ELSEVIER Journal of Public Economics 65 (1997) 89–117

JOURNAL OF PUBLIC ECONOMICS

Does welfare play any role in female headship decisions?

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Received 1 September 1995; received in revised form 1 October 1996

Abstract

Previous studies have examined whether the welfare system has contributed to the dramatic increase in single-parent families. This paper explores why the results in this literature are sensitive to the presence of state fixed effects. It considers one natural explanation, namely that the composition of the population differs across states in ways that are related to welfare program generosity. After controlling for individual effects the results provide no evidence that welfare raises the propensity to form female-headed households for either whites or blacks. These results illustrate the potential pitfalls of assuming that state factors are fixed over a long period of time. They also suggest that previous studies may have overstated the effect of welfare programs on family structure.

Keywords: Welfare; Family structure

JEL classification: I38; J12; J18

1. Introduction

During the last thirty years, the composition of families in the United States has changed dramatically. In 1960, less than 10% of families with children were headed by a single mother, while in 1990 more than 20% of families with children were (US Bureau of the Census, 1961, 1991a). These trends are common to both white and black families, although the increase among black families has been

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PII S0047-2727(97)00022-4

more dramatic. There have also been sizable increases in the number of out of wedlock births and teenage pregnancies among blacks and whites. In 1989, fully two-thirds of all births among blacks were to unmarried mothers (US House of Representatives, 1992).

One reason why this increase in the percentage of families headed by single women is of concern is because the economic well-being of single-parent families is typically below that of two-parent families. In 1990, the poverty rate among female-headed households with children was 45% compared to the rate of 8% among two-parent families (US Bureau of the Census, 1991b). Over half of all families in poverty are now female headed. Not only are female heads of household more likely to be poor in a given year, but they are also more likely to have longer spells of poverty (Bane and Ellwood, 1986) and are disproportionately represented among the persistently poor (Duncan, 1984). There is evidence that these trends are being transmitted through generations as female children raised in female-headed households are more likely to drop out of high school, have an out of wedlock birth, and become heads of household themselves (e.g. McLanahan, 1988).[1]

These trends have stimulated a large body of research exploring the potential explanations for these striking changes in family structure. Much of the research has focused on the role of the US welfare system on family structure decisions. The Aid to Families with Dependent Children (AFDC) program provides cash benefits primarily to single parents with children, and eligibility in the program is restricted to those families with both low income and low asset levels. In addition, all AFDC participants are eligible to receive benefits through the Food Stamp and Medicaid programs. Because these benefits are generally not available to two-parent families, it is argued that the US welfare system encourages divorce, separation and the delay of marriage and remarriage (for example, see Murray, 1984). Other explanations for the increase in female headship over the last two decades include the reduction in the number of marriageable men (resulting in sex ratio imbalances) through high unemployment, incarceration and mortality rates, the increase in female employment, and the fall in relative wages between men and women.[2]

An important feature of the US welfare system is that AFDC benefits are set at the state level and exhibit enormous variation across states. For example, in 1991, maximum benefits for a family of three ranged from $694 in California and $680 in Connecticut, to $288 in Indiana and $120 in Mississippi. Models of female headship are based on the proposition that a woman will choose female headship when the economic benefits (or utility) associated with female headship exceed the

[1] One should be careful in interpreting these events causally. It may be that those women who become female heads of household are more prone to poverty than those who marry, and that even if they had married, they might still have had higher propensity to be poor.

[2] See Wilson and Neckerman (1986) for a general discussion of these issues. Garfinkel and McLanahan (1986); Ellwood and Crane (1990) provide a summary of literature which explores the role of changes in labor markets on family composition.

benefits of non-headship. An important implication of this model is that women who live in states with relatively high AFDC benefits should be more likely to choose female headship then women who live in relatively low benefit states.

Studies examining the effects of welfare benefits on female headship commonly estimate cross-sectional regressions which rely on this interstate variation in benefits to identify the welfare effect (for example, see the review by Moffitt, 1992). As pointed out by Ellwood and Bane (1985), the estimated welfare effect in these studies will be sensitive to any omitted state variables which are correlated with both state welfare policy and female headship decisions. Ellwood and Bane stressed the importance of social norms, cultural effects and religious influences which are likely to be important in family structure decisions but also, through voter preferences, may affect welfare policy. To the extent that cross-sectional studies can not adequately control for these state variables, then appealing to standard omitted variable bias arguments, this approach will lead to biased estimates of the welfare effect. Furthermore, cross-sectional (and pooled cross-sectional) studies can not control for individual effects which may be correlated with the generosity of AFDC benefits through interstate migration and other changes in the sample composition over time.

In this study, I use data from the Panel Study of Income Dynamics (PSID) to estimate the impact of welfare on female headship and to assess the importance of the sources of bias discussed above. I estimate a model of female headship controlling for welfare benefits, characteristics of the woman, characteristics of the state, year effects, state effects and individual effects. Pooled cross-sectional data allow for the identification of a welfare effect and a state effect. Panel data, containing repeated observations on persons over time, allow for the identification of welfare, state and individual effects. There are two features in my data which permit identification of individual and state fixed effects. First, families may move during the panel: about 9% of the blacks and 16% of the whites in the sample move between states at some point. Second, over the two decades covered by the data, the sample of women in each state changes as different cohorts enter adulthood.

The results show that controlling for state and individual effects is very important for determining the importance of welfare in family structure decisions. In the cross-section, without controls for state or individual effects, there is a positive and significant relationship between welfare benefits and female headship for whites and blacks. As has been found in previous research (e.g. Moffitt, 1994), I find that including fixed state effects renders the estimated welfare effect for whites small and statistically insignificant, but does not affect the estimates for blacks. Once the model is specified to include individual effects, however, there is no evidence that welfare contributes to increasing propensities to form female-headed households for either whites or blacks. These results illustrate the potential pitfalls of assuming that state factors are fixed over a long period of time. They also suggest that previous studies may have overstated the effect of welfare programs on family structure.

The rest of the paper is organized as follows. Section 2 briefly reviews the existing literature on female headship. Section 3 presents the economic model and empirical implementation. Section 4 discusses the data used in the analysis. Section 5 presents the results. Concluding remarks are provided in Section 6.

2. Previous evidence on the role of AFDC on female headship

The early literature on the effects of AFDC on female headship is based primarily on state, SMSA, or city level analyses. The results from this literature are mixed and find no compelling evidence that AFDC has a significant effect on female headship decisions.[3] The more recent literature uses a variety of cross-sectional data sets and shows a significant and positive, but modest, effect of welfare on female headship (see, for example, Danziger et al., 1982; Ellwood and Bane, 1985; Moffitt, 1990a; Hoffman et al., 1991; Schultz, 1994; Winkler, 1993). These studies typically model the probability of being a female head as a function of individual characteristics, state welfare benefits, and, in some studies, other state characteristics. Danziger et al. (1982) use data from the Current Population Survey (CPS) and provide the first formalization and estimation of Becker's (Becker, 1973, 1974, 1981) model of marital formation. They estimate the earnings and income available in both marriage and female headship and find significant effects of AFDC. Schultz (1994) extends the work by Danziger et al. by modeling fertility as well as earnings and marital status. He finds a consistently positive effect of welfare on female headship for whites, but somewhat more mixed results for blacks. Cross-sectional studies based on a reduced form of Becker's model have also found significant effects of welfare on female headship (see, for example, Ellwood and Bane, 1985; Moffitt, 1990a; Hoffman et al., 1991; Winkler, 1993).

In each of these studies, the effect of welfare on female headship is identified by cross-state variation in welfare benefits. These estimates will be biased if there are any omitted individual or state level variables that are correlated with state welfare policy.[4] Ellwood and Bane (1985) point out that social norms, cultural effects and religious influences are likely to play an important role in family structure decisions, and are largely unobservable to the researcher. In addition, state preferences for welfare may also be affected by the state's economic, political and demographic condition, which may also affect family structure decisions. Besley and Case (1994) argue that if policy determinants are related to the variable of

[3] Reviews of this literature can be found in Groeneveld et al. (1983); Bishop (1980); Wilson and Neckerman (1986); Garfinkel and McLanahan (1986).

[4] While this paper focuses on the female headship decision as the outcome variable, other outcomes such as divorce, separation, out of wedlock births, and teenage pregnancy have been analyzed in this same framework. The discussion about the potential bias in the estimated benefit effect also apply to those studies. For a summary of the literature on the effect of welfare on female headship as well as other family composition decisions, see Moffitt (1992); Hoynes (1996b).

interest, their omission can lead to erroneous conclusions as to the importance of policy.[5] To the extent that cross-sectional studies can not adequately control for these state variables, then appealing to standard omitted variable bias arguments, this approach will lead to biased estimates of the welfare effect. For example, if the population in a given state believes strongly in the two-parent family, the state may not have much support for an AFDC program and, hence, it may offer low benefits. A state which is more accepting of non-traditional family structures may favor a higher level of support for female-headed households. In general, if the unmeasured effects are positively (negatively) correlated with welfare benefits then the estimated welfare effect will overestimate (underestimate) the true effect.

