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The Impact of the Earned Income Tax Credit on Marriage and Divorce: Evidence from Flow Data

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Abstract While considerable research focuses on the anti-poverty and labor supply effects of the Earned Income Tax Credit (EITC), relatively little is known about the program's influence on marriage and divorce decisions. Furthermore, nearly all work in this area uses stock measures of marital status derived from survey data. In this paper, I draw upon Vital Statistics data between 1977 and 2004 to construct a transition-based measure of marriage and divorce rates. Flows into and out of marriage are advantageous because they are more likely to capture the immediate impact of policy changes. Controlling for state-level characteristics and sources of unobserved heterogeneity, I find that increases in the EITC are associated with reductions in new marriages, although the estimated effect is economically small. I find no relationship between the EITC and new divorces. These results are robust to alternative estimation strategies, data restrictions, and the inclusion of additional policy and demographic controls.

Keywords Earned Income Tax Credit · Flow data · Marriage penalties

Introduction

During the past three decades, the Earned Income Tax Credit (EITC) has become an essential piece of U.S. anti-poverty policy. Aside from its distributional features, the EITC creates strong employment incentives for low-skilled individuals. Indeed, results from recent empirical work suggest that the EITC is responsible for a large fraction of single mothers' employment growth throughout the 1990s (Fang and Keane 2004; Grogger 2003; Looney 2005; Meyer and Rosenbaum 2001). In addition

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to altering work incentives, the EITC may influence marriage decisions within the low-income population, a topic that has received far less attention from researchers.

The EITC is similar to other means-tested programs in that it penalizes marriage for some low-income families (Dickert-Conlin and Houser 2002). For example, if an employed single mother who is eligible for the EITC marries another low-wage worker, their combined earnings could make them ineligible for the credit. However, the program creates marriage subsidies for other individuals. Consider the subsidy that occurs when a non-working mother marries a childless man whose earnings are low enough to qualify for the credit. Individually, these parties are ineligible for the credit, but they become eligible after marriage. This implies that the EITC creates marriage penalties for dual-earner families and subsidizes marriage for those with a single earner.

As this discussion suggests, economic models of marriage and divorce are indeterminate with respect to the EITC, and results from the existing empirical literature reflect this ambiguity. Although a number of studies find that the EITC expansions throughout the 1990s may have aggravated marriage penalties for low-income individuals, empirical work on the behavioral impact of the credit produces mixed results. For example, Dickert-Conlin and Houser (2002) and Ellwood (2000) find little or no effect of the EITC on marriage decisions, while Rosenbaum (2000) estimates negative effects. Finally, Eissa and Hoynes (2004) find that the EITC increases marriage among low-income families, but decreases marriage among moderate-income families.

Nevertheless, it has become increasingly important to understand the relationship between the EITC and marriage decisions. Concerns over this policy's impact on marriage are part of a larger debate on the income tax code, which contains numerous provisions that reward or penalize marriage.[1] The EITC also raises concerns over horizontal equity because it treats legally married and unmarried (or cohabitating) couples differently, even though these couples might have identical incomes and family responsibilities. Furthermore, many low-income individuals already face marriage penalties from several means-tested programs, including Temporary Assistance to Needy Families (TANF). Finally, changes to the EITC in President Bush's 2001 tax law are intended to alleviate marriage penalties by expanding eligibility for couples filing jointly. As couples adjust to the new EITC structure, it is important to know whether such changes are likely to alter marriage decisions.

In this paper, I examine the impact of the EITC on marriage and divorce using flow data from the national Vital Statistics system. Vital Statistics record the number of new marriages and divorces within the previous year, allowing researchers to construct incidence-based measures of marriage and divorce rates. These data provide a number of advantages over survey-based measures. For example, previous studies using the Current Population Survey (CPS) or the Survey of Income and Program Participation (SIPP) focus on stocks, or the total population

[1] For example, a 1996 study by the U.S. General Accounting Office (1996) found at least 59 provisions that reward or penalize marriage. Since the date of this study, several important pieces of tax legislation were passed that have implications for non-neutrality. For example, the 1997 Taxpayer Relief Act created the child tax credit, and the 2001 Economic Growth and Tax Relief Reconciliation Act (EGTRRA) altered the EITC (among other things) to alleviate marriage penalties in the tax code.

in a given marital category. As others have shown, studying entries into and out of marriage provide different estimates of the effects of policy changes because the population of recently married (divorced) differs from the overall married (divorced) population (Bitler et al. 2004; Fitzgerald and Ribar 2004; Rosenbaum 2000). Flow data also capture the immediate impact of policy changes, whereas stock measures are better equipped to examine long-term effects. Another advantage of the Vital Statistics data is that they contain virtually the complete universe of marriages and divorces, as opposed to CPS and SIPP data, which are prone to underreporting and measurement error.

Using state-level panel data on new marriages and divorces over the period 1977–2004, I find some evidence that the EITC is correlated with marriage flows. After controlling for states' socio-economic characteristics, recent social policy reforms, and unobserved heterogeneity, the preferred model suggests that a $1,000 increase in the maximum credit decreases the new marriage rate by 4.9%, although the estimates range from 3.5 to 9.2% depending on the specification. Furthermore, I find no relationship between the EITC and new divorces.

Conceptual Framework for Understanding Marriage Incentives in the EITC

EITC Structure and the Sources of Marriage Subsidies and Penalties

To understand how the EITC influences marriage decisions, it is important highlight some of its key design features. First, individuals must have positive earned income to be eligible for the credit. Second, adjusted gross income must be below some threshold, which varies by year and the presence and number of children.[2] Third, the EITC comprises three regions—phase-in, plateau, and phase-out—with the first of these regions creating a wage subsidy of 34% for families with one child and 40% for families with two or more children. The credit currently phases out at a rate of 15.98 and 21.06% for one- and multiple-child families, respectively. Finally, 1986, 1990, and 1993 tax legislation significantly raised the phase-in rate and maximum credit available to eligible tax units. These laws also increased the phase-out rate, thereby curtailing continued expansions of the EITC-eligible population.[3]

Another important development is the proliferation of state EITC programs. These programs simply "piggyback" onto the federal EITC by using its eligibility rules and credit rates. States have the option to structure their EITC programs as refundable or nonrefundable tax credits. Of the 17 states (including the District of Columbia) that implemented an EITC between 1977 and 2004, 11 made the program refundable for the entire period of operation, while two states (Illinois and Maryland) changed from nonrefundable to refundable tax credits. This distinction is

[2] Additional income from interest, dividends, and capital gains cannot exceed $2,600.

[3] The creation of a separate EITC schedule for childless workers in 1993 is another important design feature. With a phase-in rate of 7.65%, its maximum credit is substantially smaller than those for one- and multiple-child families. Nevertheless, this EITC was created to mitigate the marriage and fertility distortions imbedded in the original design.

important because many low-income families do not have income tax liabilities, and therefore cannot qualify for nonrefundable tax credits. Annual foregone revenue from state EITCs ranges from $17 million in Vermont to $591 million in New York (Nagle and Johnson 2006). Table 1 presents some key characteristics of these tax credits.

These design features create complicated marriage incentives by subsidizing it for some wage earners and penalizing it for others. For example, the phase-in region creates marriage subsidies and penalties for extremely low-income workers. When two such individuals marry (one of whom has a qualifying child), their combined earnings yield a larger EITC than the cumulative payment they would receive as single tax filers. Alternatively, a childless couple in the phase-in region can experience marriage neutrality or penalties, depending on whether their combined earnings exceed the top of the phase-in region. An important point here is that the presence and number of children can alter the nature of subsidies and penalties.

Table 1 Key features of state EITC programs, 1977–2004

State/district	Implementation year	Original rate: % of federal EITC	Current rate: % of federal EITC	Refundable
Colorado	1999, suspended after 2002	8.5	10.0	Yes
District of Columbia	2000	10.0	35.5	Yes
Iowa	1990	6.5	6.5	No
Illinois	2000	5.0	5.0	Yes, as of 2004
Indiana	2002	6.0	6.0	Yes
Kansas	1998	10.0	15.0	Yes
Maine	2000	5.0	5.0	No
Maryland	1987	50.0	20.0	Yes, as of 1998
Massachusetts	1997	10.0	15.0	Yes
Minnesota	1991	10.0	25.0–45.0	Yes
New Jersey	2000	10	20.0	Yes
New York	1994	7.5	30.0	Yes
Oklahoma	2002	5.0	5.0	Yes
Oregon	1997	5.0	5.0	No
Rhode Island	1975	17.0	25.0	No
Vermont	1988	23.0	32.0	Yes
Wisconsin	1984	30.0	4, 14, 43% for 1, 2, 3 children	Yes

Notes: Data in this table come from Dickert-Conlin and Houser (2002), Fang and Keane (2004), and various publications from the Center on Budget and Policy Priorities. Although Colorado suspended its EITC program due to insufficient funds. Maryland offers a non-refundable tax credit, set at 50% of the federal EITC. Minnesota's tax credit is not explicitly set up to follow the federal EITC. The range depicted in the table is the credit amount which depends on family income. There is also a 25% credit for childless families. New Jersey limits eligibility to its EITC program to families with incomes less than $20K. Rhode Island made its EITC partially refundable in 2003, and the refundable portion of the credit increased from 5 to 10%

Such complexities also apply to individuals within the phase-out region. Generally speaking, the phase-out region creates marriage penalties because of the implicit tax on earnings. There are, however, numerous exceptions to this. For example, if a childless worker with earnings too high to be eligible for the EITC marries a very low-income individual with one child, their cumulative earnings might place them in the phase-out region, but with an EITC that far exceeds their individual credits. Finally, a clear marriage penalty is embedded in the EITC's eligibility threshold. Two single workers with earnings low enough to be eligible for the credit could lose eligibility when they marry if their combined earnings places them above the EITC's break-even point.

