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Abstract

This paper analyzes the relationship between AFDC benefits and a single mother's propensity to reside in a subfamily--i.e., within another family rather than in her own independent household. We find that some states effectively penalize mothers for living in subfamilies. A single mother can lose a substantial amount of AFDC benefits if she chooses to reside in a subfamily rather than establish her own household. Using data from the 1984 Current Population Survey, we address the question of whether differences in AFDC benefits affect the probability that a mother will reside in a subfamily. We find that the subfamily penalty has discernible but small effects, and that the overall level of AFDC benefits has no effect.

AFDC and the Formation of Subfamilies

What effect does the Aid to Families with Dependent Children program (AFDC) have upon a young mother's living arrangements? In a sense this is an old question. There already exists an extensive literature examining the effect of AFDC benefits upon a young mother's propensity to live with a husband.[1] In another sense, however, this is a new question. Recent work by Ellwood and Bane (1985) raises the issue of whether AFDC benefits influence a young mother's propensity to live with her parents or other relatives. Their work suggests that higher AFDC benefits lead young mothers to establish their own households rather than live as a subfamily in a larger household.

It is important to learn more about this phenomenon. If the AFDC program tends to discourage mothers from residing in subfamilies, there could be implications for both the mothers and their children. A mother who lives in a subfamily may have more hands to help with child care. That could mean a better environment for children and more time for the mother to pursue schooling or work-related training. If the AFDC program discourages formation of subfamilies, it may end up contributing to the poverty problem that it seeks to alleviate.

This paper examines the effect of AFDC benefits upon a young mother's propensity to reside in a subfamily. There are two prongs to our attack. First, in Section I we analyze how state AFDC programs differ in their treatment of subfamilies versus householders. The available literature is surprisingly silent on this matter. Using information from both a telephone survey of state welfare agencies and caseload data, we find that state AFDC programs differ substantially in their treatment of the different living arrangements. In some states a mother can lose a substantial amount of AFDC benefits if she chooses to reside in a subfamily.

Second, in Section II we analyze 1984 Current Population Survey data in order to determine if differences in the level of AFDC benefits affect the probability that a mother resides in a subfamily. Our specification includes two measures of the AFDC benefits available to a mother -- one if she is a householder and one if she resides in a subfamily. In general, we find that the AFDC program has discernible, but small, effects on a young mother's propensity to reside in a subfamily.

I. INTERSTATE VARIATION IN AFDC BENEFITS PAID TO SUBFAMILIES

If a mother and her child live with "other adults" (usually her parents), some states will pay them smaller AFDC benefits than if they lived in their own household. This is justified on the grounds that the grandparents may provide economic assistance to the mother and child. In practice, such economic assistance is usually in kind and takes the form of room and board. Since there are no specific federal guidelines governing the treatment of this form of "outside" assistance, each state decides how to adjust AFDC benefits. Between May and July 1985 we contacted the relevant administrative agency for each of the 48 states in the continental U.S. plus the District of Columbia. In each case we posed the following questions:

1. What was the maximum payment for an adult mother, living independently, with no non-AFDC income and a single child under 3 years old?

Table 1

Variation in AFDC Benefits across Living Arrangements

		Two-Person Maximum When Mother Lives with Nonpoor Parents[a]			Coefficient on SHARE Variable in Regression Using 1982 Quality Control Data[c] (t-statistic)
	Two-Person Maximum (1)	Free Room and Board (2)	Nominal Payment for Room & Board (3)	Policy Type[b] (4)	(5)
Alabama	$88	$88	$88	1	-5.05 (-1.5)
Alaska	-	-	-	-	-2.75 (- .2)
Arizona	180	141	180	2	-6.45 (- .8)
Arkansas	135	135	135	1	-.37 (- .8)
California	448	283	448	2	-3.19 (- .6)
Colorado	272	198	272	2	3.39 (.5)
Connecticut	440	440	440	1	7.52 (1.6)
Delaware	212	212	212	1	.0 (.0)
D.C.	257	257	257	1	2.56 (1.1)
Florida	185	126	185	2	-2.38 (- .9)
Georgia	174	174	174	1	-.51 (-1.0)
Hawaii	-	-	-	-	-25.68 (-1.3)
Idaho	245	123	123	3	-28.27 (- .8)
Illinois	250	250	250	1	-.39 (- .1)
Indiana	196	132	132	3	-63.64 (-4.6)
Iowa	305	305	305	1	6.46 (.4)
Kansas	288	233	233	3	-22.91 (-1.4)
Kentucky	170	170	170	1	-.40 (- .9)
Louisiana	138	138	138	1	-2.19 (-1.3)
Maine	275	275	275	1	-5.54 (- .8)
Maryland	244	244	244	1	-1.25 (-1.1)
Massachusetts	328	222	328	2	-5.97 (- .4)
Michigan	416	348	416	3	-82.79 (-4.2)
Minnesota	431	431	431	1	0.55 (.0)
Mississippi	96	96	96	1	-3.10 (- .6)
Missouri	211	211	211	1	-0.31 (- .2)
Montana	279	123	123	3	-0.0 (- .5)
Nebraska	280	280	280	1	-3.86 (- .5)
Nevada	187	187	187	1	0.0 (0)
New Hampshire	320	183	183	3	20.51 (.9)
New Jersey	292	292	292	1	-4.88 (-1.3)
New Mexico	210	122	210	2	-13.96 (-1.6)
New York	486	150	150	3	-69.70 (-4.5)
North Carolina	194	194	194	1	0.34 (1.2)
North Dakota	301	226	301	2	8.77 (0.3)

Table 1, continued

	Two-Person Maximum When Mother Lives with Nonpoor Parents[a]				Coefficient on SHARE Variable in Regression Using 1982 Quality Control Data[c] (t-statistic)
	Two-Person Maximum (1)	Free Room and Board (2)	Nominal Payment for Room & Board (3)	Policy Type[b] (4)	(5)
Ohio	238	238	238	1	3.22 (1.9)
Oklahoma	218	218	218	1	-0.96 (-.4)
Oregon	328	230	230	3	2.96 (2.2)
Pennsylvania	285	285	285	1	1.60 (0.4)
Puerto Rico	-	-	-	-	-15.14 (-2.7)
Rhode Island	350	350	350	1	0.05 (0.5)
South Carolina	144	144	144	1	0.84 (1.1)
South Dakota	286	123	123	3	-96.71 (-2.8)
Tennessee	108	108	108	1	-1.27 (-0.6)
Texas	144	144	144	1	0.67 (-1.0)
Utah	301	301	301	1	-1.74 (-0.6)
Vermont	438	182	295	3	-122.05 (-3.2)
Virginia	272	231	272	2	-1.60 (-0.3)
Washington	385	263	385	2	4.96 (1.0)
West Virginia	164	98	164	2	-11.29 (-1.7)
Wisconsin	453	453	453	1	-6.55 (-0.5)
Wyoming	320	205	320	2	-57.50 (-3.9)

[a]Data from 1985 telephone survey.