The determinants of state AFDC benefits have been examined in a few studies (e.g. Darity and Myers, 1983; Moffitt, 1990b; Moffitt et al., 1996; Plotnick and Winters, 1985). Plotnick and Winters (1985) consider the determinants of state-level AFDC benefits and find that higher benefits are associated with states with higher per capita income, higher density of the poor population, higher welfare recipiency rates, lower levels of illegitimacy among recipients, and lower food stamp levels. Moffitt (1990b) finds higher state income and lower food stamp benefits to be associated with higher AFDC benefits. Moffitt et al. (1996) argue that benefits may be driven by changes in wage opportunities for low wage workers in the state. While these studies suggest important variables to include as policy determinants, these results have not been incorporated in the empirical literature on the effects of welfare on family structure.

A natural approach to deal with the omitted state variables is to include state fixed effects in the regression equation. This, in general, can not be done with a single cross-sectional data set.[6] By pooling a series of cross-sectional data sets,

[5] Besley and Case (1994) derive the bias that results from omitting these policy determinants under several standard models in the literature (e.g., fixed effect models and difference-in-difference models). They point out that controlling for observable determinants of policy may not be sufficient since unobservable determinants may be correlated with the error in the outcome equation. They propose an instrumental variables approach to estimating the effect of policy on individual outcomes. In practice, they find it difficult to find good instruments for identifying the policy determination equation. There is an extensive literature examining the endogeneity of area policies. For summaries of this literature, see Besley and Case (1994); Poterba (1994).

[6] Ellwood and Bane (1985) generate intra-state variation in welfare benefits which then allows for the inclusion of state fixed effects even when a single cross-section is used. Specifically they include an estimate of *expected* welfare benefits for each person by adjusting the state level benefit by the likelihood of receiving AFDC benefits *if* single parenthood is chosen. Because this method relies on a sample of female heads of household to estimate the participation effect, any correlation between the AFDC participation decision and female headship decision could create a bias in the estimated participation probabilities. For example, if female heads of household are more likely to participate in AFDC for unmeasurable reasons, then this method will overestimate the participation probabilities for married women. While they considered many outcome variables (divorce, female headship, out-of-wedlock childbearing) welfare was found to have the largest effect on the probability of living independently.

however, one can estimate a welfare effect while controlling for a fixed state effect. Moffitt (1994) takes this approach by pooling over twenty years of cross-sections of the Current Population Survey and finds that ignoring state effects can lead to incorrect conclusions about the impact of welfare on female headship. Moffitt finds that adding state effects changes the benefit effect for white women from positive and statistically significant to negative and statistically significant. Interestingly, for black women, adding state fixed effects does not change the estimated benefit effect.

There also may be individual effects which are important determinants of family composition decisions, such as marriage and female headship, which are not observed by the researcher. If these effects are correlated with welfare policy, then this would introduce another source of bias. In a pooled cross-sectional model with fixed state effects, these individual effects may be correlated with the generosity of AFDC benefits through interstate migration and other changes in the cohorts of welfare eligible women over time. This would result in a correlation between welfare benefit levels and the distribution of the population with respect to the propensity to be a female head. This has not been examined in the literature.

3. An empirical model of AFDC and female headship

The fundamental theory of marital formation and dissolution was developed by Becker (Becker, 1973, 1974, 1981) and most empirical studies of family formation begin with some version of his model. Becker's model is based on the proposition that a woman will choose marriage when the economic benefits (or utility) inside marriage exceed the economic benefits outside marriage. His theory implies that marriage is particularly advantageous if there is specialization between the partners. That is, one partner specializes in market work while the other specializes in home production. Implications of this model are that increases in the earnings or wages of the potential spouse will increase the probability of marriage while increases in any benefits available outside marriage (such as welfare benefits) will decrease the probability of marriage.

Becker's model can be extended to look at the decision to become a female head of household. A female head is defined as a woman who has a child and is not married. This implies two routes into female headship: an unmarried childless woman has her first child, or a married woman with children becomes divorced, separated or widowed. Both of these routes may be affected by welfare, and, therefore, one should not condition on the presence of children. Non-headship consists of childless women and married women with children. Thus, we are interested estimating the determinants of *female headship*, not the determinants of *marriage*. Conditioning on the presence of children and analyzing the determinants of marriage would miss the impact that welfare plays on the initial marital status decision.[7] [8]

In the spirit of Becker's model, consider the determinants of the discrete choice of female headship. Let the utility function,

$$U(FH, W^{f}, W^{m}, B, X) \tag{1}$$

represent the maximum utility associated with choosing female headship ($FH = 1$) or non-headship ($FH = 0$). Maximum utility is a function of the woman's wage, W^{f}, her potential spouse's wage, W^{m}, welfare benefits, B, and the woman's characteristics, X. In choosing female headship, the woman loses access to the potential spouse's wages W^{m} but gains access to welfare benefits, B.[9] The woman then chooses the state with the highest utility.[10]

If FH^{*} is defined as the difference in the maximal utility between the two states, then the woman will choose female headship if FH^{*} is greater than zero:

$$FH^{*} = U(1, W^{f}, 0, B, X) - U(0, W^{f}, W^{m}, 0, X) \tag{2}$$

$$FH = \begin{cases} 1 & \text{if } FH^{*} > 0 \\ 0 & \text{otherwise} \end{cases} \tag{3}$$

In order to evaluate the probability that a woman chooses female headship, we need to know the woman's wage and her spouse or potential spouse's wage. However, the difficulty is that we observe a spouse's wage only if the woman is married and the spouse is working and we only observe a woman's wage if she is working. If we assume that wages are a function of the woman's characteristics (X) and labor market variables (L) then we can replace both wage variables by their determinants. Therefore, the effect of earnings of the woman and her

[7] I thank an anonymous referee for making this point.

[8] The approach taken in this paper is to look at the determinants of the female headship in a *stock* concept, as opposed to examining the determinants of *flows* into female headship. An advantage of the flow approach is that you control for the economic and welfare variables at the time the headship decision is made. In the stock approach, you typically relate current headship status to current economic variables, which may be a specification error. However, using the flow approach makes the inclusion of state and individual effects much more difficult. This is a point I will return to in the conclusion.

[9] In some circumstances, welfare benefits are available to two-parent families. While the empirical implementation will take this into account, the theoretical discussion assumes, for simplicity, that benefits are available only to single mothers.

[10] This is a myopic model of female headship where the utility is re-evaluated each period. A dynamic model of based on a search model is a natural extension that may be explored in future work.

potential spouse enter implicitly through their determinants.[11] While the results from this reduced form model can not be used to determine the importance of changes in the employment and earnings of men and women, they are appropriate to explore the role of welfare benefits in the presence of regional or state effects.

Assuming a linear form for the indirect utility function and adding an error term, the difference in utility becomes:

$$FH^*_{its} = \beta_0 + \delta_1 B_{ts} + \delta_2 UP_{ts} + \beta_1 X_{its} + \beta_2 L_{ts} + \nu_{its} \quad (4)$$

where the subscripts *its* correspond to individual i in period t living in state s. Welfare benefits are captured through two variables: state level AFDC benefits (B_{ts}) and a dummy variable indicating whether state s offered AFDC benefits to two-parent families in period t (UP_{ts}).[12] Higher welfare benefits are expected to increase the probability of female headship. All else equal, we would expect that by offering AFDC-UP benefits, the economic gain to being a female head of household would be reduced. However, the eligibility rules are more restrictive for AFDC-UP families, and, thus, the program is not on a par with the program for single mothers.[13] Labor market variables, which control for wage opportunities, are captured by L_{ts}.

The error term ν_{its} is specified as

$$\nu_{its} = \lambda_t + \gamma_s + \alpha_i + Z_{ts}\eta + \epsilon_{its} \quad (5)$$

where the λ_t are year effects, γ_s are state fixed effects, α_i are individual effects, and the ϵ_{its} are assumed to be *iid* errors. Year effects are included to capture any

[11] Alternatively, if one is interested in determining the importance of labor market factors, both W^f and W^m need to be estimated. The approach used in the literature is to estimate a wage equation based on the sample of spouses of married women and to use those estimates to predict spouse's wages for the entire sample of women (e.g. Danziger et al., 1982; Schultz, 1994; Hoffman et al., 1991). The covariates used to estimate the wage equation include characteristics of the wife and local labor market variables. Estimating earnings or income in the counter-factual state, however, can be problematic. We only observe the wage of spouse for those women who are married and if there are unobservable factors that affect female headship which also affect the earnings of the potential spouse, then this method will yield biased estimates for the wage estimates. For example, if women who are married have higher marriage opportunities, then we will overestimate spouses' wages for female heads of household. While fully accounting for this correlation in the unobservable components requires estimating a simultaneous model, a two-stage estimation method has been used (Schultz, 1994).