In addition, the number and distribution of children has implications for the presence of marriage subsidies and penalties. Before the creation of a separate credit for multiple-child families, the cumulative EITC paid to two single parents, each of whom had one child, was substantially greater than the credit they would receive after marriage. Legislative changes in 1990 and 1993 mitigates some of these penalties by creating separate benefit schedules for families with two or more qualifying children as well as taxpayers without children. Finally, because they supplement the federal credit, state EITC programs—especially refundable credits—tend to magnify the marriage subsidies and penalties that already exist.

The impact of such design features on marriage subsidies and penalties is best highlighted by considering a number of hypothetical couples before and after marriage. Figure 1 examines the role of the EITC in altering marriage incentives for two illustrative families in Texas and Wisconsin between 1977 and 2004.[4] Plotted here is the differential federal/state income tax liability for a mother with two children and a potential spouse before and after marriage.[5] A marriage penalty exists when individuals have greater combined tax liabilities after marriage, relative to their individual liabilities when they are single, and a marriage subsidy exists when individuals have lower liabilities as a married couple.

A number of interesting patterns emerge from Fig. 1. First, the magnitude of marriage subsidies and penalties changed dramatically between 1977 and 2004, with sizable disparities throughout the late-1970s, a movement toward tax neutrality throughout the 1980s, and then increased subsidies and penalties beginning in the late-1980s and early-1990s. The growing non-neutrality in the most recent period coincides with the first major EITC expansion in 1986 and the third expansion in 1993. Second, a marriage subsidy exists in all years among families in which the mother does not work and the spouse's earnings place them below the eligibility

[4] I focus on Texas and Wisconsin for several reasons. Since Texas does not have a state income tax, it illustrates a pure federal EITC effect. Wisconsin, other the other hand, has relatively high state income taxes but also has one of the most generous state EITC programs. See details in text about the Wisconsin program. Therefore, these states allow one to examine the ways in which state EITCs likely amplify the marriage incentives embedded in the federal credit.

[5] These simulations, which use NBER's TAXSIM model, are conducted in the following manner. All families take the relevant standard deduction and are assumed to have zero unearned income or child care expenses. Unmarried women with children file as "head of household" and claim their children as dependents, unmarried men file as "single", and married couples file as "married filing jointly." All dollars amounts are adjusted to reflect 2004 prices. Positive numbers in Fig. 1 indicate marriage subsidies, and negative numbers indicate marriage penalties.

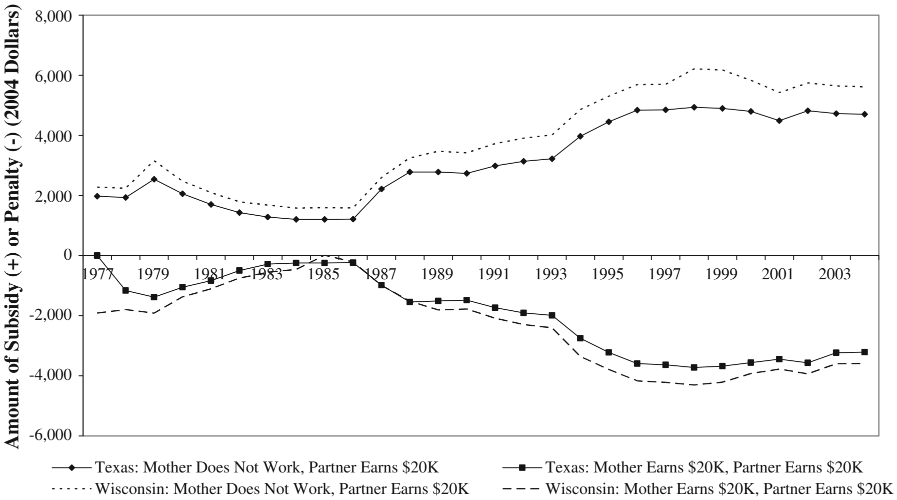

Fig. 1 The impact of the EITC (and other tax policies) on marriage penalties and subsidies, 1977–2004: Differential income tax liabilities before and after marriage for families with two children

limit. By 2004, such families received a $4,700 marriage subsidy. Conversely, families in which both spouses' earnings are equal ($20,000) experience a marriage penalty, with the penalty reaching $3,200 in 2004. Finally, Fig. 1 highlights the growing importance of state EITCs in magnifying the effects of the federal credit. Married couples in Wisconsin, a state with a generous EITC program, experience deeper marriage subsidies and penalties than those in Texas, a state without an income tax system.

Marriage and Divorce Incentives in the EITC

Economic models of marriage and divorce predict that individuals marry when the anticipated utility (benefits minus costs) of entering marriage exceeds the utility of remaining single (Becker 1973). An individual's utility from being single is specified as a function of own earnings, non-wage income (including means-tested benefits), and demographic/human capital characteristics, including age, race, and education. The utility associated with marriage depends on own earnings, the spouse's income, non-wage income, and all other individual characteristics. The likelihood of choosing a given marital status is therefore a function of the expected gains and losses associated with marriage or remaining single.[6,7]

[6] Critical to the EITC is the model's extension to account for cohabitation among partners (Clarkberg et al. 1995; Ellwood 2000; Ellwood and Jencks 2001). See Acs and Maag (2005) for a discussion of the marriage subsidies and penalties that arise in the tax and transfer system for cohabitating couples.

[7] It is important to note that several other theoretical models have rendered important insights on marriage and divorce decisions. For example, search models have been used to examine whether the quality of local marriage markets influences marriage decisions generally and the extent of assortive mating specifically (Lichter et al. 1995). Other studies focus on the importance of individual's observable

There are at least three reasons why the standard marriage model yields ambiguous predictions with respect to the EITC. First, as discussed in the previous section, the credit's design mechanically alters marriage incentives. Changes in actual behavior, however, are related to the magnitude and distribution of costs and benefits after an EITC is introduced. Numerous studies attempt to quantify the size of these aggregate shifts (Congressional Budget Office 1997; Dickert-Conlin and Houser 1998; Hoffman 2003; Holtzblatt and Rebelein 2000). For example, Holtzblatt and Rebelein (2000) show that in 2000, the EITC increased overall marriage penalties by 10.4% and reduced marriage subsidies by 1.5%, contributing a net marriage penalty of \$3.6 billion. The authors show that introducing an EITC increased marriage subsidies by 24% for low-income families (earning less than \$15,000), for a small net bonus of \$54 million. However, all other income groups suffered substantial marriage penalties as a result of the EITC, although over half of these penalties (55%) were incurred by families above the EITC's eligibility threshold.

The second factor that leads to ambiguous theoretical predictions is the interaction of marriage and labor supply decisions. The introduction of an EITC raises own-earnings for low-income workers, which simultaneously increases the utility associated with being married and single. Economic models provide unclear predictions for the impact of earnings on marriage and divorce, but the net effect depends on how increased earnings affect preferences for marriage as well as how partners share economic resources. As a result, the EITC could have an "independence effect" on women in which an increase in economic resources discourages marriage. However, the credit could also lead to a "stabilizing effect" in which the labor supply effects of the EITC increase women's exposure in the marriage market, lead to more secure marriages, and reduce the incidence of divorce. Empirical studies generally find that increased own-earnings are negatively related to marriages rates for women (implying that the independence effect dominates) and positively related to marriage rates for men (implying that the stabilizing effect dominates) (Blau et al. 2000; Brien 1997; Lichter et al. 2002). The literature is less clear, however, on the impact of earnings on divorce (Ellwood and Jencks 2001; Lichter et al. 2002; Ressler and Waters 2000).

