Institute for Research on Poverty
Discussion Paper no. 1192-99

EITC, AFDC, and the Female Headship Decision

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June 1999

The authors thank Reagan Baughman for excellent research assistance. Bob Triest, Marianne Page, participants at the 1998 National Tax Association meetings, and participants at the 1998 Association for Public Policy Analysis and Management meetings provided valuable input.

Abstract

Concerns about the incentives for female headship for low-income families have focused on Aid to Families with Dependent Children (AFDC); however, the expansion of the Earned Income Tax Credit (EITC) has brought more low-income households into the tax system, subjecting them to additional marriage nonneutralities. Theoretical predictions about the correlations between the EITC and female headship are ambiguous. This paper is the first to provide empirical evidence that the EITC is correlated with female headship decisions. Using data from the Survey of Income and Program Participation, we find no significant correlations between AFDC and female headship. However, the ambiguous effect of the EITC on female headship is evident in our empirical analysis. After controlling for individual effects, we find that higher EITCs are associated with increased female headship for white women, but with decreased female headship for black women. For a sample of white women, we find that a \$100 increase in the EITC would increase the probability of female headship by 0.1 percent. For a sample of black women, we find that a \$100 increase in the EITC would decrease the probability of female headship by 1.4 percent, although this result is not robust.

EITC, AFDC, and the Female Headship Decision

I. INTRODUCTION

The effect of taxes and transfers on marriage and other family structure decisions has received a great deal of attention from policymakers and researchers. Recent tax legislation has sought to eliminate the marriage penalty in the federal tax system, and the 1996 welfare reform contained provisions to prevent teen mothers from forming their own households by requiring them to live with their parents. Concerns about the incentives for female headship for low-income families have focused on the Aid to Families with Dependent Children (AFDC) program, but the expansion of the Earned Income Tax Credit (EITC) has brought more low-income households into the tax system, subjecting them to additional marriage incentives and disincentives. More than four times as many families now receive the EITC than receive AFDC; 18.5 million tax units received the EITC with credits totaling almost $26 billion in 1996, compared with 4.6 million AFDC families receiving total benefits of over $20 billion (U.S. Congress, 1998). Although there has been some research on the effects of the federal income tax and transfer programs on marriage, the effects of the EITC have not been explicitly included in these analyses.

The EITC is a refundable tax credit targeted primarily at low-income families with children and intended to guarantee that parents who work are not poor (Scholz, 1994). The structure of the EITC gives rise to incentives for choosing one family structure over another. The credit is calculated as a percentage of *family* earnings and is phased out as earnings increase. This implies, for example, that a single mother with no earnings who marries a man with low earnings will become eligible for the EITC, thus providing a marriage subsidy. However, a family headed by a single mother who is eligible for the EITC is likely to become ineligible for the EITC if she marries because the couple's combined income may place them beyond the phase-out range. The potential for the EITC to affect marriage decisions is clearly a concern among policymakers.[1] However, this paper is the first to provide empirical evidence that the EITC is correlated with marriage decisions.[2]

This paper also contributes to existing literature by incorporating both the EITC and AFDC into our empirical analysis of female headship decisions.[3] Like the EITC, the welfare system targets low-income families with children, and this similarity in their target populations clearly links the two systems. Unlike the EITC, single-parent families are categorically eligible for higher benefits than two-parent families in the welfare system, thereby creating large incentives for female headship over marriage. The combined effect of the two systems is ambiguous because the EITC may either mitigate or exacerbate the large incentives for female headship in the transfer system.

Understanding the role that the EITC and welfare programs play in determining family structure decisions is relevant because of evidence that the economic well-being of female-headed families is lower than other types of family structures and that children growing up in female-headed households have poor outcomes relative to their counterparts from two-parent households (Haveman and Wolfe, 1994; McLanahan and Sandefur, 1994). The role of these programs has become especially important with the passage of the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) of 1996. PRWORA ended welfare, specifically AFDC, as an entitlement in the United States by placing strict time limits on welfare recipiency. If the former welfare system encouraged female headship, we expect that eliminating the existing welfare system might reduce rates of female headship. As the income tax takes an

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[1]A bill before the 105th Congress (H.R. 3995) would minimize the penalties for marriage in the EITC by changing the structure for joint filers relative to single or head-of-household filers.

[2]Dickert-Conlin (forthcoming) explores this possibility for 1 year of data. There is some evidence that the high marginal tax rates from the EITC discourage secondary earners, primarily married women, from working (Dickert, Houser, and Scholz, 1995; Eissa and Hoynes, 1998).

[3]Meyer and Rosenbaum (1998) and Houser and Dickert-Conlin (1998) consider the interaction with respect to labor supply.

increasingly important role in income transfers, its influence on behaviors such as marriage and cohabitation will become increasingly important in policy debates.

We exploit cross-sectional and cross-time variation in the EITC and AFDC programs. Using data from the Survey of Income and Program Participation (SIPP), we find no significant correlations between AFDC and female headship. However, the ambiguous effect of the EITC on female headship is evident in our empirical analysis. After controlling for individual effects, we find that increases in the EITC are associated with small increases in female headship for white women; however, higher EITCs are associated with decreases in female headship for black women. We find that a $100 increase in the EITC would increase the probability of female headship by 0.1 percent for a sample of white women and decrease the probability of female headship by 1.4 percent for a sample of black women. As a caveat, the result for our sample of black women is not robust.

Section 2 of this paper describes the incentives implicit in the EITC and AFDC programs for choosing female headship. In Section 3 we describe the earlier literature on the female headship and marriage decision with respect to the tax and transfer systems. Section 4 describes our data, Section 5 describes our estimation procedure, and Section 6 describes our results. We discuss alternative specifications in Section 7 and conclude in Section 8.

2. INSTITUTIONAL DETAILS

Our hypothesis that welfare benefits and the EITC may influence female headship decisions is based on a generalization of Becker's (1973, 1974) model of marriage, which suggests that the probability of female headship is a function of the expected gain from female headship and the distribution of unexpected outcomes from being a female head. Like Hoynes (1997), we define a female head as an unmarried woman with children. Out-of-wedlock birth and divorce are the two routes into female headship.

Consider the following simple characterization of the utility difference (F^*) between being a female head and not, which is based on Hoynes (1997) and Moffitt (1990, 1994):

$$F^* = U^{FH}(1;\ W^W,\ 0,\ B^{FH},\ E^{FH},\ X) - U^N(0;\ W^W,\ W^M,\ B^N,\ E^N,\ X) \qquad \textbf{(1)}$$

Female headship depends on a woman's wages, W^W; wages of the potential partner, W^M; welfare benefits, B; the EITC, E; and other observable and unobservable characteristics, X. AFDC benefits and the EITC differ depending on whether a woman is female head (*FH*) or not (*N*). The following descriptions of the EITC and the AFDC program emphasize how these programs vary with female headship and describe statutory changes in the tax and transfer system over time.