[12] Starting in 1961, states have had the option to provide benefits to two-parent families. During the time period covered by my PSID sample, about half of the states offered AFDC-UP benefits with some variation in state provision over time.

[13] Two-parent families must satisfy two conditions not required of single parents. First, the primary wage earner in the family can not work more than 100 hours per month. This hours limitation is the origin for the term 'unemployed' in AFDC-UP. Second, the primary wage earner must display previous 'significant' attachment to the labor force. Significant attachment is typically satisfied if the worker was employed and earned at least \$50 in at least six of the last thirteen calendar quarters, or was eligible to receive unemployment compensation sometime in the last year.

common trends in social norms and expectations or other determinants of marital decisions. The state effects capture time-invariant factors that influence female headship which are shared by all residents of the state such as characteristics of potential spouses, state support services or cultural influences. The individual effects capture the unobserved factors at the individual level that do not change over time. The state variables, Z_{ts}, represent the possible determinants of state AFDC benefits.[14]

If either the state effects or individual effects are correlated with the benefit variable B_{st}, then omitting γ_s or α_i will lead to a biased estimate of the welfare effect, δ_1. It has been shown that the state effects are correlated with benefits, and the substantive results change when they are included (see Moffitt, 1994). The focus in this study is to examine why state effects matter. One hypothesis that will be explored is that states differ in the composition of their population which is in turn correlated with the state benefit level. Since 'state' effects represent an aggregation of the preferences of the state residents, under this hypothesis controlling for individual effects should have the same impact on the estimated welfare effect as controlling for the state effects.[15] If families move over time, however, then the state effects will not necessarily capture the same influences as the individual effects. In addition, the sample of women in the states change over time as different cohorts enter the sample. If the composition of states change enough over the course of the PSID sample period then individual effects can be identified independently of state effects.

A second hypothesis is that state benefits are influenced by economic, demographic and political variables captured in Z_{ts}. If these variables are also correlated with female headship, then omission of these variables may also lead to a bias in the estimated welfare effect.

Most of the literature follows the approach in Eq. (4), but, because of the reliance on a single cross-sectional data set, does not identify the components of the error structure in Eq. (5). In the current application, the use of panel data allows for the identification of both state and individual effects.

A linear probability model (LPM) is used to estimate the female headship equation using the error structure in Eq. (5). The error components are estimated as fixed effects using standard panel estimation procedures (for example, see Hsiao, 1986). The LPM model is used because of the difficulty in estimating

[14] The specification of the error in Eq. (5) models the individual effect as fixed over the life cycle. While we might expect that an individual's propensity to be a female head may change as they age, the controls for observed characteristics should pick this up.

[15] This was Ellwood and Bane's interpretation of the omitted state effects. Another interpretation is that state effects capture other characteristics of the state's welfare program not captured by B_{ts} such as the availability of education and training services, conditions at the welfare offices, and so on. Evidence to be presented later discounts the importance of the latter interpretation.

probit or logit models with individual fixed effects.[16] The limits to using the LPM, however, are well-known (for example, see Maddala, 1983). The results are not sensitive to the linear assumption as estimates from a mixed logit model with state fixed effects and individual random effects provided similar conclusions.[17]

4. Data

The data used for this analysis are drawn from the Panel Study of Income Dynamics (PSID). The PSID is a longitudinal data set collected by the Institute for Social Research (ISR) at the University of Michigan which began in 1968 with a sample of about 5000 households containing 18 000 individuals. All members (and descendants) of these original survey families have been re-interviewed annually such that by the twenty second year of the panel, more than 38 000 individuals have participated in, or are currently participating in, the survey. All estimates presented here are based on the 1968–1989 (or Wave XXII) sample of the PSID. The original 1968 sample consists of two subsamples: a nationally representative subsample of 3000 households (Survey Research Center or SRC subsample) and a subsample of 1900 households selected from an existing sample of low income and minority populations (Survey of Economic Opportunity or SEO subsample). To adjust for this nonrandom composition, the PSID includes weights designed to eliminate biases attributable to the oversampling of low-income groups and to attrition. All results presented here use the weights provided by the PSID.

The estimation data set includes all women aged 16–50 who are either married or household heads.[18] [19] The dependent variable in the empirical analysis is equal

[16] Chamberlain (1980) shows that a fixed effects conditional logit model can be used to estimate individual fixed effects in a discrete choice model. The conditional likelihood approach implies that the fixed effects are identified by the *switchers* (for example, those who transition between female head and non-female head status) in the data set. In this application, the sample size of switchers is not sufficient to implement this approach.

[17] The mixed logit model can not be used to examine the hypothesis that omitted individual variables are driving the state fixed effect results because, by definition, the discrete distribution describing the individual effects is assumed to be independent of the covariates in the model.

[18] Because of data limitations subfamily heads are not, in general, included in the sample. For example, if a woman has a child and is living in her parent's household, she will not be in the sample. There are exceptions, however. If a woman leaves her parent's household to form her own, she will be included in the sample for the rest of the panel (even if she ever moves back in with her parents). This may lead to sample selection problems because the decision to leave the parent's household may be endogenous (e.g. Ellwood and Bane, 1985; Hutchens et al., 1989). Dropping younger women entirely (under the assumption that they are most likely to be in subfamilies) does not change the main results in the paper.

[19] The sample is limited to women 50 or younger in order to eliminate women who are unlikely to take up AFDC benefits. In addition dropping women over age fifty reduces the cases where an older women may appear to have a child when it is actually her grandchild. The results are not sensitive to this age selection.

to one if the woman is a female head (she has a child and is not married) and zero otherwise.[20] An observation is created for each year that the woman satisfies this sample selection condition. The sample is further restricted by dropping observations for women who reside in small states.[21] Dropping small states may be important in order to identify state and individual effects simultaneously. In practice, however, this restriction is not particularly important, and in extensions of the basic model I present estimates based on the full sample. The estimation data set contains a total of 69 981 observations for 4211 white women and 3332 black women over the twenty-two year period.

The PSID data are matched with state level data on welfare benefits, economic, demographic and political variables. Two variables are used to describe the welfare benefits available in each state in each year. First, the generosity of welfare benefits are measured as the combined value of benefits from AFDC and Food Stamps for a family of four with no other income. While the above discussion has used the terms 'welfare' and 'AFDC' interchangeably, an important feature of the US welfare system is that all AFDC recipients are automatically eligible to receive Food Stamps and Medicaid, and participation rates among AFDC participants in these programs are over 90% (see US House of Representatives, 1992). Because Food Stamp benefits are calculated taking into account the family's AFDC grant and the maximum Food Stamp grant does not vary by state, the availability of Food Stamp benefits reduces the interstate variation in welfare benefits. The argument for using the combined value of AFDC and Food Stamps is further supported by evidence that states take into account the magnitude of Food Stamp benefits when setting AFDC benefit levels (see Moffitt, 1990b) and that Food Stamp coupons are valued by consumers at approximately face value (see Moffitt, 1989).

To test the sensitivity to this choice of benefits, we also present estimates using only AFDC benefits as well as the combined value of AFDC, Food Stamps and Medicaid. The case for including Medicaid with cash benefits is a difficult one because of the problems inherent in determining its cash equivalent value. Based on examination of periods of expansion in the Medicaid program, recent evidence shows that the availability of Medicaid may influence family structure decisions (see Decker, 1995; Yelowitz, 1995). As a second measure of state welfare benefits we include a dummy variable equal to one if the state had an AFDC-UP program

[20] Women who are cohabitating are assigned to the non-female head or married choice. Female heads include only those living without a spouse or partner.

[21] A state was included in the sample if there were observations for at least 75 women over the panel for whites or at least 50 women over the panel for blacks. These states account for about 78% of white and 88% of the black families with children. There are a total of 21 states in each of the black and white samples, although because of racial differences in geographic distribution of the population, they share 15 states in common.

in the particular year.[22] The state-level economic variables include the unemployment rate, average wage in manufacturing and per capita income. Demographic variables include the percentage of the population that is over age 65, and the percentage that are children. The political variables include the party of the governor, and the party composition of the state senate and house.[23]

Weighted statistics for the sample data are summarized by race and female headship status in Table 1. There are a total of 40 239 observations for whites and 29 742 observations for blacks. Over the sample period, about 9% of white women in the sample were female heads compared to 39% of blacks. For both races, female heads of household are more likely to have lower education levels and are younger than non-female heads, and for blacks, more likely to live in an SMSA. Overall, black women are more likely to have lower education levels and larger families than whites. Religion of the head, which may be correlated with headship status, is provided in the PSID. The majority of black women are Baptist while white women are more likely to be Protestant or Catholic. Female heads of household, especially blacks, are more likely to live in areas with higher unemployment rates, higher wages, higher state income and a greater Republican party presence in the state government.

