The Rise of Working Mothers and the 1975 Earned Income Tax Credit

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Abstract

The rise in female employment over the twentieth century radically changed the U.S. economy and the role of women in society. Census and Current Population Survey time-series data show a trend break for the employment of mothers beginning in the mid-1970s, but not for women without children. In the first systematic study of the 1975 introduction of the Earned Income Tax Credit (EITC), I show that this program led to a 6-percent rise in the employment of mothers and can help explain why the U.S. has a high fraction of working mothers despite having little childcare subsidies or maternity leave policies.

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The fraction of working mothers in the U.S. is strikingly high by international standards (Costa 2000; Olivetti and Petrongolo 2015) and historical standards (Goldin 1990). This is puzzling since the U.S. has little childcare subsidies, maternity leave, or other public policies that make working easier for mothers. Although female employment began rising after World War II, mothers did not enter the labor force in large numbers before the 1970s. Figure 1 shows that after falling for 70 years, the relative employment of mothers – compared to women without children – began to rapidly increase in the 1970s. Although rising education and decreasing fertility explain a small part of this trend, the explanation for the rise of working mothers has remained understudied.

This paper investigates the role of the EITC in the rise of working mothers. The EITC provides a significant subsidy for low-income working adults with children. Using Current Population Survey March Annual Social and Economic Supplement (CPS) data spanning 1969 to 1985, a dynamic difference-in-differences approach shows that the EITC led to a 6 percent (or 3.3-percentage-point) increase in the maternal employment of all 16- to 45-year-old women. I show that the effect is even larger for subgroups of mothers more likely to have earnings below the EITC limit, including mothers with relatively low education.

Figure 2 motivate the DD estimates by showing that the relative employment of mothers began to increase after 1975. From 1969 to 1975 the unadjusted employment gap between mothers and women without kids was stable at 24-percentage points. Between 1975 and 1979 the relative employment of mothers steadily increased and the gap narrowed to 18-percentage points, where it remained through 1985. Estimates in this paper suggest that the EITC explains about a third of the 1975 to 1980 rise in maternal employment and a quarter of the overall rise in female employment.[1] I find that nearly a million less mothers would have been working by 1985 if the EITC never existed.[2]

The EITC literature has consistently found that married women decrease their labor supply in response to the EITC (Dickert, Houser, Scholtz 1995; Ellwood 2000; Eissa and Hoynes 2004). However, these average effects mask substantial heterogeneity. This paper shows that the employment response of married women with low-earning spouses is positive and is similar in magnitude to that of single women. The EITC treatment effect on married women is negatively correlated with her spouse's

[1] Author's calculation based on 16 to 45 year old females in March CPS data. 57 percent of these women are mothers; maternal employment increases from 53 to 62 percent and female employment increases from 61 to 68 percent between 1975 and 1980.

[2] This is larger than the 164,000 and 350,000 mothers induced into employment by subsequent EITC expansions in 1986 and 1993. 164,000 is cited in Eissa and Liebman (1996) and reflects the authors' DD estimate of 2.8 percentage points among single mothers. Meyer and Rosembaum (2001) find a 7.2 percentage point increase in the employment of single mothers between 1984 and 1997, and about 5 percentage points of this is due to the 1993 EITC expansion, which implies the 350,000 figure. There are 52.833 million 15- to 44-year old women in the 1980 Census, 47 percent of which are mothers according to the CPS, so about 800,000 mothers were induced into employment because of the 1975 EITC

earnings. I show that this pattern is also evident during the 1986 and 1993 EITC expansions and is not particular to the 1975 program introduction. Finally, married women with spouses earning above the EITC limit were not eligible for the EITC and I estimate a precise treatment effect of zero for this placebo group.

The 1975 EITC was a 10 percent wage subsidy worth up to about $1,800 (in 2013 dollars) for low-income families. Although these benefits did not vary geographically, the employment response of mothers did. One reason relates to cost of living. On average low cost of living areas pay lower wages meaning that mothers are more likely to be under the EITC limit. Furthermore, each dollar of EITC benefits has more purchasing power in such areas (Fitzpatrick and Thompson 2010). A second explanation for the geographically varying response relates to the social stigma against working women. Using American National Election Surveys (ANES) data and General Social Survey (GSS) data, I show that state-level stigma can account for a significant part of the regional variation in the maternal employment response to the EITC, even when cost of living is controlled for.