[b]Type 1 = benefit not affected by receipt of room and board; Type 2 = benefit substantially reduced if room and board received free, but not reduced if any payment made; Type 3 = benefit substantially reduced if room and board received free, but reduction tailored to amount paid, if any.

[c]Negative values indicate penalty for living in a subfamily. See text for complete explanation.

2. Suppose the mother and child moved in with the mother's own parents who had substantial (e.g., $20,000 per year) income. How would her payment change:

   (a) If she received free room and board from her parents?

   (b) If she paid a nominal (e.g., $1 per month) amount for room and board?

   (c) If she paid a nontrivial amount (e.g., $100 per month) for room and board?

Columns (1)-(3) of Table 1 summarize the responses. The table reveals that, in some states, there is substantial variation in benefits across living arrangements, while in others there is very little.[2]

The simplest cases were those in which a recipient's payment was unaffected by in-kind income in the form of room and board. For example, Alabama paid our prototypical two-person unit $88 per month regardless of whether it received free room and board.

By contrast, in New Hampshire the two-person maximum payment was $320, composed of a $183 basic maintenance allowance and a $137 shelter allowance. If the recipient mother received free room and board, she no longer received the $137 shelter allowance and her payment dropped to $183. If she paid for room and board, her payment was adjusted upward, dollar for dollar, to the maximum of $320. In Kansas, the fact that a recipient shared a household implied a reduction of about $55 per month from the maximum. The reduction occurred regardless of any payments which might be made for room and board.

A different policy is illustrated by Colorado. There, the payment to a two-person family was reduced by 27 percent if the family received <u>free</u> room and board. As was the case in several other states, however, no distinction was made between nominal, but positive, amounts paid for room and board. Even if the family paid only $5 per month, the 27 percent reduction was restored. Obviously, there was considerable incentive to report a positive amount for room and board payments.

To summarize, the states can be divided into three categories according to how they treated the case of the prototypical family which lives with other nonpoor adults. These are the "policy types" in column 4 of Table 1.

Type 1. Some states, like Alabama, ignored this form of income entirely, so that AFDC benefits were unaffected by the receipt of room and board. There were 28 such states, constituting 52 percent of the national caseload:

Alabama	Arkansas	Connecticut	Delaware
D.C.	Georgia	Illinois	Iowa
Kentucky	Louisiana	Maine	Maryland
Minnesota	Mississippi	Missouri	Nebraska
Nevada	New Jersey	North Carolina	Ohio
Oklahoma	Pennsylvania	Rhode Island	South Carolina
Tennessee	Texas	Utah	Wisconsin

Type 2. A second group of states, similar to Colorado, considered contributions of room and board in a very lenient way. If the woman reported receiving free room and board, her AFDC payment was substantially reduced. However, if she reported paying a nominal amount, her payment was restored to the two-person maximum. The 11 states that did this (representing 28 percent of the national caseload) were

Arizona	California	Colorado
Florida	Massachusetts	New Mexico
North Dakota	Washington	West Virginia
Wyoming	Virginia	

Type 3. The last group of states, similar to New Hampshire, initially assumed that when a woman lived with "other adults" she received free room and board. If this were so, her AFDC payment was substantially reduced. If the woman provided evidence that she paid for room and board, the actual amount paid was usually considered in determining her payment. The 10 states which treat in-kind income in this fashion were

Idaho	Indiana	Kansas	Michigan[3]
Montana	New Hampshire	New York	Oregon
South Dakota	Vermont		

To examine whether actual (as opposed to prototypical) benefits vary across living arrangements within a state, we estimated regression models on data from the May 1982 AFDC Quality Control (QC) Survey. These data are collected by the federal government for purposes of checking the accuracy of state AFDC payment computations. A potential problem is small sample sizes, which ranged from 23 in South Dakota to 214 in Pennsylvania. We took the recipient mother and her children as our unit of observation, and estimated the following model separately for each state plus the District of Columbia and Puerto Rico:

$$PAYSTD_i = B_0 + B_1 SHARE_i + B_2 MAR_i + B_3 MXBEN_i + e_i ,$$

where

$PAYSTD_i$ is the monthly AFDC payment standard for the case containing the ith recipient mother. The payment standard is essentially the amount of money the case would receive if it had no other source of income (the AFDC guarantee). It is determined through a complex process that includes an assessment of the family's needs along with consideration of state maximums and percentage reductions.[4]

$SHARE_i$ is a binary variable which equals 1 when the recipient mother shares a household with an adult who could not be classified as a "husband." Otherwise, SHARE equals zero.

MAR_i is a binary variable which equals 1 when the recipient mother shares a household with a male who can reasonably be classified as a "husband" (i.e., a male who was father to the children, natural or adoptive; stepfather; nonrelative male; or unknown male). Otherwise, MAR equals zero.

$MXBEN_i$ is the "published" maximum benefit in the state for a case with the same number of children as live with the ith mother.[5]

Of course, we were primarily interested in the coefficients on the SHARE variable. Negative coefficients indicate states where mothers are effectively penalized for living with parents or relatives. Column 5 in Table 1 lists these coefficients for each of the jurisdictions. We find that in most states it is difficult to reject the null hypothesis of a zero coefficient. Thus, in most states the AFDC guarantee for a mother who lives in a subfamily is essentially the same as that for a mother who establishes her own household. There are, however, some states that unambiguously pay lower benefits to recipients who live with parents or relatives -- namely, Indiana, Michigan, New York, South Dakota, Vermont, and Wyoming.

