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TRENDS IN AFDC PARTICIPATION OVER TIME: Evidence on Structural Change

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Executive Summary

The sudden and dramatic increase in the growth rate of the caseload in the Aid to Families with Dependent Children (AFDC) program in the late 1960s and early 1970s remains the most important historical change in the caseload of our most well-known cash transfer program. The caseload in the program grew from about 1.3 million to about 2.5 million in the three years between 1967 and 1970, and increased steadily thereafter into the mid-1970s. At that point the caseload growth rate slowed and, in more recent times, has leveled off. Much of this growth resulted simply from an increase in the number of female heads of household in the United States, but this was not the only cause. Among female heads, participation rates in AFDC rose dramatically from 1967 to 1973, then grew more slowly, and have been falling since around 1979.

The question considered in this report is whether this pattern of increase and decline, particularly the period of increase, can be attributed to measurable forces: changes in the benefit levels and work incentives in the AFDC program; changes in the education, age, and other characteristics of female heads; changes in the labor market (e.g., the unemployment rate or potential earnings); changes in the level of other income available to female heads; and so on. The primary question is whether these variables can explain much of the time-series trend in AFDC participation rates. Other questions then follow. If so, which ones are more important? If not, does this constitute evidence that the "structure" of the AFDC participation decision has changed? The issue here is an old one: how much of the time-series change in AFDC participation rates has been a result of changes in economic and demographic

forces, and how much has been a result of the whole set of noneconomic forces so frequently discussed--changes in attitudes, reductions in stigma, and changes in eligibility requirements for AFDC (such as elimination of residency and man-in-the-house rules)? Such changes are almost impossible to quantify, so the approach taken here is only indirect: the shift in participation rates over time is decomposed into that portion due to the measurable economic forces and that portion due to all else, and the latter category is interpreted as an approximation of the magnitude of the noneconomic forces.

The answers to these questions are not merely of historical interest, but are highly relevant to current policy questions and to our current ability to manage the AFDC program. At issue is whether we have the capability to accurately forecast changes in the AFDC caseload in response to policy alterations (changes in the level of benefits, for example) that might be instituted. Clearly it is desirable to be able to specify with some certainty the determinants of AFDC participation rates, and we can do so only by determining whether the past experience accords with our estimates of the effects of those determinants.

The evidence obtained here provides a strong answer to the main question, for virtually all of the evidence adduced indicates that there was a major structural shift in the AFDC participation equation between 1967 and 1973. Neither AFDC benefits, earnings deductions, the characteristics of the labor market, nor characteristics of female heads can explain much of the time-series change. There is a large, unexplained residual that is plausibly ascribed to the noneconomic forces mentioned above. However, the evidence does show that AFDC participation rates are somewhat responsive to simple manipulation of AFDC benefit parameters--i.e., benefit levels and earnings deductions.

It is also shown in this report that this finding sheds light on the rather different issue of work incentives in AFDC. One of the puzzles in the evidence on work incentives in the program is the sharp reduction in work-effort levels of female heads following the provision of the 30-and-one-third earnings deductions in 1967--a provision that would have been expected to increase work effort because it permitted workers to keep more of their earnings. It is shown in the report that more than 80 percent of the reduction in the level of work effort among female heads from 1967 to 1973 can be explained by the surge of women onto the AFDC rolls, an event that (according to the main findings of the report) was unrelated to AFDC benefit parameters. The 30-and-one-third deduction itself increased participation rates slightly by raising the AFDC break-even level and thus increasing eligibility, but very little of the increase in participation rates was a result of this factor.

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I. INTRODUCTION

The policy legacy of the welfare explosion in the late 1960s and early 1970s is still strong in the 1980s. Although the growth of the caseload in the Aid to Families with Dependent Children (AFDC) program has leveled off, and although participation rates began to fall around 1979, the absolute size of the caseload is still much higher than it was in the mid-1960s. Combined with the increased budgetary pressures facing the federal government, this historical growth in the caseload creates strong policy interest in controlling caseloads and expenditures in the program.

The main tools that have been used in the past to alter the caseload of the program have been the setting of eligibility requirements and the setting of benefit levels. Other tools, such as the alteration of earnings deductions, work through both channels by changing both the level of benefits and the income eligibility point. Alterations in the maximum amount paid in the program, holding earnings deductions constant, also affect both benefit levels and eligible income levels.

It is important for policy purposes to be able to determine with reasonable certainty the magnitude of the effects of these tools on the participation rate in the program. It is also important to be able to determine the effect of other forces in the economy and society on the caseload, not only for the ability to forecast caseloads in the future but also because of its implications for more general policy toward single mothers heading households. For example, it is important to know the extent to which reductions in the unemployment rate affect the caseload, the extent to which providing female heads with more education and (hence) labor market opportunities would reduce the caseload, and other such examples.

This study reports the results of a statistical examination of the determinants of welfare participation from 1967 to 1979. This is an important historical period for the program, for the period 1967-1973 was a period of rapid growth in the caseload and in participation rates, while the period 1973-1979 was period of declining growth rates and participation rates. The study focuses on two separate questions. First, how is cross-state variation in participation rates affected by AFDC benefit levels, deductions, local labor market conditions, and differing characteristics of AFDC recipients? Second, can the time-series variation in participation rates be explained by these same variables? That is, can the increase in participation rates from 1967 to 1973 be explained by any measurable forces? Or was there a structural shift in participation propensities over this period?

The statistical examination employs data at three points in time--1967, 1973, and 1979. In each of those years, data on the female-head population as a whole from the Current Population Survey and data on the AFDC caseload from the AFDC Characteristics Surveys are combined to estimate equations for the determinants of participation rates in AFDC. Data on AFDC benefits and earnings deductions in each state in each year are employed in the equations, along with data on state unemployment rates and the sociodemographic characteristics of individual female heads.

The results of the analysis indicate that AFDC guarantees have a statistically significant positive effect on AFDC participation rates, and

AFDC benefit-reduction rates have a negative, though weak, effect on participation rates in each of the three years. The variables for educational level, number of children, and unemployment rate all have significant effects on participation rates as well--negative, positive, and positive, respectively. However, the analysis across years indicates that none of these variables explains much of the time-series variation in participation rates. In particular, the sharp increase in participation rates from 1967 to 1973 is not explained by changes in the AFDC guarantee--which fell, in any case--or by changes in the benefit-reduction rate in the program. Over the 1967-1973 period, the 30-and-one-third earnings deductions were implemented, and the evidence here indicates that the resulting increase in the break-even level increased participation rates by about four percentage points. The rest of the increase in AFDC participation rates over the period (about eighteen percentage points) is explained neither by this variable nor by the others in the analysis. This constitutes strong evidence of a structural shift in the AFDC participation rate equation, for there is a large unexplained residual in the time-series analyses.

These results are also applied to the analysis of the time-series pattern of hours of work of female heads. One of the puzzles in the analysis of the labor supply patterns of this group is the sharp reduction in hours of work from 1967 to 1973, despite the provision of work incentives by the 30-and-one-third deductions. The evidence here indicates that the reduction in hours of work over the period was primarily a result of the surge of female heads onto the rolls, a surge which independently lowered the overall labor supply of the female-head population.

The 30-and-one-third deductions were not the cause of this surge, but they did have a slight impact by preventing hours of work from falling as much as they would have otherwise. However, this impact was negligible compared to the hours reductions coming from the enormous increase in participation rates.

The outline of the report is as follows. In the next section, Section II, some background on the issue is provided. Trends in participation rates in the AFDC program and previous evidence on the determinants of these trends are discussed. In Section III the data bases employed in this report and trends in those data are presented. The following section discusses the econometric estimating techniques used in the empirical analysis. The results of the estimation of the equations are shown in Section V. A summary and conclusion ends the report in Section VI.

II. BACKGROUND

A. Trends in Participation

Table 1 shows the trends in AFDC participation rates and in related variables over the period 1967-1982. The first row shows the trend in the total participation rate, that is, the percentage of all female heads who are on AFDC (whether eligible or ineligible). As documented in the source of these figures (Moffitt, 1985), these numbers are point-in-time estimates, equal to the number of female heads on the program as of a certain date (the date of the AFDC Characteristic Surveys) divided by the number of female heads in the week of the March Current Population Survey. Section III clarifies the measurement of this variable.

As the table indicates, about 28 percent of female heads were on the program in 1967. This rose to 38 percent in 1969 and 47 percent in 1971. The participation rate peaked in 1973, when it reached over 49 percent. After a more or less stationary period from 1975-1979, it began to decline. The rate took a sudden extra downward jump in 1982, the year of the implementation of the Omnibus Budget Reconciliation Act (OBRA). The table thus shows a familiar pattern of explosive growth in the AFDC rolls in the late 1960s and early 1970s, followed by a leveling-off and then a slow decline. Note that by 1982, the participation rate had fallen three-fourths of the way back to its 1967 level from its peak in 1973.

The second row shows the participation rate among eligible female heads, from Michel (1980), whose calculations were made at a time when data were available only up to 1977, but which have been updated through 1982. Participation rates among eligibles are, of course, higher than

Table 1

AFDC Participation Rates and Related Variables

	1967	1969	1971	1973	1975	1977	1979	1981	1982
Participation rate: all female heads (%)[a]	27.9	37.5	47.2	49.4	47.7	49.0	47.7	41.6	35.1
Participation rate: eligibles (%)[b]	45.0	56.0	81.0	91.0	91.0	94.0	97.0[c]	88.0[c]	79.0[c]
AFDC program parameters[d]									
Guarantee[e]	161	156	153	149	147	141	131	113	111
Benefit-reduction rate[f]	41	42	23	22	30	33	32	24	70[g]
Break-even level[h]	393	371	665	677	490	427	409	471	158
Unemployment rate (national)	3.8	3.5	5.9	4.9	8.5	7.0	5.8	7.6	9.7
Real full-time earnings[i]	4333	4763	4864	5005	4919	5028	5003	4814	4956
Guarantee/earnings (X 100)	44	40	37	36	36	34	31	28	26
Food stamps + AFDC[j]	--	247	235	233	243	232	219	199	192

[a]From Moffitt (1985).

