Labor Supply Effects of the Earned Income Tax Credit: Evidence from Wisconsin's Supplemental Benefit for Families with Three Children

Abstract - *We examine the effect of the Earned Income Tax Credit (EITC) on labor supply, comparing outcomes in Wisconsin, which supplements the federal EITC for families with three children, to outcomes in states that do not supplement the federal EITC. Relative to previous studies, our cross–state comparison examines a larger difference in EITC subsidy rates, more similar treatment and control groups, and a policy that has been in place longer. Whereas most previous research has found significant effects of the EITC on labor force participation, we find no effect.*

INTRODUCTION

The Earned Income Tax Credit (EITC) is the largest federal means–tested antipoverty program in the United States. Federal EITC tax expenditures in 1999 were $31 billion, almost as much as for the Food Stamp and Temporary Assistance for Needy Families (TANF) programs combined. Nearly 19 million federal tax returns claimed the EITC, while six million people participated in TANF and 18 million received food stamps (Council of Economic Advisers, 2001).

In 1999, families with two children could receive a refundable 40 percent federal income tax credit for each dollar of earned income up to $9,540. Taxpayers earning between $9,540 and $12,460 received the maximum possible federal credit of $3,816 (0.40 times $9,540). Beyond $12,460, each dollar of earnings reduced the EITC by 21.06 cents. In addition, 15 states offer supplemental tax credits based on the federal EITC (Johnson, 2000).

What does the U.S. government receive for this expenditure? Two goals are typically ascribed to the EITC: redistributing income to working poor families, and encouraging labor supply. While the first is unambiguously achieved, the second is theoretically and empirically less certain.

This paper measures the labor supply consequences of the EITC, using data from the 1990 and 2000 U.S. Censuses of Population and focusing on Wisconsin's supplement to the federal EITC for families with three or more children.

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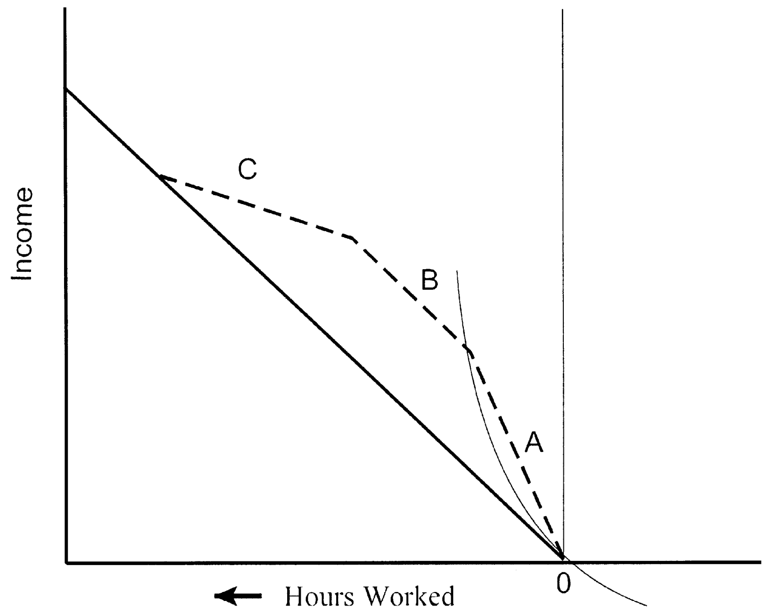
While numerous empirical studies of the EITC's labor market consequences exist, most examine changes in the federal rate over time. In many cases these changes are fairly small. For example, Eissa and Liebman (1996) investigate the 1987 expansion of the federal EITC from 11 to 14 percent.[1] In contrast, a three–child family in Wisconsin receiving the EITC will receive a tax credit that is 17 percentage points larger than that received by a comparable family in a state with no supplemental credit. More recent studies of EITC changes gradually phased in during the 1990s face the additional challenge of separating the effect of the EITC from general time trends. Finally, some previous studies rely on comparisons of women with and without children. In contrast, we use Wisconsin's EITC supplement and compare women with two and with three children, which we believe are more similar.

THEORY

The simplest theoretical effects of the EITC on labor supply, abstracting from other sources of income, are well known and can be summarized by the static labor supply diagram in Figure 1.[2] A person not working earns zero income, and the slope of the solid diagonal budget line is the wage. The dotted line in Figure 1 depicts the sum of earned income and the EITC, and so its slope along segment A is 1.4 times the wage rate. At earned income of $9,540 in 1999 the federal EITC is capped, and the dotted budget segment B runs parallel to the original budget line. Above

Figure 1. Simple Labor Supply Effects of the EITC



[1] The 1986 changes considered by Eissa and Liebman (1996) also included an increase in the maximum credit from $550 to $788 (in constant 1986 dollars) and a reduction in the phaseout rate.

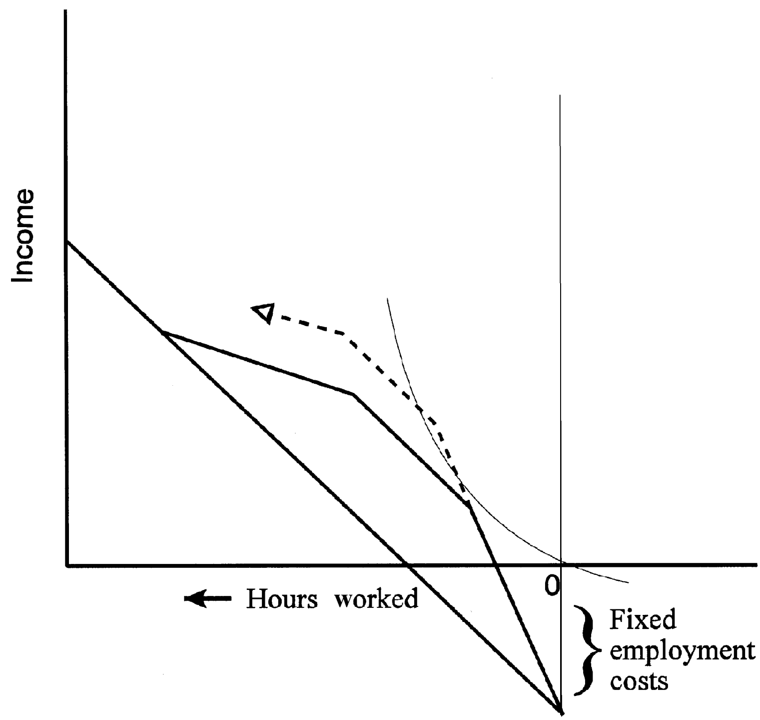
[2] This discussion assumes continuous labor supply, that people can choose to work as few or as many hours as they like, and, therefore, may overstate actual labor supply responses to employment subsidies like the EITC.