Finally, and perhaps most important, we can compare these estimates with the data from the telephone survey. Of course, the two data sets are not strictly comparable; one reflects the responses of a (perhaps harried) bureaucrat in 1985, while the other is based on caseload records for a small sample of 1982 AFDC recipients. On average, however, it is reasonable to expect a degree of correspondence between the two data sets. Above we distinguish between three types of states: (1) states that ignore any in-kind income (e.g., room and board), (2) states that count in-kind income as a resource only if the recipient obtains free room and board, and (3) states that count in-kind income as a resource and reduce payments accordingly. The average unweighted value of the coefficient on SHARE for these three types of states is as follows:

Type of State	Average Value of SHARE Coefficient
1	-0.57
2	-7.75
3	-46.26

Thus, the survey data and the regression results tell the same general story;[6] in the third type of state, mothers who live with parents or other relatives obtain smaller AFDC benefits than mothers who establish separate households.[7] In these states there are clear incentives to establish one's own household.

To conclude, in some states a mother can lose a substantial amount of AFDC benefits if she chooses to live with her parents or other relatives. In those states a contribution of room and board can trigger a reduction in the mother's benefits. One might expect such policies to have behavioral effects. The next section examines that issue.

II. THE EFFECT OF AFDC BENEFITS ON A MOTHER'S PROPENSITY TO RESIDE IN A SUBFAMILY

Theoretical Framework

Consider a single woman with children facing the problem of choosing a living arrangement. She may decide either to live independently as a householder, live with other adults (typically her parents) as a subfamily head, or live with a husband as a wife. In addition, she may decide to seek and receive public assistance. Focusing on the unmarried women,[8] we can organize these alternatives into four living arrangements, denoted as LA = 1, 2, 3, and 4, and defined as follows:

Definition of Living Arrangements Variable, LA

Household Status	Public Assistance Recipient: No	Public Assistance Recipient: Yes
Householder (H)	LA = 1	LA = 3
Subfamily Head (S)	LA = 2	LA = 4

Assume a woman has a utility function of the form

$$(1) \qquad U = U(C,L,X),$$

where C is consumption, L is leisure, and X is a vector of characteristics which affect preferences ("taste shifters"). The problem is to maximize utility subject to a budget constraint which varies over the four living arrangements. The general form of the budget constraint for living arrangement j is

$$(2) \qquad PC_j + W_jL_j < N_j + A_j + W_jL_j^*, \; j = 1,\ldots,4,$$

where the price of consumption P does not vary over living arrangements, and the wage rate W varies with public assistance status. The net (of income and payroll tax) wage rate is W_1 when the mother does not receive public assistance and $W_1(1-t)$ when she does, where t is the rate at which the public assistance program taxes earnings. The public assistance payment A is zero when she does not receive welfare, but varies with household status when she does. Nonlabor nonwelfare income N varies with household status. A woman living as a subfamily head may obtain resources from the primary family, either in cash or in kind (e.g., free room and board).[9] Further, nonlabor nonwelfare income varies with public assistance status since these programs typically tax such income. Finally, the time available to allocate to either the market or to leisure, L^*, may vary with household status (though not with public assistance status). The presence of other adults in the household may reduce the time which the woman is constrained to commit to activities such as child care.

Let $T_k = (P, W_k, A_k, N_k, L_k^*)$ denote the vector characterizing the budget constraint in living arrangement k. Then the indirect utility function corresponding to maximizing (1) subject to (2) can be written

(3) $\quad v = v(T \mid X)$

and the utility maximization problem is now

(4) choose k to maximize $v(T_k \mid X)$, $k = 1,2,3,4$.

We assume that the woman assesses the utility attainable in each living arrangement, taking as constant her current demographic characteristics (X) (e.g., the same children will live with her whether she is a subfamily head or a householder, and she will live in the same region of the country). As in Becker, Landes, and Michael (1977) and Danziger et al. (1982), we assume that the world is, at least on average, in equilibrium.[10]

Following the literature on discrete choice (McFadden 1981, for example), we assume that the indirect utility function for woman i in living arrangement k can be expressed as the sum of a representative component which depends on T_k and an idiosyncratic component:

(5) $\quad v_{ik} = v(T_{ik} \mid X_i) = v^*(T_{ik} \mid X_i) + e_{ik}.$

The distribution of the idiosyncratic taste variation e_{ik} will determine the probability that woman i attains the highest utility level in living arrangement k (conditional on T_1, T_2, T_3, T_4, and X).

Any exogenous change which increases the utility attainable in living arrangement k will increase the probability that the woman chooses that living arrangement. Of primary importance here is the effect of changes in the level of AFDC benefits. Let AFDCH denote the level of benefits available to a householder and AFDCS denote the benefits available to a subfamily head. Then the variable AFDCDIF = AFDCH - AFDCS measures the "wedge" -- the "subfamily penalty" -- between the benefits available to a woman in the two household statuses.

How do changes in AFDCDIF and AFDCH affect the probability that a young mother chooses living arrangement 1, 2, 3, or 4? An increase in the wedge (AFDCDIF), holding constant the overall level of generosity of the program (AFDCH), will decrease the utility attainable in the fourth living arrangement (subfamily head receiving public assistance) without affecting the utility attainable in the other three living arrangements. Thus, we would expect an increase in AFDCDIF to reduce the probability that a woman chooses LA = 4 and increase the probability that she chooses one of the other living arrangements. In particular, we would expect an increase in AFDCDIF to increase the probability that the woman is a householder (LA = 1 or 3).

Note that our hypothesis extends beyond what can be termed a "break-even" effect.[11] An increase in AFDCDIF, holding AFDCH fixed, implies a decrease in the subfamily guarantee, AFDCS. With program tax rates held constant, a lower guarantee implies a lower break-even, i.e., a lower level of income at which people must leave the program. Thus, if behavior did not change, an increase in AFDCDIF would imply a lower probability that a woman is a subfamily head receiving AFDC (LA = 4) and a higher probability that she is a subfamily head with no public assistance (LA = 2). This break-even effect is purely mechanical and says nothing about behavior.

Our hypothesis extends beyond that. Not only do we predict that an increase in AFDCDIF decreases the probability that LA = 4 and increases the probability that LA = 2, in addition we predict an increase in the probability that the woman is a householder (LA = 1 or 3). If all of the predictions are correct, then we are not merely observing break-even effects but also behavioral effects.

