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The Earned Income Tax Credit and abortion

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ABSTRACT

Using a panel of states between 1975 and 2005, I examine the relationship between the Earned Income Tax Credit (EITC) and abortion. Results suggest that increases in the EITC are associated with reductions in the overall abortion rate. Specifically, a $1000 increase in the maximum credit is associated with a 7.6% decrease in the overall abortion rate. This effect implies a reduction of approximately 1.6 abortions per 1000 women of childbearing age. I also provide evidence that the reduction in abortions is attributable to a drop in pregnancies, not an increase in the number of births. In other words, women appear to respond to increases in EITC generosity by altering sexual behavior, rather than shifting abortion–birth decisions after a pregnancy occurs.

1. Introduction

After peaking in 1980—7 years after the US Supreme Court decision in *Roe v. Wade*—the abortion rate declined substantially. By 2005, there were about 19 abortions per 1000 women of childbearing age, a 34% decrease from the early-1980s. Indeed, the abortion rate is currently at its lowest level since the Court's decision in 1973 (Finer and Henshaw, 2006). A sizable body of research has advanced a number of explanations for this decline, including changes in demographic trends and economic conditions and the enactment of public policies aimed at restricting women's access to abortion (Bitler and Zavodny, 2001; Blank et al., 1996; Kane and Staiger, 1996; Levine, 2003; Matthews et al., 1997).

Means-tested transfers can also influence women's fertility decisions because many of these programs target benefits at women with children. Considerable research focuses on the cash assistance program Aid to Families with Dependent Children/Temporary Assistance to Needy Families (AFDC/TANF), which not only conditions eligibility on having children but also provides families with greater benefits as the number of children increases (Moffitt, 1998).[1] Furthermore, the Earned Income Tax Credit (EITC), a wage supplement for low-income workers, potentially distorts fertility decisions for many of the same reasons. With annual expenditures and recipients that exceed the AFDC/TANF program, it is surprising that researchers have neglected a systematic examination of the relationship between the EITC and abortion.[2] The goal of this paper is to provide such an analysis.

There are several reasons to believe that the EITC may influence the decision of women to obtain an abortion. The first set of considerations deals with the program's design features. With the exception of a small credit for childless tax filers, the EITC is largely directed at families with children. This feature implies that the EITC lowers the relative cost of childbearing, leading to a possible reduction in the number of abortions and an increase in the number of births. Another design feature emphasizes variation in benefit generosity according to the number of children in the family. Although EITC benefits did not

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[1] In response, a number of states implemented a family cap policy throughout the 1990s. This policy ends the incremental increase in AFDC/TANF benefits for each additional child born into a family. As of 2005, 22 states implemented a family cap.

[2] As of 2005, federal expenditures through the EITC totaled $38.8 billion, while spending on AFDC/TANF totaled $24.8 billion. Approximately 22.8 million families claimed the credit in 2005. The comparable figure for the AFDC/TANF program is two million families (in 2003) (Urban-Brookings Tax Policy Center, 2008).

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initially vary with the number of children, a series of reforms enacted in 1990 and 1993 created a substantially greater maximum credit for families with two or more children compared to those with a single child. Therefore, prior to the early-1990s legislation, the EITC provided an unambiguous incentive to reduce abortions among childless women. Introduction of the differential benefit for families with two or more children, however, may have encouraged further reductions in abortions among one-child families.

A second set of considerations deals with the *timing* of the EITC's influence in the fertility decision-making process. In particular, the introduction of an EITC may alter abortion rates first through changes in sexual activity and contraceptive intensity, both of which ultimately affect pregnancy rates. As discussed in the next section, there are theoretical reasons to believe that the EITC can increase pregnancy rates within some groups of women while decreasing pregnancies within other groups. A key issue here is determining how those EITC-induced pregnancies are ultimately resolved by women. The EITC may also influence abortion rates by altering women's decision-making after a pregnancy occurs. However, theoretical models provide similarly ambiguous predictions for whether the policy will lead to an increase in the number of births or an increase in the number of abortions.

A final set of issues focuses on the complex employment incentives created by the EITC. If a family's earnings fall within the phase-in region of the benefit schedule, where the amount of the credit is rising with income, the policy creates strong employment incentives that may increase the opportunity costs associated with children. This implies that abortions may rise because of the EITC. For families located in the plateau region, in which the credit amount no longer increases with earnings, the ambiguous labor supply incentives suggest that only an income effect is relevant to decision-making. Assuming that children are viewed as normal goods, the income effect implies a reduction in the number of abortions. Finally, families located in the phase-out region experience employment disincentives because EITC benefits decline with earnings. In this region, the substitution effects (reduction in work) and income effects (increased income from the EITC) both imply reductions in the abortion rate. Together, these complicated employment incentives, along with the design features discussed above, suggest that the EITC has theoretically ambiguous effects on abortion decisions. Therefore, empirical analysis is necessary.

In this paper, I draw upon state-level panel data between 1975 and 2005 to examine the association between the EITC and abortion. A key advantage of the panel data approach is that it allows one to account for unobserved and difficult-to-measure determinants of state-level abortion rates that may otherwise bias the impact of the EITC. The analysis begins by using information from the Centers for Disease Control (CDC) to calculate the overall abortion rate, which is used in regressions on the combined federal and state EITC maximum credit. The key result emerging from this analysis is that increases in the EITC are associated with reductions in state-level abortion rates. Specifically, a $1000 increase in the maximum credit is associated with a 7.6% decrease in the overall abortion rate. This effect implies a reduction of approximately 1.6 abortions per 1000 women of childbearing age. I then turn to an exploration of the mechanism through which the EITC influences abortion rates. In particular, I construct state-level birth and pregnancy rates to determine whether the reduction in abortions is due to increases in the number of births or reductions in the number of pregnancies. I find that the EITC is associated with declines in birth rates *and* pregnancy rates. This provides tentative evidence that the credit reduces abortions through changes in sexual and contraceptive behavior, rather than changes in the birth–abortion decision once a pregnancy occurs.

The remainder of the paper proceeds as follows. Section 2 provides an overview of the EITC and develops a simple theoretical framework to guide the empirical analysis. Section 3 summarizes what is currently known about the influence of public policies on women's fertility decisions. The data and estimation approach are introduced in Section 4. Section 5 discusses the results, and Section 6 concludes.

2. The Earned Income Tax Credit: structure and incentive effects

2.1. Historical development, current structure, and program awareness

The 1975 Tax Reduction Act established an earned income credit for tax filers with children. It was initially conceptualized to achieve three goals: act as a "work bonus" for the working poor, offset some of the growth in payroll taxes, and respond to the 1974 recession by stimulating demand. The credit provided a 10% wage subsidy on earnings up to $4000 (for a maximum credit of $400), which was then phased out at a rate of 10% until earnings reached $8000. The EITC was structured as a refundable tax credit: if EITC benefits exceeded tax liabilities, families received a check for the difference from the Internal Revenue Service (IRS).

Major legislation on the EITC slowed for several years until the Tax Reform Act of 1986. This law raised the subsidy rate to 14% and increased the maximum credit to $851. The EITC received its second major expansion through the Omnibus Budget Reconciliation Act of 1990. The defining feature of this legislation was the creation of separate benefit schedules for one- and multiple-child families. Those with one child could receive a 16.7% wage subsidy (for a maximum credit of $1192), while families with two or more children could receive a 17.3% subsidy (for a maximum credit of $1235). A third expansion occurred through the Omnibus Budget Reconciliation Act of 1993. This legislation created another benefit schedule for childless tax filers, comprising a 7.65% subsidy over the first $4000 of earnings and a maximum credit of $306. Phase-in rates for families with children also increased substantially. Between 1994 and 1996, the phase-in rate for families with one child increased from 23.6% to 34%, while the rate for multiple-child families grew from 30% to 40%. By 1996, when the legislation was fully implemented, eligible families with children could receive maximum EITCs of $2152 (for one child) or $3556 (for multiple children).