In addition to extensive-margin employment, I also show that the EITC had positive effects on annual earnings and weeks worked. These increases are largest for groups of mothers most likely to have earnings below the EITC limit, and quantile regression shows that the middle two deciles of the income distribution saw the largest increases in annual earnings and week worked.

This is the first paper to systemically study the introduction of the EITC in 1975 and I show that this program had big effects on the labor force. The EITC helped lead to the rise of working mothers and can help explain why the U.S. has long had such a high fraction of working mothers despite having little childcare subsidies or maternity leave policies.

I. EITC Background

The EITC arose from – and was partially a response to – President Lyndon Johnson’s War on Poverty in the 1960s, where numerous laws and programs were enacted to increase the opportunity and well-being of low-income Americans (see Bailey and Danziger 2013 for a review). Although these programs succeeded in improving health and decreasing poverty (Almond, Hoynes, and Schanzenbach 2011; Hoynes, Page, and Stevens, 2011; Bailey and Goodman-Bacon 2015; Goodman-Bacon 2015), they also discouraged work (Moffitt 1992; Hoynes 1996; Hoynes and Schanzenbach 2012). After President Johnson left office, the policy debate turned to a guaranteed minimum income, which had support from economists such as Milton Friedman (1962) and James Tobin (1969).[3] In 1970 the U.S. House of Representatives passed the Family Assistance Plan with the backing of President Richard Nixon,[4] and would have guaranteed an annual income of $3,100 (2013 dollars) for each parent and $1,800 for each child; a family of four would have received $9,800. Benefits would phase out at a rate of 50 percent when household earned income surpassed $4,400 (Trattner 2007, p. 315). However, the U.S. Senate never passed the Family Assistance Plan because of disagreements about how generous the program should be and concerns about potential work disincentives. An alternative program was introduced by Louisiana Senator Russell Long in 1972 called the Work Bonus Plan. A version of this bill was eventually passed as the Earned Income Tax Credit and signed into law by President Gerald Ford on March 29, 1975 (see Liebman 1998 and Ventry 2000 for a detailed history).

Since 1975, the EITC has been expanded a number of times[5] and has consistently been supported and expanded by both the political left and right. The EITC has since grown into one of the largest anti-poverty program in the U.S. In 2013 the EITC paid out $66 billion to 28 million individuals and lifted 6.5 million people out of poverty including 3.3 million children (Center on Budget and Policy Priorities 2014).

A common misconception in the EITC literature is that the 1975 program introduction was small.[6] Appendix Figure A1 shows how the maximum benefits and income limit have evolved over four decades, for adults with one child. Appendix Figure A1 shows that the maximum benefits were actually quite high and the income cutoff was almost as high in 1975 as it was in the 2000s. Potential benefits and the income limit jumped around until the EITC schedule was pegged to inflation in 1986.

The 1975 EITC was a refundable-tax credit that provided a 10 percent income subsidy to working parents with annual earnings up to $18,000 (2013 dollars). The EITC also provided benefits to working parents with earnings above $18,000, but these benefits decreased at a rate of 10 percent and

[3] Juliet Rhys-Williams (1943) is considered to be the first to outline this type of program.

[4] See New York Times April 17, 1970.

[5] In 1979 a plateau region was added to the EITC schedule; in 1986 the phase-in rate was increased to 14 percent and the EITC parameters were indexed to inflation; in 1990 additional EITC benefits became available to parents with two or more children; in 1993 benefits were extended to adults without children (though at a low rate of 7.65 percent); between 1993 and 1996 the phase-in rate increased to 34 percent and 40 percent for women with one and women with two or more children; in 2003 the plateau region of the EITC schedule was extended to married couples to decrease the "marriage penalty"; and in 2009 additional EITC benefits became available to parents with three or more children.

