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THE EFFECTS OF THE MARRIAGE MARKET AND AFDC BENEFITS ON EXIT RATES FROM AFDC

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The Effects of the Marriage Market and AFDC Benefits on Exit Rates from AFDC

by

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Abstract

Using monthly data from the Survey of Income and Program Participation, the paper shows that roughly one-third of AFDC spells end within six months, while 40 percent last two years or more. Multivariate discrete hazard model estimates reveal that measures of spouse availability, such as sex ratios and employment ratios for single males, affect exit rates for whites but not for blacks. For blacks the state unemployment rate is important. For both races high welfare benefits slow exit, although the effect may not be significant, whereas higher education and older age speed exit.

The Effects of the Marriage Market and AFDC Benefits on Exit Rates from AFDC

I. INTRODUCTION AND BACKGROUND

A. Concern about Welfare Dependence

The Aid to Families with Dependent Children (AFDC) program provides aid for over 3.5 million families in the United States, mostly families without fathers. Policymakers have expressed concern over the extent to which providing income in the form of AFDC increases the dependence of recipients on the government. If time spent on AFDC causes a deterioration of skills or self-esteem for recipients or their children, then the program can generate significant long-term costs; these must be compared to the short-term poverty relief and other benefits offered by the program.

Recent evidence indicates that a large proportion of new AFDC cases will stay on this welfare program only a short time, yet the remaining long-term cases make up a sizable part of the caseload at any one time. In this study I follow others who have sought to understand why some persons or groups experience much longer episodes, or spells, of AFDC receipt than others. More specifically, I explore the effects of the availability of potential spouses and the AFDC benefit level on the length of AFDC spells.

As its conceptual base, this study employs a search framework; this is appropriate when acceptable jobs and spouses are not always available. I then use a new data set, the Survey of Income and Program Participation, to verify earlier work and to refine our understanding in these two key areas: the role of the "marriage market" (availability of marriageable spouses) in explaining black/white exit rate differences and the effect of AFDC benefit levels.

B. Empirical Evidence about Time on Welfare

A large body of literature addresses the effects of welfare programs on family stability and poverty. (See, for example, the survey in Garfinkel and McLanahan, 1986.) The question of welfare dependency also has a long history, but good data and methodology for studying this dependency have only recently become available. Bane and Ellwood (1983), O'Neill et al. (1984), and Blank (1986) provide us with much of what we know about the determinants of exit rates from AFDC. The earlier work of Hutchens (1981) and the work of Plotnick (1983) also provide estimates of the probability of transitions in and out of welfare. Bane and Ellwood use data on welfare spells from the Panel Study of Income Dynamics (PSID) to estimate exit rates from AFDC. Bane and Ellwood define a year of welfare receipt as one in which $250 or more of AFDC is received, and they conclude that 65 percent of those currently on welfare are involved in a welfare spell of 8 years or more. O'Neill and his colleagues used data from the National Longitudinal Survey (NLS) as well as from the PSID. They do not use the $250 exclusion, and get somewhat higher exit rates. Both Bane and Ellwood and O'Neill et al. find that education increases the rate of exit from welfare, whereas the presence of children decreases it. Obviously a yearly accounting period can lead to overestimates of actual monthly use: AFDC receipt in two nonconsecutive months, each in a different accounting year, produces a two-year spell, even if those were the only two months of AFDC receipt.

In addition to work with a yearly accounting period, O'Neill et al. (1984) also exploit monthly AFDC case records that span a long period (1965-1982). By this means they avoid the overstatement of dependence that results from using annual data. As noted by O'Neill et al. the data may understate welfare spell length to the extent that cases close for administrative reasons or because households move. Such events terminate a measured welfare spell even though the recipients return to the rolls so fast that no real interruption has occurred. For the cohort of cases opened in 1980, the authors find a relatively short median spell length of 10 months.

Blank uses monthly data as well, but from the control group of the Seattle/Denver Income Maintenance Experiment (SIME/DIME). She concludes that the monthly data do not show strong evidence of duration dependence; that is, she concludes that the rate of exit from AFDC does not decline significantly with time spent on the program.[1] This result is robust across many specifications of the hazard model. She finds that nearly 60 percent of all spells of AFDC recipt will end within two years, and the mean length of a spell is 20 months. She also finds that age, education, and nonlabor/nonwelfare income of the head increase exit rates, whereas being black decreases them.

The studies discussed above also include findings on exit rates disaggregated by type of exit. Three types of exit are considered: exit by marriage, exit by earnings increase, and the residual--other exits. All the studies conclude that marriage and earnings increases are the primary routes off of AFDC. They find that blacks are less likely than whites to leave AFDC by marriage (holding age, education, and number of children constant). The rate of exit by earnings increase is not significantly different for blacks and whites. These results imply that the lower overall rates of exit by blacks is due to their lower propensity to marry (Blank, 1986, p. 27). The studies do not attempt to measure the extent to which the lower marriage rates for black women are due to their poorer marriage prospects (i.e., a poorer marriage market in terms of availability and income of potential spouses), as opposed to noneconomic cultural influences.[2] Wilson and Neckerman (1986) suggest that economic demographics that result in a low availability of single, employed black males may be more important in explaining the large number of single female-headed households than welfare programs per se.[3] The empirical results in Section III, however, show that measures of spouse availability affect the exit rate from AFDC for whites but not for blacks.

Turning to the issue of program benefits (guarantee levels), past research has generally found a negative effect of AFDC benefit levels on exit rates, but the evidence is not strong. This is partly explained by data deficiencies. Blank's data provides some variation in real benefit rates over time, but, as she notes, lack of within site variability in the benefit data most likely accounts for the small estimated effect. O'Neill and his colleagues have better data variability, but the significance (and sign) of program benefits are quite sensitive to the specification they use for either the PSID or caseload data. Their results, using the NLS, suggest that benefit levels have a significant negative effect on exits by marriage, but not by other exit routes. My results show a negative effect: women in high benefit states have lower exit rates, all else equal. Use of a more comprehensive measure of the total welfare benefits including food stamps and Medicaid produces similar results. The SIPP provides good variation in benefit levels from the 38 identifiable states of residence over a three-year period.

C. Conceptual Framework

The conceptual model that underlies the estimation of exit rates from AFDC is generally based on a model of choice: a woman on AFDC chooses between the option of staying on or getting off welfare. Some studies further divide the choice of exiting welfare into marrying or getting a job. An early and important example of this framework is found in Danziger et al. (1982).