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Another important development is the proliferation of state EITCs. These programs "piggyback" onto the federal EITC by using its eligibility rules and calculating credits as a percentage of its 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 prior to 2005, 11 made the program refundable for the entire period of operation, while two states (Illinois and Maryland) changed from nonrefundable to refundable tax credits. Annual foregone revenue from state EITCs ranges from $17 million in Vermont to $591 million in New York (Nagle and Johnson, 2006).

Eligibility for the EITC is currently determined along several dimensions. Taxpayers must have non-zero earned income from wages or salary, business self-employment, or farm self-employment. In addition, single and married tax filers are eligible to claim the credit as long as adjusted gross income is below some threshold. This break-even point varies by year and the presence and number of children.[3] The EITC's original design extended eligibility only to families with children, but benefits did not vary with the number of children. However, as previously mentioned, legislative activity in 1993 created a small EITC for childless tax filers, and the 1990 legislation created a large and growing differential benefit for families with two or more children. With the exception of Wisconsin, all states utilize the federal credit's basic framework for the treatment of children. Wisconsin is the only state to include an incremental benefit for the third child. In particular, it supplements the federal credit by 4% for families with one child, 14% for families with two children, and 43% for families with three or more children.

A necessary condition for the EITC to alter fertility decisions is widespread awareness of and participation in the program. Scattered evidence indicates that this condition is likely to hold. Nearly two-thirds of parents with income below 200% of the federal poverty line (FPL) and over half of parents below 50% of the FPL are aware of the EITC (Phillips, 2001). Although much of this awareness comes from the growing number of public education campaigns launched by government agencies and local non-profits, low-income families increasingly use professional tax preparers to assist with claiming the credit. Indeed, ethnographic research by Romich and Weisner (2000) find that three-quarters of the families in their sample use either commercial or non-profit tax preparation services. Currently, about 62% of all EITC returns are filed with the help with a paid preparer, compared to 53% among non-EITC filers (IRS, 2000). Such broad awareness has translated into comparatively high participation rates. The best evidence to date comes from Scholz (1994), who estimates take-up rates between 80% and 86% as of the early-1990s, numbers that far exceed participation in AFDC/TANF (about 50%).

2.2. Theoretical framework

Standard economic models begin with the observation that fertility decisions are made in a series of steps, beginning with the decision about whether to have sex. Decisions are then made with respect to contraceptive intensity. If a pregnancy occurs, women must decide between aborting the pregnancy and giving birth. Two key assumptions underlie the standard model. Women are assumed to act with perfect information throughout this decision-making process, and they maintain perfect control over fertility outcomes. Based on these considerations, the model predicts that decreasing the costs associated with bearing and raising children reduces the likelihood that a pregnancy will end with an abortion. To the extent that the EITC supplements the wages of low-income families, this prediction suggests that the tax credit may lead to a decline in the abortion rate.

Some of these (relatively strong) assumptions can be relaxed by extending the model to account for the possibility that the EITC can influence abortion behavior at two points in the fertility decision-making process (Akerlof et al., 1996; Kane and Staiger, 1996; Levine, 2004).[4] First, the EITC can influence behavior when decisions are made about the level of sexual activity and contraceptive intensity. As others note, the EITC acts as insurance against income shocks by smoothing consumption when other resources are depleted (Kniesner and Ziliak, 2002). Women might therefore respond to this insurance by adopting behaviors that increase the risk of pregnancy. The central issue is determining how EITC-induced pregnancies would be resolved. For some women, the decrease in the cost of bearing and raising children would encourage a shift toward greater births. Once pregnant, other women might become increasingly aware of the logistical, financial, and emotional supports and challenges associated with obtaining an abortion (Levine, 2004). For example, some women may learn that abortion providers are highly accessible in the state of residence (leading to an increase in the number of abortions), while others could receive signals that the father is willing to support the child (leading to a reduction in the number of abortions). Overall, by encouraging women to become pregnant, the EITC can *increase* the number of abortions, although the net effect depends on how these additional pregnancies are treated.

Second, the EITC can alter abortion rates by changing women's decision-making after a pregnancy occurs. Since the credit acts as a wage supplement, an income effect is created that increases the demand for children, assuming that parents view them as normal goods. Therefore, the EITC is expected to reduce the number of abortions and increase the number of births. Also at work are price effects that operate through the EITC's differential benefit for families with different numbers of children. Prior to 1990, benefits did not vary with the number of children, and so the EITC created an incentive to reduce abortions only among childless women. The 1990 introduction of a separate benefit schedule for multiple-child families

[3] The qualifying child must be a child, grandchild, stepchild, or foster child of the taxpayer, under 19 years old (under 24 if a full-time student), or permanently disabled, and lives with the taxpayer for the entire tax year.

[4] These extensions come as a result of relaxing the assumption that women act with perfect information. Specifically, it is important to allow for the possibility that new information might be gained at several points in the decision-making process. For example, new information can arrive in the form of financial and emotional support from the father and other family members, as well as social attitudes toward non-marital childbearing.

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created an additional price incentive to reduce abortions among women with one child. Women with two or more children do not experience a positive price incentive from the current EITC structure.

This simple model is complicated by the labor supply incentives introduced by the EITC. In particular, these incentives have implications for fertility decisions that differ according to whether a tax unit's earnings are in the phase-in, plateau, or phase-out region of the EITC. If earnings fall in the phase-in region, the credit acts as a wage supplement, thereby increasing the opportunity costs associated with bearing and raising children. On the other hand, the additional income from the EITC might encourage childbearing assuming that children are normal goods. Overall, the EITC creates ambiguous fertility incentives for women in the phase-in region. If earnings fall in the plateau region, where the credit operates like a lump sum transfer, the EITC creates ambiguous labor supply incentives. However, families experience an income effect that encourages a reduction in the number of abortions and an increase in the number of births. Finally, the EITC's implicit tax on earnings is predicted to reduce employment among women in the phase-out region. As a result, the substitution and income effects are expected to reduce the number of abortions. In sum, regions of the EITC that create positive employment incentives create conflicting fertility incentives. Alternatively, regions of the credit that create ambiguous or negative employment incentives generally encourage childbearing through a pure income effect.

3. Literature review

Several strands of the empirical literature on births and abortions are relevant to the current study. In particular, a large body of research examines the impact of other means-tested programs, including AFDC/TANF benefits and recent reforms to the welfare system. In addition, a number of studies explore relationship between various tax policies (including the EITC) and fertility decisions. The discussion herein provides only a brief summary of each strand.

A review of research on AFDC/TANF benefits by Moffitt (1998) concludes that welfare generosity is likely to have a small positive and statistically significant effect on fertility. More recent work in this area focuses on the impact of welfare reform. There appears to be a convergence of evidence suggesting that welfare reform reduced fertility, particularly among teenagers. For example, Lopoo and DeLeire (2006) find that PRWORA's minor parent provisions led to a relative decline in fertility rates of 0.7% points among 15–17-year-olds. This result is corroborated by Kaestner et al. (2003) and Offner (2003), both of which find that the implementation of PRWORA led to small but statistically significant reductions in teenage childbearing.

Also relevant to the current study are evaluations of family cap policies, which deny families incremental increases in welfare benefits as the number of children increases. Studies of family caps using experimental designs have been conducted for two states. Turturro et al. (1997) finds no impact of family caps on births in Arkansas, and Camasso et al. (1999) finds that New Jersey's family cap reduced the number of pregnancies and births and increased the number of abortions. A number of studies based on observational data have been conducted as well, also producing mixed results. Horvath-Rose and Peters (2001) find that family caps are associated with lower fertility rates, while Joyce et al. (2004) detect no effects. Interestingly, recent work by Levine (2002) and Kearney (2004) find small (but largely insignificant) positive effects of family caps on births.

The final strand of empirical research focuses on the impact of the income tax code. Whittington et al. (1990) use aggregate time series data between 1913 and 1984 to examine the influence of the personal exemption. These authors determine that increases in the personal exemption are positively related to childbearing, a result confirmed Whittington (1992), who uses microdata between 1979 and 1983 to examine married couples. More recent work by Dickert-Conlin and Chandra (1999) investigates the role of taxes in the timing of fertility. Using a sample of children born between late-December and early-January, the authors find that a $500 increase in the tax benefit associated with a December birth increases the likelihood of such a birth by 17% points.