$12,460, tax credits are reduced by 21.06 cents for every dollar earned, and so the slope of the dotted line C is 0.7894 times the wage rate.

The effects of this tax program on labor supply can be broken into two parts: (a) participation, the decision whether or not to work at all, and (b) the number of hours to work. The first effect on participation is theoretically unambiguous. Any single parent who would choose to work in a world without the EITC will also choose to work in an otherwise identical world with an EITC.[3] On the other hand, some individuals not working in the absence of the EITC will prefer to work if there is an EITC. The indifference curve depicted in Figure 1 illustrates such a case. Note that in this simple case we assume workers can choose any number of hours, and only the initial subsidy rate matters for participation, not the size of the phase–in range.

Even a slight departure from the simplest static, continuous–hours case muddies this participation effect. If there are fixed costs of working (childcare, commuting, etc.), then the size of the phase–in range does matter for the decision to work. Figure 2 depicts the labor–leisure budget for a worker facing fixed employment costs. With the depicted indifference curve, the worker will choose not to work under the original (solid) EITC, and will chose to work under an expanded (dashed) EITC with the same subsidy rate but a larger phase–in range. Still, the

Figure 2. Labor Supply Market Effects of the EITC with Fixed Costs



[3] See Eissa and Hoynes (1998) for an analysis of the EITC and the labor supply of married couples, where this simple employment rule does not hold.

theoretical effect of the EITC on participation is unambiguously positive.

The second effect, on hours worked, is ambiguous even in the simplest case depicted back in Figure 1. Along segment A, the EITC increases the after–tax wage rate (by 40 percent for parents with two children), with offsetting income and substitution effects. At B, workers effectively receive a lump–sum transfer (of $3,816 in 1999 for families with two children), with pure income effects that unambiguously decrease desired work hours. At C, in the phaseout range of the EITC, workers receive a lump–sum transfer plus a decrease in the after–tax wage rate. Both the income and substitution effects unambiguously decrease desired hours.

Collectively, the labor market incentives of the EITC are mixed. The program has unambiguously positive theoretical effects on participation, but conditional on participation the program has largely negative effects on hours worked.

Finally, there are reasons to believe that all of these effects will be muted by complexities and lags in the tax code. Employment may only be offered in discrete quantity categories (e.g., part–time 20 hours per week or full–time 40 hours). Tax credits for income earned in one year are not received until taxpayers' EITC forms are filed in the following year.[4] Workers' limited understanding of the EITC may also reduce their responsiveness.[5] For all of these reasons, the size of the actual effect of the EITC on labor supply is an empirical question.

PREVIOUS STUDIES

Most existing work on the EITC relies on changes in the program's benefits, especially the 1987 and 1993 expansions. So as not to confound the effects of the EITC expansions with other changes in labor market conditions, these studies typically contrast changes in labor market behavior of eligible taxpayers before and after the EITC expansion to that of ineligible taxpayers. EITC eligibility is restricted to low–income households with earnings, and benefits are primarily targeted to households with qualifying children.[6] Eissa and Liebman (1996), for example, compare the labor supply of single mothers to that of single women without children. Meyer and Rosenbaum (2000) compare single mothers to single childless women, married women, and black men. These differences–in–differences strategies assume that any changes in labor market conditions that occur simultaneously with increases in EITC benefits do not affect single mothers and the comparison groups differently.

A second feature of these approaches is that the comparison groups typically have very high labor force participation. Of the women with no children that comprise Eissa and Liebman's primary control group, 95 percent were working at the time of the 1987 EITC expansion. Any general increase in labor force participation, therefore, is much more likely to be experienced by the women with children, of whom only 75 percent were working. Meyer and Rosenbaum restrict their sample to those with no more than a high school education, where labor force participation rates are lower and these ceiling effects are less important.

Another set of empirical papers estimates structural models of labor supply as a function of the after–tax wage, and then uses the EITC's various phase–ins

[4] Although some taxpayers can receive advanced EITC payments through their employers, only one percent of EITC recipients participated in this program in 1998 (Hotz and Scholz, 2001).

[5] See Ross Phillips (2001) for a discussion of workers' knowledge of the EITC.

[6] Beginning in 1994, households with no qualifying children were eligible for a small EITC. For example, in 1999 the maximum federal credit available to a childless household was $364.

and phaseouts to predict labor supply responsiveness to the EITC specifically. This approach assumes that administrative differences between the EITC and other tax provisions (such as the fact that the EITC credit is not realized until tax forms are filed the following year) do not affect labor supply responsiveness.

In a recent paper, Meyer and Rosenbaum (2001) estimate a structural model of labor supply in an effort to predict the effects of the EITC as well as other policies targeting low–income families, including AFDC, Food Stamps, Medicaid, and child care and training programs. In their preferred specification, they find that employment is responsive to changes in total taxes. They include all single women, allowing the effects of taxes to be identified through differences between women with and without children. They also estimate the model on a sample that includes only single mothers, with identification resting on differences across states and numbers of children. In this case, the estimated effects are substantially smaller, and are only significant using one of two data sources.

Neumark and Wascher (2001) estimate the effect of changes in the EITC on changes in employment, earned income, and official poverty status—the focus of their analysis. They match March CPS files from 1986 to 1995 in two–year pairings in order to observe changes in labor supply and earned income for individual families. Among families with no worker in the first year, increases in the EITC are associated with increased employment in the second year. However, among families with a worker in the first year, increases in the EITC are associated with declines in total hours worked.[7]

In general, these studies tend to find small and insignificant effects of EITC expansions on hours worked, but large positive effects on participation. For example, Eissa and Liebman (1996) find a statistically insignificant effect on hours worked, and a 2.8 percentage point increase in participation for single parents. Meyer and Rosenbaum (2001) conclude that 62 percent of the increase in single mothers' employment between 1984 and 1996 was due to the EITC, although the effect falls by half when the estimation includes only single mothers.

