission of important unmeasurable state-level or family size-level characteristics.[21] One method of controlling for these unmeasurable factors is to include a comparison group that would theoretically not be affected by state and family size variation in incremental AFDC benefits, but would be affected similarly by these unmeasurable factors.[22] We can then compare our estimate of the effect of incremental benefits on fertility for the AFDC sample to our estimate of the effect for the comparison group sample. The difference between the two estimated effects provides an estimate of the true effect of incremental benefits on the fertility of AFDC recipients.

To incorporate the comparison group into our regressions, we respecify equation (1) as:

$$Prob(B_{it+1} = 1) = \frac{e^{\beta' X_{it} + \delta A_{it} + \gamma_I I_{it} + \gamma_{AI} AI_{it}}}{1 + e^{\beta' X_{it} + \delta A_{it} + \gamma_I I_{it} + \gamma_{AI} AI_{it}}} \qquad (2)$$

[20] We cannot estimate a coefficient on the variable for living with parents because none of the women in this sample who live with their parents have a birth.

[21] It is likely that there exists a large number of unmeasurable factors that influence the birth decision. These factors may include cultural and religious background, attitudes toward families, perceived opportunity costs, and quality of health care and family planning services.

[22] An alternative approach is to include individual or state fixed effects in equation (2). These fixed effects remove the bias created by unmeasurable variables that differ across individuals or states, but are constant over time. A common problem with fixed-effect models, however, is that they do not perform well if there exists only a small amount of time-series variation in the independent variable of interest. In our sample, the time-series variation in incremental benefit levels is very small. For 92.9 percent of our sample, the change between 1990 and 1991 in incremental benefits is less than $6. Most of these small changes are negative and are simply due to AFDC benefits not being indexed to the consumer price index (CPI) in many states. Furthermore, using a longer time frame may not improve the performance of the model. For example, in the decade from 1981 to 1991 the average annual change in incremental benefits was $−1.40. During this period, only five states experienced an average change that was greater than $1.00 or less than $−3.00.

where A_{it} is an indicator variable denoting AFDC receipt and AI_{it} is the interaction between this indicator variable and the incremental AFDC benefit. We estimate the model using our previous sample of AFDC recipients and the comparison group sample. In this model, γ_I represents the correlation between incremental benefits and birth probabilities for the comparison group, and γ_{AI} represents the independent effect of the incremental benefit on AFDC recipients. Theoretically, in the absence of important omitted variables we should obtain an estimate of γ_I equal or approximately equal to zero and a positive estimate of γ_{AI}. Empirically, with a valid comparison group, we can assume that a nonzero estimate of γ_I represents the spurious correlation between incremental benefits and fertility and that a nonzero estimate of γ_{AI} represents the true effect of incremental benefits on the fertility of AFDC recipients.

The choice of a valid comparison group is important, and thus we estimate equation (2) using several different comparison groups. In all comparison groups, we exclude women who do not have any previous children because the determinants of first births are likely to differ greatly from the determinants of higher-order births. Our first comparison group includes all single women (ages 15 to 44) who have at least one child, but do not receive AFDC. These women are not directly affected by incremental AFDC benefits because they did not choose to receive AFDC even though they were potentially eligible for benefits. We do not rule out the possibility that a subset of these women may respond to differences in incremental AFDC benefits indirectly through responding to differences in the total guarantee (which is highly correlated with the incremental benefit). These women who do not currently receive AFDC may be affected by variation in incremental benefits because they plan to receive AFDC after having another child. However, the total effect is likely to be small because these women are only a subset of all nonrecipients and are responding indirectly to the incremental benefit.[23]

Our second comparison group removes the potential problem discussed previously. Duncan and Hoffman [1990] argue that births not associated with subsequent AFDC receipt could not have been influenced by AFDC benefits.[24] Therefore, we create the second comparison group by excluding all single mothers who receive AFDC at the end of the birth interval (time $t + 1$) from the first comparison group sample.

Our third comparison group includes only married women (ages 15 to 44) who have at least one child and who do not receive AFDC. We do not expect this group of women to directly respond to differences in incremental AFDC benefits. Finally, to create the fourth comparison group we exclude from the sample of married mothers those who receive AFDC at the end of the birth interval.