Finally, consider the effect of increasing the level of generosity of the program (AFDCH), holding constant the differential treatment of the two types of households (AFDCDIF). Since this increase raises the utility attainable by both householders and subfamilies receiving public assistance (LA = 3 and 4), without further assumptions we cannot predict how this change affects the choice between householder and subfamily status.[12]

Estimation

Our unit of analysis is a woman under 36 with at least one child who is under 19.[13] The data consist of 1599 single mothers (mothers who are not living with husbands) drawn from the March 1984 Current Population Survey (CPS), which contains data on AFDC income in calendar 1983. As noted above, we require a large cross-section data set in order to observe enough subfamilies to do meaningful empirical analysis. We use a recent CPS because prior to 1983 the CPS incorrectly coded subfamily status (Bane and Ellwood, 1983). Of these 1599 women, 608 were householders not receiving welfare, 351 were subfamily heads not receiving welfare, 464 were householders receiving welfare, and 176 were subfamily heads receiving welfare.

Using these data we estimated reduced-form models of living arrangement choice. We focused on the reduced form for two reasons. First, it answers the questions posed here: it investigates the relationship between AFDC benefits and choices among living arrangements. Second, the structural model implied by our theoretical framework requires information on the wage rate and level of nonwage income in each living arrangement for each woman in the sample. We do not have such data. For example, we do not know how much nonwage income a female head would receive were she to become a subfamily head. Instead of predicting such variables using information on the woman's age, education, race, etc., we simply estimate the reduced form.[14] Therefore, a variable like age not only controls for "tastes," but also for age-related variation in wage and nonwage income.

We use the following CPS variables in the analysis:

AGE	age, measured in years
EDUC	education, measured in years
AGESQ	age squared
EDUCSQ	education squared
AGEED	age-education interaction
DEPENDS	number of family members under 19 years of age
PRESCH	1 if there is a child aged 0-5; 0 otherwise
NWHITE	1 if woman's race is not white; 0 otherwise
SMSA	1 if the woman lives in an SMSA; 0 otherwise
SOUTH	1 if the woman lives in the South; 0 otherwise

The 1983 unemployment rate in the woman's state of residence (UNEMP), as reported in the Statistical Abstract of the United States, was also included. To measure the AFDC guarantee for householders (AFDCH) and subfamily heads (AFDCS), we estimated regression models of the AFDC monthly payment standard for each state using the 1982 Quality Control data. The regressions were similar to those described in Section I.[15] Taking these regressions as well as information on a woman's state of residence and number of children, we predicted the monthly values of AFDCH and AFDCS (denominated in thousands of dollars) for each woman in the sample.[16]

We did not use data from the telephone survey for this purpose, for three reasons. First, the telephone survey results pertain to a two-person AFDC family while the regression results encompass all family sizes. Second, it is well known that there can be a difference between stated and actual rules in the AFDC program (for example, the difference between statutory and effective tax rates on earnings; Hutchens, 1978). We view the regression results as a closer approximation to the "effective" rules of the program than the survey results. Finally, our data on living arrangements and AFDC recipiency come from the 1984 CPS and refer to 1983 income. The telephone survey results are from 1985. The regression results allow us to make predictions for 1983 on the basis of the year's published benefit schedule, while the survey results do not.

Results from Binary Logit Models

Although a multinomial logit model yields the most complete test of our hypotheses, for the sake of clarity we begin with a simple binary logit of the probability that an unmarried mother will be a subfamily head.[17] The model in the first column of Table 2 presents the key result. Holding demographic variables and the overall level of benefits (AFDCH) constant, an increase in the difference between the householder and subfamily guarantees (AFDCDIF) is associated with a decrease in the probability that the woman heads a subfamily.[18] The negative coefficient on AFDCDIF is statistically significant at the .05 confidence level.

This is consistent with the hypothesized behavioral effect of AFDCDIF. The single mothers in our sample are either subfamily heads or householders. Thus, the result indicates that holding the householder guarantee (AFDCH) and the other exogenous variables constant, an increase in AFDCDIF (which implies a decrease in the subfamily guarantee, since AFDCH is held constant) results in fewer mothers who are subfamily heads and more mothers who are householders.

Turning to the other coefficients in the first column of Table 2, we note that the coefficient on AFDCH (which indicates the overall level of AFDC benefits) is negative and not statistically different from zero. This weak result is not due to collinearity between AFDCH and AFDCDIF; as indicated in the second column of Table 2, the same result obtains when AFDCDIF is excluded from the model. Thus, our data do not support the claim that a higher overall level of AFDC benefits leads to more householders and fewer subfamilies. In particular, our column 2 result differs from that in Ellwood and Bane (1985). They find that for a single mother with a predicted probability of welfare receipt of .5, a $100 increase in the overall level of monthly AFDC benefits increases the probability that the mother establishes her own household by 15 to 20 percent. The Table 2, column 2, model predicts an increase of 5 percent for

Table 2

Binary Logit Estimates of Probability that Single Mother Is Subfamily Head
(standard errors below)

	(1)	(2)	(3)
AFDCDIF (in $000)	-3.80		
	(1.78)		
AFDCH (in $000)	-.420	-.511	
	(.698)	(.699)	
AGE	-.611	-.616	-.615
	(.141)	(.141)	(.141)
AGESQ	.011	.011	.011
	(.003)	(.003)	(.003)
EDUC	.214	.203	.202
	(.189)	(.189)	(.189)
EDUCSQ	-.0024	-.0026	-.0026
	(.005)	(.005)	(.005)
AGEED	-.0089	-.0083	-.0083
	(.006)	(.006)	(.006)
DEPENDS	-.532	-.530	-.559
	(.091)	(.091)	(.082)
PRESCH	.162	.154	.155
	(.154)	(.154)	(.154)
NWHITE	.345	.347	.344
	(.128)	(.128)	(.128)
SMSA	.114	.110	.115
	(.142)	(.142)	(.141)
UNEMP	-.064	-.063	-.062
	(.028)	(.027)	(.027)
SOUTH	.333	.376	.469
	(.184)	(.183)	(.131)
Constant	9.29	9.39	9.22
	(2.38)	(2.39)	(2.37)
Log likelihood	-852.5	-854.7	-855.0

the same $100 increase. Moreover, the Ellwood-Bane effect is statistically significant, and ours is not.[19]

