EITC and Marriage

Abstract - *The Earned Income Tax Credit (EITC) could either penalize or subsidize marriage. Using data from the Survey of Income and Program Participation and controlling for individual fixed effects and the endogeneity of the EITC, we find that, among a sample of married women with children, those who face larger increases in their EITC are less likely to remain married. However, the effect is economically insignificant. We find no relationship between the EITC and marriage for unmarried women. We conclude that the EITC expansions during the early– to mid–1990s had little or no effect on marriage decisions.*

"[T]he federal government, through the EITC, says, 'If you get married, it'll cost you. And it'll cost you big time.'"

"Thanks to Washington, marriage is disappearing"
Houston Chronicle (Nov 9, 1998).

INTRODUCTION

With the expansion of the Earned Income Tax Credit (EITC) over the 1990s, the tax system plays an increased role in transferring income to low–income families. In both recipients and dollars, the EITC (18.5 million tax units received almost $26 billion) had surpassed the traditional cash welfare program of Aid to Families with Dependent Children (4.6 million families received total benefits of over $20 billion) by 1996 (U.S. Congress, 2000). Interest in whether the EITC affects marriage decisions accompanies the EITC's larger role in aiding low–income families.

Like traditional welfare programs, the EITC penalizes marriage for many families. However, because the EITC increases over certain earnings ranges before phasing out, some families actually face a marriage subsidy. For example, a marriage penalty exists if a single mother with earnings who is eligible for the EITC marries someone with earnings, and the couple's combined income places them beyond the phase–out range of the credit. Alternatively, if a single mother with no earnings marries a man with low earnings they will become eligible for the EITC; in this way, the EITC subsidizes marriage. The EITC, like the income tax system in general, tends to penalize marriage for women who work and to subsidize marriage for women who do not work, distorting

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not only the marriage decision but also the labor supply decision of married women.

In this paper we focus on whether the EITC is correlated with changes in marital status for families with children, the specific target of the EITC. Our hypothesis that the EITC may influence marriage is based on Becker's (1973, 1974) model of marriage, which suggests that the probability of choosing a particular family structure is a function of the expected gain and the distribution of unexpected outcomes from that family structure. The EITC changes the tax price of marriage and may affect the decision to marry or to end a marriage. In addition to the inefficiencies that arise from behavioral distortions of tax policy, the marriage incentives in the EITC also raise equity concerns. Some couples receive a higher EITC when they cohabit without legally marrying than similar couples who are married. The fact that two couples who differ only by legal marital status are treated differently by the tax system violates notions of horizontal equity.

Existing research suggests that taxes influence marriage decisions (see Alm et. al., 1999, for a summary of this literature) and that the EITC can create particularly large marriage penalties for low–income families (Holtzblatt and Rebelein, 2000; Dickert–Conlin and Houser, 1998; and the Congressional Budget Office, 1997). Two studies that consider the effect of the EITC on marriage reach opposing conclusions: Ellwood (2000) finds no evidence that the EITC influences marriage decisions, but some evidence that the EITC may have encouraged cohabitation over marriage. In contrast, Eissa and Hoynes (1999) estimate that the EITC increases marriage rates for very low income taxpayers and decreases marriage rates for middle income taxpayers.

In this paper, we exploit cross–sectional and cross–time variation in the EITC to estimate the correlation between the EITC and marriage using a panel of families with children drawn from the Survey of Income and Program Participation (SIPP). After controlling for individual fixed effects and time varying covariates, we find that the EITC does not, on average, influence marriage decisions.

This paper proceeds as follows: We first describe the marriage incentives generated by the EITC. Then we summarize the earlier literature on the marriage decision with respect to the tax and transfer systems. After describing our data, we explain our estimation procedure, and describe our results. Finally, we conclude with suggestions for future research.

MARRIAGE INCENTIVES IN THE EITC

The income tax code bases EITC eligibility 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.[1] The credit increases with earnings until it reaches a maximum (see Table 1 for all EITC parameters). 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 Treasury pays the difference to the filer. This makes it particularly relevant for low–income tax filers. Almost all recipients receive their

[1] A qualified child is a natural or adopted child or stepchild of taxpayers filing a joint or head of household return.

TABLE 1
FEDERAL EITC PARAMETERS

		Credit Rate (in percent)	Phase–In Range	Maximum Credit	Phase–Out Rate (in percent)	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
1998	No Children	7.65	0–4,460	341	7.65	5,570–10,030
	1 Child	34.00	0–6,680	2,271	15.98	12,260–26,473
	2+ Children	40.00	0–9,390	3,756	21.06	12,260–30,095

Source: U.S. Congress, 2000 and Internal Revenue Service, 1997.

