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Welfare Benefits and Family-Size Decisions of Never-Married Women

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Abstract

Since the 1970s, the out-of-wedlock birthrate has been increasing rapidly in the United States and has prompted several states to propose (and in some cases, enact) legislation to deny access to higher AFDC benefits for families in which the mother gives birth while receiving AFDC. The authors investigate whether AFDC benefit levels are systematically related to the family-size decisions of never-married women. Using a Poisson regression model, applied to Current Population Survey data for the years 1980–1988, they find that the basic benefit level positively influences family size for white and Hispanic women, but not for black women. Incremental benefits for larger families, however, do not affect family-size decisions, suggesting that reducing (or eliminating) this differential will not necessarily reduce the number of illegitimate births. The basic benefit level positively affects the family-size decisions of high school dropouts, but not of high school graduates. This suggests that to discourage nonmarital births, policymakers should consider altering the AFDC benefit structure in such a way as to encourage single mothers to complete high school. However, being a high school dropout might be a proxy for some other underlying characteristic of the woman, and inducing women to complete high school who otherwise would not might have no effect whatsoever on nonmarital births.

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I. INTRODUCTION

Over the entire post–World War II period, the birthrate among unmarried women has increased tremendously. In 1950, there were approximately fourteen live births per thousand unmarried women. By 1989, this figure had tripled, to forty-two live births per thousand (see Figure 1). Except for a slight decline between 1970 and 1975, the upward trend in the rate of illegitimate births has continued unabated, and has even accelerated since 1975.

The rise in out-of-wedlock births is the result of two underlying phenomena. First, the birthrate among unmarried white women almost quintupled during this period, from six births per thousand in 1950 to twenty-nine births per thousand in 1989. Second, black women, who have a much higher birthrate than white women, have been constituting an ever increasing proportion of the total population of unmarried women. In 1970, black women constituted 15 percent of the population of unmarried women; by 1989, they constituted 20 percent. Interestingly, the birthrate among unmarried black women fell considerably between 1970 and 1984 (from ninety-six live births to seventy-seven), but since 1984, it has been increasing like the white rate and has almost returned to its 1970 level.[1]

Many casual observers have asserted that the Aid to Families with Dependent Children (AFDC) program encourages out-of-wedlock childbearing because AFDC guarantee levels increase with family size. As indicated in Table 1, the average (unweighted) monthly guarantee level in the United States in 1991 for a four-person family (parent and three children) was roughly 45 percent higher than the average guarantee level for a two-person family ($463 versus $320).[2] The average differential has fallen by more than 40 percent in real terms since the late 1960s, mirroring the decline in average guarantee levels.

Figure 1 here

TABLE 1

Average Monthly AFDC Guarantee Levels, by Number of Children: 1968–1991

Year	Number of Children 1	2	3	Average Differential between Benefit for Family with One Child and Benefit for Family with Three Children
1968	$516	$641	$767	$126
1969	521	641	761	120
1970	515	638	762	124
1971	514	635	756	121
1972	510	631	749	120
1973	501	616	731	115
1974	498	608	718	110
1975	480	595	703	112
1976	472	587	694	111
1977	462	572	676	107
1978	452	560	662	105
1979	425	521	618	97
1980	374	459	544	85
1981	361	449	530	85
1982	345	427	504	80
1983	349	430	507	79
1984	337	414	490	77
1985	341	420	497	78
1986	355	436	514	80
1987	348	428	505	79
1988	346	424	499	77
1989	336	413	487	76
1990	326	401	472	73
1991	320	393	463	72

Sources: Background Material and Data on Programs within the Jurisdiction of the Committee on Ways and Means, Washington, D.C., U.S.G.P.O., various years, and private communications from Evelyn Mills, Department of Health and Human Services, Family Support Administration.

Note: Unweighted average guarantees across states, 1991 dollars, deflated using the Consumer Price Index.

Guarantee differentials by family size vary considerably across the states, ranging from a 22 percent average differential in Wyoming to a 70 percent average differential in Louisiana (see Table 2). The change in the differential over time also has varied considerably across states, ranging from a $200 real decline in Florida to a $7 real increase in Hawaii.

Although it has been argued that the scaling of AFDC benefits with family size fosters childbearing (Murray, 1984, 1993), the logic behind the family-size differentials is based on the notion that larger families are more costly to support than smaller families.[3] Nevertheless, the fact that there is a family-size differential, together with the fact that the out-of-wedlock birthrate has been increasing rapidly since the mid-1970s, has prompted several states to propose (and in some cases, enact) legislation to deny access to higher AFDC benefits for families in which the mother gives birth while receiving AFDC.[4] The purpose of such legislation is apparently to discourage out-of-wedlock childbearing among AFDC recipients and to correspondingly reduce the costs of the AFDC program. This rationale is somewhat curious given the real decline in benefit differentials over time. It may be that the desire is to reduce the rate of increase of nonmarital births, which are due primarily to noneconomic factors. It is also possible that such legislation is meant to be simply symbolic, sending a signal to the public that out-of-wedlock childbearing is socially undesirable.

Although it has been asserted that denying benefit increases for additional children will discourage out-of-wedlock childbearing (or at least slow it down), the empirical validity and quantitative importance of this hypothesis have not been adequately tested. Many studies have examined the relationship between benefit <u>levels</u> and family size, but few studies have examined the relationship between benefit <u>differentials</u> and family size. The purpose of this paper is to investigate whether both AFDC benefit levels and differentials are systematically related to the family-size decisions of never-married women.