2.1 The Earned Income Tax Credit

EITC eligibility is based on the earnings of a tax-filing unit. Legal marital status determines who is included in the tax unit; married couples must file a joint return, and unmarried individuals file either head-of-household or single returns, depending on whether or not they have dependents. The largest EITCs are available to filing units with qualifying children, although a small credit is available to childless filers.[4] The credit increases with earnings until it reaches a maximum. Over a range of income, taxpayers receive the maximum credit, and then the credit is phased out with additional income above a certain amount. Unlike other credits, the EITC is refundable; that is, if a filing unit's credit is greater than its tax liability, the difference is paid to the filer by the Treasury. This makes it particularly relevant for low-income tax filers. Although eligible workers have the option of receiving the credit in advance with their earnings, almost all

[4]A qualified child is a natural or adopted child or stepchild of taxpayers filing joint or head-of-household returns.

recipients receive their EITC in a lump sum with their tax return.[5] The timing of the payment may imply that the EITC is more likely to be perceived as a transfer from the government than as a wage subsidy.

Although the amount of the EITC does not vary with filing status per se, the EITC may subsidize or penalize marriage relative to cohabitation or living independently, depending on the total family income and the distribution of income between the spouses. If a single mother with no income marries a man with low earnings, the new tax unit may become eligible for the EITC based on the man's earnings and the presence of children. For example, if a single mother of two with no earnings marries a man who earns $15,000 in 1997, the couple would receive an EITC of $3,014. Likewise, a two-earner married couple's joint income may exceed the EITC maximum, but if they were unmarried, one partner with income below the EITC maximum may file a head-of-household or single return and be eligible for the EITC. For example, a married couple who have two children and who each earn $15,000 would receive no EITC in 1997 because their combined income would make them ineligible. If the couple separates, the partner with the children would receive an EITC of $3,014.[6] Therefore, the effect of an increase in the EITC on the female headship decision is ambiguous. Because the EITC does not explicitly depend on female headship, an increase in the credit may simultaneously increase the attractiveness of female headship and the attractiveness of marriage. This is illustrated in Equation 1, in which E enters both U^{FH} and U^{N}.

The credit has changed a great deal in the past decade (see Table 1). Between 1986 and 1991, the federal EITC was indexed to inflation. The maximum benefit was $953 in 1990, conditional on having at least one child. With the passage of the 1990 Omnibus Budget Reconciliation Act, the maximum EITC value differed for families with one child versus families with two or more children. In addition, the

[5]According to the General Accounting Office (1992), only 0.5 percent of EITC recipients get the credit in advance.

[6]In one sense, the EITC is marriage neutral because a married couple with two children and $15,000 in total earnings receives the same credit as a single parent with two children and $15,000 in earnings.

TABLE 1
Federal EITC Parameters

		Credit Rate	Phase-In Range	Maximum Credit	Phase-Out Rate	Phase-Out Range
1989	1+ child	14.00%	$0–$6,500	$910	10.00%	$10,240–$19,340
1990	1+ child	14.00	0–6,810	953	10.00	$10,750–$20,264
1991	1 child	16.70	0–7,140	1,192	11.93	11,250–21,250
	2+ children	17.30	0–7,140	1,235	12.36	11,250–21,250
1992	1 child	17.60	0–7,520	1,324	12.57	11,840–22,370
	2+ children	18.40	0–7,520	1,384	13.14	11,840–22,370
1993	1 child	18.50	0–7,750	1,434	13.21	12,200–23,050
	2+ children	19.50	0–7,750	1,511	13.93	12,200–23,050
1994	No children	7.65	0–4,000	306	7.65	5,000–9,000
	1 child	26.30	0–7,750	2,038	15.98	11,000–23,755
	2+ children	30.00	0–8,425	2,528	17.68	11,000–25,296
1995	No children	7.65	0–4,100	314	7.65	5,130–9,230
	1 child	34.00	0–6,160	2,094	15.98	11,290–24,396
	2+ children	36.00	0–8,640	3,110	20.22	11,290–26,673
1996	No children	7.65	0–4,220	323	7.65	5,280–9,500
	1 child	34.00	0–6,330	2,152	15.98	11,610–25,078
	2+ children	40.00	0–8,890	3,556	21.06	11,610–28,495
1997	No children	7.65	0–4,340	332	7.65	5,430–9,770
	1 child	34.00	0–6,500	2,210	15.98	11,930–25,750
	2+ children	40.00	0–9,140	3,656	21.06	11,930–29,290

Source: U.S. Congress, 1998

maximum benefit levels increased—to $1,434 in 1993 from $1,192 in 1991 for a family with one child and to $1,511 in 1993 from $1,235 in 1991 for a family with at least two children.[7] Another EITC expansion followed the 1993 Omnibus Budget Reconciliation Act. A small benefit was made available to childless tax units, and there were large increases in the maximum credit. In 1997, the maximum benefit was $3,656 for a family with two or more children, $2,120 for a family with one child, and $332 for a family without children. The Joint Committee on Taxation estimates that over 18.7 million tax units will have received the EITC with credits totaling over $27.7 billion in 1998, an increase from 12.5 million families receiving credits totaling only $7.5 billion in 1990 (U.S. Congress, 1998).

Currently, ten states have earned income tax credits: Iowa, Kansas, Massachusetts, Maryland, Minnesota, New York, Oregon, Rhode Island, Vermont, and Wisconsin. These credits are all calculated as some percentage of the federal EITC. Therefore, the state EITCs provide the same incentives as the federal EITC for choosing one family structure over another. The parameters for the state EITCs are shown in Table 2. The state EITC is nonrefundable in three of these states, making it less well targeted toward low-income families.[8] In 1989, the first year of our data, only four states had EITCs. By 1995, the last year of our data, seven states had EITCs. The credit rates vary greatly across states over this period. For example, in 1993, Wisconsin's EITC was 25 percent of the federal EITC for families with two children and Minnesota's EITC was 15 percent of the federal EITC for families with children, regardless of the number.

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[7]The EITC in 1991 to 1993 was greater for filing units with children less than 1 year old. This so-called "wee-tots" credit increased the maximum credit by $388 in 1993.

[8]Nonrefundability implies that the tax unit must have positive tax liability to receive any EITC benefit and tax liability that is greater than the EITC to receive the entire benefit. We do not differentiate between these two cases in our empirical work.