EITC in a lump sum transfer with their tax return.[2]

Because the EITC does not explicitly depend on filing status, an increase in the credit may simultaneously increase the attractiveness of marriage and single–parenthood. For example, a mother of two children who earns $16,000 and a married couple with two children in which one spouse earns $16,000 in 2000 would both receive an EITC of $3,185. Increasing the EITC makes both marital states more financially attractive. Generally, the EITC subsidizes marriage for single–earner families and penalizes marriage for two–earner families.

The credit has changed a great deal in the past decade (see Table 1). The Tax Reform Act of 1986 legislated that the EITC be indexed for inflation so that by 1990, the nominal maximum benefit was $953, conditional on having at least one child. Following the passage of the 1990 Omnibus Budget Reconciliation Act, enacted in 1991, the maximum EITC value differed for families with one child versus families with two or more children. In addition, the maximum benefit levels increased; for example, the maximum benefit increased from $1,192 in 1991 to $1,434 in 1993 for a family with one child.[3] Another EITC expansion began with the 1993

[2] Although eligible workers have the option of receiving the credit in advance with their earnings, according to the General Accounting Office (1992), only 0.5 percent of EITC recipients do so.

[3] In 1991, 1992, and 1993 the EITC was greater for filing units with children under one year of age. This so–called "wee–tots" credit increased the maximum credit by $388 in 1993.

Omnibus Budget Reconciliation Act. Childless tax units became eligible for a small benefit, and there were large increases in the maximum credit for families with children. In 2000, the maximum benefit was $3,888 for a family with two or more children, $2,353 for a family with one child, and $353 for a family without children. The Joint Committee on Taxation estimates that over 18.4 million tax units will receive the EITC with credits totaling over $30.0 billion in 2000, an increase from 12.5 million families receiving credits totaling only $7.5 billion in 1990 (U.S. Congress, 2000).

The expansion of the federal EITC was accompanied by many similar credits in the states.[4] In the 2000 tax year, 14 states and the District of Columbia have earned income tax credits: Colorado, Illinois, Iowa, Kansas, Maine, Massachusetts, Maryland, Minnesota, New Jersey, New York, Oregon, Rhode Island, Vermont, and Wisconsin. All states calculate their EITCs 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 non–refundable in five of these states, making it less well–targeted toward low–income families than refundable EITCs.[5] In 1989, the first year of our data, only four states had EITCs and by 1995, the last year of our data, seven states had EITCs. The credit rates vary across states over this period. For example, in 1993, Iowa's EITC was 6.5 percent of the federal EITC and Wisconsin's EITC varied with the number of children and was 75 percent of the federal EITC for families with three children.

LITERATURE REVIEW

The literature that focuses specifically on the EITC and marriage includes descriptive evidence of the size of the EITC and empirical evidence of the effect of the EITC on marriage. In the descriptive literature, Dickert–Conlin and Houser (1998) document the source of penalties and subsidies implicit in the income tax and transfer systems. They show that the EITC in 1990 was a large contributor to subsidizing marriage for poor families and penalizing marriage for near–poor families; however, they also show that the EITC expansions were not able to offset the large marriage penalties that arise from the transfer system. More recently, Holtzblatt and Rebelein (2000) characterize the size of the EITC marriage penalties and subsidies under a variety of scenarios. They estimate that, in the aggregate, the 2000 EITC will increase total marriage penalties in the tax system by 10 percent and reduce bonuses by 1.5 percent. Like Dickert–Conlin and Houser, they find that the EITC will increase bonuses for taxpayers with low income—for taxpayers with incomes below $15,000 bonuses will increase by 24 percent.

Two recent papers that focus on the behavioral effect of the EITC with respect to marriage reach opposite conclusions. Eissa and Hoynes (1999) calculate family level income tax costs of marriage in repeated cross–sections of the 1985 to 1998 Current Population Surveys (CPS) and estimate how these tax costs affect the probability of marriage. They find that raising the tax cost of marriage, which includes the EITC, by $1,000 lowers the probability of marriage by 1.3 percentage points. Using simulations, they estimate that the EITC increases marriage rates by 1 to 5 percent for families with income below $25,000 and reduces marriage rates by 1 percent for families with incomes between $25,000 and $75,000. Despite a detailed calculation of the income tax costs of marriage, Eissa and Hoynes ignore the

[4] Nick Johnson from the Center on Budget and Policy Priorities provided us with the state EITC parameters.

