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HAS STATE REDISTRIBUTION POLICY GROWN MORE CONSERVATIVE? AFDC, FOOD STAMPS, AND MEDICAID, 1960-1984

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HAS STATE REDISTRIBUTION POLICY GROWN MORE CONSERVATIVE?

I. INTRODUCTION

The decline in real AFDC benefits over the 1970s and 1980s is one of the most widely noted trends in the U.S. welfare system in recent years. It has been used as part of the explanation for the reversal in the historic decrease in poverty rates that occurred around 1980, when poverty rates started rising. The benefit decline is often cited in the popular press as evidence for a growing conservative climate, and it is currently playing a role in welfare reform discussions in Washington over arguments for a federally mandated minimum state AFDC benefit. It has also been used as an argument against the thesis that AFDC destabilizes marital unions, since the rise in the divorce rate and in female family headship in the United States has occurred over the same period that AFDC benefits have fallen.

The alternate explanation explored in this paper is that the decline reflects a substitution of federally funded Food Stamp benefits for partially state-funded AFDC benefits. State legislatures, which set the level of AFDC benefits, must pay approximately 40 percent of the marginal costs of benefit increases after federal matching. On the other hand, Food Stamp benefits are set by the U.S. Congress and are an externally provided extra benefit from the point of view of state legislatures. If it is the total transfer to its poor that enters the state utility function (or the utility function of the median voter), Food Stamp benefits will displace AFDC dollars on a one-for-one basis. The decline in real AFDC benefits thus could be a result of the increase in Food Stamp benefits that occurred over the 1970s and 1980s.

This hypothesis has been examined previously by Hulten et al. (1982), Gramlich (1982), Orr (1979), and Plotnick and Winters (1985). Orr found that Food Stamps substituted for AFDC benefits on a one-for-one basis, while Gramlich and Plotnick and Winters found no substitution. Hulten et al. found evidence of substitution but stressed that the models that have been used have been extremely nonrobust and sensitive to specification. The variance in results in these studies may be, in part, a reflection of a key difficulty in testing the hypothesis. Because the Food Stamp benefit schedule is uniform in the nation as a whole, a cross-sectional regression of state-specific AFDC benefits on a standardized Food Stamp benefit is not possible. Moreover, the number of time periods available over the 1970s and 1980s is too small to conduct a reliable time-series analysis, and in any case the evidence is strong that the states were not in equilibrium in those years. The previous studies attempted to circumvent this problem by using various sources of cross-sectional variation in the Food Stamp benefit actually paid out in a state to identify the Food Stamp substitution effect. However, the validity of using such variation to measure the substitution effect is subject to question and, in any event, the amount of such variation is small.

In this paper a more direct method of attack is taken. Cross-sectional regressions of AFDC benefits are estimated at a point in time (1960) prior to the introduction of Food Stamps, and the results are then used to forecast the sum of AFDC and Food Stamps at a later time (viz., in 1984). If the substitution hypothesis is correct, the 1960 regression should correctly forecast the later sum. This method, while direct, is also fairly heroic. Time-series forecasts from cross-sectional regressions are notoriously poor and, in this case, there is an additional difficulty created by the transformation of the U.S. welfare system between 1960 and 1984, making it perhaps unlikely that there has been no structural change in the AFDC benefit equation. But the possibly quixotic nature of the exercise also makes it a much stronger test of the hypothesis than has been previously attempted.

The results indicate surprisingly strong support for the substitution hypothesis. Forecasts of the real AFDC-Food Stamp sum to 1984 are only about $60 per month higher than the actual sum, as compared to a drop in the real AFDC benefit of $120 per month since 1960 (1982 dollars). When the Medicaid program is introduced, a stronger set of results is obtained. The AFDC benefit in 1960 was $200 per month lower than the sum of AFDC, Food Stamps, and Medicaid in 1984; but a prediction of the 1960 AFDC benefit from a 1984 regression with the benefit sum as the dependent variable comes within $9 of the actual 1960 AFDC benefit. Finally, almost all tests conducted on 1960 and 1984 differences find that the null hypothesis of no structural change cannot be rejected.

In the next section of the paper, the background time-series trends in the AFDC benefit, and in variables that might have caused its decline, are discussed. The models and econometric methods used to test the Food Stamp hypothesis are discussed in Section III, followed by a presentation of the main results in Section IV. A supplementary analysis of the AFDC-U program (that for which unemployed men are eligible) is reported in Section V. A summary and a discussion of the policy implications of the paper are provided in the last section.

II. THE TIME-SERIES EVIDENCE AND THE INSTITUTIONAL BACKGROUND

A relative brief graphic exposition of the relevant trends should provide a proper context for the econometric work. The dramatic reduction in real AFDC benefits since the late 1960s is illustrated in Figure 1. As the figure indicates, the real benefit grew steadily into the early 1960s and accelerated slightly in the mid-1960s. But around 1967 or 1968, the increase came to a halt and benefits took a sharp nosedive, setting off a decline that continued all the way to 1981. Since 1981 the benefit has leveled off and has remained essentially constant.

That changes in the U.S. political climate leading to more conservative policies occurred at about the same time as the benefit decline leads to the obvious hypothesis that the benefit reduction has resulted from changes in preferences toward redistribution. State legislatures, which set AFDC benefits, are traditionally more conservative than the Congress and could be argued to be particularly susceptible to changes in the attitudes of voters. Nevertheless, there could be economic causes of the change as well, and these clearly need to be explored.

One such alternative hypothesis is suggested by the trend in the AFDC caseload, also shown in Figure 1. The increase in the benefit in the early 1960s was followed shortly thereafter by an explosion in the AFDC caseload, for the number of AFDC families per capita almost tripled over the six years between 1966 and 1972. Although the caseload has since leveled off, the subsequent reduction in the benefit may simply have been a response to the

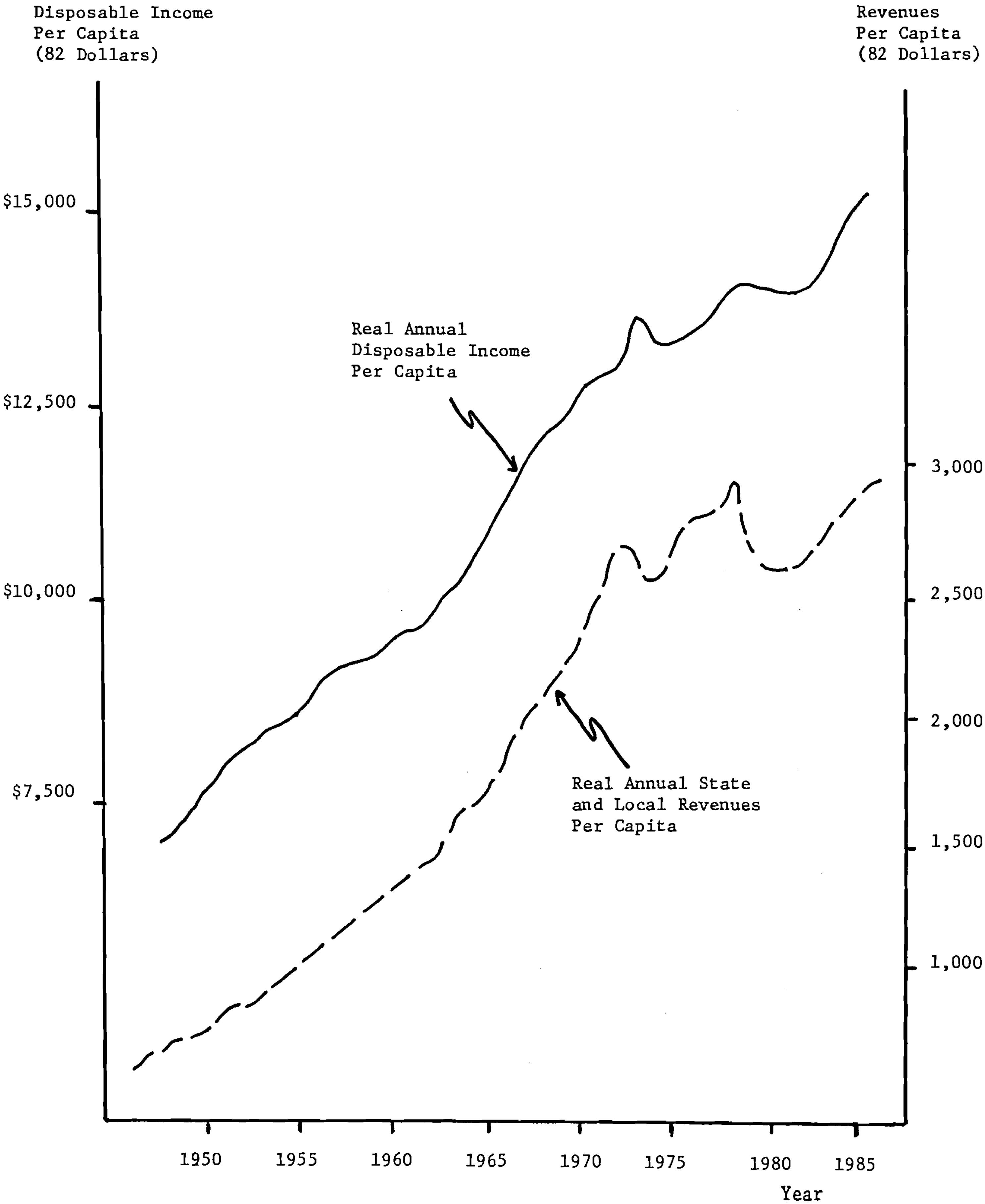

AFDC Families
Per Thousand Pop.
Real
Benefit
(82 Dollars)
20
15
10
5
400
350
300
250
200
Real Monthly AFDC
Benefit Per
Family
Average Monthly No.
AFDC Families Per
Capita
1940
1945
1950
1955
1960
1965
1970
1975
1980
1985
Year

Figure 1

caseload increase. As should be quite intuitive, and as will be demonstrated formally below, the caseload is effectively the price of the benefit; consequently, the caseload explosion represented a 300 percent increase in the price of AFDC benefits.[1]

This caseload increase suggests that the states may have allowed real benefits to decline simply in order to keep real AFDC expenditures by the states constant, or at least growing in line with income. However, the benefit decline exceeded what was necessary to achieve this. AFDC expenditures leveled off in the late 1960s and early 1970s and then declined in absolute terms after about 1973, as should be clear from the caseload and benefit trends in Figure 1. A fortiori, AFDC expenditures declined as a fraction of state revenues (more on this momentarily). Whether this should be expected or not depends upon whether the price elasticity of the benefit does or does not exceed one.