[b]From Michel (1980).

[c]Updates provided by Richard Michel.

[d]From Fraker et al. (1985).

[e]Family of four in 1967 dollars (monthly).

[f]Effective tax rate on earned income (percentage).

[g]Without 30-and-one-third.

[h]Guarantee divided by benefit-reduction rate.

[i]Annual, for women working full-time, full year, in 1967 dollars. From U.S. Bureau of the Census, Current Population Reports, Series P-60.

[j]Monthly, in 1967 dollars.

those in the total female-head population, but both show the same pattern over the period 1967-1982. By the late 1970s, participation rates among eligibles appear to have come fairly close to 100 percent according to these calculations, but have since declined. Michel's calculations were made to replicate and improve upon those of Boland (1973).

It should be noted at this point that these estimates of participation rates are based upon estimates of the number of AFDC-eligible female-headed families from the Current Population Survey. As Ellwood and Bane (1985) have recently noted, many female-headed subfamilies were miscoded in the Survey. Proper coding began in 1982, so the estimates of participation-rate trends from 1981 to 1982 in Table 1 may be in error. It is probably the case that the trends in participation rates prior to 1982 were not affected in a major way by the miscoding. Current research on determining the extent of the miscoding prior to 1982 should shed more light on the seriousness of the problem.

Our primary interest here is in explaining the determinants of AFDC participation, including the trends in Table 1. The remaining rows of Table 1 show the trends in some of the leading candidates for explanation. The real AFDC guarantee (the amount paid to a family of four with no other income) fell continuously in real terms over the entire period. Thus, while the AFDC guarantee may have contributed to the decline in participation rates in the latter half of the period--the rate of decline in the guarantee did accelerate with time--there is no prima facie evidence for its having contributed to the increase in participation in the early part of the period.

It should be noted, however, the benefits in the AFDC program did increase to a significant extent during the period prior to 1969. From

1960 to 1968 the real guarantee rose by 8 percent, with most of this increase coming in the four-year period between 1964 and 1968 (U.S. House of Representatives, Committee on Ways and Means, 1986, p. 578). Though seemingly small in percentage terms, these increases were considerably larger than real growth rates of benefits in the 1950s over the same lengths of time. Since the most rapid increase in the participation rate occurred between 1967 and 1971 (see Table 1), a lag of about three years in the responsiveness of participation to benefit jumps could explain the participation-rate increase. The plausibility of this type of lag is heightened by subsequent experience, for the rate of increase in the participation rate began slowing in 1971, just three years after the real guarantee seems to have begun to decline. However, for the lag to be a complete explanation, the responsiveness of participation to the benefit would have to be fairly large--a mere 8 percent increase in the real guarantee would have to induce a rise in the participation rate of over 19 percentage points (1967-71).

The benefit-reduction rate (BRR) declined from 1967 to 1973, then began to rise, declined in 1981, and took a sharp jump in 1982 as a result of OBRA. The initial decline, which was a result of the earnings deductions provided in the 1967 Social Security Amendments (which provided the well-known "30-and-one-third" deduction), could have contributed to the increase in the total participation rate simply because the break-even level rose. This is shown explicitly in the table, which indicates that the break-even level was much higher in 1973 than in 1967.[1] However, the fact that the participation rate among eligibles rose simultaneously suggests that not all of the increase in the total participation rate was a result of the decline in the BRR, for there is no a priori reason to expect the BRR to affect the participation rate among eligibles.[2] Nevertheless, the subsequent increase in the BRR and the consequent decline in the break-even level may certainly have played a role in the decline in participation rates from 1977 to 1979. The low participation rate in 1982 is clearly a result of the low break-even level in that year, which in real terms was much lower than that in 1967 (though participation had already fallen in 1981, despite an increase in the break-even level).[3]

The responsiveness of the AFDC participation rate to the state of the national economy is also frequently discussed. As the trends in the unemployment rate in Table 1 illustrate, while there is slight evidence of the expected procyclical trend in participation rates, it is quite mild. Indeed, unemployment rates were considerably higher in the latter half of the period than in the former, while participation rates were the opposite. Thus the unemployment rate does not appear to be a likely candidate for explanation of the time-series trend.

The next two rows of Table 1 show how labor market indicators of earning power changed over the period. The real earnings of full-time, full-year working women rose from 1967 to 1973, more or less leveled off until 1979, and then declined to 1981. Thus there is no evidence of any decline in the attractiveness of the labor market that could explain the increase in participation rates in the early years. Also, as the table indicates, the ratio of the AFDC guarantee to the earnings variable shows that the relative attractiveness of AFDC declined over the early years.

The last row of Table 1 shows the trend in the sum of real AFDC and Food Stamp benefits. Food stamps became available in the late 1960s and were provided in a gradually increasing number of counties into the early 1970s. In 1974 the federal government mandated that all counties provide such benefits. Subsequently, food stamps were indexed to the inflation rate, thereby protecting their real value from price inflation. As Table 1 shows, the addition of food stamps implies that the benefit from the welfare option appears to have increased by 45 percent from 1967 to 1971 ($161 to $235), the period of the greatest participation-rate increase as well. Although the presence of food stamps was not sufficient to keep the sum of AFDC and food stamps from declining over the rest of the 1970s and early 1980s, it does seem that food stamps may provide an explanation for the participation-rate increases in the early period.

Unfortunately, there are several serious difficulties with this explanation. First, a free food commodities program was in existence in the 1960s even before the Food Stamp program began. There are no available data on the value of such commodities, but this must imply that the increase in available food benefits was less than the 45 percent implied by Table 1. Second, and more important, one of the most important characteristics of the Food Stamp program is that its benefits are made available to all income- and asset-eligible families, regardless of their participation or lack of same in other transfer programs, such as AFDC. Consequently, there is no logical reason for the introduction of food stamps to have increased the attractiveness of being on AFDC. Although there are no data from the early 1970s on the number of non-AFDC female heads receiving food stamps, data from 1979 indicate that almost one quarter of all female heads receiving food stamps were not, in fact, on AFDC (Weinberg, 1985, p. 72; Fraker and Moffitt, forthcoming).

Third, and also quite important, the Food Stamp program "taxes" the AFDC benefit by reducing the Food Stamp bonus by 30 cents for every dollar of AFDC received by a family. As a result, the increase in the real benefit of the welfare option was only 25 percent instead of 45 percent over the period 1968-1972 (U.S. Committee on Ways and Means, 1986). Moreover, since the Food Stamp benefits of non-AFDC female heads are not so reduced, the introduction of the Food Stamp program should, if anything, have increased the relative attractiveness of not being on AFDC. For all these reasons, the use of the introduction of food stamps as an explanation for the participation-rate increase must be heavily discounted.

Table 2 shows the trends in some of the demographic characteristics of the AFDC caseload and of female heads in general. None of the trends in these characteristics is likely to provide a satisfactory explanation for participation-rate trends. Over time, female heads have become slightly younger and have headed families with fewer children, both of which run counter to participation-rate trends. The percentage of the female-head population that is white has declined (the reverse is true of recipients, however), and if white women exhibit lower participation rates, this would tend to increase participation rates. But the magnitude of the trend is far too small to explain the explosive participation growth in the early periods in Table 1. Likewise, while the upward trend in education could potentially explain the decline in participation rates in the later years, its rate of increase was no smaller in the early period, when participation rates were rising.

Table 2

Demographic Characteristics of Female Heads

	1967	1969	1971	1973	1975	1977	1979	1981	1982
AFDC Recipients									
Age	33.1	32.7	32.1	31.3	30.4	29.9	30.0	30.5	29.5
Race[a]	.48	.48	.48	.44	.48	.52	.52	NA	NA
Education (grades completed)	9.1	9.4	9.6	9.1	10.2	10.4	10.4	NA	NA
No. children under 18	3.01	2.85	2.71	2.57	2.37	2.12	2.09	2.02	1.98
All Female Heads									
Age	37.8	37.5	37.1	36.1	35.3	35.9	35.9	35.5	34.4
Race[a]	.68	.68	.66	.63	.65	.64	.64	.65	.65
Education (grades completed)	10.3	10.3	10.6	10.7	10.9	11.0	11.2	11.4	11.6
No. children under 18	2.36	2.32	2.31	2.21	2.11	2.05	1.94	1.82	1.80

Source: Moffitt (1985).

[a]Proportion white.

Table 3 shows trends in the monthly income received by female heads in the United States. The trends for AFDC recipients are not of direct interest for present purposes because the income levels of recipients should naturally fluctuate as the break-even level of the program does so. We are interested instead in discerning whether alternative income sources in the entire female-head population might have fluctuated in a manner that could explain the trends in participation rates. The portion of the table showing trends for all female heads indicates that little explanation is to be found here. Earnings of female heads generally rose, then fell, over the period, opposite to the pattern that would explain participation rates. The earnings of others in the family did decline over the period, but the pace of the decline in the early years was insufficient to explain the large increase in participation rates. Likewise, while both real transfers from non-AFDC sources and real income from other non-transfer sources fell in the early years, they continued to fall in the later years and hence are not consistent with the full pattern of participation rates.