TABLE 2

AFDC Monthly Guarantee Differentials by Family Size in 1991 and Change from 1968 (1991 Dollars)

State	Guarantee Level 1 Child 1991	Guarantee Level 3 Children 1991	Average Differential Per Child 1991	Change in the Average Differential 1968–1991
Alabama	$93	$155	$31	$-28
Alaska	792	990	99	-58
Arizona	233	353	60	-46
Arkansas	162	247	43	3
California	560	824	132	-11
Colorado	280	432	76	-47
Connecticut	549	792	122	-131
Delaware	270	407	69	-53
District of Columbia	336	522	93	-32
Florida	225	346	61	-202
Georgia	235	330	48	-66
Hawaii	504	760	128	7
Idaho	254	357	52	-58
Illinois	268	414	73	-84
Indiana	229	346	59	-39
Iowa	361	495	67	-113
Kansas	338	470	66	-96
Kentucky	196	285	45	-104
Louisiana	138	234	48	-22
Maine	337	569	116	4
Maryland	317	489	86	-29
Massachusetts	446	628	91	-97
Michigan	429	635	103	-40
Minnesota	437	621	92	-84
Mississippi	96	144	24	-11
Missouri	234	342	54	-40
Montana	295	445	75	-109
Nebraska	293	435	71	-46
Nevada	270	390	60	-77
New Hampshire	451	575	62	-40
New Jersey	322	488	83	-111
New Mexico	247	373	63	-56
New York	468	687	110	-102

(table continues)

TABLE 2, continued

State	Guarantee Level 1 Child 1991	Guarantee Level 3 Children 1991	Average Differential Per Child 1991	Change in the Average Differential 1968–1991
North Carolina	236	297	31	-22
North Dakota	326	491	83	-72
Ohio	274	413	70	-9
Oklahoma	264	423	80	-28
Oregon	380	541	81	-82
Pennsylvania	330	514	92	-49
Rhode Island	449	632	92	-49
South Carolina	167	252	43	-40
South Dakota	340	429	45	-89
Tennessee	150	238	44	-15
Texas	158	221	32	-51
Utah	323	470	74	-18
Vermont	571	762	96	-61
Virginia	294	410	58	-46
Washington	428	624	98	-31
West Virginia	201	312	56	-50
Wisconsin	440	617	89	-48
Wyoming	320	390	35	-37
U.S. average	320	463	72	-54

Sources: Background Material and Data on Programs within the Jurisdiction of the Committee on Ways and Means, Washington, D.C., U.S.G.P.O., 1991, and private communications from Evelyn Mills, Department of Health and Human Services, Family Support Administration.

II. PAST STUDIES

The empirical basis for the argument that welfare programs increase fertility and illegitimacy dates back to Becker (1960), who developed a cost-benefit model for evaluating fertility decisions (see also Becker [1981] and Fuchs [1983]). Increases in the illegitimacy rate have stimulated numerous empirical studies to determine if the AFDC program, which lowers the cost of bearing an out-of-wedlock child, has an effect on childbearing. Table 3 contains a summary of several of the more prominent studies, dating back to 1970.[5]

In general, the studies can be classified into two groups: macro studies that use aggregate data from states or cities, and micro studies that use data on individuals. The early macro studies (Cutright, 1970; Winegarden, 1974) did not find any effect of AFDC benefits on fertility rates, possibly because they used small samples. Cutright only used twenty state observations, while Winegarden used fifty-one. However, studies by Janowitz (1976) and Southwick (1978) also used small samples (seventy-one SMSA and thirty-one state observations, respectively) and found that AFDC benefits have a significant effect on the percentage of AFDC families with illegitimate children. Unfortunately, one cannot draw any conclusions from these two studies because Janowitz found a positive effect and Southwick found a negative effect. Freshcock and Cutright (1979), using almost eight hundred observations from Statistical Areal Units, found that AFDC benefit levels have no effect on the teenage out-of-wedlock birthrate; however, they found that benefit levels positively affect the out-of-wedlock birthrate of white adult females and negatively affect the out-of-wedlock birthrate of black adult females. Three additional macro studies from the 1980s also yielded mixed results. Using state data from 1960 to 1980, Bernstam and Swan (1986) found a negative effect on the illegitimacy rate, yet Winegarden (1988), using Granger causation on annual time-series data from 1947 to 1983, found a positive effect for nonwhites. Hyatt and Milne (1991) used time-series data

(text continues on p. 13)

TABLE 3
Studies of the Effect of Welfare on Childbearing Decisions

Study	Data Set	Unit of Observation	Dependent Variables	Independent Variables	Estimation Technique	Results
Cutright (1970)	1950–66 NCHS	States (n=20)	Illegitimacy rates	Annual AFDC benefit per recipient	Cross-tabs	No effect
Placek & Hendershot (1974)	1972 AFDC payees in Tennessee	Welfare mothers, aged 15–44 (n=300)	The notion that welfare mothers (1) favor welfare status, (2) are extremely active sexually, (3) don't use contraceptives, and (4) want to become pregnant	Participation in AFDC	Cross-tabs	No effect
Polgar & Hiday (1974)	1965–67 NORC-NY	Married women with kids and single parents, aged 18–39 (n=304)	Additional births in given period	Participation in AFDC	Cross-tabs	No effect
Winegarden (1974)	1967 DHEW	States (n=51)	Number of recipient children aged 0–1 per 100 AFDC mothers	(1) AFDC benefit level for one adult and three children	2SLS	No effect
				(2) Increment to the grant for a three-child family relative to a one-child family		No effect
Presser & Salsberg (1975)	1970–72 NORC-NY	Women with recent first child (n=408)	Desire to have more children	Participation in AFDC	Cross-tabs	Negative significant effect

(table continues)