Another potential source of the benefit reduction is the well-known reduction in the growth rate of real income over the 1970s. As shown in Figure 2, real income per capita in the United States grew during the 1970s, but at a slower rate than in the 1960s and earlier. Whether this income slowdown is sufficiently large to explain the benefit decline is an empirical question, of course, and will be examined below. But it is consistent with the growth pattern of per capita state and local revenues, also shown in the figure, which flattened out markedly in the 1970s. Naturally, along with the slowdown in revenue growth came a slowdown in expenditure growth. Figure 3 shows the trend in the per capita budget surplus, which indirectly reflects expenditure trends. Surprisingly, the budget surplus in the state and local



Disposable Income
Per Capita
(82 Dollars)
Revenues
Per Capita
(82 Dollars)
$15,000
$12,500
$10,000
$7,500
3,000
2,500
2,000
1,500
1,000
Real Annual
Disposable Income
Per Capita
Real Annual State
and Local Revenues
Per Capita
1950
1955
1960
1965
1970
1975
1980
1985
Year

Figure 2

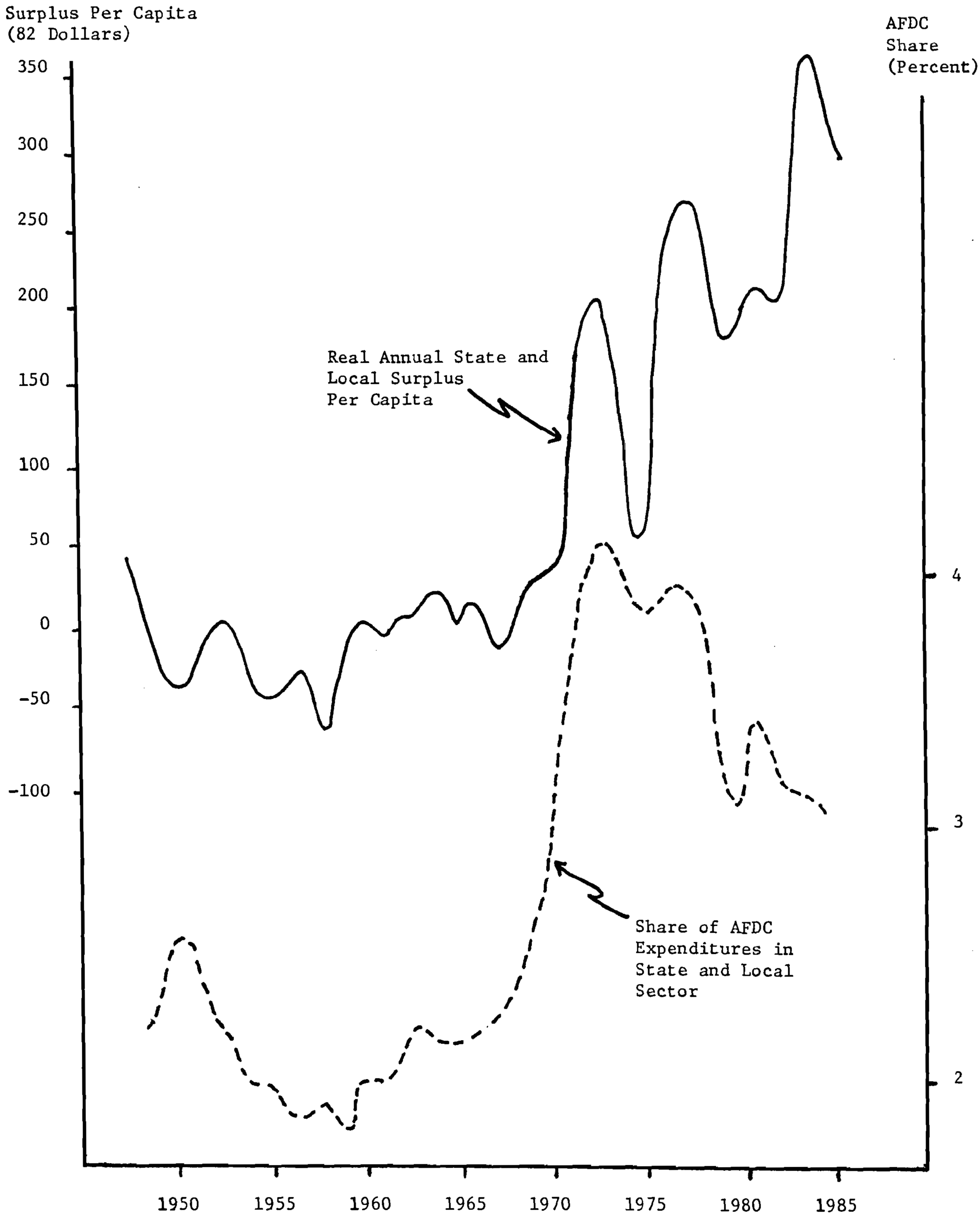

Surplus Per Capita
(82 Dollars)
AFDC
Share
(Percent)
350
300
250
200
150
100
50
0
-50
-100
4
3
2
Real Annual State and
Local Surplus
Per Capita
Share of AFDC
Expenditures in
State and Local
Sector
1950
1955
1960
1965
1970
1975
1980
1985

Figure 3

sector actually increased over the 1970s, implying that expenditure growth declined by more than that of revenues (and that there was no long-term change in "fiscal distress," as it is called in the literature, in that sector). Nevertheless, despite this decline in growth of state and local expenditures, expenditures on AFDC fell even more and even in absolute terms. As a consequence, AFDC expenditures fell as a share of total state and local expenditures, as shown in Figure 3. Thus the decline in the AFDC benefit could not have been wholly a result of general revenue and expenditure decline.

An additional hypothesis for the decline in the AFDC benefit is a reduction in the generosity of federal matching for AFDC benefits. Matching rates for AFDC did indeed decline over the period, but the reductions were quite small in absolute terms--from a mean of 58 percent to one of 56 percent for the regular AFDC matching rate and from 62 percent to 59 percent for the Medicaid matching rate. It will require a large price elasticity for these reductions to generate the drastic decline in the AFDC benefit observed.

The alternative hypotheses to be examined in detail here are those relating to the possible substitution of non-AFDC benefits for the AFDC benefit. The most prominent source of such effects is the possible substitution of federally financed Food Stamp benefits for state-financed AFDC benefits. The Food Stamp program was introduced in the mid-1960s as an option for the states and grew slowly until 1974, when Congress mandated that all states implement the program in all their counties. As shown in Table 1, the AFDC guarantee--the amount actually set by state legislatures--declined rapidly after the introduction of Food Stamps; in fact, the real AFDC

Table 1
Trends in U.S. Monthly Public Assistance Transfer Benefits (1982 dollars)

	1960	1964	1968	1972	1974	1976	1978	1980	1982	1984
AFDC Guarantee[a]	$483	$471	$506	$512	$507	$500	$475	$436	$394	$388
Food Stamp Guarantee[a]	0	0	n.a.	227	212	259	237	236	233	234
AFDC & Food Stamps	-	-	n.a.	739	719	759	712	672	627	622
(.7) AFDC & Food Stamps	-	-	n.a.	586	567	609	570	541	509	505
Medicaid[b]	0	0	n.a.	142	191	196	192	176	178	163
AFDC & Food Stamps & Medicaid	-	-	n.a.	881	910	955	904	848	805	785
(.7) AFDC & Food Stamps & Medicaid	-	-	n.a.	728	758	804	762	717	687	669

[a]Maximum amount paid for a family of four with no other income.

[b]Insurance value, equal to medical expenditures per AFDC family of four.

n.a. = not available.

guarantee was 25 percent lower in 1984 than it had been in 1960.[2] On the other hand, real Food Stamp benefits, while fluctuating over the 1970s and 1980s, have essentially followed no trend and have been approximately constant, no doubt because they were indexed to inflation by Congress in 1972. In any case, the sum of real AFDC and Food Stamp benefits in 1984 was $622 per month, almost 30 percent higher than that for AFDC alone in 1960.

The hypothesis that states have allowed Food Stamps to substitute for AFDC benefits in a one-for-one ratio has several elements supporting it. First, until 1979 AFDC recipients were automatically eligible for Food Stamps, regardless of income, and hence almost all AFDC recipients received them. Since 1979, when AFDC recipients began to be certified for Food Stamps on the basis of their income and assets, a high proportion have still received them. Second, in most states Food Stamp certification of AFDC recipients takes place physically in the AFDC offices themselves, providing strong motivation for integration of the programs. Third, the cash equivalent value of Food Stamps is essentially the same as their market value, making it unlikely that recipients would be worse off by having food transfers substituted for cash transfers.

Nevertheless, the details of the the Food Stamp benefit formula imply that simply summing AFDC and Food Stamp benefits together is not quite correct. The Food Stamp program taxes AFDC benefits at approximately a 30 percent rate, lowering the net income increment provided by the program by 30 percent of the AFDC amount. Hence, an increase in the AFDC benefit of $1 would raise disposable income of the recipient by only 70 cents. As shown in Table 1, the net sum of AFDC and Food Stamps in 1984 was $505 per month, only

5 percent higher than the value of AFDC alone in 1960. This is a rather small increase over the twenty-four years, given the much higher levels of taxpayer income in 1984.

This taxation also implies that a reduction in the state AFDC benefit by one dollar would lower the net transfer to female-headed families in the state by only 70 cents. Thus an additional incentive for state legislatures to let the real AFDC benefit decline is provided.