This brief review of the background trends in participation rates and in the possible determinants of those rates thus leaves us without any obvious explanation for the participation-rate trends. There are two major alternative explanations that are difficult to quantify but are probably important. First, a series of court decisions in the late 1960s and early 1970s considerably liberalized eligibility for the program by eliminating residency requirements, "man-in-the-house" rules, and other restrictions (Lurie, 1973; Michel, 1980). Second, although it is difficult to document formally, it is widely thought that there was a

Table 3

Sources of Monthly Income
(1967 dollars)

	1967	1968	1969	1970	1971	1972	1973	1974	1975	1976	1977	1978	1979	1980	1981	1982
AFDC Recipients																
Earnings of head	$ 22.6	—	$ 29.2	—	$ 31.0	—	$ 34.6	—	$ 32.9	—	$ 26.8	—	$ 24.7	—	$ 20.9	$ 5.8
Earnings of others	2.3	—	2.6	—	0.9	—	0.7	—	0.2	—	0.2	—	0.1	—	0.2	0.1
AFDC	160.3	—	171.5	—	151.3	—	146.0	—	137.7	—	133.9	—	115.7	—	107.8	107.8
Other transfers	5.5	—	5.1	—	4.7	—	4.1	—	3.9	—	2.4	—	1.9	—	1.9	1.7
Other income	19.2	—	15.8	—	14.8	—	8.5	—	5.6	—	2.1	—	1.1	—	0.6	0.4
Total income	$209.9	—	$224.2	—	$202.7	—	$143.9	—	$180.3	—	$165.4	—	$143.5	—	$131.4	$115.8
All Female Heads[a]																
Earnings of head	$152.9	$153.2	—	$165.4	—	$157.1	—	$164.7	—	$171.7	—	$192.3	—	$191.6	$176.4	—
Earnings of others	43.2	43.8	—	42.0	—	38.9	—	33.2	—	35.1	—	39.8	—	33.6	27.5	—
AFDC	32.0	36.9	—	47.4	—	54.0	—	50.1	—	48.2	—	41.7	—	33.9	31.8	—
Other transfers	38.4	37.1	—	37.6	—	36.6	—	38.0	—	40.5	—	38.3	—	35.3	29.4	—
Other income	43.9	40.8	—	38.0	—	35.5	—	39.3	—	38.7	—	36.8	—	36.6	35.2	—
Total income	$310.4	$311.8	—	$330.4	—	$322.1	—	$325.3	—	$334.2	—	$348.9	—	$331.0	$300.3	—

Source: Moffitt (1985).

[a]These income amounts are totals received in the year, divided by 12.

significant decrease in the stigma of welfare receipt over this same period, which would also lead to increased participation rates.

Although the impact of these two forces cannot be quantified, the lack of explanatory power of the measurable and quantifiable factors discussed above gives a strong measure of support to their possible importance in the time-series participation-rate trends. However, the review given above is, of course, extremely crude, and it would be desirable to conduct a more rigorous set of multivariate analyses to estimate more formally the importance of those possible explanators. This is necessary before the hypothesis of court-induced impacts and attitudinal changes can be accepted. The work reported in the later sections of this report is intended to provide one such analysis.

Past Economic Studies of Participation-Rate Trends. There have been a number of econometric studies of determinants of AFDC participation. Willis (1979) surveys many of the early studies by Bluestone, Holmer, Wiseman, and others. More recent studies include those of Willis (1980), Barr and Hall (1981), and Moffitt (1981, 1983). These later econometric studies were primarily concerned with the estimation of cross-section participation-rate equations and not with their potential for explaining time-series trends. There has also been interest more recently in the dynamics of AFDC turnover (Bane and Ellwood, 1983; Hutchens, 1981; O'Neill et al., 1984), but these studies are not directly germane to the issues under discussion here.

A fairly detailed examination of the determinants of time-series trends in the participation rate of eligibles from 1967 to 1977 was conducted by Michel (1980). He divided his period into the 1967-1972 and the 1973-1977 subperiods, and compared the growth rate of participation in the two subperiods with the average unemployment rates, wage rates, income levels, AFDC benefit levels, and other such variables in those subperiods. Michel's conclusion was quite similar to that reached above, which was that none of the variables seemed to provide a strong and consistent explanation for the trends in participation. Michel also provided a good discussion of the various court decisions made at that time, and speculated that they could have been a major contributor.

In a report to the U.S. Department of Health and Human Services prior to this one (Moffitt, 1985), the author made a preliminary attempt to determine if the structure of the AFDC participation equation has changed over time. Participation rates of all female heads were constructed for most states for several years over the period 1967-1982, using data from the biennial AFDC Characteristic Surveys (AS) and the March Current Population Surveys (CPS). The AS data provide counts of recipients in each state for each year in which the AS is available, and the CPS provides counts of the number of female heads in each state in those same years. Participation rates were calculated as the ratios of the two. AFDC guarantee levels and BRR values were taken from Fraker et al. (1985), although, because those variables were not available for all states in all AS years, the sample size was reduced somewhat. The CPS data were used to provide estimates of mean age, education, and other such variables for all female heads in each state in each relevant AS year. Participation equations were then estimated with ordinary least squares on these data. The grouped, state data were used rather than the underlying individual observations because of certain econometric difficulties that arise in the latter case; this is discussed in Section IV below.

Tables 4 and 5 show the results of the estimations.[4] The year-by-year estimates in Table 4 show consistently positive effects of the guarantee level on participation rates, but no obvious trend in the effects. However, the imprecision of the estimates--clearly a result of the small sample sizes and the grouping of the data by state--makes it difficult to draw any conclusions from the results. The effects of the benefit-reduction rate (BRR) are generally negative, as expected, but sometimes positive and almost never significant at conventional levels. The same applies to the coefficients in the rest of the table which, though often showing consistent signs across the years in some of the coefficients, also show almost no significance.

Table 5 shows the result of pooling all the years into one regression and including time dummies, thus constraining all parameters except the intercept to equality across time. Three different econometric techniques were applied--random effects, fixed effects, and between estimators--each of which treats the panel nature of the data (a time series of cross sections of state observations) differently. Details can be found in the report. Excepting the unusual results of the fixed-effects estimator, the table shows significantly positive effects of the AFDC guarantee and negative effects of the AFDC BRR. An increase in the guarantee of $100 in 1967 dollars increases the participation rate from 5 to 13 percentage points and a 25 percent increase in the BRR lowers the participation rate by from 6 to 12 percentage points. Age, education, the fraction white, the number of children, other income, and living in the South all have negative effects on participation, though generally not at statistically significant levels.

Table 4

Year-by-Year Estimates of AFDC Participation Equation
(standard errors in parentheses)

	1967	1969	1971	1973	1975	1977	1979	1981	1982
Guarantee[a]	0.083**	0.190	0.115	0.045	0.084*	0.091**	0.060	0.070	0.129*
	(0.039)	(0.165)	(0.073)	(0.050)	(0.051)	(0.030)	(0.049)	(0.045)	(0.083)
BRR[b]	−0.235	−0.568	−0.788	0.050	0.116	−0.377*	0.077	−0.300	0.543
	(0.229)	(0.800)	(0.484)	(0.240)	(0.421)	(0.219)	(0.205)	(0.219)	(0.342)
Age	−0.002	−0.051	0.020	−0.013	0.124	0.006	−0.013	0.018	—
	(0.017)	(0.082)	(0.040)	(0.016)	(0.536)	(0.013)	(0.017)	(0.017)	—
Education	−0.009	−0.096	0.047	−0.088	0.036	0.013	−0.163**	−0.007	−0.127
	(0.041)	(0.184)	(0.120)	(0.062)	(0.079)	(0.040)	(0.058)	(0.065)	(0.113)
Race[c]	−0.166	−0.044	0.185	−0.513**	−0.058	−0.124	−0.011	−0.173	—
	(0.170)	(0.825)	(0.483)	(0.213)	(0.335)	(0.136)	(0.191)	(0.214)	—
No. children	−0.001	0.054	0.034	0.052	0.127	0.037	−0.128	0.089	1.297
	(0.094)	(0.586)	(0.270)	(0.153)	(0.185)	(0.116)	(0.139)	(0.144)	(1.196)
Real hourly wage rate	−0.006	−0.013	0.007	0.015	0.034	−0.007	0.039	−0.024	—
	(0.028)	(0.135)	(0.097)	(0.035)	(0.043)	(0.028)	(0.045)	(0.066)	—
Real other income[a]	−0.065	−0.065	−0.138	−0.053	−0.111	0.001	0.143**	0.053	—
	(0.051)	(0.173)	(0.141)	(0.084)	(0.160)	(0.006)	(0.072)	(0.075)	—
South	−0.055	−0.091	0.046	−0.263**	−0.020	−0.099*	−0.134*	−0.179**	—
	(0.081)	(0.350)	(0.148)	(0.095)	(0.113)	(0.057)	(0.074)	(0.083)	—
Unemployment rate	0.037	0.095	0.018	−0.012	0.001	−0.015	0.010	0.016	—
	(0.027)	(0.113)	(0.044)	(0.023)	(0.027)	(0.014)	(0.026)	(0.012)	—
Constant	0.361	2.593	−1.168	2.089**	−0.908	0.117	2.500**	−0.322	−1.092
	(1.040)	(5.057)	(2.604)	(1.113)	(2.911)	(0.852)	(1.142)	(1.371)	(2.708)
Standard error	0.088	0.341	0.195	0.083	0.100	0.098	0.101	0.079	0.094
R-squared	0.552	0.461	0.395	0.785	0.742	0.584	0.636	0.771	0.618
No. of observations	29	26	27	22	20	46	37	29	9

Source: Moffitt (1985).
*Significant at the 10 percent level.
**Significant at the 5 percent level.

[a]Divided by 100.
[b]Benefit-reduction rate.
[c]Fraction white.