[5] Non–refundability 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.

TABLE 2
STATE EITC PARAMETERS[a]

	First Tax Year	Refundable	Rate as Percent of Federal EITC	Without Qualifying Children[b]
CO	1999	Yes	1999: 8.5; 2000: 10	
DC	2000	Yes	2000: 10	Yes
IA	1990	No	1990–2000: 6.5	Yes
IL	2000	No	2000: 5	Yes
KS	1998	Yes	1998–2000: 10	Yes
ME	2000	No	2000: 5	Yes
MA	1997	Yes	1997–2000: 10; 2001: 15	Yes
MD	1987	Yes	1987–1997: 50 nonrefundable 1998–1999: { 50 *refundable*; 10 *nonrefundable* } 2000: { 50 *refundable*; 15 *nonrefundable* }	No
MN	1991	Yes	1991–1992: 10 1993–1997: 15 qualifying children 1998: { 15 no qualifying children; 20+ qualifying children } 1999–2000: { 15 no qualifying children; 20–46 qualifying children }	Yes
NJ	2000	Yes	2000: 10 (if income <$20,000)	No
NY	1994	Yes	1994: 7.5 1995: 10 1996–1999: 20 2000: 22.5	Yes
OR	1997	No	1997–2000: 5	Yes
RI	1975	No	1987: 23.46 1988–1990: 22.96 1991–1997: 27.5 1998: 27 1999: 26.5 2000: 26	Yes
VT	1988	Yes	1988: 23 1989: 25 1990–1993: 28 1994–1999: 25 2000: 32	Yes
WI[c]	1989	Yes	1989–1993: { 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 } 1996–2000 { 4 : 1 qualifying child; 14 : 2 qualifying children; 43 : 3+ qualifying children }	No

[a]Many thanks to Nick Johnson at the Center on Budget and Policy Priorities for providing 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.

potential endogeneity of their tax cost measure with the marriage decision.

In part to avoid this endogeneity issue, Ellwood (2000) uses a different identification strategy for considering whether the EITC affects marriage decisions. He posits that the EITC expansions during the 1990s increased the incentives for marriage among very low wage women and decreased the incentives for marriage among higher wage women. To test whether this change in incentives affected marriage decisions, he compares the marriage rates of women with low predicted wages to women with high predicted wages using data from the 1975 to 1999 CPS. Controlling for year effects, he finds no evidence of the marriage rates of low–wage women increased relative to those of high–wage women. Along another margin, Ellwood hypothesizes that the expansions in the EITC should have encouraged cohabiting couples facing marriage subsidies (i.e., couples with a single earner) to marry. He finds that cohabiting couples facing a marriage subsidy had an increase in marriage rates in 1996, shortly after the EITC expansions, suggestive of the EITC influencing marriage. However, he is cautious about drawing conclusions because the timing is not closely linked to the EITC expansions. He suggests that, "[A] more convincing test [of how the EITC affects marriage] will await longitudinal data to examine how marriage patterns of individuals changed over time as the incentive they faced changed."

More broadly, research on the relationship between income taxes and marriage decisions suggests that the tax system has small but statistically significant effects on marriage and divorce decisions. Alm et. al. (1999) summarize the literature as consistently finding evidence that as the tax penalty on marriage increases, individuals are less likely to marry and couples are more likely to divorce. The estimated magnitudes are small. For example, Alm and Whittington (1999) find that a 10 percent rise in the marriage penalty leads to a 2.3 percent reduction in the probability of first marriage.

An earlier literature on the negative income tax (NIT) experiments may be also relevant because the NIT targeted low–income families. Groeneveld, Tuma, and Hannan (1980) hypothesize that the NIT could either stabilize marriages by providing an income effect to married couples *or* weaken marriages by making NIT benefits available to unmarried individuals. Their statistical analysis concludes that participants had higher marital dissolution rates than non–participants over the duration of the experiments. However, Cain and Wissoker (1990) discount these results, suggesting that the temporary nature of the NIT experiments and weaknesses in the data make the NIT experiments inappropriate for making such conclusions.