A second source of benefit substitution that may have occurred is the substitution of Medicaid benefits for AFDC. The Medicaid program was introduced by the U.S. Congress in 1965 and grew rapidly over the late 1960s and 1970s, at the same time that AFDC benefits were declining. AFDC recipients are categorically eligible for Medicaid benefits so that, even though not all receive medical care in any given time interval, all are essentially covered by health insurance and hence should be thought of as receiving a transfer. As in the AFDC program, state legislatures pay for Medicaid expenditures, but payments are matched by the federal government at the Medicaid matching rate referred to earlier. However, unlike AFDC, the basic set of medical services provided to recipients is mandated by the federal government. Although states can supplement the basic set, and can even in some circumstances put restrictions on the basic set (e.g., by limits on hospital days), it is nevertheless the case that the core of Medicaid expenditures are mandated by the federal government and hence are not under the control of the states.

As Table 1 shows, the Medicaid benefit (i.e., its insurance value) grew in the early 1970s and declined in the later 1970s. The decline was the result of sharp increases in medical care inflation and consequent reductions in service delivery. By 1984, the sum of AFDC, Food Stamps and Medicaid was $785 per month, over 60 percent greater than the value of AFDC alone in 1960. Taking into account the taxation of AFDC by the Food Stamp program, the net sum of the three benefits was $669 per month in 1984, 39 percent higher than the value of AFDC alone in 1960.

A further source of possible benefit substitution, also related to the Medicaid program, is the substitution of non-AFDC Medicaid benefits for AFDC-related benefits. AFDC recipients account for only 25 percent of Medicaid expenditures, the other 75 percent consisting predominately of the aged and the disabled. The average Medicaid benefit for the aged is about double that for AFDC families and that for the disabled is about triple that for AFDC families. The explosion in medical care prices in the 1970s led to tremendous growth in non-AFDC Medicaid expenditures, particularly for nursing home care. As shown in Figure 4, non-AFDC Medicaid expenditures per capita grew strongly all the way into the late 1970s, until medical care inflation once more generated service reductions. In regard to AFDC benefits, the simple implication is that non-AFDC Medicaid expenditures may have crowded out AFDC Medicaid expenditures and the AFDC benefit itself in the state budget; that is, the two may be substitutes in the state utility function.[3]

Plan of Analysis. The goal of the analysis in the next few sections is to test these substitution hypotheses against the data, and to determine whether the other possible causes of benefit decline (caseload growth, matching rates, real income growth slowdown) are sufficient by themselves, either alone or in combination, to explain it. The central difficulty in testing the Food Stamp portion of the hypothesis, which is the primary one, is

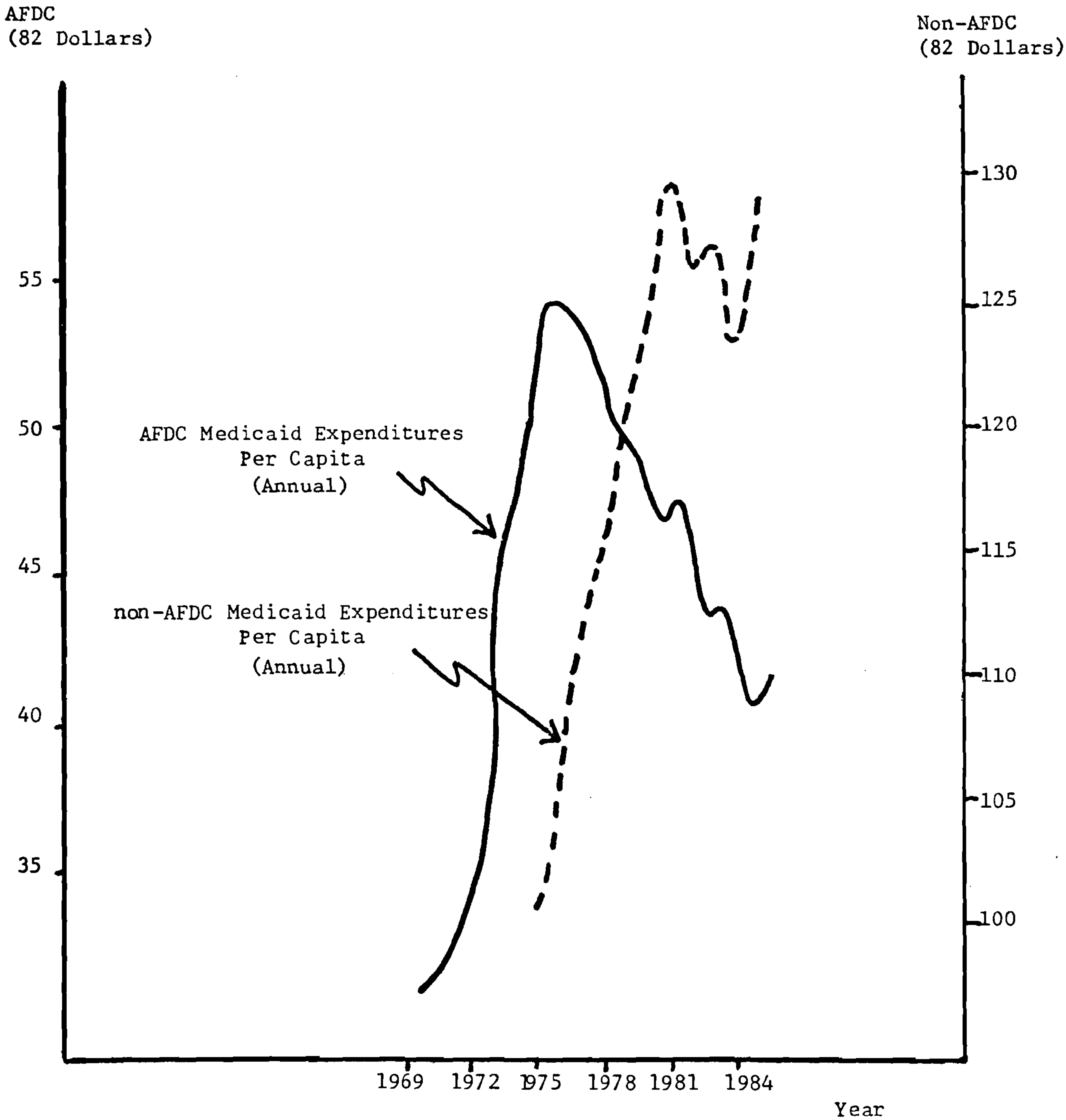

AFDC
(82 Dollars)
Non-AFDC
(82 Dollars)
55
50
45
40
35
130
125
120
115
110
105
100
AFDC Medicaid Expenditures
Per Capita
(Annual)
non-AFDC Medicaid Expenditures
Per Capita
(Annual)
1969
1972
1975
1978
1981
1984
Year

Figure 4

that the Food Stamp benefit schedule is set by Congress and therefore does not vary across the states. Consequently, no cross-sectional correlation between AFDC and Food Stamp benefits can be estimated.

In past studies of the substitution hypothesis, this difficulty has been circumvented in various ways. Orr (1979) conducted a cross-sectional analysis in the mid-1970s and regressed the AFDC benefit on the average Food Stamp benefit actually paid in each state. The Food Stamp benefit actually paid differs across states because of differences in family size, nonwelfare income, and the amounts of various deductions in the benefit formula.[4] However, legislatures are well aware of the difference between a guarantee set in a benefit formula--which holds family size, other income, etc. constant--and an actual benefit paid out. In setting the guarantee in the AFDC program, state legislatures typically consider only the guarantee in the Food Stamp program in their calculations. In addition, econometrically speaking, identifying the effect of Food Stamps by using variation in family size and other variables requires the assumption that those variables do not affect state AFDC actions directly, which is unlikely to be the case. Moreover, the variation in family size and other variables across states is not very large, which could lead to unstable results. A reanalysis of the Orr data by Hulten et al. (1982) is consistent with this possibility, for Hulten et al. found the Food Stamp coefficient in the Orr model to be quite sensitive to the inclusion of additional state-specific variables. Finally, regardless of the legitimacy of family size and other such variables as instruments, the cross-sectional variation in the Food Stamp benefit they induce is unlikely to have the same effect on the AFDC benefit as will an upward shift in the entire Food Stamp schedule, as has occurred over time.

Plotnick and Winters (1985) (see also Plotnick, 1986) used cross-sectional variation in the Food Stamp program in 1971 and 1972, when the program was not in place in all counties. The Food Stamp benefit was multiplied by the fraction of the counties in the state that had instituted the program. However, as discussed by Orr (1979), the states that adopted the program first were the more liberal, high-benefit states; thus the Plotnick-Winters variable runs the risk of some degree of endogeneity. On the other hand, Gramlich (1982) used time-series variation in the Food Stamp benefit from 1974 to 1981 to estimate the substitution effect.[5] Gramlich found his results to be quite sensitive to the specification assumed. This is not too surprising for, as Table 1 above shows, the AFDC benefit fell from 1974 to 1981 and the Food Stamp benefit fluctuated with little or no pattern. In fact, given the long-standing decline of the AFDC benefit prior to 1974 and its leveling off after 1981, it is unlikely that the states were in equilibrium over the 1974-1981 period.

To avoid these difficulties, this study takes a more direct approach to the essentially time-series nature of the question (i.e., why did AFDC benefits decline over a specific calendar period?) by using cross-sectional AFDC benefit equations estimated prior to the introduction of Food Stamps and Medicaid to forecast benefits forward to a period in the future when states had fully adjusted to Food Stamps and Medicaid. Pre-1965 benefit regressions are used to forecast the effects of changes in the caseload, matching rates, and state income on the benefit, and comparisons of the forecasted mean benefit and the actual mean benefit are then used to test for structural change in the AFDC benefit equation over time, and hence for the substitution hypothesis. Some "backcasts" are performed as well by estimating cross-sectional regressions in 1984 and predicting the AFDC benefit in 1960, and some direct pooling across years is conducted to test for structural change directly. This approach avoids the requirement of artificially generating cross-sectional variation in Food Stamp amounts or of having to use AFDC-Food Stamp time-series correlations in the 1970s to test for substitution.

As a method of testing the weak version of the substitution hypothesis--namely, that there was some substitution though not necessarily on a dollar-for-dollar basis--the approach is rather weak, for it implies that any significant difference between actual and forecasted AFDC benefits be taken as a sign of substitution. Obviously other factors could have been at work. However, as a method of testing the strong version of the hypothesis--that the substitution was actually one-for-one--the approach is correspondingly strong, for it implies that the forecasted AFDC benefit should equal a precise dollar amount. The fact that cross-sectional regressions generally have tracked time-series variables rather poorly in most past applications strengthens the nature of the test even more, especially when the drastic transformations of the welfare system in the late 1960s and early 1970s are recognized.