Table 5

Estimates of the Participation Equation
(standard errors in parentheses)

	Random Effects (1)	Fixed Effects (2)	Between (3)
Guarantee[a]	0.051**	0.001	0.139*
	(0.021)	(0.036)	(0.024)
BRR[b]	-0.263**	-0.189	-0.504**
	(0.100)	(0.115)	(0.175)
Age	-0.0046	-0.007	-0.002
	(0.0078)	(0.009)	(0.016)
Education	-0.0004	0.050	-0.079**
	(0.025)	(0.031)	(0.038)
Race[c]	-0.191*	-0.120	-0.032
	(0.103)	(0.186)	(0.124)
No. children	-0.0096	-0.031	0.117
	(0.062)	(0.070)	(0.170)
Real hourly wage rate	0.0126	-0.0025	0.004
	(0.016)	(0.018)	(0.031)
Real other income[a]	-0.0012	0.009	-0.055
	(0.031)	(0.036)	(0.048)
South	-0.120**	--[d]	-0.079
	(0.048)		(0.044)
Unemployment rate	0.0048	-0.0003	0.004
	(0.008)	(0.009)	(0.013)

Table Continues

Table 5, Continued

	Random Effects (1)	Fixed Effects (2)	Between (3)
Year Dummies[e]			
1969	0.104**	0.081**	0.073
	(0.039)	(0.038)	(0.063)
1971	0.133**	0.132**	-0.155
	(0.043)	(0.044)	(0.523)
1973	0.164**	0.141**	0.093
	(0.044)	(0.048)	(0.202)
1975	0.138**	0.104	0.300
	(0.058)	(0.066)	(0.220)
1977	0.152**	0.087	0.233
	(0.052)	(0.057)	(0.219)
1979	0.185**	0.083	0.498*
	(0.055)	(0.061)	(0.285)
1981	0.130*	0.013	0.166
	(0.068)	(0.077)	(0.260)
1982	0.120	-0.016	0.380
	(0.094)	(0.106)	(0.273)
Constant	0.488	--[b]	0.769
	(0.526)		(1.07)
Standard error	0.124	0.116	0.076
Rho	0.200	--	--
R-squared	0.327	0.216	0.783

[a]Divided by 100.
[b]Benefit-reduction rate.
[c]Fraction white.
[d]Coefficients on variables that are constant over time cannot be estimated.
[e]1967 omitted.
*Significant at the 10 percent level.
**Significant at the 5 percent level.

The coefficients on the year dummies are in some ways the most important, for they show that there is considerable variation in AFDC participation rates independent of the other variables included in the equation. This implies necessarily that those other variables are not explaining the entire time-series trend. Moreover, the coefficients on the time dummies roughly follow the quadratic pattern of participation rates shown in Table 1 above, indicating that the other variables in the equation are also not doing well at explaining that particular pattern. However, the significance levels of the time patterns are not always high, generating some doubt as to the robustness of this result. As with Table 4, it would be desirable to obtain more precise estimates of both the time-dummy coefficients as well as those on the other variables. If the same data set is to be used, this must involve employing the underlying individual data rather than the grouped state data. Use of the individual data would also permit separate year-by-year estimation, which is clearly of little or no reliability with the grouped data. The analysis reported in the later sections of this report will follow such an approach.

B. Participation and Work Incentives

There is a close connection between participation in the AFDC program and the level of earnings and labor supply of a female-headed family. Participation in the program necessitates low income, which implies on average low earnings and work effort. An important and long-standing issue of considerable policy importance concerns the possible effects of the AFDC benefit formula on the work incentives of female household heads. While both economic theory and the existing empirical evidence indicate that there is some work disincentive associated with participation in AFDC, it has been frequently hypothesized that a low BRR can minimize such disincentives by providing welfare recipients with an inducement to work. In the report previous to DHHS (Moffitt, 1985), the author demonstrated that, while the basic economic theory of labor supply does not support such a hypothesis, most of the available evidence on female heads does so. However, it was pointed out in that report that the time-series evidence in the late 1960s and early 1970s contradicts this finding.

Table 6 provides some evidence illustrating this issue. The table shows the hours of work and employment rates of AFDC recipients, which are, as should be expected, extremely low. Both rose in the early years but have fallen more recently. However, these figures are of little interest for the issue at hand because their patterns may simply be a result of the quadratic pattern in the AFDC break-even level shown in Table 1. An increase in the break-even level should raise the earnings levels, hours of work, and employment rates of recipients, and a decrease in the break-even level should do the opposite, even if there are no work disincentives of AFDC (i.e., even if not a single woman changes her work effort in going on or off the program).

The more relevant trends are those for the hours of work and employment rates of all female heads, which should respond negatively to the guarantee and to the BRR according to the work-disincentive hypothesis. The table shows that the trends in these two labor supply measures are exactly the opposite of those for recipients, falling in the early

Table 6

Trends in Labor Supply of Female Heads

	1967	1969	1971	1973	1975	1977	1979	1981	1982
AFDC Recipients[a]									
Hours of work per week[b]	5.0	5.0	5.3	5.8	6.1	4.9	4.9	4.3	1.7
Employment rate (%)	16.0	15.0	17.0	18.0	18.0	15.0	16.0	14.0	7.0
All Female Heads[a]									
Hours of work per week[b]	18.8	18.6	17.6	17.7	17.2	18.2	20.6	20.4	19.3
Employment rate (%)	52.0	51.0	49.0	49.0	48.0	56.0	56.0	55.0	53.0
Other Variables[c]									
Unemployment rate (national)	3.8	3.5	5.9	4.9	8.5	7.0	5.8	7.6	9.7
Guarantee (monthly 1967 $)	161	156	153	149	147	141	131	113	111
Benefit-reduction rate (%)	41	42	23	22	30	33	32	24	70
Break-even level ($)	393	371	665	677	490	427	409	471	158

[a]From Moffitt (1985).
[b]Includes zeros for nonworkers.
[c]From Table 1.

period and rising in the latter. This pattern is in flat contradiction to the notion that the BRR is negatively related to labor supply. Also, although the declining guarantee level could at least partially explain the increasing labor supply levels in the latter period, it could not in the early period. In addition, as Table 1 indicated, potential wages rose over the late 1960s and early 1970s, which could not have generated the hours decline. Thus the most anomalous aspect of the table is its demonstration that the labor supply of female heads fell, rather than rose, after the implementation of the 30-and-one-third deductions of the 1967 Social Security Amendments.

A possible explanation for this puzzling piece of evidence is provided by the trends in participation discussed previously and shown in Table 1. As stressed there, the simple review of the evidence indicates that participation rates in the program may have shifted upward in the late 1960s and early 1970s for reasons not explainable by any measure of economic or demographic change. If this is true and if there was indeed a sharp increase in the participation rate for exogenous reasons, this could explain the sharp drop in labor supply of female heads shown in Table 6, for participation in AFDC always is associated with labor-supply reductions. According to this hypothesis, the reduction in the BRR that occurred over the same period was only coincidental.

Table 7 shows the results of a few time-series regressions on the nine observations in Table 6, with hours of work as the dependent variable. As shown in columns (1) and (2)--equations with and without a time trend--the time-series data imply that the BRR is positively, not negatively, correlated with hours of work (holding all other variables

Table 7

Time-Series Regressions (1967-1982)
(H = hours of work; standard
errors in parentheses)

	H (1)	H (2)	ΔH (3)	H (4)	ΔH (5)
BRR	0.76 (2.05)	0.46 (1.66)	--	-2.12 (2.20)	--
Guarantee	-0.07 (0.05)	-0.09 (0.19)	--	0.09 (0.10)	--
Year	0.06 (0.20)	--	--	0.07 (0.15)	--
Unemployment rate	-0.59* (0.22)	-0.55* (0.17)	--	-0.32 (0.22)	--
ΔBRR	--	--	-0.18 (2.04)	--	-2.87 (1.89)
ΔGuarantee	--	--	-0.05 (0.71)	--	0.04 (0.06)
ΔUnemployment rate	--	--	-0.41* (0.19)	--	-0.47* (0.14)
D75[a]	--	--	--	--	1.55* (0.70)
D75*(Year-75)	--	--	--	0.99* (0.52)	--
Intercept	27.8	34.5	0.04	1.31	-.23
R-squared	.82	.82	.61	.92	.85

Source: Moffitt (1985).
[a]D75 = 1 if YEAR $\geq$ 75, 0 if not.
*Significant at 10 percent level.

constant). The first-difference regression in column (3) shows a negative, but extremely small, effect of BRR on hours of work. However, in columns (4) and (5), the time-trend in the equations is allowed to differ before and after 1975, the year at which the participation rate in AFDC was near its peak. With the two periods thus separated in the regressions, the BRR coefficient becomes negative. Moreover, the magnitude of the coefficient is very close to that obtained in the other analyses in Moffitt (1985). Of course, the evidence in Table 7 is extremely weak by its nature. Therefore, in Section V of this report, individual data are used to test this hypothesis more rigorously.

III. DATA BASES AND VARIABLE CONSTRUCTION

The two primary data sources used in this study are the biennial AFDC Characteristics Survey (AS) and the annual March Current Population Survey (CPS). The AS data are obtained from samples of the AFDC caseload in each state, and are available from 1967 to 1979 at two-year intervals. In addition to data on benefits received, demographic characteristics of the head of the household are available. The sample sizes range considerably from year to year, but are no smaller than 16,000 (1969) and are as large as 67,000 (1967). The March CPS is a nationally representative survey of the U.S. population which gathers economic and socio-demographic information on the respondents in the week of the survey, and income data for the prior year from retrospective questions. Approximately 3000 female heads are available in the CPS in each year.

These two data sets provide information on AFDC recipients and on the total female-head population, respectively. Fortunately, the AS data were generally collected in the spring of each year, not far from the March CPS survey date. Consequently, by matching a March CPS with each AS, a point-in-time estimate of the participation rate can be calculated simply by dividing the weighted sum of AS observations by the weighted sum of female heads in the CPS. Note that AFDC status as of the survey week is not available in the CPS.[5]

There are two minor problems with this matching procedure. First, the 1967 AS was administered in November of 1967, closer to the March 1968 CPS than the March 1967 CPS. Therefore the March 1968 CPS is used with the 1967 AS. Second, for the study of labor supply, the AS has the disadvantage of providing information only on part-time and full-time status (defined as hours of work below and above 35 per week, respectively), unlike the CPS, which provides information on hours of work in the survey week. However, this difference can be incorporated into the estimation procedure, as discussed in Section IV.