TABLE 3, continued

Study	Data Set	Unit of Observation	Dependent Variables	Independent Variables	Estimation Technique	Results
Janowitz (1976)	1968 DHEW and other U.S. government sources	SMSAs (n=71)	Illegitimate birth rate per 1,000 unmarried women, by age group	Welfare payment for a one-parent, one-child family	OLS	Positive significant effect for young nonwhites
Southwick (1978)	1973 DHEW	States (n=31)	Percentage of AFDC families with illegitimate children	AFDC guarantee level for a mother and three children	OLS	Negative significant effect
Freshcock & Cutright (1979)	1969 U.S. Census	SAUs (n=778)	Illegitimate births per 1,000 women, by age group	Maximum AFDC benefit in late 1967 in the appropriate state, divided by mean size of families on AFDC	OLS	No effect for teenagers, positive significant effect for whites, negative significant effect for blacks
Keefe (1983)	1970–71 AL-CA	Women on AFDC (n=3,155)	Whether or not mother in sample AFDC case became pregnant and carried child to term	(1) Basic welfare benefit, and (2) marginal benefit if had additional child	Logit	No effects

(table continues)

TABLE 3, continued

Study	Data Set	Unit of Observation	Dependent Variables	Independent Variables	Estimation Technique	Results
Ellwood & Bane (1985)	1976 SIE	Unmarried women (n=1,094 to 24,228)	Had a child in last year	AFDC benefit for one adult and three children	OLS	No effect
Bernstam & Swan (1986)	1960–80 DHHS and other U.S. government sources	States (n=500)	Illegitimacy rate	AFDC benefit level for one adult and three children	OLS	Negative effect
Winegarden (1988)	1947–83 DHHS and other U.S. government sources	Time-series (n=37)	Proportion of all births occurring out of wedlock, by race	Net gain for average adult female participant	Granger causation	Positive significant effect for nonwhites
Gonul (1988)	1979–83 NLSY	Females aged 14–22 without multiple births (n=1,407)	Marital status, work status, and fertility status	Basic AFDC guarantee level	Stepwise hazard function	Weak positive effect
Plotnick (1990)	1979–84 NLSY	Unmarried, childless females aged 14–15 in 1979 (n=1,184)	(1) Probability of having an out-of-wedlock birth by age 19 (2) Conditional probability of having an out-of-wedlock birth in a given year	Five indicators of state welfare policy including the monthly AFDC benefit plus the value of food stamps	(1) Logit (2) Discrete hazard model	Positive significant effect for whites, no effect for nonwhites

(table continues)

TABLE 3, continued

Study	Data Set	Unit of Observation	Dependent Variables	Independent Variables	Estimation Technique	Results
Duncan & Hoffman (1990)	1968–85 PSID	Black teenagers (n=874)	Probability of an out-of-wedlock birth associated with receipt of AFDC	(1) Guarantee level for a family of two	Logit	No effect
				(2) Potential earned income at age twenty		Significant negative effect
An, Haveman, & Wolfe (1990)	1987 PSID	Unwed young women (n=892)	Probability of AFDC receipt conditional on out-of-wedlock birth	Maximum AFDC benefit level for three kids and one adult	Bivariate probit with sample selection	No effect
Lundberg & Plotnick (1990)	1979–86 NLSY	Never-married, childless women aged 14–16 in 1979 (n=1,718)	Had a premarital birth	AFDC benefit level for one adult and three children and food stamp benefit	Nested logit	Positive significant effect for whites, no significant effect for blacks
Hyatt & Milne (1991)	1948–86 Various Canadian government sources	Canadian time-series (n=39)	Log of total fertility rate	Annual family allowance payments and child tax credit	OLS with IV	Small positive effect
Allen (1993)	1986 COC	Low-income women under age 45 (n=8,009)	Children born out of wedlock	Welfare income, liquid asset exemption level	Logit	Significant positive effects

(table continues)

TABLE 3, continued

Study	Data Set	Unit of Observation	Dependent Variables	Independent Variables	Estimation Technique	Results
Murray (1993)	1954–88 DHHS and other U.S. government sources	States over time (n=1,014)	Illegitimacy ratio (illegitimate births as a proportion of all live births)	Average AFDC benefit level plus food stamp benefit level for a family of four	Bivariate correlations over time	Significant positive effect for whites, significant negative effect for blacks (positive when controlling for population density)
Acs (1993)	1979–88 NLSY	Women aged 14–16 in 1979 (n=1,814)	First and second births by age 23	Maximum AFDC benefit (family of two for first births, family of three for second births); extra benefit for second birth	Logit model for hazard rates	Significant effect on first births for whites; no effect on second births

Acronymns: NCHS = National Center for Health Statistics; NORC-NY = National Opinion Research Center, New York data; DHEW = Department of Health, Education, and Welfare; SAU = Statistical Areal Unit; SMSA = Standard Metropolitan Statistical Area; AL-CA = Alameda County, California, data; NLSY = National Longitudinal Survey of Youth; PSID = Panel Study of Income Dynamics; IV = Instrumental Variable; SIE = Survey of Income and Education; COC = Census of Canada.

from Canada and found that government programs that reduced the cost of bearing a child had a small positive effect on the total fertility rate. Finally, Murray (1993) used a time series of data on states and examined bivariate correlations between AFDC plus food stamp income and the illegitimacy ratio (the ratio of illegitimate births to all live births). He found a significant positive correlation for whites but not for blacks. However, when he controlled for a measure of black population density, Murray found a positive correlation for blacks.

In general, studies using micro data have also yielded mixed results. Placek and Hendershot (1974) tested the Brood Sow hypothesis (the idea that women on welfare favor welfare status, are extremely active sexually, do not use contraceptives, and want to become pregnant). Their study of welfare mothers in Tennessee rejected the Brood Sow hypothesis. A study by Presser and Salsberg (1975) found that women on welfare do not have a desire to have more children. In fact, they found that participation in AFDC had a negative effect on the desire to have additional children. Ellwood and Bane (1985) also found no effect using 1976 data from the Survey of Income and Education.