III. MODELING THE EFFECT OF FOOD STAMPS AND MEDICAID ON AFDC

Since most of the regressions to be estimated will be based upon only 48 observations, the models must be kept as simple as possible. In the simplest, the median voter of each state allocates his income to expenditure on the AFDC benefit and on other goods, conditional upon a fixed Food Stamp benefit provided by the federal government. Unfortunately, the data do not contain information on the income or the tax price faced by the median voter, so the mean voter must be used instead, as an approximation.

Model I

$$\text{Max} \quad U(B + \psi F, Z), \tag{1}$$

$$\text{s.t.} \quad Y = PB + Z, \tag{2}$$

where B is the AFDC guarantee for a fixed family size (e.g., four); F is the Food Stamp guarantee for the same family size; Z is the per capita amount of some other composite good; Y is per capita income in the state after federal taxes but before state taxes;[6] and P is the price of the AFDC benefit. P is equal to $(C/N)(1-s)$, where C is the AFDC caseload, N is state population, and s is the federal matching rate for AFDC expenditures. Approximating the solution to the maximization problem for B with a linear demand equation, we have

$$B = \alpha + \beta P + \gamma \hat{Y} - \psi F, \tag{3}$$

where $\hat{Y} = Y + \psi PF$ is virtual income, incorporating the income effects arising from the federal gift of Food Stamps. Conventional theory predicts that $\beta < 0$ and $\gamma > 0$. Here interest centers on tests of the null hypothesis H_0: $\psi = 1$, under which the demand equation simplifies to the following:

$$B + F = \alpha + \beta P + \gamma(Y + PF). \tag{4}$$

Thus an increase in F of \$1 will lower B by \$1, controlling for income effects. Not controlling for such effects will generate a reduction of B of less than \$1.

As noted in the previous section, a fully rational voter will realize that the Food Stamp program taxes AFDC benefits, leading to a variation on this model.

Model IA

$$\text{Max} \quad U(B + \psi F', Z), \tag{5}$$

$$\text{s.t.} \qquad Y = PB + Z, \tag{6}$$

$$F' = F - .3B, \tag{7}$$

which leads to the demand equation

$$B = (\alpha/\omega) + (\beta/\omega^2)P + (\gamma/\omega)\hat{Y} - (\psi/\omega)F, \tag{8}$$

where $\omega = 1 - .3\psi$ and $\hat{Y} = Y + (\psi/\omega)PF$. Under the null of $\psi = 1$, the equation reduces to

$$.7B + F = \alpha + \beta(1.43)P + \gamma[Y + (1.43)PF]. \tag{9}$$

The introduction of the tax rate has, surprisingly, ambiguous effects on the level of the benefit. To illustrate, let t be the tax rate ($t = .3$ currently). Substitution of t for .3 in (9) and differentiation of (9) w.r.t. the tax rate can be shown to lead to

$$\left.\frac{1}{B}\frac{\partial B}{\partial t}\right|_{t=0} = 1 + \frac{\beta P + \gamma PF}{B_0}, \tag{10}$$

where $B_0 = \alpha + \beta P + \gamma(Y + PF) - F$. The price effect ($\beta < 0$) tends to make the tax effect negative, as is intuitive, but the income effect ($\gamma > 0$) moves it in the opposite direction. Moreover, the benefit increases by one percent from a unit increase in t because the utility function now contains $B(1-t) + F$ as its first argument, implying that the benefit must be increased in order to leave utility at the same level as previously.

Perhaps more important for present purposes, the introduction of this tax rate has ambiguous effects on the substitution effect of F on B. While the tax rate increases the income effects of PF in (9), the same multiplicative effect of $(1-t)$ on B implies that the reduction in B will be greater than before.

Medicaid benefits can be introduced to this model in a similar fashion, leading to Model II.

<u>Model II</u>

$$\text{Max} \quad U(B + \psi F + \phi M, Z), \tag{11}$$

$$\text{s.t.} \qquad Y = P(B + QM) + Z, \tag{12}$$

where M is the insurance value of the Medicaid benefit for AFDC recipients and Q is the relative price of medical care.[7] The resulting demand equation for B can be written:

$$B = \alpha + \beta P + \gamma\hat{Y} - \psi F - \phi M, \tag{13}$$

where $\hat{Y} = [Y + P(\psi F + \phi M) - QPM]$. The income effects in the virtual-income terms are in this case partly negative $(-QPM)$ because the federal "gift" of Medicaid is not free--states must still pay a share. If, in fact, the positive income effects are equal but opposite in sign $(\phi = Q)$, there are no income effects of Medicaid.

Under the null of $\psi = \phi = 1$, the demand equation becomes

$$B + F + M = \alpha + \beta P + \gamma[Y + P(F + M) - QPM]. \tag{14}$$

As with Model I, a variation can be introduced to allow for rational perception of the taxation of AFDC by the Food Stamp program:

Model IIA

$$\text{Max} \quad U(B + \psi F' + \phi M, Z), \tag{15}$$

$$\text{s.t.} \qquad Y = P(B + QM) + Z, \tag{16}$$

$$F' = F - .3B, \tag{17}$$

which leads to a demand equation under the null of the form:

$$.7B + F + M = \alpha + \beta(1.43)P + \gamma[Y + (1.43)P(F + M) - QPM]. \tag{18}$$

Finally, the influence of non-AFDC Medicaid expenditures can be incorporated by modifying the utility function to allow a third argument representing the transfer to non-AFDC Medicaid recipients, primarily the aged and disabled. A separate argument is required because the marginal utility of transfers to the aged and disabled is likely to be quite different from that of transfers to female heads of family. As noted in the previous section, Medicaid transfers to the aged and disabled are much greater than to female family heads, which could be interpreted as evidence that the marginal utility of transfers to the former are higher than to the latter.

Model III

$$\text{Max} \quad U(B + \psi F + \phi M,\ M_N,\ Z), \tag{19}$$

$$\text{s.t.} \qquad Y = P(B + QM) + QP_N M_N + Z, \tag{20}$$

where M_N is the Medicaid benefit to non-AFDC recipients and P_N is the price of that benefit, equal to the product of $(1-s)$ and the per capita caseload in the non-AFDC portion of the Medicaid program.[8] The demand equation for B now becomes

$$B = \alpha + \beta P + \delta(QP_N) + \gamma\hat{Y} - \psi F - \phi M, \tag{21}$$

where $\hat{Y}$ is the same as in equation (13). If Medicaid and non-Medicaid expenditures are gross substitutes, $\delta > 0$. Under the null of $\psi = \phi = 1$, the equation becomes[9]

$$B + F + M = \alpha + \beta P + \delta(QP_N) + \gamma[Y + P(F+M) - QPM]. \tag{22}$$

Once again, incorporating the taxation of AFDC results in a modification. Without stating the maximization problem, suffice it to state that the demand equation under the null in this case is (Model IIIA)

$$.7B + F + M = \alpha + \beta(1.43)P + \delta(QP_N) + \gamma[Y + P(F+M) - QPM]. \tag{23}$$

Testing the Null Hypotheses. As discussed in the last section, the approach taken here to testing the nulls of dollar-for-dollar substitution is based upon forecasts from cross-sectional regressions for B estimated prior to the introduction of Food Stamps and Medicaid. The earliest year for which the AFDC guarantee for a family of four is available is 1960 and the latest year for which it and the independent variables are available is 1984. Both years can be reasonably argued to be equilibrium years, for in 1960 the AFDC system had been stable in structure and in caseload growth for over a decade and in 1984 the AFDC benefit appears to have settled down after the transitional years in the 1970s.[10] Thus cross-sectional regressions of B on P and Y in 1960 can be used to forecast the AFDC benefit under Models I-II to 1984, and significance tests can be conducted on the difference between the forecasted and actual mean AFDC benefit (Model III cannot be estimated on 1960 data because data on P_N are unavailable then). Such tests will indirectly determine the extent to which changes in the level of the caseload, matching rates, and disposable income between the years are capable of explaining the benefit decline.

This test is a very strong one because it does not utilize information in 1984, and because it therefore tests the joint null of $\psi = \phi = 1$ and of no structural change in the equation. It is implicitly a test of the rational-expectations forecast under the null of dollar-for-dollar substitution. A statistically symmetrical alternative is to conduct a set of tests by estimating Models I-III on the 1984 cross-section and by using them to predict the 1960 benefit. Such estimates obviously incorporate different information, particularly that on Medicaid, and Model III can be estimated. Conditional upon the outcome of this analysis, an obvious further set of tests can be conducted by pooling the two years and testing directly for structural change in the parameters, thereby using all the statistical information in the data.

The means of the variables used in the analysis are shown in Table 2.[11] As noted previously, the mean AFDC guarantee fell markedly between 1960 and 1984, but Food Stamps and Medicaid outweighed the AFDC decline. The caseload more than tripled, leading to a large effective price increase.[12] The matching rate appears to have increased over the period, but this is a result of the nonlinearity of the matching schedule in 1960, to be discussed momentarily. Real disposable income about doubled from 1960 to 1984. Although standard errors are not shown, the data reveal that the addition of Food Stamps has lowered the cross-state variance in transfers but Medicaid has raised it back to the same level as that of AFDC alone.

The nonlinearity of the matching rate schedule in 1960 requires modifying the estimation procedure for the estimates in that year.[13] In 1960 the federal government matched state AFDC expenditures at 83 percent of low benefit levels, at a state-specific "federal" matching rate at medium benefit levels (the mean of which is shown in Table 2), and at zero rates for high benefit levels. As is well known from the analysis of piecewise-linear budget sets (Hausman, 1985; Moffitt, 1986), the demand equation along each segment in 1960 can be written in this case as

$$B = \alpha + \beta\hat{P} + \gamma\hat{Y}, \tag{24}$$

Table 2
Means of the Variables Used
in the Analysis

	1960	1984
B (AFDC guarantee)	\$465	\$344
F (Food Stamp guarantee)	0	\$242
M (insurance value of Medicaid benefit)	0	\$185
$C/N^{a,c}$ (caseload/state pop.)	13.8	42.8
s^{b} (matching rate)	57.9	59.6
P^{c} (price of AFDC benefit)	5.6	17.6
Y (per capita income)	\$5,719	\$10,185
Q (price of medical care)	--[d]	1.14
$P_N^{a,c}$ (price of non-AFDC Medicaid benefit)	0	12.5
QP_N^{c}	0	14.8

Notes:
Data Sources: See Appendix B.
N = 48
All dollar figures in 1982 dollars.