Another disadvantage of the CPS for the study of AFDC is that the measurement of subfamilies is in error prior to 1982 (Ellwood and Bane, 1985). According to their estimates, about 10 percent of the children in subfamilies were incorrectly assigned to the parent family, resulting in an undercount of subfamilies in the CPS of unknown magnitude (but less than 10 percent). This problem is ignored here. It should not seriously affect the results, for the focus of the project is on the 1967-1973 change, a period when the coding procedures were not altered.[6] Results of forthcoming research on the miscoding problem should help determine its seriousness.

The CPS and AS data sets are reduced in several ways before obtaining an analysis sample. First, and most important, to reduce the analysis burden on the project to a reasonable magnitude, only the years 1967, 1973, and 1979 are examined. These years are chosen to bracket the major trends in AFDC: the participation rate peaked in 1973, so the 1967-1973 period is the one most in need of explanation, while the 1973-1979 period saw the slackening of the growth of participation. Second, because of the inordinately large size of the AS surveys, they are subsampled to bring their sizes down to approximately 3000, the size of the CPS female-head samples. In addition, when the analysis was begun, it was found that even these sample sizes (6000 per year for three years, or 18,000 observations) were too large for feasible estimation on a mainframe IBM computer; consequently, a further one-third subsample is drawn. Finally, a few exclusions are made for outliers and missing values, including unavailable AFDC guarantees (see below). The final sample size is 5225.

A few other variables are obtained from other sources. AFDC guarantee and benefit-reduction rates by state for the three years are obtained from Fraker et al. (1985), who estimated effective guarantees and benefit-reduction rates from the same AS data employed in this project. However, because sample sizes were too small in some states for estimation, no estimates are available for some of the smaller states in the nation. Observations in these states are deleted from the sample, as mentioned in the last paragraph. The unemployment rate for each state in each year is obtained from the U.S. Department of Labor. Finally, regional CPI indicators are used with a base year of 1977 to deflate the monetary variables.

The means of the variables are shown in Table 8. The fraction of CPS female heads who are eligible for AFDC increased from 1967 to 1973, then decreased. Interestingly, this follows the same pattern as participation rates, suggesting that they may both be affected by common variables. Hours of work of female heads in the sample dropped from 1967 to 1973 and then rose, following the same pattern as that discussed previously. The full-time and part-time work trends among AFDC recipients also followed the previously discussed pattern. The other variables in the table are self-explanatory. Note that no hourly wage is included in the model; such a variable would have to be imputed for nonworkers.[7]

Table 9 shows the estimated participation rates in the analysis sample. Participation rates in the total female-head population rose

Table 8

Means of the Variables Used in the Analysis of Female Household Heads

	1967		1973		1979	
	CPS	AS	CPS	AS	CPS	AS
Fraction eligible for AFDC	0.84	1.00	0.89	1.00	0.80	1.00
Hours of work per week	20.03	-	17.76	-	20.65	-
Proportion employed full time	-	0.09	-	0.13	-	0.11
Proportion employed part time	-	0.09	-	0.07	-	0.08
Monthly guarantee[a]	3.14	3.40	3.03	3.20	2.63	2.81
Benefit-reduction rate	0.31	0.32	0.21	1.20	0.30	0.30
Monthly nonwage income[b]	0.87	0.34	1.03	0.19	0.51	0.05
Age/10	3.62	3.20	3.53	3.08	3.44	3.01
Education	10.51	9.26	10.70	9.11	11.22	10.52
Race (1=white)	0.67	0.48	0.63	0.50	0.64	0.52
No. children < 18	2.41	2.94	2.17	2.79	1.94	2.11
Unemployment rate	3.64	3.76	4.95	5.11	5.92	6.04
South dummy	0.35	0.25	0.36	0.28	0.36	0.28
No. observations/100	7.89	7.69	9.21	8.73	10.92	7.81

[a]For a family of four in 1977 dollars, divided by 100.

[b]Sum of nontransfer nonwage income and earnings of others in 1977 dollars, divided by 100.

Table 9

AFDC Participation Rates in the Sample

	1967	1973	1979
Total female-head population	0.26	0.49	0.46
Eligible female-head population	0.31	0.55	0.58
Michel (1980)	0.45	0.91	0.97

from 26 percent in 1967 to 49 percent in 1973, then fell back to 46 percent in 1979. Among eligibles, the participation rate rose from 31 percent in 1967 to 55 percent in 1973, and then rose a bit more to 58 percent in 1979. These estimates are considerably below those of Michel (1980), also shown in the table. The major differences between the two estimates lie in the estimates of the number of eligible female heads in the CPS, which Michel used as well for this purpose. The differences are that (1) Michel employed an assets test to screen for eligibles, whereas here a female-headed family in the CPS is judged to be eligible merely if the sum of the prior year's family earnings and nonearned income is less than the AFDC break-even level (guarantee divided by BRR); (2) Michel assumed that the 30-and-one-third disregard did not apply for determining eligibility, whereas here not only are they assumed to apply, but the effective BRR values used implicitly assume all deductions are applied for eligibility; and (3) Michel reduced his estimate of eligibles by a part-year, split-sample treatment of the prior year's earnings, whereas no such procedure is followed here. Thus a much broader definition of eligibility is used in this report. For the purposes of the behavioral models estimated below, the participation rate of the total female-head population is the more relevant.[8]

IV. ESTIMATING TECHNIQUES

The object of the estimation in the report is the determination of whether there was a structural shift in AFDC participation over the period 1967 to 1979. To examine this question, three separate participation equations will be estimated, one in each year. Tests for the significance of the difference in the coefficients across years will then be performed.

The main estimation issue concerns the manner in which the AS and CPS samples can be combined to estimate participation equations. If a single sample of recipients and nonrecipients were available, it would be straightforward to apply probit or logit to the estimation of the equation. However, in the present case the CPS sample contains some fraction of recipients who cannot be identified. Thus there is an overlap in the sampling frames of the two data sets. Nevertheless, intuitively one can see that the effects of a set of independent variables on participation should be estimable with these data. The aggregate participation rate can be estimated simply by dividing the sample size of the AS data by that of the CPS data (appropriately weighted). But a similar procedure could be followed within any arbitrary stratum of the population defined by particular values of all the independent variables. If the participation rate is estimable within all strata of independent variables, then the effect of changing values of those variables on participation is also identified.

Consistent estimation of binary-choice coefficients in this circumstance has been discussed in the literature on choice-based sampling (Manski and Lerman, 1977; Cosslett, 1981). The closest case to that here is discussed by Cosslett (p. 91). Assume first that the probability of participating in AFDC by individual i is $P_i = F(X_i^{'}\beta)$, where F is the normal cumulative distribution function, X_i is a vector of observations on the independent variables for individual i, and β is the associated vector of coefficients. If there are N_1 observations in the AS sample and N_2 observations in the CPS sample, then Cosslett shows that consistent estimates of β can be obtained by maximizing the following log likelihood function w.r.t. β and the parameter λ:

$$L = \sum_{i=1}^{N_1} \log\ [\lambda P_i/(\lambda P_i + H_0)] + \sum_{i=1}^{N_2} \log\ [1/(\lambda P_i + H_0)]. \tag{1}$$

Here H_0 is the ratio of the number of observations in the CPS sample to the number of observations in the AS sample. In Cosslett's case, the aggregate participation rate is unknown, and it is the parameter λ that is the estimate of its inverse. However, in the present case the aggregate participation rate is known, and can be substituted into the likelihood function. But since H_0 equals the inverse of the aggregate participation rate as well (when the ratios of weighted counts are used), both λ and H_0 disappear from the likelihood function, leaving us with the following:

$$L = \sum_{i=1}^{N_1} \log\ [P_i/(1 + P_i)] + \sum_{i=1}^{N_2} \log\ [1/(1 + P_i)]. \tag{2}$$

This formulation of the likelihood function has considerable intuitive appeal. Each of the probabilities represents the conditional probability of observing one type of observation in the total, combined sample. The probability of being in the AS sample is P_i and the probability of being in the CPS sample is 1 (ignoring sampling weights for a moment), so the conditional probabilities for the two samples are simply P/(1+P) and 1/(1+P) for the AS and CPS, respectively. Also, it is easy to show that maximization of L in (2) w.r.t. to the aggregate participation rate ($P=P_i$) yields a solution of $P=N_1/N_2$, as should be the case.

Adding relative sampling weights (w_i) to the likelihood function, we have

$$L = \sum_{i=1}^{N_1} w_i \log [P_i/(1 + P_i)] + \sum_{i=1}^{N_2} w_i \log [1/(1 + P_i)], \tag{3}$$

which, when maximized w.r.t. β, yields consistent estimates of the parameters. The maximization algorithm of Berndt et al. (1974) is used for the estimation.

Endogeneity of Eligibility. Maximization of (3) can be employed to obtain probit estimates of either the total participation equation or of the eligible participation equation. In the latter case, ineligibles are simply deleted from the sample. However, this deletion raises the problem of selectivity bias, for if there are any unobserved determinants of eligibility which are correlated with unobserved determinants of participation conditional upon eligibility, the probit estimates of the eligibles-only participation equation will be inconsistent. This selectivity bias is intrinsically a problem of sample composition, for if a change in a right-hand-side variable brings into (or takes out of) the eligible sample a group with systematically different participation propensities than those initially in the sample, the participation rate will change simply because the sample composition changes.