There is a much larger body of research on the effect of the transfer system on family structure decisions. Moffitt (1998) concludes that a majority of recent studies find a positive correlation between welfare and female headship. The most common empirical analysis relies on the cross–state variation in the levels of AFDC benefits. However, a frequent criticism of much of the existing literature is that failure to control for unobserved state characteristics that are correlated with both welfare and marriage decisions may bias the coefficients on the welfare covariates. Two later studies by Moffitt (1994) and Hoynes (1997) that address this issue find little effect of AFDC on female headship.[6]

[6] In addition to the research on welfare and female headship, a smaller literature considers the inframarginal decisions of whether welfare influences the decision to cohabit or legally marry. Hu (1998) finds no consistent effect of welfare benefit levels on the likelihood of marriage relative to cohabitation. Hu (1998), Winkler (1995), and Schram and Wiseman (1988) find that generous welfare benefits for two–parent families are not positively correlated with the incidence of two–parent families. Moffitt, Reville, and Winkler (1998) find that states with lenient rules toward cohabitants have higher rates of combining cohabitation with welfare.

This paper extends the existing literature by analyzing the relationship between marriage and the EITC using longitudinal data in a framework that controls for individual fixed effects. As Ellwood (2000) suggests, we will draw our conclusions based on the changes in generosity of the EITC that individuals face over time.

DATA

We use a sample of women drawn from the 1990, 1991, 1992, and 1993 panels of the SIPP. The SIPP has relatively large sample sizes for a panel data set, although the panels are quite short. The SIPP divides households into four staggered rotation groups that are interviewed once every four months about their experiences during the past four months. A wave of the survey is completed when each of the rotation groups has been interviewed. The 1990 and 1991 panels each contain 8 waves; the 1992 panel contains 10 waves, and the 1993 panel contains 9 waves. These overlapping panels cover the period from October 1989 to December 1995.[7]

We include observations on women in December of each year so that we can identify their tax filing status; therefore, we have a maximum of three observations per person. Because we rely on cross state variation as part of identification, we drop observations from those states that are not uniquely identified in the SIPP.[8] To focus on women who are most likely to make marriage decisions, we limit our sample to women between the ages of 18 and 50. 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.[9] We drop 240 observations on women who become widowed and 147 observations in which marital status changes from married to never married during the panel. We also exclude women who report an increase of more than two years of education in a one year period or a decrease in years of education.[10]

We focus on women with children because families with children receive the largest EITCs and are most likely to be affected by changes in the credit. Table 3 shows the incidence of transitions in our data with the baseline year being the first year the individual was in our sample. Conditional on being married in the first year we observe them, there are 21,520 women with 56,250 person year observations (2.6 years per person on average). Between the first and the second year, 2.0 percent of the married women have divorced and by the third year we observe them, 3.3 percent have divorced. Conditional on being unmarried in the first year we observe them, there are 13,577 women with 32,866 person year observations (2.4 years per person on average). A total of 5.3 percent have married after one year and 9.9 percent have married by the third year.

ESTIMATION ISSUES

Based on Becker (1973, 1974) we hypothesize that marriage decisions depend on opportunities available within and outside of the marriage including the EITC, and other observable and unobserv-

[7] A 1996 panel of the SIPP exists. However, the years covered by the 1996 SIPP include vast state–level changes in the welfare system that hinder our ability to isolate behavioral changes from the EITC.

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

[9] Brien, Dickert–Conlin, and Weaver (2000) find compelling evidence that widows who face large penalties in the Social Security system for remarrying before age 60 are less likely to do so.

[10] This is 5,522 person years. We assume these are reporting errors, but including them does not change our results.

TABLE 3
ENTRIES AND EXITS INTO MARRIAGE
(IN PERCENT)

Year 1 \ Year 2	Married	Unmarried
Married with children (n = 20,164)	98.0	2.0
Unmarried with children (n = 9,755)	5.3	94.7

Year 1 \ Year 3	Married	Unmarried
Married with children (n = 16,484)	96.7	3.3
Unmarried with children (n = 7,307)	9.9	90.1

Source: Sample of women aged 18 to 50 from December of the 1990–1993 panels of SIPP. Year 1 is the first year we observe them in the data. Some women are not in the sample all for three years because of attrition or because they entered SIPP households after the start of the panel. We use SIPP weights.