Hotz and Scholz (2001) argue that "probably the most powerful way to look at EITC labor market effects is to look at differences in labor market patterns for families with one and two–or–more children starting in the mid 1990s (when the discrepancies began to get large)." Hotz, Mullen and Sholz (2002) do that by examining the 1994 expansion of the federal EITC from 18.5 percent to 34 percent for families with one child, and from 19.5 percent to 40 percent for families with two or more children. They note that the difference between the EITC for one– and two–child families increased from one percentage point (19.5 minus 18.5 percent) to six percentage points (40 minus 34), and that controlling for other family characteristics this difference was associated with a six–percent increase in employment. The employment elasticity they estimate (1.2) falls at the high end of previous estimates.

A second study that exploits the change in family–size differences in EITC benefits is Grogger (2003). That paper uses data from the March CPS each year from 1979 to 2000. It regresses welfare use, employment status, and weeks worked on policy variables, time, state and family size dummies, and the federal EITC maximum credit, which varies over time and across family sizes starting in 1991. Grogger finds that a $1,000 increase in the maximum federal EITC credit increases participation by

[7] One puzzling aspect of Neumark and Wascher's results is that they sometimes find positive effects of state EITC credits on employment, but negative effects of the federal EITC on employment.

3.6 percentage points and increases weeks worked per year by 1.2 weeks.

WISCONSIN'S EITC SUPPLEMENT

We take the Hotz et al. (2002) and Grogger (2003) strategy one step further by looking at the even larger differences in EITC subsidy rates provided by Wisconsin's third–child supplement. Wisconsin provides the largest state EITC supplement, and the only one that differentially affects families with three children. Since 1995, Wisconsin has supplemented the federal EITC by four percent for families with one child, 14 percent for families with two children, and 43 percent for families with three children, where the Wisconsin supplement is calculated as a fraction of the federal credit.

The Wisconsin supplement rate (17.2 percentage points in this case—43 percent times 40 percent), is larger than the variation exploited by previous studies. Even when we subtract Wisconsin mothers' two–child benefit (45.6 percent) from Wisconsin's three–child benefit (57.2 percent), that difference (11.2 percentage points) is twice as large as the difference exploited by Hotz et al. (2002) and Grogger (2003a), and three times that of Eissa and Liebman (1996). If we focus on the maximum benefit, rather than the subsidy rate, the Wisconsin third–child supplement is comparable in magnitude to policy differences studied previously. The 1999 Wisconsin EITC benefit for three–child families was \$1,641 larger than for three–child families in other states, and \$1,107 larger than for two–child families in Wisconsin.[8]

Finally, we believe that an advantage of our approach is that we can use cross–sectional differences to identify the EITC effect. Prior studies have relied on changes in EITC subsidy rates that have been phased in over time, and eligible workers may take several years to learn about and respond to the new policies. Estimates relying on changes in differences over time likely have serial correlation in the error terms, which can generate spurious significant estimated effects of ineffective policies (Bertrand, Duflo and Mullainathan, 2002). By using cross–state variation, we avoid this bias. While we also compare 1990 and 2000 outcomes, our main analysis exploits the variation in EITC available to two– and three–child families in 2000. Moreover, Wisconsin's 43 percent EITC supplement has been in place since 1995, long enough for us to interpret data from the 2000 Census as an equilibrium response to the policy differences we study.[9]

The actual size of the Wisconsin supplement is depicted in Figure 3. Most papers on the EITC contain a figure similar to Figure 1, where the vertical axis is exaggerated for the sake of exposition. Instead, Figure 3 displays the actual federal EITC and the Wisconsin supplements on a set of axes that are not exaggerated. The Wisconsin supplement, though noticeable on the graph, is not dramatic. It seems plausible to us that the budget differences depicted in Figure 3, received as lump sums the following year, might have no effect on labor supply.

Because we need a large sample of low–income single mothers in Wisconsin, we use as data for this study the five–percent Public Use Microdata Sample (PUMS) of the 1990 and 2000 U.S. Censuses of Population. Our principal sample includes single mothers with a high school education or less.

[8] A single parent with three children in Wisconsin earning \$9,540 in 1999 would be eligible to receive the maximum federal EITC credit of \$3,816 (40 percent), and a state credit of another \$1,641 (43 percent of the federal credit). By comparison, Eissa and Liebman (1996) study the 1987 EITC expansion, when the maximum federal benefit increased by \$301. Hotz et al. (2002) and Grogger (2003a) study the 1990's family–size changes, where the difference between one– and two–child families increased from \$0 in 1990 to \$1,404 in 1996.

[9] We do, of course, appreciate the irony in touting the benefits of not using panel data.