Only two studies examine the impact of the EITC on births. The first, by Duchovny (2001), uses a difference-in-differences approach to evaluate the 1990 tax legislation that introduced an incremental increase in EITC benefits for multiple-child families. Using as the comparison group families with two or more children, the author finds that the tax reform increased fertility only among married white women. The second study, by Baughman and Dickert-Conlin (2007), draws upon aggregate natality data between 1990 and 1999 from the National Center for Health Statistics. These authors estimate birth rate models that include the EITC maximum benefit and the incremental benefit. Interestingly, results suggest that increases in both are associated with *reductions* in first and higher-order births, especially among white women.

4. Empirical approach

4.1. Primary data source: CDC counts of state-level abortion occurrences

Abortion data for this research are drawn from the CDC's annual publication *Abortion Surveillance*.[5] A data collection initiative that began in 1969, the CDC receives on a voluntary basis state-level abortion counts from public health agencies, hospitals, and other medical facilities. Many of the public health agencies obtain information directly from abortion providers. The CDC abortion counts reflect only legally-induced abortions, defined as those performed by a licensed physician or an individual

[5] Detailed information on these reports, as well as the abortion data themselves, can be found at http://www.cdc.gov/reproductivehealth/Data_Stats/Abortion.htm.

acting under the supervision of a physician. I use this information to construct a state-by-year dataset of abortions between 1975 and 2005.[6] The CDC abortion counts are available for all state-year combinations except in Alaska, California, Louisiana, New Hampshire, Oklahoma, and West Virginia.[7] These missing data reduce the number of state-year combinations by 26, leaving a total of 1555 observations in the analysis.

These data contain a number of noteworthy characteristics. Perhaps the most important is that, until recently, states' public health agencies reported data according to the state in which abortions were performed, rather than the state in which women resided. This artificially increases the abortion rate in states with large numbers of out-of-state residents crossing the border to obtain an abortion.[8] States vary considerably in the degree to which non-residents use in-state abortion services. As of 2005, for example, about half of abortions in the District of Columbia were performed on out-of-state residents, whereas none of the abortions in Wyoming were performed on such residents. Another issue deals with the well-known undercount of CDC's abortion occurrences. Analyses by Blank et al. (1996) find that CDC counts are consistently below those of the Alan Guttmacher Institute (AGI), another key source of aggregate abortion data in the US.[9] Indeed, I find that in the period covered in the current analysis, annual AGI abortion counts exceed the CDC counts by approximately 22%, on average.

4.2. Empirical model

Using state-level data on abortion occurrences between 1975 and 2005, I estimate permutations of the following OLS regression model:

$$y_{st} = E_{st}\beta + \mathbf{P}'_{st}\psi + \mathbf{D}_{st}\gamma + \boldsymbol{\mu}_s + \boldsymbol{\upsilon}_t + \varepsilon_{st}, \tag{1}$$

for $s = 1, \ldots, S$; $t = 1, \ldots, N$, where s indexes states and t indexes years. The dependent variable, y, represents one of several measures of aggregate abortion occurrences. I rely primarily on the abortion rate, defined as the number of abortions within each state-year cell divided by the number of women ages 15–44 (and multiplied by 1000). This variable should be interpreted as the number of abortions per 1000 women of childbearing age. In specification checks, I estimate models in which the dependent variable is the number of abortions, adding the state's female population ages 15–44 as a control. The spread of abortions across states and over time is not normally distributed, so all outcome variables are expressed in logarithmic form.[10]

The key right-hand-side variable in (1) is denoted by E, the combined federal and state EITC maximum credit.[11] Although E is defined as the contemporaneous EITC, I experiment with models that include a one-period lag of the credit. This lag structure accounts for the gap that exists between the time fertility decisions are influenced by the EITC and when those decisions manifest in the data. By parameterizing E as the maximum credit, I capture the income effect associated with the introduction and subsequent expansion of the EITC.

I rely on several sources of geographic (cross-state) and temporal (over time) variation to identify the impact of the EITC. The study period includes three major federal 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 increased from 18.5% to 34% for families with one child and from 19.5% to 40% for families with multiple children. Furthermore, the 1990 reform created the incremental benefit for families with two or more children. The federal credit was adjusted for inflation in 1986, approximately 10 years after its implementation, providing another source of policy variation during the observation period. There is also substantial cross-state variation in EITC benefits. As of 2005, 16 states (and the District of Columbia) operated their own EITC program, thus becoming an important mechanism for offsetting state income taxes.[12] Previous studies rely on this

[6] Between 1970 and 1974, only eight to 37 state public health agencies reported abortion information, although many hospitals contributed to the data collection. Beginning in 1975, 41 agencies reported this information, reaching a peak of 49 agencies during the period 2000 to 2005. Given data quality concerns in the years immediately after *Roe v. Wade*, I follow other researchers by omitting 1973 and 1974 from the analysis. Therefore, the first year of my observation period coincides with the introduction of the EITC.

[7] Abortion data are missing for Alaska between 1998 and 2002, for California between 1998 and 2005, for Louisiana in 2005, for New Hampshire between 1998 and 2005, for Oklahoma during 1998 and 1999, and for West Virginia during 2003 and 2004.

[8] This also creates an inconsistency in the abortion rate measure: whereas as the numerator commingles abortions to in-state and out-of-state residents, the denominator includes only in-state residents of childbearing age (ages 15-44) (Blank et al., 1996).

[9] The AGI collects abortion data in a different manner. Specifically, it conducts periodic surveys of abortion providers, and is known to maintain an extensive list of such providers. Although the AGI is widely regarded as providing the most accurate abortion data, it too is not without its drawbacks. First, unlike the CDC reports, the AGI survey is not conducted annually, and has become less frequent over time. During the period 1975–2005, AGI abortion data are only available for the following years: 1975–1982, 1984–1985, 1987–1988, 1991–1992, 1995–1996, 2000, and 2005. The increased data gaps after 1990 are especially problematic, given that the EITC experienced its largest expansions during this time. In addition, analyses by Michael (1999) and Joyce and Kaestner (1996) reveal somewhat contradictory but nevertheless problematic characteristics of AGI data. Michael (1999) finds that AGI may in fact overstate the number of abortions, while Joyce and Kaestner (1996) argue that overall reporting of AGI abortion counts has declined over time. Finally, AGI data do not disaggregate abortion counts by the number of previous live births or demographic sub-group.

[10] For 2000 and 2001, Wyoming reports zero abortions among women with one previous live birth. To construct the logged abortion rate for this group, I omit these two observations from the analysis. I also experimented with imputing a value of one abortion for these state-year cells, and the results are not sensitive to this imputation.

[11] I draw from several sources to construct the EITC variables. Federal EITC parameters are from the *Green Book* (US House of Representatives, 2004) and various publications from the Center on Budget and Policy Priorities. The state parameters are also taken from the Center on Budget and Policy Priorities, as well as Fang and Keane (2004) and Dickert-Conlin & Houser (2002).

[12] In 1999, Colorado implemented an EITC, but it was suspended in 2002 because of insufficient funds.

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federal and state variation to identify the credit's impact on employment (Grogger, 2003; Herbst, 2008), welfare participation (Fang and Keane, 2004; Looney, 2005) and fertility (Baughman and Dickert-Conlin, 2007). The key assumption in using the combined federal and state maximum credit is that all legislative changes are exogenous to fertility decisions. This assumption is plausible, especially for the federal component of the EITC, since tax units in all states are exposed to the same eligibility rules and benefit schedules.[13]

Eq. (1) incorporates a large vector of additional policy controls ($\mathbf{P}'$). Specifically, I include three variables measuring restrictions on abortion access: federal and state bans on the use of Medicaid funds to cover abortions, parental notification or consent laws for minors seeking an abortion, and mandatory waiting periods before an abortion can be performed.[14] I follow previous researchers by coding each abortion restriction according to the fraction of the year it was in effect (Bitler and Zavodny, 2001; Blank et al., 1996). This is particularly important for the Medicaid funding restriction, given that its implementation was delayed and enjoined in the years following the 1976 Congressional ban on the use of federal funds to pay for abortions (Hyde Amendment).[15] To account for the possibility that abortions in a given state are influenced not only by its own restriction policies but also those in nearby states, I incorporate dummy variables that equal unity if a border state enacts Medicaid funding, parental involvement, and mandatory delay laws. Inclusion of these variables also helps to mitigate the bias that may result from using abortion data based on state of occurrence.