For completeness we describe the findings on the demographic variables. Age is clearly an important determinant of a mother's propensity to live in a subfamily; older single mothers are less likely to live in subfamilies than younger single mothers, ceteris paribus. Note also that even after controlling for age and AFDC benefits (which vary with numbers of children), an increase in the number of dependents raises the probability that mothers are householders.[20] In addition, the large and statistically significant positive coefficients on NWHITE and SOUTH indicate a greater propensity for nonwhites and southerners to live in subfamilies, ceteris paribus. Similarly, the negative coefficient on UNEMP indicates that areas with higher unemployment rates tend to have fewer subfamilies, ceteris paribus.[21] Finally, note that these results are quite robust to the inclusion or exclusion of the AFDC variables.[22]

Results from Multinomial Logit Models

While more difficult to interpret than a binary logit, a multinomial logit provides a more complete test of our hypotheses. Consider the four alternatives introduced above:

(1) householder not receiving welfare,
(2) subfamily head not receiving welfare,
(3) householder receiving welfare, and
(4) subfamily head receiving welfare.

As indicated in the theoretical discussion, we expect an increase in AFDCDIF to reduce the probability that a single mother is in the fourth living arrangement and to increase the probability that she is in the other three.

Table 3 presents the key results. In a multinomial logit the log of the ratio of two probabilities is a linear function of the explanatory variables. We let alternative 4 (subfamily head receiving welfare) be the basis for comparison, and estimate three sets of coefficients. Thus, the first column of coefficients refers to the log of the ratio of the probability of choosing alternative 4 versus alternative 1, the second refers to the relative probability of choosing alternative 4 versus alternative 2, and the third to the relative probability of choosing alternative 4 versus alternative 3. A negative coefficient in the first column then implies that an increase in the value of the associated explanatory variable is related to a decrease in the relative probability of choosing alternative 4 versus alternative 1. An analogous interpretation applies to the negative coefficients in the second and third columns of the table.

Readers who are familiar with multinomial logits in the literature on transport mode choice (e.g., McFadden, 1981) may find Table 3 somewhat confusing. In that literature each choice is associated with a <u>different</u> vector of characteristics, and the analyst estimates a <u>single</u> vector of coefficients. Such models assume independence of irrelevant alternatives. In the Table 3 model (which is also called a multinomial logit model), each choice is associated with the <u>same</u> vector of characteristics, and the analyst estimates <u>multiple</u> vectors of coefficents (three vectors in this case). As demonstrated in Appendix B, one can place restrictions on the Table 3 model in order to arrive at a model identical to that used in the transport literature. We do not impose

Table 3

Four-Category Logit Estimates with Both AFDC Variables
(standard errors below)

	Subfamily on Welfare vs. Head Not on Welfare (1)	Subfamily on Welfare vs. Subfamily Not on Welfare (2)	Subfamily on Welfare vs. Head on Welfare (3)
AFDCDIF (in $000)	-3.57	-.716	-4.38
	(2.66)	(2.95)	(2.40)
AFDCH (in $000)	1.27	1.22	-.495
	(1.09)	(1.14)	(1.04)
AGE	-.523	.187	-.718
	(.237)	(.199)	(.226)
AGESQ	.011	-.0016	.012
	(.004)	(.004)	(.004)
EDUC	1.46	1.30	.478
	(.453)	(.425)	(.457)
EDUCSQ	-.057	-.054	-.020
	(.016)	(.016)	(.016)
AGEED	-.021	-.012	-.003
	(.010)	(.009)	(.010)
DEPENDS	-.181	.342	-.466
	(.138)	(.155)	(.133)
PRESCH	.588	.260	.039
	(.266)	(.281)	(.273)
NWHITE	1.19	.574	.195
	(.206)	(.201)	(.199)
SMSA	.034	-.165	.002
	(.221)	(.221)	(.222)
UNEMP	.019	.036	-.106
	(.045)	(.045)	(.044)
SOUTH	-.241	-.337	.612
	(.295)	(.294)	(.299)
Constant	1.46	-9.64	9.11
	(4.46)	(3.66)	(4.31)
Log likelihood	-1710.3		

these restrictions because our data firmly reject them. As such, the Table 3 model does not imply independence of irrelevant alternatives.

Consistent with expectations, all three coefficients on AFDCDIF in Table 3 are negative.[23] Thus, an increase in AFDCDIF reduces the relative probability of choosing alternative 4 versus alternatives 1, 2, and 3. Note, however, that none of the coefficients attain a high level of statistical significance. Of the three coefficients, the coefficient on AFDCDIF in the third column is the most negative and most precisely estimated. This is a plausible outcome. The column 3 coefficient indicates that an increase in AFDCDIF reduces the relative probability that a mother chooses to be a subfamily head on welfare versus being a householder on welfare. One would anticipate that if AFDCDIF affects behavior, its principal impact would be on choices between these two alternatives.

To help interpret these results, and to assess the magnitude of the AFDCDIF effect, it is useful to compute actual probabilities. In Table 4 we simulate the effect of a 10 percent decrease in the subfamily guarantee (AFDCS) from the sample mean of $318 to $286. (To reiterate, a decrease in AFDCS with AFDCH held constant implies an increase in AFDCDIF.) Column 1 displays the probability that an unmarried mother chooses each of the four alternatives when probabilities are computed at the sample means. Column 2 displays the same probabilities when the subfamily guarantee is reduced. As expected, the decrease in the subfamily guarantee decreases the probability of being a subfamily head receiving welfare (4) and increases the probability of being in either of the two householder states (1 or 3). Contrary to expectations, however, the probability of being a subfamily head not receiving welfare (2) decreases. Since a decrease in the subfamily guarantee does not affect utility attainable as a subfamily head without welfare, this is a surprise.