TABLE 2
State EITC Parameters[a]

	Year Enacted	First Tax Year	Refundable	Rate as Percentage of Federal EITC	Without Qualifying Children[b]
IA	1989	1990	No	1990–1998: 6.5	Yes
KS	1998	1998	Yes	1998: 10	Yes
MA	1997	1997	Yes	1998: 10	Yes
MD	1987	1987	Yes[a]	1987–1997: 50 nonrefundable 1998: { 50 *refundable* 10 *nonrefundable*	No
MN	1991	1991	Yes	1991–1992: 10 1993–1997: 15 qualifying children 1998: { 15 no qualifying children 25 qualifying children	Yes
NY	1994	1994	Yes	1994: 7.5 1995: 10 1996: 20	Yes
OR	1997	1997	No	1997–1998: 5	Yes
RI	1975	1975	No	1987: 23.46 1988–1990: 22.96 1997–1997: 27.5 1998: 27	Yes
VT	1988	1988	Yes	1988: 23 1989: 25 1990–1993: 28 1994–1998: 25	Yes
WI[c]	1989	1989	Yes	1998–1993: { 5: 1 qualifying child 25: 2 qualifying children 75: 3 qualifying children 1994–1995: { 4: 1 qualifying child 16: 2 qualifying children 50: 3 qualifying children 1996–1998: { 4: 1 qualifying child 14: 2 qualifying children 43: 3 qualifying children	No

[a]Nick Johnson at the Center for Budget and Policy Priorities provided us with these data.
[b]This only applies to years following the 1993 tax law change in the federal EITC that made childless tax units eligible.
[c]Wisconsin had a nonrefundable EITC from 1983 to 1985. It was repealed in 1985 when the legislature eliminated income tax burdens on working poor. In 1994, the Wisconsin EITC used its own schedule, rather than a percentage of the federal EITC. However, the average credit was similar to the 1995 levels.

2.2 Aid to Families with Dependent Children[9]

AFDC-Basic provides cash benefits to low-income families in which the children are deprived of parental support because at least one parent is absent or incapacitated. Benefits increase with the number of people in the unit and decrease with income. If the natural or adoptive parents are married or cohabiting, the family is not eligible for basic AFDC, but may be categorically eligible for AFDC-Unemployed Parent (AFDC-UP) if the primary wage earner is unemployed. Welfare benefits may actually increase with a union because benefits increase with the number of people in the unit. Conceptually, an increase in welfare benefits may stabilize a union and simultaneously increase the attractiveness of independence. However, in practice, an increase in welfare benefits for single parents is not equivalent to an increase in benefits for married couples because the eligibility requirements are often more strict for AFDC-UP than for basic AFDC. For example, AFDC-UP requires the unemployed parent to show previous attachment to the labor force and to work fewer than 100 hours per month. In addition, states adopting AFDC-UP programs after the 1988 Family Support Act mandated that all states adopt UP programs may limit participation to 6 months per calendar year.

The above discussion applies if the partners are the natural parents of the children. Generally, cohabiting with an unrelated male typically does not decrease the size of the AFDC grant.[10] In most states, marriage to someone other than the natural parent lowers the AFDC grant based on only a portion of the

[9]PRWORA eliminated AFDC and replaced it with Temporary Aid for Needy Families. Our data are prior to 1996, so we refer to AFDC.

[10]In a 1993 survey of the 50 states and Washington, DC, Moffitt, Reville, and Winkler (1998) found that most state AFDC programs treat unrelated (nonparent) cohabitants quite leniently. Specifically, 35 states disregard any rent contributions made by the cohabitant. Another 13 states decrease the AFDC benefit by the amount of the shelter allowance if the cohabitant pays all the rent, but do not reduce the AFDC benefit at all if the cohabitant pays only a portion of the rent. The remaining three states reduce the AFDC benefit if the cohabitant pays the rent, but the reduction is lessened if the AFDC unit pays any part of the rent. All states count cash contributions made by the cohabitant to "meet the needs of the woman and her children." However, 27 states do not reduce the grant if these contributions are made for shared expenses, such as household supplies or food.

stepparent's income.[11] In summary, the highest AFDC benefits are available to a single parent who lives independently of the other parent. AFDC is not available to childless individuals, and benefits are often lower for cohabiting or married parents because the eligibility requirements are more strict for two-parent families.

The size of the penalties for cohabitation or marriage varies greatly by state because AFDC benefit levels differ across states. For example, the 1996 maximum benefit for a family of three in California was $607 but only $120 in Mississippi.

3. LITERATURE REVIEW

Research on how taxes affect marriage decisions has progressed independently of research on how transfer benefits affect family structure decisions. Research on the relationship between income taxes and marriage decisions suggests that the tax system has small but significant effects on marriage and divorce decisions. Recently, Alm and Whittington (1994, 1995) and Whittington and Alm (1997) have used microdata to examine the explicit effect of taxes on marital decisions. They find statistically significant evidence that the larger the tax penalty on marriage (an increase in joint tax liability relative to the sum of individual tax liabilities) the less likely an individual is to marry and the more likely a couple is to divorce. Although Alm and Whittington include the EITC in their calculations of the marriage penalty, they do not focus explicitly on low-income families.

Earlier literature on the negative income tax (NIT) may be more relevant to our analysis of the EITC because the NIT affected low-income families. Groeneveld, Tuma, and Hannan (1980) consider the effect of participation in an NIT experiment on marital dissolution. They proposed that the NIT may

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[11]Seven states count stepparents as natural parents, and the family's categorical eligibility for AFDC in these seven states would only be through the more restrictive UP program.

stabilize marriages by providing an income effect to married couples or weaken marriages because the NIT benefits were available to unmarried individuals as well. Their statistical analysis concludes that participants had higher marital dissolution rates than nonparticipants over the duration of the experiments. Cain and Wissoker (1990) discount these results by suggesting that the design of the NIT experiments and the data collected were inappropriate for making such conclusions. Cain and Wissoker's criticisms, which included the temporary nature of the NIT, are irrelevant for the EITC.

There is a much larger body of research on the effect of the transfer system on family structure decisions, including a thorough review by Moffitt (1992). Moffitt summarizes the findings from the empirical studies conducted in the 1980s, which consistently find a small but positive and significant effect of welfare on female headship. More recent articles control for additional economic variables, such as own earnings and earnings of potential spouses. Schultz (1994) finds that welfare benefits have a modest but statistically significant negative effect on the probability of living with a spouse, after controlling for the earnings of actual or potential spouses. Hoffman and Duncan (1995), controlling for husband's earnings, wife's wage rate, and child support benefits, find that AFDC benefits slightly increase marital dissolution rates.