able characteristics. Couples choose to marry if their utility within marriage is greater than the utility outside of marriage. If we assume that the indirect utility function is linear, we can express the underlying difference in utility between married and not married as:

$$M^*_{it} = \alpha + \gamma_1 e_{it} + X_{it}'\gamma_2 + y_t'\gamma_3 + v_{it},$$

where e_{it} is the combined federal and state EITC for state i in year t, X_{it} is a vector of individual and state characteristics, and y_t is a set of year dummies. M^*_{it} is empirically unobservable; however, we do observe whether individuals are married, M_{it}:[11]

$$M_{it} = \begin{cases} 1\ (\textit{married}) \text{ if } M^*_{it} > 0 \\ 0\ (\textit{not married}) \text{ if } M^*_{it} \le 0\,. \end{cases}$$

We use a tax calculator to estimate the individual's combined state and federal EITC, e, given income for the calendar year and demographic characteristics of the woman's family in December.[12] We adjust the EITC and other variables to 1995 dollars.

The vector of covariates, X, includes age, education levels, and a dummy for whether the woman lives in an urban area. The vector X also includes state–year characteristics that might be correlated with marriage decisions, such as the generosity of the welfare system, as captured by the maximum AFDC and food stamp benefits in the state;[13] the average real wage in manufacturing; real average per capita income; and the unemployment rate (U.S. Department of Commerce, various years). Finally, we include year dummies to capture any effects that affect all individuals in a single year, such as other changes in federal tax policy and business cycles. Descriptive statistics for our sample are in Appendix Table 1.

ENDOGENEITY

Two issues deserve discussion in our estimation. The first is our choice of the independent variable of interest, e_{it}, a measure of the individual's EITC. The actual EITC for which the family is eligible is endogenous with the marriage decision. In particular, the EITC depends on the

[11] We treat women who we believe are cohabiting as unmarried because the tax system would do so.

[12] We estimate annual income for those individuals who are not in the sample for the entire year by inflating reported income in the months that we observe to reflect 12 months. For example, if a woman is only in the sample for three months, we multiply her reported income by four as an estimate of annual income.

[13] Some states used waivers to alter their welfare programs before the 1996 reform. However, Meyer and Rosenbaum (2000) report that most waivers applied to only small parts of states and by 1994, fewer than 2 percent of single women were covered by major waivers.

family earnings, and earnings are clearly endogenous to marriage. For example, divorce is likely to lower family earnings, which may mechanically lead to an increase in a woman's EITC, although the EITC had no independent affect on her decision to divorce. In addition, high marginal tax rates on secondary earners in the phase out range of the EITC are likely to reduce married women's labor supply and mechanically increase the family's EITC (see Dickert, Houser, and Scholz, 1995; and Eissa and Hoynes, 1998 for evidence of this). If having a single earner reduces stress in the marriage and decreases the likelihood of divorce, we would observe a spurious positive correlation between the EITC and marriage.

Given this potential endogeneity, we need an instrument that is correlated with actual changes in the EITC but exogenous to the marriage decision. Assuming that the state and federal law changes are exogenous to individual behavior, we construct such an instrument in the following way: for each woman, we hold constant her family demographic and income characteristics from the first year she is in the sample and calculate what her EITC would have been in later years with those characteristics (adjusting income only for inflation).[14] The source of variation for identifying the effect of the EITC changes over time in EITC parameters and not any behavioral changes. In our estimations, we use a two–stage least squares approach.

FIXED EFFECTS

The second issue that deserves discussion in our estimation is the use of individual fixed effects. Although we observe a number of variables that are likely to influence the marriage decision it is likely that unobserved individual characteristics, such as attitudes toward marriage or divorce, also affect the marriage decision. Failure to control for these characteristics will bias our results. As an example, suppose women who are more committed to their marriages are less likely to work because they have more confidence in their husbands' financial support. Because single–earner families tend to have higher EITCs, we would observe a spurious positive relationship between the EITC and marriage.

To account for these fixed effects, we estimate:

$$M^*_{it} = \alpha_i + \gamma_1 e_{st} + X_{it}'\gamma_2 + y_t'\gamma_3 + v_{it}$$

where α_i is the individual time–invariant effect. The γ_i parameter estimates the effect of changes in the individual's EITC on the probability of marriage.

RESULTS

All Women with Children

Table 4 shows our results for the linear probability model with the dependent variable equal to 1 if the women is married.[15] In all cases, the standard errors reflect a correction for heteroskedasticity and the fact that individuals appear in the data multiple times. The first column shows the results when we pool the panel data and use each woman's calculated value of the EITC as our measure of the credit's generosity. This measure of the

[14] Carroll et. al. (2000) and Milligan (2000) construct similar instruments that fix individual characteristics in order to isolate the effects of policy changes.