In Table 4, we estimate equation (2) using our AFDC recipient sample and each of the comparison groups. We only report estimates of δ, γ_I and γ_{AI}, but we include the same individual- and state-level variables as those included in

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It appears that the variation across states in 1990 is substantially larger than the variation within states over the 1980s.

[23] Approximately 6 percent of our sample of single mothers who are nonrecipients in time t receive AFDC in time $t + 1$.

[24] They note the possibility that some women who do not subsequently receive AFDC may view AFDC as insurance, thus allowing benefit levels to affect fertility for these women. However, they do not find evidence suggesting that this is the case. The insurance effect is likely to be quite small implying that our more restrictive comparison group should remain valid.

Table 4. Logit regressions for probability of birth for women with children (ages 15–44), 1990 Survey of Income and Program Participation (SIPP).

	Specification			
	1	2	3	4
AFDC receipt	0.2893	0.3958	0.0178	0.0183
	(0.3994)	(0.4379)	(0.3897)	(0.3923)
Incremental AFDC benefit	0.0075	0.0068	0.0075	0.0081
	(0.0047)	(0.0051)	(0.0021)	(0.0021)
AFDC receipt * incremental benefit	−0.0006	0.0001	−0.0067	−0.0068
	(0.0047)	(0.0051)	(0.0038)	(0.0038)
Sample birthrate	0.0511	0.0465	0.0694	0.0689
Sample size	3854	3699	10778	10673
Log likelihood	−639.32	−561.04	−2315.84	−2279.37

Notes: See Table A.2 in the Appendix for more details and coefficient estimates for all included variables. The comparison groups are single mothers who do not receive AFDC in time t (Specification 1); single mothers who do not receive AFDC in time t or $t + 1$ (Specification 2); married mothers who do not receive AFDC in time t (Specification 3); and married mothers who do not receive AFDC in time t or $t + 1$ (Specification 4).

the logit regressions of Table 3 (see Table A.1 for mean values and Table A.2 for estimates).[25] We include our first comparison group, single mothers who do not receive AFDC in time t, in Specification 1. The coefficient on the incremental benefit (γ_I) is 0.0075 and the coefficient on the incremental benefit interacted with AFDC participation (γ_{AI}) is −0.0006. These estimates suggest that nonrecipients respond more than AFDC recipients to differences in incremental AFDC benefits. Because there is no theoretical justification for a large effect among nonrecipients, at least part of the positive estimate of γ_I must be due to a positive spurious correlation between incremental benefits and fertility among this group. In addition, the negative estimate of γ_{AI} indicates that the positive correlation among AFDC recipients is smaller. These two results taken together suggest that the true effect of the incremental benefit on the fertility of AFDC recipients is much smaller than that implied by our original coefficient estimate (γ) using only the AFDC sample.

In Specification 2, we use our second and more restrictive comparison group which excludes those single women who receive AFDC in time $t + 1$. This restriction reduces the sample size by 4 percent. Our estimate of γ_I has the same sign and magnitude as the estimate in Specification 1, while our estimate of γ_{AI} is now positive. Although our estimate of γ_{AI} is positive, it differs only slightly from our previous estimate and implies a negligible effect. It appears as though the exclusion of the subgroup of single women who are potentially most responsive to differences in incremental AFDC benefits does not change our results.

The estimated coefficients from the logit regression that uses married women with children as the comparison group are reported in Specification 3. We find a large positive correlation between incremental benefits and higher-order birth probabilities among married women. Again, there is no theoretical justi-

[25] We do not report mean values for the two more restrictive comparison group samples because of their similarity to the original single and married mother samples.

fication for married women to respond greatly to differences in incremental benefits suggesting that this coefficient estimate mainly represents a spurious correlation. Furthermore, our negative estimate of γ_{AI} implies that AFDC recipients respond even less to differences in incremental benefits. Taken together, these two findings severely undermine the hypothesis that incremental benefit levels have a large positive effect on fertility among AFDC recipients. We also estimate equation (2) using a more restrictive comparison group which excludes married women who receive AFDC in time $t + 1$ (reported in Specification 4). This reduces the comparison group sample size by only 1 percent and produces very similar results.