In these discrete choice models, a woman chooses to stay on welfare, take a job, or marry by choosing the option that delivers the highest present value of utility. To make this static equilibrium model explain the dynamics of exit and entry, one allows the returns on the various options or tastes to vary over time and posits that the woman reevaluates her options as the changes occur and may choose different options over time. (See, for example, Blank, 1986.)

Information about options can be considered to be one feature of the environment that changes over time.[4] Changes in job opportunities or spouse availability can cause a woman to reevaluate her options and choose to leave AFDC. If we think of "information about job opportunities" as wage offers, and "information about spouse availability" as marriage offers, we can employ a search framework. Restated, a woman on AFDC is engaged in search for a spouse and/or a job. Hutchens (1979) develops one marital search model, without job search, and finds that AFDC benefits reduce remarriage probabilities as predicted.

The key insight of search theory is that a woman does not face a single, deterministic wage offer or a marriage offer from a spouse with known characteristics. Rather she faces a distribution of wages or spouse offers from which she may draw. In the face of this uncertainty, she develops a strategy for deciding whether to accept or reject job and marriage offers when she receives one. Of course, the benefits of search need not outweigh the costs, and a woman may decide not to search. In any case, in a search framework the rate at which women leave AFDC depends on the rate at which offers are received and the probability that those offers are acceptable and cause her to leave AFDC. This probability depends on the relative evaluation of the expected present value of utility of accepting or rejecting the offers. Other changes in the environment, such as children leaving home, will, of course, also produce exits from AFDC.

To help clarify the search framework, consider a simple model where a woman searches for spouses and jobs costlessly. If she accepts either type of offer she leaves AFDC. Over some short time period, let $\lambda_j(t)$ be the probability of a job offer and $\lambda_m(t)$ be the probability of a marriage offer. Let S(t) be the expected present value of utility of remaining single and staying on AFDC during period t, assuming an optimal search strategy. Let J(t) and M(t) be similarly defined utility (value) functions for accepting the current job offer or marriage offer, respectively, if an offer is available. The woman's goal is to maximize utility:

If a job offer occurs, she wants to Max $[J(t), S(t)]$. She accepts the current job offer if $J(t) > S(t)$.

If a marriage offer occurs, she wants to Max $[M(t), S(t)]$. She accepts the current marriage offer if $M(t) > S(t)$.

If j and m offers are rare events then we can ignore the probability that both events occur in a short time period. We have:

$$\text{Prob (leave AFDC at time t)} =$$

$$\lambda_j(t) \text{ Prob } (J(t) > S(t)) + \lambda_m(t) \text{ Prob } (M(t) > S(t))$$

where $J(t)$, $S(t)$, $M(t)$ are the present value of discounted utility assuming an optimal search strategy is employed. Spouse availability affects λ_m but also $S(t)$, $M(t)$, and $J(t)$, since it may affect job search strategy and spouse search strategy (i.e., the reservation wages).[5] This formulation suggests that the rate of exit from AFDC conditional on the time spent on AFDC, the hazard rate, is determined by the rates λ_j and λ_m and the determinants of the value functions S, M, and J. I later discuss the empirical form of the hazard.

This paper is concerned in particular with two comparative statics: How does the level of AFDC benefits affect the exit rate? And how does spouse availability affect the exit rate.[6] In a simple model where taking a job causes you to leave AFDC,[7] higher AFDC benefits raise $S(t)$ relative to $M(t)$, and $S(t)$ relative to $J(t)$. This should reduce the exit rate. The effect of greater spouse availability can be thought to raise the probability of receiving an offer, λ_m, which by itself increases exits, but it also may change reservation wages and reservation spouse quality and thus have an ambiguous effect. Increases in the quality of potential spouses increases the proportion of "good draws" in the offer distribution, but a woman may become more choosy [and raise her standards] when prospects get better.

A formal model of the joint search process is beyond the scope of this paper. This section is intended to suggest that search framework can guide our decisions about relevant variables for the empirical work. Education level, availability of other income (nonearned, nontransfer), number and age of children, mother's age, and area unemployment rates will all affect the cost and benefits of search just as do AFDC program parameters and spouse availability.

II. DATA FROM THE SIPP

A. The SIPP and Defining Monthly Spells of AFDC Receipt

This study uses monthly data on AFDC receipt from the 1984 panel of the Survey of Income and Program Participation (SIPP). The SIPP includes monthly data on approximately 20,000 households over 32-36 months. My sample is restricted to single female heads of AFDC units. The AFDC unit was defined as the family or subfamily of which the woman was the head. The "family" definition may differ from the administrative AFDC unit when the family includes other relatives of the head but there is no subfamily. A spell of AFDC receipt will be defined as a length of time during which the head of an AFDC unit continuously receives AFDC income. (Some recoding was done, as explained below.) In order to make best use of the data on the length of AFDC spells, the beginning date of an AFDC spell must be known. This restricts the sample to spells observed to begin after the first month of the panel--thus the maximum length of a completed spell of AFDC receipt will be 30 months. I include data from cases in which the end of a spell is not observed--right censored--and use only the first observed spell of AFDC receipt for each head.

Many heads had multiple recorded spells of receipt during the 32-month period.[8] Persons who had one-month gaps between two spells of AFDC receipt were recoded as one continuous spell. This was done to limit the effects of any administrative churning.[9] These one-month spells of nonreceipt do not represent less dependency on the program. Thus a completed spell can occur in one of two ways: a two-month period of nonreceipt, or a one-month period of nonreceipt followed by exit from the sample. Those persons who leave the sample immediately following a month of AFDC receipt are treated as censored. When data were imputed to a person who missed an interview, I ignored the imputation and treated the person as leaving the sample.

