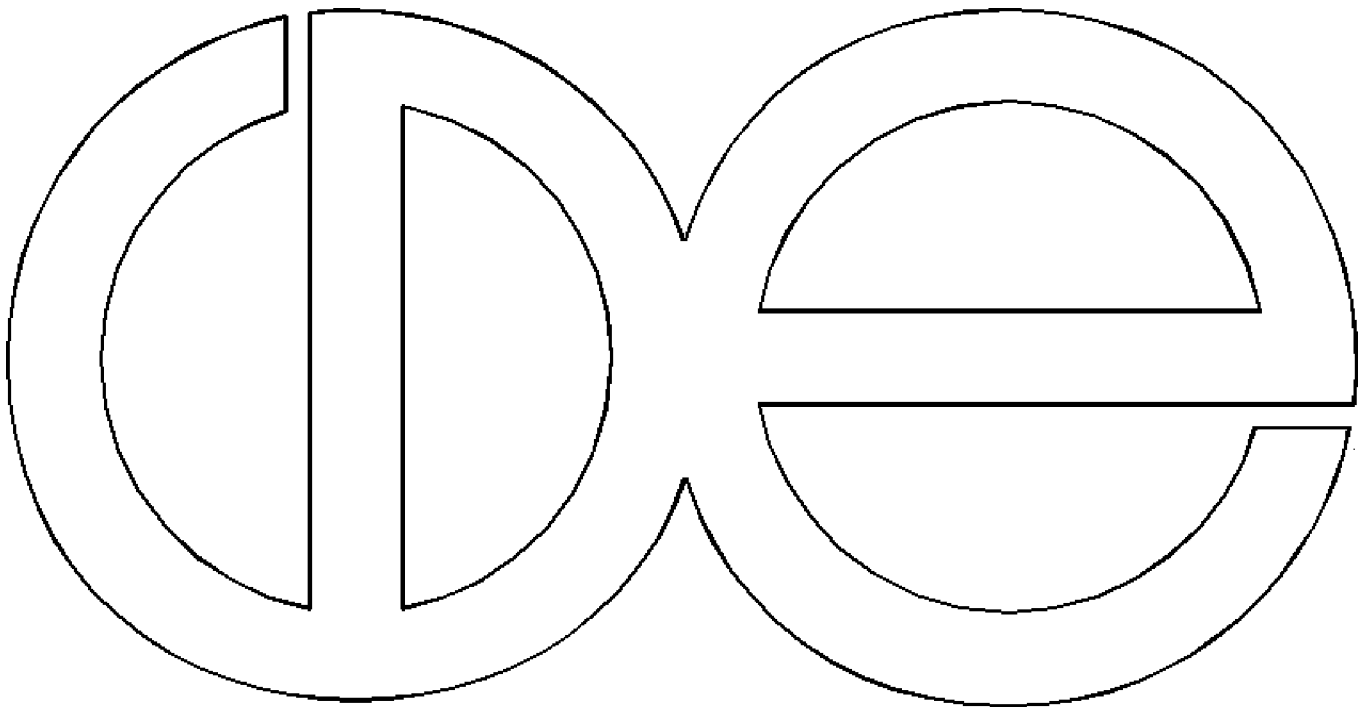# Center for Demography and Ecology

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Modeling Living Arrangements and Welfare Participation Choices Among Single Mothers

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ABSTRACT

Using data from the Survey of Income and Program Participation, this study analyzes the effects of AFDC benefits on living arrangements in combination with welfare participation. The study finds high rates of cohabitation and marriage among women receiving welfare. As well, results from the multivariate models indicate that the welfare system influences the decisions to cohabit and receive welfare. However, as the evidence partially depends upon the specific model adopted, results remain suggestive rather than conclusive. Nonetheless, the findings clearly show that, for women on welfare, cohabitation is a preferred alternative to marriage. If the relationship between the two living arrangements is ignored, estimates of the impact of the welfare system on living arrangements and welfare-participation decisions are potentially flawed.

1. Introduction

Family structure in the United States has radically changed over the last few decades. Two of the most notable changes in family structure are the large numbers of families that are now headed by single mothers and the large share of families headed by cohabiting (i.e., unmarried) couples.[1]

Isolating the causes for these two changes has proved a more difficult task than documenting them. Among ongoing efforts aimed at understanding why these changes have occurred is research evaluating the role that the Aid to Families with Dependent Children (AFDC) plays. Though the results are inconclusive, current evidence suggests that the program modestly encourages the formation of female-headed households and discourages marriages (Bishop, 1980; Danziger et al., 1982; Ellwood and Bane, 1985; Moffitt, 1990a; Hoffman et al., 1991; Winkler, 1993; Schultz, 1994).

With a few exceptions, most of the studies that have examined how the U.S. welfare system affects the growth in female headship and the decline in the traditional two-parent family have neglected the relationship between cohabitation and welfare participation.[2] Yet, as Moffitt, Reville, and Winkler (1995a) and Bumpass and Raley (1995) show, many single parents cohabitate with another adult who is sometimes the parent of their children, sometimes not. By definition, that cohabitation creates a two-adult household. Moreover, if the cohabiting adult is unrelated to the children in the household, the mother often receives welfare as well (Moffitt, Reville, and Winkler, 1995a).