We now examine the sensitivity of our results to alternative specifications of the birth probability model. The samples used in Specifications 1 and 3 are used for these robustness checks. First, we estimate our equations using the log incremental AFDC benefit, a quadratic specification for the incremental benefit, and the incremental AFDC plus food stamp benefits as independent variables. We find that our results are not sensitive to the chosen measure of incremental benefits. Second, we estimate specifications that include additional or different subsets of state-level controls. Our results are not sensitive to these alternative specifications. Third, we include regional indicator variables. Our estimates of γ_I become smaller in absolute value, but our estimates of γ_{AI} do not noticeably change.[26] Therefore, the reduction in sample variation across states in incremental benefits from the inclusion of the regional indicators affects our estimates of γ_I, but not our estimates of γ_{AI}. Fourth, we estimate our equations excluding those women who had a child in the year prior to the beginning of the one-year birth interval. Our estimates of γ_{AI} are very similar using this restriction. Fifth, we estimate our equations using a two-year time frame instead of the one-year time frame to check for the occurrence of a birth. The estimate of γ_{AI} using married mothers as the comparison group is similar to the one reported in Specification 3. The estimate of γ_{AI} using single mothers as the comparison group is now positive, although small and statistically insignificant. Finally, we find either negative or small positive estimates of γ_{AI} when we impose separate restrictions of including only prime childbearing-age women (ages 20 to 30) or including only women with one child.

The results presented in this section demonstrate that the correlation between incremental benefits and higher-order birth probabilities among AFDC recipients is smaller than the correlation among each of our comparison groups. This finding suggests that a large part of the positive relationship between incremental benefits and fertility among AFDC recipients that we find in the previous section is spurious. None of the results generated from our comparison group analyses provide evidence of a strong positive effect of incremental benefits on higher-order births among AFDC recipients. Furthermore, our sensitivity analysis suggests that these results are quite robust to alternative specifications of the birth probability model.

The Effect of Incremental Benefits among Subgroups of the AFDC Population

To determine if subgroups of the AFDC population respond differently to incremental benefits, we estimate equation (2) separately by race and marital

[26] The coefficient estimates for the single mother and married mother comparison groups are -0.0005 and -0.0064, respectively.

Table 5. Logit regressions for probability of birth for women with children (ages 15–44), 1990 Survey of Income and Program Participation (SIPP).

	Specification				
Sample	1 (White)	2 (Black)	3 (Hispanic)	4 (Never married)	5 (Divorced or separated)
AFDC receipt	0.6823	0.6750	0.2126	0.5514	0.1765
	(0.8900)	(0.6014)	(1.0553)	(0.4758)	(0.7911)
Incremental AFDC benefit	0.0086	0.0068	0.0048	0.0120	−0.0027
	(0.0075)	(0.0090)	(0.0119)	(0.0064)	(0.0074)
AFDC receipt * incremental benefit	−0.0091	−0.0029	0.0056	−0.0030	0.0030
	(0.0105)	(0.0075)	(0.0105)	(0.0056)	(0.0089)
Sample birthrate	0.0308	0.0633	0.0941	0.0894	0.0259
Sample size	1983	1311	457	1465	2242
Log likelihood	−207.55	−260.77	−112.49	−390.53	−217.49

Notes: See Table A.3 in the Appendix for more details and coefficient estimates for all included variables. For all specifications, the comparison groups are single mothers who do not receive AFDC in time t and who are in the appropriate subgroup.

status. First, we estimate logit regressions using three racial groups: whites, blacks, and Hispanics. In Specifications 1, 2, and 3 of Table 5, we report the results from these regressions using single mothers who do not receive AFDC as the comparison group. Our estimate of γ_{AI} is negative for whites and blacks, but is positive for Hispanics. The coefficients for all of these groups, however, are imprecisely measured. Although not reported, we estimate logit regressions using our other comparison groups. Using the more restrictive comparison group of single mothers, the estimates of γ_{AI} have the same sign, but are larger in absolute value. The results for the comparison group of married mothers are also qualitatively similar.[27]

Second, we estimate separate logit regressions for the two main marital status categories for single mothers: never-married and divorced or separated. In Specification 3, we include only never-married women with children and use the same restrictions for our comparison group as those used for our first comparison group. Similar to the results using the full sample, our estimate of γ_I is positive and our estimate of γ_{AI} is negative. We find similar results using the more restrictive never-married comparison group which excludes women receiving AFDC in time $t + 1$. These results support the finding of Robins and Fronstin [1996] that there is no strong positive effect of the incremental benefit on higher-order births among all never-married women.[28]

In Specification 5, we include divorced and separated women. These two groups of women comprise 38.1 percent of the AFDC population. In this regression, we obtain a negative estimate of γ_I and a positive estimate of γ_{AI}. The

[27] Estimates from the more restrictive married mother comparison group are similar to these.