[6] This is one reason that this program introduction has never been studied (according to a conversation I had with Nada Eissa). The 1975 EITC is only ever mentioned in passing as the following quotes indicate. "The EITC began in 1975 as a modest program aimed at offsetting the social security payroll tax for low-income families with children. There were few changes in the credit in the first 10 years" (Eissa and Hoynes 2004). "Between its beginning in 1975 and the passage of the Tax Reform Act of 1986, the EITC was small, and the credit amounts did not keep up with inflation" (Meyer and Rosenbaum 2001). "The earned income tax credit began in 1975 as a modest program aimed at offsetting the social security payroll tax for low-income families with children. After major expansions in the tax acts of 1986, 1990, and 1993, the EITC has become a central part of the federal government's antipoverty strategy" (Eissa and Liebman 1996).

reached zero for earnings above $36,000. Appendix Figure A2 illustrates the budget constraint for EITC-eligible adults. To be EITC eligible, married couples had to file taxes jointly, and families had to have at least one child living in their home for more than half the year ("residency test") that was under 19, or under 24 if a full-time student, or any age if disabled. Until 1990 taxfilers also had to demonstrate that they provided at least half the costs of maintaining the household ("support test"), implying that cash and in-kind public assistance had to be less than half of the household budget (Holtzblatt 1991; Holtzblatt, McCubbin, and Gillette 1994).[7] At this time there were no additional EITC benefits for having more than one child, and benefits did not vary at the state level.

The EITC has raised maternal employment (Dickert, Houser, and Scholtz 1995; Eissa and Liebman 1996; Meyer and Rosenbaum 2000, 2001; Hotz and Scholz 2006; Eissa, Klevin, and Kreiner 2008), earnings (Dahl, DeLeire, and Schwabish 2009), and health (Evans and Garthwaite 2014). The EITC has also decreased rates of poverty (Scholz 1994; Neumark and Wascher 2001; Meyer 2010; Hoynes and Patel 2015) and has helped children of EITC recipients, improving infant health (Hoynes, Miller, and Simon 2013), test scores (Chetty, Friedman, Rockoff 2011; Dahl and Lochner 2012), educational attainment (Manoli and Turner 2014; Bastian and Michelmore 2016), health (Averett and Wang 2015), as well as employment and earnings (Bastian and Michelmore 2016).[8] See Nichols and Rothstein (2015) for a recent review of the EITC literature.

II. Theoretical Framework: The EITC Had an Important Role in the Rise of Working Mothers

This paper is the first systematic study of the EITC introduction in 1975 and investigates the role of this program in the rise of working mothers. The 1975 EITC is a wage subsidy for low-income parents and should have a positive effect on the employment of mothers and no effect on women without children.[9] Intuition for this can be formalized in a simple labor-leisure model where a woman's decision to work is a function of her consumption c_i, leisure L_i, and cost of working $g_i(.)$.

$$V(c_i(.), L_i, g_i(.)) = \max[f(c_i(.), L_i, g_i(.))] \quad (1)$$

[7] I am not able to account for the residency test or support test in the data and ignore these relatively minor issues.

[8] The EITC may also have had a few unintended consequences, driving down the pre-tax equilibrium wages of low-skill workers (Leigh 2010; Rothstein 2010; Neumark and Wascher 2011) and discouraging marriage (Dickert-Conlin and Houser 2002; Herbst 2011. Although Bastian (2016) finds evidence that the EITC has increased marriages and fertility).

[9] This implicitly assumes that employers do not lay off women in the control group and hire their treatment group counterparts. I assume this throughout, but see Neumark and Wascher (2011) for a discussion.

$c_i(.)$ is determined by her after-tax wage w_i and non-labor income n_i (e.g. spousal wage, welfare benefits), while $g_i(.)$ is determined by childcare and opportunity costs, as well as psychological costs that come from a social stigma against women working (Fortin 2005; Charles, Guryan, and Pan 2009) or from her own beliefs about gender roles and identity (Akerlof and Kranton 2000). A woman will work if the additional consumption benefits of working outweigh the costs of working, and working can be thought of as either a binary or continuous decision.