Finally, the EITC could influence marriage decisions through its impact on fertility. By targeting most benefits to families with children, the credit lowers the costs associated with childrearing, thereby creating pro-natalist incentives. However, because the EITC acts as a wage subsidy and increases the returns to market work, the credit raises the opportunity costs associated with bearing and raising children. Therefore, one might also expect the EITC to reduce fertility. Several strands of the empirical literature are relevant to this issue. There is a growing body of work on the relationship between income taxes and fertility, which generally finds pro-natalist effects of the personal exemption and other tax provisions (Whittington 1992; Whittington et al. 1990). Other studies examine the impact of the EITC on fertility. This research provides mixed evidence, with Duchovny

Footnote 7 continued
characteristics (e.g., race and ethnicity and educational attainment) in driving marriage patterns (Blackwell and Lichter 2000).

(2001) finding increased fertility among married, white women and Baughman and Dickert-Conlin (2009) finding reductions in higher-order fertility among white women. Finally, a large empirical literature examines the effect of childbearing on marriage and divorce. These studies find that non-marital childbearing reduces the likelihood of subsequent marriage (Bennett et al. 1995; Joyce et al. 2001; Qian et al. 2005), while marital childbearing tends to have a stabilizing effect on married couples (Cherlin 1977; Joyce et al. 2001; Steele et al. 2005).

Direct evidence on the impact of the EITC on marriage and divorce decisions reflect this theoretical ambiguity.[8] Eissa and Hoynes (2004) use CPS samples between 1985 and 1998 to characterize the income tax "costs" associated with marriage. Using as the dependent variable the current marital status of individuals (i.e., marriage stocks), the authors find that a $1,000 increase in the cost of marriage decreases the marriage rate by 1.3% points. However, simulation results suggest that increases in the EITC raise marriage rates among low-income families by as much as 5%. Ellwood (2000) also uses a stock measure of marriage in CPS samples between 1975 and 1999 to compare marriage rates between women with low and high predicted earnings. He finds no evidence that the EITC is responsible for the secular decline in marriage over this period.

Only two studies have explored transition-based (i.e., flow) measures of marriage outcomes, both using individual-level survey data. A working paper by Rosenbaum (2000) uses the CPS and SIPP to explore stock and flow measures of marriage outcomes. Results suggest that the EITC can have large negative effects on marriage, especially marriage entries, but the estimates are sensitive to the way tax costs are specified in the model. Finally, a study by Dickert-Conlin and Houser (2002) uses several panels of SIPP data to explore the impact of individuals' combined federal and state EITC on marriage rates and transitions. These authors find that increases in the EITC lead to small negative effects on the probability of marriage but inconsistent effects on marriage flows.

In sum, the anticipated effect of the EITC on marriage flows is ambiguous. The credit's design could either reward or penalize marriage depending on the partner's combined earnings, number of children, and state of residence. Economic theory suggests that the labor supply effects of the EITC could discourage marriage through an independence effect, or it could promote marriage through a stabilizing effect. Finally, the EITC distorts fertility decisions, which could further influence

[8] A large literature examines the impact of various components of the income tax system on marital behavior. See Alm and Whittington (1995, 1997), Dickert-Conlin (1999), and Whittington and Alm (1997) for a sampling of work on this topic. This work generally finds that taxes affect both the decision to enter and end a marriage, as well as the timing of these decisions. Also relevant to this study is the large literature on the effects of traditional cash assistance programs on marriage and living arrangements. Moffitt's (1998) review concludes that there appears to be a small positive relationship between welfare benefits and female headship. However, as noted in Moffitt (1994) and Hoynes (1997), much if not all of the positive effect is driven by unobserved heterogeneity. Including fixed effects in such models renders the welfare effect insignificant. More recent evidence by Fitzgerald and Ribar (2004) and Dickert-Conlin and Houser (1999) confirm findings from earlier work. Another cluster of studies examines in the impact of welfare reform on marriage, divorce, and headship decisions, with some finding positive effects of reform (Schoeni and Blank 2000), others finding negative effects (Bitler et al. 2004; Horvath-Rose and Peters 2001), and still others finding inconsistent or insignificant effects (Fitzgerald and Ribar 2004; Ellwood 2000; Kaushal and Kaestner 2001).

flows into and out of marriage. Therefore, the overall impact of the EITC depends on which of these design features and incentive-effects dominate.

Empirical Implementation

Key Data Source: National Vital Statistics

I construct a state-by-year dataset of the number of marriages and divorces using Vital Statistics between 1977 and 2004. This period covers nearly the entire history and every major legislative change associated with the EITC.[9] Vital Statistics on the number of marriage occurrences are available for all state-year combinations except in California and Oklahoma, while data on the number of new divorces are missing for California, Colorado, Georgia, Hawaii, Indiana, Louisiana, Nevada, and Oklahoma.[10] Given these missing data, I estimate marriage and divorce regressions using the balanced panel of states.[11] The marriage analyses are limited to 1,372 state-year combinations, and the divorce analyses are limited to 1,204 combinations.

Nearly all studies on the impact of the EITC on marriage and divorce examine the stock of married and divorced women with children, and all studies use individual-level survey data (e.g., CPS and SIPP). These studies estimate the prevalence of such outcomes within a given population and average the effects of the EITC over that entire population. Vital Statistics data allow me to construct a flow measure of marriage and divorce, or recent entries to and exits from marriage. Flow data capture the incidence of such transitions, and as a result, the effects of the EITC are averaged over a population whose observable characteristics differ from those in prevalence-based samples. Analyses by Bitler et al. (2004) and the U.S. Census Bureau (2004) confirm that the newly married and divorced are younger, less-skilled, less likely to be employed, and more likely to be poor. Therefore, Vital Statistics data capture marital behavior within a population that is more likely to respond to changes in social policy reforms.

Flow data have been used extensively by researchers in similar contexts, and they offer several advantages over survey-based measures of marriage and divorce stocks (Bitler et al. 2004; Fitzgerald and Ribar 2004; Moffitt and Rendall 1995; Blank 1999; Gittleman 2001; Wolfers 2006). First, CPS and SIPP measures of marriage

[9] I omit 1975 and 1976 from the analysis because the data used for the denominator in the marriage and divorce rates (number of unmarried and married women, respectively, ages 15 and over) are drawn from the CPS, which does not have these data available for these years.

[10] California is missing new marriages for 1991, and Oklahoma is missing these data for 2001–2003. Divorce data are missing for the following state-years: California (1991–2004), Colorado (1995–2000), Georgia (2004), Hawaii (2003–2004), Indiana (1991–2004), Louisiana (1991–2004), Nevada (1991–1993), and Oklahoma (2001–2003).

[11] Using an unbalanced panel of state-years causes a few concerns. If states with missing marriage/divorce data are trending differently, it would be difficult to distinguish the impact of federal EITC expansions or the implementation of a state EITC from other state-specific factors. Moreover, a number of states are missing information during a period when important expansions to the federal credit were implemented (1991–1993; 1991–2004; 1995–2000). In a later section, however, I test the sensitivity of the main results by estimating the model on the unbalanced panel of states.

and divorce are underreported, and these datasets contain measurement error arising from misreporting of marital status, incomplete or missing data, and respondent confusion over survey questions (Goldstein 1999; Thorton and Rodgers 1987). In contrast, Vital Statistics provide nearly the entire universe of marriage and divorce occurrences within the previous year, and these counts are not affected by the biases found in survey datasets. Second, given that flow data record the number of marriages and divorces within the previous 12 months, they are well-equipped for studying the short-run impact of policy changes.

These data are, however, not without their drawbacks. Divorce counts are missing for a nontrivial number of states, three of which have state EITC programs (Colorado, Indiana, and Oklahoma). The loss of these states reduces the amount of plausibly exogenous variation in the EITC. Marriage and divorce occurrences are measured at the state-level, and so it is not possible to model individual behavior. Furthermore, Vital Statistics provide little information on background characteristics, precluding sub-group analyses by race or level of educational attainment. These data also report new marriages and divorces by the state of occurrence, rather than residence. Hawaii and Nevada, for example, contain a large number of marriages to nonresidents. Generally speaking, an implication of this cross-state marriage migration is that migration decisions may be correlated with states' policy reforms, leading to biased estimates of those reforms. However, this is not likely to be a problem for the present analysis, given that the federal component of the EITC exposes individuals to identical program rules regardless of the state of residence.

Following Bitler et al. (2004), I express marriage and divorce flows as a rate by dividing by the population of women who could plausibly transition to a marriage or divorce: unmarried women for new marriages and married women for new divorces. I calculate these population estimates for women ages 15 and over using March CPS samples between 1977 and 2004. I expand the denominator to include women as young as age 15 because two states do not have a statutory minimum age (Arizona and California) and 11 states (Georgia, Hawaii, Missouri, Indiana, Michigan, Mississippi, New Hampshire, New York, Pennsylvania, Texas, and Utah) permit marriages among 15-year-olds (and younger) with parental consent.