[28] They only include never-married women in their analysis because the CPS does not contain information on marital histories. By including only this group, they are assured that all births occurring in the past were out-of-wedlock births and, thus potentially affected by AFDC benefits. The CPS also does not allow them to distinguish between AFDC recipients and non-AFDC recipients. Therefore, they cannot estimate the independent effect of incremental benefits on fertility among AFDC recipients.

estimate of 0.0030 for the interaction coefficient, however, is much smaller than our original coefficient estimate (γ) using only the AFDC sample. The use of a more restrictive comparison group which excludes those divorced and separated women who receive AFDC in time $t + 1$ provides a larger estimate of the interaction coefficient (equal to 0.0078). In sum, there is some evidence of a positive effect of incremental benefits on the fertility of previously married AFDC recipients, but this evidence is based on imprecisely measured coefficients.

The results for whites, blacks, and never-married women are similar to the results using the full sample. We find that for each of these three groups of women, those who receive AFDC respond substantially less to differences in incremental AFDC benefits than those in the comparison groups. The results for Hispanics and divorced or separated women are less clear. Overall, the results in this section must be interpreted with some caution because many of the relevant coefficients are imprecisely measured for these subgroups of our sample.

CONCLUSION

The existing variation in higher-order birth probabilities and incremental AFDC benefit levels across states and family sizes provides an obvious source of data to test the potential impact of family cap policies. However, some caution is warranted as the familiar adage that correlation does not necessarily imply causation may be especially appropriate in this case. It may be impossible with existing data sources to control for all of the factors that affect fertility differently across states and family sizes and are correlated with incremental AFDC benefits. Our use of comparison groups that are unlikely to respond to differences in incremental benefits offers a potential remedy for this problem.

Using only AFDC recipients to estimate our logit regressions of the birth decision, we find a positive coefficient on the incremental AFDC benefit. This coefficient, although imprecisely measured, it shown to imply a large effect of these benefits on fertility in the direction predicted by economic theory. We argue, however, that the actual effect is much smaller and is probably negligible. We find a large positive coefficient on incremental benefits for each of our comparison groups (which by design should not theoretically respond to this variable) and a negative coefficient on the variable that measures the interaction between AFDC receipt and incremental benefits. These results imply that a substantial portion of the estimated positive relationship between incremental benefits and fertility among AFDC recipients is spurious. We also argue that the alternative explanation for these findings, that our comparison groups respond more to differences in incremental benefits than AFDC recipients, is implausible.

Overall, our estimates suggest that the potential effect of family cap policies on the fertility of AFDC recipients is likely to be small at best. There are many possible reasons for this finding. First, incremental benefit levels may appear very small relative to the costs of raising an additional child, thus weakening AFDC participants' sensitivity to changes in incremental benefits.[29] Second, a large proportion of the AFDC population receives AFDC for less than two years.[30] Because these women have to pay the costs of raising their children

[29] Updating older estimates from Espenshade [1984], Haveman and Wolfe [1995] estimate that the average annual direct cost per child in the United States was $7579 in 1992.

[30] Using data from the Panel Study of Income Dynamics (PSID), Bane and Ellwood [1994] estimate that 36.4 percent of women beginning a first spell of AFDC have an expected total time on AFDC of two years or less.

for a much longer period of time, the effect of changes in incremental benefits on fertility is likely to be inconsequential to them. Third, economic theory proposes that higher incremental benefits increase the probability of planned births; however, many births are unplanned among certain subgroups of the AFDC population.[31] Finally, past ethnographic research shows that nonpecuniary factors are very important in the birth decision for at least a portion of the AFDC population.[32]

The implementation of family cap policies in most states in the United States appears to be inevitable in coming decades. Perhaps the most vehement argument used to promote these policies is that they will remove the economic incentive for single women receiving AFDC to have additional out-of-wedlock children. Based our findings, however, policymakers may need to find another policy lever if the ultimate goal is to reduce fertility among AFDC recipients.