To fix ideas, consider a less general form of equation (1):

$$(2)\quad V(c_i(.), L_i, g_i(.)) = \max_{l,L\in[0,1]} \{\alpha \log[l * w_i(k_i, COL_i) + n_i] + \beta \log[L] - \gamma_{s,t}[l * (k_i + S_s)]\}$$

Where α and β weight the relative importance of consumption and leisure, and $\gamma_{s,t}$ weights the cost of working, which can vary by state and year, and increases with work hours l and number of kids k_i. Wages $w_i(k_i, COL_i)$ are a function of children since having kids is a requirement of the EITC and also depend on geography and cost of living. See Appendix C for first order conditions, comparative statics, and conditions required for higher wages w_i to induce an increase in work hours l.

To estimate the effect of the EITC on maternal employment, I compare the employment rates of mothers and women without children in the years before and after 1975, with a dynamic difference-in-differences (DD) approach. I approximate equation (2) with the following probit model that models the probability that a woman works based on a number of observable characteristics:

$$(3)\qquad P(E_{i,s,t}) = \Phi(\beta_1 Mom_{i,s,t} + \beta_2 Post_{i,s} + \beta_3 Mom\ x\ Post_{i,s} + \beta_4 X_{i,s,t} + \gamma_s + \delta_t + \epsilon_{i,s,t})$$

where *i* indexes individuals, *s* states, and *t* years. $E_{i,s,t}$ is binary for whether a woman is employed,[10] $Mom_{i,s,t}$ is binary for whether a woman has any children, $Post_{i,s}$ is binary for whether the year is after 1975 and the EITC exists, and $Mom\ x\ Post_{i,s}$ is the DD variable of interest. The EITC treatment effect is estimated by β_3. $X_{i,s,t}$ contains a set of controls – described in detail below – that vary at the individual, state, and year level. γ_s and δ_t account for state[11] and year fixed effects, and $\epsilon_{i,s,t}$ is an

[10] Employed is defined as having positive earned income. Appendix Table A1 shows that results are robust to various definitions of employed or labor-force participation.

[11] Prior to 1977 (working year 1976) March CPS did not uniquely identify all states of residence. Groups of smaller states were bunched together. In order to provide a balanced panel of consistently defined geographical units, states are merged into the smallest consistent unit. This yields the following 21 state-groups that I refer to as "states": CA, CT, DC, FL, IL, IN, NY, NJ, OH, PA, TX, and AL-MS, AK-HI-OR-WA, AR-LA-OK, AZ-CO-ID-MT-NE-NM-NV-UT-WY, DE-MD-VA-WV, GA-NC-SC, KY-TN, IA-KS-MN-NE-ND-SD, ME-MA-NH-RI-VT, and MI-WI.

idiosyncratic error term. Since probit is a nonlinear model, all results shown are average-marginal effects and each coefficient is measured in percentage points.[12] Standard errors are calculated using the delta method, robust to heteroskedasticity, and clustered at the state-level to account for any correlation of state-level unobserved characteristics.[13] March CPS weights are used throughout, although unweighted results are very similar.

III. March CPS Data and Necessary Conditions for Difference-in-Differences

To estimate the effect of the EITC on maternal employment, I use 1972 to 1981 March Current Population Survey data[14] (Ruggles et al. 2015) – corresponding to employment in 1971 to 1980 – and the sample of *all* 16- to 45-years-old women. The treatment group consists of mothers and the control group consists of women without children.[15] Table 1 shows summary statistics for the whole sample in column 1, as well as for the treatment and control groups separately in columns 2 and 3. Women in the sample average 29 years old with 12 years of education, 11 percent are Black, 8 percent are Hispanic, 64 percent are employed and have average annual earnings (in 2013 dollars) of $12,262 (or $19,282 conditional on working), and 87 percent have income below the EITC limit (or 50 percent conditional on working).[16] Table 1 also shows that mothers tend to be older, have less years of education, are more likely to be non-white, and are less likely to work.

Figure 2 shows unadjusted CPS employment trends for mothers and women without children between 1969 and 1985, and previews the regression-adjusted DD results. From 1969 to 1975 the employment gap between mothers and women without kids was stable at 24-percentage points. Between 1975 and 1979 the relative employment of mothers steadily increased and the gap narrowed to 18-percentage points, where it remained from 1979 through 1985. Although the level of employment differed for these two groups, and Table 1 shows that their average characteristics differ, Figure 2 shows that the employment trends of these two groups were parallel *before* 1975.[17] Parallel trends are a

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[12] Results are nearly identical for a logit and are slightly larger for OLS; see discussion below and Appendix Table A2.