Figure 3. Actual Federal EITC and WI 3rd Child Supplement

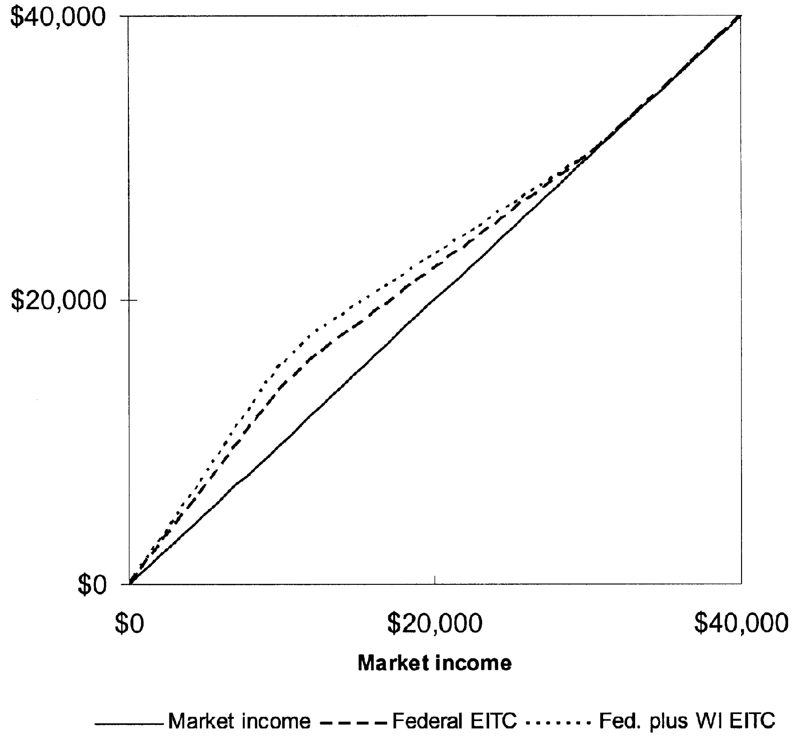


Table 1 begins to sketch the empirical strategy we use to identify the labor supply effects of the EITC, in a simple differences–of–means framework without controlling for other demographic characteristics of families or other state policy differences. Table 1A presents the employment rates for single women aged 19 to 44, with a high school education or less. (The construction of our sample is described in appendix table A1.) We show results for two measures of employment status: current employment, and whether the respondent worked at any time last year. As discussed above, the EITC has a theoretically unambiguous positive effect on the decision to work, but has ambiguous effects on total hours worked. The clearest test of the effect on the decision to work is, therefore, an analysis of any work in a given year, since a change in hours may result in changes in hours worked in a given week or in the weeks worked in a given year. (The EITC is calculated with reference to annual earnings.) On the other hand, a very high proportion of mothers work at some point during the year, potentially leaving less room to observe an EITC effect on that measure of employment.[10]

Low–income single mothers with two children in Wisconsin were eligible for up to a 45.6 percent tax credit on earnings (40 percent federal credit plus 14 percent state

[10] Meyer and Rosenbaum (2001, p. 1082) argue that a measure of current employment (in their case, whether a woman worked in the last week) is more policy–relevant, since it gives a measure of the proportion of all women working at a given time.

TABLE 1A
EMPLOYMENT PARTICIPATION OF SINGLE MOTHERS
(19–44 years old with 2 or 3 children and a high school degree or lower)

	Currently Employed			Worked Last Year		
	Two Children (1)	Three Children (2)	Difference (2) – (1)	Two Children (4)	Three Children (5)	Difference (5) – (4)
(1) Wisconsin (2000 Census)	0.709	0.648*	–0.060†	0.872	0.830	–0.041
	(0.016)	(0.026)	(0.031)	(0.012)	(0.021)	(0.024)
	n = 780	n = 330		n = 780	n = 330	
(2) States w/out EITC supplements (2000)	0.631	0.566	–0.065*	0.794	0.744	–0.050*
	(0.002)	(0.004)	(0.004)	(0.002)	(0.003)	(0.004)
	n = 39,640	n = 18,938		n = 39,640	n = 18,938	
(3) Difference (1) – (2)	0.078*	0.083*	0.005	0.078*	0.087*	0.009
	(0.016)	(0.027)	(0.031)	(0.012)	(0.021)	(0.024)
(4) Difference (1990 Census)	0.089*	0.122*	0.033	0.086*	0.094*	0.008
	(0.018)	(0.027)	(0.033)	(0.016)	(0.026)	(0.030)
(5) Diff–in–Diff–in–Diff (3) – (4)	–0.011	–0.039	–0.028	–0.008	–0.007	0.0003
	(0.024)	(0.038)	(0.045)	(0.020)	(0.033)	(0.039)

Standard errors are in parentheses.
*Difference of proportions is statistically significant at 5 percent; †Difference of proportions is statistically significant at 10 percent.
Source: U.S. Census of Population, 2000 and 1990, Public Use Micro Sample (5 percent).
Notes: "Other states" are those without EITC supplements: AL, AR, AZ, CA, CT, DE, FL, GA, HI, ID, IN, KY, LA, ME, MO, MS, MT, NC, ND, NE, NH, NM, NV, OH, OK, PA, SC, SD, TN, TX, UT, VA, WA, and WV.

supplement, where the state supplement is a fraction of the federal credit). By contrast, working single parents with three children in Wisconsin were eligible for up to a 57.2 percent credit (40 percent federal credit plus 43 percent state supplement). The first columns of Table 1A show that among single mothers with a high school degree or less, 71 percent of those with two children, and 65 percent of those with three children were working at the time of the 2000 census. Though the difference (–0.060) seems to suggest that the EITC supplement for three children discourages employment, we must recognize the unequal private work incentives for two–child and three–child families.

To provide a basis for comparison, the second row of Table 1A examines employment rates for similar families in states that do not supplement the federal EITC. Employment in those states is lower than in Wisconsin by about eight percentage points. More importantly, in states without an EITC supplement, mothers with three children are 6.5 percentage points less likely to work than mothers with two children. The slightly smaller difference between the employment rates in row (1), where having a third child increases the EITC from 14 to 43 percent of the federal level, and in row (2), where a third child adds nothing to the EITC, is consistent with Wisconsin's EITC supplement increasing employment for eligible parents with three children by 0.5 percentage points, a difference that is statistically insignificant.

Row (4) of Table 1A shows the same difference (between employment rates in Wisconsin and in states that in 2000 did not have an EITC supplement) for 1990, when families in Wisconsin were eligible to receive the same EITC as those in other states and there was no Wisconsin third–child supplement. In 1990, comparing low–education single mothers in Wisconsin to those elsewhere, employment rates were about nine percentage points higher among mothers with two children in Wisconsin, and 12 percentage points higher among mothers with three children in Wisconsin. This difference, three percent, indicates that when compared to mothers with the same number of children living in states where there was never a supplement, Wisconsin mothers with three children were more likely to work *prior* to the implementation of Wisconsin's third–child EITC supplement, though the difference is not statistically significant. The final row (5) shows the difference–in–difference–in–differences—a comparison that accounts for any time–invariant state–specific differences in two– and three–child families. The results are statistically insignificant (and the estimated difference is negative).

The second part of Table 1A presents a parallel analysis of employment, in this case considering the proportion of mothers who worked at any time during the previous year. Employment levels by this definition are higher, but the differences by number of children and state are similar. In Wisconsin, single mothers with three children were 4.1 percentage points less likely to work during 1999 than those with two children. In states that do not have an EITC supplement, mothers with three children were 5.0 percentage points less likely to work. Again the smaller difference between the employment rates in row (1) than in row (2) is consistent with Wisconsin's EITC supplement increasing employment for eligible parents with three children, but also by a small and statistically insignificant amount. The difference (0.008) is similarly small and insignificant in 1990, and the difference between the 1990 difference and the 2000 difference (0.0003) is also small and insignificant.