Before proceeding to the results of the regressions, consider the simple time series trends in program generosity and female headship. Fig. 1 presents trends in real benefits in the AFDC, Food Stamps and Medicaid programs over the last 25 years. The line labeled 'AFDC' shows the maximum AFDC benefits for a family of four, averaged over all states. The maximum combined benefit from AFDC and Food Stamps is shown by 'AFDC and FS'. Finally, the generosity of Medicaid is measured by average state Medicaid expenditures among female headed households, shown by 'Medicaid'.[24] The most striking fact in this figure is the dramatic decline in AFDC benefits since late 1960s. The real value of the AFDC guarantee dropped by almost 50% during this period, with benefits continually in decline, aside from the 1982–1988 period when benefits were largely unchanged. The

[22] Over the time period covered by the PSID, there was only minor variation in the number of states participating in the AFDC-UP program (see Hoynes, 1996a). In 1968, 21 states offered AFDC-UP benefits. Through the mid-1970s state participation increased, then decreased in the early 1980s. State participation has increased since the early 1980s. The states that offer AFDC-UP benefits tend to be higher benefit, higher caseload states. The Family Support Act of 1988 mandates that all states extend AFDC-UP benefits by 1990.

[23] These political variables were generously provided by Anne Case.

[24] The combined benefit from AFDC and Food Stamps is equal to 70% of the maximum AFDC benefit plus the Food Stamp maximum benefit reflecting the fact that AFDC income is taxed in calculating the Food Stamp benefit. The AFDC data came from unpublished tables from the Family Support Administration, Department of Health and Human Services. The food stamp data came from unpublished tables from the Food and Nutrition Service, Department of Agriculture. Medicaid benefits are average benefits by state for a family of four and were provided by Robert Moffitt. All values in Fig. 1 are weighted averages of the relevant state variable, using the state's AFDC caseload as the weight.

Table 1
Means of PSID sample by race and female headship status

	Whites		Blacks	
	Female Head	Not Female Head	Female Head	Not Female Head
Age	33.815	34.173	32.444	34.172
	(8.183)	(8.804)	(8.064)	(9.123)
Education				
<9 years	0.073	0.041	0.066	0.088
9–11 years	0.293	0.135	0.361	0.271
12 years	0.410	0.479	0.387	0.401
>12 years	0.219	0.342	0.177	0.233
Presence of children	1.000	0.645	1.000	0.602
Catholic	0.255	0.287	0.057	0.071
Baptist	0.170	0.150	0.655	0.620
Protestant	0.375	0.348	0.183	0.209
Jewish	0.019	0.048	–	–
Other religion	0.054	0.057	0.020	0.024
No religion	0.128	0.109	0.085	0.075
SMSA	0.627	0.641	0.775	0.732
AFDC and FO guarantee	578.0	580.4	523.0	512.7
	(130.9)	(134.2)	(138.0)	(137.5)
AFDC guarantee	471.2	474.9	395.9	380.9
	(175.9)	(179.3)	(186.5)	(185.2)
AFDC-UP	0.689	0.676	0.572	0.460
State unemployment rate	7.023	6.630	7.066	6.620
	(2.281)	(2.226)	(2.237)	(2.310)
State average wage	8.963	8.841	8.615	8.293
	(1.236)	(1.212)	(1.386)	(1.468)
State income per capita (1000)	11.709	11.574	11.428	10.852
	(1.687)	(1.785)	(1.988)	(2.049)
State pop. over 65 (%)	0.114	0.114	0.110	0.108
	(0.020)	(0.020)	(0.018)	(0.020)
State pop. kids (%)	0.211	0.215	0.212	0.220
	(0.028)	(0.029)	(0.028)	(0.031)
State pop. black (%)	0.124	0.123	0.193	0.201
	(0.055)	(0.057)	(0.113)	(0.107)
Rep. state house (%)	0.397	0.384	0.329	0.283
	(0.129)	(0.144)	(0.168)	(0.176)
Rep. state senate (%)	0.402	0.390	0.332	0.282
	(0.158)	(0.170)	(0.193)	(0.198)
Rep. Governor (%)	0.540	0.528	0.473	0.418
Weighted percent of sample (by race)	9.38%	90.62%	38.5%	61.5%
Number of observations	3113	37 126	10 242	19 500

Notes: Sample is comprised of women aged 16–50 from the 1968–1989 Panel Study of Income Dynamics. A female head is a woman who has a child and is not married. See text for details. Standard errors in parentheses.

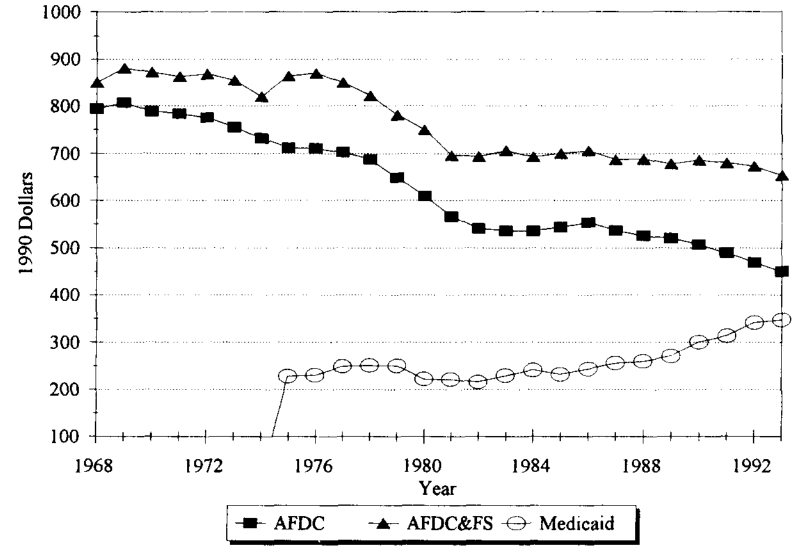

Fig. 1. Maximum welfare benefits for a family of four, 1968–1993 (1990 dollars).

introduction of in-kind benefit programs in the late 1960s and early 1970s moderated the decline in AFDC benefits in the early part of the period. The cash value of AFDC and Food Stamp benefits declined by about 30% over the period. This is in part due to the fact that Food Stamp benefits are adjusted annually for changes in food prices, where changes in AFDC have to be authorized by state legislatures. Despite the fact that real wages have also declined over much of this period, benefit-to-wage ratios exhibit similar trends to real benefits shown in Fig. 1 (see Hoynes and MaCurdy, 1994). Average state Medicaid expenditures for female-headed households have increased somewhat over the period, which, if valued by households as cash, would further moderate, but not reverse, the fall in AFDC benefits.[25]

Fig. 2 shows female-headed households as a percent of all families with children over the period 1968–1993. In 1968, about 8% of white families with children were headed by a single mother, while in 1993 almost 17% of white families with children were female-headed households. These trends are even more dramatic for black families where the rate of female headship increased from about 30% in 1970 to over 50% in 1993. Comparing the trend in benefits to the

[25] If the value of Medicaid to families is equal to the average expenditure then the combined benefits in the three programs increased somewhat up until the mid 1970s, declined until the late 1980s, and increased somewhat at the end of the period.

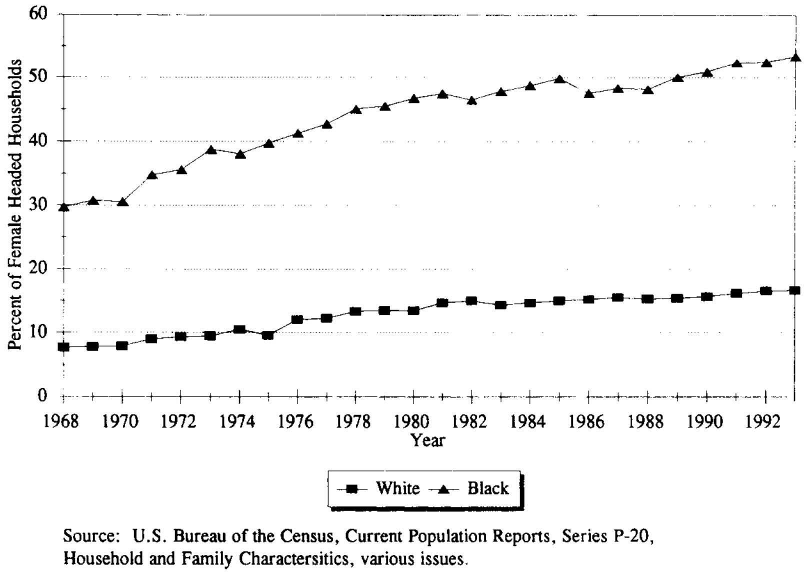

Fig. 2. Female headed households as a percent of all families with children, 1968–1993 (by race).

trend in female headship, it appears that benefits tracked female headship quite closely until the mid-1970s. Since then, real benefits have declined while the headship rate has increased. Clearly, however, there may be other factors that may have changed since the mid-1970s; the simple time-series correlation therefore does not rule out a role for welfare.