Moffitt (1994) claims that a primary shortcoming in much of this literature is the use of cross-state variation in AFDC benefits as the exogenous variation in welfare benefits. This criticism arises from the possibility that unmeasured cross-state cultural differences may be captured by the welfare benefit variable. Failure to control for unobserved state characteristics that are correlated with both welfare and marriage decisions may bias the coefficients on the welfare covariates. For example, states that are generally accepting of single parenthood may also have generous welfare programs; therefore, the measured effect of the welfare benefit will be overstated.

Moffitt (1994) uses repeated cross sections from the Current Population Survey to generate large enough samples to include state fixed effects. He finds that the correlation between welfare benefits and female headship is negative but insignificant when controlling for state fixed effects. Hoynes (1997) also notes that unobservable individual preferences that are constant over time may be correlated with measures of the welfare benefits. For example, women who have a propensity to be on welfare may move to states with generous welfare benefits. When she includes state and individual effects in a panel analysis of the female headship decision using data from the Panel Study of Income Dynamics, she finds that welfare benefits are negatively and insignificantly related to the choice of female headship.

Dickert-Conlin and Houser (1998a, 1998b) and Dickert-Conlin (forthcoming) consider the interaction of the tax and transfer systems with regard to marriage disincentives. Dickert-Conlin and Houser (1998a, 1998b) describe the marriage penalties or subsidies created by the interactions between the tax and transfer systems. They find that tax penalties or subsidies for separation among low-income families, which are also likely to be affected by the transfer system, may be quite substantial relative to their total income. In addition, those facing the largest marriage penalties in the transfer system also face the largest marriage subsidies in the tax system. Dickert-Conlin (forthcoming) incorporates tax and transfer penalties or subsidies into a study of whether these incentives are related to the probability of divorce and finds that couples with more to gain from separating, in the form of lower tax liability, are more likely to separate.

A smaller literature considers the inframarginal decisions of whether welfare influences the decision to cohabit or legally marry.[12] Hu (1998) considers the possibility that the marriage penalties implicit in the AFDC program encourage cohabitation over marriage. Hu finds that women are very unlikely to marry a man who is not the father of their children, but there is no consistent effect of welfare benefit levels on the likelihood of marriage relative to cohabitation. He concludes that the fact that AFDC-

[12]Willis and Michael (1994) use data from the 1986 interview of the National Longitudinal Study of the High School Class of 1972 and conclude that economic variables influence the decision of whether to begin a partnership with cohabitation or legal marriage. They do not consider any welfare or tax variables in their analysis.

Basic and AFDC-UP benefit levels change together creates (or reinforces) an incentive for some women on AFDC-UP to leave a marriage and go onto AFDC-Basic, despite AFDC-UP's intended goal of keeping couples together. These results are consistent with Winkler (1995), who finds that generous welfare benefits for two-parent families (she defines two-parent families based on biological relationship to the child rather than marital status) are not positively correlated with the incidence of two-parent families. Schram and Wiseman (1988) find similar results.

Moffitt, Reville, and Winkler (1998) investigate whether cohabitation and marriage rates respond to AFDC benefit levels and to the rules regarding cohabitation. They find that states with generous polices toward in-kind contributions of cohabitants have higher rates of combining cohabitation with welfare. Unlike previous work, their study finds that the presence of a state AFDC-UP program is positively correlated with being on welfare and being married.

Our study controls for individual effects in a panel analysis, based on the work of Moffitt (1994) and Hoynes (1997), and extends the existing literature by analyzing the relationship between the female headship decision and the EITC.

4. DATA

We use a sample of women from the 1990, 1991, 1992, and 1993 panels of the SIPP. These overlapping panels cover the period from October 1989 to December 1995. The SIPP has relatively large sample sizes for a panel data set, although the panels are quite short. SIPP households are divided into four staggered rotation groups that are interviewed once every 4 months about their experiences during the past 4 months. A wave of the survey is completed when each of the rotation groups has been interviewed. The 1990 and 1991 panels each contain eight waves, the 1992 panel contains ten waves, and the 1993 panel contains nine waves.

SIPP does not uniquely identify nine states; therefore, we drop observations from those states.[13] To focus on women who are most likely to participate in the EITC and AFDC, we limit our sample to women between the ages of 18 and 50.[14] Excluding very young and elderly women also avoids the complicated family structures of teenage mothers and the different set of transfer programs for the elderly. Our primary sample has 33,267 white women and 4,411 black women. We take the observations from December of each year so we have a maximum of three observations per person.[15] We choose December because tax filing status is based on the marital status during this month. This gives us a total of 89,173 person-year observations for white women and 11,553 person-year observations for black women.

The means for selected variables are in Table 3. Generally, female heads face higher welfare guarantees than nonheads. Female heads face lower EITCs than women who are not female heads, although the difference is only statistically significant for the sample of white women. Female heads are generally younger with substantially lower incomes and higher welfare participation than nonheads.

5. ESTIMATION ISSUES

We linearize the underlying difference in utility between whether or not to be a female head from Equation 1 as

$$F_{it}^{*} = \alpha + X_{it}{}'\gamma_3 + \gamma_1 b_{it} + \gamma_2 e_{it} + v_{it}. \quad \textbf{(2)}$$

[13]SIPP aggregates Alaska, Idaho, Iowa, Maine, Montana, North Dakota, South Dakota, Vermont, and Wyoming into three groups.

[14]We also show the results of estimations using a sample that excludes women with fewer than 12 years of education.

[15]In future work, we hope to take advantage of the monthly observations on individuals. Data that observe individuals once in a year may not capture changes that occur in a relatively shorter time frame.