Empirical Model

Using state-level Vital Statistics data on new marriages and divorces between 1977 and 2004, I estimate permutations of the following OLS regression model:

$$y_{st} = E_{st-2}\,\beta + \mathbf{W}'_{st}\,\psi + \mathbf{M}'_{st}\,\gamma + \mathbf{D}'_{st}\,\theta + \mathbf{P}'_{st}\,\eta + \mu_s + \upsilon_t + (\text{trend} \times \mu_s) + \varepsilon_{st}, \quad (1)$$

for $s = 1, \ldots, S$; $t = 1, \ldots, N$, where s indexes states and t indexes years. The dependent variable, y_{st}, is either the new marriage or new divorce rate, defined as the number of marriages (divorces) within the previous year divided by the number of unmarried (married) women ages 15 and over. In specification checks, I estimate models in which the dependent variable is the number of new marriages (divorces), adding state population as a control, and models that correct for autocorrelation using a panel data generalization of the Prais-Winsten method. Indeed, tests reveal that the errors in the marriage and divorce regressions are serially correlated within states.[12] All outcome variables are expressed in logarithmic form.

The key variable in this model is denoted by E_{st-2}, the combined federal and state EITC maximum credit, lagged 2 years. This lag structure properly accounts for two important time gaps: one occurring between the time the EITC is earned and when it is actually paid by the Internal Revenue Service (IRS), and another occurring between the time individual marriage and divorce decisions are influenced by the EITC and when the implications of those decisions manifest in the data.[13] The parameter of interest is β, which captures the average change in the log of the new marriage and divorce rate when the EITC maximum credit increases by \$1,000.

The $\mathbf{W}'_{st}$ denotes a vector of state welfare policies that may influence marriage and divorce decisions, including the maximum AFDC/TANF benefit for a three-person family, a dummy variable that equals unity in the year after the implementation of a statewide welfare waiver, and a dummy variable that equals unity in the year after the implementation of TANF. In the year of implementation, I follow the standard practice in the literature by setting both welfare reform variables equal to the fraction of the year in which they were “turned on.” The $\mathbf{M}'_{st}$ is a vector of state-level macroeconomic indicators, including the average annual unemployment rate, combined employment in the service and retail industries, and average annual earnings in the service and retail industries. I also include a number of demographic and human capital characteristics, $\mathbf{D}'_{st}$, that may shift preferences for marriage and divorce. These variables include population density, percent African American population, female age structure, and educational attainment. Differences over time and across states in the political and policy-making environments are controlled for by $\mathbf{P}'_{st}$, which includes a dummy variable that equals unity for all state-years in which there is a Republican governor, as well as a continuous variable measuring the fraction of each state’s population voting Republican in the previous Presidential election.

[12] A Wooldridge (2002) test of no serial correlation in panel data yields a highly significant F-statistic in the marriage ($F = 34.41$) and divorce ($F = 27.26$) regressions. The Prais-Winsten regression assumes that the autocorrelation follows an AR(1) process that is common across all states. Standard errors in these models are also corrected for heteroskedasticity.

[13] In particular, the EITC is typically paid by the IRS in the calendar year after the tax year in which it was earned (Dickert-Conlin and Houser 1999). Also, there are a number of mechanisms used by states to slow the decision-making process prior to a marriage or divorce. For example, three states (Arizona, Arkansas, and Louisiana) have covenant marriage options, which stipulate that couples must receive counseling before seeking a divorce. A number of states specify a waiting period after a divorce before an individual can re-marry. There are also so-called “cooling off” periods between the time a divorce is petitioned and when it is actually granted by a judge. These waiting periods vary dramatically by state, and can range from no statutory requirement to 18 months (Arkansas, Connecticut, and New Jersey). On the other hand, states have also passed no fault divorce laws, which allow partners to seek a divorce without gaining the consent of the spouse. See research by Friedberg (1998) and Wolfers (2006) for an evaluation of the impact of no fault laws on divorce. To examine the robustness of the main results to the choice of lag structure, I experiment with a one-period lag of the EITC. Results from this specification are discussed in a subsequent section. There are, however, reasons to be believe that a 2-year lag structure is not long enough. Low-income couples are more likely to separate for substantial periods, and hence stop filing taxes jointly, before a divorce is formalized. It is difficult to know whether an EITC-induced divorce would catalyze a long separation period, but suffice it to say that a 2-year lag on the EITC may not pick up all divorces that occur after a long-term separation. I thank an anonymous referee for pointing out this possibility.

The regression model also includes controls for sources of unobserved heterogeneity, which may simultaneously influence the generosity of the EITC and marriage propensities. State fixed effects, μ_s, control for time-invariant differences between states, year fixed effects, υ_t, filter out year-to-year changes in marriage propensities that are common to all states, and state-specific time trends, (trend $\times$ μ_s), control for factors that are trending linearly within states over the observation period. These controls are essential because they allow us to distinguish between the impact of the EITC and changes in marriage markets or unobserved attitudes toward marriage. All regressions are weighted by the number of women ages 15 and over, and standard errors are corrected for arbitrary heteroskedasticity using robust standard errors. The impact of the EITC is well identified in this model. The study period includes three major expansions (in 1986, 1990, and 1993), each one substantially increasing the maximum credit. The largest of these expansions came in 1993, when the phase-in rate gradually increased from 19.5 to 40.0% for families with two or more children. Growth in the federal credit has been strengthened by the proliferation of state EITCs. As shown in Table 1, 16 states operated their own program as of 2004, thus becoming an increasingly important mechanism for offsetting state income taxes. Twice-lagging the EITC should also help mitigate the influence of omitted factors that are correlated with the contemporaneous benefit and marriage decisions. Finally, the federal credit was adjusted for inflation in 1986, approximately ten years after its implementation, providing another source of policy variation during the observation period. The key assumption in using the combined federal and state EITC maximum credit is that all legislative changes are exogenous to state-level marriage and divorce decisions.[14] This assumption is plausible, especially for the federal component of the EITC, since low-income workers in all states are exposed to the same eligibility rules and benefit schedules.

Results

Main Results

Table 2 presents summary statistics for the state-level panel data between 1977 and 2004. It shows that 5.6% of unmarried women transition into marriage and 1.9% of married women transition out of marriage each year. Rates of new marriages and divorces declined monotonically during the study period, at an average annual rate of 1.8 and 0.7%, respectively.[15] The maximum value for the new marriage rate is

[14] The potential endogeneity of the EITC is a manifestation of the more general policy endogeneity problem. The concern here is that the timing of EITC expansions coincides with, or responds to, broader societal trends that influence marriage and divorce propensities. For example, if states that enact EITC programs to reduce poverty or strengthen work incentives have populations with different underlying tastes for marriage, then EITC policies will be correlated with these unobserved preferences and the estimated effect of the credit on marriage and divorce will be biased. As noted in the text, that the EITC is primarily a federal program, however, makes policy endogeneity less likely.

[15] To calculate this number, I regressed the log of the new marriage (divorce) rate on a linear time trend (with robust standard errors and weighted by the number of women ages 15 and over).

Table 2 Summary statistics for state panel data, 1977–2004

Variable	Mean	SD	Minimum	Maximum
Marriage rate (per 1,000 unmarried women)	0.056	0.048	0.014	1.369
Divorce rate (per 1,000 married women)	0.019	0.004	0.008	0.045
Number of new marriages (vital statistics)	76,575	50,769	2,372	210,978
Number of new divorces (vital statistics)	38,758	25,654	1,043	106,354
Federal/state EITC maximum credit ($1,000)				
One child	1.294	0.835	0.400	3.315
Two or more children	1.801	1.512	0.400	5.484
Implementation of statewide welfare waiver	0.059	0.223	0.000	1.000
Implementation of TANF	0.311	0.457	0.000	1.000
Maximum monthly AFDC/TANF benefit ($1,000)	0.373	0.150	0.067	0.924
State unemployment rate	0.062	0.019	0.022	0.180
Combined service/retail employment (per 1,000)	2,706	2,304	68.13	9,988
Combined service/retail earnings ($1,000)	15.530	5.905	4.545	48.880
Population density (persons psm, per 100)	2.475	5.431	0.006	109.744
Percent black population ages 15+	0.111	0.074	0.002	0.669
Percent of female population ages 15–39	0.476	0.042	0.370	0.700
Percent of female population ages 40–64	0.344	0.030	0.260	0.456
Percent ages 25+ with a BA degree	0.212	0.052	0.098	0.464
Fertility rate (births per 1,000 women ages 15–44)	66.50	7.104	47.90	124.90
Republican governor	0.507	0.500	0.000	1.000
Percent republican vote in presidential election	0.481	0.088	0.090	0.826
Adjacent state has an EITC program	0.362	0.480	0.000	1.000

Notes: Summary statistics in this table come from state panel data over the period 1977–2004 and are weighted by the number of women ages 15 and over. Marriage and divorce rates are derived from the balanced panel of state-year combinations, which for the marriage rate is 1,372 observations and for the divorce rate is 1,204 observations. All other summary statistics are calculated for the full observation period ($N = 1{,}428$). The denominator in the marriage rate is the number of unmarried women ages 15 and over. The mean number of unmarried women for the balanced panel is 1,520,133 (unweighted mean: 780,003). The denominator in the divorce rate is the number of married women ages 15 and over. The mean number of married women for the balanced panel is 2,018,566 (unweighted mean: 1,041,834). Both figures are derived from the 1977–2004 March CPS. The EITC variables are lagged 2 years

1.4, which applies to Nevada. Many out-of-state residents come to Nevada for a wedding, raising the state's marriage rate above what it would be if marriages were counted by state of residence.[16] As will be shown, the regression results are not sensitive to excluding Nevada from the analysis. Table 2 also presents information on the combined EITC maximum credit, disaggregated by the number of children to which the credit applies. The real value of the EITC declined between 1977 and 1986, but increased substantially after the 1986, 1990, and 1993 reforms.