In the state fixed effects model, the welfare effect is identified by within state variation in benefits over time. An examination of state trends in AFDC benefits (not shown here) shows that there are substantial differences in the trends across the states. Illinois was a very high benefit state in 1968 (4th highest ranked state) and is now one of the lowest benefit states outside of the south (ranked 22nd overall). California was an average state in 1968 (ranked 23rd) but has risen to become one of highest benefit states by the end of the period.[26] Texas has fallen from modest benefits (ranked 12th in 1968) to one of the lowest benefit states in the country (ranked 3rd from the bottom above Alabama and Mississippi). An analysis of variance, however, shows that a permanent state component is by far the largest contributor to the variance of AFDC benefits.

[26] During much of this period, California was the only major state where AFDC benefits were automatically adjusted for changes in the cost of living. Typically, benefits are set nominally by state legislatures and benefits remain fixed in nominal terms for periods of several years.

5. Results

This section presents estimates for the female headship model described in Section 3. The dependent variable is equal to one if the woman is a female head, and equal to zero otherwise. Because headship patterns differ quite substantially for blacks and whites (see, for example, Ellwood and Crane, 1990; Danziger et al., 1982; Hoffman et al., 1991; Moffitt, 1990a, 1994), separate equations are estimated for white and black women. To account for non-random sample composition, all regressions are estimated using the sample weights. Standard errors are adjusted for arbitrary correlation over time using the correction of Huber (1967).[27]

5.1. Estimates for white women

The main estimates for white women are provided in Table 2. The bottom panel of the table describes whether time, area or individual effects are being controlled for in each regression, and presents the implied elasticity of the probability of female headship with respect to the welfare benefit, evaluated at the mean. The model in Eq. (1) provides estimates for the basic model which includes welfare benefit variables, characteristics of the woman, labor market variables, division dummies and time effects. This is the standard specification used in the cross-sectional studies. Consistent with recent evidence, welfare benefits have a positive and significant, but modest, effect on female headship for white women. The results imply that if the combined value of AFDC and Food Stamps increases by $100, female headship would increase by 0.9 percentage points, or an increase of about 10% generating an elasticity at the mean of 0.56. As expected, living in a state that offers AFDC-UP benefits is estimated to have a negative effect on female headship. The other covariates included in the model show that female headship is higher for younger women, who have lower education levels, and live in urban areas with higher unemployment rates and higher wages. Female headship also varies by religious affiliation: relative to the omitted group of Catholics, Jewish women have lower propensities to be female heads of household while other groups have higher propensities. The effects of education and living in an urban area are particularly large. White women with less than a high-school education are more than twice as likely to be female heads than high-school

[27] The linear probability model was also estimated accounting for heteroscedasticity due to the use of a discrete dependent variable in a linear model (see Maddala, 1983). In order to calculate the weight to correct for heteroscedasticity, observations with probabilities outside the (0,1) range have to be dropped or set to predefined constants. Dropping observations outside the (0,1) range yielded somewhat non-robust results across specifications due to differential sample selection. Using an arbitrary constant to form the weights for these observations with predicted values outside the (0,1) range yield results very similar to those reported here. Both sets of estimates are available from the author.

Table 2
Parameter estimates for female headship model (white women)

	(1)	(2)	(3)	(4)
AFDC and FO guarantee (100)	0.009	−0.001	0.0002	0.002
	(0.002)	(0.003)	(0.002)	(0.003)
AFDC-UP	−0.008	−0.008	−0.006	−0.009
	(0.004)	(0.007)	(0.005)	(0.006)
Age	−0.001	−0.001	0.003	0.003
	(0.0002)	(0.0002)	(0.002)	(0.002)
Education 9–11	0.012	0.014		
	(0.008)	(0.008)		
Education 12	−0.092	−0.085		
	(0.007)	(0.007)		
Education >12	−0.115	−0.107		
	(0.007)	(0.007)		
SMSA	0.008	0.003	0.011	0.015
	(0.003)	(0.003)	(0.004)	(0.004)
Baptist	0.025	0.022		
	(0.005)	(0.005)		
Jewish	−0.019	−0.018		
	(0.007)	(0.008)		
Protestant	0.018	0.014		
	(0.004)	(0.004)		
Other religion	0.013	0.010		
	(0.007)	(0.007)		
No religion	0.021	0.016		
	(0.005)	(0.005)		
Unemployment rate	0.007	0.005	0.003	0.002
	(0.001)	(0.001)	(0.001)	(0.001)
Average wage	0.004	0.007	0.002	0.001
	(0.002)	(0.005)	(0.003)	(0.003)
Intercept	0.038	0.059		
	(0.018)	(0.038)		
Year dummies	Yes	Yes	Yes	Yes
State dummies	No	Yes	No	Yes
Indi. fixed effects	No	No	Yes	Yes
Elasticity of Pr(FEMHD) wrt				
Welfare benefit	**0.56**	**−0.06**	**0.01**	**0.12**
Mean of FEMHD	0.094	0.094	0.094	0.094
No. of observations	40 239	40 239	40 239	40 239

Notes: Dependent variable is one if woman is a female head. Sample is based on 1968–1989 PSID. Estimates are based on linear probability model. Standard errors are in parentheses.

graduates. Living in an SMSA almost doubles the probability of female headship.[28] Time effects are included to control for changes in social norms, and show a consistent upward trend. These time effects are jointly significant at the 1% level.

[28] In general, one should take care in interpreting these effects as causal. Female heads may prefer to live in SMSAs because of availability of services, jobs or other factors.

The remaining models in Table 2 examine the role of state and individual effects. Adding the fixed effects for the states, as presented in Eq. (2), substantially changes the estimated welfare effect. These results imply that the unmeasured state effects are quite influential as the estimated welfare effect changes from being positive and statistically significant to being close to zero and statistically insignificant. These state effects are jointly significant at the 1% level. To explore what causes this reversal, Fig. 3 plots the estimated state fixed effects against average welfare benefits for each of the states in the sample. The decline in the estimated welfare effect is a result of the state effects being *positively* correlated with state welfare benefits for female heads of household. For example, high benefit states such as California, Minnesota and New York have high benefits and relatively large state effects. The correlation coefficient for these series is 0.36. These results are consistent with the idea that unmeasured state effects influence white headship decisions and welfare benefits. For example, a state may have a strong two-family tradition which results in fewer female headed households and less support for the AFDC program. Not taking into account state effects attributes this difference in preferences to a welfare incentive.

The remainder of the table provides strong evidence that for whites, state effects are capturing differences in the composition of the population across the states. Adding individual fixed effects, as shown in Eq. (3), has virtually the same impact on the estimated welfare effect as adding state fixed effects—the coefficient on AFDC benefits becomes small and insignificant.[29] Once individual effects are

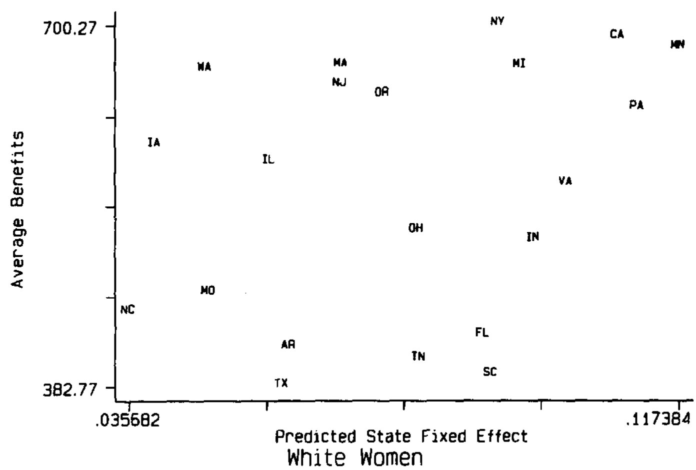

Fig. 3. Correlation between state fixed effects and average welfare benefits.