TABLE 3
Descriptive Statistics of Pooled Data Means and Standard Deviations

	White Not Female Head	White Female Head	Black Not Female Head	Black Female Head
Monthly maximum AFDC + Food Stamps for family of three (1996$)	726 (139)	742 (139)	675 (142)	680 (143)
Annual maximum EITC for family with two children (1996$)	1608 (520)	1598 (514)	1571 (500)	1564 (493)
Percentage eligible for the EITC, based on predicted wages	47.2% (49.9)	69.4% (46.1)	56.0% (49.6)	79.4% (40.4)
AFDC-UP program (1=yes)	14.8% (35.5)	16.2% (36.8)	14.1% (34.8)	15.8% (36.5)
Age	36.7 (8.2)	35.3 (7.3)	38.1 (8.1)	33.9 (7.2)
Education (years)	13.3 (2.8)	12.2 (2.8)	12.9 (2.5)	12.1 (2.2)
Monthly household earnings (1996$)	4137 (2907)	1580 (1696)	3041 (2339)	1073 (1392)
Monthly household income (1996$)	4475 (3044)	2100 (1779)	3293 (2325)	1437 (1338)
Percentage in urban	76.9% (42.2)	77.0% (42.1)	83.8% (36.9)	81.4% (38.9)
Percentage with children	57.7% (49.4)	100.0% (0.0)	50.5% (50.0)	100.0% (0.0)
Number of children	1.1 (1.2)	1.8 (1.0)	1.0 (1.3)	2.0 (1.1)
Percentage receiving AFDC	1.6% (12.4)	24.2% (42.8)	6.4% (24.5)	41.0% (49.2)
Percentage receiving Food Stamps	4.0% (19.7)	32.1% (46.7)	12.3% (32.9)	53.7% (49.9)

(table continues)

TABLE 3, continued

	White Not Female Head	White Female Head	Black Not Female Head	Black Female Head
State real per capita income (1996$)	21255 (3021)	21412 (2961)	20984 (3372)	20919 (3432)
State unemployment rate (1996$)	6.6 (1.9)	6.6 (1.8)	6.5 (1.3)	6.6 (1.2)
State manufacturing wage (1996$)	12.75 (1.55)	12.83 (1.47)	12.32 (1.73)	12.41 (1.91)
Person years	78,660	8,890	7,358	3,938
Individuals	33,267		4,411	

Source: Sample of women between 18 and 50 years old from the 1990–1993 panels of SIPP. We take observations from December of each year and restrict the sample to women who are heads of household or spouses of heads.

F_{it}^{*} is empirically unobservable, but we do observe[16]

$$F_{it} = \begin{cases} 1(\textit{female head}) \ \textit{if} \ F_{it}^{*} > 0 \\ 0(\textit{not female head}) \ \textit{if} \ F_{it}^{*} \leq 0 \end{cases} \tag{3}$$

Our covariates, *X*, include age, a dummy for whether the woman lives in an urban area, and year dummies. We also include state characteristics that might be correlated with female headship decisions, such as the average real wage in manufacturing, real average per capita income, unemployment rate, and whether the state had an AFDC-UP program in 1989 and 1990 (U.S. Department of Commerce).[17] Because earlier work has found significant differences between black and white women (Hoynes, 1997; Moffitt, 1994; Schultz, 1994), we estimate separate regressions for black and white women.

5.1 Endogeneity

Two issues deserve discussion in our estimation of Equation 2. The first is our choice of the independent variables of interest, b_{it}, a measure of the welfare guarantee, and e_{it}, a measure of the generosity of the EITC. The actual welfare benefit or EITC for which the family is eligible is likely to be endogenous with the female headship decision. In particular, the actual AFDC benefit and EITC depend on the number of children, but fertility decisions are endogenous to the female headship decision. In addition, the AFDC benefits and EITC depend on the relative incomes of husband and wife, and income is clearly endogenous. High tax rates on earnings for welfare recipients and on secondary earners in the progressive tax system may influence women's labor supply. Therefore, we use the maximum monthly combined

[16]We initially treat women who we believe are cohabiting as unmarried. In specification tests, we experiment with defining these women as married. The tax and transfer system do not treat cohabitation symmetrically—the tax unit is based on legal marital status, while the AFDC unit is typically based on kin relationships.

[17]The 1998 Family Support Act mandated that all states adopt an AFDC-UP program by October 1990. In 1990, we code all states that did not have an UP program until October as not having an UP program in that year.

AFDC and Food Stamp benefit for a family of three and the maximum combined annual federal and state EITC for a family with two children (U.S. Congress, various editions).[18] These vary by state and time, and we adjust them to reflect 1996 dollars. We lag the EITC value by one year because it is most often paid as a tax refund in the calendar year following the tax year in which the credit is earned.

5.2 Fixed Effects

The second issue that deserves discussion in our estimation of Equation 2 is based on the recent work by Moffitt (1994) and Hoynes (1997) showing the importance of controlling for individual and/or state fixed effects. Their argument for controlling for state or individual effects in a study of welfare and female headship also holds for the EITC. For example, women with an unobserved propensity to work may move to states that give higher earned income credits. Because the data have short panels, we cannot identify separate individual and state effects (less than 4 percent of the sample moves between states), and, therefore, we only include individual effects. If the composition of the state is constant over our 7-year period, then the state and individual effects should be identical (Hoynes, 1997).[19]

A fixed-effects model assumes that the individual effect is a fixed variable and does not require strict assumptions about the exogeneity of the individual effects and the other regressors. The fixed effects model estimates Equation 2 as

$$F_{it}^{*} = \alpha_1 + \gamma_1 b_{it} + \gamma_2 e_{it} + X_{it}'\gamma_3 + v_{it} \qquad \textbf{(4)}$$

[18]Nick Johnson at the Center for Budget and Policy Priorities provided us with the state EITC parameters.

[19]Hoynes (1997) also notes that controlling for both individual and state effects does not change the coefficient on the welfare variable relative to controlling for only one or the other.

where α_i is the time-invariant individual effect.[20] Despite our discrete dependent variable, we follow Hoynes (1997) and estimate a linear probability model because of the computational difficulties in estimating a discrete choice fixed-effects model.[21]

6. RESULTS

6.1 Sample of White Women

Our results for the sample of white women are shown in Table 4. In column 1 we show that if the data are pooled, the coefficient on the welfare guarantee is positive and statistically significant. This positive correlation between welfare and female headship is consistent with earlier literature that does not control for individual or state effects. On the other hand, the coefficient on the EITC is negative and statistically significant, suggesting that the EITC has more of a marriage-stabilizing effect than an independence effect.

As shown in column 2, when we treat individual effects as fixed, the coefficients on the individual fixed effects are jointly significant, and the coefficient on the welfare variable is negative and statistically insignificant. This result is consistent with the recent findings of Hoynes (1997) and Moffitt (1994) and suggests that female headship decisions are uncorrelated with the generosity of welfare benefits after controlling for unobserved fixed effects. The coefficient on the EITC is negative and statistically insignificant, indicating that the EITC is also uncorrelated with female headship decisions.

[20]An alternative would be to model the individual effects as random. We present the results of specifications using random effects in Section 7.

[21]See Greene (1997) for a discussion of the limitations of a linear probability model for a binary dependent variable. The coefficients in the pooled linear probability model are similar to the marginal effects in a pooled probit regression.