Numerous studies have used longitudinal data to determine the effects of welfare benefit levels and other socioeconomic characteristics on the probability that young women will have an out-of-wedlock birth and join the welfare ranks. The results of these studies have also been inconclusive. Allen (1993), using Canadian data, found a significant positive effect. Gonul (1988), using data from the National Longitudinal Survey of Youth (NLSY), found a weak positive effect. Plotnick (1990), Acs (1993), and Lundberg and Plotnick (1990), also using NLSY data, found a positive effect for whites, but no effect for nonwhites. Duncan and Hoffman (1990) and An, Haveman, and Wolfe (1990), using Panel Study of Income Dynamics (PSID) data, found no effect. Different models were analyzed in each of these studies. Gonul (1988), Plotnick (1990), and Acs (1993) analyzed hazard models; Plotnick also analyzed logit models, as did Duncan and Hoffman (1990) and Allen (1993);

An, Haveman, and Wolfe (1990) analyzed a bivariate probit model with sample selection; and Lundberg and Plotnick (1990) analyzed a nested logit model.

Only two studies (Keefe, 1983; Acs, 1993) directly tested the hypothesis that increases in benefit differentials affect fertility decisions. Keefe (1983) analyzed a sample of over 3,100 AFDC families in California before and after the 1970 California AFDC benefit level increase to determine if the increase for additional children had an effect on fertility. Keefe found no evidence to support the hypothesis. Acs (1993), using NLSY data, examined whether the additional benefit for a second child influenced the probability of a second birth. Acs found no effect of either the basic benefit level or the additional benefit for a second child on the probability of a second birth. Acs did not test the hypothesis that the benefit differential might have an influence on first births.

In general, the results of previous studies have been mixed, but generally indicate no direct relationship between AFDC benefit levels (or differentials) and family size. Small sample sizes have plagued most of the analyses. In this study, we use a much larger sample and a somewhat different methodology to analyze the relationship between the AFDC benefit structure and the childbearing decisions of never-married women.

III. THEORETICAL CONSIDERATIONS

The decision to have children is influenced by a large number of economic and noneconomic factors. In the economics literature, modeling fertility decisions is a complex undertaking, particularly when a life-cycle framework is used and when labor supply (and other behavior) is considered to be jointly determined with fertility decisions. In principle, an economic model of fertility and labor supply decisions could be used to predict the behavioral effects of economic factors. As discussed in Blau and Robins (1989), in a standard model of fertility and labor supply behavior, the key economic variables are the wage rate and the "cost" of children (expenses incurred in bearing and raising children).

Within this framework, AFDC benefit levels and differentials can be thought of as "subsidies" that reduce the cost of children. In addition, AFDC benefit levels and differentials have a work disincentive effect that reinforces their positive effect on fertility decisions, because a reduction in labor supply reduces child care costs and hence reduces the cost of children.

A fully developed model of fertility and labor supply decisions is beyond the scope of this paper. Our objective is to specify a tractable reduced form model that focuses on the effects of economic variables on fertility decisions. In particular, we wish to determine whether AFDC benefit levels and differentials influence family-size decisions and whether these effects vary by a number of economic and demographic factors.

IV. EMPIRICAL FRAMEWORK

A. Data

The data used to estimate the effects of AFDC benefit levels and differentials on family size are drawn from the March Current Population Survey (CPS) for the years 1980 through 1988. Each March CPS contains demographic data for individuals at the time of the survey, and data on employment and income for these individuals during the calendar year prior to the survey. The CPS contains detailed responses to survey questions on the number and ages of children in the family, the age of the individual, race/ethnicity, and the education, earnings, and nonwage income of each member of the family.

The strategy adopted in this paper is to determine whether AFDC benefit levels and differentials that existed prior to the birth of a woman's first child are systematically related to the number of children she has at the time of the survey.[6] Note how this strategy differs from that taken in some previous studies. Several previous studies examined the effects of current AFDC benefits on family size at the time of the survey. However, these latter studies misspecify the AFDC benefit variables because family-size decisions are presumably based on past, rather than present, AFDC benefit levels. If benefit levels and differentials change over time in unanticipated ways, use of the current values will result in a misspecified model.

Our approach is to create a series of cohorts, based on cross-sectional data, and to construct, for each woman, variables based on the AFDC benefit levels that existed in her current state of residence prior to the birth of her first child. For purposes of this paper, we use the benefit levels and differentials in effect when the woman was eighteen years old. It should be noted that it would be possible to define the AFDC benefit variables precisely before the birth of the first child for each woman with children, because we know each woman's age. However, as will be described below, we include women without children in the sample and need to define a comparable set of AFDC benefit variables for them in a way that will not bias the results. We can't use current benefit levels for women without children because that would generate systematic positive correlation between family size and the benefit levels (the correlation would be positive because benefit levels have been falling uniformly over the time period covered in our sample). While benefit levels and differentials prevailing at age eighteen will be subsequent to the birth of the first child for many women, it should be reasonably close to the values that were used in the decision-making process.

One might argue that using benefit levels in the woman's current state of residence is not correct and that benefit levels in the state she actually resided in at age eighteen would be theoretically more appropriate. Of course, given the structure of the CPS, we do not know which state the woman lived in at age eighteen. However, using the AFDC benefit levels in her current state of residence may be theoretically justifiable because it incorporates any induced effects on migration for women that originally resided in another state. Thus, we implicitly assume in our analysis that when the woman plans her fertility decisions, she bases it, in part, on the benefit structure in her future state of residence.[7]

The CPS does not contain marital history information. Thus, we restrict our analysis to the sample of never-married women, to ensure that all sample members were categorically eligible for AFDC benefits at the time their first child was born (only unmarried women are eligible for benefits under the regular AFDC program).[8] Never-married mothers constitute almost one-half of the AFDC caseload and are perhaps the group of greatest concern to policymakers. Theoretically, the same woman could appear in our sample more than once (if she was interviewed in more than one survey year), but family size, which is measured as of the date of the survey, could change over time.