Also included in $\mathbf{P}'$ is a number of social policy controls. I account for the flurry of welfare reform legislation implemented throughout the 1990s by including dummy variables for welfare waivers and the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) of 1996. Only statewide welfare waivers for work requirements, benefit termination time limits, work requirement time limits, and sanctions are considered in this analysis. I code both variables as equal to the fraction of the year in which the policy is "turned on." I also control for the generosity of states' welfare benefits through the maximum monthly AFDC/TANF benefit paid to non-working recipients (in a three-person family). Finally, I incorporate a variable capturing the fraction of the year in which a family cap is implemented through welfare waivers or PRWORA. Inclusion of these social policy controls is important because they alter the costs and benefits of participating in a cash assistance program (AFDC/TANF) that conditions eligibility on having one or more children.[16]

The final set of observables included in (1) focuses on states' economic, political, and demographic environments ($\mathbf{D}'$). Changes in fertility behavior due to economic conditions are controlled for by the average, annual state unemployment rate. Differences over time and across states in the political and policy-making environments are accounted for through a dummy variable that equals unity for state-years in which there is a Republican governor. Also included is a continuous variable measuring the fraction of each state's population voting Republican in the previous Presidential election. Finally, I incorporate several demographic and human capital characteristics that may shift preferences for bearing and raising children. These variables include population density, the overall marriage rate, the fraction of the population ages 25 and over with at least a bachelor's degree, and indicators of the female age structure.

As previously stated, an advantage of the panel data used in this study is the ability to control for many unobserved attributes of states, which may simultaneously influence EITC policymaking and abortion decisions. One concern is that states implementing a tax credit could be systematically different from those that do not in terms of political and cultural attitudes toward abortion. If the processes leading to a state EITC are not properly accounted for, coefficients on the EITC will be biased. Therefore, state fixed effects, μ_s, are added to control for unobserved time-invariant differences between states, including differences in abortion attitudes and preferences for certain types of social policies. I also add year dummy variables, υ_t, to filter out year-to-year changes in fertility behavior that are common to all states, driven for example by nationwide economic shocks or downturns, Supreme Court decisions related to abortion policy, and federal social policy reforms (e.g., welfare reform and the child care tax credit). Finally, I experiment with models that add state-specific time trends, which control for unobserved factors that trend linearly within states over the observation period. All regressions are

[13] 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 abortion propensities. For example, if states that enact EITC programs to reduce poverty or strengthen work incentives have populations with different underlying views on abortion, then EITC policies will be correlated with these unobserved preferences and the estimated effect of the credit on abortion will be biased. However, as noted in the text, this is not likely to be a major problem in the current study. Given that most of the variation in the EITC comes from federal policy changes—and hence all individuals, regardless of the state of residence and attitudes toward abortion, are exposed to the same benefit rules—it is not likely that EITC policymaking is related to state-specific policy environments or cultural shifts related to abortion. Nevertheless, if state legislators use the EITC to respond to changes in employment and poverty or gain electoral advantage by providing tax relief to low-income individuals, then one might expect future changes in the EITC to be correlated with current abortion rates. Therefore, I conduct specification checks that include a one-year lead of a state EITC dummy variable in the abortion regression. In no case do I find that future state policy changes are related to current abortion rates.

[14] Data for these variables are drawn from a number of sources, including Levine (2002; 2003; 2004), Merz et al. (1995), Sollom (various years), and the National Abortion Rights Action League (NARAL) annual publication *Who Decides?*. I am also grateful to Rebecca London for generously sharing the abortion restriction codes used in Blank et al. (1996).

[15] Note that my coding does not make a distinction between periods during which abortion restrictions are present but nevertheless enjoined and periods during which such laws do not exist. In this sense, my coding is similar to that of Joyce and Kaestner (2001), and Levine (2002; 2003). A number of studies create separate variables for enjoined abortion restrictions. However, they apply to a very small number of state-year combinations and either have not been found to influence abortion decisions or do so in a way that conflicts with theoretical predictions (Bitler and Zavodny, 2001; Blank et al., 1996).

[16] Data on the timing and nature welfare waiver and TANF reforms are drawn from Crouse (1999), US DHHS (1997), and US GAO (1997). Maximum welfare benefits are drawn from the *Green Book* (US House of Representatives, 2004) and the Urban Institute's *Welfare Rules Databook* (Rowe and Versteeg, 2005). To code states' family cap policies, I use the information provided in Kearney (2004) and the *Welfare Rules Databook* (Rowe and Versteeg, 2005).

Table 1
Variable definitions and summary statistics.

Variable	Definition	Mean	*SD*
Abortion rate	Number of abortions per 1000 women ages 15–44, based on CDC data	20.94	11.05
Maximum EITC	Combined maximum federal/state EITC for families with two or more children ($1000)	1.967	1.633
Federal EITC	Federal maximum EITC for families with two or more children ($1000)	1.900	1.538
State EITC	State EITC, based on state-specific and federal credit rates ($1000)	0.595	0.384
Medicaid	=1 if a state implements a restriction on using Medicaid funds to pay for abortions	0.556	0.489
Involve	=1 if a state implements a parental notification or consent law for minors seeking an abortion	0.316	0.465
Delay	=1 if a state implements a mandatory delay law	0.123	0.328
Medicaid border	=1 if a border state implements a restriction on using Medicaid funds to pay for abortions	0.871	0.335
Involve border	=1 if a border state implements a parental notification or consent law for minors seeking an abortion	0.692	0.461
Delay border	=1 if a border state implements a mandatory delay law	0.328	0.469
Welfare waiver	=1 if a state implements a statewide welfare waiver	0.057	0.219
TANF	=1 if a state implements TANF	0.277	0.442
Welfare benefit	Maximum monthly AFDC/TANF benefit for a three-person family ($1000)	0.358	0.145
Family cap	=1 if a state implements a family cap on incremental increases in welfare benefits	0.144	0.347
UR	Average annual state unemployment rate (/100)	0.063	0.019
Governor	=1 if the current state governor is Republican	0.496	0.500
President	Percent Republican vote in the previous Presidential election	0.493	0.092
Population density	Number of individuals per square mile (/100)	2.469	5.704
Marriage	Marriage rate, per 1000 in the population (/100)	0.094	0.070
BA	Percent of population ages 15 and over with at least a bachelor's degree	0.207	0.053
Females ages 15–19	Percent of female population ages 15–44 between ages 15 and 19	0.166	0.024
Females ages 15–44	Female population ages 15–44 (/1000)	2321.660	1777.079

Notes: Summary statistics in this table are derived from state panel data between 1975 and 2005 in which state-year combinations of CDC abortion data are non-missing (unbalanced panel $N = 1555$). There are two state-year combinations with zero one-child abortions (Wyoming, 2000 and 2001); these are dropped from the regressions, given that the log of all abortion rates is taken. Summary statistics for the State EITC are based those state-years with a state EITC implemented. All summary statistics are weighted by the female population ages 15–44.

weighted by the female population ages 15–44, and standard errors are corrected for arbitrary heteroskedasticity using robust standard errors.