[29] In the regressions with individual effects, individual covariates with minimal time variation such as religion and education are dropped.

included, adding state effects does not significantly change the parameter of interest. The state effects are still important (e.g., they are still jointly significant at the 1% level with the individual effects present in the regression) but they do not influence the estimated welfare effect. This result is good news for applications when only pooled cross-sectional information is available in that controlling for fixed state effects is sufficient to account for population heterogeneity across states. As a side point, controlling for area effects larger than the state level do not appear to be sufficient to handle the issue of omitted state variables. Adding dummies for the nine census divisions does not change the estimated welfare effect relative to model Eq. (1).

It is worth considering more formally why we might expect state and individual effects to have a different impact on the estimated welfare effect. Suppose that the PSID was a balanced panel and that no families moved over state lines over the course of the panel. Then controls for state of residence, through state fixed effects, would have the same influence on the welfare effect as the individual effects. This is because welfare benefits vary only by state and, in the event that no households move across states, all that we have to identify the state effects is the individual effects of its residents. In that case, we would expect the specifications in Eqs. (2) and (3) to have the same coefficient on welfare benefits. In other words, the correlation between welfare benefits and state effects would be equal to the correlation between benefits and individual effects. We do, however, observe families moving between states in the PSID data. Among the sample of female heads of household, about 9% of black and 16% of whites move at some time over the two decades covered by the PSID. Further, the panel is not balanced. The composition of a state changes over the twenty year period as different cohorts of women enter adulthood. These two factors imply that the composition of the population within a state is not fixed, and, therefore, individual effects and state effects can have different impacts on the estimated welfare effect.

In addition, the models in Eqs. (1) and (2) were estimated adding the state level economic, demographic and political variables in Table 1. These estimates, which are presented in Hoynes (1995), imply that higher rates of female headship are found in states with proportionally fewer children and fewer elderly. The political variables (dummy for Republican governor and proportion of state houses held by Republicans) do not appear to be important. Adding these demographic and political variables as additional policy determination variables did not change the estimated welfare effect. Thus, despite their ability to explain variation in female headship patterns (they are jointly significant) the state variables are not highly correlated with AFDC benefits. Thus, omitting them does not bias the estimated welfare effect.[30]

[30] As noted by Besley and Case (1994), even if the omission of these state level determinants of AFDC benefits do not bias the estimated welfare effect, correlation between the unobservable elements could still lead to a bias. This is unlikely due to the lack of evidence based on observable influences. However, attempts to find an instrument for AFDC benefits were not successful.

5.2. Estimates for black women

The main results for blacks are presented in Table 3. The estimates in the model in Eq. (1), without geographic controls, show that the determinants of female headship for blacks differ somewhat from the estimates reported for whites. The benefit effect is significantly larger among blacks than it is for whites. For blacks, a $100 increase in welfare benefits increases the headship probability by about two

Table 3
Parameter estimates for female headship model (black women)

	(1)	(2)	(3)	(4)
AFDC and FO guarantee (100)	0.019	0.011	−0.002	0.007
	(0.003)	(0.006)	(0.004)	(0.005)
AFDC-UP	0.051	0.077	0.014	0.023
	(0.009)	(0.016)	(0.010)	(0.011)
Age	−0.006	−0.006	0.003	0.003
	(0.0003)	(0.0003)	(0.003)	(0.003)
Education 9–11	0.015	0.006		
	(0.011)	(0.011)		
Education 12	−0.092	−0.093		
	(0.011)	(0.011)		
Education >12	−0.178	−0.187		
	(0.012)	(0.012)		
SMSA	0.007	−0.041	−0.011	−0.007
	(0.007)	(0.008)	(0.011)	(0.011)
Baptist	0.064	0.047		
	(0.011)	(0.012)		
Protestant	0.026	0.001		
	(0.012)	(0.013)		
Other religion	−0.006	−0.059		
	(0.023)	(0.023)		
No religion	0.029	−0.004		
	(0.014)	(0.015)		
Unemployment rate	0.011	0.015	0.007	0.008
	(0.002)	(0.002)	(0.002)	(0.002)
Average wage	0.023	0.016	−0.006	−0.003
	(0.003)	(0.007)	(0.004)	(0.004)
Intercept	0.137	0.253		
	(0.032)	(0.063)		
Year dummies	Yes	Yes	Yes	Yes
State dummies	No	Yes	No	Yes
Indi. fixed effects	No	No	Yes	Yes
Elasticity of Pr(FEMHD) wrt				
Welfare benefit	**0.26**	**0.15**	**−0.03**	**0.09**
Mean of FEMHD	0.385	0.385	0.385	0.385
No. of observations	29 742	29 742	29 742	29 742

Notes: Dependent variable is one if woman is a female head. Sample is based on 1968–1989 PSID. Estimates are based on linear probability model. Standard errors are in parentheses.

percentage points. However, due to a much higher propensity for female headship, the estimated elasticity is about 0.26 compared to 0.56 among whites. Previous studies have found greater welfare effects for among lower education groups (see Moffitt, 1990a, 1994; Winkler, 1993) but have mixed evidence on racial differences in welfare effects (see Ellwood and Bane, 1985; Moffitt, 1990a, 1994; Hoffman et al., 1991). Contrary to expectations, living in a state that offers AFDC-UP benefits is estimated to have a positive effect on female headship.[31] The effect of other covariates is similar to that found for whites.

In sharp contrast to the estimates for whites, controlling for state effects, as shown in the model in Eq. (2), has little effect on welfare benefit estimates for black women. The coefficient on AFDC benefits is 0.011 compared to 0.019 without state effects. Fig. 4 shows the lack of correlation between the estimated state fixed effects and average welfare benefits by state. The correlation coefficient is an insignificant −0.07. It is possible that the state effects are capturing other characteristics of the state welfare program not being captured by the welfare benefit. However, this would imply similar values for the state effects in the black and white regressions; these values are not similar in my estimates. Among the 15 states common to the black and white state subsamples, the correlation between the state effects is weakly positive but insignificant.

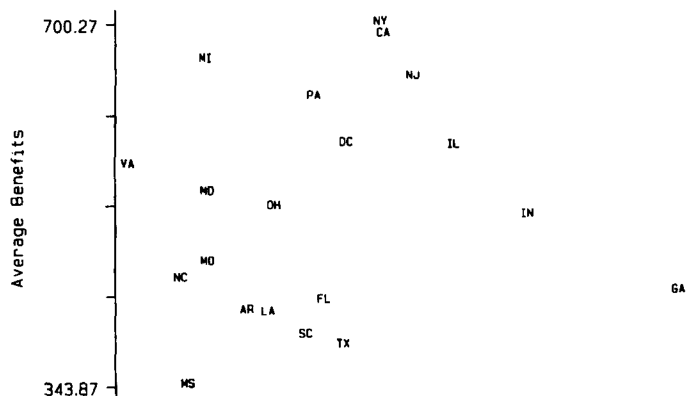

Fig. 4. Correlation between state fixed effects and average welfare benefits.

[31] In UP states, benefits are potentially available both inside and outside of marriage. Consequently, it is possible that the effect of benefits on family structure will differ in states offering AFDC-UP benefits versus non-UP states. In results not shown, interacting the AFDC-UP dummy with the benefit variable results in the expected negative coefficient on the AFDC-UP dummy. This was also found by Winkler (1993). The implications for the benefit variable are unchanged.

The remaining specifications in Table 3 show that unobserved individual effects are correlated with the state benefit level. Adding individual effects in Eq. (3) reduces the magnitude of the welfare effect from a positive and significant 0.011 to a statistically insignificant −0.002. By omitting a control for unobserved individual characteristics, we overestimate the role of welfare on the propensity to be a female head of household. Adding the state effects after the individual effects are added causes the estimated benefit effect to increase from −0.002 to 0.007. This result, while not statistically significant, implies that there are some omitted state factors which are negatively correlated with state benefit variables. It is not a robust finding, however, as sensitivity tests shown below find that in each case adding state effects once individual effects are included does not change the estimated benefit effect.

There are two important points to draw from the results for blacks. First, omitting individual effects can lead—erroneously—to the conclusion that AFDC benefits matter for female headship decisions. The composition of black women across the states (with respect to their propensity to choose female headship) is correlated with the state welfare benefit. Thus, the results from previous studies may be biased because of the omission of individual effects. Second, state effects for blacks are not correlated with welfare benefits. Thus, black women are influenced by social, cultural or religious norms at the state level (e.g. the state effects are jointly significant) but these norms do not affect the policy process leading to higher welfare benefits. This may be because of a lack of political power in this group.