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[1]In 1990, over 20 percent of families with children were female-headed households (U.S. Bureau of the Census, 1991a). That number is double the figure for 1960 (U.S. Bureau of the Census, 1961).

[2]Studies that have investigated the relationship include: Edin (1991), Gabe (1992), and Moffitt, Reville, and Winkler (1995a).

Apparently, cohabitation is a common alternative to marriage and female headship for many single mothers. However, the extent of cohabitation among single women with children and the effect of AFDC on the choice of cohabitation remains unresolved (Moffitt, Reville, and Winkler, 1995a).

This study examines these issues and builds on Moffitt, Reville, and Winkler's (1995a) research. In contrast to their study and others, however, I use an underutilized source of data and I adopt a particular modeling technique that, I argue, more realistically depicts single mothers' choices about living arrangements and welfare participation.

I conclude from the findings that cohabitation among low-income women is under-appreciated and that correctly estimating the impact of AFDC on choice of living arrangement requires more information than is currently available or presumed.

I organize the paper as follows. Section 2 summarizes recent research on cohabitation and marriage rates among women receiving welfare and reviews new findings on determinants of those women's living arrangements. Section 3 discusses modeling living arrangements and the welfare participation choice, describes a particular estimation approach, and overviews the data used in this study. Section 4 presents the findings, and the implications of the research are contained in Section 5.

2. Background

2.1 Cohabitation among AFDC Recipients: Evidence from Existing Data

Cohabitation and marriage rates are surprisingly high among women receiving AFDC. Moffitt, Reville, and Winker (1995a), henceforth abbreviated to MRW, used four data sets to highlight the rates of cohabitation and marriage. The four data sets they used were the National

Longitudinal Survey of Youth (NLSY), the Current Population Survey (CPS), the National Survey of Households and Families (NSFH), and the Panel Study of Income Dynamics (PSID).[3]

Using another data source, the Survey of Income and Program Participation (SIPP), I also find high rates of marriage and cohabitation among women receiving welfare. I compare my results with those reported by MRW (1995a) in Tables 1 and 2. (For details on the advantages and disadvantages of using SIPP see Section 3.2.)

The percentages in Table 1 for women age 18-55 resemble those from the MRW (1995a) study. Indeed, cohabitation rates across all five data sets display much uniformity. My findings strengthen MRW's (1995a) argument that "the AFDC caseload is more diverse than is usually perceived and contains significant numbers of both cohabitors and married couples." Moreover for several measures of cohabitation across age groups, the SIPP showed its superiority over the CPS. Its estimate of the proportion of women under 30, who are high school dropouts and with children under 18 closely paralleled those from the PSID and the NSFH, the latter having the most dependable estimates on cohabitation.

Marriage rates across the five data sets were also similar. The lower panel of Table 1 shows that a large share of women receiving AFDC were married. SIPP provides more evidence that marriage rates are quite high among women with less schooling--close to one-quarter.

Percentages in Table 1 are somewhat inaccurate due to measurement error, i.e., to produce results presented in Table 1, neither MRW's (1995a) study nor mine needed to exactly map the overlap between living arrangements and welfare receipt. However, by exploiting the panel nature

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[3]MRW (1995a) concluded that no one data set was superior to another, each had its own advantages and disadvantages.

of SIPP I could exactly identify monthly overlaps among cohabitation, marriage, and welfare receipt and thereby correct the measurement bias. Table 2 contains those results. (MRW did likewise using the PSID and NLSY panels.)

Table 2 shows that cohabitation rates from SIPP fall between those of the PSID and the NLSY, while marriage rates were higher than those reported for the PSID and the NLSY.

Such differences may reflect differences across sampling frames. As the SIPP, like the NLSY, includes all women, regardless of headship status, and the PSID only includes women who were heads of households or spouses, the PSID may undercount the proportion. And, as the NLSY only includes women aged 22-29, an age range in which cohabitation rates may rise above those of others, it may overcount the proportion. Nevertheless, once I restricted my sample to women aged between 18 and 40, marriage rates fell substantially, squaring better with rates from the PSID and the NLSY.

Results from the SIPP in Tables 1 and 2 echo two themes underscored by MRW (1995a). First, incorrectly measuring the timing of marriages or cohabitations with receipt of AFDC produces misleading estimates. Second, women receiving welfare often cohabit with unrelated males, the proportion among some of them reaching as high as 16.4 percent. Furthermore, I find that some of the women receive welfare while cohabiting in males' households, not their own.[4] Like MRW (1995a), I too judged that these high rates of cohabitation deserved more investigation.

2.2 Effects of AFDC on Cohabitation: Current Evidence

The rules of the AFDC program impose few penalties on women who receive welfare while cohabiting with men who are unrelated to their children (MRW 1995a; 1995b). Usually, the

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[4]Of the 1,626 women in the sample, 103 of them lived in households headed by the unrelated male. Among the 103, the number who received welfare was 35.

cohabitor's income, contributions towards rent, and gifts of cash do not affect the amount of a woman's AFDC benefit, as they would if the couple were married. Plainly, therefore, a monetary incentive to cohabit while on AFDC should exist. Except for MRW (1995a) and the present study, however, the literature has ignored this incentive effect.