To avoid possible selection biases, we include never-married women without children in the sample and, like never-married women with children, we use AFDC benefit levels and differentials for them that existed in their current state of residence at age eighteen. Thus our empirical model captures the effects of AFDC benefit levels and differentials on both the probability of having children and the number of children born for those with children.[9] The model takes advantage of state and time variation in AFDC guarantee levels and differentials as a means of identifying their influence on family size decisions.

We restrict our analysis to never-married women between the ages of eighteen and thirty at the time of the survey. This is done for two reasons. First, we want a representative sample of women during their primary childbearing years. Second, we are only able to collect AFDC benefit levels by family size from 1968 to the present, so that women who were age thirty in 1980 were the oldest for whom we could construct the AFDC benefit variables (they were age eighteen in 1968). The resulting sample size is 74,355, of which 66,965 (90 percent) do not have children and 7,390 (10 percent) have one or more children. Table 4 shows the distribution of the sample with respect to family size in the nine survey years covered in our analysis.

TABLE 4

Distribution of the Sample, by Survey Year and Number of Children

Survey Year	Number of Children 0	1	2	3	4	5+	All Women	Mean Number of Children per 1,000 Women
1980	7,871	423	209	73	26	14	8,616	144
1981	8,066	458	202	90	20	13	8,849	146
1982	7,360	416	222	87	26	11	8,122	158
1983	7,510	428	209	78	31	15	8,271	156
1984	7,436	479	212	79	29	10	8,245	159
1985	7,416	487	241	88	32	15	8,279	174
1986	7,128	458	266	89	33	16	7,990	184
1987	7,137	497	267	115	31	15	8,062	197
1988	7,041	467	239	116	39	19	7,921	196
All years	66,965	4,113	2,067	815	267	128	74,355	168

Source: Current Population Surveys, March 1980–March 1988.

B. Econometric Specification

In our model, the dependent variable represents the number of children in the family and it is measured as an integer that varies between zero and nine. To analyze the distribution of this variable, we utilize the Poisson regression model, which is appropriate for models analyzing count data. The econometric specification for the Poisson model is given by the following pair of equations:

$$(1) \quad Prob(N_i = n_i) = \frac{e^{-\lambda_i} \lambda_i^{n_i}}{n_i!}, \quad n_i = 0,1,2,...$$

$$(2) \quad \ln \lambda_i = \beta x_i$$

where N_i is a discrete random variable representing the number of children in a family. N_i has observed frequencies, n_i. The vector of regressors (described below) is represented by $\mathbf{x}_i$.

It can be shown that λ_i represents the mean and the variance of the Poisson distribution.[10] In addition, the probability of having at least one child is given by 1-exp(-λ_i), and the mean number of children for those with children is given by λ_i/(1-exp(-λ_i)).

The Poisson model may not adequately describe the distribution of family size in our sample. To test the robustness of our results to the distributional assumption, we estimate two additional models. One is a simple ordinary least squares (OLS) model. The other is a probit model, in which the dependent variable is equal to one if the woman has at least one child, and zero otherwise. Given the definitions in the previous paragraph, the probit model results can be directly compared to the Poisson model results.

C. Variables

All of the models reported in this paper include the following explanatory variables: dummy variables for geographic region (NE, NC, WEST, SOUTH), dummy variables for race/ethnicity (WHITE, BLACK, HISPANIC), the woman's age at the time of the survey (AGE), dummy variables for whether the woman is a high school graduate (HS) or dropout (NHS),[11] the woman's nonwage income (PNONWAGE),[12] the effective AFDC tax rate on nonwage income in the current state of residence at age eighteen (TAX),[13] and a series of dummy variables for each survey year (Y81 to Y88, omitting Y80).[14]

For all women in the sample, an average guarantee level differential is calculated for the year the woman reached age eighteen. The guarantee level differential is measured as the average difference between the guarantee level for three children and the guarantee level for one child in the relevant current state of residence (G31).[15] The differential is measured in real 1991 dollars. In addition to the average differential variable (G31), the guarantee level for one child (G1) is included to capture state differences in the basic level of AFDC benefits.[16] Interaction variables are also included in some specifications to allow for different effects by race/ethnicity, whether or not the woman finished high school, and her age. Significance tests are performed for each set of interactions.

V. EMPIRICAL RESULTS

Table 5 presents the empirical results for the AFDC guarantee variables for the three models (Poisson, OLS, and probit). The full set of parameter estimates for the models without any interaction terms is presented in the appendix.