[13] I only observe 21 states-groups; such few clusters may inflate the standard errors (Cameron, Gelbach, and Miller 2008, 2011; Angrist and Pischke 2008, p. 319), but results are highly statistically significant and I ignore this issue.

[14] Although the CPS May Outgoing Rotation Group is in some ways more attractive than the March CPS (Lemieux 2006), these data do not extend before 1973. Furthermore, results in this paper are primarily based on employment at the extensive margin and concerns about measuring hourly wages are largely tangential.

[15] Since I do not observe tax filing, I assume all single women file taxes as household head, all married couples file joint taxes, and all family members under 19 (or 24 if a full-time student) are dependent children. I also treat all subfamilies within a household as separate tax-filing units and potentially EITC-eligible.

[16] Appendix Figure A3 shows the income distribution of women in this sample: 25 percent have zero earnings, 58 percent earn within the EITC-eligible range, and about 17 percent earn above the original EITC limit.

[17] Figure 2 also shows parallel trends *after* 1979, which may suggest that it took mothers a few years to fully respond to the

necessary condition required to claim that the EITC *caused* an increase in maternal employment.

Another necessary condition for DD is that no other confounding events or policies affected the relative employment of mothers. Although the 1970s were a period of inflation, oil and food price shocks, and two recessions, in the following paragraphs I do not find any confounders that would threaten the empirical strategy of this paper.

The first oil shock began in October 1973 when the Organization of Arab Petroleum Exporting Countries proclaimed an oil embargo against the West in response to support for Israel in the Yom Kippur War against Egypt. This led to a quadrupling of oil prices by March 1974, double-digit inflation and increases in food prices, and a US recession marked by stagflation lasting from November 1973 to March 1975. A few years later the second oil shock began when oil production sharply decreased in Iran due to the Iranian Revolution. This preceded the double-dip recession that occurred between January 1980 and July 1980, and between July 1981 and November 1982.

This first recession ended around the same time that the EITC was introduced. If the relative employment of mothers was strongly pro-cyclical, perhaps this could have led to an increase in relative employment of mothers. However, Figure 2 shows that this did not occur after the double-dip recession of the early 1980s, which suggests that unless the 1973 to 1975 recession was somehow unique, it is likely not responsible for the increase in working mothers. Theoretically, if prices permanently jumped in 1975 then the income effect could make families worse off and induce many mothers to begin working.[18] However, even if this happened it would take a complex story to explain how this could affect mothers but not women without children. Since the price shocks of the 1970s were temporary, they should not have led to a permanent change in the relative employment of mothers.

Another threat to identification operating through the income effect could involve cuts to public programs that differentially affect mothers. This has often been discussed in the context of the 1993 EITC expansion when welfare was also being dramatically cut back, and is a reason that the 1970s provides an arguably cleaner policy environment. Appendix Figures A5, A6, and A7 show trends in the number of family recipients, total payments, and average benefits per recipient for Women, Infants, and Children (WIC),[19] Food Stamps[20], and Aid to Families with Dependent Children. Fortunately, I do not

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employment-incentives of the 1975 EITC.

[18] The same logic would apply to a tax increase, but Appendix Figure A4 rules out this potential confounder.

[19] WIC provides in-kind benefits to pregnant women and mothers with children up to age 5, began rolling out at the county-level in 1972, and had small negative labor-supply effects (Moffitt and Fraker 1988; Hagstrom 1996; Moffitt and Keane 1998; Currie 2003). Hoynes, Page, and Stevens (2011) shows that the weighted percentage of counties with WIC in place rose from 0 percent in 1973, to 60 percent in 1975, to 100 percent in 1979.

[20] Food Stamps began rolling out in 1964 at the county level, and a 1973 amendment required that all counties adopt the program by 1975. During the 1970s the number of families on Food Stamps increases from about 13 million to about 20

find any evidence of public program cuts, and if anything, public programs were expanded during the 1970s which would bias the estimated EITC treatment effects in this paper downward.