The first 32 months of the SIPP panel contain 1602 spells of AFDC receipt including multiple spells and spells that were ongoing at the time a person was first observed in the panel. When restricted to cases with an uncensored spell beginning (i.e., where we observe at least one month of nonreceipt prior to the spell beginning) and when one-month gaps are recoded (15 cases), the sample contains 619 cases. I then eliminated multiple spells and took the first observed spell for each person. This leaves 527 cases. Within this group there appear to be many suspect cases, and I used the following additional restrictions. First, I used only female heads, eliminating men, married women, and some inconsistent cases of multiple-recipient households.[10] Second, of the remaining 398 cases, I eliminated those who had no eligible children (age 18 or younger) at some point during the spell. Third, for the remaining 368 cases, I checked the income eligibility. The results reported in the text are for those 350 who are not likely to be ineligible because their income is too high (see Appendix A for details). As described later I also used a less restrictive and a more restrictive definition of income eligibility to check the sensitivity of results and found the results robust. Fourth, for the hazard models, I included only those cases with identifiable state of residence and complete age and education data. I excluded races other than black and white to clarify interpretation of the race variable. This adjustment left 329 cases for the hazard models.

B. Nonparametric Estimates of Time on Welfare

Table 1, Panel A, shows the distribution of first observed spells in the SIPP data. Included in the final sample of 350 spells of all races are 13 spells that include a recoded one-month gap. The median estimated spell length is about 18 months.

The table also shows a nonparametric (Kaplan Meier) estimate of the probability of having a spell of at least length T (the survivor function). Figure 1 graphs this survivor function against time. Note that 45 percent of AFDC spells are expected to last longer than 22 months, while 30 percent end within six months.

In Panel B these results are compared to the estimated survivor function from Blank, who used monthly data from the SIME/DIME control group for 1970 to 1976. The survivor functions are directly comparable up to 24 months.[11] The differences between the two studies are quite small past six months. This is remarkable because there are differences in sample design and data collection, minor differences in spell

Table 1

Distribution of First Observed AFDC Spells:

Panel A: SIPP (1984 Panel)[a]

Time (Months)	Number Completed	(Cumulative Percentage)	Number Censored at End of Period	(Cumulative Percentage)	Survivor Function[b]
2	38	(28)	25	(12)	.889
4	76	(55)	65	(30)	.768
6	94	(68)	84	(39)	.700
8	105	(76)	107	(50)	.654
12	123	(89)	143	(66)	.562
16	127	(92)	165	(76)	.533
20	132	(96)	188	(87)	.482
22	134	(100)	191	(88)	.449
31	134	-	216	(100)	-

Estimated median spell length: 18 months

Panel B: Survivor Function for First Observed AFDC Spells from SIME/DIME[c]

Time (Months)	Survivor Function[b]
2	.907
6	.750
12	.588
24	.433
36	.323

[a]Sample consists of first observed spells by female heads with children. Sample excludes those likely to be income ineligible.

[b]Kaplan Meier estimates of the survivor function.

[c]This panel was constructed from data kindly supplied to me by Rebecca Blank from her 1986 paper.