The results from the base model (no interaction terms) suggest that the differential for additional children does not significantly influence childbearing decisions of never-married women. In fact, the coefficient of the differential variable is negative in every model. The results do indicate, however, that childbearing decisions are influenced by the basic benefit level. In the Poisson model, the coefficient implies an elasticity for the basic benefit level of .135 (.00031*434.32). A comparison of the Poisson and probit results reveals that the basic benefit level mainly affects the probability of

TABLE 5

Empirical Results for the Effects of AFDC Benefits on Family Size (Standard Errors in Parentheses)

	OLS	Poisson	Probit
No Interaction Terms			
G31	-.000087	-.00021	-.00032
	(.00011)	(.00047)	(.00036)
G1	.000053**	.00030***	.00040***
	(.000024)	(.00011)	(.000082)
Race/Ethnicity Interactions			
WHITE*G31	-.00054***	-.0015**	-.0011***
	(.00012)	(.00071)	(.00043)
WHITE*G1	-.000033	.00042***	.00024**
	(.000026)	(.00016)	(.000096)
BLACK*G31	.00034	-.000078	.00072
	(.00030)	(.00068)	(.00075)
BLACK*G1	.00032***	-.000093	.00031*
	(.000065)	(.00015)	(.00017)
HISPANIC*G31	.0017***	.0019	-.00074
	(.00046)	(.0013)	(.0013)
HISPANIC*G1	.00018*	.00091***	.0013***
	(.00010)	(.00030)	(.00028)
Test statistic for interactions	F = 130.7***	χ^2 = 104.5***	χ^2 = 92.1***
Education Interactions			
NHS*G31	.0022***	.00049	.00092
	(.00027)	(.00072)	(.00076)
NHS*G1	.00022***	.00049***	.00075***
	(.000059)	(.00016)	(.00016)
HS*G31	-.00048***	-.00075	-.00072*
	(.00011)	(.00058)	(.00040)
HS*G1	.000010	.00013	.00027***
	(.000025)	(.00013)	(.000092)
Test statistic for interactions	F = 285.1***	χ^2 = 37.5***	χ^2 = 80.4***

(table continues)

TABLE 5, continued

	OLS	Poisson	Probit
Race/Ethnicity and Education Interactions			
WHITE*G31*NHS	0.00024	0.0014	0.00052
	(0.00033)	(0.0012)	(0.00094)
WHITE*G1*NHS	0.00014**	0.00088***	0.00060***
	(0.000070)	(0.00026)	(0.00020)
BLACK*G31*NHS	0.00079	-0.0021**	0.00017
	(0.00057)	(0.0010)	(0.0014)
BLACK*G1*NHS	0.0010***	0.00016	0.00047
	(0.00013)	(0.00022)	(0.00031)
HISPANIC*G31*NHS	0.0068***	0.0058***	0.0035*
	(0.00078)	(0.0016)	(0.0020)
HISPANIC*G1*NHS	-0.00044**	0.00015	0.00068
	(0.00018)	(0.00037)	(0.00045)
WHITE*G31*HS	-0.00069***	-0.0022***	-0.0014***
	(0.00012)	(0.00085)	(0.00048)
WHITE*G1*HS	-0.000061**	0.00026	0.00015
	(0.000027)	(0.00019)	(0.00011)
BLACK*G31*HS	0.00041	0.00096	0.00072
	(0.00034)	(0.00087)	(0.00085)
BLACK*G1*HS	0.000048	-0.00029	0.00022
	(0.000075)	(0.00020)	(0.00019)
HISPANIC*G31*HS	-0.00068	-0.0046**	-0.0037**
	(0.00054)	(0.0020)	(0.0016)
HISPANIC*G1*HS	0.00042***	0.0020***	0.0017***
	(0.00012)	(0.00045)	(0.00037)
Test statistic for interactions	F = 146.3***	χ^2 = 519.3***	χ^2 = 237.0***
Mother's Age Interactions			
G31	.00016	.00077	-.0021
	(.00063)	(.0032)	(.0023)
G1	.00023	.0012*	.0017***
	(.00014)	(.00069)	(.00049)
AGE*G31	-.000011	-.000038	.000070
	(.000026)	(.00012)	(.000089)
AGE*G1	-.0000077	-.000035	-.000054***
	(.0000058)	(.000026)	(.000020)
Test statistic for interactions	F = 3.8**	χ^2 = 9.4***	χ^2 = 13.9***

Source: Authors' calculations based on Current Population Surveys, March 1980–March 1988.
*Significant at the 10 percent level.
**Significant at the 5 percent level.
***Significant at the 1 percent level.

having at least one child; it does not appear to affect the number of children for women who already have children.[17]

For all three models, the test statistic indicates that the effects of the AFDC benefit variables differ by race/ethnicity. In the Poisson model, the differential is significantly negative for whites, insignificantly negative for blacks, and insignificantly positive for Hispanics. The basic benefit level is significantly positive for whites (with an elasticity of .197), insignificantly negative for blacks (although it is significantly positive in the OLS model), and significantly positive for Hispanics (with an elasticity of .422). The effect for whites is consistent with the findings of Freshcock and Cutright (1979), Plotnick (1990), and Lundberg and Plotnick (1990), but the effect for Hispanics is not. Thus, our results for the Poisson model imply that the basic benefit level is exerting a significant influence on family-size decisions of whites and Hispanics, but not of blacks. Our results for the probit model, on the other hand, imply that the basic benefit level is exerting a significant influence on the decision of whether or not to have children for all three racial/ethnic groups.

The test statistics also indicate that the effects of the AFDC benefit variables are different for high school dropouts and graduates. In the OLS model, both the differential and the basic benefit level have a significant positive effect on family size for high school dropouts (about 17 percent of the women in our sample reported that they did not finish high school). For high school graduates, the differential has a significant negative effect and the basic benefit level has an insignificant positive effect.

In the Poisson and probit models, the effect of the differential is not statistically significant for high school dropouts and is negative and statistically significant for high school graduates. Generally, the directions of the effects are the same as those in the OLS model. The effect of the basic benefit level for dropouts is significantly positive in all three models, while for high school graduates it is statistically significant only in the probit model. In the Poisson model, the elasticity of the basic benefit level for dropouts is .223, while for graduates it is .065 (but not statistically significant). For the differential, the elasticity for dropouts is .047, a very small effect that is also not statistically significant.