Other policies to consider involve abortion and divorce laws. Abortion is legalized in 1973 with Roe v Wade, but four states (Alaska, Hawaii, New York, and Washington) legalized abortion in 1970. Appendix Figure A8 shows that legalized abortion did not result in any detectable trend break in the relative employment of mothers. Divorce has been rising in the US since 1960 and was propelled higher with the introduction of unilateral-divorce laws (Wolfers 2006). These laws began in 1970 in California and existed in all jurisdictions by 1985.[21] Isolating California shows a trend in maternal employment similar to other states (analysis not shown).

Another possible confounder that affected mothers is the 1976 Child Care Tax Credit that allowed working families to claim a credit against taxes owed for up to 20 percent of child care expenditures.[22] Averett, Peters, and Waldman (1997) show that government subsidies to child care increases the labor supply of married women. However, since the tax credit is not refundable, this program mainly benefits upper-middle class families (Tax Policy Center 2011).[23] The placebo test in Table 4 column 5 shows a null response from married women with higher family income that would have benefitted from this program. Furthermore, the triple differences analysis in Appendix B would net out any employment effects of the Child Care Tax Credit since it applied to both women in the high-impact sample and to women with family income above the EITC limit.

Finally, I show that the maternal employment response after 1975 is not due to a decline in male earnings. Appendix Figure A9 shows no noticeable change in average male – or married male – earnings during this period.

I conclude that there does not seem to be any macroeconomic event or policy confounder that would cast doubt on the analysis. The 1970s provides a nice environment in which to study the impact of the EITC. If anything, policies during the 1970s may have been discouraging mothers from working, which would imply that the results in this paper should be interpreted as lower-bound estimates of the EITC effect on female employment. This stands in contrast to studies of EITC expansions in the 1980s and 1990s when policymakers had begun explicitly cutting public benefits, setting time limits, and

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million. As with WIC, this program likely had a negative effect on employment (Hoynes and Schanzenbach 2012).

21 See Peters (1986) and Parkman (1992) for more details. Johnson and Skinner (1986) estimate that 2.6-percentage points of the 15-percentage-point increase in married women's employment between 1960 and 1980 can be attributed to increased divorce risk and future income risk. However, these policies had a gradual impact on women that began well before the 1975 EITC and is not concentrated between 1976 and 1979.

22 The credit applied to the first $2,000 for the first child and the first $4,000 for two or more children. Source: http://www.iff.org/resources/content/3/1/documents/century_of_caring.pdf

23 http://www.taxpolicycenter.org/press/quickfacts_cdctc.cfm#two

nudging low-income women into the labor force.

IV. Results: Average Effect of the EITC on the Employment of Mothers

Table 2 uses equation (3), the sample of all 16- to 45-year-old women, and 1971 to 1980 employment data to estimate the average impact of the 1975 on maternal employment. If mothers did not fully respond to the EITC until 1979 – as Figure 2 suggests – then the estimated effects in Table 2 likely under-estimate the full impact of the EITC on maternal employment. However, extending the sample into the 1980s increases the chance that estimates partly reflect unobserved or unaccounted for trends or events. As a check, I extend the sample through 1985 in analysis below.

Each regression in Table 2 controls for whether a woman is a mother and whether the observation occurs after 1975. The DD variable of interest is the interaction of these two variables ($Mom\ x\ Post_{i,s}$) since only parents were eligible for the 1975 EITC. Estimates are similar if I define Mom to be a woman with a child born before 1975 to account for potentially endogenous fertility responses to the EITC. Columns in Table 2 cumulatively add covariates left to right: Column 1 uses no additional control variables and column 2 adds state- and year-fixed effects to account for idiosyncratic state characteristics and annual shocks affecting all women. Column 3 adds demographic controls that include an age cubic, an education quadratic, number of children, welfare income, and binary variables for nonwhite, married, and having a child under 5 years old. Controlling for these demographic controls is important since mothers are on average older, have less education, and more likely to be married and nonwhite. Column 4 adds annual state-level employment-to-population ratios and federal unemployment rates to control for business cycles and local economic conditions. Finally, column 5 adds the following interactions: nonwhite-mom, nonwhite-post, age-mom, married-post, unemployment-married, and unemployment-mom. These interaction terms allow for a more flexible model that controls for economic conditions and policies that may differentially impact women that are married, nonwhite, or mothers. For these reasons column 5 is my preferred specification and all regressions in this paper use this set of controls unless otherwise specified.