One explanation for the importance of individual fixed effects is differential migration among blacks. About 9% of black families in the sample move at some time during the twenty-two years in the PSID sample. On average, these families move to higher benefit states. Those who moved found their benefits to be 1.5% higher than they would have been if they had stayed. Further, those with a higher propensity to be female head of household (a higher value for α_i) experience a larger increase in real benefits relative to those with a lower propensity to be a female head of household. About 16% of whites in the sample moved across state lines during the course of the panel. But the whites are not, on average, moving to higher benefit states. This difference in the migration patterns of female heads of household in the PSID is consistent with the broad patterns of migration among whites and blacks during this time. Up until the mid- to late 1970s, blacks (especially blacks in poverty) were moving out of the South (with low welfare benefits) and into the Northeast and Midwest (with higher benefits). Whites, on the other hand, were net migrants from the Northeast and Midwest to the South and West (Long, 1988).

On the surface, these results seem at odds with the literature that examines the importance of welfare benefits in migration decisions. While this literature has generated somewhat mixed evidence on the welfare magnet hypothesis (for example, see the reviews by Walker, 1994; Moffitt, 1992), recent work finds no evidence of welfare induced migration to high benefit states (e.g. Walker, 1994). There are several important reasons why the present results may not be inconsistent with prior findings. First, Walker's study does not include controls for race. Second, his results are based on (local) migration to three states considered welfare magnets, migration flows which may not be representative of migration patterns of blacks.[32] Third, and most importantly, my results imply a *correlation* between those who move and the benefit level, and this correlation explains the importance of controlling for individual fixed effects. Families may be moving to take advantage of stronger labor market conditions not benefits, yet the state's labor market opportunities may be correlated with the state's generosity of benefits. Thus, these results are not a direct test of welfare-induced migration. Regardless of the causal explanation, however, the main point is that omitting individual effects can lead to a bias through differences in the composition of states over time.

I also added the state level economic, demographic and political variables to the models in Eqs. (1) and (2). Adding these variables does not affect the estimated welfare effect (Hoynes, 1995); this is similar to the finding for whites.

5.3. Extensions of the main results: Full state sample

Tables 4 and 5 present estimates for several alternative specifications, one panel for each sensitivity test. Table 4 presents the results for whites and Table 5 presents the results for blacks. While they are included in the estimation, all parameters other than the benefit effect are dropped for brevity.

The first panel of Tables 4 and 5 present estimates where the sample is expanded to include the women residing in all states including the small states that were dropped from the main results. The sample sizes increase to 51 160 for whites and 30 685 for blacks. These results are very similar to the main results, especially after state effects are added. The one difference is that for whites, in the model in Eq. (1) without controls for state effects, the estimated welfare effect is 0.006 for an elasticity of 0.36 compared to 0.009 and an elasticity of 0.56 in the state subsample.[33]

[32] Walker considers moves to Wisconsin (from Illinois), Michigan (from Indiana and Ohio) and Virginia (from Tennessee). This omits the northeast and most of the south, important sources of movements for blacks (Long, 1988).

[33] As stated earlier, small samples within some states make it difficult to estimate both state and individual effects in the sample with the full set of states. The benefit effect, however, can be identified and is similar to those presented in the main specifications.

Table 4
Sensitivity tests for female headship model (whites)

	(1)	(2)	(3)	(4)
Sample includes all states				
AFDC and FO guarantee (100)	0.006	0.0003	0.001	
	(0.002)	(0.003)	(0.002)	
Elasticity of Pr(FEMHD) wrt welfare benefit	**0.36**	**0.02**	**0.06**	
Mean of FEMHD	0.095	0.095	0.095	
Number of observations	51 160	51 160	51 160	
Welfare benefit = AFDC guarantee				
AFDC guarantee (100)	0.006	−0.001	−0.0004	0.0005
	(0.001)	(0.002)	(0.002)	(0.002)
Elasticity of Pr(FEMHD) wrt welfare benefit	**0.30**	**−0.05**	**−0.02**	**0.03**
Mean of FEMHD	0.094	0.094	0.094	0.094
Number of observations	40 239	40 239	40 239	40 239
Women with kids				
AFDC and FO guarantee (100)	0.012	−0.0002	−0.001	−0.002
	(0.002)	(0.004)	(0.003)	(0.003)
Elasticity of Pr(FEMHD) wrt Welfare benefit	**0.51**	**−0.01**	**−0.04**	**−0.08**
Mean of FEMHD	0.138	0.138	0.138	0.138
Number of observations	27 681	27 681	27 681	27 681
Year dummies	Yes	Yes	Yes	Yes
State dummies	No	Yes	No	Yes
Indi. fixed effects	No	No	Yes	Yes

Notes: Dependent variable is one if woman is a female head. Sample is based on 1968–1989 PSID. Estimates are based on linear probability model. Standard errors are in parentheses.

5.4. Extensions of the main results: Using the AFDC guarantee

The welfare benefit variable used in the regressions for blacks and whites is defined to be the combined value of AFDC and Food Stamps. As noted earlier, Food Stamps are available to all eligible persons regardless of family structure. Therefore, one could argue that the AFDC guarantee is the correct variable to use. This is done in the second panel of Tables 4 and 5. For both whites and blacks, the estimated elasticities are about 40% lower when AFDC benefits are used. The general pattern for whites in terms of the importance of state and individual effects for this benefit specification is consistent with the earlier results. For blacks, comparing the models in Eqs. (1) and (2) shows the lack of correlation between unobserved state effects and AFDC benefits, and adding individual effects, in Eq. (3), generates a statistically insignificant welfare effect. Adding state fixed effects

Table 5
Sensitivity tests for female headship model (blacks)

	(1)	(2)	(3)	(4)
Sample includes all states				
AFDC and FO guarantee (100)	0.018	0.011	0.002	
	(0.003)	(0.006)	(0.004)	
Elasticity of Pr(FEMHD) wrt				
Welfare benefit	**0.25**	**0.15**	**0.03**	
Mean of FEMHD	0.380	0.380	0.380	
Number of observations	30 685	30 685	30 685	
Welfare benefit = AFDC guarantee				
AFDC guarantee (100)	0.015	0.010	0.004	0.004
	(0.003)	(0.005)	(0.004)	(0.004)
Elasticity of Pr(FEMHD) wrt				
Welfare benefit	**0.15**	**0.10**	**0.04**	**0.04**
Mean of FEMHD	0.385	0.385	0.385	0.385
Number of observations	29 742	29 742	29 742	29 742
Women with kids				
AFDC and FO guarantee (100)	0.030	0.028	−0.002	−0.006
	(0.004)	(0.007)	(0.004)	(0.004)
Elasticity of Pr(FEMHD) wrt				
Welfare benefit	**0.30**	**0.28**	**−0.02**	**−0.06**
Mean of FEMHD	0.510	0.510	0.510	0.510
Number of observations	23 788	23 788	23 788	23 788
Year dummies	Yes	Yes	Yes	Yes
State dummies	No	Yes	No	Yes
Indi. fixed effects	No	No	Yes	Yes

Notes: Dependent variable is one if woman is a female head. Sample is based on 1968–1989 PSID. Estimates are based on linear probability model. Standard errors are in parentheses.

to the model with individual effects, in Eq. (4), does not change the welfare effect. The female headship models were also estimated using the combined value of AFDC, Food Stamps and Medicaid, generating very similar results to those based on AFDC&FS.[34]

5.5. Extensions of the main results: Limiting sample to women with children

There are two main routes into female headship: a married woman with children becomes separated or divorced or an unmarried woman without children has a

[34] Various measures of the combined benefit of AFDC, food stamps and Medicaid with alternative valuations of Medicaid were examined. In general the estimated welfare effect is somewhat larger in those specifications then that shown here for AFDC and FS. See Hoynes (1995) for more detail on those results.

nonmarital birth. In order to capture both of these routes into female headship, the full sample includes all women regardless of marital status or presence of children. To examine the sensitivity to this sample construction, the last panel of Tables 4 and 5 presents estimates based on a sample of women aged 16–50 with children. Thus, these estimates focus on the marriage part of the female headship decision. These results show somewhat larger elasticities for blacks but fairly comparable elasticities for whites, compared to the results for the full sample. This taken with the earlier evidence suggests that welfare does not play a major role in either route in female headship: fertility or marriage decisions.[35]

6. Conclusion

This paper examines what impact unmeasured state and individual effects have on estimates of the effect of welfare benefits on female headship decisions. Using over twenty years of data from the Panel Study of Income Dynamics, I specify a model of female headship that not only includes controls for characteristics of the woman, state characteristics, year effects and welfare variables, but also controls for state of residence and individual effects.