[16] Nevada's new marriage rate exceeded one between 1977 and 1980, but remained far above the national average after that. Summary statistics for the new marriage rate after dropping Nevada are the following: mean = 0.053; SD = 0.016; minimum = 0.014; maximum = 0.137.

Results from the new marriage and divorce regressions are presented in Tables 3 and 4. For each outcome, I discuss five specifications. Model 1 shows the coefficient on the twice-lagged EITC variable without any controls added to the specification. Model 2 is the full model, incorporating all other policy, economic, demographic, and political variables, as well as controls for unobserved heterogeneity. Models 3, 4, and 5 are specification checks that explore alternative functional forms and correct for autocorrelation. Model 3 uses as the dependent variable the log number of new marriages (divorces), and Models 4 and 5 re-estimate the full specifications using the Prais-Winsten method.

Models 1 and 2 in Table 3 consistently show that the EITC is negatively associated with entries into marriage. The estimate in the first column suggests that a $1,000 increase in the combined EITC maximum credit reduces new marriages by 9.0%. However, this model does not control for other factors that impact marriage decisions. When these controls are added in Model 2, the estimated effect of the EITC declines dramatically, although it remains statistically significant at conventional levels. The coefficient on the EITC implies that a $1,000 increase in the EITC is associated with a 4.9% decrease in new marriages.

Parameter estimates in Models 3–5 suggest that these results are not specious. Model 3 explores the robustness of the main results to changes in functional form, specifically using the log number of new marriages as the dependent variable. The coefficient on the EITC implies that a $1,000 increase in the maximum credit leads to a 6.1% decrease in the number of new marriages. This coefficient is statistically significant at the 0.01 level. Results presented in Models 4 and 5 suggest that the EITC estimates are robust to the Prais-Winsten correction for autocorrelation. Model 4 provides autocorrelation-corrected estimates using the new marriage rate as the dependent variable, and Model 5 provides autocorrelation-corrected estimates using the number of new marriages as the dependent variable. In both cases, the parameter estimate on the EITC is virtually identical to the uncorrected models and continues to be statistically significant.

Several points are raised by these results. First, it is important to note the large decline in the magnitude of the EITC effect after fixed effects and time trends are included in the model. Such declines have been found elsewhere (Bitler et al. 2004; Dickert-Conlin and Houser 1999; Fitzgerald and Ribar 2004; Hoynes 1997; Moffitt 1994), and are due to unobservable differences between geographic areas and over time that are correlated with marriage and headship decisions. Results in this study suggest that without controls for unobservable factors, the estimated effect of the EITC is biased upward. Second, the EITC appears to have an economically small impact on transitions into marriage. Using the coefficient from Model 2, I find that each one-dollar increase in the EITC maximum credit is associated with 0.02 fewer marriages among unmarried women.[17] This finding is corroborated by other studies, which estimate small or insignificant impacts of the EITC on marriage and headship

[17] The impact of a one-dollar increase in the EITC is calculated by the following: [β(unmarried)/EITC], where β is the parameter estimate from Table 3, Model 2 (divided by 1,000); *unmarried* is the number unmarried women ages 15 and over, averaged over the observation period; and *EITC* is the maximum credit for a family with two or more children, averaged over the observation period. See the notes in Table 2 for summary statistics associated with *unmarried* and *married*.

Table 3 The impact of the EITC on new marriages

Variable	Model 1	Model 2	Model 3	Model 4	Model 5
(Federal/state EITC maximum credit)$_{t-2}$	−0.090***	−0.049**	−0.061***	−0.050**	−0.073***
	(0.005)	(0.024)	(0.021)	(0.025)	(0.022)
Implementation of welfare waiver		−0.025	−0.032**	−0.016	−0.025*
		(0.016)	(0.012)	(0.016)	(0.014)
Implementation of TANF		−0.147**	−0.150**	−0.169***	−0.195***
		(0.065)	(0.061)	(0.038)	(0.034)
Maximum monthly AFDC/TANF benefit		0.166	0.074	0.162	0.060
		(0.115)	(0.088)	(0.108)	(0.092)
State unemployment rate		−1.151***	−0.670***	−0.987***	−0.434*
		(0.262)	(0.209)	(0.311)	(0.261)
Log of combined service/retail employment		0.367***	0.622***	0.344**	0.585***
		(0.115)	(0.117)	(0.146)	(0.136)
Combined service/retail earnings		−0.003	0.0001	−0.002	0.002
		(0.005)	(0.004)	(0.005)	(0.004)
Population density		0.003	−0.034*	−0.005	−0.044
		(0.020)	(0.017)	(0.033)	(0.038)
Percent black population ages 15+		2.418**	0.371	1.972*	0.569
		(1.179)	(1.107)	(1.051)	(0.848)
Percent of female population ages 15–39		0.993	1.272	1.385	1.119
		(1.022)	(0.960)	(1.136)	(1.228)
Percent of female population ages 40–64		−1.051	0.227	−0.482	0.099
		(1.025)	(0.965)	(1.169)	(1.161)
Percent ages 25+ with a BA degree		0.848**	0.394	0.778**	0.534**
		(0.395)	(0.392)	(0.324)	(0.253)
Republican governor		−0.003	−0.006	0.000	−0.001
		(0.005)	(0.004)	(0.006)	(0.004)
Percent republican vote in presidential election		0.072	0.222***	−0.025	0.101
		(0.090)	(0.067)	(0.104)	(0.089)
Log of state population ages 15+			0.406*		0.441
			(0.235)		(0.270)
Dependent variable: ln(new marriage rate)	×	×		×	
Dependent variable: ln(no. of new marriages)			×		×
Corrected for serial correlation	No	No	No	Yes	Yes
State and year fixed effects	No	Yes	Yes	Yes	Yes
State-specific time trends	No	Yes	Yes	Yes	Yes
R^2	0.164	0.962	0.995	0.998	0.999
Number of state-year combinations	1,372	1,372	1,372	1,372	1,372

Notes: Regression coefficients in this table come from the balanced panel of states between 1977 and 2004, and are weighted by the number of women ages 15 and over. Standard errors, reported in parentheses, are corrected for arbitrary heteroskedasticity using Huber-White robust standard errors. These analyses omit California and Oklahoma due missing information on new marriages for some years. The EITC variable is the combined federal/state maximum credit for families with two or more children, lagged 2 years. Estimates in column (4) and (5) come from a Prais-Winsten regression to correct for serial correlation, assumed to take an AR(1) process that is common across all states. ***, **, * Indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