To determine whether the effects of educational attainment on family size differ by race/ethnicity, the race/ethnicity and education variables are interacted with the differential and basic benefit level variables. Once again the test statistics indicate that the interactions are statistically significant. The Poisson results suggest a significantly positive effect of the differential for Hispanic high school dropouts, but not for white or black high school dropouts (the effect is significantly negative for black high school dropouts). In the case of the basic AFDC benefit level, the effect is significantly positive for white dropouts (with an elasticity of .392), but not for black or Hispanic dropouts. The effect of the basic benefit level is significantly positive, however, for Hispanic graduates, with an elasticity of .204.

The final set of interactions we tested for was age of the woman. The age interactions are jointly statistically significant in all three models, but the test statistics are not as large as they are for the other interactions and only one of the individual interaction terms is statistically significant (AGE*G1). The results show no significant variation in the effect of the differential with age, but do show a significant effect of the basic benefit level with age in the probit model. The results imply that the effect of the basic benefit level on family-size decisions is strongest for younger women. Since the age variable is measured as of the date of the survey, this result suggests that the effects of the basic benefit level have been increasing over time.

VI. CONCLUSIONS

Using a much larger sample than has been used in previous studies, we have been able to obtain more precise estimates of the influence of AFDC benefit levels and differentials on family-size decisions of never-married women. Our estimates of a Poisson regression model suggest that the basic AFDC benefit level for a family of two (one adult and one child) exerts a significant influence on the family-size decisions of white and Hispanic women, but not on those of black women. We also find that the additional benefits for larger families do not appear to be exerting an overall effect on family size, suggesting that eliminating benefit differentials for larger families will not influence family-size decisions.

When the AFDC benefit variables are interacted with the educational attainment of the woman, a number of significant positive effects of both the basic benefit level and the incremental benefit for larger families are found for high school dropouts. For high school graduates, the effects are generally smaller in magnitude than for high school dropouts, and the effects of the benefit differentials for high school graduates are generally negative. These results suggest that cutting welfare benefits for larger families, as some states are now doing, may be a misguided policy; a better strategy might be to alter the benefit structure in such a way as to encourage single mothers to complete high school. Such a strategy would give single mothers better labor market opportunities, which in turn might discourage nonmarital births and encourage greater self-sufficiency through employment. It should be noted, however, that being a high school dropout might be a proxy for some other underlying characteristic of the woman and that inducing women to complete high school who otherwise would not might have no effect whatsoever on nonmarital births. On the other hand, there might be a direct causal relationship between completion of high school and these other factors that lead to nonmarital births.

Some states seem to be recognizing the importance of completing high school on reducing welfare dependency. A recent program put into effect in Ohio called LEAP (Learning, Earning, And Parenting) provides an incentive for welfare mothers to complete high school by giving them an extra $62 (or 22 percent) in additional benefits for every month they stay in school and by cutting benefits by $62 for every month they do not stay in school.[18] Our analysis of CPS data suggests that this approach might be a more desirable social policy than simply cutting the benefits of welfare mothers who have additional children.

APPENDIX

Full Results
(Base Model with No Interaction Terms)

	OLS	Poisson	Probit	Means
Intercept	-0.35***	-5.08***	-2.90***	--
	(0.046)	(0.16)	(0.14)	
NE	0.048***	0.20***	0.097***	0.25
	(0.0080)	(0.037)	(0.028)	(.43)
NC	0.077***	0.37***	0.19***	0.24
	(0.0072)	(0.031)	(0.025)	(0.43)
WEST	0.043***	0.25***	0.14***	0.24
	(0.0078)	(0.036)	(0.028)	(0.43)
BLACK	0.41***	1.72***	1.03***	0.16
	(0.0058)	(0.022)	(0.017)	(0.37)
HISPANIC	0.18***	1.14***	0.53***	0.10
	(0.0070)	(0.028)	(0.021)	(0.30)
AGE	0.031***	0.14***	0.078***	22.12
	(0.00083)	(0.0036)	(0.0028)	(3.39)
PNONWAGE	-0.0000028***	-0.000046***	-0.000020***	422.10
	(0.00000085)	(0.0000079)	(0.0000048)	(2353.5)
FS31	-.00014	-.0066*	-.0082**	75.99
	(.0012)	(.0040)	(.0036)	(3.70)
FS1	-.00018	.0026*	.0025*	185.65
	(.00046)	(.0015)	(.0014)	(9.68)
TAX	-0.061***	-0.45***	-0.30***	0.40
	(0.015)	(0.075)	(0.056)	(0.19)
HS	-0.31***	-1.16***	-0.73***	0.83
	(0.0055)	(0.019)	(0.017)	(0.38)
Y81	-0.00010	0.033	0.027	0.12
	(0.0083)	(0.040)	(0.030)	(0.32)
Y82	0.011	0.090**	0.059**	0.11
	(0.0085)	(0.041)	(0.031)	(0.31)
Y83	0.0084	0.071*	0.049	0.11
	(0.0087)	(0.042)	(0.031)	(0.31)
Y84	0.012	0.082*	0.098***	0.11
	(0.0089)	(0.043)	(0.032)	(0.31)
Y85	0.023**	0.16***	0.14***	0.11
	(0.0091)	(0.042)	(0.032)	(0.31)
Y86	0.035***	0.23***	0.18***	0.11
	(0.0094)	(0.043)	(0.033)	(0.31)
Y87	0.045***	0.30***	0.21***	0.11
	(0.0095)	(0.043)	(0.033)	(0.31)
Y88	0.045***	0.27***	0.20***	0.11
	(0.0097)	(0.044)	(0.034)	(0.31)
G31	-0.000087	-0.00021	-0.00032	100.63
	(0.00011)	(0.00047)	(0.00036)	(36.53)

(table continues)

APPENDIX, continued

	OLS	Poisson	Probit	Means
G1	0.000053** (0.000024)	0.00030*** (0.00011)	0.00040*** (0.000082)	434.32 (169.46)
Log likelihood		-30472.11	-19568.04	
Adj R^2	.146			
Mean of dependent variable	0.1677 (0.5906)	0.1677 (0.5906)	0.0994 (0.2992)	

Source: Authors' calculations based on Current Population Surveys, March 1980–March 1988.