5. Results

5.1. Summary statistics

Table 1 presents summary statistics for the key abortion variables in this analysis. The overall abortion rate is approximately 21 abortions per 1000 women of childbearing age. Stated another way, 2.1% of such women have an abortion in a given year. These figures mask substantial year-to-year variation in abortion rates: after peaking at about 24 abortions per 1000 women throughout the early-1980s, the abortion rate declined to a low of 16 between 2000 and 2005. Indeed, a regression of the (log) abortion rate on a linear time trend suggests that the abortion rate between 1975 and 2005 declined about 1.5% per year. Table 1 also shows that the standard deviation around the mean abortion rate (for the entire study period) is 11.1 abortions per 1000 women of childbearing age. Within-state, over time variation in abortion rates is comparatively low, ranging from 1.4 standard deviations in New Mexico to 7.1 standard deviations in California.[17] Thus, most of the variation in abortion rates comes from cross-state differences.[18]

Descriptive results for the EITC and other policy variables are also provided in Table 1. The real value of EITC maximum credit for families with one child increased about 94% between 1975 and 2005, while the value of the credit for multiple-child families increased over 200%. Given that the impact of the EITC is identified using primarily within-state variation, Table 1 disaggregates the combined maximum credit into its federal and state components. Among states with an EITC, there is sizeable variation in benefits (*SD* = $384) around the mean benefit ($595).[19] A Medicaid restriction was in effect for 56% of the study period, parental involvement laws were in place for 32% of period, and mandatory delay legislation covered only 12% of the sample period. Table 1 reveals that border state policies are potentially important influences on fertility behavior: fully 70% of adjacent states implemented a parental involvement law between 1975 and 2005, and another 33% of adjacent states enacted mandatory waiting periods. Finally, welfare reform legislation through AFDC waivers was in place for 6% of the study period, and PRWORA was in effect for approximately 28% of the study period.

[17] The District of Columbia is a clear outlier in these data, with an average abortion rate of 109 per 1000 women of childbearing age and a standard deviation of 45. Robustness checks with the District omitted leave the main results intact.

[18] Indeed, a regression of the abortion rate on state and year fixed effects produces an R^2 of 83%, suggesting that only 17% of the variation in abortion rates is due to factors that change within states over time.

[19] Another way to measure the amount of within-state variation in the EITC is to count the number of times states altered the credit rate between 1975 and 2005. Across all states, the average state changed its credit rate 1.02 times (the minimum number of policy changes is zero and the maximum is eight). Among states with an EITC, the average number of changes is approximately three. This suggests a modest amount of within-state policy variation, although it should be considered a lower bound, as the actual state credit *amount* changed each year due to indexing for inflation.

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5.2. Estimation results

Main results from the abortion regressions are shown in Table 2. It presents estimates from OLS regressions of the overall abortion rate on the EITC maximum credit. Columns (1)–(3) show results from the baseline model, and columns (4)–(7) present a number of specification checks. Specifically, I test the sensitivity of the basic results to the inclusion of state-specific time trends, the log number of abortions as the dependent variable, a one-period lag of the EITC maximum credit.

The estimate in column (1) of Table 2 suggests that a $1000 increase in the combined EITC maximum credit is associated with a decrease in the abortion rate of 8%. However, this model does not account for observable state factors, as well as unobserved heterogeneity, that may influence individual fertility and state-level social policy decisions. When the full set of policy, economic, demographic, and political controls are added in column (2), the estimated effect of the EITC increases dramatically. The coefficient implies that a $1000 increase in the EITC is associated with a 17% decrease in the abortion rate. Adding state fixed effects and year dummies [column (3)] reduces the impact of the EITC to a 7.6% reduction in the abortion rate. The coefficient remains statistically significant at conventional levels. This effect implies a reduction of approximately 1.6 abortions per 1000 women of childbearing age.

Results presented in columns (4)–(7) suggest that the EITC effect is fairly robust to a number of specification checks. Including state-specific time trends further reduces the magnitude of the EITC coefficient, rendering it imprecisely estimated. However, substituting the log number of abortions for the log abortion rate produces a statistically significant coefficient on the EITC.[20] The estimate implies that a $1000 increase in the maximum EITC leads to a 7.7% reduction in the number of abortions. Columns (6) and (7) explore the impact of a one-period lag in the maximum credit. Column (6) uses as the dependent variable the log abortion rate, and column (7) uses the log number of abortions. State-specific time trends are included in both models. Parameter estimates reveal that the impact of the EITC is robust to changes in the lag structure. If anything, the lagged results imply slightly larger effects, especially in the model using the log number of abortions. Results from this model show an 11% reduction in the number of abortions in response to a $1000 increase in the EITC's maximum credit.[21]

Turning to the remaining policy variables, I find mixed evidence for the impact of abortion restrictions and other policy reforms. Bans on Medicaid funding are associated with reductions in abortion, but the effect becomes statistically insignificant once state-specific time trends are added to the model. On the other hand, the impact of mandatory delay laws doubles in magnitude and becomes statistically significant once the time trends are included. Consistent with previous studies, border state policies have conflicting and counter-theoretical effects on abortion. For example, border state restrictions on teen abortions are associated with decreased in-state abortion rates, a result contrary to theoretical predictions but nevertheless mirrored by other studies (Blank et al., 1996; Bitler and Zavodny, 2001). Finally, the results provide mixed evidence on the impact of social policy reforms. The negative coefficient on welfare waivers is inconsistent with a theoretical model of welfare reform (Klerman, 1998). Although family caps are predicted to increase abortion rates, the policy is associated with reductions in abortion. This result is consistent with previous work by Kearney (2004) and Levine (2003), who find that family caps lead to small but imprecisely estimated increases in birth rates.

5.3. Births and pregnancies

The results in Table 2 suggest that expansions to the EITC are associated with reductions in the overall abortion rate. As the theoretical framework makes clear, the EITC has implications for abortion decisions at two points in the fertility decision-making process: when choices are made regarding sexual and contraceptive practices and after a pregnancy occurs. If the EITC does not alter sexual and contraceptive behavior, a reduction in the number of abortions should be attributable to an increase in the number of births. On the other hand, if the EITC influences the intensity with which women use contraception, the reduction in abortion rates can be traced to a decline in pregnancies.

The results presented in Table 3 allow us to examine these pathways. Estimated in columns (1) and (2) are OLS regressions of state-level birth rates on a one-period lag of the EITC maximum credit. The mean birth rate between 1975 and 2005 is about 67 births per 1000 women of childbearing age. Columns (3) and (4) show results from OLS regressions of pregnancy rates on the identical set of variables. Following Levine (2002, 2003), I construct the pregnancy rate by summing the birth and abortion rates.[22] The mean pregnancy rate over the study period is 88 pregnancies per 1000 women of childbearing age. The odd-numbered columns estimate models without state-specific time trends, and the even-numbered columns add these controls to the model. Although birth rate data are available for all state-years during the study period, for consistency I limit

[20] Defining the outcome variable in terms of the log number of abortions is tantamount to moving the denominator of the abortion rate (number of women ages 15 to 44) from the left-hand-side of the equation into the right-hand-side. Therefore, in all models that include the log number of abortions, I add a control for the log number women ages 15 to 44.

[21] In a further robustness check, I estimate a model in which all of the right-hand-side variables are lagged one year. Doing so does not change the estimated effects of the EITC on abortion rates.

[22] Creating a pregnancy rate based on birth and abortion rates necessarily misses pregnancies that end in a spontaneous abortion (i.e., a miscarriage). Therefore, the pregnancy rate used in this analysis is likely to be below the actual rate. Determining the rate at which pregnancies are spontaneously aborted, however, is a difficult task because many miscarriages occur very early in a pregnancy, often before a woman realizes she is pregnant. Estimates by Wilcox et al. (1999) and Wang et al. (2003) suggest that 25% of pregnancies are miscarried by the 6th week, while clinical miscarriages—those occurring after the 6th week—occur in about 8% of pregnancies (Wang et al., 2003).

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Table 2
The impact of the EITC on abortion.

Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Maximum EITC	−0.080*** (0.011)	−0.172*** (0.028)	−0.076** (0.035)	−0.055 (0.048)	−0.077* (0.046)		
Maximum $EITC_{t-1}$						−0.086* (0.045)	−0.112*** (0.043)
Medicaid		−0.135*** (0.030)	−0.068*** (0.022)	−0.027 (0.022)	−0.026 (0.022)	−0.026 (0.022)	−0.026 (0.022)
Involve		−0.240*** (0.028)	0.001 (0.021)	0.025 (0.018)	0.023 (0.017)	0.025 (0.018)	0.023 (0.017)
Delay		−0.027 (0.041)	−0.022 (0.019)	−0.043** (0.020)	−0.050** (0.020)	−0.044** (0.020)	−0.052** (0.020)
Medicaid border		−0.053 (0.035)	0.064* (0.037)	0.052 (0.043)	0.047 (0.042)	0.055 (0.043)	0.050 (0.042)
Involve border		−0.125*** (0.035)	−0.023 (0.020)	−0.044** (0.020)	−0.044** (0.020)	−0.045** (0.021)	−0.044** (0.020)
Delay border		−0.026 (0.048)	0.001 (0.020)	0.017 (0.017)	0.006 (0.017)	0.017 (0.017)	0.005 (0.017)
Welfare waiver		0.056 (0.067)	0.012 (0.023)	−0.042** (0.021)	−0.044** (0.021)	−0.040* (0.020)	−0. 041* (0.021)
TANF		0.258*** (0.069)	0.013 (0.043)	0.049 (0.041)	0.027 (0.044)	0.051 (0.042)	0.027 (0.045)
Welfare benefit		0.412*** (0.114)	0.169 (0.166)	0.266 (0.162)	0.194 (0.168)	0.261 (0.161)	0.185 (0.166)
Family cap		−0.065 (0.046)	−0.030 (0.020)	−0.029 (0.025)	−0.031 (0.025)	−0.030 (0.025)	−0.033 (0.025)
DV: ln(abortion rate)	√	√	√	√		√	
DV: ln(No. of abortions)					√		√
State characteristics	No	Yes	Yes	Yes	Yes	Yes	Yes
State fixed effects	No	No	Yes	Yes	Yes	Yes	Yes
Year dummies	No	No	Yes	Yes	Yes	Yes	Yes
State-specific time trends	No	No	No	Yes	Yes	Yes	Yes
R^2	0.066	0.493	0.877	0.928	0.987	0.928	0.987
State-year combinations	1555	1555	1555	1555	1555	1555	1555

Notes: Regression coefficients in this table come from the unbalanced panel of state-year combinations between 1975 and 2005, and are weighted by the female population ages 15–44. Standard errors, reported in parentheses, are corrected for arbitrary heteroskedasticity using Huber–White robust standard errors. Models (1)–(4) and (6) use the log abortion rate as the dependent variable, while models (5) and (7) use the log number of abortions as the dependent variable. All models rely on CDC-reported abortion figures. Models (1)–(5) parameterize the EITC by using the contemporaneous 2+-child maximum credit. Models (6) and (7) parameterize the EITC variable by once-lagging the 2+-child maximum credit. Models (2)–(7) control for UR, governor, president, population density, marriage, BA, and females ages 15–19. Models (5) and (7) control for the log female population ages 15–44. See Table 1 for an explanation of the variable definitions.

*** Indicate statistical significance at the 0.01 levels.
** Indicate statistical significance at the 0.05 levels.
* Indicate statistical significance at the 0.10 levels.

Table 3
The impact of the EITC on births and pregnancies.

Variable	(1)	(2)	(3)	(4)
Maximum $EITC_{t-1}$	−0.037*** (0.010)	−0.082*** (0.029)	−0.063*** (0.009)	−0.068*** (0.018)
Medicaid	0.004 (0.008)	−0.001 (0.009)	−0.012* (0.006)	−0.010 (0.006)
Involve	−0.025*** (0.004)	−0.010*** (0.003)	−0.025*** (0.004)	−0.011*** (0.003)
Delay	−0.028*** (0.005)	0.013** (0.006)	−0.018*** (0.005)	0.003 (0.006)
Medicaid border	−0.021** (0.009)	0.013 (0.021)	0.006 (0.009)	0.022 (0.013)
Involve border	−0.012 (0.009)	0.007 (0.007)	−0.023*** (0.007)	−0.000 (0.006)
Delay border	0.013** (0.005)	0.008 (0.006)	0.015** (0.006)	0.012* (0.006)
Welfare waiver	0.031*** (0.010)	−0.000 (0.007)	0.026** (0.011)	−0.015* (0.008)
TANF	−0.009 (0.019)	0.042 (0.031)	−0.000 (0.022)	0.048 (0.036)
Welfare benefit	0.428*** (0.065)	0.266*** (0.055)	0.410*** (0.065)	0.357*** (0.066)
Family cap	0.017*** (0.006)	0.018* (0.009)	0.008 (0.006)	0.017** (0.008)
DV: ln(birth rate)	√	√		
DV: ln(pregnancy rate)			√	√
Mean of birth rate	66.51			
Mean of pregnancy rate			87.46	
State characteristics	Yes	Yes	Yes	Yes
State fixed effects	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes
State-specific time trends	No	Yes	No	Yes
R^2	0.834	0.906	0.900	0.948
State-year combinations	1.555	1.555	1.555	1.555

Notes: Regression coefficients in this table come from the unbalanced panel of state-year combinations between 1975 and 2005, and are weighted by the female population ages 15–44. Although fertility data are available for all state-years, for consistency the analysis is restricted to state-years with non-missing CDC abortion data. Standard errors, reported in parentheses, are corrected for arbitrary heteroskedasticity using Huber–White robust standard errors. Models (1) and (2) and (6) use the log fertility rate as the dependent variable, while models (3) and (4) use the log pregnancy rate as the dependent variable. The pregnancy rate is constructed by summing the CDC abortion rate and the fertility rate (per 1000 women ages 15–44). All models parameterize the EITC variable by once-lagging the 2 + -child maximum credit. All models control for UR, governor, president, population density, marriage, BA, and females ages 15–19. See Table 1 for an explanation of the variable definitions.
*** Indicate statistical significance at the 0.01 levels.
** Indicate statistical significance at the 0.05 levels.
* Indicate statistical significance at the 0.10 levels.

the analysis to the observations for which abortion data are available (N = 1555). Both dependent variables are expressed in logarithms.

A number of interesting findings emerge from Table 3. Increases in the maximum EITC are associated with *reductions* in aggregate birth rates. As shown in column (2), a $1000 increase in the maximum credit is expected to decrease birth rates by about 8%. Such results find some support in the empirical literature. For example, Baughman and Dickert-Conlin (2007) determine that EITC expansions are associated with decreases in first and higher-order births among white women. Moreover, recent work by Kearney (2004) finds that lowering incremental welfare benefits through family caps leads to small increases in birth rates. Based on the theoretical model, if the EITC is associated with reductions in birth and abortion rates, the credit must encourage women to alter sexual and contraceptive behavior, as reflected in pregnancy rates. Results in column (4) corroborate this: a $1000 increase in the maximum credit is associated with a 7% decline in pregnancy rates. Therefore, it appears that the EITC changes fertility outcomes by altering earlier decisions about the level of sexual activity and contraceptive intensity.

5.4. Specification checks

The main results are subjected to a number of sensitivity tests, as shown in Tables 4 and 5. Table 4 examines the impact of the EITC across several demographic sub-groups. I also explore the sensitivity of the main results to changes in the measurement and functional form of the EITC variable, alternative weights and data restrictions, and a correction for autocorrelation. Table 5 compares results using CDC abortion counts with those drawn from the AGI. As others note, the impact of abortion restrictions is somewhat sensitive to the abortion data source, making such a comparison important for the EITC as well (Bitler and Zavodny, 2001; Levine, 2003; Matthews et al., 1997).