In our view there are two plausible explanations for this anomaly. First and most likely, it may simply reflect imprecise parameter estimates. The coefficient on AFDCDIF in the second column of Table 3 is estimated as -.716 with a standard deviation of 2.95. If that parameter is reduced by one standard deviation (to -3.666), the anomaly disappears. The first explanation is thus that the anomaly is a random event. The second explanation focuses on nonreporting of AFDC income. As recognized for some time, there is a problem of nonreporting of welfare income in the CPS. Some of the people who receive welfare income do not report receipt. As such, it is likely that some of the people who are classified as nonrecipients in these data are in fact recipients. In that case a decrease in the subfamily guarantee would decrease the probability that a mother is in either of the subfamily living arrangements. Indeed, in Table 4 the probability of residing in a subfamily falls from .330 to .306.

Given this, does AFDCDIF have a behavioral effect? In our judgment the weight of the evidence indicates that it does. Both the binary logit and the multinomial logit provide evidence to support the hypothesis of a behavioral effect, and the one anomalous finding does not necessarily contradict that hypothesis. Perhaps more important, however, the evidence also indicates that if there is a behavioral effect, that effect is quite small. As shown in Table 4, a 10 percent decrease in the subfamily

Table 4

The Effect of a 10 Percent Decrease in the Subfamily Guarantee (AFDCS)

	Subfamily Guarantee Set at $318	$286
Probability that an unmarried mother is a		
(1) Householder not receiving welfare	.380	.389
(2) Subfamily head not receiving welfare	.220	.206
(3) Householder receiving welfare	.290	.305
(4) Subfamily head receiving welfare	.110	.100
Total	1.000	1.000

guarantee has a small effect on headship and subfamily probabilities. We conclude that AFDCDIF probably does affect behavior, but not in a major way.

Turning to the other coefficients in Table 3, note that none of the coefficients on AFDCH attain a high level of statistical significance. Thus, as with the binary logit, we find no evidence that the overall level of benefits influences choices between subfamily and householder status.[24] Table 5 presents a simulation of the effect of raising the overall level of benefits. This is a simulation of a 10 percent increase in AFDCH holding all other variables (including AFDCDIF) at their sample means. As in Table 4, the effect on living arrangements is small.

Next, consider the demographic variables in Table 3. As in the binary logit, age is important. Increases in age are associated with a lower relative probability of being a subfamily head on welfare versus being a householder in either welfare status (columns 1 and 3). Moreover, additional dependents, high local unemployment rates, and not living in the South all reduce the relative probability of being a subfamily head receiving welfare versus being a householder receiving welfare (column 3). Finally, nonwhites have a higher relative probability of being a subfamily head receiving welfare versus being in either household status and not receiving welfare (columns 1 and 2). As indicated in Appendix Tables A.1 and A.2, these results are quite robust to inclusion or exclusion of the AFDC variables.

Table 5

The Effect of a 10 Percent Increase in the Householder Guarantee (AFDCS), Holding AFDCDIF Constant

	Householder Guarantee Set at $318	$350
Probability that an unmarried mother is a		
(1) Householder not receiving welfare	.380	.372
(2) Subfamily head not receiving welfare	.220	.216
(3) Householder receiving welfare	.290	.300
(4) Subfamily head receiving welfare	.110	.112
Total	1.000	1.000

Conclusion

This paper addresses two questions. First, how do state AFDC programs differ in their treatment of subfamilies versus householders? We find that in some states single mothers suffer a substantial reduction in AFDC benefits if they choose to head a subfamily in a larger household. Such states may determine that the mother is receiving an in-kind contribution of room and board from the larger household, and adjust benefits accordingly.

The second question concerns the effect of the AFDC program on a single mother's propensity to reside in a subfamily. In general, we find discernible but small effects. An increase in the overall level of benefits (benefits paid to both householders and subfamilies) had no statistically significant effect on "living arrangements." A decrease in the subfamily guarantee, which increases the income "wedge" between subfamilies and householders, slightly decreased the propensity to reside in a subfamily.

To conclude, at least in some states the AFDC program discourages mothers from residing in subfamilies. Since the behavioral effect of the differential treatment of householders and subfamily heads appear to be small, it is unlikely that this is a major cause of welfare dependency.

Notes

[1]For example, Danziger, Jakubson, Schwartz, and Smolensky (1982), Honig (1974, 1976), Hutchens (1979), Ross and Sawhill (1975).

[2]In surveying the states by telephone we also sought information on two related issues. First, we asked about the effect of a change to shared living arrangements when both units were poor. For example, what would happen if the mother and child moved in with a poor grandparent? In most states benefits were unaffected by such a change in living arrangement. Second, we asked how the treatment of a minor mother differs from that of an adult mother. In many states the treatment of minors was the same as the treatment of adult mothers. In some states, however, a minor mother who moved in with her nonpoor parents may (a) have had the parents' income deemed to the case, or (b) been excluded from the AFDC case (although the child may be an AFDC recipient), or both (a) and (b). It should be emphasized that with the prospect of federal rules in this area, the state provisions are in considerable flux.

[3]Michigan was really a hybrid of types 2 and 3. For small rent payments the AFDC guarantee was increased dollar for dollar, but after reaching a threshold (around $50 per month) the guarantee went back up to $416.

[4]We also estimated models that used actual benefits received as the dependent variable. Such models must include the recipient family's earned and unearned income as explanatory variables, thereby raising the problem of truncation bias (see Hutchens, 1978). Since truncation bias is not an issue when the payment standard is used as the dependent variable, and since the two models yield similar results, we focus here on the payment standard models.

[5]These data are published in U.S. Department of Health and Human Services (1984). We include MXBEN in the specification for two reasons. First, the variable contains information on the determinants of AFDC payment standards--information that is external to the 1982 Quality Control data. By including this variable, we are able to improve the model's predictive power. Second, although the model is estimated with the 1982 QC data, for purposes of the Section II empirical work it is used to impute benefits to women in different living arrangements in 1983. With this specification the model can be estimated using 1982 values of MXBEN, and the imputations can subsequently be based on 1983 values.

[6]A regression of the SHARE coefficient on the state policy type yields the folowing results:

$$\text{SHARE coefficient} = \underset{(7.60)}{22.50} - \underset{(4.18)}{20.83}\ \text{Policy type} \qquad R^2 = .35,$$

where we show standard errors below the estimates. The Spearman correlation between the two is -.39 and Kendall's Tau is -.33. These latter two are more appropriate than the usual Pearson correlation because state type is ordinal rather than cardinal. Both correlations are significantly different from zero at the 1 percent level.