If cohabitation is a viable alternative for women on AFDC, then existing models of the incentive effects of the AFDC program on living arrangements have been misspecified. Those models only considered the choice between female headship and marriage, when really the choice was among three alternatives: female headship, marriage, and cohabitation. Theoretically, relatively lenient rules about cohabitation suggest that the incentive to cohabit is increased relative to both marriage and headship.

To redress the misspecification problem and to better understand the effect of AFDC on cohabitation, MRW (1995a) estimated several discrete choice models. In one set of multinomial logit (MNL) models, which used data from the PSID and the NLSY, they examined determinants of "partner status" (cohabiting, married, or neither) and determinants of welfare participation. The results were consistent across both sources of data and mirrored findings in the literature (see Moffitt, 1992). In another set of MNL models, again using the same data, they combined partner status and the welfare participation choice into a single, six-category dependent variable to study determinants of those combinations.

Compared to the first set of MNL models that predicted partnership status, the second set of MNL models, which predicted partner status and welfare receipt, were more compelling. The second set detected some effect of welfare benefits on cohabitation for the samples drawn from both sources of data. (See Tables 4, 5, and 6 from MRW, 1995a).

The estimated effects of welfare benefits on living arrangements and welfare receipt suggested that the incentive effect of AFDC on cohabitation was plausible but needed further inquiry.

Notwithstanding differences across data sets, I possessed a comparable sample which also permitted estimating models analogous to those that MRW (1995a) estimated. Indeed, if the rules of the AFDC program truly encourage cohabitation relative to marriage and female headship, then the sample I possessed from SIPP should produce estimates corresponding to those that MRW (1995a) generated. Table 3 defines my variables, Table 4 details the descriptive statistics for the sample, and Tables 5 and 6 present estimates for a sample of women aged between 18 and 40.[5]

Based on data from SIPP, Table 5 shows MNL estimates for the three category partnership status variable and a binary welfare participation variable. Some of my findings are consistent with MRW's results, others are inconsistent. Like MRW (1995a), I find that black women have lower rates of marriage and cohabitation, that more educated women have lower cohabitation rates, and that those rates are higher among women with more preschool-aged children. Unlike their results, though, I found that marriage rates are lower among older women and women with more preschool-aged children and that higher benefits are not related to significantly lower marriage rates. In spite of the coefficient's statistical insignificance, I found, again contrary to MRW's (1995a) results, that welfare benefits increase rates of cohabitation, as predicted.

Results from the logistic regressions for welfare participation in Table 5 also agreed with those reported in the literature. (See Moffitt, 1992, for a review of findings on determinants of welfare participation.)

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[5]MRW's (1995a) sample had a broader age range than my sample. I restricted mine to women younger than 40, as it seemed that these women still considered remarriages, still cared for preschoolers, and were more inclined to cohabit.

Table 6 shows estimates of MNL models after I combined partnership status and the welfare participation choice into a single, six-category dependent variable. Like the findings in Table 5, results in Table 6 suggest that higher AFDC benefits are related to increased chances of receiving welfare and that in states with lax policies[6] about in-kind contributions by cohabitors rates of cohabitation, female headship, and marriage are increased among women receiving welfare. I suspect that the last model catches a subfamily effect as women living in subfamilies face the same rules as women cohabiting, yet not living in subfamilies. After that policy variable was added, effects of AFDC benefits strengthened.

The MNL models imply that the incentive effects of the AFDC program on living arrangements and welfare participation are veritable. However, the authenticity of the estimated effect is unclear, since these types of "discrete choice" models dismiss equal cross-substitution between any pair of living arrangement alternatives.[7] A general pattern of statistical dependence among living arrangement alternatives, if important, would decisively affect results from the MNL models, possibly challenging any effect attributed to the AFDC program. I examine this important methodological point so that there is some assurances that this modeling approach provides unbiased, consistent estimates.

Next, I discuss problems associated with the MNL model, I outline its implications for modeling cohabitation and welfare participation choices, and I supply an alternative to the MNL approach that still possesses the advantages of applied discrete choice modeling.

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[6]By lax, I mean the woman's AFDC eligibility was not affected in anyway by shelter contributions or cash assistance.

[7]The probit is free of this condition but is subject to computational intractability for more than three alternatives (Maddala, 1983; McFadden, 1982).

3. Empirical Framework and Data Description

3.1 Motivation for the Empirical Model

MNL models discussed in the last section were one-level (simultaneous) specifications assuming that the unobserved attributes of each alternative living arrangement and welfare participation choice were statistically independent of each other.[8] (See Figure 1.) However, if mothers choosing joint household headship (JHH) felt that cohabitation and marriage were more alike than different, then those similarities between the alternatives invalidate the joint MNL model of choice. This lack of correlation in the MNL model is largely responsible for its possessing a property known as the "independence from irrelevant alternatives" (IIA). What is needed is a discrete choice model--a "companion" to the MNL modeling approach--that would account for perceived similarity between alternatives.

