Did Expanding the EITC Promote Motherhood?

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During the 1990's the Earned Income Tax Credit (EITC) emerged as a primary means of providing income support for low-income families in the United States. In an effort to keep the program well targeted, the credit largely restricts eligibility to tax filers with children. One potentially unintended consequence of this design is that it might encourage childbearing. We raise the question of whether the EITC, through its generous benefits to families with children, actually increases fertility.

We approach this topic for three reasons. The first is to expand upon an existing literature of economic incentives and fertility using the EITC expansion as a large exogenous variation in the price of childbearing. Findings in the welfare literature are inconclusive (Robert A. Moffitt, 1998), and the income tax literature typically finds small, but statistically significant effects of the income-tax system on fertility behavior (e.g., Leslie A. Whittington et al., 1990). Second, declining fertility rates in many Western counties raise the general issue of whether the tax system can be used as a tool for encouraging fertility. Finally, by considering the link between the EITC and fertility, we question a common, yet untested, assumption in the literature on the EITC and the labor supply of single parents: that the presence of a child is exogenous to the value of the EITC.

I. Pro-Natalist Features of the EITC

Eligibility for and value of the EITC both changed a great deal in the past decade, but recipients have always been predominantly parents. Initially the credit was only available to tax units with qualifying children, but its value did not differ by number of children in a family. Beginning in 1991, the maximum credit value for families with two or more children was set higher than the credit for those with just one child. In 1994 a small benefit to childless tax units with very low earnings was added.

Fertility incentives changed as the value of the EITC expanded during the 1990's. The largest expansion in the credit's history, authorized by the Omnibus Budget Reconciliation Act (OBRA) of 1993, phased in between 1994 and 1996. Between 1990 and 1999 the maximum credit available to an individual with one child grew from $953 to $2,312, and the incremental credit for a second child grew from $0 to $1,504. Additionally, implementation of EITC programs in many states accompanied the expansion of the federal EITC. In 1990 only five states (Iowa, Maryland, Rhode Island, Vermont, and Wisconsin) had EITC's, and by 1999 six additional states (Colorado, Kansas, Massachusetts, Minnesota, New York, and Oregon) had EITC's; additionally, four of the five states with a credit in 1990 expanded it during this ten-year period. All states calculate their EITC's as some percentage (ranging between 5 and 50 percent), of the federal EITC, so the state credits provide the same fertility incentives as the federal EITC.

The simplest fertility incentive in the design of the EITC is that income-eligible women benefit by having a first child because the large credit is only available to families with children. Unlike most welfare programs, eligibility for the EITC is not conditional on marital status, so this fertility incentive exists for both married and unmarried childless women. For women who already have a child, the fertility incentives in the EITC are more complex.[1] In this analysis

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[1] Specifically, an ambiguous fertility incentive exists for higher-order births to women in the phase-in range of the credit. Although the income effect of the earnings subsidy makes all normal goods more attractive for all women, the

we consider only the credit's most straightforward incentive for fertility: that its design encourages first births. Our hypothesis is that expansions in the credit increased the probability that lower-income women became mothers.

II. Data and Empirical Approach

To test the hypothesis that the EITC influences the decision to have children, we examine birthrates over the course of the 1990's, controlling for state and demographic characteristics and exploiting variation in state EITC programs over time to identify the effect of the credit on fertility. To construct the birthrates we use data from U.S. birth certificates between 1990 and 1999.[2] The Natality Detail File, maintained by the National Center for Health Statistics, documents information on state birth certificates for the approximately 40 million births over these ten years.

Each birth certificate record contains the following basic information:

(i) Birth year;
(ii) Mother's state of residence;
(iii) Age, race, education level and marital status of mother;
(iv) Birth order of this child.

We select only first births and restrict our sample to women with less than a college education as an indicator for those most likely to be affected by the EITC expansion.[3]

Because it would be inappropriate to compare birth counts across states without accounting for population differences, we group the data into cells and normalize them by a measure of the *at-risk* population. Using the 1990 Census, we estimate cell denominators by a method (which accounts for migration) described in Baughman and Dickert-Conlin (2002).

The birthrate for characteristics i, the dependent variable in this analysis, is:

$$(\text{Birthrate})_i$$

$$= (\text{No. First Births})_i / (\text{No. Childless Women})_i$$

We break birthrates into cells according the following demographic categories of the mother:

(i) Age group (5): 15–24, 25–29, 30–34, 35–39, 40–44;
(ii) Race group (2): white, nonwhite;
(iii) Education group (2): Fewer than 12 years, 12–15 years;
(iv) State (50);
(v) Year (10): 1990, ... , 1999.