APPENDIX

Table A.1. Means of variables used in logit regressions for women with children (ages 15–44), 1990 Survey of Income and Program Participation (SIPP).

Variable	Single AFDC recipients	Single non-recipients	Married non-recipients
Birth probability	0.094	0.034	0.067
1991 observation	0.483	0.482	0.484
Incremental AFDC benefit	77.664	73.953	73.817
Black	0.464	0.289	0.085
Hispanic	0.152	0.105	0.011
Asian, Native American, or other race	0.042	0.020	0.044
Age	29.030	32.893	33.737
High school graduate	0.529	0.816	0.860
With a job	0.118	0.818	0.657
Never married	0.588	0.295	0.000
Live with parents	0.169	0.153	0.010
Log other family income	2.451	4.440	7.521
Number of children (ages ≤18)	2.136	1.649	1.999
Number of children (ages ≤5)	1.009	0.485	0.756
Age of youngest child	4.454	7.658	5.987
Restrictions on teen abortions	0.192	0.214	0.213
No Medicaid funding for abortions	0.583	0.647	0.641
Abortion providers/women (ages 15–44)	0.042	0.040	0.040
Log expenditures on contraceptive services	2.136	2.137	2.121
Unemployment rate	6.309	6.135	6.106
Female labor force participation rate	56.928	57.356	57.549
Log median household income	10.288	10.282	10.286
Male-to-female ratio (ages 15–44)	1.004	1.005	1.008
Log AFDC incremental benefit	4.239	4.194	4.197
Incremental AFDC benefit + food stamps	126.779	126.592	124.375
Sample size	1120	2734	10,778

Note: All mean values are unweighted.

[31] See Zelnik and Kantner [1980] for evidence.

[32] For example, Anderson [1989] states that “among many young poor ghetto women, babies have become a sought-after symbol of status, of passage to adulthood, or being a ‘grown’ woman” (p. 69).

Table A.2. Logit regressions for probability of birth for women with children (ages 15–44).

	Specification			
	1	2	3	4
AFDC receipt	0.2893	0.3958	0.0178	0.0183
	(0.3994)	(0.4379)	(0.3897)	(0.3923)
Incremental AFDC benefit	0.0075	0.0068	0.0075	0.0081
	(0.0047)	(0.0051)	(0.0021)	(0.0021)
AFDC receipt * incremental benefit	−0.0006	0.0001	−0.0067	−0.0068
	(0.0047)	(0.0051)	(0.0038)	(0.0038)
Black	0.5482	0.6145	0.0979	0.1175
	(0.1985)	(0.2167)	(0.1376)	(0.1394)
Hispanic	1.1286	1.1843	0.0934	0.1015
	(0.2340)	(0.2561)	(0.1340)	(0.1351)
Asian, Native American, or other race	0.9465	1.0156	0.0675	0.0908
	(0.4204)	(0.4357)	(0.1903)	(0.1922)
Age	0.2053	0.1977	0.2941	0.2918
	(0.1096)	(0.1163)	(0.0676)	(0.0690)
Age squared	−0.5683	−0.5509	−0.6377	−0.6349
	(0.2003)	(0.2150)	(0.1152)	(0.1174)
High school graduate	−0.2742	−0.3824	−0.1960	−0.2119
	(0.1662)	(0.1785)	(0.1103)	(0.1112)
With a job	−0.4030	−0.5432	−0.1488	−0.1590
	(0.2157)	(0.2408)	(0.0857)	(0.0868)
Never married	0.0294	−0.1245	0.2903	0.2836
	(0.2091)	(0.2229)	(0.2620)	(0.2627)
Live with parents	−0.5384	−0.4952	−0.4568	−0.4341
	(0.2318)	(0.2464)	(0.2435)	(0.2461)
Log other family income	−0.0269	−0.0152	−0.0375	−0.0377
	(0.0290)	(0.0307)	(0.0220)	(0.0226)
Number of children (ages ≤18)	−0.1981	−0.2137	−1.3510	−1.3471
	(0.3058)	(0.3256)	(0.1221)	(0.1226)
Number of children (ages ≤18) squared	−0.0385	−0.0433	0.1627	0.1625
	(0.0659)	(0.0708)	(0.0180)	(0.0180)
Number of children (ages ≤5)	−0.4412	−0.4061	0.6578	0.6684
	(0.3352)	(0.3418)	(0.2001)	(0.2046)
Number of children (ages ≤5) squared	0.1922	0.2045	−0.0922	−0.0952
	(0.0868)	(0.0914)	(0.0605)	(0.0626)
Age of youngest child	−0.0079	0.0427	0.1445	0.1454
	(0.0819)	(0.0841)	(0.0408)	(0.0411)
Age of youngest child squared	−0.0049	−0.0085	−0.0148	−0.0147
	(0.0056)	(0.0054)	(0.0034)	(0.0034)
Restrictions on teen abortions	0.1881	0.0547	0.0398	0.0485
	(0.2120)	(0.2353)	(0.1159)	(0.1177)
No Medicaid funding for abortions	0.1626	0.1489	0.0641	0.0687
	(0.2755)	(0.2988)	(0.1336)	(0.1340)
Abortion providers/women (ages 15–44)	−3.0698	−1.9041	−0.2156	−0.5246
	(5.9551)	(6.2104)	(2.6823)	(2.6734)
Log expenditures on contraceptive services	0.1511	0.0510	0.0568	0.0721
	(0.2285)	(0.2438)	(0.1204)	(0.1210)
Unemployment rate	−0.1389	−0.1505	0.0115	0.0028
	(0.1017)	(0.1136)	(0.0488)	(0.0492)
Female labor force participation rate	0.0324	0.0480	0.0176	0.0192
	(0.0289)	(0.0316)	(0.0155)	(0.0156)