To derive the model, first assume that each woman chooses the living arrangement that maximizes her conditional indirect utility. Also, partition the full choice set of living arrangements into K disjoint headship types H_k, $K = 1,...K$. Each arrangement is indexed by a double subscript (k,s) where K denotes headship type and s, $s \epsilon H_k$, denotes specific attributes of each living arrangement within a headship type. The indirect utility is[9]

$$V_{ks} = \overline{V}_{ks} + \epsilon_{ks} = Z_{ks} \beta + y_k \alpha + \epsilon_{ks} \quad k = 1, \ldots k, \ s \epsilon H_k$$

(1)

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[8]That is, unobservable components of indirect utility are independently and identically distributed.

[9]I only discuss the empirical specification of the random utility model. McFadden (1981) reviews the theory of the random utility model.

Equation (2) shows the MNL specification of Equation (1). It is obtained by assuming that the ϵ_{ks} are independently and identically distributed (iid), with the extreme-value distribution given in Equation (3).

$$P_{ks} = e^{V_{ks}} / \sum_{b=1}^{K} \sum_{m=1}^{S_b} e^{V_{bm}},$$

(2)

$$\text{Prob}[\epsilon_{ks} \leq \epsilon] = \exp(-e^{-\epsilon}).$$

(3)

Here, independence among the ϵ_{ks} is an unreasonable assumption, however, because a high utility from marriage should imply a high utility from cohabitation. Calculated under the IIA assumption, the MNL would underestimate the true probability of marriage relative to single motherhood as it ignores the event that preferring cohabitation to single motherhood makes preferring marriage to single motherhood even more likely.

I solve the problem and allow for reinterpretations of empirical results by recognizing that the joint choice model (2) can be rewritten as $P_{s|k}$ for living arrangement, given headship, and a marginal choice probability P_k for headship type. This "tiered" MNL is often called a "nested" MNL, or nested logit (NL). Figures 2 and 3 display these nested structures.

If the joint distribution of $\{\epsilon_{ks}\}$ is instead given by

$$F\{\epsilon_{ks}\} = \exp\sum_{k=1}^{k} - \left[\sum_{S\in H_R} \exp(-\epsilon_{ks}/\theta_k)\right]^{\theta_k} \quad 0 < \theta_k \leq 1$$

(4)

then the probability of choosing alternative s attached to headship type k is

$$P_k = P_{k|s} P_s$$

$$\text{where } P_{k|s} = \exp\{Z_{ks}\beta/\theta_k\}/\left[\sum_{s\in H_k} \exp\{Z_{ks}\beta/\theta_k\}\right]$$

$$P_k = \exp\{Y_k\alpha + \theta_k I_k\}/\left[\sum_{k=1}^{H} \exp\{Y_k\alpha + \theta_k I_k\}\right]$$

$$\text{and } I_k = \log\left[\sum_{s\in H_k} \exp\{Z_{ks}\beta/\theta_k\}\right]$$

(5)

If $0<\theta_k\leq1$ for all k, these probabilities can be derived from a random utility model in which V_k is the expected maximum utility for alternative k (McFadden, 1981).[10] This utility, Equation (1), is a linear function of observable characteristics, Z_k, Y_k, describing both the alternative and the individual making the choice. If $\theta_k = 1$, then few similarities among alternatives exist in each subset, which suggests collapsing the nest into a single MNL model. I_k is called the inclusive value corresponding to headship type k, and θ_k is the dissimilarity coefficient. The inclusive value I_k is a measure of the expected maximum utility of headship type k. The dissimilarity coefficient θ_k

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[10]That is, the partitioning is consistent with individual utility maximization.

measures the degree of correlation of $\{\epsilon_{ks}\}$, $s\epsilon H_k$ tends to an iid extreme-value distribution with choice probabilities given by the MNL model. When $\theta_k \to 0$ the $\{\epsilon_{ks}\}$ become perfectly correlated so that the conditional on choosing type H_k with probability (1) chooses the alternative $s^*\epsilon H_k$ that has the highest strict utility $Z_{ks}\beta$.

Consistent estimates of the parameters defining these probabilities are obtained by sequential maximum likelihood. In a two-tiered NL, if a sample member is superscripted by n, the alternative chosen by the member by k^n, and the corresponding node by s^n, then the log likelihood function is

$$L = \sum_n \log P^n_{k^n|s^n} + \sum_n \log P^n_{s^n} \equiv L_1 + L_2.$$

(6)

The dissimilarity coefficient θ_k appears in L_1 only through the estimate $\hat{\beta}/\theta_k$. The sequential estimator first estimates $\hat{\beta}/\theta_k$ by maximizing L_1 (first stage); it then uses this estimate to compute I_k and estimates $\hat{\theta}_k$ by maximizing L_2 (second stage).

The advantage of sequential estimation is that both L_1 and L_2 have the form of logit log-likelihoods with coefficients $\hat{\beta}/\hat{\theta}_k$ and $\hat{\theta}_k$, respectively. Hence each stage can be estimated with ordinary MNL estimation across alternatives.