TABLE 4
Primary Specification for White Women

	Pooled	Fixed Effects	Fixed Effects
Maximum AFDC + Food Stamps (1000)	0.093**	-0.022	-0.016
	(0.011)	(0.028)	(0.028)
Maximum EITC (1000)	-0.022**	-0.001	-0.012
	(0.009)	(0.019)	(0.019)
Max EITC * Eligibility	—	—	0.014**
			(0.002)
AFDC-UP program (1=yes)	0.001	0.000	0.000
	(0.005)	(0.003)	(0.003)
Age (years)	-0.002**	0.000	0.000
	(0.000)	(0.005)	(0.005)
High school (1=yes)	-0.073**	-0.014**	-0.010*
	(0.004)	(0.006)	(0.006)
More than high school (1=yes)	-0.108**	-0.013*	-0.004
	(0.004)	(0.007)	(0.007)
Urban (1=yes)	0.003	0.006	0.010*
	(0.003)	(0.006)	(0.006)
State real per capita income	-0.001**	0.000	0.000
	(0.0005)	(0.001)	(0.001)
State unemployment rate	0.002**	0.000	0.000
	(0.001)	(0.000)	(0.000)
State manufacturing wage	0.000	0.000	0.000
	(0.001)	(0.001)	(0.001)
1990 (1=yes)	-0.005	-0.002	-0.001
	(0.005)	(0.005)	(0.005)
1991 (1=yes)	-0.005	0.001	0.001
	(0.005)	(0.010)	(0.010)

(table continues)

TABLE 4, continued

	Pooled	Fixed Effects	Fixed Effects
1992 (1=yes)	-0.001	0.002	0.002
	(0.006)	(0.009)	(0.009)
1993 (1=yes)	0.004	0.002	0.001
	(0.007)	(0.012)	(0.012)
1994 (1=yes)	0.011	0.001	0.002
	(0.008)	(0.015)	(0.015)
1995 (1=yes)	0.040**	—	—
	(0.017)		
Constant	0.219**	0.149	0.129
	(0.015)	(0.149)	(0.149)

Source: Sample of women between 18 and 50 years old in December of each year from the 1990–1993 panels of SIPP. 33,267 individuals; 89,173 observations.

*Statistically significant at the 10 percent level.
**Statistically significant at the 5 percent level.

Because the EITC is targeted to families with low earnings, it may have a significant effect on the female headship decisions of only those women who are likely to be eligible for the credit. As discussed above, eligibility for the EITC is likely to be endogenous with both the female headship decision and income. As an instrument for eligibility, we use whether a woman would be eligible for the EITC assuming that she worked 2,000 hours at her predicted wage.[22] This instrument proxies for whether the woman would be eligible if she were single and, assuming assortative mating, whether her family would be low-income if she were married. We then interact predicted eligibility with the maximum EITC as a measure of the additional effect that the credit has on those women who are likely to be eligible for the EITC.

The results of the estimation including the EITC-eligibility interaction term are shown in the third column of Table 4. Again, the estimated coefficients on both the welfare and the EITC variables are negative and not statistically significant. However, the coefficient on the maximum EITC interacted with our instrument for EITC eligibility is positive and statistically significant, implying that the independence effect of the EITC dominates the marriage-stability effect for those women who are likely to be eligible. The coefficients on the EITC and the interaction term are jointly significant, and the combined marginal effect is positive but small. The magnitude of the coefficient suggests that a $100 increase in the EITC would increase the probability of an EITC-eligible woman of being a female head by 0.02 of a percentage point, or 0.1 percent.

One possible reason for the difference between the coefficients on the EITC and the welfare guarantee variables is that eligible families are more likely to participate in the EITC than they are to participate in transfer programs. Our measures of the penalties and subsidies for female headship reflect 100 percent participation, while the actual participation in transfer programs is between 25 and 75 percent

[22]Predictions are based on log wage regressions, dropping the self-employed and adjusting for selection. Covariates include education, quadratics in age and experience, and dummies for African American, residence in a metropolitan area, and region. We use the state unemployment rate for women, number of children, and a dummy for the presence of children under 6 as exclusion restrictions in the probit regression for being in the labor force.

(see Dickert, Houser, and Scholz, 1995).[23] The EITC participation rate is also not 100 percent, but there is evidence that the participation rate is higher than among transfer programs. Scholz (1994), for example, finds that among those eligible for the EITC in 1990, when the EITC was much smaller than its current level, the participation rate was approximately 85 percent. Stigma, transaction costs, or information costs are considered to be higher for transfer programs than for the EITC and may explain the difference in participation.

In each of the specifications, higher levels of educational attainment are negatively correlated with female headship at statistically significant levels. Living in an urban area is also positively correlated with female headship in our third specification.

Another way to isolate the effects of the EITC on the population of women who are most likely to participate is to restrict the sample to women with low levels of education. This is the approach taken by Moffitt (1990) in his analysis of AFDC, in which he notes that it may be inefficient to use all women because so few participate in AFDC. As an additional measure of the EITC on female headship, we repeat the previous specifications of the fixed-effects model using a sample of white women with less than high school education.

The results using the low-education sample are shown in Table 5. Column 1 shows that the coefficient on the welfare variable is positive and not statistically significant, consistent with Hoynes (1997) and Moffitt (1994). The coefficient on the EITC is positive and statistically significant for the low-education sample, indicating that the independence effect of the EITC dominates for this group of women. The magnitude of the coefficient suggests that a $100 increase in the EITC would increase the probability of being a female head by 2.2 percentage points, or 12 percent. Given our particularly restrictive sample

[23]According to Dickert, Houser, and Scholz (1995), the participation rate (number of participants divided by the number of eligibles) for AFDC is 76 percent for single parents and 25 percent for married couples. The Food Stamp participation rate is 67 percent for singles and 32 percent for married couples.

TABLE 5
White Women with Less than High School Education

	Fixed Effects	Fixed Effects
Maximum AFDC + Food Stamps (1000)	0.100 (0.113)	0.107 (0.113)
Maximum EITC (1000)	0.222** (0.098)	0.206** (0.099)
Max EITC * Eligibility	—	0.014 (0.012)
AFDC-UP program (1=yes)	0.010 (0.010)	0.010 (0.010)
Age (years)	-0.058** (0.025)	-0.057** (0.025)
Urban (1=yes)	-0.028 (0.024)	-0.028 (0.024)
State real per capita income	0.001 (0.005)	0.001 (0.005)
State unemployment rate	0.000 (0.002)	0.000 (0.002)
State manufacturing wage	0.000 (0.003)	0.000 (0.003)
1990 (1=yes)	0.066** (0.026)	0.066** (0.026)
1991 (1=yes)	0.128** (0.051)	0.128** (0.051)
1992 (1=yes)	0.117** (0.048)	0.116** (0.048)
1993 (1=yes)	0.153** (0.062)	0.152** (0.062)

(table continues)

TABLE 5, continued

	Fixed Effects	Fixed Effects
1994 (1=yes)	0.188**	0.187**
	(0.077)	(0.077)
1995 (1=yes)	—	—
Constant	1.819**	1.802**
	(0.762)	(0.762)

Source: Sample of women between 18 and 50 years old from the from 1990–1993 panels of SIPP. We take observations from December of each year and restrict the sample to women with fewer than 12 years of education. 4,075 individuals; 10,828 observations.