For white women, the results show that welfare benefits are positively correlated with both individual and state effects. Models excluding both measures result in a positive and statistically significant welfare effect while adding individual or state fixed effects lead to a small negative (or zero) and statistically insignificant welfare effect. This gives a natural interpretation to the state effects as capturing the composition of the state residents. Among blacks, however, the results are quite different. There is virtually no correlation between state effects and welfare benefits and omitting state effects has no impact on the estimated welfare effect. Omitting the individual effects, however, does generate a substantial bias. Once we control for individual effects, the estimated welfare effect is small and statistically insignificant. One explanation for this is a higher propensity to move to higher benefit states among women with greater likelihoods of becoming female heads of household. These results do not provide direct evidence in support of the welfare-magnet hypothesis. Instead they show a correlation between migration and benefits among blacks that explains the importance of controlling for individual fixed effects. This does suggest, however, that a study examining the determinants of inter-state migration patterns of black women would be useful. More generally,

[35] The female headship model was also estimated separately for a sample of women with a high-school education or lower. Somewhat surprisingly, the results show that the sensitivity to welfare benefits in this low-education sample is no greater than that found in the full sample. Less than a high-school education might be a better definition of low education but because of small sample sizes in the PSID, this was not feasible. In general, the pattern of the results are very similar to those presented above. These results are available from the author.

these results illustrate the potential pitfalls of assuming that state factors are fixed over a long period of time.

The results imply that AFDC benefits do not play a role in female headship decisions. A few caveats apply, however. First, in the wake of the passage of welfare reform, many states are considering reforms which eliminate eligibility for AFDC for certain groups such as teen mothers. The estimates in this study imply that these policies are not likely to generate large effects. Using these results to evaluate reforms like this, however, requires out-of-sample predictions which should be used cautiously. Second, the study analyzes female headship as a stock concept and assumes that current benefits affect current headship status. However, modeling the flows into female headship (from marriage or childlessness) may imply that current headship status is a function of prior benefit levels. Furthermore, the social or cultural norms may themselves be functions of past welfare benefits. All of this implies that lagged benefits may matter. I explored this using 2–5 year lags in welfare benefits and the results were very close to those presented here. Nonetheless, an important extension of this research would involve directly modeling the evolution of these norms, which would build in a history of benefits. Finally, if welfare benefits are persistent over time then the state fixed effect may be absorbing part of the welfare effect. While the variance decomposition implies a significant amount of persistence in state benefits, Fig. 1 shows that benefits dropped over 30% over the period. If headship status is responsive to these financial incentives, it should show up in this period of significant declines in generosity.

Acknowledgments

I have benefitted from comments from Anne Case, Nada Eissa, Bruce Meyer, Robert Moffitt, Tom Mroz, Jim Poterba, Jim Powell, Paul Ruud, Glenn Sueyoshi, two anonymous referees and seminar participants at Berkeley, Michigan, Northwestern, San Diego, the Institute for Research on Poverty Workshop, and the NBER Summer Institute. Cindy Gustafson and Gary Painter provided excellent research assistance. Financial support was received from the National Institute for Child Health Development, and the Institute for Industrial Relations at Berkeley. Computing support was provided by the Econometrics Laboratory at UC Berkeley. A previous version of this paper was circulated under the title 'Female Headship and AFDC Benefits: State Effects or Welfare Effects?'

References

Bane, M., Ellwood, D., 1986. Slipping into and out of poverty: The dynamics of spells. Journal of Human Resources 21(4) 1–23.

Becker, G., 1973. A theory of marriage: Part I. Journal of Political Economy 81, 813–846.
Becker, G., 1974. A theory of marriage: Part II. Journal of Political Economy 82, 511–526.
Becker, G., 1981. A treatise on the family (Harvard University Press, Cambridge).
Besley, T., Case, A., 1994. Unnatural experiments? Estimating the incidence of endogenous policies, NBER Working Paper 4956, December 1994.
Bishop, J., 1980. Jobs, cash transfers and marital stability: A review and synthesis of the evidence. Journal of Human Resources 15, 301–334.
Chamberlain, G., 1980. Analysis of covariance with qualitative data. Review of Economic Studies 47, 225–238.
Danziger, S., Jakubson, G., Schwartz, S., Smolensky, E., 1982. Work and welfare as determinants of female poverty and household headship, Quarterly Journal of Economics August, 519–534.
Darity, W., Myers, S., 1983. Changes in black family structure: Implications for welfare dependency. American Economic Review 73(2) 59–64.
Decker, S., 1995. Medicaid, AFDC and female headship, Mimeo, New York University.
Duncan, G., 1984. Years of poverty, years of plenty (Institute for Social Research, The University of Michigan, Ann Arbor, MI).
Ellwood, D., Bane, M., 1985. The impact of AFDC on family structure and living arrangements, in: Ronald Ehrenberg, ed. Research in Labor Economics 7, 137–207.
Ellwood, D., Crane, J., 1990. Family change among black Americans: What do we know?. Journal of Economic Perspectives 4(4) 65–84.
Garfinkel, I., McLanahan, S., 1986. Single mothers and their children (The Urban Institute Press, Washington, DC).
Groeneveld, L., Hannan, M., Tuma, N., 1983. Final report of the Seattle/Denver income maintenance experiment, Vol. I, Design and results (US Government Printing Office, Washington, DC).
Hoffman, S., Duncan G., Mincy, R., 1991. Marriage and welfare use among young women: Do labor market, welfare and neighborhood factors account for declining rates of marriage among black and white women?, mimeo.
Hoynes, H., 1995. Does welfare play any role in female headship decisions?, National Bureau of Economic Research, Working Paper No. 5149.
Hoynes, H., 1996. Welfare transfers in two-parent families: The case of AFDC-UP. Econometrica 64(2) 295–332.
Hoynes, H., 1996b. Work, welfare and family structure: What have we learned?, in: A. Auerbach, ed., Fiscal policy: Lessons from economic research (MIT Press, Cambridge), in press.
Hoynes, H., MaCurdy, T., 1994. Has the decline in benefits shortened welfare spells?. American Economic Review 84(2) 43–48.
Huber, P., 1967. The behavior of maximum likelihood estimates under non-standard conditions, Proceedings of the Fifth Berkeley Symposium on Math Statistics and Probability 1, 221–233.
Hutchens, R., Jakubson, G., Schwartz, S., 1989. AFDC and the formation of subfamilies. Journal of Human Resources 24, 599–628.
Hsiao, C., 1986. Analysis of panel data (Cambridge University Press, Cambridge).
Long, L., 1988. Migration and residential mobility in the United States (Russell Sage Foundation, New York).
Maddala, G.S., 1983. Limited-dependent and qualitative variables in econometrics (Cambridge University Press, Cambridge).
McLanahan, S., 1988. Family structure and dependency: Early transitions to female household headship. Demography 25(1) 1–16.
Moffitt, R., 1989. Estimating the value of an in-kind transfer: The case of food stamps. Econometrica 57, 385–409.
Moffitt, R., 1990. The effect of the US welfare system on marital status. Journal of Public Economics 41, 101–124.
Moffitt, R., 1990. Has state redistribution policy grown more conservative?. National Tax Journal 43, 123–142.

Moffitt, R., 1992. Incentive effects of the US welfare system: A review. Journal of Economic Literature 30, 1–61.
Moffitt, R., 1994. Welfare effects on female headship with area effects. Journal of Human Resources 29(2) 621–636.
Moffitt, R., Ribar, D., Wilhelm, M., 1996. The decline of welfare benefits in the US: The role of wage inequality, mimeo.
Murray, C., 1984. Losing ground: American social policy, 1950–1980 (Basic Books, New York).
Poterba, J., 1994. State responses to fiscal crises: The effects of budgetary institutions and politics. Journal of Political Economy 102(4) 799–821.
Plotnick, R., Winters, R., 1985. A politico-economic theory of income redistribution. American Political Science Review 79, 458–473.
Schultz, T.P., 1994. Marital status and fertility in the United States: Welfare and labor market effects. Journal of Human Resources 29(2) 637–669.
US Bureau of the Census, 1961. Household and family characteristics: March 1960, P-20, No. 106.
US Bureau of the Census, 1991a. Household and family characteristics: March 1990, P-20, No. 458.
US Bureau of the Census, 1991b. Poverty in the United States: 1990, P-60, No. 175.
US House of Representatives, 1992. Background materials and data on programs within the jurisdiction of the Committee on Ways and Means (Government Printing Office, Washington, DC).
Walker, J., 1994. Migration among low-income households: Helping the witch doctors reach consensus, Institute for Research on Poverty Discussion Paper No. 1031-94.
Winkler, A., 1993. AFDC, cohabitation and the Family Support Act of 1988, mimeo (University of Missouri, St. Louis, MO).
Wilson, W., Neckerman, K., 1986. Poverty and family structure: The widening gap between evidence and public policy issues, in: S. Danziger and D. Weinberg, eds., Fighting poverty. What works and what doesn't (Harvard University Press, Cambridge) pp. 232–259.
Yelowitz, A., 1995. Will extending Medicaid to two-parent families encourage marriage? mimeo (UCLA).