Note: Standard errors in parentheses.

***Significant at the 1 percent level.
** Significant at the 5 percent level.
* Significant at the 10 percent level.

Notes

[1]It should be noted that the proportion of all women who are unmarried has been increasing since 1965 (from 30 percent to 39 percent in 1989) and that the rising trend in nonmarital births over this period may be partially reflecting a higher birthrate among women who would have been previously married.

[2]The differential increases somewhat less than proportionately for larger families. For example, in 1991, the benefit for a family with five children was roughly 30 percent higher than the benefit for a family with three children.

[3]See, for example, Bassi et al. (1990) for a summary of studies of expenditures on children.

[4]On November 1, 1992, New Jersey enacted a law in which a welfare recipient with two children who has a child at least ten months after enrolling in AFDC will no longer receive the usual $64 increase in benefits. In California, a proposed law will also eliminate benefit increases. The family would have to be off AFDC for twenty-four consecutive months before they would be eligible for a cash grant increase for additional children. Other states having similar proposals are Arkansas, Wisconsin, Connecticut, Florida, and Georgia. Additionally, Maryland's governor has proposed offering free contraceptives to women on welfare; the birthrate in Baltimore among teenage girls is triple the U.S. rate.

[5]Some of these studies, plus a few additional ones, are discussed in Hofferth (1987) and Murray (1993).

[6]It is important to point out that we are not examining <u>completed</u> family size for all families. Variation in family size in our sample arises through variation in the ages of the women in our sample, variation in fertility decisions across individuals, and variation in fertility decisions over time.

[7]This may be an unduly restrictive assumption for an older woman with several children, particularly if she moved several times between the birth of her first child and the survey year.

[8]During our survey period, two-parent families in about half of the states were eligible for the AFDC-UP (Unemployed Parent) program, but the eligibility criteria (unemployment of both heads) was very restrictive and few two-parent families participated in the program.

[9]Because we restrict our analysis to a sample of never-married women, another possible selection bias may be present in that the benefit levels and differentials would be influencing the decision to be unmarried rather than the decision to have children, given that a woman is unmarried. Prior research suggests, however, that welfare benefit levels are not systematically related to marital status decisions (Moffitt, 1992). Nonetheless, a more complete analysis would incorporate marriage decisions along with family-size decisions, but such an analysis is not possible with the CPS data because of the absence of marital history information.

[10]A distribution with the mean equal to the variance may be an unduly restrictive assumption. The negative binomial model, which allows the mean and variance to differ, is an attractive alternative to the Poisson model. We utilized both specifications, but the parameter estimates from both were virtually identical. For simplicity, we report the Poisson regression model results in this paper.

[11]The high school graduate variable is used as a proxy for the woman's potential wage rate at age eighteen. It should be a reasonable proxy for most sample members. Whether the mother is currently a high school graduate will be a poor proxy for the potential wage rate at age eighteen in the cases of women who obtained their high school diploma after age eighteen. Due to the structure of the CPS questionnaire, the high school graduate variable is based on reported years of education rather than whether a diploma was actually received.

[12]Nonwage income (PNONWAGE) is calculated by summing the individual's reported values of social security income, interest income, dividend income, rental income, and pension income. Public assistance income is not included in nonwage income. This variable is expressed in 1991 dollars, and is deflated using the Consumer Price Index.

[13]We use the effective tax rates in the AFDC program reported in Fraker, Moffitt, and Wolf (1985). Values are extrapolated for years not covered by their data.

[14]All models were estimated with and without the survey year dummies and the results are available from the authors on request. Although the survey year dummies were often statistically significant, the coefficients of the other variables were virtually unaffected.

[15]We had originally planned to use the average differential between one and five children, but guarantee levels for four and five children were not available for the early years in our sample. We estimated models using the actual average differential for one and five children for years we had this information and an extrapolated differential for one and five children for the other years. The results, which are available on request from the authors, produced results similar to those reported in the text. This is mainly because the differentials remain close to proportionate for more than three children.

[16]Murray (1993) used the combination of AFDC and food stamps benefits as a more comprehensive measure of welfare income. All models reported in this paper were estimated using a measure of food stamp income in addition to AFDC income. Two specifications were tested: (1) a model with a separate variable for the food stamp basic benefit level for a family with one child (FS1) plus the average food stamp differential for families with one and three children (FS31), and (2) a model with the food stamp basic benefit level and differential added to the AFDC basic benefit level and differential. Only the results from the first specification are reported here; the results from the second specification are available on request from the authors. The results including the food stamp variables were virtually identical to the results using the AFDC variables only. The food stamp data were obtained from D. H. Moritz of the Food and Nutrition Service, U.S. Department of Agriculture. The food stamp benefit levels were deflated by both the overall CPI and by the food component of the CPI; the results were insensitive to which deflator was used.

[17]We also estimated the OLS and Poisson models using only the sample of women with children. Although potentially subject to selectivity biases, the results confirmed that the effect of the basic benefit level is operating primarily on the probability of having at least one child and not on the number of children for those with children. In fact, in the Poisson model on the restricted sample of women with children, the effect of the basic benefit level on family size is negative and statistically significant. The effect of the differential is not statistically significant in any model.

[18]See Bloom et al. (1993) for a discussion of the LEAP program and its impacts on school attendance during the first three years of the program.

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