The first two rows in Table 4 explore heterogeneity in the impact of the EITC across sub-groups of women defined by race.[23] I use as the dependent variable the log number of abortions among women in the relevant sub-group. The coefficient on the maximum credit shows a statistically significant 20% reduction in the number of abortions for black women, but a smaller and imprecisely estimated reduction for white women. These findings are contrary to those reported in Duchovny (2001) and Baughman and Dickert-Conlin (2007), both of which find greater fertility responses among white women. In results not reported, I estimate models that substitute the incremental EITC for the maximum credit. Across all sub-groups, increases in

[23] As previously noted, these analyses are constrained to the period 1990–2005. These sub-group analyses include state and year fixed effects, state-specific time trends, and a control for the log number of women ages 15–44.

Table 4
Specification checks.

Specification	Coefficient on the EITC (1)	(2)
DV: Log of abortions to black women	–	−0.200*** (0.072)
DV: Log of abortions to white women	–	−0.139 (0.093)
Abortion rate and the number of abortions(no logs)	−3.045*** (0.528)	−15,453*** (3524)
Log of maximum EITC	−0.221* (0.130)	−0.177 (0.133)
Maximum EITC for families with one child	−0.088 (0.058)	−0.062 (0.061)
Analysis of 1990–2005 data, refundable state EITCs	−0.104*** (0.038)	−0.113*** (0.038)
Balanced panel data	−0.090* (0.046)	−0.093** (0.046)
Unweighted	−0.179*** (0.048)	−0.165*** (0.047)
Prais–Winsten correction for serial correlation, AR(1) process	−0.088** (0.040)	−0.076* (0.040)
Prais–Winsten correction for serial correlation, AR(1) process, and panel-corrected standard errors	−0.088* (0.045)	−0.076* (0.046)
Exclude states with at least 20% of abortions performed on non-residents	−0.095** (0.037)	−0.100** (0.039)
Exclude states with at least 10% of abortions performed on non-residents	−0.084** (0.037)	−0.086** (0.038)
Add control for whether border state has an EITC program	−0.070* (0.036)	−0.059 (0.038)
Include additional demographic controls	−0.072* (0.037)	−0.064* (0.038)
DV: ln(abortion rate)	√	
DV: ln(No. of abortions)		√

Notes: Unless otherwise noted, regression coefficients in this table come from the unbalanced panel of state-year combinations between 1975 and 2005, and are weighted by the female population ages 15–44. Standard errors, reported in parentheses, are corrected for arbitrary heteroskedasticity using Huber–White robust standard errors. In addition, unless otherwise noted, the coefficient presented is the estimate on Maximum EITC, the combined federal/state EITC for families with two or more children. The dependent variable in the first two models is the log number of abortions black ($N = 579$), and white ($N = 585$) women. Abortion data by demographic group are available between 1990 and 2005. The number of observations in the balanced panel regressions is 1395. This model includes state-specific time trends. The Prais–Winsten models assume an AR(1) process that is common across all states. The model with additional demographic controls includes the percent of women in 5-year age-groupings between ages 15 and 44. All models control for UR, governor, president, population density, marriage, BA, and females ages 15–19, while the estimates in (2) control for the log female population ages 15–44. All models include state fixed effects and year dummies. Models one through four and model nine include state-specific time trends.
*** Indicate statistical significance at the 0.01 levels.
** Indicate statistical significance at the 0.05 levels.
* Indicate statistical significance at the 0.10 levels.

Table 5
Comparison of empirical results using CDC and AGI abortion data.

Variable	(1a)	(1b)	(2a)	(2b)	(3a)	(3b)
Maximum EITC	−0.070*** (0.017)	−0.075*** (0.016)	−0.242*** (0.043)	−0.267*** (0.041)	0.002 (0.067)	0.036 (0.048)
Medicaid			−0.157*** (0.037)	−0.156*** (0.033)	−0.027 (0.028)	−0.022 (0.022)
Involve			−0.208*** (0.037)	−0.253*** (0.035)	0.016 (0.025)	0.030* (0.018)
Delay			0.045 (0.075)	0.009 (0.070)	−0.063* (0.037)	−0.055* (0.030)
Medicaid border			−0.019 (0.045)	−0.012 (0.044)	0.038 (0.047)	0.004 (0.040)
Involve border			−0.109*** (0.039)	−0.077** (0.034)	−0.045* (0.026)	−0.065*** (0.018)
Delay border			0.015 (0.085)	0.110 (0.076)	0.063* (0.034)	0.010 (0.028)
Welfare waiver			0.033 (0.102)	−0.038 (0.086)	−0.026 (0.036)	−0.020 (0.028)
TANF			0.268** (0.114)	0.222** (0.107)	0.005 (0.160)	0.158 (0.121)
Welfare benefit			0.448*** (0.152)	0.299** (0.132)	0.282 (0.252)	0.115 (0.192)
Family cap			−0.040 (0.092)	0.077 (0.077)	−0.044 (0.043)	0.042 (0.033)
DV: ln(abortion rate)	√	√	√	√	√	√
CDC/AGI Data	CDC	AGI	CDC	AGI	CDC	AGI
State characteristics	No	No	Yes	Yes	Yes	Yes
State fixed effects	No	No	No	No	Yes	Yes
Year dummies	No	No	No	No	Yes	Yes
State-specific time trends	No	No	No	No	Yes	Yes
R^2	0.039	0.047	0.497	0.557	0.912	0.943
State-year combinations	912	912	912	912	912	912

Notes: Regression coefficients in this table come from the unbalanced panel of state-year combinations between 1975 and 2005, and are weighted by the female population ages 15–44. Specifically, the estimates are derived from state-years in which CDC and AGI abortion data are both available. These years include 1975–1982, 1984–1985, 1987–1988, 1991–1992, 1995–1996, 2000, and 2005. Standard errors, reported in parentheses, are corrected for arbitrary heteroskedasticity using Huber–White robust standard errors. All models use as the dependent variable the log CDC or AGI abortion rate. All models parameterize the EITC by using the contemporaneous 2 + -child maximum credit. Models (2a) through (3b) control for UR, governor, president, population density, marriage, BA, and females ages 15–19. See Table 1 for an explanation of the variable definitions.
*** Indicate statistical significance at the 0.01 levels.
** indicate statistical significance at the 0.05 levels.
* indicate statistical significance at the 0.10 levels.

the incremental benefit for one child and multiple children generate larger abortion responses than the maximum credit. These results are consistent with those reported in Baughman and Dickert-Conlin (2007).

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The next several rows investigate changes to the parameterization of the EITC variable, construction of the panel dataset, and the weighting procedure. Generally speaking, the results are robust to these specification checks. Measuring the dependent variable in levels does not substantially change the results. The estimate in column (1) suggests that increasing the EITC maximum credit by $1000 is associated with a reduction in the abortion rate of about three abortions per 1000 women of childbearing age. This estimate translates to a 15% (3.045/20.94) decrease in the overall abortion rate, a somewhat larger effect than the one from the log model.

This study relies in part on variation in state credits to identify the impact of the EITC. Two concerns are raised by this approach. First, several states structure their EITC programs as nonrefundable tax credits, thereby reducing the likelihood that low-income families will qualify for benefits. Second, given that most states implemented their EITC programs throughout the 1990s, a large portion of this study's observation period contains very little cross-state variation in the EITC.[24] One might therefore question whether the main results are robust to a shortened panel beginning in the early-1990s. To explore these concerns, I create an EITC variable that uses variation in the state credit only if it is structured as a refundable credit. I then test this variable in a regression using state-year combinations between 1990 and 2005. As shown in Table 4, results from this specification check are quite similar to the main results.

Using the balanced panel of state-years does not appreciably change the results. When the six states with missing abortion data (Alaska, California, Louisiana, New Hampshire, Oklahoma, and West Virginia) are omitted from the analysis, the coefficient on the maximum EITC changes only slightly, to a 9% reduction in the abortion rate [column (1)]. Following Bitler and Zavodny (2001), I explore the robustness of the main results to a correction for serial correlation. Indeed, tests reveal that the errors in the abortion regressions are serially correlated within states.[25] To correct for within-state autocorrelation, I use a panel data generalization of the Prais–Winsten method that assumes the autocorrelation follows a common AR(1) process across all states. Column (1) provides autocorrelation-corrected estimates using the log abortion rate as the dependent variable, and column (2) provides autocorrelation-corrected estimates using the log number of abortions. In both cases, the parameter estimate on the EITC is similar to the uncorrected models and continues to be statistically significant. In addition, correcting for the possibility that the errors are contemporaneously correlated across panels (using panel-corrected standard errors) does not appreciably change the precision with which the EITC coefficients are estimated.