Table 1B conducts the same exercise for hours worked, for only those women who did work (women with positive weekly hours worked). Low–education working single mothers in Wisconsin with two children worked an average of 38.3 hours per

TABLE 1B
WEEKLY HOURS WORKED BY SINGLE MOTHERS
(19–44 years old with 2 or 3 children and a high school degree or lower)

	Two Children (1)	Three Children (2)	Difference (2) – (1)
(1) Wisconsin (2000 Census)	38.31 (9.62) n = 680	38.30 (9.95) n = 274	–0.02 (0.70)
(2) States without EITC supplements (2000 Census)	38.13 (9.76) n = 31,467	37.55 (9.99) n = 14,083	–0.58* (0.10)
(3) Difference (1) – (2)	0.18 (0.38)	0.75 (0.61)	0.56 (0.72)
(4) Difference (1990 Census)	–1.66* (0.48)	–2.21* (0.80)	–0.55 (0.93)
(5) Diff–in–Diff–in–Diff (3) – (4)	1.85* (0.61)	2.96* (1.01)	1.11 (1.18)

Standard errors are in parentheses.
*Difference of means is statistically significant at 5 percent.
Source: U.S. Census of Population, 2000 and 1990, Public Use Micro Sample (5 percent).
Notes: Includes only women working at the time of the Census (i.e., Weekly Hours Worked > 0). "Other states" are those without EITC supplements: AL, AR, AZ, CA, CT, DE, FL, GA, HI, ID, IN, KY, LA, ME, MO, MS, MT, NC, ND, NE, NH, NM, NV, OH, OK, PA, SC, SD, TN, TX, UT, VA, WA, and WV.

week; working single mothers with three children in Wisconsin also worked an average of 38.3 hours per week. For comparison, in row (2) of Table 1B, low–education single working women with two children in states without EITC supplements worked 38.1 weekly hours, while women with three children worked 37.6 hours. One interpretation of these results would be that without the third–child EITC supplement, working Wisconsin mothers would have worked fewer hours. Mothers in Wisconsin with three children worked the same as mothers with two children; in other states they work less. However as with the employment results in Table 1A, this difference–in–differences (+0.56 hours) is statistically insignificant.

The next row of Table 1B shows the difference in hours worked by mothers in Wisconsin and other states in 1990. In that year, mothers in Wisconsin with two children worked 1.7 hours less, while those with three children worked 2.2 hours less than mothers in other states. However, the difference (–0.55 hours) is small and statistically insignificant, as is the difference–in–difference–in–differences (1.1 hours) shown in the final cell.

The cross–sectional difference–in–difference results in Tables 1A and 1B do not control for other demographic differences between small and large families or between Wisconsin families and those of other states. Nor do they control for differences among states other than their EITC schedules. For this reason, we have also estimated versions of:

[1] $$Y_i = \alpha + \beta_1(WI) + \beta_2(3\ kids) + \beta_3(WI) * (3\ kids) + \beta_4 X_i + \varepsilon_i,$$

where Y_i is the outcome of interest (e.g., employment) for household i, *WI* is a dummy variable equal to one for households in Wisconsin, *3 kids* is a dummy variable for households with three children, and X_i are characteristics of households and states, including age, education, health, race, state unemployment rates, and state welfare policies. Including those other characteristics estimates the differential effect of Wisconsin's large third–child supplement while controlling for other important family and policy differences.

One concern with comparing Wisconsin to other states is that there may be some feature of Wisconsin policy or Wisconsin residents that makes labor supply behave differently in that state. By comparing the labor supply of women with two children to that of women with three children in Wisconsin and in comparison states, we ameliorate some of that problem. To the extent that Wisconsin's economy or policies influence mothers' employment in general, these state–specific effects will not bias our estimates of the difference in employment rates for mothers with two or three children. Even if features of Wisconsin's economy or policy environment differentially influence mothers' employment for two– and three–child families, so long as these state–specific and family–size–specific features are time–invariant, they will not bias our estimates that rely on a comparison of 1990 and 2000. However, there may be reasons beyond the EITC supplement why having a third child has different effects on labor supply in Wisconsin than in the comparison states. Two obvious candidates are state welfare policy and state child care policy.

On both counts, Wisconsin's policies potentially exaggerate the differential labor supply of low–income women with larger families. The first concern is accounting for differences in state AFDC and TANF benefit adjustments for larger families. Under AFDC, cash benefits increased with family size in every state, including Wisconsin. With the implementation of TANF, all but a few states continued to pay larger cash benefits to families with more children. In Wisconsin, however, TANF cash benefits do not depend at all on the number of children. If we were to find Wisconsin mothers with three children more likely to work in 2000, that result might have been due to the generosity of Wisconsin's EITC supplement or to the lack of a family–size adjustment in its cash welfare program.[11] To account for this, we include in our estimates of equation [1] a measure of maximum state AFDC/TANF benefits that varies with family size.

A second factor that may systematically alter the work incentives of families of different sizes is the availability and cost of subsidized child care. In the absence of subsidized care, families with more young children face higher work–related child–care expenses. Wisconsin offers relatively generous child–care subsidies and has high rates of subsidized child–care use. Since reducing the cost of child care should be particularly important for larger families, Wisconsin's child–care policy may further exaggerate our estimates of the effect of the EITC supplement on the labor supply of women with more children. Thus, in our estimates of equation [1], we include two measures of the availability and generosity of child–care subsidies: total expenditures on child care

[11] There is substantial variation in benefit levels across states, and in the absolute and proportional increase in benefits for larger families. Prior to TANF implementation, median AFDC benefits were $80 (and 21 percent) higher for families with three children than for those with two. Wisconsin AFDC benefits were $517 for a family of three and $617 for a family of four—a difference of $100 (and 19 percent). In 2000, eight states had maximum monthly TANF benefits for families with three children that were at least $100 higher than maximum benefits for those with two children. At the same time, 15 states had benefits for two– and three–child families that varied by $50 or less. While many states with higher overall benefit levels also had greater increases for larger families, the pattern was inconsistent: some states increased benefits for families with a third child by more than 25 percent, while others included adjustments of less than ten percent. Wisconsin is an extreme case in this regard. Under Wisconsin's TANF program, benefits do not vary with family size: most women qualify for a maximum cash payment of about $650, regardless of the number of children.

per poor child under the age of 13, and total preschool and Head Start spending per child under age six.[12]

While our model includes state and family–size–specific measures of cash benefits and child–care subsidies, we cannot be confident that our measures perfectly capture the influence of these policies on the labor supply of two– and three–child families in Wisconsin and other states. To the extent that we fail to fully account for these policies, our estimates of the labor supply effect of the EITC may be upwardly biased, since both child–care and cash–benefit policies would also create a greater incentive for families with more children to work at higher rates in Wisconsin than in most other states. In the end, however, none of this will matter. Wisconsin's lack of TANF family size adjustment and generous child–care policies will bias our findings in favor of measuring a large EITC effect on labor supply. But in a departure from the published literature to date, we find no EITC effect on either hours worked or participation.