3.2 Description of Data

The data I use come from the 1986, 1987, and 1988 panels of SIPP. SIPP is a longitudinal survey of a random sample of the U.S. population. The panels I use each contain four rotation groups spanning the period from October 1985 through March 1990. Each rotation group provides information on 24 or 28 consecutive months.[11] Each wave of the survey was conducted every four months, so each participant was interviewed three times a year about his or her monthly experiences over the previous four months. Thus, these data provide monthly information on household composition, labor market behavior, and income sources.

Combining the three panels yielded a sample of 12,017 black and white mothers caring for children under the age of 18.[12] About 90 percent of them (N = 10,743) reported no AFDC participation. The remaining 1,274 mothers reported AFDC participation for one or more months of the panel period. Furthermore, about 28 percent of the 12,017 (N = 3,370) reported a spell of single motherhood during the survey.

From the 3,370 mothers, I drew the sample I needed for the NL models. For that sample, numbering 1,626 mothers, I created a record of their experiences in the labor market, changes in their household composition, changes in their marital status, and shifts in their sources of income. Durations of jobs, occupations, and housing arrangements, as well as numbers of coresiding children, were added to information collected on their demographic attributes. Together, the variables portray the experiences of single mothers, many of whom cohabited or married while receiving welfare. As stated, Table 4 contains descriptive means for the group.

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[11]Rotation group 1 of the 1986 panel was followed for only 24 months instead of 28 months.

[12]Hispanics and other ethnic groups in the SIPP are excluded from these analyses.

Combining panels provided much data about a nationally representative sample of single mothers, some of whom choose to receive welfare. Still, because of the construction of SIPP, I was limited to 24 or 28 months of data on each individual and I lost a few individuals living in states that the SIPP had grouped together, (see Appendix B). The weakness of SIPP is that it disallows analyses of long-term welfare participation or cohabitation. Many cohabited or received AFDC when they were first interviewed (N = 287 and 198, respectively). Not knowing if these were extended spells or if they were one of many short spells may bias, to some degree, my estimates of AFDC participation or cohabitation.

In retrospect, each SIPP survey should have collected information on the timing and duration of past use of AFDC, even if the women were not currently receiving it. So, the trade-off is between either possessing richer amounts of monthly data on mothers, their children, earnings, and living arrangements in the short-term or gaining a longer perspective with incomplete data.

For cohabitation, the SIPP has advantages and disadvantages similar to those of the CPS. Like the CPS, SIPP has weak measures of cohabitation, leaving relationships among unrelated adult household members of the opposite sex undetermined. (See Appendix A for a description of the data collected on household composition in SIPP.) Nonetheless, SIPP shares with the CPS an advantage of possessing a large sample size that is representative of the U.S. population.

Unlike the CPS, however, and more like the PSID and NLSY, the SIPP is a short-term panel providing much information on individuals. Because, it contains much economic and demographic information on both adult males and females who share households, it permits me to identify the months in which cohabitations, marriages, and welfare spells coincide.

Notwithstanding its flaws, the monthly SIPP offers a unique opportunity to learn about the effects of AFDC on the living arrangements and the welfare participation choice of single mothers.

Besides using SIPP data, I utilize information on state AFDC benefit levels and industry-specific average hourly wages by state. I assembled those data and modified them so that they spanned the same period as the combined SIPP panels. (See Appendices B and C.) Integrating these multiple data sources resulted in one data source that contained state AFDC benefit levels and the timing of changes in mothers' AFDC and living arrangements.

4. Findings from SIPP Data

Using a hierarchical NL model to analyze the effects of AFDC on cohabitation among single mothers does not require showing that the IIA is violated (Maddala, 1983). However, since the IIA is such a restrictive property, testing its validity for a MNL model of single mothers' decisions about living arrangements and welfare participation is useful.

I test the IIA assumption applying Hausman and McFadden's (1984) procedure. First for both the three-category and six-category dependent variables, I estimate models with all choices. Then, for each dependent variable I estimate models with restricted sets of choices. Thus, I estimate models with restricted sets of alternatives while keeping the same covariates for each model. The test statistic is $\psi = (\Lambda_u - \Lambda_r)' [V_u - V_r]^{-1} (\Lambda_u - \Lambda_r)$.

Table 7 presents the first set of results from this procedure for the three-category dependent variable. It shows results for a truncated sample that excludes those mothers choosing to cohabit, i.e., restricting the sample to those who were deciding between female headship and marriage.[13]

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[13]Under the null hypothesis, the population displays the IIA property with degrees of freedom equal to the number of the elements Λ, thirteen in this case. The test is weaker when alternatives have no natural grouping but in this case they do.

As a result of conducting the IIA test on the three-category dependent variable, the critical value of the χ^2 statistic was 14.25, shown at the bottom of Table 7. It indicates that the Hausman and McFadden (1984) test cannot reject the IIA assumption for the MNL model of partner status; therefore the MNL model gives consistent but inefficient estimates of living-arrangement choice probabilities.

Table 8 displays results for the six-category dependent variable of welfare participation and living arrangement decisions. In this analysis, the sample is truncated to exclude those who choose marriage and not receive welfare, thereby leaving five categories. At the bottom of the table the critical value of the χ^2 statistic is 51.1 is displayed. Although again failing to reject the IIA assumption, the magnitude of this value suggests a possible underlying correlation among the errors which can be decomposed into its observable and unobservable components.