If a single data set with observations on AFDC participation and eligibility were available, a sample-selection model of the following type could be estimated:[9]

$$P_i^* = X_i'\beta + \varepsilon_i,$$
$$P_i = 1 \text{ if } P_i^* \geq 0;\ P_i = 0 \text{ if not, and} \tag{4}$$

$$E_i^* = Z_i'\delta + \nu_i,$$
$$E_i = 1 \text{ if } E_i^* \geq 0;\ E_i = 0 \text{ if not,} \tag{5}$$

where P_i is an AFDC participation dummy variable, E_i is a dummy equal to one if eligible and zero if not, P_i^* and E_i^* are their latent variables, X_i and Z_i are vectors of observed covariates for the participation and eligibility equations (respectively), and where ε_i and ν_i are assumed to be error terms distributed bivariate normal with correlation ρ. The difficulty in estimation is that P_i^* is not observed for ineligibles, so one "cell" of the design is missing.[10] The parameters of (4) and (5) can be estimated with maximum likelihood subject to the identification restriction that there be one variable in X_i which is not in Z_i.

The direction of bias in estimates of the participation equation alone can be seen as follows. In the eligible-only sample, the expected value of P_i^* is

$$E(P_i^* | E_i = 1) = X_i'\beta + E(\varepsilon_i | \nu_i \geq -Z_i'\delta)$$
$$= X_i'\beta + \rho\lambda_i, \tag{6}$$

where

$$\lambda_i = \frac{f(-Z_i'\delta)}{1-F(-Z_i'\delta)} , \tag{7}$$

and where f is the unit normal density function. Taking the partial derivative of (6) w.r.t. the j^{th} variable (assumed to be in Z_i as well), one obtains

$$\frac{\partial E(P_i^* | E_i=1)}{\partial X_{ij}} = \beta_j + \rho \frac{\partial \lambda_i}{\partial X_{ij}} . \tag{8}$$

Since it can be shown that $\partial\lambda_i/\partial X_{ij}$ takes on the sign of $-\delta_j$, the bias in the coefficient β_j takes on the sign of $-\rho\delta_j$.[11]

An example of such bias should illustrate the problem. Suppose that the right-hand-side variable of interest is the guarantee level. Assume as well that $\rho > 0$, for it must surely be the case that eligibles have higher propensities to participate in AFDC than ineligibles. Then an increase in the guarantee may raise the participation rate of those initially eligible, but will simultaneously raise the break-even level and bring into the eligible population a relatively high-income group whose participation propensities are lower than those of the initially eligible group. This will lower the participation rate in the new population of eligibles, biasing the guarantee effect downward.

Estimation of the parameters of the model in (4)-(5) with the overlapping AS-CPS data set requires once again conditioning on the probability of being in the sample $(1+P_i)$. Applying this principle leads to the log likelihood function:

$$L = \sum_{AS} w_i \log [Q_i/(1 + Q_i)] + \sum_{\substack{CPS \\ \text{Eligibles}}} w_i \log [R_i/(1 + Q_i)]$$
$$+ \sum_{\substack{CPS \\ \text{Ineligibles}}} w_i \log [(1 - R_i)/(1 + Q_i)], \tag{9}$$

where

$$\begin{aligned} Q_i &= \text{Prob}(P_i^* \geq 0,\ E_i^* \geq 0) \\ &= \text{Prob}(\varepsilon_i \geq -X_i'\beta,\ \nu_i \geq -Z_i'\delta) \\ &= B(X_i'\beta,\ Z_i\delta;\rho), \text{ and} \end{aligned} \tag{10}$$

$$\begin{aligned} R_i &= \text{Prob}(E_i^* \geq 0), \\ &= \text{Prob}(\nu_i \geq -Z_i'\delta), \\ &= 1 - F(-Z_i'\delta), \end{aligned} \tag{11}$$

where B is the bivariate normal cumulative distribution function with correlation ρ: $B(a, b;\rho) = \text{Prob}(\varepsilon_i < a, \nu_i < b)$. The reversal of the inequalities follows from the symmetry of the normal distribution.

Incorporating Labor Supply. As discussed in Moffitt (1985), the standard structural model of labor supply and AFDC participation posits that hours of work are a function of the AFDC guarantee and BRR actually faced by the individual--that is, hours of work are a function of AFDC participation. The probability of participation is in turn a function of the guarantee and BRR as well as of the individual's tastes for work, which appears in the error term of the hours-worked relationship. Rather than estimate this structural model, only a reduced-form model will be considered here. In the reduced-form model, hours of work are a function of the AFDC guarantee and BRR and other variables, AFDC participation is a function of the same variables, and the error terms in the two equations are correlated. In such a model, the coefficients in the labor supply equation combine the effects of the independent variables on participation as well as on hours of work conditional upon participation.

There are three econometric complications in estimating the model on the present data set. First, there is once again the problem of overlapping samples, but this can be addressed by appropriate conditioning of the probabilities. Second, there is the problem that many female heads in the CPS and AS do not work (most in the AS do not--see Table 6 in Section II). This problem can be addressed by application of the Tobit technique. Third, there is the problem that only part-time and full-time work status is observed for the AS observations. This can be addressed by the application of ordered probit, in which it is assumed that the part-time and full-time observations represent groupings of the underlying hours equation.

The model to be estimated is the following:

$$P_i^* = X_i'\beta + \varepsilon_i, \tag{12}$$

$$H_i^* = Z_i'\delta + \nu_i, \tag{13}$$

$$H_i = 0 \text{ if } H_i^* < 1, \tag{14}$$

$$H_i = H_i^* \text{ if } H_i^* \geq 1 \text{ and CPS}, \tag{15}$$

$$H_i = \text{part-time if } 35 \geq H_i^* \geq 1 \text{ and AS}, \tag{16}$$

$$H_i = \text{full-time if } H_i^* > 35 \text{ and AS}. \tag{17}$$

A cutoff point of one hour of work instead of zero is used to accord with the AS questions and to maintain consistency in the work probability between the AS and the CPS. Assuming that ε_i and ν_i are distributed bivariate normal with correlation ρ, and applying the necessary conditioning for the overlap in the sampling frames, the log likelihood function becomes

$$L = \sum_{\substack{AS \\ H_i=0}} w_i \log [Q_{1i}/(1 + P_i)] + \sum_{\substack{AS \\ \text{Part-time}}} w_i \log [Q_{2i}/(1 + P_i)]$$

$$+ \sum_{\substack{AS \\ \text{Full-time}}} w_i \log [Q_{3i}/(1 + P_i)] + \sum_{\substack{CPS \\ H_i=0}} w_i \log [Q_{4i}/(1 + P_i)]$$

$$+ \sum_{\substack{CPS \\ H_i>0}} w_i \log [Q_{5i}/(1 + P_i)], \tag{18}$$

where

$$P_i = \text{Prob}(P_i^* \geq 0) = 1 - F(-X_i'\beta),$$

$$Q_{1i} = B(X_i'\beta,\ z_{1i};\ -\rho),$$

$$Q_{2i} = B(X_i'\beta,\ z_{35i};\ -\rho) - B(X_i'\beta,\ z_{1i};\ -\rho),$$

$$Q_{3i} = B(X_i'\beta,\ z_{35i};\ \rho),$$

$$Q_{4i} = F(z_{1i}),$$

$$Q_{5i} = f(z_i)/\sigma_\nu,$$

$$z_{1i} = (1 - Z_i'\delta)/\sigma_\nu,$$

$$z_{35i} = (35 - Z_i'\delta)/\sigma_\nu,$$

$$z_i = (H_i - Z_i'\delta)/\sigma_\nu,$$

and where σ_ν is the standard deviation of ν_i.

V. ESTIMATION RESULTS

A. Participation Equations

Total Female-Head Population. Table 10 shows the results of the estimation of the probit participation equation on the total female-head population separately for the three years. As the table indicates, the guarantee coefficients fall into a fairly narrow range across the three years, ranging from .25 to .29. These imply that a $100 increase in the guarantee would increase the probit index by about .27, which, at the means of the data, implies an increase in the probability of participating in AFDC of about .10. The three guarantee coefficients are also all statistically significant at the 10 percent level.

The coefficients on the benefit-reduction rate are also fairly stable over time, ranging from -1.00 to -1.11. However, none is significant at conventional levels. But the point estimates are negative in sign, implying that there is some response in participation to changes in the BRR, and that policy changes in the BRR can be expected to alter the size of the caseload. At the means of the variables, the coefficient estimates imply that a 25 percent increase in the BRR would lower the participation rate (and therefore the caseload) by about .08. It is also interesting to note that these guarantee and BRR effects are quite similar to those obtained in the grouped-state data of Moffitt (1985), discussed above in Section II in connection with Table 5. Thus the grouping in that report does not appear to have seriously affected the quality of the estimates.

The other coefficients in Table 10 show that the participation rate is significantly affected by many of the other variables in the analysis.

Table 10

Estimates of Total-Participation-Base Equations
(standard errors in parentheses)

	1967	1973	1979
Guarantee/100	0.29* (0.12)	0.29* (0.15)	0.25* (0.15)
BRR	-1.05 (0.73)	-1.11 (1.09)	-1.00 (0.83)
Nonwage income/100	-0.22* (0.05)	-0.57* (0.05)	-2.76* (0.14)
Age/10	-0.34* (0.06)	-0.38* (0.07)	-0.43* (0.07)
Education	-0.19* (0.03)	-1.25* (0.09)	-0.62* (0.05)
Race (1=white)	-0.58* (0.13)	-0.42* (0.15)	-0.10 (0.16)
No. children	0.15* (0.04)	0.17* (0.06)	0.11 (0.07)
Unemployment rate	0.17* (0.08)	0.03 (0.07)	-0.10 (0.08)
South dummy	-0.39* (0.21)	0.04 (0.26)	-0.42 (0.30)
Intercept	1.44* (0.59)	13.85* (1.24)	9.08* (1.11)
Log LF	-1057	-1255	-1436

*Significant at 10 percent level.