[7]Given the differences in the data, it is no surprise that the correspondence between the survey results and the regression results is not exact. For example, the survey results for Oregon suggest lower benefits for subfamily heads, while the regression coefficient is positive and statistically significant.

[8]Neither the theoretical predictions nor the empirical results presented below differ substantively from those in a model that includes married women (see Hutchens, Jakubson, and Schwartz, 1986).

[9]In the last case we consider N to include the cash equivalent of the in-kind transfer. Note that, in general, we would expect $N_S > N_H$. However, if independent living is preferred to sharing a household, other things equal, then we may still observe a woman choosing a sacrifice income in order to live as a household head.

[10]The alternative is to model the transitions among living arrangements. We choose not to do this for two reasons. First, there are relatively few subfamilies and few transitions between subfamily and householder status in a given year. Panel data like the PSID do not contain enough observations for an analysis of the effects of the AFDC program on the number of subfamilies and households. (A key advantage to using cross-section data is that one observes larger numbers of subfamilies.) Second, we would have to calculate the steady state associated with the dynamic model in order to answer questions concerning the effect of AFDC on the number of subfamilies. Small errors in the dynamic model will be compounded in the calculation of the steady state, and hence our conclusions might be seriously misleading. For a detailed discussion of the equilibrium assumption, see our longer paper (Hutchens, Jakubson, and Schwartz, 1986). The issue is also discussed in an appendix which is available on request.

[11]The distinction between behavioral effects and break-even effects is most prominently discussed in Ashenfelter (1983).

[12]We can predict that an increase in AFDCH will increase the probability that an unmarried mother is a welfare recipient. That hypothesis is tested in our longer paper (Hutchens, Jakubson, and Schwartz, 1986), and found to be valid.

[13]While AFDC may also have an effect on women without children, we ignore that effect. Trying to account for AFDC effects on fertility in our simple model seems overly ambitious. Further, we feel that the focus on women with children makes the analysis more relevant for policy purposes.

[14]We did, however, experiment with predicted wage and nonwage income variables (both from simple linear regression and "selectivity-adjusted" regressions), as well as full structural models. See Hutchens, Jakubson, and Schwartz (1986), Chapter 5. In part because we predict the income variables with large error, our most meaningful results come from the simple reduced forms presented here.

[15]In particular, we estimated models of the form,

$PAYSTD_i = B_0 + B_1\ SHARE_i + B_2\ MAR_i + B_3\ MXBEN_i + B_4\ MXBEN_i x SHARE_i + B_5\ MXBEN_i x MAR_i + e_i$. See Section I for definitions of variables.

[16]We do not explicitly consider other welfare benefits available to the family, in particular Food Stamps and Medicaid. The problems in valuing Medicaid are well known (see Smeeding, 1982, for a discussion). Food Stamps also present a problem, primarily because the filing unit may differ from the AFDC filing unit for a subfamily. If the whole household files together, the entire unit is likely to be ineligible for Food Stamps, and hence AFDCDIF will understate the differential treatment of householders and subfamily heads. If the whole household files together

and is eligible, then we must impute (i) the number of people in the Food Stamp filing unit, (ii) the earned, and (iii) the nonearned income of that unit. Given our data, accurate imputations are impossible. If the AFDC subfamily establishes itself as a separate Food Stamp unit, on the other hand, then AFDCDIF will overstate the differential treatment by household status due to the leveling effect of Food Stamps. The magnitude of AFDCDIF is small, in general, and we therefore do not feel that ignoring Food Stamps introduces a substantial bias. Further, if we assume that the AFDC and Food Stamp filing units are identical, and impute the cash value of Food Stamps, there is little change in the empirical results. For a further discussion of those issues see Hutchens, Jakubson, and Schwartz (1986).

[17]Inclusion of married women in the sample does not substantively alter the results presented here. In particular, the AFDCDIF coefficient remains negative and statistically significant. See Hutchens, Jakubson and Schwartz (1986).

[18]There is an equivalent parameterization of the model. Since AFDCDIF = AFDCH - AFDCS, we have b1 AFDCH + b2 AFDCDIF = (b1 + b2) AFDCH + (-b2) AFDCS = c1 AFDCH + c2 AFDCS. For this model we have c1 = -4.22(1.89) and c2 = 3.80(1.78).

[19]While it is beyond the scope of this study to fully assess why the results differ, two reasons seem potentially important. First, the data are drawn from different sources. Ellwood and Bane use the 1976 Survey of Income and Education (SIE), and we use the 1984 Current Population Survey. The reason that this may be important is that prior to 1983 the Census Bureau incorrectly coded subfamilies, so that Ellwood and Bane

were forced to develop a procedure for identifying subfamilies. Second, they estimate linear regressions for both subfamily status and welfare receipt. Their measure of the AFDC effect in the subfamily status regression takes the form of an interaction between the published state AFDC benefit for a four-person family (G_s) and the predicted probability of welfare receipt (P_{is}). In addition, they include a vector of state dummy variables (d_s) in the subfamily status regression. Thus, their regression takes the form $y_{is} = \ldots + b_1 P_{is} G_s + b'_s d_s + \ldots$, where b_s is a vector of coefficients. As they note (p. 167), "the AFDC benefit level varies only across the states and thus is completely collinear with the state dummies." As such, it can be shown that the coefficient on their interaction variable is determined by within-state variation in the predicted probability of welfare receipt. It reveals only that poorer women (i.e., women who are more likely to be AFDC recipients) are more likely to live in subfamilies; it does not reveal the effect of variation in AFDC benefit levels. In consequence, we employ a different specification. Finally, some readers have suggested that by using the published AFDC benefit for a four-person family, Ellwood and Bane used a more accurate measure of AFDC benefits than we did. Since we use the same published data in predicting AFDCH and AFDCS, that seems unlikely. Just to be sure, however, we re-estimated our models with AFDCH replaced by the published measure. The coefficient remained small and statistically insignificant.

[20]The variation in AFDCH and AFDCS comes from variation among states and from nonlinearities in the benefit formula as a function of family size.

[21]This should not be used as indicating that a lack of job opportunities leads to female headship. UNEMP is an average over all men and women in the state and is influenced by industrial mix, population composition, unemployment insurance, etc. It indicates little about the job opportunities confronting a particular woman. UNEMP is simply a control variable. This coefficient may arise out of unobserved characteristics of states with comparatively high unemployment rates. Results on the AFDC variables are robust to exclusion of the UNEMP variable.