[a]Caseload lagged 3 years.

[b]Multiplied by 100.

[c]Multiplied by 1000.

[d]Not required for analysis.

where $\hat{P}$ and $\hat{Y}$ are, respectively, virtual price and income on a segment. Their definitions are straightforward and are not written out for brevity. Monte Carlo evidence indicates that OLS estimates of (24) can give extremely biased estimates of the effects of grants-in-aid (Megdal, 1987), for $\hat{P}$ and $\hat{Y}$ are endogenous. Although the most efficient method of estimation of such models is maximum likelihood (see Moffitt, 1984, for an application to grants-in-aid), the main analysis presented below will instead use the instrumental variable technique of evaluating the schedule at the mean benefit for all observations. The mean benefit in 1960 was in the middle segment of the schedule, with the matching rate shown in Table 2. The implied virtual price and income are therefore used in all 1960 regressions. To test the sensitivity of the results to this procedure, maximum likelihood piecewise linear constraint (PLC) estimates are obtained as well on a subset of the models.

IV. MAIN RESULTS

Table 3 shows the results of estimating the 1960 AFDC benefit equations. Column (1), the simplest model, shows a significant and positive income effect and a negative, though insignificant, price effect. At the means of the data the coefficients imply price and income elasticities of -.17 and .98, respectively. The near-unity income elasticity implies that the share of income devoted to AFDC should stay approximately constant as income increases.

Columns (2) and (3) show the effects of entering additional state variables for region, urbanization, educational level of the population, and other factors. Region appears to be moderately important in explaining benefits, with the South showing the lowest benefits, as expected, and the West and Northeast showing the highest. However, neither the urbanization variables nor the variables added in column (3) are very significant, and an F-test strongly rejects the significance of the incremental variables in that column. Perhaps more important for present purposes, the inclusion of these variables has no quantitatively important effect on the income coefficient, but it does reduce the magnitude of the price effects greatly and renders them completely insignificant. Thus the 1960 data provide weak evidence of price effects at best.

Forecasts to 1984 are shown in Table 4 for Models I and II (Model III cannot be forecasted because no estimate of the parameter δ is possible

Table 3
1960 Benefit Regressions

	(1)	(2)	(3)
β	-14117.5 (8342.0)	-3032.9 (8614.9)	-1962.7 (10070.3)
γ^a	79.6* (18.9)	75.1* (27.1)	63.0* (45.4)
NE	-	131.4* (58.6)	136.8 (76.0)
NC	-	110.4 (55.7)	100.0 (74.0)
W	-	131.9* (55.6)	112.4 (83.5)
METPCT	-	-105.3 (100.0)	-93.8 (108.7)
PCHS	-	-	1.6 (6.1)
UN	-	-	-2.7 (20.7)
RW	-	-	0.2 (1.2)
Constant	88.9 (132.0)	34.0 (132.6)	1.5 (227.3)
R-squared	.41	.54	.54
Standard error	125	116	120

Notes:

Standard errors in parentheses.
* Significant at 5 percent level.

[a]Multiplied by 1000.
Variable definitions: NE, NC, W are regional dummies for Northeast, North Central, and West, respectively (South omitted); METPCT = percent of state population in metropolitan areas; PCHS = percent of population with a high school degree; UN = state unemployment rate; RW = real weekly manufacturing wage. Means given in Appendix Table C-1.

Table 4
1960 Forecasts to 1984

Model I	
Actual B+F	585
Forecast B+F	645
Forecast error	60
F-statistic	.15
(p-value)	(.70)
Standard error	156
Model IA	
Actual .7B+F	482
Forecast .7B+F	537
Forecast error	55
F-statistic	.07
(p-value)	(.80)
Standard error	214
Model II	
Actual B+F+M	770
Forecast B+F+M	645
Forecast error	-125
F-statistic	.64
(p-value)	(.43)
Model IIA	
Actual .7B+F+M	667
Forecast .7B+F+M	537
Forecast error	-130
F-statistic	.37
(p-value)	(.55)
Standard error	214

with the 1960 data). All forecasts use the coefficients in column (1) in Table 3.[14] The forecasted benefit in 1984 is $645 per month, far above the actual AFDC benefit of $344. But this forecast is only $60 above the sum of AFDC and Food Stamps in 1984, a very close forecast. The forecast is insignificantly different from the actual value, though this is partly a result of a fairly high standard error of the prediction ($156). The forecast of the model under the assumption that voters recognize the taxation of AFDC by the Food Stamp program (Model IA) is quite similar, overpredicting the net benefit sum by approximately $55, again insignificantly different from zero. When used to predict for models including Medicaid (Models II and IIA), the 1960 regressions underpredict because the Medicaid benefit is about $180. The forecast error ranges from $125 to $130, but is again insignificantly different from zero.

These 1960-based forecasts provide considerable support for the full substitution hypothesis, for the forecasts are fairly close given the major changes that occurred in the system between 1960 and 1984, and given the fairly large absolute magnitude of the implied increase in the benefit sum between the years. Even for Model II, for example, the implied increase in the benefit sum is $305 ($770 - $465), and the 1960 regression predicts two-thirds of that increase.

The estimates of Models I-III on the 1984 data are presented in Table 5.[15] The results are for the most part quite similar to those in 1960. Price effects, though negative, are on the borderline of significance at conventional levels in most of the models. Even in those models (IIA and IIIA) where price effects are significant, the implied elasticities are quite

Table 5
1984 and Pooled Regressions

	1984 Regressions						Pooled	
	I	IA	II	IIA	III	IIIA	III	IIIA
β	-854.4 (1989.9)	-420.8 (974.6)	-4569.1 (2561.4)	-3023.9* (1414.5)	-5075.0 (3081.2)	-2946.9* (1702.9)	-6629.8* (2453.1)	-4348.0* (1623.4)
δ	-	-	-	-	789.5 (2610.1)	-171.9 (2061.7)	1285.2 (2381.4)	-180.7 (2194.0)
γ^a	50.2* (14.7)	35.1* (10.2)	71.4* (18.9)	56.4* (14.7)	71.0* (19.1)	56.4* (15.1)	81.5* (8.3)	67.7* (7.8)
α	89.1 (128.9)	134.8 (90.2)	123.4 (166.0)	168.9 (131.0)	123.6 (167.7)	168.9 (132.5)	37.1 (54.4)	101.8 (50.3)
R^2	.29	.29	.26	.25	.26	.25	.66	.55

Notes:

Standard errors in parentheses.

* Significant at the 5 percent level.

[a]Multiplied by 1000.

low (about 12 percent). Income effects, on the other hand, are positive and significant in all models. The implied elasticities center around .94, once again quite close to unity and therefore again implying constant shares with respect to income. The effect of non-AFDC expenditures, shown in the estimates for Model III, are completely insignificant, thus indicating no substitutability between the expenditures of the two programs.

Predictions back to 1960 using the coefficients for each model separately are shown in Table 6. The forecast errors in general are once again very small, especially so for the models including Medicaid. Whereas Models I and IA underpredict the 1960 AFDC benefit by $94 to $132, Models II, IIA, III, and IIIA overpredict by no more than $41. The best models for prediction are those which assume rational voters who take account of both Medicaid and the taxation of AFDC by the Food Stamp program (Models IIA and IIIA), models which predict the AFDC benefit to be only $9-$10 over its actual value! Moreover, the F-statistics reject Models I and IA most strongly and favor Models IIA and IIIA the most. The standard errors of the estimates are quite small for all equations. The 1984 equations are thus even more supportive of the full substitution hypothesis than those for 1960, especially for the full rationality models.

Estimates for Models III and IIIA, the most complete models, obtained by pooling the 1960 and 1984 data are shown in the last two columns of Table 5.[16] Price effects are negative and imply elasticities at the mean of about -12 percent, and income effects are once again positive and significant, implying income elasticities of about 1.06. An F-test for similarity of the coefficients overwhelmingly fails to reject the null of no structural change

Table 6
1984 Backcasts to 1960

	I	IA	II	IIA	III	IIIA
Predicted B[a]	371	333	506	474	502	475
Forecast error	-94	-132	41	9	37	10
F-statistic (p-value)	2.96 (.09)	12.47 (.00)	.35 (.56)	.03 (.87)	.26 (.61)	.03 (.86)
Standard error	54	37	70	54	72	56

Notes:

[a]Actual B = 465 in all cases.

(F-statistics of .39 and 1.2 in Models III and IIIA, respectively). Thus the use of all the information at hand again strongly supports the full substitution hypothesis.

Taking the pooled estimates for Model III as valid, the sources of the increase in the benefit sum from $465 to $770 between 1960 and 1984 can be deduced. Using the means for the variables indicates that the increase in price pushed down the benefit by $80, but that the increase in disposable income pushed it up by $364, the residual ($19) accounted for by the introduction of the non-AFDC Medicaid program. The relatively weak effect of the caseload explosion in the late 1960s and early 1970s in pushing down the benefit is consistent with the consistently weak price elasticities estimated in both the 1960 and 1984 data. Thus the bulk of the evidence clearly supports an interpretation that income increases over the past three decades have indeed increased the amount of transfers to female heads of families, and that the decline in the AFDC benefit is virtually entirely a result of the substitution of Food Stamps and Medicaid.

PLC Estimates. To test the sensitivity of the estimates to the instrumental variable procedure used to address the piecewise-linearity of the 1960 matching rate formula, the full maximum likelihood PLC model was estimated on the 1960 data and on the pooled 1960-1984 data for Models III and IIIA. The exact specification of the model is given in Appendix A, along with the definition of the log likelihood function.