Table 4 The impact of the EITC on new divorces

Variable	Model 1	Model 2	Model 3	Model 4	Model 5
(Federal/state EITC maximum credit)$_{t-2}$	−0.041***	−0.007	0.032	−0.034	0.004
	(0.005)	(0.026)	(0.026)	(0.031)	(0.031)
Implementation of welfare waiver		−0.040**	−0.055***	−0.053***	−0.064***
		(0.016)	(0.014)	(0.017)	(0.015)
Implementation of TANF		−0.093**	−0.100**	−0.120**	−0.132***
		(0.046)	(0.045)	(0.050)	(0.048)
Maximum monthly AFDC/TANF benefit		−0.118	−0.150	−0.139	−0.182
		(0.128)	(0.120)	(0.138)	(0.130)
State unemployment rate		−1.090***	−0.681***	−0.853**	−0.519
		(0.267)	(0.248)	(0.386)	(0.362)
Log of combined service/retail employment		0.441***	0.558***	0.510***	0.652***
		(0.139)	(0.138)	(0.168)	(0.179)
Combined service/retail earnings		−0.012**	−0.010	−0.012*	−0.010
		(0.006)	(0.006)	(0.006)	(0.006)
Population density		−0.084***	−0.031	−0.087*	−0.038
		(0.029)	(0.029)	(0.046)	(0.044)
Percent black population ages 15+		2.087*	−0.013	2.865**	1.100
		(1.243)	(1.315)	(1.180)	(1.193)
Percent of female population ages 15–39		1.009	1.110	0.704	0.630
		(1.096)	(1.221)	(1.548)	(1.643)
Percent of female population ages 40–64		1.272	1.782	1.108	1.347
		(1.316)	(1.278)	(1.703)	(1.661)
Percent ages 25+ with a BA degree		−0.689**	−0.433	−0.416	−0.299
		(0.348)	(0.351)	(0.361)	(0.348)
Republican governor		0.010**	0.009*	0.005	0.005
		(0.005)	(0.004)	(0.006)	(0.006)
Percent republican vote in presidential election		−0.311**	−0.155	−0.181	−0.010
		(0.131)	(0.122)	(0.131)	(0.125)
Log of state population ages 15+			0.479		0.443
			(0.308)		(0.359)
Dependent variable: ln(new divorce rate)	×	×		×	
Dependent variable: ln(no. of new divorces)			×		×
Corrected for serial correlation	No	No	No	Yes	Yes
State and year fixed effects	No	Yes	Yes	Yes	Yes
State-specific linear time trends	No	Yes	Yes	Yes	Yes
R^2	0.063	0.935	0.994	0.999	0.999
Number of state-year combinations	1,204	1,204	1,204	1,204	1,204

Notes: Regression coefficients in this table come from the balanced panel of states between 1977 and 2004, and are weighted by the number of women ages 15 and over. Standard errors, reported in parentheses, are corrected for arbitrary heteroskedasticity using Huber-White robust standard errors. These analyses omit California, Colorado, Georgia, Hawaii, Indiana, Louisiana, Nevada, and Oklahoma due missing information on new divorces for some years. The EITC variable is the combined federal/state maximum credit for families with two or more children, lagged 2 years. Estimates in column (4) and (5) come from a Prais-Winsten regression to correct for serial correlation, assumed to take an AR(1) process that is common across all states. ***, **, * Indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively decisions (Dickert-Conlin and Houser 1999, 2002; Fitzgerald and Ribar 2004). Finally, these results indicate that the independence effect dominates the stabilizing effect for recent entries into marriage. By increasing the returns to market work, the EITC improves labor market opportunities for unmarried women, leading to reductions in marriage.

Turning to the other policy variables, I find that welfare reform is negatively associated with entries into marriage. Implementation of a statewide welfare waiver is marginally significant, depending on the model, and the coefficients imply a reduction in new marriages between 1.6 and 3.2%. Implementation of TANF, on the other hand, is strongly related to new marriages, implying a reduction in new marriages between 14.7 and 19.5%.[18] These estimates are somewhat smaller than those appearing in Bitler et al. (2004); however, they provide further evidence that the independence effect dominates the stabilizing effect for transitions into marriage. Finally, increases in AFDC/TANF benefits reveal a *positive*, but largely insignificant, relationship with recent marriages. In results not shown here, the coefficient on welfare benefits is negative and statistically significant in models without fixed effects, a common finding in marriage literature (Moffitt 1998).

Table 4 presents estimates of the relationship between the EITC and entries into divorce. Model 1 shows that a $1,000 increase in the maximum credit is associated with a 4.1% decrease in the new divorce rate. When the full set of controls are added, the point-estimate declines and the standard error increases substantially, rendering the coefficient statistically insignificant. Interestingly, when the model omits the state-specific time trends (and includes only state and year fixed effects in addition to the measured characteristics), the EITC coefficient is positive and statistically significant, indicating that a $1,000 increase leads to a 7.3% increase in new divorces. Adding time trends to the model absorbs most of the variation in new divorces, which may explain why the EITC effect disappears when the trends are added.[19]

None of the robustness checks provide statistically significant EITC estimates, confirming the results in Model 2. Using the log number of new divorces as the dependent variable (Model 3) causes the sign on the EITC coefficient to become positive. Correcting the estimates for autocorrelation (Models 4 and 5) using the Prais-Winsten method does little to change the story. As in the uncorrected models,

[18] As noted by Bitler et al. (2004) and Blank (2002), the identification strategy for TANF is weaker than for the waiver variable. TANF was phased-in by all states over a relatively short period (September 1996 to January 1998, or 16 months), while welfare waivers were experimented with by a smaller number of states over 5-year period (1992–1996). Therefore, the effect of TANF is identified by substantially smaller cross-state variation in the timing of implementation. It is also important to note that both welfare reform variables roll several policy reforms into a one-dimensional variable, and that TANF, in particular, is likely to generate different impacts on marriage and divorce if one were to estimate these policy components separately.

[19] It is worth reiterating here the potential for the EITC to generate both marriage subsidies and penalties. Many of these differences depend on where individuals reside on EITC benefit schedule and the number of children included in the tax unit. Unfortunately, it is not possible using aggregate Vital Statistics data to tease out differential EITC effects across families with different observable characteristics. I merely point out these potentially offsetting effects as another explanation for why the EITC does not appear to be correlated with aggregate divorce flows.

the EITC coefficients are of conflicting signs and not statistically significant at conventional levels.

The mixed evidence on the relationship between the EITC and entry into divorce is evident in other studies. For example, Dickert-Conlin and Houser (2002) find that the EITC reduces the propensity to stay married in some models, but increases the propensity in other models. Likewise, Rosenbaum (2000) determines that increases in the tax cost of marriage are unrelated to exiting marriage in a given year. Finally, Fitzgerald and Ribar (2004) find a positive but insignificant relationship between the EITC and the probability of exiting from female headship.

Welfare reform is negatively associated with transitions into divorce. States that implemented a welfare waiver in the early-1990s experienced a 4.0% reduction in new divorces, and while TANF further reduced divorces by approximately 9.0%. The magnitude of these estimates matches closely those of Bitler et al. (2004). Furthermore, these results suggest that welfare reform may have a stabilizing effect on marriage among married women. Finally, contrary to theoretical predictions, welfare benefits are negatively related to divorce rates, although the coefficient is not statistically significant in models that include controls for unobserved heterogeneity.

Specification Checks and Additional Tests of Robustness

The results discussed above are subjected to three broad types of specification checks: further changes to the measurement and functional form of the EITC variable; alternative weights and data restrictions; and the inclusion of additional policy and demographic variables. Table 5 presents these additional analyses, using the EITC coefficient from Model 2 in Tables 3 and 4 as the baseline estimate.

Measuring the dependent variable in levels does not qualitatively change the results. The estimate for the marriage equation suggests that raising the EITC maximum credit by $1,000 is associated with a 0.004-point reduction in the new marriage rate. This estimate translates to a 7.1% (0.004/0.056) decrease in new marriages, approximating the result from the log model. The EITC coefficient in the level divorce model remains statistically insignificant. A criticism of the baseline model is that a one-period lag of the EITC is sufficient to account for the delay in EITC payments and marital decisions. Twice-lagging the EITC may also be unnecessary given that the IRS allows claimants to receive advance payments along with their regular paychecks.[20] Therefore, Table 5 presents estimates for the impact of a one-period lag in which the dependent variable is measured in log and level forms. In both cases, once-lagging the maximum credit reduces somewhat the estimated EITC effect in the marriage model, but the point estimates remain statistically significant at conventional levels. Results in the divorce model are never statistically significant.

[20] It should be noted, however, that the advance payment option does not fully warrant a change in the EITC lag structure. An analysis by the U.S. General Accounting Office (1992) finds that few potential claimants are aware of this option, and only 0.5% of EITC recipients chose to obtain their benefits via this method.