*Statistically significant at the 10 percent level.
**Statistically significant at the 5 percent level.

selection for this regression, we are not surprised that the magnitude from this regression is so different from our primary specification.

The coefficient on the interaction of the EITC with our instrument for eligibility is positive but not statistically significant (these results are shown in the second column of Table 5). This may be due to the fact that the selection of the sample on educational level dilutes the power of this instrument for EITC eligibility. However, the estimated coefficients on the maximum EITC and the interaction term are jointly significant. The coefficients indicate that a $100 increase in the EITC would increase the probability of being a female head among EITC-eligible women in this sample by 11.5 percent.

6.2. Sample of Black Women

As in the estimations for white women, the coefficient on the welfare guarantee in the pooled specification for black women is positive and statistically significant (Table 6, column 1). The sign on the coefficient of the EITC variable is negative and statistically significant. This is the opposite of what we found in the sample of white women and suggests that the incentives for stable marriages implicit in the EITC outweigh the incentive for female headship. When we treat individual effects as fixed (column 2), the coefficients of both the welfare guarantee and the EITC have the same sign but are not statistically different from zero. This result is also different from what we found in the sample of white women. The fixed-effects model indicates that the EITC is not correlated with female headship in our sample of black women. Unlike the sample of white women, education is positively correlated with female headship for black women, and the probability of being a female head is lower in later years.

We also estimate the fixed-effects model for black women including the interaction between the maximum EITC and our instrument for EITC eligibility (Table 6, column 3). The coefficient on the EITC is still negative and statistically insignificant. However, the coefficient on the interaction term is positive and significant at the 10 percent level. The two coefficients on the maximum EITC and the interaction term

TABLE 6
Primary Specification for Black Women

	Pooled	Fixed Effects	Fixed Effects
Maximum AFDC + Food Stamps (1000)	0.088*	0.114	0.123
	(0.052)	(0.107)	(0.108)
Maximum EITC (1000)	-0.101**	-0.068	-0.075
	(0.033)	(0.064)	(0.064)
Max EITC * Eligibility	—	—	0.013*
			(0.007)
AFDC-UP program (1=yes)	0.062**	-0.003	-0.003
	(0.019)	(0.011)	(0.011)
Age (years)	-0.015**	0.018	0.018
	(0.000)	(0.017)	(0.017)
High school (1=yes)	-0.130**	0.042**	0.043**
	(0.012)	(0.021)	(0.021)
More than high school (1=yes)	-0.222**	0.066**	0.071**
	(0.012)	(0.025)	(0.025)
Urban (1=yes)	-0.031**	0.002	0.003
	(0.013)	(0.034)	(0.034)
State real per capita income	-0.002	-0.003	-0.003
	(0.002)	(0.004)	(0.004)
State unemployment rate	0.010**	0.002	0.002
	(0.004)	(0.004)	(0.004)
State manufacturing wage	0.009**	0.000	0.000
	(0.003)	(0.002)	(0.002)
1990 (1=yes)	0.002	-0.028	-0.026
	(0.018)	(0.018)	(0.018)
1991 (1=yes)	0.025	-0.061*	-0.058*
	(0.020)	(0.034)	(0.034)

(table continues)

TABLE 6, continued

	Pooled	Fixed Effects	Fixed Effects
1992 (1=yes)	0.055**	-0.068**	-0.066**
	(0.023)	(0.033)	(0.033)
1993 (1=yes)	0.089**	-0.084*	-0.082*
	(0.026)	(0.044)	(0.044)
1994 (1=yes)	0.123**	-0.088*	-0.084
	(0.029)	(0.053)	(0.054)
1995 (1=yes)	0.268**	—	—
	(0.063)		
Constant	0.946**	-0.255	-0.244
	(0.056)	(0.513)	(0.513)

Source: Sample of women between 18 and 50 years old in December of each year from the 1990–1993 panels of SIPP. 4,411 individuals; 11,553 observations.

*Statistically significant at the 10 percent level.
**Statistically significant at the 5 percent level.

are not jointly significant. The overall marginal effect of the EITC is negative and suggests that a $100 increase in the EITC would lower the probability of being a female head by 1.4 percent. Given the insignificance of the coefficient on the EITC variable, we do not have confidence in this interpretation. If we reinterpret the results, focusing only on the interaction term, we predict that a $100 increase in the EITC would raise the probability of being a female head by 0.1 of a percentage point, or 0.3 percent.

One explanation for why the EITC and female headship are inconsistent between our samples of black and white women relates to our general model of female headship. Recall that a priori we do not have a prediction on the sign of the correlation between the EITC and female headship because an increase in the EITC raises utility inside and outside of marriage. It may be the case that the labor market opportunities for single black mothers are insufficient for the independence effect to dominate. It may also be the case that the spouses of black women eligible for the EITC have sufficiently low earnings (again, this may be a result of poor labor market opportunities) that even if the women have high preferences for work, their combined income as a couple does not make them ineligible for the EITC. These qualities may not hold true for the sample of white women and, therefore, the independence effect dominates.

We also reestimate the fixed-effects models for black women and restrict the sample to those with less than high school education (Table 7). The coefficient on the welfare guarantee and the maximum EITC are both positive and statistically insignificant, indicating that the EITC is not correlated with female headship for the low-education sample. The coefficient on the interaction between the maximum EITC and our instrument for EITC eligibility is also negative and insignificant when included in the estimation.