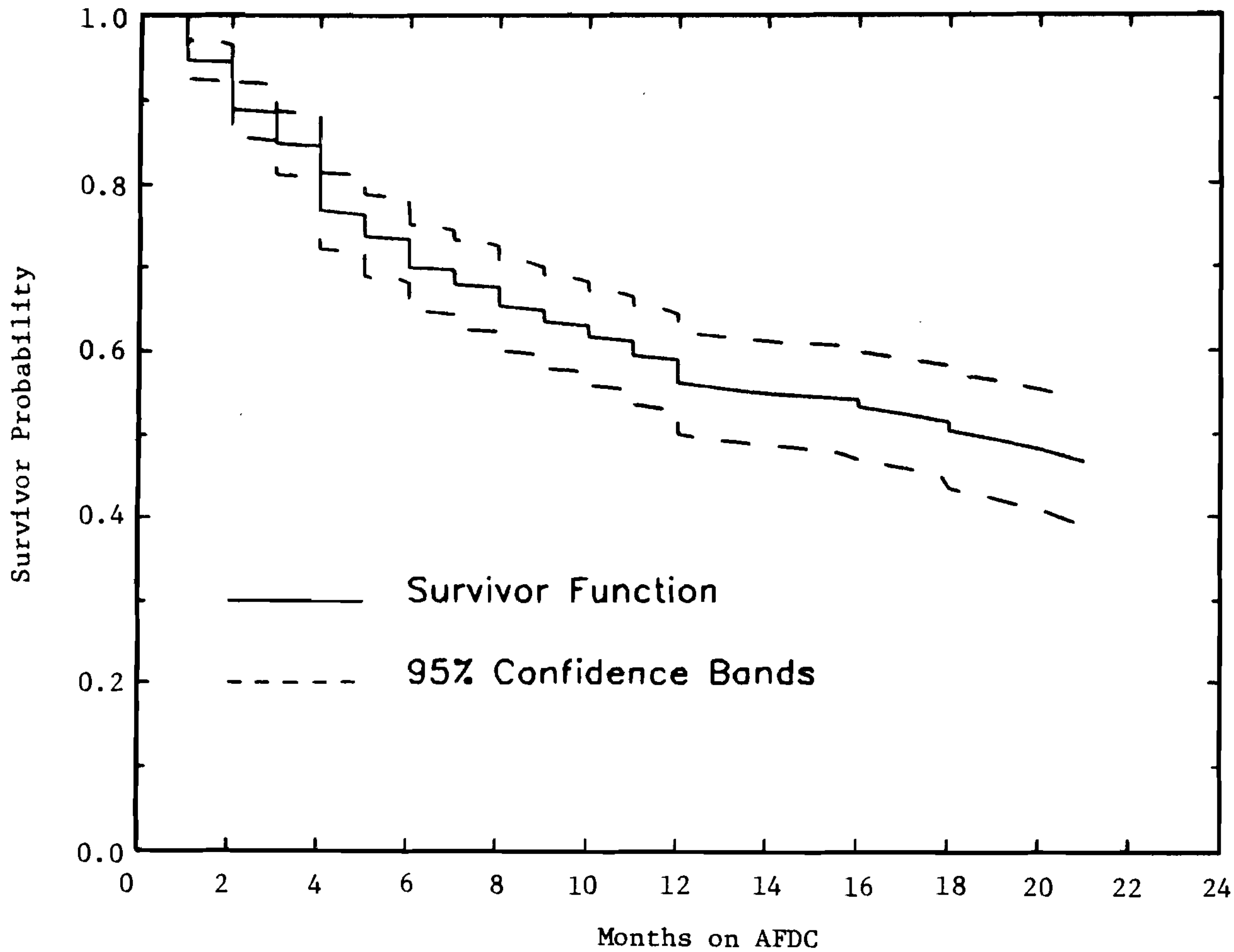

Figure 1

Survivor Function: All Races

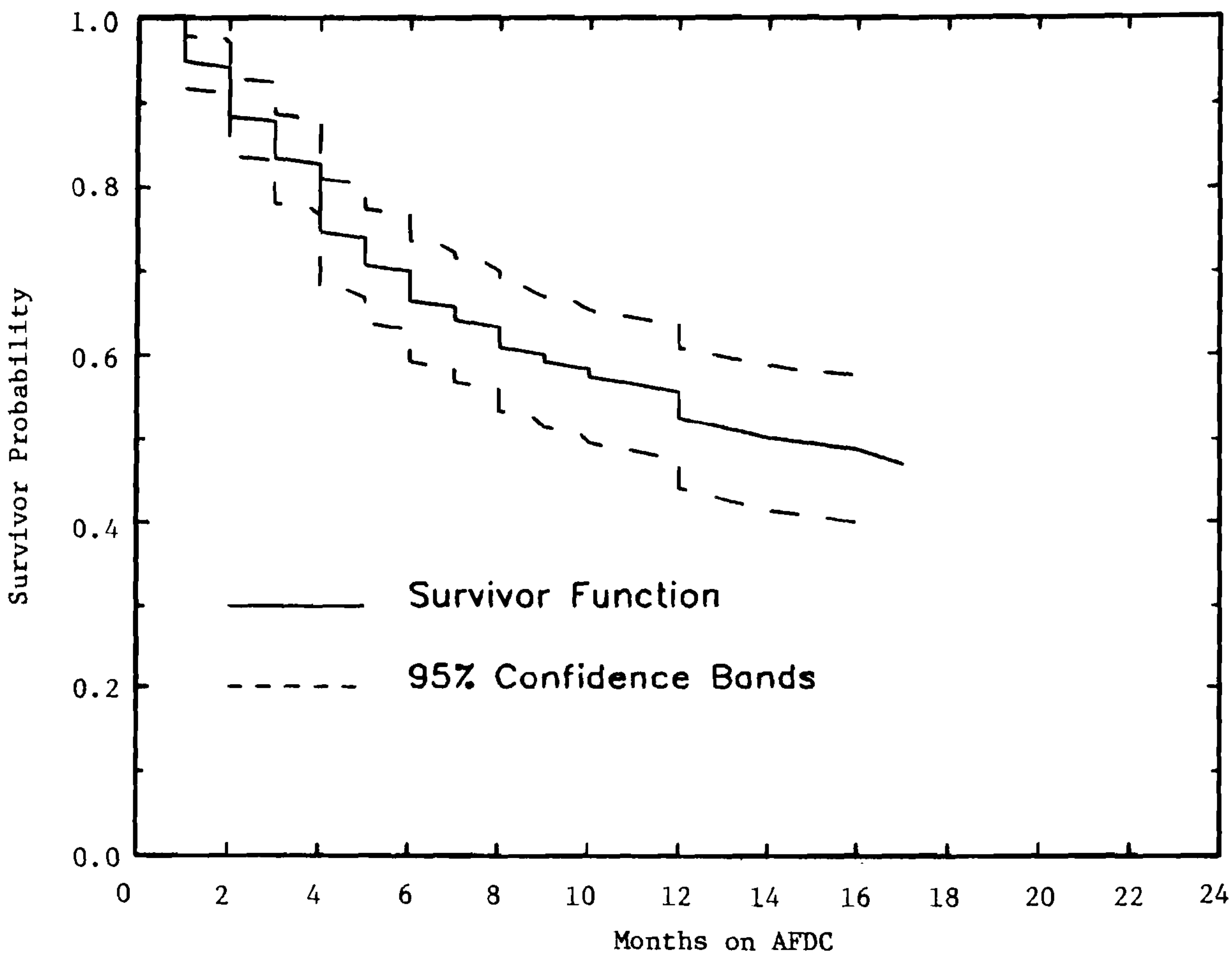

Figure 2

Survivor Function: Whites

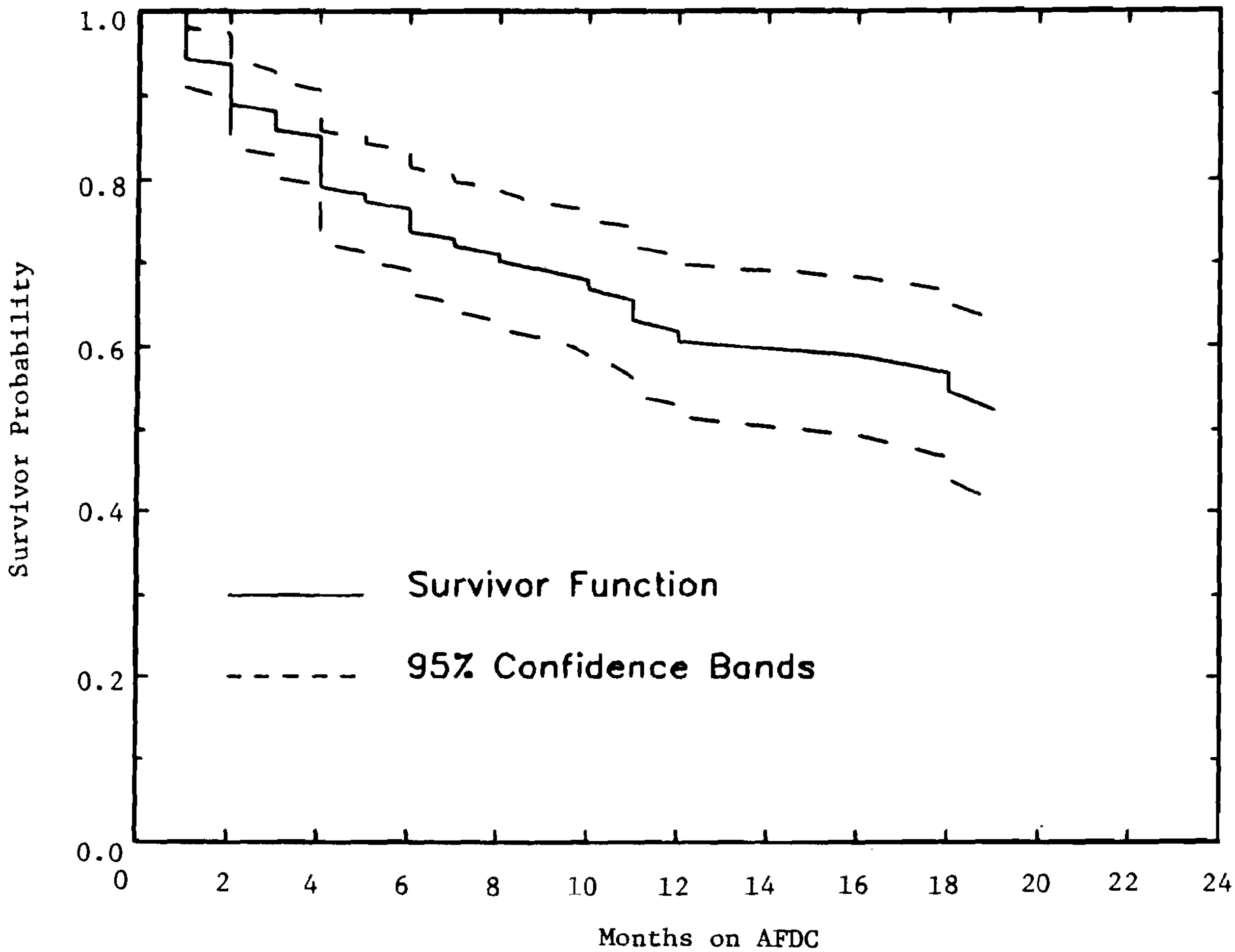

Figure 3

Survivor Function: Blacks

state of residence. These variables are intended to capture the relevant components of a state's welfare package. Obviously, they may also pick up the effects of other unobserved state-specific attributes, as noted by Ellwood and Bane (1985).

Table 2 shows the definitions and means of the explanatory variables. Some vary over the AFDC spell and others do not. All dollar denominated variables are adjusted to real values by the monthly CPI (to January 1984 dollars).

The variable labeled other income available, OTHINC, includes property income and private transfers (alimony and child support). The unemployment rate, UNEMP, is the monthly rate by state.

Two measures of spouse availability are computed. The first, SEXRATIO, is the ratio of single males to single females of the same race and in a relevant age group by state of residence. The key assumption is that this ratio approximates the availability of a marriage partner for each woman in a particular state. Demographers Goldman, Westoff, and Hammerslough (1984) point out that sex ratios aggregated by age and race do not adequately represent the availability of potential spouses. The method in the next paragraph attempts to incorporate their ideas, but it is much rougher. The second marriage market measure is EMPMALE, the ratio of employed single males to all single males, by age group, state, and race. This is in the spirit of Wilson and Neckerman's argument that the quality of potential spouses is important.

SEXRATIO is calculated from the 1980 decennial Census by race, state, and age group. Goldman, Westoff, and Hammerslough (1984) present evidence that there is a fairly large variation in age differences at

Table 2

Definitions and Means of Explanatory Variables

Variable	Mean
A. Fixed during the spell	
AGE (age at spell beginning)	26.7
RACE (0 = white, 1 = black)	.437
SEXRATIO (ratio of single men to single women of same race and age group, by state)	.922
EMPMALE (ratio of single employed men to single men of same race and age group, by state)	.676
B. Vary during the spell	
EDUC (head's highest grade completed)	12.0
NKIDS (number of children younger than age 18 in the AFDC unit)	1.73
YKID (dummy set to 1 if any children younger than age 6 are in the AFDC unit)	1.01
AFDC4MAX (AFDC benefit maximum for family of 4, by state, in dollars)	377
TBEN (measure of total welfare package for family of 4 including AFDC, food stamps and cash value of Medicaid, by state, in dollars)	595
OTHINC (private transfers and property income of female head, in dollars)	179
UNEMP (monthly state unemployment rate, in percentages)	7.89
AFDCU (1 = state has AFDC-Unemployed Parent program; 0 = no program)	.592

Notes: Mean values taken during the first month of the spell for first observed spells of female heads in SIPP. Sample Size = 329 persons (spells). The sample requires the presence of children and excludes cases likely to be income ineligible. All dollar amounts in January 1984 dollars.