[15] We choose to estimate a linear probability model because of our interest in instrumental variables and fixed effects. See Greene (1997) for a discussion of the limitations of a linear probability model for a binary dependent variable. As a specification check, we found that the coefficients in the pooled linear probability model are almost identical to the marginal effects in a pooled probit regression. We also estimated conditional logit models instead of linear probability fixed effects models and found qualitatively consistent results. Results are available upon request.

TABLE 4
LINEAR PROBABILITY MODEL FOR ALL WOMEN WITH CHILDREN
DEPENDENT VARIABLE: MARRIED[a]

	Pooled	Instrumental Variable	Individual Fixed Effect	Individual Fixed Effect Instrumental Variable
EITC (100)	–0.0096***	–0.0134***	–0.0038***	0.0009
	(0.0006)	(0.0009)	(0.0005)	(0.0045)
Age (years)	0.0174***	0.0172***	0.0034**	0.0019
	(0.0003)	(0.0003)	(0.0014)	(0.0019)
High School (1 = yes)	0.0352***	0.0306***	0.0142	0.0200**
	(0.0075)	(0.0076)	(0.0133)	(0.0119)
More than High School (1 = yes)	0.0146*	0.0058	0.0213	0.0320**
	(0.0075)	(0.0077)	(0.0143)	(0.0153)
Urban (1 = yes)	–0.0409***	–0.0426***	–0.0142	–0.0142
	(0.0058)	(0.0058)	(0.0184)	(0.0144)
State Maximum AFDC + Food Stamp (100)	0.0048*	0.0046*	–0.1092**	–0.0106***
	(0.0025)	(0.0025)	(0.0552)	(0.0043)
State Real Per Capita Income	–0.0053***	–0.0055***	0.0002	0.0023
	(0.0012)	(0.0012)	(0.0002)	(0.0016)
State Unemployment Rate	–0.0020*	–0.0019*	–0.0006	–0.0007
	(0.0011)	(0.0011)	(0.0006)	(0.0005)
State Manufacturing Wage	–0.0018	–0.0020	–0.0010	–0.0010
	(0.0015)	(0.0015)	(0.0010)	(0.0008)
1990 (1 = yes)	–0.0013	–0.0011	–0.0004	0.0009
	(0.0043)	(0.0043)	(0.0029)	(0.0027)
1991 (1 = yes)	0.0056	0.0069	0.0046	0.0061*
	(0.0050)	(0.0050)	(0.0038)	(0.0033)
1992 (1 = yes)	0.0182***	0.0203***	0.0027	0.0053
	(0.0064)	(0.0064)	(0.0041)	(0.0042)
1993 (1 = yes)	–0.0023	–0.0002	0.0080	0.0115
	(0.0070)	(0.0070)	(0.0080)	(0.0072)
1994 (1 = yes)	0.0014	0.0079	0.0104	0.0104*
	(0.0073)	(0.0075)	(0.0079)	(0.0061)
1995 (1 = yes)	0.0140	0.0223*		
	(0.0116)	(0.0117)		

[a]Sample of women between 18 and 50 years old in December of each year from the 1990–1993 panels of SIPP. 35,097 individuals; 89,116 observations. The standard errors reflect a correction for heteroskedasticity and the fact that the same individual is in the data more than one time.
*Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
*** Statistically significant at the 1 percent level.

EITC does not account for the likely endogeneity of the EITC with marital status. In this specification, the coefficient on the EITC variable (–0.0096) is negative and statistically significant (se = 0.0006) at standard levels, suggesting that the EITC discourages marriage. However, the magnitude of the relationship is small. A $100 increase in the EITC lowers the probability of being married by 1 percentage point. Given that the probability of being married is 67 percent and the average EITC is $186 for the women is our sample, this translates into an elasticity of –0.028 at the mean.

When we instrument for the actual EITC (results shown in Column 2 of Table 4) the coefficient on the EITC variable is still negative and statistically significant (coef. = –0.0134, s.e.= 0.0009). Although the coefficient on the EITC variable is larger in magnitude when using the instrumental variables approach, indicating that the endogeneity results in a positive bias, the magnitude of the relationship between the generosity of the EITC and marriage is still small. At the mean the estimated elasticity is –0.037.

In these specifications, age, education and welfare benefits are positively correlated with marriage at statistically significant levels. Living in an urban region, state per capita income and the unemployment rate are negatively correlated with marriage.