Table A.2. (*Continued*)

	Specification			
	1	2	3	4
Log median family income	−1.0914	−1.2633	−0.2269	−0.2443
	(1.0341)	(1.1020)	(0.4904)	(0.4939)
Male-to-female ratio (ages 15–44)	−0.3089	−1.5132	1.1572	1.2673
	(3.0382)	(3.1820)	(1.5966)	(1.6139)
Sample birthrate	0.0511	0.0465	0.0694	0.0689
Sample size	3854	3699	10778	10673
Log likelihood	−639.32	−561.04	−2315.84	−2279.37

Notes: The dependent variable is equal to 1 if the woman has a child during the one-year interval. Standard errors are in parentheses below the coefficient estimates and are corrected for the possibility of including two observations per individual. All equations include an intercept and an indicator variable denoting the year of the observation. The comparison groups are single mothers who do not receive AFDC in time t (Specification 1); single mothers who do not receive AFDC in time t or $t + 1$ (Specification 2); married mothers who do not receive AFDC in time t (Specification 3); and married mothers who do not receive AFDC in time t or $t + 1$ (Specification 4).

Table A.3. Logit regressions for probability of birth for women with children (ages 15–44).

	Specification				
Sample	1 (White)	2 (Black)	3 (Hispanic)	4 (Never married)	5 Divorced or separated)
AFDC receipt	0.6823	0.6750	0.2126	0.5514	0.1765
	(0.8900)	(0.6014)	(1.0553)	(0.4758)	(0.7911)
Incremental AFDC benefit	0.0086	0.0068	0.0048	0.0120	−0.0027
	(0.0075)	(0.0090)	(0.0119)	(0.0064)	(0.0074)
AFDC receipt * incremental benefit	−0.0091	−0.0029	0.0056	−0.0030	0.0030
	(0.0105)	(0.0075)	(0.0105)	(0.0056)	(0.0089)
Black	—	—	—	0.2536	1.2338
				(0.2337)	(0.3560)
Hispanic	—	—	—	1.0143	1.1540
				(0.3070)	(0.3908)
Asian, Native American, or other race	—	—	—	0.4605	1.5838
				(0.5197)	(0.6707)
Age	0.2401	0.2784	0.1728	0.1312	0.1831
	(0.2135)	(0.1801)	(0.2024)	(0.1757)	(0.2389)
Age squared	−0.5526	−0.7350	−0.4596	−0.4391	−0.5340
	(0.3762)	(0.3301)	(0.3505)	(0.3572)	(0.3781)
High school graduate	−0.8679	−0.0304	−0.3374	−0.1076	−0.6131
	(0.2887)	(0.2412)	(0.4146)	(0.1977)	(0.3352)
With a job	−0.1439	−0.5632	−0.1824	−0.5192	−0.2000
	(0.3690)	(0.3680)	(0.5932)	(0.2571)	(0.4006)
Never married	0.4651	−0.3818	0.1168	—	—
	(0.3709)	(0.3310)	(0.4088)		
Live with parents	−2.0467	−0.0247	−0.1997	−0.2002	−1.2507
	(0.6297)	(0.3264)	(0.4919)	(0.2712)	(0.5434)