marriage so I chose 11-year age groups. I assume that grooms are on average two years older than their brides, also based on Goldman and his colleagues (1984). Thus, for a woman aged 30, I computed the number of unmarried men of age 27 through 37 and divided it by the number of unmarried women aged 25 through 35 to get SEXRATIO. This was done for each race, state, and woman's age from 18 through 54. These ratios were then associated with sample women by race, state, and age. Appendix B provides details.

The employment ratio, EMPMALE, was computed from the 1980 Census, then updated to 1985 by an adjustment to reflect changes in employment between 1980 and 1985. Again, details are in the Appendix B.

For white women in the sample, the mean SEXRATIO is .97 and the mean EMPMALE is .74. For black women the means are .87 and .59, respectively. As noted by Wilson and Neckerman, the relative availability of employed single black males is lower. Inspection of the ratios also reveals wide variation across states and within each race.

III. EMPIRICAL RESULTS

A. Estimated Exit Rates from AFDC

This section describes the estimation of a discrete time hazard model for exits from AFDC. For an individual, let T index the month that the person leaves AFDC (without censoring). Define the discrete time hazard rate as

$$P(t) = \text{Prob}(T=t \mid T \geq t,\ X(t)),$$

where X(t) is the vector of explanatory variables at time t. I use a proportional hazards model with a complementary log-log form:

$$\log[-\log(1-P(t)] = a(t) + b'X(t).$$

This form arises from calculating the hazard for a continuous time proportional hazard model where the data has been grouped into discrete data points. (See Allison, 1982, p. 73). The parameter vector b tells the effect of the explanatory variables on the rate of exit, and these effects are assumed constant over time. The parameters a(t) can represent an arbitrary function of time, producing an essentially nonparametric underlying hazard.

The model below specifies the time dependence of the hazard a(t) as a step function with seven steps (time dummies). The coefficients on the steps are shown in Appendix C Tables C-1 and C-2 corresponding to particular specifications described below. The steps show a peak at months 2-4, then a decline.

Table 3 presents two models. Model 1 uses the AFDC benefit level for a family of four, AFDC4MAX, as the welfare benefit measure, while model 2 uses the more comprehensive measure, TBEN. In model 1 the AFDC benefit level has a fairly precisely estimated negative effect on exit rates. The results suggest that the marriage market variables are important in determining exit rates. The SEXRATIO has a positive effect that is statistically significantly different from zero at a 10 percent level. The precisely estimated coefficient on EMPMALE suggests that the availability of employed single males, a proxy for spouse quality, speeds exits.