Table 5 Specification checks

Coefficient on the EITC in the following model	Dependent variable	
	ln(new marriage rate)	ln(new divorce rate)
(Federal/state EITC maximum credit)$_{t-2}$ (main estimates)	−0.049**	−0.007
	(0.024)	(0.026)
New marriage and divorce rate (no logs)	−0.004***	0.000
	(0.001)	(0.000)
(Federal/state EITC maximum credit)$_{t-1}$	−0.039*	0.010
	(0.023)	(0.023)
(Federal/state EITC maximum credit)$_{t-1}$ and new marriage/divorce rate	−0.003***	0.0001
	(0.001)	(0.0003)
Use variation in state EITC only if it is structured as a refundable tax credit	−0.092***	0.037
	(0.034)	(0.039)
Weight marriage/divorce regressions by the number of unmarried/married women	−0.056**	−0.004
	(0.024)	(0.026)
Use total female population ages 15+ as the denominator in marriage/divorce rates	−0.071***	0.023
	(0.020)	(0.024)
Exclude Nevada and Hawaii	−0.045*	−0.007
	(0.025)	(0.026)
Include additional tax and social policy controls	−0.046**	−0.014
	(0.023)	(0.026)
Include overall fertility rate (lagged one year)	−0.044*	−0.044
	(0.025)	(0.028)

Notes: The coefficient presented is the twice-lagged federal/state EITC maximum credit for a family with two or more children. Regression coefficients in this table come from the balanced panel of states for the respective new marriage and divorce outcomes (lagged 2 years) and are weighted by the total number women ages 15 and over. All models include the full set of controls in Tables 3 and 4, including state and year fixed effects and state-specific time trends. The additional tax and social policy controls referred to in the table include a dummy variable for the onset of the Child Tax Credit, a dummy variable for the implementation of a family cap, a dummy variable for the onset of the 2001 Economic Growth and Tax Relief Reconciliation Act (EGTRRA), and spending through the Child Care and Development Fund (CCDF). Standard errors, reported in parentheses, are corrected for arbitrary heteroskedasticity using Huber-White robust standard errors. ***, **, * Indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

As previously sated, states have the option of structuring their EITC programs as refundable or nonrefundable tax credits. Given that low-income families are more likely to qualify for refundable credits, it is important to distinguish between these structures in the estimation. One approach is to create an EITC variable that uses variation in the state credit only if it is refundable. Adding this variable to the basic specification nearly doubles the estimated impact of the EITC, to a 9.2% reduction in the new marriage rate. The coefficient remains insignificant in the divorce model. The larger estimate in the marriage model could reflect the loss of identifying variation in the EITC variable, but it is suggestive of a greater behavioral response to refundable EITC structures.

In addition, weighting the new marriage and divorce regressions by the number of unmarried and married women, respectively, does not affect the results. One concern with the measurement of the marriage and divorce rate is that the denominators are derived from survey data, and the population estimates are somewhat volatile for small states. To test whether the results are sensitive to these denominators, I construct alternative transition rates in which new marriages and divorces are conditioned on the total female population ages 15 and over. The estimated EITC coefficient becomes somewhat larger in this marriage model, implying a 7.1% reduction in new marriages, and the coefficient remains statistically insignificant in the divorce model. Conditioning new marriages on a larger population, which leads to smaller transition rates, could explain the larger EITC effect.

A criticism of the Vital Statistics data is that marriages and divorces are measured by the state of occurrence, rather than the state of residence. A rough way to deal with this potential problem is to drop state-years from Nevada and Hawaii. Dropping either or both of these states does not appreciably change the results. When Nevada and Hawaii are dropped from the marriage model, the magnitude of the EITC coefficient declines slightly but is still statistically significant at the 10% level. The EITC coefficient remains insignificant in the divorce model.[21]

Finally, I estimate several models that include additional tax, social policy, and demographic variables. Specifically, I estimate parameterizations of the Child Tax Credit (CTC), Economic Growth and Tax Relief Reconciliation Act of 2001 (EGTRRA), family cap policies, and child care subsidy expenditures through the Child Care and Development Fund (CCDF).[22] Adding a control for the EGTRRA is particularly important because it attempts to minimize the EITC's marriage penalties by expanding incrementally the flat and phase-out regions for married taxpayers who file jointly. Also included is the overall fertility rate, which captures state-specific preferences for children. Parameter estimates on the EITC are unchanged by the inclusion of these variables.[23]

[21] Another solution is to use Vital Statistics Detailed Data, which provide counts of marriages and divorces by state of residence, along with a number of basic demographic characteristics. However, the Detailed Data contain a number of serious flaws. First, due to budget constraints and concerns about quality, data collection was discontinued in 1995. When the system was in place, only 41 states and the District of Columbia participated in marriage data collection and 31 states and the District of Columbia participated in the divorce data collection. Only 77% of marriages and 49% of divorces were captured by the Detailed Data. Furthermore, perhaps the most important demographic variable for the purposes of the EITC, educational attainment, has not been collected since 1989.

[22] The CTC is a dummy variable that equals unity in the years after its implementation. The EGTRRA is a dummy variable that equals unity for all state-year combinations after 2001. Family caps are measured by a dummy that equals unity in the state-years after its implementation. The CCDF variable is the combined federal and state expenditures per child ages 0–12.

[23] A final specification check should be mentioned. Using unbalanced panel data reduces the magnitude of the EITC coefficient in the marriage model, rendering it imprecisely estimated ($\beta_{\text{new marriage}} = -0.035$; SE $= 0.026$ and $\beta_{\text{new divorce}} = 0.006$; SE $= 0.028$). However, estimating level measures of the dependent variable on the unbalanced panel leads to a statistically significant effect of the EITC on new marriages ($\beta_{\text{new marriage}} = -0.003$; SE $= 0.001$ and $\beta_{\text{new divorce}} = 0.0003$; SE $= 0.0004$). In neither case is the EITC coefficient significant in the new divorce model.

Further Issues: Legislative Endogeneity

It was argued earlier that legislative activity on the EITC is a plausibly exogenous source of variation with which to identify its impact on marriage and divorce flows. Confidence in this assumption is greater for federal EITC expansions, in which changes to rules and benefits do not differentially apply to states with certain underlying labor supply or marriage propensities. However, this leaves open the question of state EITC programs. States implementing a tax credit could be systematically different from those that do not in terms of economic and political conditions, as well as cultural attitudes toward work and family. It is therefore important to control for the processes that lead to their implementation to arrive at an unbiased estimate of the EITC (Besley and Case 2000).

I explore two sources of policy endogeneity. First, it is reasonable that changes to marriage behavior created by a policy shift in one state could have an impact on behavior in an adjacent state. This could occur through the transfer of social norms from reform states to nearby non-reform states, thereby producing a de-facto policy shock in non-reform states. Another possibility is through legislative diffusion, whereby non-reform states adopt EITC programs because they border others that have already done so. The cross-state diffusion of policy innovation is a well-documented phenomenon that occurs when states learn from or compete with each other to adopt policy reforms. Indeed, previous studies of policy diffusion have examined mother's pensions (Gray 1973), abortion regulation (Mooney and Lee 1995), anti-smoking laws (Shipan and Volden 2008), among others. Evidence of such diffusion with respect to the EITC is compelling given the geographic clustering of state credits. Fully 81% of states with an EITC are adjacent to another with the credit, compared to 31% of states without an EITC. A second source of policy endogeneity arises from state legislators' strategic use of EITC policymaking to respond to changes in employment and poverty or gain electoral advantage by providing tax relief to low-income individuals. If this is the case, it would be expected that future changes to states' EITC programs are correlated with current employment and marriage behavior.

Results presented in Tables 6 and 7 examine these sources of policy endogeneity. Table 6 evaluates the cross-state transfer of policy effects by including a dummy variable that equals unity for states that border one with an EITC. The coefficient on the border dummy is negative in the marriage and divorce models but is only statistically significant for marriages, a result that is broadly consistent with the social norms and legislative diffusion stories. When the EITC maximum credit is added to the marriage and divorce regressions (Model 3), the magnitude of the coefficient increases substantially in both models. As measured, however, the border state dummy does not distinguish between states with and without an EITC; it only distinguishes between states that border others with the program. Confidence in these findings would therefore increase if the negative effects are concentrated among non-EITC states that border one with the program. Models 4 and 5 include dummy variables capturing the full set of state/border state EITC combinations. The results indicate that any cross-state transfer of policy effects—through de facto policy shocks or legislative diffusion—flows into states without an EITC from their

Table 6 Cross-state transfer of EITC effects

Variable	Model 1	Model 2	Model 3	Model 4	Model 5
Dependent variable: ln(new marriage rate)					
(Federal/state EITC maximum credit)$_{t-2}$	−0.049**		−0.063**		−0.068***
	(0.024)		(0.025)		(0.025)
Adjacent state has an EITC program		−0.021**	−0.027***		
		(0.009)	(0.009)		
EITC in state/EITC in adjacent state				−0.011	−0.012
				(0.022)	(0.022)
EITC in state/no EITC in adjacent state				0.009	0.021
				(0.018)	(0.018)
No EITC in state/EITC in adjacent state				−0.020**	−0.026***
				(0.009)	(0.009)
Dependent variable: ln(new divorce rate)					
(Federal/state EITC maximum credit)$_{t-2}$	−0.007		−0.012		−0.013
	(0.026)		(0.027)		(0.026)
Adjacent state has an EITC program		−0.008	−0.009		
		(0.009)	(0.009)		
EITC in state/EITC in adjacent state				0.006	0.006
				(0.025)	(0.025)
EITC in state/EITC not in adjacent state				−0.006	−0.004
				(0.018)	(0.017)
No EITC in state/EITC in adjacent state				−0.010	−0.011
				(0.010)	(0.010)
State controls	Yes	Yes	Yes	Yes	Yes
State and year fixed effects	Yes	Yes	Yes	Yes	Yes
State-specific linear time trends	Yes	Yes	Yes	Yes	Yes
Number of state-year combinations	Marriage model: $N = 1{,}372$; Divorce model: $N = 1{,}204$				