The coefficient on our variable of interest changes dramatically when we control for individual fixed effects. Using our estimate of the actual EITC as the independent variable, Column 3 shows a negative (–0.0038), and statistically significant (s.e. = 0.0005) correlation between the EITC and marriage. However, the economic effect is essentially zero: at the mean, the elasticity is –0.008. Finally, when we instrument for the EITC in the fixed effects model (the fourth column of Table 4), the coefficient on the EITC becomes positive (0.0009) but is statistically insignificant (s.e. = 0.0045).

Controlling for individual fixed effects, there is the expected negative and statistically significant effect of welfare generosity on the probability of marriage. Changes in education are positively correlated with marriage and later years are generally positively correlated with marriage.

Controlling for individual fixed effects and the endogeneity of the actual EITC leads us to the conclusion that the EITC does not affect the marriage decision. However, suppose that some variables have opposing effects on the decision to marry for single persons relative to the decision to stay married for married persons. Then, we may be obscuring that variation by pooling all women together. This seems particularly relevant for our variable of interest. For example, if large changes in the EITC increase a single woman's ability to live independently and simultaneously increase the joint income in a marriage and thereby stabilize the marriage, the effect of the EITC may appear to be zero in our pooled sample. To explore this possibility, we run similar regressions on two subsamples of women with children: a sample of women with children who were married in the first period of their SIPP panel and a sample of women with children who were unmarried in the first period of their SIPP panel.[16]

Sample of Married Women and Unmarried Women

Table 5 provides some evidence that the generosity of the EITC does not have symmetric effects on the decision of married persons to remain in a marriage (not divorce) and the decision of unmarried persons to marry. In this discussion, we focus on the EITC coefficients.[17] In both samples, our dependent variable is still one if the woman is married and zero otherwise.

In the Panel A of Table 5 we see that the EITC is negatively correlated with marriage for married women when we do not control for the endogeneity of the EITC or individual fixed effects. Although statistically significant (coef.= –0.0024, s.e.= 0.0003), the relationship between the EITC and marriage is not economically significant. A $100 increase in the EITC decreases

[16] An alternative sample split would be by some exogenous measure of earnings. See Ellwood (2000) for an example of this tactic.

[17] Full results for the two samples are available upon request.

TABLE 5
LINEAR PROBABILITY MODEL FOR WOMEN WHO WERE INITIALLY MARRIED WITH CHILDREN DEPENDENT VARIABLE: MARRIED[a]

	Pooled	Instrumental Variable	Individual Fixed Effect	Individual Fixed Effect Instrumental Variable
	Panel A: Initially Married Sample			
EITC (100)	–0.0024***	–0.0004	–0.0033***	0.0079**
	(0.0003)	(0.0003)	(0.0005)	(0.0040)
	Panel B: Initially Unmarried Sample			
EITC (100)	–0.0019***	–0.00004	–0.0043***	0.0026
	(0.0003)	(0.0004)	(0.0008)	(0.0053)

[a]Sample of women aged 18 to 50 in December of each year from the 1990–1993 panels of SIPP Married Sample: 21,520 individuals; 56,250 observations. Unmarried Sample: 13,577 individuals; 32,866 observations. The standard errors reflect a correction for heteroskedasticity and the fact that the same individual is in the data more than one time.
*Statistically significant at the 10 percent level.
**Statistically significant at the 5 percent level.
***Statistically significant at the 1 percent level.

the probability of marriage by 0.2 percentage points (an elasticity of –0.003 at the mean). However, as we show in column 2, the EITC is essentially uncorrelated (coef.= –0.0004, s.e.= 0.0003) with the decision to stay married when we use an instrumental variables estimation. The coefficient is negative but statistically insignificant and trivial in an economic sense. The fixed effects specification using the calculated EITC (column 3) yields results that are similar to the pooled specification; the correlation between the EITC and the probability of marriage is negative, statistically significant and very small (elasticity= –0.005 at the mean).

However, when we include individual fixed effects and instrument for the actual EITC (column 4), the coefficient on the EITC is positive (0.0079) and statistically significant (s.e.= 0.0040). That is, changes in the EITC are likely to encourage married women with children to remain married, but the economic effect remains small. Given that the probability of remaining married over the panel is approximately 97 percent (see Table 2), a \$100 increase in the EITC increases this probability by 0.8 percentage points. At the mean EITC of \$138 for this sample, this translates into an elasticity of only 0.011.