This procedure results in approximately 20,000 cell observations. Of these, 13 percent contain no births. Zero births are most likely in the nonwhite race category in smaller states.

The mean first birthrate is 6.4 births per 100 women across all cells and years, with considerable variance across cells. The mean birthrate for nonwhite women is 7.2 and for white women is 6.2. Overall, the birthrate for married women (16.2) is significantly higher than that for unmarried women (3.7), but the gap is much smaller for nonwhite women.

Figure 1 plots birthrates by race and whether or not a state ever has a state EITC for women having a first birth between 1990 and 1999. Obviously, fertility behavior differs significantly by race. Among white women, states with EITC's have lower first birth rates than the states without EITC's in all years, and neither trend changes significantly. In contrast, for nonwhite women, first birth rates are consistently higher in states with an EITC, and birthrates in EITC states grow relative to those in non-EITC states over the 10 years, suggesting of a policy response to the EITC.

We turn to regression analysis to see if measurable demographic or policy factors account

EITC increases the net wages of the woman in the phase-in range, and this raises the opportunity cost of childbearing (V. Joseph Hotz et al., 1997). In Baughman and Dickert-Conlin (2002), we present an analysis of the impact of the EITC on higher-order births.

[2] We use birthrates to study fertility (as do Theodore Joyce et al. [1998]) because lower-income new mothers are disproportionately likely to drop out of nationally representative microdata sets.

[3] Sample selection based upon education level creates a second data issue: birth certificates do not contain mother's education for some births in our 10-year sample. We exclude Washington and Connecticut entirely from our analysis and exclude New York and New Jersey for 1990, because the problem is pervasive in these state-years.

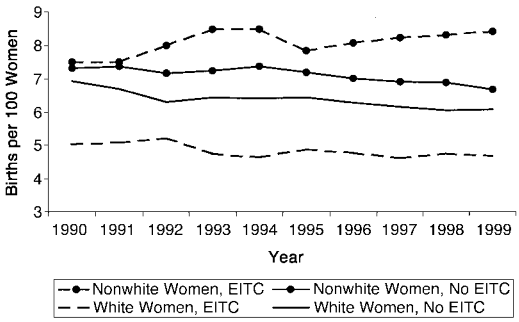

FIGURE 1. FIRST BIRTHS PER 100 WOMEN AT RISK (AGES 15–44)

for these differences among nonwhite women and if these measurable factors mask responses to the EITC among white women. We analyze birthrates within states as state supplemental EITC credits changed over time, holding constant observable demographic characteristics, using the following empirical model:

$$(\text{Birthrate})_{ist} = \alpha_s + \alpha_t + \beta \ln(\text{EITC}_{s,t-1}) + \boldsymbol{\gamma}\mathbf{X}_{it} + \boldsymbol{\pi}\mathbf{Z}_{s,t-1} + \varepsilon_{it}$$

where i indexes cell groups 1 through n, α_s is a state fixed effect, α_t is a year effect that is constant across all states and ε_{it} is an error term. The state fixed effects control for unobservable time-invariant state characteristics, such as "religiosity."

Our variable of interest, $\ln(\text{EITC}_{t-1})$, is the natural log of the real maximum state plus federal credit for a family with two children. We lag the EITC value one year, assuming that families make decisions about having children using information about the current year's tax code, and that most births occur in the next year.

The vector **X** consists of a set of dummy variables identifying demographic characteristics of the cells: age, race, and education. The vector **Z** contains a set of time-varying policy and economic variables that may influence the birthrate: state unemployment rates, the maximum AFDC/TANF (Temporary Assistance for Needy Families) benefit for a three-person family, an indicator for the presence of a state family cap after welfare reform, a measure of children's public-health-insurance eligibility, and the state's minimum marginal income-tax rate. We lag all policy control variables one year.