As previously stated, one of the drawbacks of the CDC data is that abortion counts are reported by state of occurrence, rather than state of residence. Therefore, to examine the influence of non-state residents crossing state borders to receive an abortion, I estimate models that exclude states with at least 20% and 10% of its abortions performed on out-of-state residents.[26] Results presented in Table 4 suggest that any cross-state abortion migration is not driving the results. Specifically, excluding the 14 states that perform 10% or more of its abortions on non-state residents leads to an EITC effect implying an 8% reduction in the abortion rate and a 9% reduction in the number of abortions. These estimates remain statistically significant despite a 28% reduction in the sample size (from $N = 1555$ to $N = 1123$).

As a final specification check, I compare regression estimates derived from CDC and AGI abortion data. Results from this exercise are presented in Table 5. To maintain consistency, all estimates are derived from state-year combinations in which CDC and AGI abortion data are both available.[27] This reduces the sample to 912 state-year combinations. All models use as the dependent variable the log CDC or AGI abortion rate. The AGI data tend to yield coefficients on the EITC and other policy variables that are consistent with those from the CDC data. In models without controls, the coefficient on the EITC implies a 7% and 7.5% reduction in the CDC- and AGI-based abortion rate, respectively. Adding states' observable characteristics leads to similar increases in the magnitude of the EITC effect (CDC estimate: −0.242; AGI estimate: −0.267). Finally, adding the full set of controls for state-level unobserved heterogeneity also leads to similar EITC estimates (CDC estimate: 0.002; AGI estimate: 0.036). Although they are not statistically significant, the positive coefficients in the final model reflect the fact that important gaps appear in the AGI data collection throughout the 1990s and 2000s, a period in which several reforms to the EITC were enacted. For example, AGI abortion counts are reported only three times between 1996 and 2005. During this 10-year period, 10 states implemented their own EITC.[28] Such data gaps therefore reduce substantially the amount of identifying variation in the EITC. With the exception of the coefficients on family caps, CDC- and AGI-based abortion data produce qualitatively similar results for the remaining policy variables. The estimate for family caps is negatively signed when using the CDC data and is positively signed when using the AGI data, although in neither case is the estimate statistically significant.

[24] Indeed, 13 of the 17 state EITCs that operated between 1975 and 2005 were implemented on or after 1990. These states are: Colorado (1999, but suspended after 2002), District of Columbia (2000), Iowa (1990), Illinois (2000), Indiana (2002), Kansas (1998), Maine (2000), Massachusetts (1997), Minnesota (1991), New Jersey (2000), New York (1994), Oklahoma (2002), and Oregon (1997). Of the 13 states, 10 structured the program as a refundable tax credit. These states are: Colorado, District of Columbia, Illinois, Indiana, Kansas, Massachusetts, Minnesota, New Jersey, New York, and Oklahoma.

[25] A Wooldridge (2002) test of no serial correlation in panel data yields a highly significant F-statistic in the log abortion rate ($F = 57.76$) and log number of abortions ($F = 56.67$) regressions.

[26] I use CDC (2008) to make these classifications. States excluded because at least 20% of its abortions are performed on out-of-state residents include the following: District of Columbia, Kansas, North Dakota, Rhode Island, and Tennessee. States excluded because at least 10% of its abortions are performed on out-of-state residents include the following: District of Columbia, Kansas, North Dakota, Rhode Island, Tennessee, Alabama, Arkansas, Georgia, Nebraska, North Carolina, Oregon, South Dakota, Vermont, and West Virginia.

[27] As previously stated, AGI abortion data are available for the following years: 1975–1982, 1984–1985, 1987–1988, 1991–1992, 1995–1996, 2000, and 2005. These data are also reported by the state of occurrence.

[28] These states are: Colorado (1999, but suspended after 2002), District of Columbia (2000), Illinois (2000), Indiana (2002), Kansas (1998), Maine (2000), Massachusetts (1997), New Jersey (2000), Oklahoma (2002), and Oregon (1997).

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6. Conclusion

Using a panel of states between 1975 and 2005, I develop and test a model of the relationship between the EITC and abortion. The theoretical model predicts that the EITC has implications for abortion behavior at two points in the fertility decision-making process: when women are making decisions about the level of sexual activity and contraceptive intensity and after a pregnancy occurs. While the model highlights the possibility that the EITC can increase the number of abortions, the income and price effects associated with the credit suggest that reductions in abortion are highly likely. The empirical results generally support these theoretical propositions, and can be summarized as follows. Increases in the EITC are associated with statistically significant reductions in the overall abortion rate. Main results imply that a $1000 increase in the EITC maximum credit is expected to reduce the abortion rate by 7.6%. I also find that reductions in abortion are not offset by increases in births. Rather, the EITC appears to be reducing pregnancy rates.

These findings warrant some discussion. Generally speaking, results in this paper are broadly consistent with and may help to explain some counter-theoretical findings in previous studies. In particular, a recent paper by Baughman and Dickert-Conlin (2007) determines that increases in the EITC are associated with small *reductions* in births. Moreover, Kearney (2004) and Levine (2002) find that policies decreasing incremental welfare benefits (through family caps) are associated with *increases* in the birth rate.

One interpretation of the results in this study and Baughman and Dickert-Conlin (2007) focuses on the labor supply effects of the EITC. Previous research finds consistent evidence that the credit is associated with greater employment levels among disadvantaged women (Eissa and Liebman, 1996; Grogger, 2003; Herbst, 2008; Meyer and Rosenbaum, 2001). By increasing the returns to work, however, the EITC raises the opportunity costs associated with bearing and raising children. Such employment effects are important for understanding a key result in this paper: reductions in abortion appear to originate in decisions regarding sexual behavior and contraceptive intensity, as opposed to after a pregnancy occurs. In other words, the channel through which the EITC reduces abortions is a decrease in risky sexual behavior (perhaps through more diligent contraceptive use), and not a shift from abortions to births conditional on being pregnant. Future research should subject this proposition to more rigorous testing, however.

As Baughman and Dickert-Conlin (2007) point out, another interpretation of these results emphasizes Becker's (1991) theory on the demand for children. This model suggests that positive income shocks create trade-offs for families between the number (quantity) and development (quality) of children. If children are assumed to be normal goods, the introduction of an EITC is predicted to increase the demand for children. Alternatively, families might be concerned with the quality of children. In this case, the introduction of an EITC might lead to reductions in higher-order childbearing and an increase in human capital investments for existing children. In this conceptualization, the EITC acts as a mechanism by which parents substitute greater investments in higher-quality children for increases in the quantity of children. Results in this study and others suggest that the latter prediction is more appropriate for explaining fertility responses to the EITC.

Understanding the role of the EITC in shaping women's fertility decisions is important for a number of reasons. Previous studies find that other factors related to fertility—such as abortion legalization and the availability of contraception—have powerful effects on women's lifecycle fertility, marriage decisions, and labor market outcomes (Ananat et al., 2007; Angrist and Evans, 1999; Bailey, 2006; Goldin and Katz, 2002). Policies affecting fertility also have implications for child outcomes, including the propensity to commit crime and use illicit drugs, educational attainment, welfare use, and employment (Ananat, et al., 2006; Charles and Stephens, 2006; Donohue and Levitt, 2001; Gruber et al., 1999). For disadvantaged children, the additional income a family receives from the EITC can be particularly important. Recent studies suggest that the credit lifts 2.7 million children above the poverty line each year, and that EITC-induced increases in family income are associated with gains in children's cognitive test scores (Center on Budget and Policy Priorities, 2005; Dahl and Lochner, 2008). Therefore, by potentially influencing short-run fertility decisions, the EITC has the ability to shape long-term outcomes for adult recipients and their children.

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