The results are shown in Table 7. The first column shows the 1960 estimates, which appear to be moderately close to those in Table 3 (at least given some of the large standard errors in both equations), but are not quite

Table 7
Maximum Likelihood PLC Estimates

	1960	1960, 1984 Pooled Model III	1960, 1984 Pooled Model IIA
β	-5513.5*	-5267.7*	-3986.7*
	(2869.7)	(1660.1)	(1397.4)
γ^a	71.3*	70.7*	59.1*
	(22.0)	(9.7)	(8.2)
δ	--	813.6	-240.6
		(2688.7)	(2556.6)
α	132.2	131.3	174.8*
	(153.1)	(74.4)	(60.1)
σ_ε	123.1*	126.4*	114.3*
	(17.6)	(10.8)	(10.6)
σ_ν	10.4	10.5	9.4
	(13.0)	(13.3)	(14.9)
Log Likelihood	-297.6	-599.47	-590.09
Unrestricted Log Likelihood[b]	--	-599.45	-588.85

Notes: Standard errors in parentheses.

* Significant at 10-percent level.

[a]Multiplied by 1000.

[b]Allows separate coefficients for 1968 and 1984.

the same. The second and third columns of Table 7 show the estimates of Models III and IIIA, respectively, on the pooled 1960-1984 data. The estimates in this case are very close to those in the corresponding columns of Table 5 and, consequently, generate the same types of predictions back in time shown in Table 6. Moreover, tests for a change in structure between the years--obtainable from the unrestricted log likelihoods shown in the table--are strongly rejected (chi-squared statistics of .04 and 2.48 in the two equations, respectively). Thus the instrumental variable procedure used in the previous sections appears to have generated sufficiently good estimates to allow all the conclusions reached previously to be retained.

Effects of Federal Matching. Given the primary result of the analysis--that lump-sum federal transfers appear to generate dollar-for-dollar displacement by the states--it should be of interest to determine whether traditional federal matching for AFDC may be an alternative mechanism to increase transfers by the states. As is well known, the efficacy of matching in general depends upon the magnitude of price elasticities and, in related fashion, on the degree to which federal grants are substituted into other areas of expenditure by the states. The fraction of the AFDC grant substituted into other areas of expenditure is calculated as one minus the ratio of the stimulus (i.e., the effect of matching on the AFDC benefit) to the federal grant.

In a prior analysis (Moffitt, 1984), it was found that almost 80 percent of federal AFDC grants in 1970 were spent on non-AFDC purposes, implying that federal matching stimulates relatively little additional AFDC expenditure by the states. A similar calculation here implies that out-substitution was 85 percent in 1960 but only 38 percent in 1984.[17] The much lower out-substitution effect in 1984 is a result of two factors. First, the caseload tripled between 1960 and 1984; therefore, a given federal matching rate has a much greater stimulative effect on the AFDC benefit than it did previously. Second, the mean federal grant has fallen because the mean benefit has fallen, thereby again lowering the percentage substituted out. Thus it appears that, currently, almost two-thirds of the federal AFDC grant is indeed used to increase state AFDC expenditures, making this strategy considerably more attractive than using Food Stamps to increase the total transfer.[18]

V. EFFECTS ON THE AFDC-U PROGRAM

The AFDC-Unemployment Parent program is a supplement to the regular AFDC program and provides benefits to low-income families even if an able-bodied male parent is present in the household. Eligibility for such families is based not only upon the usual income and asset conditions in the regular AFDC program but also upon whether the principal earner is unemployed and has had a history of sufficiently strong attachment to the labor force. The AFDC-U program was enacted by Congress and made optional to the states in 1961, with matching set at the same rate as that for the regular AFDC program. Although the number of states adopting the program rose quickly in the early 1960s, it leveled off at around 50 percent in the late 1960s and has remained approximately at that level since.

The Food Stamp supplementation hypothesis explored in the last section would appear to be of relevance to the growth of AFDC-U as well. The Food Stamp program is notable in the U.S. transfer system for its provision of benefits to husband-wife families, unlike the regular AFDC program, and hence the Food Stamp program is closer to being a universal transfer program than any other in the United States. However, the provision of federally funded benefits to husband-wife families, the same group covered by the AFDC-U program, may, by a similar logic to that discussed for the regular AFDC program, have discouraged the adoption of AFDC-U program in the states. It should be noted that AFDC benefits to female-headed families and to husband-wife families are required to be the same for equivalent family sizes and other circumstances, so no leeway is possible in lowering the benefit in the AFDC-U program separately from that in the AFDC program; the only decision is to adopt or not to adopt.

A simple model that captures the major factors influencing the AFDC-U adoption decision is as follows.

Model IV

$$\text{Max} \quad U(B + \theta DB, Z), \tag{25}$$

$$\text{s.t.} \quad Y = (P_1 + DP_2)B + Z, \tag{26}$$

where B is the common benefit for AFDC and AFDC-U families, D is a dummy variable equal to one if the state adopts AFDC-U and 0 if not, and P_1 and P_2 are the prices for the AFDC and AFDC-U programs, respectively (caseload times own-payment share). The parameter θ measures the marginal utility of transfers to husband-wife couples. The solution to the maximization problem can be written as follows:

$$B = \frac{\alpha}{1+\theta} + \frac{\beta}{(1+\theta)^2} P_T + \frac{\gamma}{(1+\theta)} Y \quad \text{if} \quad D = 1, \tag{27}$$

$$B = \alpha + \beta P_1 + \gamma Y \qquad \text{if} \quad D = 0, \tag{28}$$

$$D = 1 \quad \text{if} \quad D^* > 0; \quad D = 0 \quad \text{if} \quad D^* < 0, \tag{29}$$

$$D^* = V[\frac{P_T}{1+\theta}, Y] - V[P_1, Y], \tag{30}$$

where $P_T = P_1 + P_2$ is the total price if AFDC-U is adopted (total per capita caseload times own-payment share) and $V[P,Y]$ is the indirect utility function.

Our interest here is less in the benefit equations than in the AFDC-U choice equation (30).[19] The interesting implication of (30) is that AFDC-U is chosen iff $[P_T/(1+\theta)] < P_1$, or

$$D = 1 \quad \text{iff} \quad \theta > (P_2/P_1). \tag{31}$$

Since P_1 and P_2 are both per capita caseloads times one minus identical matching rates, we can say that AFDC-U is adopted only if the marginal utility of transferring funds to husband-wife families is greater than the ratio of the husband-wife caseload to the regular female-head caseload. This result implies that increases in the regular AFDC caseload and decreases in the actual or potential AFDC-U caseload raise the probability of adopting AFDC-U, both of which are testable implications.

The criterion in (31) is notable for its implication that the income level of the state has no effect on the probability of adopting AFDC-U. This would seem contrary to any notion of AFDC-U, and transfers to husband-wife families in general, as normal goods. Clearly this property of (31) is a result of the specification of the utility function and the linear indifference curves between the benefits to female-headed families and husband-wife families, a restriction which may be violated. However, this restriction does make the model in (25)-(26) testable by simply determining whether income does or does not affect AFDC-U adoption probabilities independent of P_1 and P_2.[20]

Incorporating Food Stamps in a manner similar to that in the last section, we have

Model V

$$\text{Max} \quad U[B + \theta DB + \psi_1 F + \psi_2 \theta F, Z], \tag{32}$$

with the budget constraint as in (26). Note that Food Stamps are provided to husband-wife families regardless of whether the state has adopted an AFDC-U program. Under the null of $\psi_1 = \psi_2 = 1$, the demand equations are the following:

$$B + F = \frac{\alpha}{1+\theta} + \frac{\beta}{(1+\theta)^2} P_T + \frac{\gamma}{1+\theta} [Y + P_T F] \quad \text{if} \quad D = 1, \tag{33}$$

$$B + F = \alpha + \beta P_1 + \gamma [Y + P_1 (1+\theta) F] - \theta F \quad \text{if} \quad D = 0, \tag{34}$$

$$D^* = V[\frac{P_T}{1+\theta}, Y + P_T F] - V[P_1, Y + P_1 (1+\theta) F]. \tag{35}$$

Here the demand equations for benefits (33) and (34) imply that the Food Stamp program lowers the AFDC benefit in non-AFDC-U states more than in AFDC-U states, at least ignoring income effects, as a result of the term $- \theta F$ in (34). The source of this effect is simply that non-AFDC-U states are initially providing nothing to husband-wife families at all, and hence the introduction of Food Stamps has an overly strong substitution effect on the AFDC benefit.

More important for present purposes, equation (35) implies that, contrary to expectations, the AFDC-U adoption decision is actually independent of the Food Stamp benefit, aside from income effects. The latter are certain to be insignificant, as the increase in virtual income created by the Food Stamp program is about one-tenth of one percent of income. Thus the AFDC-U criterion function is essentially equivalent to that in the prior model, as given by (31)--only relative prices matter. The independence of the AFDC-U decision from the Food Stamp benefit again arises from the linear indifference curves in the particular utility function postulated, and is once again empirically testable.

A modification to incorporate Medicaid benefits leaves the model virtually unchanged.

Model VI

$$\text{Max} \quad U[(B + M)(1+\theta D) + F(1+\theta), Z], \tag{36}$$

$$\text{s.t.} \quad Y = (P_1 + DP_2)(B + QM) + Z, \tag{37}$$

under the null of full substitution of Medicaid and Food Stamps for AFDC. Note that the provision of Medicaid benefits is tied to the provision of AFDC-U, for husband-wife families are generally eligible for Medicaid only if they are AFDC recipients. The criterion function for AFDC-U adoption is the following:

$$D^* = V[\frac{P_T}{1+\theta}, Y + P_T(F+M) + P_T QM]$$

$$- V[P_1, Y + P_1(F(1+\theta) + M) - P_1 QM]. \qquad (38)$$

Ignoring income effects, which will be trivial, the criterion function for AFDC-U adoption is once again the same as that in (31).

Econometric Tests. To test the various hypotheses implied by the model, a before-and-after strategy similar to that in the last section is taken. The "pre" year is, in this case, taken to be 1968. The year 1960 is not appropriate because AFDC-U was not available at that date, and earlier dates are too close to 1961 to have allowed the slower-acting states to adopt the program if they so desired. The "post" year is again taken as 1984, the latest year for which all data are available. In 1968 44 percent of the states had adopted AFDC-U, and in 1984 48 percent had; thus there was virtually no change between the two dates.