Among the remaining model 1 coefficients, only two are very well estimated: higher education speeds exits, as does older age. The remaining coefficients are imprecisely estimated and none is significantly different from zero at a 5 percent level. The

Table 3

AFDC Exit Rates: Estimated Hazard Rate Regressions
Complementary Log-Log Specification for Proportional
Hazard Model with Time Dummies

	Model 1	Model 2
EDUC = 12	-.179 (.219)	-.179 (.219)
EDUC > 12	.606* (.255)	.587* (.257)
OTHINC (in $1000's)	.812 (1.18)	.641 (1.18)
NKIDS	-.0782 (.103)	-.0729 (.102)
YKID (=1 if present)	-.292 (.253)	-.270 (.253)
AGE	.0432* (.0158)	.0406* (.0157)
RACE (=1 if black)	.157 (.287)	.158 (.287)
UNEMP (%)	-.0548 (.548)	-.0406 (.573)
AFDCU	.160 (.243)	-.0135 (.218)
SEXRATIO	.831 (.473)	.673 (.467)
EMPMALE	2.54* (1.08)	2.46* (1.08)
TBEN ($1,000's)	-	-2.53 (1.37)
AFDC4MAX ($1,000's)	-2.40* (.102)	-
Log likelihood	-490.2	-491.4
Sample size		
Person-months	2,989	2,989
Persons	329	329

Notes: Standard errors are in parentheses. Sample consists of first observed spells by female-headed households with children. Cases likely to be income ineligible are excluded. Specification includes a constant and 7 time dummies for spell duration to that month: dummies for 3-4, 5-6, 7-8, 9-12, 13-16, 17-20, 20+ months (see Table C-1).
*Statistically signficiant at the 5% level.

unemployment rate has a negative coefficient, as one would expect for a measure of job availability, but it is imprecisely estimated. Note also that the residual effect of race has an insignificant effect when all of the other variables are included.

Model 2 shows that replacing AFDC4MAX with TBEN changes the results very little. The negative coefficient on TBEN is less well estimated (but significant at the 10 percent level) and roughly the same magnitude. Thus the results are not very sensitive to the benefit measure. The significant effects of higher education, age, and the male employment ratio come through.[14] Table 5 illustrates the magnitudes of the coefficients by simulating the survivor function. I computed the survivor function for each person based on the person's characteristics at the beginning of the spell, then averaged across persons at each month to produce a mean survivor function.[15] Comparing the first line of the table to the Kaplan Meir estimate of the survivor function in Table 1 shows that the model fits quite well. The table also shows that education and age have large effects relative to the overall average--the base case. Of course in all models the benefit level and AFDC-U dummy proxy for all the relevant components of a state's welfare package and enforcement policies and should be interpreted with care. The availability of AFDC-U has no impact; it always has a small and imprecise coefficient.

B. Separate Black and White Models and Other Specifications

The lack of significance of race in the above specifications does not imply that the hazards for the two groups are the same, since other coefficients may differ by race. Table 4 explores this question by showing the model 2 specification separately for whites and blacks, and panels B and C of Table 5 show simulations for these specifications. The sample sizes for these hazards are relatively small: 185 whites of whom 76 have complete spells, and 144 blacks of whom 52 have complete spells. Given the large number of parameters--17 including time dummies--caution is in order. The number of time dummies was reduced to 6 to facilitate convergence for these models.

For whites, the marriage market variables have positive, statistically significant effects on exit rates. The simulation shows that a 10 percent rise in SEXRATIO or EMPMALE lowers the proportion of survivors at 24 months from .48 to .40, a drop of 17 percent. For blacks, these variables have nearly negligible, statistically insignificant effects. For blacks, lower unemployment rates significantly hasten exits, decreasing the proportion of survivors at 24 months by almost 20 percent. For whites the coefficient on unemployment is less well estimated (although significant at 10 percent) but has an unexpected positive sign. One interpretation consistent with this result and other studies is that marriage markets may be important for whites, but the labor market is more important for blacks.

For both races, age has a well-estimated positive effect, while TBEN has a negative, but imprecise coefficient. For whites, a 10-percent rise in TBEN raises the proportion surviving at 2 years by 10

Table 4

AFDC Exit Rates: Estimated Hazard Rate Regressions
Complementary Log-Log Specification for Proportional
Hazard Model with Time Dummies

	Whites	Blacks
EDUC = 12	-.465 (.307)	-.258 (.367)
EDUC > 12	.348 (.342)	.516 (.473)
OTHINC ($1,000's)	-1.93 (2.06)	3.30* (1.64)
NKIDS	-.0984 (.132)	-.110 (.190)
YKID (=1 if present)	-.507 (.334)	.589 (.456)
AGE	.0629* (.0235)	.0756* (.0265)
UNEMP (%)	.156 (.889)	-.275* (.823)
AFDCU	-.0982 (.344)	.105 (.352)
SEXRATIO	3.11* (.732)	-.489 (.740)
EMPMALE	4.34* (1.93)	.0721 (1.75)
TBEN ($1,000's)	-2.96 (1.83)	-2.35 (2.79)
Log likelihood	-275.6	-197.4
Sample size		
Person-months	1,570	1,419
Persons	185	144

Notes: Standard errors are in parentheses. Sample consists of first observed spells by female-headed households with children. Cases likely to be income ineligible are excluded. Specification includes a constant and 6 time dummies for spell duration to that month: dummies for 3-4, 5-6, 7-8, 9-12, 13-16, 16+ months (see Table C-2).

*Significant at the 5% level.