Given that the NL model decomposes the error structure into its observed and unobserved components (Amemiya, 1986) and since such an analysis produces inclusive values, (see Section 3), which can provide insights into the relationship among alternative living arrangements, I pursue this modeling approach.

Such analyses also permit me to contrast estimated coefficients from the NL model with those from equivalent binomial logits which ignore the underlying correlation across living arrangements. (Incidently, a comparison between results reported in Table 5 and those from the NL are invalid because the former contains results form a MNL model of partnership choice where the comparison (base) group is female heads.)

Table 9 contains results for the NL model of partnership status and results from the standard binomial logit equivalent.

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As drawn in Figure 2 and summarized in Table 9, the first step of the estimation sequence of the NL model yields consistent estimates of the parameters affecting the choice between cohabitation and marriage. (See Column 3 in Table 9.) This first stage suggests that AFDC benefits do not significantly encourage a woman to cohabit rather than marry. As I expect, the estimated coefficients from the NL model match those from the standard binomial logit, (Column 1, Table 9).

The more influential factors determining the type of joint living arrangement are demographic ones, such as her age, the number of preschool-aged children she raises, and her race. All increase the probability she cohabits relative to marrying. Two other variables, educational level and labor market attachment, are also stronger predictors of living arrangements than state AFDC benefit levels. However, these other two variables, along with the dummy variable for living in the South, ("South"), decrease the odds of cohabiting relative to marrying.

This first step of the NL model of partner status generates the inclusive value needed in the second step. Using the inclusive value to control for the possible unobserved similarities between marriage and cohabitation, the second step of the estimation procedure predicts joint headship relative to female headship. The estimated coefficient for the inclusive value provides information on the appropriateness of the specification, as well as an estimate of the relative gain from choosing cohabitation over female headship.

In Table 9, the sequentially estimated coefficient of the inclusive value, θ, is -0.17. Such an estimate is consistent with results from the IIA test in Table 7 which suggests an underlying correlation among the errors. Though statistically insignificant, the magnitude of θ is different from unity, suggesting that the NL model is not misspecified and that the inclusive value contains information about similarities across living arrangements. The sign of the coefficient indicates that

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choosing to cohabit relative to marriage is a "net lose" and that that lowers the relative odds of choosing joint headship over female headship.

After controlling for the correlation between cohabitation and marriage, the second stage of the NL model shows that increasing AFDC benefits are not significantly associated with higher levels of female headship. It is possible, however, that this statistically insignificant results is due to the correction for the error variance in the NL model. The result echoes some past findings showing no overwhelming evidence that the welfare system encourages female headship. (See Moffitt, 1992.)

Again, variables lowering odds of joint headship are age and race and labor market characteristics of mothers. Women who never worked are more likely to share a household with an unrelated male while those with either higher earnings, or higher household assets, are more likely to head their own households.

Tables 10 and 11 contain results for the NL model of living arrangement choice and welfare participation. In these tables, the six-category dependent variable from the previous table (Table 6) is decomposed into a discrete hierarchy of mutually exclusive decisions. (See Figure 3.) Table 10 shows estimated parameters at each stage of the NL estimation procedure whereas Table 11 only focuses on the estimated coefficients for AFDC benefits at each stage. I restrict Table 11 to only AFDC results so that I can show those results alongside ones produced from alternative binomial logit models that do not account for similarities across living arrangements.

The first three columns of Table 10 show that higher AFDC benefits increase welfare participation among females heading households, females cohabiting, and females marrying. Presumably, the reason why I find an positive and strong effect of welfare on the decision to marry is that the rules of the AFDC Basic program do not penalize a single mother for marrying a man who

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is unrelated to her children. In other words, with the exception of seven states (MRW, 1995a), the AFDC program does not automatically penalize a women for marrying and thereby providing a stepfather for her children.

Furthermore, Panel A of Table 11 confirms, as it should, that the NL and binomial logit models produce identical magnitudes for the effects of welfare on living arrangements and welfare participation choices. The correspondence occurs because this is only the first stage of the sequential NL estimation strategy which aims to produce inclusive values that test for dissimilarities across living arrangements at higher levels in the decision tree. (It would not make sense to produce inclusive values attempting to test dissimilarities across types of welfare participation.)

The same three columns of Table 10 provide evidence, the strongest of which being for female heads, that higher earnings discourage welfare participation. The third column of Table 10 also shows that among those same women heading families that those who never worked also received AFDC more often. Again, an expected result.

The NL model of welfare participation and choice of living arrangement generates other results in agreement with the literature. According to results in Table 10, welfare participation is lower for older women and for women who are more educated but higher for black women. Furthermore, the model indicates, like past MNL models and duration models, that women with greater household wealth have lower rates of welfare participation. (See Moffitt, 1992.)

Besides displaying estimated effects of the above-mentioned variables on the welfare-participation choice, Table 10 presents estimated effects of the variables on living arrangements, <u>with</u> corrections, (i.e., the inclusive values), for serially correlated errors across alternatives.