Notes: Regression coefficients in this table come from the balanced marriage and divorce panel of states between 1977 and 2004 and are weighted by the number of women ages 15 and over. The EITC variable is the combined federal/state maximum credit for families with two or more children, lagged two years. The EITC estimates reported in Model 1 come from Table 3 (new marriages) and Table 4 (new divorces), Model 2. All models include the full set of state-level demographic and economic variables presented in Tables 3 and 4. Standard errors, reported in parentheses, are corrected for arbitrary heteroskedasticity using Huber-White robust standard errors. The three-variable combination of state and adjacent state EITC programs are dummies, with the omitted category being states without an EITC program and not adjacent to one with an EITC. ***, **, * Indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

neighbors with the credit. Taken together, these findings suggest that a failure to control for border state policy places a downward bias on the EITC effect.[24]

[24] The negative border state coefficients could be an artifact of unobserved regional factors that influence EITC policymaking and marriage norms. To investigate this possibility, I add region dummies to all models in Table 6. Doing so does not alter the sign, magnitude, or significance of the coefficients on the border state dummy and EITC maximum credit.

Table 7 Test for pre-reform differences in new marriage and divorce rates

Variable	Model 1	Model 2	Model 3
	Dependent variable: ln(new marriage rate)		
(Federal/state EITC maximum credit)$_{t-2}$			−0.050**
			(0.024)
(State has an EITC program)$_{t+1}$	0.029**	0.021	0.021
	(0.012)	(0.017)	(0.017)
	Dependent variable: ln(new divorce rate)		
(Federal/state EITC maximum credit)$_{t-2}$			−0.008
			(0.026)
(State has an EITC program)$_{t+1}$	0.022	0.009	0.009
	(0.016)	(0.018)	(0.018)
State controls	Yes	Yes	Yes
State and year fixed effects	Yes	Yes	Yes
State-specific linear time trends	No	Yes	Yes
Number of state-year combinations	Marriage model: N = 1,372; Divorce model: N = 1,204		

Notes: Regression coefficients in this table come from the balanced marriage and divorce panel of states between 1977 and 2004 and are weighted by the number of women ages 15 and over. The EITC variable is the combined federal/state maximum credit for families with two or more children, lagged 2 years. Some models include a one-period-lead dummy variable that equals unity if the state has an EITC program. All models contain the full set of state-level demographic and economic variables presented in Tables 3 and 4. Standard errors, reported in parentheses, are corrected for arbitrary heteroskedasticity using Huber-White robust standard errors. ***, **, * Indicate statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

Results in Table 7 attempt to control for policymakers' strategic use of the EITC. Specifically, I include a one-period lead of a state EITC dummy variable in the marriage and divorce regressions. Model 1 includes state and year fixed effects, and Model 2 adds state-specific time trends. Although the coefficient on the lead is statistically significant in the fixed effects marriage regression, adding time trends renders it insignificant. The lead is never significant in the divorce regressions. In addition, the coefficient on the maximum EITC is unaffected by the inclusion of the lead variable. Thus, it appears that states with an EITC are systematically different from those without the program, but most of these differences are handled by controls for unobserved heterogeneity.

Conclusion and Discussion of Policy Implications

President Ronald Reagan once lauded the EITC as "best antipoverty, the best pro-family, the best job-creation measure to come out of Congress." A sizeable research literature has examined the anti-poverty and labor supply effects of the credit. On both accounts, the EITC is an effective policy tool. However, significantly less is known about the marriage distortions created by the program. The EITC's design could either reward or penalize marriage depending on the partner's combined earnings, number of children, and state of residence. Economic theory suggests that the labor supply effects and fertility distortions introduced by the EITC could further influence marriage and divorce decisions.

Constructing a flow measure of marriage and divorce rates from national Vital Statistics over the period 1977 to 2004, I find that increases in the EITC are associated with declines in new marriage. The preferred model suggests that a $1,000 increase in the maximum credit decreases the new marriage rate by 4.9%, although the estimates range from 3.5 to 9.2% depending on the specification. Furthermore, I find no evidence of a relationship between the EITC and transitions into divorce.

To put these findings in perspective, it is important to consider the larger policy environment in which the EITC operates. Throughout the 1990s, the program received two major expansions that occurred roughly contemporaneously with the implementation of welfare reform, the expansion of child care subsidies, and changes to Medicaid. Each of these policy changes is expected to raise the work effort of single women with children, and may therefore alter marriage behavior as well, either through an "independence" or "stabilizing" effect. Results in this study indicate that the EITC and welfare reform promote an "independence" effect.

Marriage non-neutrality in the EITC raises a number of concerns. First, by treating couples differently before versus after marriage, the EITC violates the notion of horizontal equity. Second, when people perceive to be economically worse off because of the tax system, illicit behaviors and tax avoidance can increase.[25] In addition, low-income families that already experience marriage penalties from the transfer system are further burdened by those of the tax system. With state EITC programs on the rise, such burdens warrant attention. Finally, considerable policy debate focuses on the role of social policy in increasing out-of-wedlock births. Although the precise influence of the tax and transfer system on family structure is controversial, the rise in female-headed families raises concerns over the healthy development of children.

In recent years, policymakers have sought to reduce the marriage penalties embedded in the EITC, and some of these proposals were enacted in the 2001 EGTRRA. Specifically, for married couples filing joint returns, the beginning and ending points of the phase-out region were extended by $1,000 between 2002 and 2004, $2,000 between 2005 and 2007, and $3,000 beginning in 2008 (U.S. House of Representatives 2004). Results in this study suggest that the behavioral effects of the 2001 EITC changes are likely to be small. These are tentative conclusions, however, and future research in this area should attempt to isolate the behavioral and fiscal impact of the 2001 EITC changes on marriage and divorce.

Appendix

[25] EITC noncompliance has been a topic of substantial interest. A study of the 1994 tax year found that over 20% of EITC payments were made erroneously, most of them in overclaims, and one-third of these overclaims were due to misreporting of filing status by married partners (McCubbin 2000; Scholz 1994, 1997). Moreover, a recent analysis of 1999 tax returns found that between $8.5 and $9.9 billion of EITC payments (or 27.0 to 31.7%) were paid in error (U.S. Department of the Treasury 2002). Although the report cited qualifying child requirements as the most important factor behind the erroneous payments, married taxpayers filing as single (or head of household) was another key factor.

Variable	Source
State-level numbers of new marriages and divorces	National Center for Health Statistics: *Vital Statistics of the United States* and *Monthly Vital Statistics Report*. Available at: http://www.cdc.gov/nchs/mardiv.htm
Population estimates for married and unmarried women ages 15 and over	March Current Population Survey, 1977–2004
Federal/state EITC maximum credit	U.S. Congress, Committee on Ways and Means: *2004 Green Book*; Dicket-Conlin and Houser (2002); Center on Budget and Policy Priorities: *A Hand Up: A How State Earned Income Tax Credits Help Working Families Escape Poverty*, various years
Welfare waiver and TANF implementation dates	Crouse (1999), DHHS (1997), and U.S. GAO (1997)
Maximum monthly AFDC/TANF benefits, by family size	U.S. Congress, Committee on Ways and Means: *2004 Green Book*. Available at: http://www.gpoaccess.gov/wmprints/green/index.html
State unemployment rate	U.S. Department of Labor, Bureau of Labor Statistics: Local Area Employment Statistics. Available at: http://www.bls.gov/lau/home.htm
Combined service/retail industry employment and earnings	U.S. Department of Commerce, Bureau of Economic Analysis: Regional Economic Information System 1969–2004. Available at: http://www.bea.gov/
Population density	*Statistical Abstracts of the United States*, various years Available at: http://www.census.gov/compendia/statab/. Available for the census years only; intercensal years are interpolated
Black and female population estimates	U.S. Census Bureau. Available at: http://www.census.gov/popest/estimates.php
Percent of population with a BA degree	*Statistical Abstracts of the United States*, various years. Available at: http://www.census.gov/compendia/statab/
Fertility rate	National Center for Health Statistics: *Vital Statistics of the United States* and *Monthly Vital Statistics Report*. Available at: http://www.cdc.gov/nchs/mardiv.htm; *Statistical Abstracts of the United States*, various years. Available at: http://www.census.gov/compendia/statab/
Republican governor	*Statistical Abstracts of the United States*, various years. Available at: http://www.census.gov/compendia/statab/
Republican vote in Presidential elections	*Statistical Abstracts of the United States*, various years. Available at: http://www.census.gov/compendia/statab/

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