For each specification the DD estimate is stable between 3.3- and 3.5-percentage points (or 6.2 and 6.6 percent from a baseline of 53 percent).[24] Since there are 52.833 million 15- to 44-year old women in the 1980 Census, and 47 percent of these are mothers according to the March CPS, this

[24] DD results in Table 2 are slightly larger if I exclude women earning above EITC limit.

means that about 800,000 mothers were induced into employment because of the 1975 EITC.[25] The binary dependent variable employed $E_{i,s,t}$ is defined as having positive earned income. Appendix Table A1 shows that the estimated treatment effect is similar across alternative definitions of working based on earnings, weeks worked, or labor-force participation. Appendix Table A2 shows that the DD results are robust to a probit, logit, or OLS model and a "kitchen-sink" set of controls that includes a number of additional interactions including interactions of each control with year, state, and race.[26]

Table 2 shows that the EITC had a large effect on the employment of all mothers. However, since 75 percent of married mothers and 15 percent of unmarried mothers already had household earnings above the EITC limit in 1974, the estimated effects in Table 2 are averaged over a sample where about half of mothers were EITC ineligible. Table 3 verifies that the EITC had a larger employment effect on mothers more likely to be EITC eligible and a null effect on mothers with household earnings above the EITC limit. Each regression in Table 3 uses equation (1) and the full set of controls from Table 2 column 5.

In Table 3 column 2 I use a "high-impact" sample, which contains all single women, all married women with spouses earning below the EITC limit (representing the bottom quartile of male earners)[27], and women not in school full time, disabled, or retired. I find that the EITC increased the employment of this group by 3.9-percentage points (or 6.8 percent). Using this "high-impact" sample, I show that the DD result is robust to the model choice and which year to end the sample in Appendix Figure A10.

The sample of women in Table 3 columns 3 and 4 are women with less than 12 years of education, and women with no more than 12 years of education.[28] For these two groups I find that the EITC led to a 4.1-percentage point (or 9.0 percent) and 3.5-percentage point (or 6.8 percent) increase in employment. Women with less education are more likely to have earnings below the EITC limit and benefit from the EITC if they choose to work. On the other hand, women with a college degree are unlikely to be eligible for the EITC and make a nice placebo test: if the employment of these women

[25] This is larger than the 164,000 and 350,000 mothers induced into employment by subsequent EITC expansions in 1986 and 1993.

[26] "Kitchen-sink" controls only explain about 10 percent of the DD estimate (OLS estimate falls from 0.040 to 0.037). Since a probit model does not converge with all these covariates, this regression in Appendix Table A2 uses OLS.

[27] Household labor supply is a joint decision between spouses, and if it becomes more attractive for a mother to begin working, it is possible that the other spouse would decrease his labor supply. However, there is evidence that male labor-supply was fairly inelastic at this time (Blundell and MaCurdy 1999), and thus a married woman deciding whether to work would likely take her spouse's labor supply as given. This approach follows Eissa and Hoynes (2004), which compares a married woman's labor-supply decision to a second mover in a two-person sequential game. In this framework the primary earner does not adjust his labor supply in response to his spouse's labor supply.

[28] Although it is possible that some women might adjust their educational attainment due to the EITC (perhaps dropping out of school to work and take advantage of the EITC earnings subsidy), this is likely to be a small effect since most women in the sample would have already completed their education.

increases after 1975, it is likely for reasons unrelated to the EITC. In Table 2 column 5 I verify that the estimated effect of the EITC on this group is statistically indistinguishable from zero.[29]

A simple DD estimator averages employment rates from years before and after 1975. To ensure that the DD result is not being driven by outliers (on either side of 1975), or a general trend, I estimate the following equation:

$$(4) \qquad P(E_{i,s,t}) = \Phi(\beta_1 Mom_{i,s,t} + \beta_2 Year_t + \textstyle\sum_{y\in[1969,1985]} \beta_{3,y}\, Mom\ x\ Year_{i,s,y} + \beta_4 X_{i,s,t} + \epsilon_{i,s,t})$$

which is identical to equation (1) except that instead of estimating the average regression-adjusted effect of having children after 1975 ($Mom\ x\ Post_{i,s}$) on employment, I estimate the *annual* effect of having children ($Mom\ x\ Year_{i,s,y}$) on the likelihood of being employed.[30] This serves as a test of whether the DD results reflect a permanent trend break. I also extend the sample period from 1971 to 1980 (used for Table 2), to 1970 to 1985, which enables a direct comparison with the unadjusted-employment trends in Figure 2.