TABLE 7
Black Women with Less than High School Education

	Fixed Effects	Fixed Effects
Maximum AFDC + Food Stamps (1000)	0.254	0.253
	(0.343)	(0.343)
Maximum EITC (1000)	0.015	0.020
	(0.217)	(0.224)
Max EITC * Eligibility	—	-0.005
		(0.049)
AFDC-UP program (1=yes)	-0.011	-0.011
	(0.024)	(0.024)
Age (years)	0.001	0.001
	(0.057)	(0.057)
Urban (1=yes)	0.287**	0.287**
	(0.123)	(0.123)
State real per capita income	-0.012	-0.012
	(0.011)	(0.011)
State unemployment rate	0.003	0.003
	(0.009)	(0.009)
State manufacturing wage	0.004	0.004
	(0.005)	(0.005)
1990 (1=yes)	-0.046	-0.046
	(0.058)	(0.058)
1991 (1=yes)	-0.065	-0.065
	(0.113)	(0.113)
1992 (1=yes)	-0.077	-0.077
	(0.107)	(0.107)
1993 (1=yes)	-0.130	-0.130
	(0.143)	(0.143)

(table continues)

TABLE 7, continued

	Fixed Effects	Fixed Effects
1994 (1=yes)	-0.117 (0.176)	-0.117 (0.177)
1995 (1=yes)	—	—
Constant	0.293 (1.730)	0.296 (1.731)

Source: Sample of women between 18 and 50 years old from the from 1990–1993 panels of SIPP. We take observations from December of each year and restrict the sample to women with fewer than 12 years of education. 879 individuals; 2,242 observations.

*Statistically significant at the 10 percent level.
**Statistically significant at the 5 percent level.

7. SPECIFICATION TESTS

7.1 Cohabitation

In our initial regressions we treat all unmarried women as female heads if they have children and as nonheads if they do not have children. However, the prevalence of cohabitation and the treatment of cohabitation by the tax and transfer systems leave this as a potentially important misspecification. Bumpass and Sweet (1995) find that about half the population under age 40 has lived with an unmarried partner at some point, with even higher rates for individuals with low education levels. Bumpass, Raley, and Sweet (1995) find that one-half of all cohabiting couples have children, which makes the tax and transfer system's treatment of children particularly relevant.

The welfare system and the EITC may provide incentives for cohabitation over marriage or living independently. With respect to the tax system, women with a propensity to work might be better off cohabiting rather than being married because a second income in a family would put the family in the phase-out range or make them ineligible for the EITC. Moffitt, Reville, and Winkler (1998) show that the AFDC program is much more lenient toward cohabitation than toward marriage with respect to grant generosity (see footnote 10).

We are therefore interested in determining how our estimates change if we define a female head as a woman who is unmarried *and* does not have a cohabiting partner. Unfortunately, SIPP does not explicitly identify cohabiting partners. Based on work by Dickert-Conlin, Houser, and Baughman (1998), we define cohabiting couples as those who are unmarried, of the opposite sex, and with a maximum age difference between partners of 13 years.[24] We recode a total of 2,016 observations (717 individuals) in which women

[24]Thirteen years is the mean age difference plus two standard deviations among explicitly identified cohabitors in the decennial census.

who were female heads in our first definition are not considered female heads because of their cohabiting status.

In the sample of black and white women, the coefficients in the fixed-effects model are basically unchanged. Our results are not sensitive to the definition of female headship.

7.2 Women with Children

The main specification does not distinguish between the two routes into female headship—women who are not married can become female heads by having a child, and married women with children can become female heads by divorcing or separating. We can isolate the marriage part of the female headship decision by dropping women without children from the sample.[25] For white women, the results are similar to the main specification shown in Table 4. For black women, the size of the coefficient on the interaction between the EITC and eligibility increases relative to the estimate for the full sample, and the coefficient is statistically significant. The coefficients on the EITC and the interaction term are jointly significant. The marginal effect of the EITC on female headship is still negative for the sample of black women with children, but the magnitude is smaller than that for all black women.

7.3 Random Effects

The choice between a random-effects and fixed-effects estimation of panel data is not straightforward. Hsaio (1986) suggests that treating individual effects as random variables is more appropriate because we are interested in making inferences about the entire population, not just our sample. Treating individual effects as random effects implies that the error term in Equation 2 can be represented as

$$v_{it} = \gamma_i + \mu_{it} \ . \qquad \textbf{(5)}$$

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[25]Hoynes (1997) uses this approach as an alternate specification in her analysis of AFDC and female headship.

The component γ_i represents the individual effect that is constant over time but randomly distributed. In this specification μ_{it} is the random disturbance particular to both the individual and time. Although conceptually preferable, estimating a random-effects model requires the strict assumption that the unobserved individual effects are uncorrelated with the other covariates.

Hausman (1978) provides a specification test for the random-effects model's assumption that the individual effect is exogenous to all of the regressors for an ordinary least squares regression. We use this specification test to explore the applicability of the fixed-effects and random-effects models. In most of our specifications, the Hausman test rejects the hypothesis that the individual effects are uncorrelated with the other regressors, indicating that the random-effects model may not be appropriate.

However, when we use the sample of black women with low education, the Hausman test supports the necessary assumption of the random-effects model. Using the random-effects model, the coefficient on the welfare guarantee is positive and statistically insignificant, and the estimated coefficient on the maximum EITC is negative and statistically significant. This change in significance for the EITC variable may due to the ability of the random-effects model to capture variation within an individual as well as between individuals, unlike the fixed-effects model, which captures only variation within an individual. The magnitude of the coefficient suggests that a \$100 increase in the EITC would decrease the probability of being a female head by 2.8 percentage points, or 5 percent.

The random effects specification of the model that includes the interaction between the maximum EITC and our instrument for eligibility is also not rejected by the Hausman test. The coefficient on the EITC is again negative while the coefficient on the interaction term is positive; the coefficients are both individually and jointly significant. The combined effect of the two variables is very similar (a \$100 increase in the EITC would decrease the probability of being a female head by 6 percent) to the effect of the EITC in the previous specification.

8. CONCLUSIONS

The primary goal of this paper is to consider whether the EITC is correlated with female headship decisions, while controlling for the other major income transfer to low-income families, AFDC. The vast expansion in the EITC during the 1990s has made this question a focus of policy debates and provides us with variation over time for the empirical study. Based on the statutory structure of the EITC, the effect of the EITC on female headship is ambiguous. The EITC is available to families with single heads and with married heads, and the credit may increase or decrease with marriage or separation. This ambiguity is not completely resolved in the empirical analysis. We find that an increase in the EITC is positively correlated with female headship for a sample of white women. Our best estimate suggests that a $100 increase in the EITC would increase the probability of an EITC-eligible woman of being a female head by 0.1 percent. However, the EITC is not robustly correlated with female headship for a sample of black women.

There are several possible reasons for the difference in the correlation of the EITC and female headship across race. One likely explanation is that the effect of the EITC varies with the route into (or out of) female headship. We have begun to explore this possibility in some of the alternate specifications. These results suggest that the correlation of the EITC with marriage decisions may differ from the correlation of the EITC with female headship. Further work needs to be done to estimate the effect of the EITC on divorce or separation for married women with children and marriage for unmarried women with children. Because nonmarital childbearing is correlated with race, the different results for black and white women may be driven by different routes into female headship.

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