Table 5

Simulated Survivor Functions

	Survivor Function at		
	6 Months	12 Months	24 Months
A. All races			
1. All persons (Table 3, Model 2)	.70	.57	.48
2. High Educ (>12)	.51	.34	.25
3. High school educ (=12)	.74	.62	.53
4. Low educ (<12)	.72	.60	.51
5. Age > 25	.63	.48	.38
6. Age ≤ 25	.78	.67	.59
B. Whites			
1. Base case (Table 4)	.67	.57	.48
2. TBEN increased 10%	.71	.62	.53
3. UNEMP decreased 10%	.70	.60	.52
4. SEXRATIO increased 10%	.60	.48	.40
4. EMPMALE increased 10%	.59	.48	.39
C. Blacks			
1. Base case (Table 4)	.74	.61	.51
2. TBEN increased 10%	.76	.64	.55
3. UNEMP decreased 10%	.66	.51	.40
4. SEXRATIO increased 10%	.75	.62	.52
5. EMPMALE increased 10%	.74	.61	.51

Notes: Simulated mean of the survivor function is based on estimated hazard from the indicated table. See text and note 14 for explanations. Means are for 329 cases.

percent; for blacks the effect is somewhat smaller. The level of OTHINC (property income plus private transfers) has a significant positive effect for blacks--it may indicate the existence of other options to welfare. For whites, OTHINC has a statistically insignificant effect.

Several other specifications were run to check the robustness of the results. Since income eligibility is a potential problem, I ran two additional versions of model 2 on different samples. One used a more restrictive definition of income eligibility, excluding anyone who had other income in excess of the family-size adjusted AFDC benefit level or who had gross income (earnings plus other income) in excess of 1.85 times the need standard of the person's state of Table A-1 residence and family size. This is shown as model in Table A-1. Model uses a less restrictive sample that included all female heads with children regardless of income eligibility status. Recall that the sample used in the text falls in between these two extremes in its restrictiveness.

The results in Table A-1 show the model to be robust with respect to the key variables SEXRATIO, EMPMALE, and TBEN. The coefficients on these variables generally improve in precision relative to model 2 of Table 3 for either the more or less restrictive sample. The coefficients on SEXRATIO and TBEN rise in absolute value. The remaining coefficients are remarkably stable across the three specifications, implying that the different income eligibility criteria do not alter the results. These results provide some evidence that the SIPP furnishes adequate data on AFDC recipiency for the female heads with eligible children.

IV. CONCLUSION

This paper uses a sample of first observed spells of AFDC receipt by female heads from the SIPP. Nonparametric estimates of survivor function based on these monthly data reveal that roughly one-third of AFDC spells will end within six months; two-fifths will last two years or more. I also find a significant and sizable negative effect of AFDC benefit levels on the length of time spent on AFDC. A more comprehensive measure of total welfare benefit, TBEN, shows a robust, but less well estimated negative effect. From the multivariate hazard model, I find that higher education and older age increase exit rates from AFDC.

Measures of spouse availability do affect exit rates when blacks and whites are examined together. The ratio of single employed males to single males, a proxy for spouse quality, has a significant positive effect on exit rates when entered along with a measure of the sex ratio. The SEXRATIO itself has a less significant positive effect. The residual effect of race becomes statistically insignificant. When hazards are run separately by race, the marriage market variables are important for whites but not for blacks. For blacks the unemployment rate matters. This result runs counter to the Wilson and Neckerman hypothesis, since differences in spouse availability do not matter for blacks.

Further work using a competing-risk framework, where exits from AFDC by marriage and by earnings are distinguished, could potentially help clarify the role of spouse availability. But when the competing risks are not independent, as seems likely here, identification of a model becomes problematic. Moreover, the extent to which better marriage prospects would affect women's human capital is not addressed here. One also wonders about the adequacy of the measures of spouse availability used here. Nevertheless, this paper uses more refined measures of the marriage market than have been tried before, and finds that they do matter.

Appendix A: Income Eligibility Checks

I checked two types of income eligibility. First, is the woman's other income (property income plus private transfers) larger than her AFDC benefit adjusted for family size? Second, is the woman's gross income (other income plus earnings) larger than 1.85 times the need standard (by family size) for her state? If all incomes, benefit levels, and need standards were accurately known, then a "yes" answer to either question would indicate ineligibility. Due to potential measurement problems, however, a more lenient standard was applied.

Three measurement problems are relevant. First, the AFDC benefit level and need standard were assumed to apply to a July-June fiscal year. Thus any intrayear changes within a state can lead to assuming the wrong benefit level or need standard for a few months. Second, property income is collected as a four-month aggregate in SIPP, then one-fourth of the amount is assigned back to each month. Thus, a large rise in property income in the last month of the SIPP reporting period could produce overestimates of the property income in the three previous months. This could produce the appearance, but not the reality, of ineligibility. Finally, with retrospective budgeting in the administration of an AFDC case, there is a lag between income determination and AFDC benefit payment.

The more lenient standard excludes cases that are likely to be income ineligible. These cases have other income in excess of $50 over the AFDC benefit level, or have gross income in excess of $100 over 1.85 times the need standard. This leaves 329 cases. Samples using a more restrictive criterion that excludes any possible ineligibles (leaving

304 cases) and a less restrictive criterion excluding no one for income ineligibility (345 cases) are used in the models of Appendix Table A-1.

Table A-1

AFDC Exit Rates: Complementary Log-Log Specification with Time Dummies: Different Income Eligibility Criteria Used

	Model 1 (Excludes Any Possible Income Ineligibles)	Model 2 (No Income Eligibility Exclusion)
EDUC = 12	-.290	-.152
	(.236)	(.208)
EDUC > 12	.501	.595*
	(.273)	(.240)
OTHINC ($1,000's)	1.38	1.18
	(1.42)	(.623)
NKIDS	-.0263	-.0605
	(.111)	(.0974)
YKID (=1 IF PRESENT)	-.356	-.250
	(.282)	(.226)
AGE	.0395*	.0395*
	(.0174)	(.0146)
RACE (=1 if black)	.260	.0749
	(.327)	(.271)
UNEMP (%)	-.281	-.365
	(.680)	(.545)
AFDCU	-.0364	.140
	(.242)	(.211)
SEXRATIO	.990*	.694
	(.501)	(.451)
EMPMALE	2.95*	1.95*
	(1.23)	(.956)
TBEN ($1,000's)	-2.79	-3.41*
	(1.47)	(1.30)
Log likelihood	-424.6	-528.2
Sample size		
Person-months	2,772	3,132
Persons	304	345
Complete spells	109	139

Table A-1, Continued

Notes: Standard errors are in parentheses. Sample consists of first observed spells by female-headed households with children. These are comparable specifications to Table 3, Model 2, but with different samples. Specification includes a constant and 7 spell dummies for for Model 1, 6 spell dummies for Model 2.

*Significant at the 5% level.