TABLE 1—WEIGHTED LEAST-SQUARES REGRESSIONS BY MOTHER'S RACE (DEPENDENT VARIABLE: FIRST-BIRTH BIRTHRATE)

	Unmarried		Married	
Variable	White	Nonwhite	White	Nonwhite
$\ln(\text{EITC})_{t-1}$	−0.0004†	0.0011**	0.0023	0.0076**
	(0.0002)	(0.0004)	(0.0017)	(0.0016)
Less than high-school education	0.0002	−0.0282**	0.0881**	−0.0895**
	(0.0003)	(0.0005)	(0.0033)	(0.0019)
(State unemployment rate)$_{t-1}$	−0.0009	0.0029	0.0104	0.0166*
	(0.0015)	(0.0023)	(0.0110)	(0.0084)
(Log AFDC benefits)$_{t-1}$	0.0028**	0.0083**	0.0049	0.0009
	(0.0006)	(0.0042)	(0.0056)	(0.0047)
(AFDC/TANF family cap)$_{t-1}$	−0.0011**	0.0005	0.0059	0.0148**
	(0.0005)	(0.0009)	(0.0039)	(0.0034)
(Medicaid child eligibility)$_{t-1}$	−0.00001	−0.00002	−0.00004	−0.0001
	(0.00003)	(0.00004)	(0.0002)	(0.0001)
(Minimum state income-tax rate)$_{t-1}$	−0.0002	0.0006	0.0005	0.0022
	(0.0003)	(0.0006)	(0.0021)	(0.0024)
Year fixed effect	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes
R^2:	0.64	0.86	0.72	0.73
Number of observations:	4,780	4,320	4,760	4,320

Note: Regressions also include dummies for four age categories: 15–24, 25–29, 35–39, and 40–44.
Source: Authors' calculations.
† Statistically significant at the 10-percent level.
* Statistically significant at the 5-percent level.
** Statistically significant at the 1-percent level.

III. Results

Table 1 presents results of a weighted least-squares model in which weights control for differences in the cell sizes and the precision of estimated birthrates. We split the sample by race because findings in the welfare literature and our descriptive results in Figure 1 suggest differential fertility responses to policy by race. We also split the sample by marital status because the policy implications of marital and nonmarital fertility are likely to be different.

The first two columns of Table 1 present results for unmarried women. Surprisingly, an increase in the EITC is negatively correlated with first births for white women, but it is only statistically significant at the 10-percent level.

In addition, with mean birthrates of 3.12 per 100 women, the coefficient implies an economically small elasticity of −0.01. For nonwhite women, the coefficient has the expected positive sign and is statistically significant at the 1-percent level. However, the economic significance is still small, with an estimated elasticity of only 0.02. The only policy variable that is consistently statistically significant among unmarried women is the state maximum AFDC/TANF benefit. Higher benefits are correlated with higher first birthrates. These elasticities are larger than the estimated elasticities for the EITC: 0.09 for white women and 0.14 for nonwhite women.

For the sample of married women, the EITC is positively correlated with first birth rates for both white and nonwhite women. The coefficient is statistically significant only for nonwhite women. With a mean birthrate of 13.6 per hundred women, this implies an elasticity of 0.06. While larger than the estimates for unmarried nonwhite women, this estimate is less than half the size of those found by Whittington et al. (1990) for the U.S. income tax. However, these estimates suggest that the more than 500-percent increases in average state EITC's over the 1990's may be responsible for increases in the fertility rates of 34 percent, or 4.9 percentage points.

Our regression results confirm our findings in Figure 1: the EITC is positively and statistically significantly associated with first birth rates among nonwhite, married women. While most researchers in the welfare literature find larger effects of transfers on fertility for white families (Moffitt, 1998), we find a larger effect for nonwhite families. One explanation for these differential results is that the EITC works through the labor market, and white and nonwhite families may have differential labor-force attachments and earnings distributions between spouses. For example, if nonwhite women live in low-income families that are more likely to be eligible for the EITC whether or not they work after childbirth, the EITC may be a more relevant source of income for them. Or, if nonwhite women have fewer marriage options, financial incentives for additional children may matter more (Jeff Grogger and Stephen G. Bronars, 2001).

IV. Conclusion

Using state-level birthrate data between 1990 and 1999, and exploiting the variation in state EITC programs, we find that very large increases in the income support provided by the EITC encouraged first births. Although not the intended consequence of the EITC program, these results suggest that government income transfers serve as a pro-natalist policy tool, although probably a blunt one, given that our estimated effects are small.

Our results cast some doubt on the assumption made by researchers of the effect of the EITC on labor supply that childbearing decisions are exogenous to EITC benefits. However, because the effects we estimate are small and concentrated among nonwhites, the magnitude of the bias introduced by this assumption is not likely to be large. Finally, although our results are generally consistent with empirical research on fertility effects of traditional welfare programs, one major difference is that the effects of the EITC appear to be larger for nonwhite families, rather than for white families. There is clearly much that remains to be understood about economic policy and fertility behavior with respect to race, and this is a topic for future research.

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