Higher amounts of available nonwage income lower the rate of participation as should be expected. Participation rates also fall significantly with age, with greater levels of education, and for white female heads relative to nonwhite. The number of children has a positive effect on the participation rate as well. The unemployment rate has a positive effect on participation in 1967 but has an insignificant effect in 1973 and 1979. This is an interesting finding, for it has been suggested elsewhere (principally by Senator Daniel Moynihan) that the AFDC caseload in the 1970s lost its procyclical behavior. This evidence supports that hypothesis.

Another fairly interesting result discernible from the table is that the general level of significance of the variables included in the equations falls over time. This is not a result of falling sample sizes (they increase, in fact), but instead seems to signal a weakening of the relationship between many demographic characteristics and the probability of participation. This weakening may arise from the increase in the participation rate itself: as the rate increases toward 100 percent, the small amount of nonparticipation may be due to increasingly random factors.

With regard to the question of whether there was a structural change in the equation from 1967 to 1973, the statistical tests unequivocally show that there was indeed such a change. The chi-squared value under the null of no structural change (i.e., the same intercept and coefficients in 1967 and 1973) is 327, far in excess of the critical value for 10 degrees of freedom (23.2). Thus a stable relationship is overwhelmingly rejected. In quantitative terms, it is clear from the results that most of this change is centered in the intercept term, for the guarantee and BRR coefficients, for example, are fairly stable. Consequently, there appears to have been a significant, unexplained upward shift in participation propensities between 1967 and 1973 that is unaccounted for by any of the independent variables in the equation.

The negative signs on the BRR variables indicates that the reduction in the BRR generated by the 1967 Social Security Amendments made some contribution to the increased participation rates over the period. However, its contribution is slight. As noted previously, the participation rate in the total female-head population rose from 26 percent to 49 percent between 1967 and 1973. Using a BRR coefficient of -1.0, it can be shown that the reduction in the BRR alone would have raised the participation rate only to 31 percent. Thus the fall in the BRR can explain only about one-fifth of the increase in the participation rate.

As noted previously, the real guarantee did increase by 8 percent in the middle 1960s, possibly generating a lagged increase in the participation rate. But, once again, the magnitude of the relevant coefficient is insufficient to explain any sizable portion of the participation-rate increase from this source. Since the 8 percent increase corresponds to an increase in the real guarantee of from $25 to $32 (depending upon the base year used), and since a $100 increase in the guarantee appears to increase the participation rate by .10, a lagged response would have increased participation only by 2.5 to 3.2 percentage points. This effect is dwarfed by the 23-percentage-point increase in the actual participation rate. Of course, the estimated coefficient is based upon the responsiveness of participation to the current, not the lagged, guarantee. But to explain the 23-point increase would require the coefficient on a lagged benefit variable to be about 10 times the size of the coefficient on the current benefit variable, an implausibly large difference.

It should also be noted that food stamps have been omitted from the estimating equations and their influence ignored. Since the food stamp program is national in scope, its benefit levels are constant across the states and, hence, it contains no cross-sectional variation and cannot be used in the equations. However, presuming the AFDC guarantee effect in the equation to also apply to food stamps, the 45 percent increase in the benefit from AFDC and food stamps combined from 1967 to 1971 (see previous discussion) would have generated an increase in the participation rate of .07, almost one-third of the increase from 1967 to 1973. However, for the reasons given earlier--not the least of which is that the effective benefit increase was 25 percent, not 45--this estimate must be regarded as a serious overestimate of the effect of food stamps.

<u>Eligible Female-Head Population</u>. Table 11 shows the results of deleting the ineligible CPS observations from the sample and estimating the probit participation equation on eligibles alone. As the table indicates, the guarantee coefficient in 1967 is slightly below that in Table 10 but the guarantee coefficients in 1973 and 1979 are about the same. The coefficients on all the other variables except the BRR are also fairly close to those in Table 10. However, the BRR coefficients are much different in 1967 and 1979 than previously, and in 1979 the coefficient is even positive. This result would seem to be implausible, and the first hypothesis to explore for this anomalous finding is the selectivity-bias hypothesis discussed in the last section.

Table 11

Estimates of Eligible-Participation-Rate Equation
(standard errors in parentheses)

	1967	1973	1979
Guarantee/100	0.20*	0.27*	0.27*
	(0.11)	(0.15)	(0.16)
BRR	-0.28	-1.39	0.58
	(0.70)	(1.08)	(0.86)
Nonwage income/100	-0.27*	-0.49*	-1.06*
	(0.07)	(0.06)	(0.07)
Age/10	-0.38*	-0.37*	-0.45*
	(0.06)	(0.08)	(0.08)
Education	-0.15*	-1.36*	-0.59*
	(0.02)	(0.10)	(0.05)
Race (1=white)	-0.60*	-0.47*	-0.44*
	(0.13)	(0.16)	(0.16)
No. children	0.17*	-0.01*	-0.05
	(0.07)	(0.11)	(0.09)
Unemployment rate	0.15*	0.01	-0.08
	(0.08)	(0.07)	(0.08)
South dummy	-0.47*	0.06	-0.23
	(0.20)	(0.25)	(0.32)
Intercept	1.59*	15.87*	8.82*
	(0.53)	(1.40)	(1.03)
Log LF	-1045	-1277	-1453

*Significant at 10 percent level.

Table 12 shows the results of estimating probit eligibility equations on the CPS sample alone. The results show clearly that the probability of being eligible is strongly affected by most of the same variables that are in the participation equation in Table 11. Of special interest is the fact that the BRR has a strong and negative effect on the probability of being eligible. This is consistent with the results of Table 11, for, as hypothesized in the previous section, the selectivity bias involved should bias the BRR coefficient in a positive direction (since its coefficient in the eligibility equation is negative), assuming that p is positive.

When attempts were made to estimate the joint participation-eligibility model discussed in the previous section, no stable estimates could be obtained. Specifically, the correlation parameter p was estimated to be so close to 1.0 that the two equations, in effect, could not be distinguished. However, to illustrate the importance of the correlation, the model was estimated by fixing the value of p at a high value, namely, .95. Although this is not the correct estimate of the model, it demonstrates the effect of allowing a nonzero correlation on the estimates of the participation equation.

Table 13 shows the results. As the table shows, when the sample selectivity of eligibility is incorporated into the model, the BRR coefficients in the participation equations in 1967 and 1979 are once more negative. In 1967 its value is -.81, a bit below that in Table 10, and in 1979 it is -1.14, a bit above, but these values are sufficiently close that the difference can be ascribed to sampling error. The guarantee effects in Table 13 are somewhat stronger than have been obtained

Table 12

Estimates of Probit Eligibility Equations on
CPS Sample
(standard errors in parentheses)

	1967	1973	1979
Guarantee/100	0.79*	0.87*	0.52*
	(0.15)	(0.22)	(0.10)
BRR	-6.04*	-10.38*	-4.68*
	(0.86)	(1.58)	(0.42)
Nonwage income/100	-0.21*	-0.23*	-0.38*
	(0.03)	(0.02)	(0.04)
Age/10	-0.09*	-0.05	-0.11*
	(0.05)	(0.07)	(0.04)
Education	-0.16	-0.22	-0.15*
	(0.02)	(0.02)	(0.01)
Race (1=white)	-0.46*	-0.19	-0.21*
	(0.15)	(0.16)	(0.09)
No. children	0.18*	-0.15	0.30*
	(0.09)	(0.15)	(0.07)
Unemployment rate	-0.02	-0.02	0.01
	(0.08)	(0.08)	(0.04)
South dummy	0.13	-0.21	-0.36*
	(0.20)	(0.24)	(0.16)
Intercept	3.51*	4.98*	3.31*
	(0.55)	(0.64)	(0.44)
Log LF	-355	-268	-591

*Significant at 10 percent level.

Table 13

Estimates of Joint Participation-Eligibility Model
(ρ = 0.95; standard errors in parentheses)

	1967		1973		1979	
	Participation	Eligibility	Participation	Eligibility	Participation	Eligibility
Guarantee/100	0.46* (0.07)	0.78* (0.10)	0.32* (0.06)	0.83* (0.18)	0.48* (0.09)	0.52* (0.08)
BRR	−0.81 (0.65)	−6.02* (0.55)	−1.79 (1.11)	−9.74* (1.33)	−1.14* (0.64)	−4.61* (0.35)
Nonwage income/100	−0.27* 0.067	−0.21* (0.02)	−0.54 (0.81)	−0.23* (0.02)	−1.16* (0.08)	−0.37* (0.03)
Age/10	−0.32* (0.06)	−0.09* (0.03)	−0.36* (0.08)	−0.08 (0.06)	−0.42* (0.07)	−0.14* (0.04)
Education	−0.17 (0.26)	−0.16* (0.01)	−0.96* (0.06)	−0.29* (0.03)	−0.51* (0.05)	−0.15* (0.01)
Race (1=white)	−0.59 (0.13)	−0.47* (0.10)	−0.34 (0.13)	−0.49* (0.13)	−0.44* (0.14)	−0.25* (0.07)
No. children	—	0.18* (0.06)	—	0.23* (0.13)	—	0.29* (0.05)
Unemployment rate	—	−0.02 (0.05)	—	0.03 (0.07)	—	−0.01 (0.03)
South dummy	—	0.13 (0.13)	—	−0.22 (0.19)	—	−0.39* (0.14)
Intercept	1.66* (0.41)	3.52* (0.35)	11.41* (0.85)	5.68* (0.51)	6.91* (0.67)	3.46* (0.37)
Log LF	−1562		−1631		−2195	

*Significant at 10 percent level.

previously, possibly suggesting that guarantee effects are stronger for eligibles than for ineligibles.

For present purposes, it seems clear from Table 11 and Table 13 that a structural shift occurred in participation, given eligibility, as well. The constant terms in the participation equations invariably shift upward dramatically between 1967 and 1973, and tests for structural change once again reject the null hypothesis of no change between the years. Thus the conclusion reached previously remains unchanged.