Appendix B: Computing Marriage Market Variables

To compute SEXRATIO from the 1980 Census I used the 1 percent sample for whites and the 5 percent sample for blacks. I included only noninstitutionalized civilians. For each state and race I computed the ratio of unmarried males to unmarried females by 11-year age groups as follows. For a woman age X, I divided the number of unmarried men age X-3 to X+7 by the number of unmarried women age X-5 to X+5. My census extract only included unmarried persons aged 18 through 54, so I adjusted the size of the groups at the endpoints to keep the same number of years for men and women. For example, for women age 18 the ratio is unmarried men age 20 through 25 divided by unmarried women age 18 through 23. For women age 19, the ratio is unmarried men age 20 through 26 divided by unmarried women age 18 through 24. Thus groups near the endpoints are less than 11 years, while groups in the middle (woman's age 23 through 47) are 11-year groups. These ratios were then assigned to women based on age, state, and race. Women younger than 18 were given the 18 ratio while women older than 54 were given the 54 ratio.

The employment ratios were computed using the same groups from the 1980 Census, then updated as follows: Let EMPSINGLE80 denote the ratio of employed single males to single males for a particular state, race, and age cell from the 1980 Census. To compute EMPMALE, I adjust this as follows:

$$\text{EMPMALE} = \text{EMPSINGLE80} \cdot \frac{\text{EMPCPS85}}{\text{EMPCPS80}}$$

where EMPCPS85 is the employment ratios for all men (regardless of marital status) computed for the same state, race, and age cells from the 1985 CPS. EMPCPS80 is computed similarly from the 1980 CPS. Thus I adjust the single employment ratio within each cell by a quotient reflecting the change in employment of the total male population. The state, race, and age cells are too small to use the CPS to directly calculate these measures for single persons.

Appendix C

Table C-1

Stepwise Hazard for Model 2 of Table 3:
Coefficients for the Time Dummies

	Estimate	Standard Error
Constant	-4.30*	1.44
Months 3-4	.246	.243
Months 5-6	-.111	.304
Months 7-8	-.393	.374
Months 9-12	-.361	.300
Months 13-16	-1.38*	.541
Months 17-20	-.686	.506
Months 20+	-2.02	1.06

Note: These coefficients are estimated simultaneously with the coefficients in Table 3, Model 2.

*Significant at the 5% level.

Table C-2

Stepwise Hazards for Blacks and Whites in Table 4:
Coefficients for the Time Dummies

	Whites	Blacks
Constant	-9.67*	-1.68
	(2.29)	(2.21)
Months 2-4	.511	.161
	(.330)	(.402)
Months 5-6	.212	-.220
	(.410)	(.508)
Months 7-8	-.991	-.475
	(.495)	(.669)
Months 9-12	-.253	-.138
	(.414)	(.468)
Months 13-16	-.910	-1.65
	(.644)	(1.18)
Months 16+	-1.15	-.625
	(.849)	(.632)

Note: Standard errors are in parentheses. These coefficients are estimated simultaneously with the coefficients in Table 4.

*Significant at the 5% level.

NOTES

[1]She finds roughly constant exit rates for the first eight months, declining rates from 8 to 18 months, and then a constant rate. Duration dependence is less strong when one controls for heterogeneity.

[2]O'Neill et al. do use the state average manufacturing wage as a proxy for potential husband's income, but they do not disaggregate this average by race or take other account of spouse availability.

[3]Guttentag and Secord (1983) present some time series evidence that low male/female ratios reduce the marriage rate, but they do not consider the quality of available mates.

[4]Blank makes this explicit when she considers that time spent on AFDC may decrease information about job opportunities. She cites this as a possible explanation of duration dependence.

[5]Fitzgerald (1987) develops a simple job search model where spouse availability affects labor market reservation wages. More generally, but not in an explicit search context, Grossbard-Shechtman (1984) links the marriage and labor markets in a model where spouse availability affects both the value of home production and market wages.

[6]See Lippman and McCall (1976) for a survey that shows the comparative statics for standard search models.

[7]The current 100% tax on earnings after 4 months on AFDC removes the incentive to stay on AFDC and work, at least in the short term. Nevertheless women may decide to stay on AFDC and work to maintain labor market experience, hoping to get long-term returns, or work to reduce the stigma of welfare receipt.

[8]Data on multiple spells are available for some cases, but not all. These multiple spells are necessarily short, since a person must get on AFDC, leave, and get back on within a 30-month period. The available multiple spell data are not exploited here.

[9]Persons who are paid weekly may get five paychecks in some months and lose their AFDC eligibility for that month, yet have average income that would always qualify them. Analogous reasoning applies for biweekly checks. Bernard Stumbras alerted me to this problem.

[10]Some of these are likely legitimate AFDC-U cases (AFDC program for unemployed two-parent households), while others may represent a misreporting of General Assistance or child support as AFDC.

[11]Blank used slightly different rules for recoding spells--she did not remove one-month gaps if she could identify a cause, such as an income increase.

[12]Disaggregating by SMSA would be useful in this context, but in SIPP not all SMSA residents are identified as such. Also small cell sizes for age and race would become a problem for sex ratios and employment ratios within SMSAs.

[13]The TBEN sums 70 percent of the AFDC guarantee, the food stamp guarantee, plus 36.8 percent of the insurance value of Medicaid. Only 70 percent of the AFDC guarantee is used, since food stamp benefits are reduced by 30 percent of the AFDC benefit. Smeeding (1982) estimated 36.8 percent as the conversion to the cash equivalent value of Medicaid. These data were kindly provided to me by Robert Moffitt and are discussed more fully in Moffitt (1988).

[14]I also ran a specification, not shown, that used the AFDC benefit level adjusted for family size. The results were unchanged except that

the coefficient on the size-adjusted benefit level is even more precisely estimated than in model 1. The AFDC benefit data was taken from the U.S. House of Representatives, Committee on Ways and Means, 1983, 1984, 1985, 1986.

[15]The hazard P(t) for each person was computed using the estimated coefficients and the beginning of spell covariates. These were then converted to survivor functions: $S(t) = \Pi_{j<t}P(j)$. The mean survivor function across persons was then computed for each t. Simulations of policy changes involved changing each individual's covariates and recomputing the survivor function. For example, to simulate raising benefits 10 percent, each person's benefits were raised 10 percent and the new survivor function computed.

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