The estimated inclusive value for "Incvwelc" in column 4 of Table 10, (0.60), suggests that the welfare-participation choice, within the set of joint-headed households, are alike and that accounting for the cross-substitution between living arrangement and welfare participation choices is required. Once, I control for that relationship, I find that higher AFDC benefits across states are positively associated with higher cohabitation rates, as predicted. In this analysis, however, they are statistically insignificant. Panel B in Table 11 also shows that once "Incwelc" is added to the model, the magnitude of the "AFDC benefit" variable increases by over 50 percent, (0.03 to 0.05), though still remaining statistically insignificant.

Obviously at this second level of the nested structure I can examine the effects of other variables as well, (see Figure 3). Beyond choosing welfare less frequently, shown in columns 1 through 3, column 4 of Table 10 shows that women in the South are less inclined to cohabit, possibly reflecting a regional bias against this type of living arrangement. In addition, Black women are more likely to cohabit, as are women who never worked. Not only do higher earnings among women lower their chances of receiving welfare, (Columns 1 and 2), they also appear to increase their chances of cohabiting relative to marrying. The same argument holds for those living in Western states, ("West"), and those women who are older, ("Age").

The third and final level of the nested structure models the basic decision to form a joint-headed household. At this third stage of the analysis, I find two important pieces of evidence, one relating to the effects of AFDC benefits on living arrangement choice, the other relating to the underlying relationships among types of living arrangements.

Taking the latter first, the estimated coefficients for the two inclusive values in Table 10, "Incvmar" and "Incvwels", suggest that potential unobserved "common denominators" (i.e.,

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undetected violations of the IIA assumption) exist among living arrangements and welfare-participation choices and that accounting for such correlations may improve the model.

For instance, I interpret the sign of the coefficient for "Incvwels" to suggest that joint-headship will occur less often as the rate of AFDC participation among single mothers increases; on the other hand, however, the direction of the sign of the coefficient for "Incvmar" indicates that the rate of joint-headship increases if it means marriage. In other words, the net gain to well-being (utility) from choosing marriage over its alternatives increases the likelihood of joint-headship.

The other notable finding, contained in column 5 of Table 10 and in Panel C of Table 11, is that AFDC benefits are positively yet weakly associated with decisions to form joint-headed households with unrelated males. Intuitively, the rules of the AFDC program and higher benefit levels should increase the gains from joint headship relative to female headship, however, the estimated coefficients for "Kind", (-.02), and "AFDC benefit", (.04), provide little support for this hypothesis.

Finally, according to the results in the last column of Table 10, joint living arrangements less often occur among older women, Black women, unemployed women, and women with higher earnings. The dummy variable I added indicating lax AFDC rules concerning cohabitors ("Kind") indicates that lenient rules regarding shelter contributions encourage higher rates of welfare participation in particular living arrangements--specifically those characterized as either female-headed households or cohabitor-headed households but do not encourage a decision to cohabit instead of marrying or to form a joint-headed household, regardless of whether that joint-headed household is characterized by cohabitors or a married couple.

5. Conclusions

Understanding the determinants of living arrangements among single-mother families and the effects of the welfare system on those arrangements deserves more attention than it currently receives. Gaining knowledge about the intricacies of single mothers' living arrangements, how one type of living arrangement is related to another one, and single mothers' welfare participation choices is especially important as it becomes increasingly clear that any growth of the AFDC caseload now results from the growing number of female-headed families rather than from increased AFDC participation among female-headed families; that female headship does not lower work effort; and, that AFDC entries and exits stem more from changes in family structure than from changes in earnings or labor supply (Moffitt, 1992). This study has generated such added knowledge and it shows the inherent significance of family structure to the study of the welfare system.

My findings clearly imply along with previous ones reported by MRW (1995a) that cohabitation and marriage rates among women on AFDC are substantial. Furthermore, whereas MRW (1995a) could only provide weak evidence that AFDC benefits and states' AFDC rules towards cohabitors are related to the welfare participation decision, this study produces stronger empirical evidence verifying such a relationship. Indeed, these effects persisted even when these data were subjected to more rigorous empirical models that accounted for possible cross-substitution across similar living arrangements. Though perhaps the real contributions of those more complicated NL models are (1) demonstrating the difficulties in modeling the complex social decisions and (2) exposing potential biases in reduced-form MNL models.

As little is known about the connection between the welfare system and choice of living arrangement, policymakers should cautiously interpret my findings. My results nevertheless suggest

that a positive link between the two exists, though like other studies, mine fails to account for several important factors.

The largest weakness of the study is its failure to distinguish among relationships between the cohabiting men and women in the SIPP. Clearly establishing paternity would greatly assist in isolating the true effect of AFDC benefits and rules on cohabitation, as would better knowing the economic relationships among household members. A topical module which collects data at Wave 2 of the panel partially redresses this weakness but it does not overcome the problem for those who cohabited after that period, not does is help understand the permanency of relationships or income allocation within households (U.S. Bureau of the Census, 1991b).