In the sample of unmarried women, Panel B of Table 5, the results are generally consistent with those for married women. As we showed for married women, the correlation between the EITC and marriage is negative, statistically significant, and small for initially unmarried women in the pooled model and the fixed effects model when we do not instrument for the EITC, and instrumenting for the EITC yields a correlation that is not different from zero. In addition, combining individual fixed effects with an instrumental variables approach for the EITC results in a positive coefficient for the initially unmarried women as for our married sample, but the coefficient is statistically insignificant at standard levels.

CONCLUSIONS

The goal of this paper is to consider whether the EITC is correlated with marriage decisions. The significant expansion in the EITC during the 1990s made this question a focus of policy debates and provides us with variation over time for the empirical study. The statutory structure of the EITC provides ambiguous predictions about the effect of the EITC on marriage. The EITC is available to families with married partners and single household heads, and an increase in the credit may increase the incentives for both states.

In our empirical work, we follow individuals in the Survey of Program Participation over time and consider how changes in the EITC affect their marriage behavior. Because the estimated EITC for individuals may be endogenous to the marriage decision, we instrument for the EITC with the EITC the individuals would get if only the law, and not their personal characteristics (income, demographics) changed over time. Our panel data strategy also allows us to account for characteristics that are fixed over time but unobservable to the researcher, such as attitudes toward marriage. Accounting for endogeneity and fixed effects, we find the EITC discourages divorce among married women. However, the elasticity of marriage with respect to the EITC is only 0.011. On the other hand, we find no relationship between the EITC and marriage for unmarried women using this same specification. Alternative models indicate that the EITC is negatively correlated with marriage for both married and unmarried women; however, the relationship is not economically significant. Our result that the EITC expansions during the early- to mid-1990s had little or no effect on the marriage decision is consistent with the work by Ellwood (2000) who uses different data and identification strategy.

A question for future research is whether couples are choosing to cohabit rather than marry. This issue has implications for whether the tax system is horizontally equitable. In our analysis we treated cohabiting couples as unmarried because the SIPP, prior to 1996, did not explicitly identify cohabiting couples (see Baughman et. al. 2000 for a discussion of identifying cohabitors in the SIPP). Along this same line, an additional direction for future research is an investigation into the misreporting of marital status. A 1994 report by the Treasury found that "25.8 percent of total EITC claimed, exceed the amount to which taxpayers were eligible" (Scholz, 1999). Among taxpayers with children, 31 percent of the errors were due to misreporting of filing status (Scholz, 1997). It may be those who face the largest penalties from marriage are least likely to report a change.

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APPENDIX TABLE 1
MEANS AND (STANDARD DEVIATIONS) FOR SELECTED VARIABLES

	Full Sample	Initially Married with Children Sample	Initially Unmarried with Children Sample
Married	0.67 (0.47)	0.99 (0.09)	0.04 (0.20)
EITC	186 (435)	138 (381)	268 (503)
Age (years)	34.58 (8.79)	36.83 (7.49)	30.72 (9.49)
High School (1 = yes)	0.41 (0.49)	0.42 (0.49)	0.40 (0.49)
More than High School (1 = yes)	0.44 (0.50)	0.45 (0.50)	0.42 (0.49)
Urban (1 = yes)	0.78 (0.42)	0.76 (0.43)	0.80 (0.40)
AFDC + Food Stamp (max family of 3)	716 (142)	715 (141)	720 (143)
State Real Per Capita Income	211 (31)	21.03 (3.08)	21.23 (3.18)
State Unemployment Rate	6.6 (1.9)	6.57 (1.86)	6.60 (1.85)
State Manufacturing Wage	12.62 (1.57)	12.61 (1.55)	12.62 (1.59)
1990 (1 = yes)	0.16 (0.36)	0.15 (0.36)	0.16 (0.37)
1991 (1 = yes)	0.23 (0.42)	0.23 (0.42)	0.24 (0.43)
1992 (1 = yes)	0.22 (0.42)	0.23 (0.42)	0.22 (0.41)
1993 (1 = yes)	0.18 (0.38)	0.18 (0.38)	0.17 (0.38)
1994 (1 = yes)	0.11 (0.31)	0.11 (0.31)	0.10 (0.30)
1995 (1 = yes)	0.02 (0.14)	0.02 (0.14)	0.02 (0.14)
N	89,124	56,252	32,866

Source: Sample of women between 18 and 50 years old in December of each year from the 1990–1993 panels of SIPP.