At each year, probit equations for AFDC-U adoption are estimated using as independent variables the regular AFDC price variable (P_1), evaluated in 1968 on the mean segment, and proxies for the AFDC-U caseload. Since that caseload is observed only for states actually adopting the program, standard selectivity bias problems would arise if caseload data were directly used. Two proxy variables are employed instead: (1) the male unemployment rate in the state, assumed to be positively related to the actual or potential AFDC-U caseload, and (2) the real weekly manufacturing wage in the state (most workers in manufacturing are male), taken to be inversely related to the actual or potential AFDC-U caseload. Other variables that may be of interest (male wage rates, earnings, etc.) are not available. The means of the variables for the two years are shown in Appendix Table C-2.

Table 8 shows the results of the probit exercises. Column (1) shows a probit equation estimated on the 1968 data alone. The price of the regular AFDC caseload has a strong and significant positive effect on the AFDC-U probability, as predicted by the models. The magnitude of the coefficient implies that a 10-percent increase in the AFDC price increases the adoption probability by 9 percentage points at the mean. The manufacturing wage and the male unemployment rate both have the expected signs--increases in male wages and decreases in male unemployment make the cost of adoption lower--but both effects are weak statistically, largely a result of the small sample sizes. The state income variable is also included and, interestingly, is positive but insignificant, implying that income may have no effect independent of the prices of AFDC and AFDC-U benefits. Thus there is tentative support for the model in the 1968 data.

Columns (2) and (3) show results obtained by pooling the 1968 and 1984 observations.[21] Column (2) shows that the price and income variables retain their same signs and significance levels in the pooled data, but a dummy for 1984 is strongly significant and negative. A test for a change in the parameters other than the intercept from 1968 to 1984 was rejected, but the significant change in the intercept is consistent with a substitution effect of the Food Stamp program and hence inconsistent with the restrictive parameterization of the utility functions in Models IV-VI. The source of the downward shift in AFDC-U probabilities is primarily the strong increase in

Table 8
AFDCU Probit Regressions

	1968	Pooled 1968, 1984 (1)	Pooled 1968, 1984 (2)
P_1^a	.190*	.221*	.229*
	(.081)	(.065)	(.073)
Real Manufacturing Wage	.011	.008	.014*
	(.009)	(.006)	(.008)
Male Unemployment Rate	-.088	-.050	-.103
	(.259)	(.157)	(.159)
Y^b	.191	.319	.168
	(.287)	(.217)	(.231)
1984 Dummy	--	-2.160*	-1.759*
		(0.668)	(0.689)
Northeast Dummy	--	--	1.034*
			(0.584)
Intercept	-6.361	-6.789	-7.412
Log Likelihood	-22.34	-35.00	-33.17

Notes:
Standard errors in parentheses.
* Significant at 10 percent level.
[a]Variable multiplied by 1000.
[b]Variable divided by 1000.

the caseload over the 1970s, which, according to the strong and positive caseload effects, should have generated an increase in AFDC-U adoption from 1968 to 1984. The failure of the fraction of states adopting the program to grow past the late 1960s--about the time Food Stamps were introduced--is thus attributed in the regression to the intercept and hence to a structural shift downward.

Since AFDC-U is commonly strongly associated with the northeastern industrial states, which also have high caseloads, column (3) reports the results of adding a regional dummy for the Northeast. The addition of the dummy does reduce the magnitude of the downward structural shift, though it remains significant. However, the caseload effect is unchanged and the effects of manufacturing wages and male unemployment rates increase in both magnitude and significance, thus strengthening their support of the relevant hypotheses.

VI. SUMMARY AND IMPLICATIONS

This study has examined the causes of the decline in real AFDC benefits over the 1970s and 1980s and has focused on testing the hypothesis that states have allowed Food Stamps and Medicaid to substitute for AFDC in the total benefit package provided to female family heads in the United States. The analysis is conducted by forecasting benefits in the 1980s from cross-sectional regressions in the 1960s, predicting benefits back to the 1960s from cross-sectional regressions in the 1980s, and pooling data from both periods to test directly for structural change. The results support the strong version of the substitution hypothesis, according to which substitution occurs on a dollar-for-dollar basis. In other words, for every dollar granted in the form of Food Stamps or medical coverage, AFDC benefits were reduced by a dollar. The evidence is stronger in support of the hypothesis that both Food Stamps and Medicaid have substituted for AFDC than that Food Stamps alone has done so, for some tests reject the Food-Stamps-only model. Additional results suggest that the Food Stamp program has also slowed the adoption of the AFDC-U program.

There are several implications of the findings of the paper. First, the basic result--that the transfer package has increased over time in line with the growth of income--suggests that the benefit will continue to grow in the future, at least to the extent that income growth also maintains its past pattern. Thus there is no reason in these results to expect any decline in redistribution.

Second, the results imply that lump-sum transfer programs enacted by Congress have no effect on total transfers, and therefore on the net incomes, of low-income female family heads. Instead, they merely replace cash transfers with in-kind transfers. As a matter of perhaps naive political speculation, one may wonder why Congress has enacted such a program; the answer may be that there are stronger political lobbies (agriculture, hospitals) behind in-kind transfers than behind cash transfers. Be that as it may, the implication is that Congress simply does not have the ability to increase transfers, at least not in this fashion. Instead, it can only provide a large measure of budget relief to the states.

There are, of course, other policies available to Congress should it truly wish to increase the level of transfers. As discussed in the paper, heavier use of matching rates would provide nontrivial price incentives, at least on average. Alternatively, federalization of the AFDC program would directly eliminate the ability of states to counter federal transfer policy, or the establishment of a minimum benefit would constrain that ability. Optimal federal policy under these conditions should be a topic for future research.

APPENDIX A

LIKELIHOOD FUNCTION FOR THE PLC MODEL

The PLC model for 1960 is the following:

$$B_i^* = \alpha + \beta\hat{P}_{ji} + \gamma\hat{Y}_{ji} + \varepsilon_i, \tag{A1}$$

$$B_i = B_i^* + \nu_i, \tag{A2}$$

where B_i is the observed benefit, B_i^* is the "desired" (i.e., utility-maximizing) benefit, $\hat{P}_{ji}$ is the virtual price on segment j, $\hat{Y}_{ji}$ is virtual income on segment j, ε_i is heterogeneity error, and ν_i is "random" error. The error terms ε_i and ν_i are assumed to be normally and independently distributed with respective variances σ_ε^2 and σ_ν^2. Virtual price and income are equal to

$$\hat{P}_{1i} = (.17)(C_i/N_i), \tag{A3}$$

$$\hat{P}_{2i} = (1-s_i)(C_i/N_i), \tag{A4}$$

$$\hat{P}_{3i} = (C_i/N_i), \tag{A5}$$

$$\hat{Y}_{1i} = Y_i, \tag{A6}$$

$$\hat{Y}_{2i} = \hat{Y}_{1i} + (\hat{P}_{2i} - \hat{P}_{1i})\hat{B}_{1i}, \tag{A7}$$

$$\hat{Y}_{3i} = \hat{Y}_{2i} + (\hat{P}_{1i} - \hat{P}_{2i})\hat{B}_{2i}, \tag{A8}$$

where C_i and N_i are the state caseload and population, respectively, and $\hat{B}_{1i}$ and $\hat{B}_{2i}$ are the two kink points in the constraint. The log likelihood function to be maximized w.r.t. the parameters α, β, γ, σ_ε, and σ_ν is

$$L = \sum_i \log g(B_i), \tag{A9}$$

where g is the density function for B_i. When 1984 is included in the estimation, an extra term for a linear regression equation is added to (A9). The density function g is defined as the following:

$$\begin{aligned} g(B_i) = {} & \text{Prob}\ (\varepsilon_i \leqslant \hat{B}_{1i} - \alpha - \beta\hat{P}_{1i} - \gamma\hat{Y}_{1i},\ \varepsilon_i + \nu_i = B_i - \alpha - \beta\hat{P}_{1i} - \gamma\hat{Y}_{1i}) \\ & + \text{Prob}\ (\hat{B}_{1i} - \alpha - \beta\hat{P}_{1i} - \gamma\hat{Y}_{1i} < \varepsilon_i \leqslant \hat{B}_{1i} - \alpha - \beta\hat{P}_{2i} - \gamma\hat{Y}_{2i},\ \nu_i = B_i - \hat{B}_{1i}) \\ & + \text{Prob}\ (\hat{B}_{1i} - \alpha - \beta\hat{P}_{2i} - \gamma\hat{Y}_{2i} < \varepsilon_i \leqslant \hat{B}_{2i} - \alpha - \beta\hat{P}_{2i} - \gamma\hat{Y}_{2i}, \\ & \qquad \varepsilon_i + \nu_i = B_i - \alpha - \beta\hat{P}_{2i} - \gamma\hat{Y}_{2i}) \end{aligned}$$

$$+ \text{Prob}\ (\hat{B}_{2i} - \alpha - \beta\hat{P}_{2i} - \gamma\hat{Y}_{2i} < \varepsilon_i \leqslant \hat{B}_{2i} - \alpha - \beta\hat{P}_{3i} - \gamma\hat{Y}_{3i},$$

$$\nu_i = B_i - \hat{B}_{2i})$$

$$+ \text{Prob}\ (\varepsilon_i > \hat{B}_{2i} - \alpha - \beta\hat{P}_{3i} - \gamma\hat{Y}_{3i},\ \varepsilon_i + \nu_i = B_i - \alpha - \beta\hat{P}_{3i} - \gamma\hat{Y}_{3i}),$$

$$= \frac{1}{\sigma_\omega} f(z_{1i})F(r_{1i}) + \frac{1}{\sigma_\nu} f(u_{1i})[F(e_{12i}) - F(e_{11i})]$$

$$+ \frac{1}{\sigma_\omega} f(z_{2i})[F(r_{2bi}) - F(r_{2ai})] + \frac{1}{\sigma_\nu} f(u_{2i})[F(e_{23i}) - F(e_{22i})]$$

$$+ \frac{1}{\sigma_\omega} f(z_{3i})[1 - F(r_{3i})], \tag{A10}$$

where f and F are the unit normal density and distribution functions, respectively, and

$$z_{1i} = (B_i - \alpha - \beta\hat{P}_{1i} - \gamma\hat{Y}_{1i})/\sigma_\omega, \tag{A11}$$