Table A.3. (*Continued*)

Sample	Specification 1 (White)	2 (Black)	3 (Hispanic)	4 (Never married)	5 Divorced or separated)
Log unearned family income	0.0072	−0.0849	0.0062	−0.0676	0.0379
	(0.0518)	(0.0440)	(0.0665)	(0.0363)	(0.0548)
Number of children (ages ≤18)	1.0232	−0.1247	−1.0601	−0.2143	−0.3387
	(1.0297)	(0.4391)	(0.7634)	(0.3734)	(0.5686)
Number of children (ages ≤18) squared	−0.4284	−0.0435	0.1252	−0.0207	−0.0360
	(0.3120)	(0.0847)	(0.1270)	(0.0837)	(0.1105)
Number of children (ages ≤5)	−0.6021	−0.1462	−1.4387	−0.4405	−0.0944
	(0.7158)	(0.5239)	(0.7694)	(0.4189)	(0.6648)
Number of children (ages ≤5) squared	0.3019	0.1338	0.5153	0.1796	0.1069
	(0.2667)	(0.1205)	(0.1941)	(0.1119)	(0.1927)
Age of youngest child	−0.0818	0.0779	−0.1790	0.0162	0.0397
	(0.1614)	(0.1248)	(0.2132)	(0.0954)	(0.1692)
Age of youngest child squared	−0.0095	−0.0073	0.0070	−0.0033	−0.0080
	(0.0128)	(0.0081)	(0.0113)	(0.0062)	(0.0113)
Restrictions on teen abortions	0.9139	−0.3132	1.2491	0.2526	0.2883
	(0.3926)	(0.3447)	(1.0891)	(0.2553)	(0.3923)
No Medicaid funding for abortions	0.0088	−0.3905	2.2672	0.0697	0.1312
	(0.5272)	(0.4248)	(1.0672)	(0.3821)	(0.4596)
Abortion providers/ women (ages 15–44)	1.0830	−18.3189	34.4952	−6.4948	0.6648
	(14.7372)	(11.0483)	(17.8799)	(9.6804)	(8.9506)
Log expenditures on contraceptive services	−0.2446	0.2131	2.2389	0.3838	−0.3296
	(0.3605)	(0.4024)	(1.0399)	(0.2990)	(0.4151)
Unemployment rate	−0.4926	−0.0709	0.1774	−0.2029	−0.1145
	(0.2114)	(0.1537)	(0.3070)	(0.1144)	(0.2541)
Female labor force participation rate	−0.0304	0.0025	0.1305	0.0268	0.0104
	(0.0588)	(0.0457)	(0.1066)	(0.0329)	(0.0695)
Log median family income	−0.5369	−0.2191	4.0172	−1.8782	0.1140
	(2.1588)	(1.4909)	(4.6055)	(1.2917)	(1.9518)
Male-to-female ratio (ages 15–44)	6.3254	−1.1730	−4.1222	−0.3728	4.5302
	(6.3888)	(5.8067)	(7.2523)	(3.6651)	(6.2770)
Sample birthrate	0.0308	0.0633	0.0941	0.0894	0.0259
Sample size	1983	1311	457	1465	2242
Log likelihood	−207.55	−260.77	−112.49	−390.53	−217.49

Notes: The dependent variable is equal to 1 if the woman has a child during the one-year interval. Standard errors are in parentheses below the coefficient estimates and are corrected for the possibility of including two observations per individual. All equations include an intercept and an indicator variable denoting the year of the observation. For all specfications, the comparison groups are single mothers who do not receive AFDC in time *t* and who are in the appropriate subgroup.

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