Furthermore, full impact of state AFDC rules on the sample of households needs further verification. I take data reported by MRW (1995a) and use it in these models. However, as they forthrightly point out, the enforceability of these rules is dubious and their data cannot account for unreported income. Thus, this study only begins to clarify the relationship between living arrangements and the welfare-participation choice among low-income mothers. Much more remains to be explored.

HOW THE SIPP COLLECTS DATA ON HOUSEHOLD RELATIONSHIPS

SIPP collects monthly data and quarterly data on the relationship of each household member to the household reference person (HRP). All persons, including children, are expressed in terms of their relationship to the HRP. Family relationships among other household members are recorded and, together with relationships to the HRP, define families, subfamilies, unrelated subfamilies, and unrelated individuals. Data collected on family relationships also include marital status and designated parent or guardian.

If a person enters or leaves the household during the panel period (through birth, marriage, death, institutionalization, or migration), the date of the entry or exit is recorded so that most changes in household and family composition are pinpointed to the specific month. For more information see the SIPP Users' Guide (U.S. Bureau of the Census, 1991c). Exhibit 1 below is a facsimile of one of two questions recording data on household relationships. (I abbreviate household reference person to HRP.)

Exhibit 1. Facsimile of One of Two SIPP Items on Household Relationships

```
U All persons, including children
V      0 .Not a sample person this month
V      1 .HRP, living with relatives
V      2 .HRP, living alone or with only nonrelatives (primary individual)
V      3 .Spouse of HRP
V      4 .Child of HRP
V      5 .Other relative of HRP
V      6 .Nonrelative of HRP but related to others in the household--member of an
              unrelated sub-(secondary)-family
V      7 .Nonrelative of HRP not related to anyone else in the household (secondary)
              individual
```

APPENDIX B

THE STATE AFDC PANEL DATA SET

Panel data on AFDC benefit levels have one observation for each state and protectorate for the same time period as the SIPP panel data set: October 1985 until January 1990. For this period of time, I constructed a chronology of changes in each state's AFDC benefit levels by family size using information disseminated by the Administration of Children and Families (U.S. Department of Health and Human Services, 1992) and data published in the *Green Book* (U.S. House of Representatives, 1994). From these sources, I created variables that measured the length of time that each state's AFDC benefit level for each family size was valid, as well as variables that measured nominal levels of benefits for families of different sizes. Obviously, the procedure identified the months when states' AFDC benefit levels increased, if they did go up during the period when SIPP data were collected. The maximum number of benefit level changes in the five-year period is nine and the maximum family size is five.[14]

Before merging these data with the SIPP data, I deflated AFDC benefits into 1987 real dollar amounts. The deflator I chose was a rebased version of the Consumer Price Index, excluding food and energy, as reported in Table B-62 of the *1994 Economic Report of the President* (Council of Economic Advisors, 1994). This deflator, or its variants, have been commonly used (see Moffitt, 1992).

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[14]Data I possessed on state AFDC benefits went up to a family size of 14. However, the marginal increases in benefits above family sizes of five are paltry and my sample contained few single mothers who had families sized greater than five. For computing efficiency, I appended benefit levels for a family size of five to those few who had family sizes greater than five.

When joining these state-level panel data to the SIPP data, I again encountered the problem of several of the least populated states having been combined to protect respondents' privacy. I had to discard mothers living in those states (N = 35) because I was unable to correctly match AFDC benefit levels to their welfare and employment histories.

THE STATE HOURLY WAGE PANEL DATA SET

Panel data on average hourly real wages by state by industry include observations covering 50 states and the District of Columbia for the same time period as the SIPP panel.

To create these data, I used the Bureau of Labor Statistics (BLS) publication, *Employment and Wages: Annual Averages* (U.S. BLS, 1994b) for the years 1985 through until 1989. Except for the government sector, industry classifications match those of SIPP. For the government sector, I used average weekly hours reported in Table 6 of *Employment and Wages* for public administration workers in state government. I calculated an hourly wage measure by dividing average weekly wages by average weekly hours by industry at the national level. These data come from *Employment and Earnings*, another BLS publication (U.S. BLS, 1994a). Also, hours data for calculating hourly wages pertains only to private nonfarm payrolls; data on average weekly hours of work for farming and government workers are unavailable. As a proxy, average weekly hours for total private nonfarm payrolls has been used instead.[15]

Before merging these data with the SIPP data, I deflated average hourly wages into 1987 real dollar amounts, as described in Appendix B for state AFDC benefit levels.

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[15]The steps introduce measurement error into analyses but I do not believe that compromises results. First, few mothers reported working in the public sector and none worked in agriculture. Second, no published state/industry breakdowns are available for hours so I had to exploit national sources of data. I could have used state hours data for manufacturing only but the proportion of mothers in manufacturing was modest; thus, manufacturing hours would have been a poor proxy for all industries. Besides, manufacturing is heavily unionized relative to other industries and using state hours data for manufacturing would have narrowed variation in average hourly wages.

BLS standards on confidentiality and SIPP protections against respondent identification generated missing data when I merged these two sources of data. The problem of combining the least populated states to protect respondents' privacy has been discussed already. The BLS generated some additional missing data to prevent exposing small firms at a point in time within an industry and state. These observations were coded as -99 for missing. The problem caused no additional attrition from the sample.

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