[22]Other tests indicated that all of the results presented here are also robust to exclusion of the SMSA, South, and unemployment rate variables; to changes in the specification of the education variable; to inclusion of an AFDC tax rate variable; to inclusion of a "never married" dummy variable; to changing from one to three region variables; to replacing AFDCDIF with the ratio AFDCH/AFDCS; and to restricting the sample to single mothers under 25.

[23]As in the binary logit models, we can parameterize the model in terms of AFDCH and AFDCS rather than AFDCH and AFDCDIF. If we do so, we obtain the following results:

	column 1	column 2	column 3
AFDCH (in $000)	−2.30	.50	−4.88
	(2.84)	(3.14)	(2.61)
AFDCS (in $000)	3.57	.72	4.38
	(2.66)	(2.95)	(2.40)

[24]To check whether this is due to collinearity between AFDCDIF and AFDCH, we estimated the multinomial logit without AFDCDIF. As indicated in Appendix Table A.1, this does not alter the results on AFDCH.

Appendix A

Table A.1

Four-Category Estimates without AFDCDIF
(standard errors below)

	Subfamily on Welfare vs. Head Not on Welfare (1)	Subfamily on Welfare vs. Subfamily Not on Welfare (2)	Subfamily on Welfare vs. Head on Welfare (3)
AFDCH (in $000)	1.21	1.22	-.587
	(1.09)	(1.14)	(1.04)
AGE	-.530	.183	-.730
	(.237)	(.200)	(.226)
AGESQ	.011	-.002	.012
	(.0043)	(.0040)	(.0041)
EDUC	1.44	1.29	.456
	(.453)	(.425)	(.456)
EDUCSQ	-.057	-.054	-.020
	(.016)	(.016)	(.016)
AGEED	-.020	-.012	-.003
	(.010)	(.009)	(.010)
DEPENDS	-.175	.347	-.459
	(.139)	(.155)	(.133)
PRESCH	.554	.261	.031
	(.266)	(.280)	(.272)
NWHITE	1.19	.577	.201
	(.206)	(.201)	(.199)
SMSA	.031	-.165	-.001
	(.221)	(.222)	(.222)
UNEMP	.020	.037	-.106
	(.044)	(.045)	(.044)
SOUTH	-.197	-.325	.667
	(.294)	(.292)	(.298)
Constant	1.60	-9.59	9.31
	(4.47)	(3.66)	(4.31)
Log likelihood		-1712.5	

Table A.2

Four-Category Logit Estimates with No AFDC Variables
(standard errors below)

	Subfamily on Welfare vs. Head Not on Welfare (1)	Subfamily on Welfare vs. Subfamily Not on Welfare (2)	Subfamily on Welfare vs. Head on Welfare (3)
AGE	-.532	.184	-.720
	(.237)	(.200)	(.226)
AGESQ	.011	-.002	.012
	(.004)	(.004)	(.004)
EDUC	1.44	1.29	.471
	(.453)	(.425)	(.456)
EDUCSQ	-.057	-.054	-.020
	(.016)	(.016)	(.016)
AGEED	-.020	-.012	-.003
	(.010)	(.009)	(.010)
DEPENDS	-.104	.415	-.500
	(.125)	(.143)	(.120)
PRESCH	.546	.253	.026
	(.265)	(.280)	(.272)
NWHITE	1.20	.584	.204
	(.206)	(.201)	(.199)
SMSA	.017	-.178	.017
	(.220)	(.221)	(.221)
UNEMP	.015	.031	-.105
	(.044)	(.045)	(.044)
SOUTH	-.426	-.541	.808
	(.213)	(.210)	(.221)
Constant	2.01	-9.24	8.95
	(4.45)	(3.65)	(4.29)
Log likelihood		-1716.2	

Appendix B

The Reduced-Form Multinomial Logit Model

For notational simplicity, consider a model with three alternatives. Each alternative k is characterized by a vector of attributes x_k, k = 1,2, and 3. The multinomial logit in the transport mode choice literature is

(1) $$P(\text{choose } j) = \exp\{\beta' x_j\} \;/\; \Sigma_k \exp\{\beta' x_k\}.$$

Now stack all attributes of all alternatives into a single vector z:

(2) $$z = (x_1', x_2', x_3')',$$

and consider the polychotomous choice model which allows choice to depend freely on all the elements of z:

(3) $$P(\text{choose } j) = \exp\{\alpha_j' z\} \;/\; \Sigma_k \exp\{\alpha_k' z\}.$$

If the equation (1) model were appropriate, then we would have

(4a) $$\alpha_1 = (\beta' \; 0' \; 0')',$$

(4b) $$\alpha_2 = (0' \; \beta' \; 0')', \text{ and}$$

(4c) $$\alpha_3 = (0' \; 0' \; \beta')',$$

where 0' indicates a row vector of zeros. However, since probabilities must add to one, there are only two free vectors of coefficients α_k. We normalize the model by dividing the numerator and denominator of each probability in equation (3) by $\exp\{\alpha_3' Z\}$, in effect choosing the last alternative as a "base case." We then rewrite the probabilities in equation 3 as

(5a) $P(\text{choose } 1) = \exp\{\gamma_1{}'Z\}/D$,

(5b) $P(\text{choose } 2) = \exp\{\gamma_2{}'Z\}/D$, and

(5c) $P(\text{choose } 3) = 1 / D$,

where

(6a) $D = \exp\{\gamma_1{}'Z\} + \exp\{\gamma_2{}'Z\} + 1$,

(6b) $\gamma_1 = \alpha_1 - \alpha_3$, and

(6c) $\gamma_2 = \alpha_2 - \alpha_3$.

If the restrictions of the model in equation (1) are correct, then we would have:

(7a) $\gamma_1 = (\beta' \; 0' \; -\beta')'$ and

(7b) $\gamma_2 = (0' \; \beta' \; -\beta')'$.

A Wald test of these restrictions is easy to compute once the reduced-form model in equations (5) has been estimated, and is the way in which we tested the model. For more details on the procedure see Hutchens, Jakubson, and Schwartz (1986), Appendix 4b, pages 110-123.

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