B. Participation-Labor-Supply Model

Table 14 shows the results of estimating two versions of the participation-labor-supply model. In both cases all three years are pooled into a single equation and the years are dummied out. Thus the coefficients on any other variables in the equation are constrained to equality. In the first model estimation, only dummies for the time periods are included. The results show that, as expected, participation propensities shifted upward from 1967 to 1973 and then came back down a bit in 1979, while hours of work shifted downward from 1967 to 1973, then came back up considerably in 1979. The estimate of the cross-equation correlation coefficient is -0.98, showing an extremely strong negative correlation between participation and labor supply. The simple means shown in Table 6 in Section II illustrate that same negative correlation in tabular form.

While the estimates in model (1) are intended primarily for illustration, those in model (2) include three of the most important variables in the analysis. As the results show, participation is positively affected by the guarantee and negatively by the BRR, though in weaker form than previously. The results of the hours equation indicate that guarantees have a significantly negative impact on hours of work. The BRR also has a negative impact on hours of work, but not at a significant level. However, the negative sign on BRR does corroborate the grouped-state findings of Moffitt (1985), who also found negative partial correlations between BRR and hours of work.

The time dummies in the two equations indicate once more that there are unexplained shifts in participation over the period, but also that there are unexplained changes in hours of work as well. Thus the results in model (2) of Table 14 offer preliminary evidence that, indeed, the drop in hours of work from 1967 to 1973 occurred for reasons independent of the BRR.

A more formal way to examine this question is to decompose the change in mean hours of work between 1967 and 1973 into that portion due to the change in participation rates and that part due to changes in hours of work conditional upon participation. The mean value of hours of work at time t (H_t) can be decomposed as follows:

$$E(H_t) = P_t H_{1t} + (1 - P_t)H_{0t}, \tag{19}$$

where

P_t = probability of being on AFDC at time t,

H_{1t} = expected value of hours of work if on AFDC at time t,
$= E(H_t | P_t = 1)$,

H_{0t} = expected value of hours of work if not on AFDC at time t,
$= E(H_t | P_t = 0)$.

Table 14

Pooled Estimates of the Participation-Labor-Supply Model
(standard errors in parentheses)

	(1)		(2)	
	Participation	Hours	Participation	Hours
Guarantee/100	—	--	0.07* (0.03)	-2.25* (0.53)
BRR	—	—	-0.27 (0.24)	-2.01 (4.91)
Unemployment rate	—	--	0.04* (0.02)	-1.79* (0.46)
D73[a]	0.33* (0.05)	-7.58* (1.06)	0.27* (0.06)	-5.38* (1.29)
D79[b]	0.19* (0.05)	-3.23* (1.00)	0.15* (0.07)	0.09 (1.50)
Intercept	0.62* (0.76)	10.62* (0.76)	-0.63* (0.10)	24.20* (2.10)
σ_ν	—	32.21* (1.10)	--	32.15* (1.09)
ρ	-0.98* (0.54)		-0.97* (0.37)	
Log LF	-14710		-14680	

[a]D73=1 if observation is in 1973 sample, = 0 if not.

[b]D79=1 if observation is in 1979 sample, = 0 if not.

*Significant at the 10 percent level.

By standard decomposition manipulations it can be shown that

$$E(H_{t+1}) - E(H_t) = A_t + B_t, \tag{20}$$

where

$$A_t = [(1 - P_t)H_{0,t+1} + P_tH_{1,t+1}] - [(1 - P_t)H_{0t} + P_tH_{1t}], \tag{21}$$

$$B_t = (P_{t+1} - P_t)(H_{1,t+1} - H_{0,t+1}). \tag{22}$$

The term A_t represents the change in hours of work that would occur if the participation rate were held constant but hours of work of participants and nonparticipants changed, while the term B_t represents the change in hours solely due to the change in the participation rate. The weight in the B_t term is the difference in the hours of work of participants and nonparticipants.

The terms A_t and B_t can be evaluated using the parameter estimates in Table 14 to calculate the probabilities and expected hours values in the decomposition (the calculation of the conditional hours means involves using standard truncated normal formulas). When the simple dummy-variable estimates in model (1) of Table 14 are applied, the term B_t is almost five times the size of the term A_t. Consequently, the shift in participation rates explains about 83 percent of the drop in hours of work between 1967 and 1973. When the estimates in model (2) are used, and the mean guarantee, BRR, and unemployment rate in each of the years are used to calculate the relevant terms, the term B_t is more than six times the size of A_t; thus the participation-rate change explains about 86 percent of the drop in hours of work. However, in this latter case, part of the increase in the participation rate is explained by the change in the BRR itself. But, as noted previously, the change in the BRR explains only about one-fifth of the change in participation probabilities between 1967 and 1973. Thus the remaining "unexplained" shift upward in participation still explains the bulk of the drop in hours of work. The negative sign on the BRR in the hours equation in Table 14 implies that hours of work would have fallen further from 1967 to 1973 had the 30-and-one-third deductions not been in place.

VI. SUMMARY AND CONCLUSIONS

The main analysis in this report has been aimed at providing some new evidence on an old issue--was there a "structural change" in the propensity to participate in AFDC in the late 1960s and early 1970s? The most likely sources of such structural shifts are attitudinal changes over the period and the impacts of a number of liberalizing court decisions. Or was the increase in participation a result of changes in the labor market, changes in the characteristics of female household heads in the United States, increases in benefit levels, or increases in earnings deductions? Earnings deductions could be expected to increase participation rates because they raise the AFDC break-even level and make a greater portion of the population eligible for the program.

The results of the analysis in this report provide strong evidence that there was indeed a structural shift. The labor market changes over the period explain little of the change in participation, nor was there any change in demographic characteristics of the female-head population significant enough to induce such an increase in participation. In addition, the guarantee in the program, which is shown in the analysis here to have a statistically significant positive effect on participation rates, fell in real terms over the period 1967 to 1973. Thus it also can provide no explanation. On the other hand, the benefit-reduction rate in AFDC declined between 1967 and 1973, primarily because of the earnings deductions (the 30-and-one-third) legislated in the 1967 Social Security Amendments. The analysis here shows that these deductions indeed increased the participation rate over the period, yet only about one-fifth of the increase in participation can be attributed to this source.

Thus the evidence gives indirect support to the hypothesis that the welfare explosion of the late 1960s and early 1970s was primarily a result of changes in attitudes and the impacts of court decisions.

A secondary analysis of hours worked among all female heads of household indicates that the upward shift in participation over the period resulted in a sharp drop in the work-effort levels of female household heads. However, the earnings deductions implemented by the 1967 Amendments retarded this drop somewhat and provided some work incentives to the population.

The implications of these results for policy are somewhat mixed. On the one hand, attitudinal changes are difficult if not impossible to control with available policy tools, so there is no reason that dramatic changes in the caseload could not occur again in the future. Court decisions are, of course, more amenable to policy, but are not under the control of either the executive or the legislative branch. However, the most positive implications for policy of the study are the rigorous demonstrations that alterations in benefit levels and earnings deductions can affect the AFDC caseload.

Notes

[1]The 1967 Amendments were implemented in 1969, but at different times in different states. The data used for the calculation of the 1969 figures in the table were collected before most of the states had implemented the legislation.

[2]If anything, because the participation rate should be expected to be lower in the portion of the female-head population with high earnings, one should expect the participation rate among eligibles to fall with an increase in the break-even level.

[3]The benefit-reduction rates in Table 1 are "effective" benefit-reduction rates, i.e., incorporating all deductions in the program. They are thus smaller, sometimes considerably so, than nominal benefit-reduction rates. The trends in this variable between 1971 and 1981, when the nominal rate was constant, are apparently a result of changes in the generosity of allowable deductions provided by state legislatures. See Fraker et al. (1985).

[4]The regressions in Table 5 were reported in Moffitt (1985) but those in Table 4 were not, even though both were estimated at that time. The 1985 report was focused primarily on work effort rather than participation.

[5]The retrospective income questions in the CPS do allow one to determine whether a female head received any AFDC income in the prior year. But matching an annual measure of AFDC receipt to the point-in-time AS data would involve unacceptable approximation error, as would the assumption that any female head who was on AFDC in the prior year is also on AFDC as of the following March. One could, of course, employ the CPS

only and study the determinants of participation over the year, but such a participation variable is of less interest than that used here. In addition, it is difficult to conduct a study of work incentives in this case, for, although the CPS data provide information on earnings and weeks worked in the year, they do not allow one to determine earnings and weeks worked separately for the periods of AFDC participation and non-participation.

[6]Also, as a practical matter, the data used here were drawn from the extracts made in the prior DHHS project. Recreating the extracts from the CPS would have entailed greater project cost than gain in accuracy.

[7]To do so would require specifying a wage equation with variables that are not in the participation or hours-of-work equation, a difficult task. The equations to be estimated in the report should be thought of as reduced-form equations derived from participation and labor supply equations into which a wage equation has been substituted.

[8]In addition, it is not obvious whether the broad or narrow definition of eligibility is appropriate in general. For example, while it is true that the earnings disregards are not applied for eligibility determination, the fact that an eligible woman can receive those disregards after becoming eligible lessens the importance of the rule. In fact, if women do adjust their earnings voluntarily, the constraint should never be binding in the long run--that is, the caseload would be unaffected by the restriction. Also, the application of the assets test, given the poor CPS data on eligibility, is fairly problematical. There is also the question of whether, behaviorally, assets are not endogenous. In sum, it is fair to say that different definitions of eligibility are appropriate for different purposes.

[9]The reader unfamiliar with such models may wish to refer to Maddala (1983) for a discussion of the relevant literature.

[10]Although $P_i=0$ is necessarily observed for ineligibles, we have no information on their propensities to participate if they were made eligible.

[11]If the j^{th} element of X_{ij} is not in Z_i, its coefficient will still be inconsistent if it is correlated with the elements of Z_i.

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