Table 2 presents estimates of equation [1]. In column (1) we show the means and standard deviations of the control variables for the entire population of working and non–working single mothers. (Again we limit the sample to women with a high school degree or lower, with either two or three children, and living in either Wisconsin or a state without an EITC supplement.) Columns (2) and (3) contain estimates for a probit regression of current employment, defined as working at the time of the census. Columns (4) and (5) contain estimates for a probit regression of annual employment, defined as having worked in the year prior to the census. Finally, column (6) has results from an OLS regression of weekly hours worked among working women.

Turning to the first row of Table 2, we see that 1.9 percent of this sample lives in Wisconsin. The Wisconsin mothers are more likely to work and work more weekly hours than their non–Wisconsin counterparts. The probit coefficients suggest that mothers in Wisconsin are three to four percentage points more likely to be working than otherwise similar women in the comparison states—though the difference in employment at a point in time is not statistically significant.

Thirty–two percent of the sample has three children, with the remainder having two children. Women with a third child are less likely to be working at a point in time, or at any time during the year, and work fewer weekly hours. The probit coefficients suggest that having a third child reduces the probability that a single mother works by three to four percentage points.

The key coefficient is that on the interaction between the Wisconsin dummy and the third–child dummy, because only in Wisconsin does the state EITC supplement increase with the addition of a third child. The coefficient on hours worked in column (6), 0.142, is both small and statistically insignificant. This is unsurprising, given the ambiguous theoretical effects of the EITC on hours worked.[13]

The EITC does, however, have unambiguous theoretical effects on whether or not people work. The relevant interaction coefficients from columns (2) and (4) of Table 2 are small, statistically insignificant and *negative*: –0.044 for *currently employed*,

[12] We thank Marcia Meyers for providing these state–level measures of child–care expenditures. Because these data omit Washington DC and Wyoming, we omit those states from all of the analysis here. We have tried the analyses with DC and WY, and dropping the child care variables, with no discernable change. For a detailed analysis of child–care policy and single mothers' employment, see Bainbridge, Meyers, and Waldfogel (2002).

[13] We have also estimated column (6) as a Tobit, and as a simple OLS including both working and non–working mothers. In no case is the coefficient on the interaction term large or statistically significant.

TABLE 2
LABOR SUPPLY AND PARTICIPATION EFFECTS OF THE EITC
(Single mothers, 19–44 years old with 2 or 3 children and a high school degree or lower)

		Currently Employed		Worked Last Year		
	Means (1)	Probit Coefficients (2)	Marginal Effects (3)	Probit Coefficients (4)	Marginal Effects (5)	Weekly Hours Worked OLS[A] (6)
Wisconsin dummy	0.019	0.074 (0.051)	0.028	0.157* (0.061)	0.043	0.705† (0.399)
3 children dummy	0.323	–0.104* (0.012)	–0.040	–0.101* (0.013)	–0.030	–0.266* (0.106)
Wisconsin × 3 children dummy	0.006	–0.044 (0.087)	–0.017	–0.054 (0.102)	–0.016	0.142 (0.710)
Preschool age	0.441	–0.140* (0.013)	–0.053	–0.180* (0.014)	–0.052	–0.494* (0.110)
Age of mother	32.53 (6.407)	0.098* (0.009)	0.037	0.054* (0.009)	0.016	0.330* (0.076)
Age of mother squared	1,099.2 (412.79)	–0.0014* (0.0001)	–0.0005	–0.0009* (0.0001)	–0.0003	–0.004* (0.001)
Completed high school	0.598	0.449* (0.011)	0.172	0.452* (0.012)	0.135	1.047* (0.100)
Has health problem that disrupts work	0.170	0.123* (0.015)	0.046	0.026† (0.016)	0.008	–0.137 (0.122)
Hispanic origin	0.104	0.194* (0.023)	0.072	0.126* (0.024)	0.035	0.911* (0.203)
White	0.506	0.309* (0.012)	0.118	0.174* (0.013)	0.050	0.498* (0.104)
Other race	0.044	0.144* (0.028)	0.054	0.057† (0.030)	0.016	0.325 (0.249)

TABLE 2 (*continued*)
LABOR SUPPLY AND PARTICIPATION EFFECTS OF THE EITC
(Single mothers, 19–44 years old with 2 or 3 children and a high school degree or lower)

		Currently Employed		Worked Last Year		
	Means (1)	Probit Coefficients (2)	Marginal Effects (3)	Probit Coefficients (4)	Marginal Effects (5)	Weekly Hours Worked OLS[A] (6)
Immigrant	0.123	–0.057* (0.019)	–0.022	–0.0005 (0.0204)	–0.0001	0.341† (0.178)
State unemployment level	4.278 (0.759)	–0.087* (0.010)	–0.033	–0.126* (0.011)	–0.036	–0.431* (0.085)
State maximum AFDC benefit ($1,000)	0.382 (0.173)	–0.072† (0.039)	–0.027	–0.114* (0.043)	–0.033	–2.665* (0.351)
Federal child care spending per poor child ($1,000)	0.733 (0.302)	–0.067* (0.023)	–0.025	–0.068* (0.026)	–0.020	–0.611* (0.199)
Head start and pre–K spending per young child ($1,000)	0.248 (0.095)	–0.083 (0.063)	–0.032	–0.137* (0.069)	–0.040	–0.014 (0.522)
Constant		–1.28* (0.145)		0.456* (0.157)		34.25* (1.28)
N	59,688	59,688		59,688		46,504

Standard errors in parenthesis. (Heteroskedastic–consistent standard errors in (2) through (6)).
* Statistically significant at 5%. †Statistically significant at 10%.
[A]Column (6) includes only working women.

and –0.054 for *worked last year.* These suggest that having a third child in Wisconsin does not increase the probability of working, relative to having a third child in a state without an EITC supplement, and that Wisconsin's large EITC supplement has no effect on labor supply.