$$z_{2i} = (B_i - \alpha - \beta\hat{P}_{2i} - \gamma\hat{Y}_{2i})/\sigma_\omega, \tag{A12}$$

$$z_{3i} = (B_i - \alpha - \beta\hat{P}_{3i} - \gamma\hat{Y}_{3i})/\sigma_\omega, \tag{A13}$$

$$u_{1i} = (B_i - \hat{B}_{1i})/\sigma_\nu, \tag{A14}$$

$$u_{2i} = (B_i - \hat{B}_{2i})/\sigma_\nu, \tag{A15}$$

$$e_{11i} = (\hat{B}_{1i} - \alpha - \beta\hat{P}_{1i} - \gamma\hat{Y}_{1i})/\sigma_\varepsilon, \tag{A16}$$

$$e_{12i} = (\hat{B}_{1i} - \alpha - \beta\hat{P}_{2i} - \gamma\hat{Y}_{2i})/\sigma_\varepsilon, \tag{A17}$$

$$e_{22i} = (\hat{B}_{2i} - \alpha - \beta\hat{P}_{2i} - \gamma\hat{Y}_{2i})/\sigma_\varepsilon, \tag{A18}$$

$$e_{23i} = (\hat{B}_{2i} - \alpha - \beta\hat{P}_{3i} - \gamma\hat{Y}_{3i})/\sigma_\varepsilon, \tag{A19}$$

$$r_{1i} = (e_{11i} - \rho z_{1i})/\sqrt{1-\rho^2}, \tag{A20}$$

$$r_{2ai} = (e_{12i} - \rho z_{2i})/\sqrt{1-\rho^2}, \tag{A21}$$

$$r_{2bi} = (e_{22i} - \rho z_{2i})/\sqrt{1-\rho^2}, \tag{A22}$$

$$r_{3i} = (e_{23i} - \rho z_{3i})/\sqrt{1-\rho^2}, \tag{A23}$$

$$\sigma_\omega = (\sigma_\varepsilon^2 + \sigma_\nu^2)^{1/2}, \tag{A24}$$

$$\rho = \sigma_\varepsilon/\sigma_\omega. \tag{A25}$$

APPENDIX B

DATA SOURCES

B: Real AFDC Guarantee for a Family of Four. 1960: U.S. House of Representatives, Committee on Ways and Means, Background Material and Data on Programs within the Jurisdiction of the Committee on Ways and Means, 1987 Edition, pp. 660-662. 1984: Unpublished data, Office of Family Assistance, Department of Health and Human Services. Nominal guarantees deflated by the personal consumption expenditure deflator (PCE) for the GNP accounts.

F: Real Food Stamp Guarantee for a Family of Four. 1984: U.S. House of Representatives, Committee on Ways and Means, Background Material ..., p. 504. Deflated by the PCE.

M: Real Medicaid Insurance Value. 1984: Calculated by multiplying the average Medicaid expenditure for an AFDC family of four by the ratio of AFDC families having Medicaid expenditures to all AFDC families. The average Medicaid expenditure for an AFDC family of four is calculated by summing the average Medicaid expenditure per AFDC adult and the average Medicaid expenditure per AFDC dependent times three. Medicaid expenditures obtained from unpublished data of the Health Care Financing Administration; AFDC data source described below. Deflated by Q (see below).

C: AFDC and AFDC-U Caseloads. 1960: Number of families and number of recipients from Social Security Administration, Social Security Bulletin, November 1960, p. 55. 1968: National Center for Social Statistics, Public Assistance Statistics, Series A-2, July 1968, Tables 6 and 7. 1984: Office of Research and Statistics, Quarterly Public Assistance Statistics, July-September 1984, Tables 1 and 8.

N: Population. 1960, 1968: Bureau of Economic Analysis, State Personal Income: 1929-1982, Table 8. 1984: Survey of Current Business, August 1985, Table 3.

s: Matching Rate. 1960: unpublished data provided by Larry Orr. 1984: Social Security Administration, Annual Statistical Supplement of the Social Security Bulletin, 1983.

Y: Real Disposable Income per Capita. Real per capita income in 1960, 1968, 1984: Same as population. 1960, 1968: Federal taxes obtained from Bureau of the Census, Statistical Abstract of the United States, 1964 and 1971 Editions. 1984: Federal taxes obtained from Internal Revenue Service, SOI Bulletin, Fall 1986. Taxes divided by population. Deflated by PCE.

P_N: Price of Non-AFDC Medicaid Benefit. Caseload of non-AFDC Medicaid recipients obtained from unpublished HCFA data; population as referenced above; matching rate as referenced above.

Q: Price of Medical Care. State-specific medical care index for 1980 obtained from T. Grannemann and M. Pauly (1983, pp. 109-110). Converted to 1984 value using the medical care component of the CPI.

Real Manufacturing Wage. 1968: Bureau of Labor Statistics, Handbook, 1974, Table 105. 1984: 1.01 times the 1983 value in BLS Handbook, 1985, Table 87. Deflated by PCE.

Male Unemployment Rate. 1968: 1970 value taken from Bureau of the Census, CPR Series P-20, No. 334, Demographic, Social, and Economic Profile of States: 1976, Table 31. 1984: U.S. Bureau of the Census, State and Metropolitan Area Data Book, 1982 Edition, Table C.

Appendix Table C-1
Means of Other Variables
Used in Main Analysis

	1960	1984
NE	.19	.19
NC	.25	.25
W	.23	.23
PCHS	41.40	66.91
METPCT	.61	.61
UN	5.41	7.32
RW	265.15	323.06

Appendix Table C-2
Means of Variables in the
AFDC-U Analysis

	1968	1984
D	.44	.48
P_1[a]	8.58	17.6
Real Manufacturing Wage	303.03	323.06
Male Unemployment Rate	3.91	6.46
Y[b]	7.28	10.19
Northeast	.19	.19

Notes:

n = 48

[a]Multiplied by 1000.

[b]Divided by 1000

NOTES

1. An additional question is whether the changes in the caseload are explainable by the changes in the benefit. Although an interesting question, it will not be examined here. However, some research has indicated that cross-sectional AFDC participation equations do not provide sufficiently large benefit elasticities to explain more than a small fraction of the tripling of the caseload. Put differently, there seems to have been a structural change in the AFDC participation equation over the period 1966-1972.

2. The AFDC benefit per family shown in Figure 1 peaked earlier because there was a steady decline in the mean AFDC family size over the period. The guarantee is the appropriate variable to examine.

3. These considerations suggest quite naturally that a state may lump all income-maintenance-related expenditures into a single pot, which constitutes a single argument in the utility function. The state may thus be concerned in setting only the share of its budget that is devoted to income maintenance in general.

4. The Food Stamp benefit is also affected by the AFDC benefit itself, but Orr used an instrumental-variables technique to remove this source of endogeneity.

5. It appears that only time-series variation was utilized because Gramlich (1982) regressed the AFDC benefit (or a weighted sum of AFDC and the Food Stamp benefit) on the Food Stamp benefit and a set of state dummies, using a cross-section of states from 1974 to 1981. The inclusion of

state dummies sweeps out any cross-sectional variation in Food Stamps, leaving only time-series variation.

6. It is assumed that federal taxes are exogenous to the median state voter, being set instead by the median U.S. voter.

7. Equation (12) assumes that AFDC and Medicaid benefits are matched at the same rate, as will be the case for the years used in the analysis.

8. The matching rate is the same in both portions.

9. One could also test nulls on the substitutability of non-AFDC Medicaid expenditures and AFDC Medicaid expenditures, but this is not directly germane to the question of interest here. In addition, of course, perfect substitutability does not correspond to a value of 1 for the utility parameter δ.

10. Although it would be desirable to improve the efficiency of the estimates by pooling a set of years either before or after the transition period, this is not practical. The AFDC guarantee is next available in 1964, then in 1968, both years uncomfortably close to the explosive transformations of the 1960s; and the years in the 1980s earlier than 1984 are less assuredly post-adjustment years.

11. Alaska and Hawaii are excluded because some of the variables are not present for them in 1960. Arizona is also excluded because it has no Medicaid program. The District of Columbia is included.

12. The caseload variable is lagged three years to avoid potential endogeneity. However, from past work (Moffitt, 1984) it appears that there is little difference in the estimates when the current caseload is used.

13. The benefit schedule in 1984 was linear, for the states were then employing the open-ended Medicaid matching schedule instead of the nonlinear, convex "federal" matching schedule. The Medicaid schedule created a nonconvex kink in the overall budget set. Another reason for avoiding the late 1960s and 1970s for estimation is that those years were a period of gradual switching from one schedule to the other, and states often did not switch immediately when it was advantageous for them to do so. See Orr (1976, 1978) and Moffitt (1984) for discussions.

14. Forecasts for the estimates in columns (2) and (3) of that table were also calculated. Their point estimates were very close to those for column (1), but their standard errors of forecast were much higher, no doubt because of the larger number of insignificant coefficients in those regressions.

15. Since M is included in virtual income, there is some endogeneity in the variable. However, the Food Stamp and Medicaid terms in the virtual income variable are minuscule in magnitude, constituting less than one-tenth of one percent of income. Consequently, the coefficients of the equations are virtually identical when they are left out of virtual income.

16. To test for random effects, a sample estimate of the variance of a permanent state effect was calculated for Model III. The estimated variance was less than zero (negative variances are possible in such ANOVA calculations), from which it was concluded that the data from the two years could be safely pooled and the equations estimated with OLS.

17. This calculation uses the coefficients in Model III in Table 5, but the mean benefits, matching rates, and so on for 1960 and 1984 separately.

18. The price elasticity in the equation is -.13 as compared to the -.08 elasticity in Moffitt (1984). This is responsible for part of the difference between the results in the earlier study and those here, but not the major part.

19. Partly for this reason, only the AFDC-U choice equation will be estimated below, and not the benefit equations. But, in addition, there are too few observations to estimate separate benefit equations on AFDC-U and non-AFDC-U states.

20. Allowing the transfer to husband-wife families to constitute a separate argument in the utility function--e.g., U(B,B,Z)--makes the derivation of demand functions for B and properties of D* intractable, the difficulty arising because the first two arguments are constrained to equality.

21. Given the evidence of weak state-specific effects obtained in the last section, no error components models were estimated.

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