Table 3 summarizes estimates from several alternative specifications, all of which are consistent with our base results. The first row of Table 3 replicates the key coefficient from Table 2, for ease of comparison. The second row of Table 3 presents that same key coefficient from

TABLE 3
ALTERNATIVE SAMPLES

	Estimated Coefficients on Wisconsin × 3 Children Interaction		
	Currently Employed Probit (1)	Worked Last Year Probit (2)	Weekly Hours Worked OLS Regression[A] (3)
(1) Base Sample from Table 2 (mothers with two or three children, high school education or less)	–0.044 (0.087) [–0.017] n = 59,688	–0.054 (0.102) [–0.016] n = 59,688	0.142 (0.710) n = 46,504
(2) Any number of children (compares mothers with three or more to those with two or fewer); High school education or less	–0.076 (0.080) [–0.029] n = 107,686	–0.047 (0.093) [–0.013] n = 107,686	0.320 (0.652) n = 85,837
Mothers with two or three children			
(3) Includes states with and without EITC supplements (only WI has 3rd child supplement); High school education or less[B]	–0.063 (0.087) [–0.024] n = 77,325	–0.068 (0.102) [–0.021] n = 77,325	–0.007 (0.707) n = 59,711
(4) With income is less than 300% of the federal poverty line (no education restriction)	–0.026 (0.068) [–0.009] n = 96,021	–0.100 (0.082) [–0.026] n = 96,021	0.164 (0.554) n = 79,001
(5) With income is less than 200% of the federal poverty line (no education restriction)	–0.019 (0.074) [–0.007] n = 80,352	–0.080 (0.087) [–0.023] n = 80,352	0.016 (0.634) n = 63,694
(6) With income is less than 100% of the federal poverty line (no education restriction)	–0.112 (0.101) [–0.045] n = 45,618	–0.123 (0.109) [–0.045] n = 45,618	0.283 (1.134) n = 30,582
(7) Include state fixed effects (and drop time–invariant state variables); High school education or less	–0.095 (0.093) [–0.037] n = 59,688	–0.132 (0.107) [–0.040] n = 59,688	0.945 (0.759) n = 46,504
(8) Pooled 1990 and 2000 Censuses. Coefficient on WI × 3 children × 2000 interaction; High school education or less	–0.187 (0.122) [–0.074] n = 111,694	–0.119 (0.136) [–0.040] n = 111,694	0.715 (1.162) n = 79,964

Heteroskedastic–consistent standard errors in parentheses. Marginal effects for probit models in square brackets.
[A]Column (3) regression is conditional on positive hours worked.
[B]States with EITC supplements are CO, IL, IA, KS, MA, MD, ME, MN, NJ, NY, OR, RI, VT, and WI. Washington DC also has an EITC supplement, but is excluded here.

a specification including single mothers with any number of children. (The third–child dummy here is for mothers with three *or more* children.) This almost doubles the sample size, while blurring somewhat the distinction between the two groups. The larger sample size does not yield a more statistically significant coefficient in any of the specifications, and the coefficients remain small. Next, we include in the comparison groups states that do have an EITC supplement (though none differentiates between two–child and three–child families). Again, the effects are small and insignificant.

Following Neumark and Wascher (2001), we also estimated our model on subsamples of women with low incomes, rather than with low education. Rows (4) through (6) of Table 3 show the coefficient estimates for the effect of the interaction of Wisconsin residence and having three children for subsamples with incomes below 300 percent, 200 percent, and 100 percent of the federal poverty line. Unlike Neumark and Wascher's (2001) results, we do not find larger employment elasticities for those with lower incomes. Rather, the coefficients all remain small, statistically insignificant and negative.

In row (7) of Table 3, we estimate a version of our basic specification that includes state fixed effects. The prior versions have only a Wisconsin dummy, and variables that describe states (unemployment rate, maximum AFDC benefits, child–care spending, and head–start and pre–K spending). All of these drop out of the fixed effects version, except for AFDC benefits, which vary by family size. The key coefficients in row (7) remain small and statistically insignificant.

Finally, row (8) of Table 3 estimates the three–way difference on a pooled sample of the 1990 and 2000 censuses. We include dummy variables for Wisconsin, three children, and the 2000 census, three two–way interactions between each pair of dummies, and the three–way interaction between all three. Row (8) reports the coefficient on this three–way interaction, analogous to the difference–in–difference–in–differences reported in Tables 1 and 2. Here again the estimates suggest no significant effects on participation or hours.

COMPARISONS WITH PREVIOUS RESULTS

Eissa and Liebman (1996) found that women with children increased their employment after the 1987 EITC expansion by 1.9 percentage points. That was in response to an increase in the federal EITC from 11 to 14 percent, or a 2.7 percent increase in total labor compensation. By contrast, we find a (statistically insignificant) *decrease* in employment of about –1.6 percentage points in response to an eight–percent increase in labor compensation.[14] Roughly speaking, Eissa and Liebman (1996) estimate a statistically significant employment elasticity of 0.70, while our insignificant point estimate of that same elasticity is about –0.21.[15] Meyer and Rosenbaum (2001, 1089–92) estimate that a $1,000 decline in annual taxes increases employment by 2.7 to 4.5 percentage points, implying elasticities of 0.83 to 1.07.

Grogger (2003) focuses on the maximum EITC benefit, and estimates that a $1,000 increase in that maximum increases

[14] Women with two children in Wisconsin receive a total EITC benefit of 0.456 percent (federal credit of 0.4 plus Wisconsin's 14–percent supplement). Women with three children receive 0.572 percent (the federal credit plus Wisconsin's 43–percent supplement). The ratio 1.572/1.456 equals 1.08.

[15] Hotz and Scholz (2001) report elasticities with respect to net incomes, which relies on an assumption about the typical work hours of a labor market entrant. We report elasticities with respect to net wages, which is equivalent so long as hours worked are fixed and entrants have not reached the EITC cap. Our calculation is –0.017/(1.572/1.456–1)= –0.21. For Eissa and Liebman (1996) the equivalent calculation is 0.019/(1.14/1.11–1).

employment by 3.2 percentage points. Hotz et al. (2002) estimate that same $1,000 increases employment by five percentage points. Like our study, both Hotz et al. (2002) and Grogger (2003a) use reduced form approaches, and cannot say whether their effects come from the increase in maximum benefits or the increase in subsidy rates. Nevertheless, we find that the Wisconsin third–child supplement, which amounts to a $1,107 increase in the maximum benefit relative to two–child families, did not change participation or hours worked.

Our analysis compares the employment of single mothers with two and three children. We include measures of child–care subsidies and welfare benefits because these are the two state policies that we particularly suspect would differentially affect families of different sizes, since child–care costs generally increase with the number of children and since there is substantial state variation in the extent to which welfare benefits vary with family size. In other respects we expect that single–mother families with two and three children are more comparable than, for example, single women with and without children. A comparison of basic demographic characteristics of women by maternal status and number of children confirms this expectation (see appendix Table A2). In this context it is noteworthy that when Meyer and Rosenbaum (2001) restrict their analysis to single mothers, relying on variation across states and number of children to identify the effects of taxes, the estimated effects of the EITC are smaller and only significant for one of the two samples they use.

There are several reasons why our results might be expected to differ from previous estimates. We rely on a comparison of mothers with two and three children, and argue that there are less likely to be other unmeasured differences between these groups than, for example, between women with and without children. Meyer and Rosenbaum (2001) also find smaller (or no) effects when they compare mothers with different numbers of children. Thus, our results may correctly measure the general failure of the EITC to increase participation or hours worked. On the other hand, the employment decisions of mothers with three children may be less sensitive to the EITC. It may be, for example, that the non–pecuniary costs and benefits of employment are more important to mothers with larger families.

Finally, note that we use two measures of employment—current employment and worked last year. These are similar to the measures used by Meyer and Rosenbaum (2001), who point out that any employment in the last year should provide a sharper test of the theory, since the EITC has a theoretically unambiguous effect on ever working in a tax year, but an ambiguous effect on employment in any given week during the year. Despite the theoretical predictions, we estimate statistically insignificant effects using both definitions. One possible explanation is that annual employment rates are so high, particularly in Wisconsin, that it is difficult for the EITC to have a discernible effect. Meyer and Rosenbaum generally find larger effects for their annual measure, and when they restrict their analysis to single mothers, only find statistically significant effects using the annual measure.[16]

CONCLUSION

A key goal of the EITC is to redistribute income to working poor families. In practice, the EITC is an important income

[16] In our analysis, both measures are for the same sample and from the same data source. Meyer and Rosenbaum (2001) use a measure of work in the last week from the larger Outgoing Rotation Group File of the CPS, and a measure of annual employment from the March CPS.

source for many vulnerable families, including many single–mother families making the transition from welfare to work under recent welfare reforms (Johnson, 2000; Cancian, Haveman, Meyer and Wolfe, 2002). To many analysts, the EITC is preferable to other programs aimed at low–income families because it is tied to work. For families with a single worker earning low wages, the more hours they work, the greater is their EITC. Thus, given its basic structure, the EITC unambiguously targets resources to low–income working families.

A less certain advantage of the EITC is its ability to increase labor supply. We use the 1990 and the 2000 Censuses of Population to examine the labor market consequences of the EITC by comparing the labor market behavior of eligible parents in Wisconsin, which supplements the federal EITC for families with three children, to the labor market behavior of otherwise similar parents in states that do not supplement the federal tax credit. We find no evidence of increased employment: for all of our samples and specifications, the effect of the EITC on employment appears to be small and statistically insignificant.

Our conclusion that the EITC has no effect on participation or hours worked departs from previous published results, despite the fact that we have a larger sample size (the five–percent sample of the 2000 Census), identify the effect of the EITC using larger subsidy rate variation (Wisconsin's 43–percent third–child supplement), and use treatment and control groups that are more similar in other dimensions (low–education single women with two or three children). However, a finding of no EITC effect on labor supply is not altogether surprising. The program is complex, its subsidies are paid out long after the eligible labor is supplied, many workers are not even aware of its existence, and jobs may not have flexible hours. In the end, we should not condemn the EITC's possible failure to stimulate participants' labor supply; rather, we should credit its ability to support low–income working families without deterring participants' labor supply.

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Appendix

TABLE A1A
SAMPLE CONSTRUCTION FROM 2000 CENSUS

Sample Criterion	Observations
Single female household heads	1,471,967
Age 19–44	558,896
Exclude DC and WY	555,338
With a high school diploma or lower	223,588
Living in Wisconsin or a state without an EITC supplement	171,069
With two or three dependent children	59,688

Source: 5% Public Use Microsample (PUMS) of the 2000 Census of Population.

TABLE A1B
SAMPLE CONSTRUCTION FROM 1990 CENSUS

Sample Criterion	Observations
Single female household heads, Age 19–44	473,181
Exclude DC and WY	469,716
With a high school diploma or lower	205,626
Living in Wisconsin or a state without an EITC supplement	153,189
With two or three dependent children	52,006

Source: 5% Public Use Microsample (PUMS) of the 1990 Census of Population

TABLE A2
DEMOGRAPHIC CHARACTERISTICS OF HOUSEHOLD HEAD
(Sample: female household head, 19–44 years old, with a high school degree or lower)

	With No Children in All States (1)	With 2 Children in Wisconsin (2)	With 3 Children in Wisconsin (3)	With 2 Children in States with No EITC Supplement (4)	With 3 Children in States with No EITC Supplement (5)
Age	33.60	32.75	33.18	32.57	32.43
	(8.17)	(6.71)	(5.69)	(6.64)	(5.88)
High school degree	0.713	0.709	0.652	0.623	0.540
Unhealthy	0.187	0.128	0.179	0.170	0.172
Hispanic origin	0.055	0.024	0.055	0.096	0.126
Black	0.215	0.185	0.251	0.328	0.390
White	0.679	0.745	0.624	0.533	0.438
Other	0.051	0.046	0.070	0.043	0.046
Immigrant	0.104	0.028	0.048	0.115	0.147
Mother's 1999 earnings ($)[A]	20,477	18,018	18,687	16,576	15,162
	(20,313)	(16,753)	(25,171)	(19,395)	(20,132)
Observations	70,388	780	330	39,640	18,938

Note: Standard deviations of continuous variables in parentheses.
Source: 5% Public Use Microsample (PUMS) of the 2000 Census of Population.
[A]Mother's 1999 earnings conditional on working.