The annual estimates of $Mom\ x\ Year_{i,s,t}$ in Figure 3 resemble the trend in Figure 2 and show that the 1970 to 1975 annual estimates are stable and near zero, become increasingly positive until 1979, and are stable between 4.5 and 6 percentage points between 1979 and 1985. The gradual increase in maternal employment between 1975 and 1979 in Figure 3 may imply that it took women a few years to learn about or fully respond to the employment-incentives of the EITC.[31] These annual estimates show that DD estimates are not driven by outliers or general trends.

V.A. Heterogeneous Effects of the EITC on the Employment of Mothers: Spousal Income

Empirical analysis above shows that the EITC had a positive average effect on the employment of

[29] Appendix Figure A11 shows some evidence that not only did lower-education women respond more, but that they also responded faster. Since lower-education women are more likely to earn in the EITC range and benefit from the EITC, information about the EITC likely traveled by word of mouth within this group. Chetty, Friedman, and Saez (2013) show that this local EITC knowledge leads to increased EITC benefits. That the highest-educated women did not respond faster than lower-education women suggests that EITC response does not primarily require an understanding of public policy, law changes and the tax code.

[30] State and year fixed effects are also included but excluded in equation (4) for visual reasons.

[31] Eissa and Liebman (1996) also find that it takes a few years for the new equilibrium to be reached after the 1986 EITC expansion. Liebman (1998) suggests that this gradual ramp up might be due to the fact that the EITC does not pay off until the tax refund in the following year; therefore it would be at least a year before EITC recipients became aware of and responded to the EITC. Even if taxpayers do not fully understand the EITC, those that are on the margin of working and not working might try working for one year, discover that they ended up better off than they expected, and decide to remain employed.

mothers, especially for subgroups of mothers more likely to have potential earnings under the EITC limit. This section incorporates spousal income n_i – that appeared in equation (2) – and shows that this is negatively correlated with responding to the EITC. See Appendix C for comparative statics and conditions required for higher spousal income n_i to induce a decrease in labor supply. The effect of spousal income can also be seen in the budget constraint depicted in Appendix Figure A2. In this figure, higher non-labor income (point A) increases the likelihood that the highest indifference curve a woman can attain intersects kink point A where her labor supply is zero. Since married women have higher non-labor income on average – and are eligible for less EITC benefits – than most single women, they should have a relatively smaller response to the EITC.[32]

Table 4 columns 1 and 2 verify this empirically. Using equation (1) and dividing the sample of all 16- to 45-year-old women into single and married, the estimated employment-effect of the EITC was 2.6-percentage points for single women and a statistically insignificant 0.6-percentage points for married women. The DD estimates of these two groups are statistically different at the 99.9-percent level.[33] This aligns with the EITC literature that has consistently found that single women respond positively to the EITC (Eissa and Liebman 1996; Meyer and Rosenbaum 2000, 2001; Keane and Moffitt 1998; Ellwood 2000; Grogger 2003; Hotz, Mullin, and Scholtz 2006; Eissa, Kleven, Kreiner 2008) but married women do not (Dickert, Houser, Scholtz 1995, Ellwood 2000, Eissa and Hoynes 1998 and 2004). However, these average effects ignore heterogeneity among married women.

Appendix Figure A2 shows that married women with spouses earning below the EITC kink point would be eligible for EITC benefits, while those with spouses earning above the EITC kink point were not. The likelihood that a mother responded to the EITC and began working should be negatively correlated with her spouse's earnings. During the 1970s male labor supply was fairly inelastic, and a married woman deciding whether to work would likely take her spouse's labor supply as given, much like a second mover in a two-person sequential game.[34] With this approach in mind, I estimate the following probit model:

$$(5) \qquad P(E_{i,s,t}) = \Phi(\beta_1 Mom\ x\ Post_{i,s} + \beta_2 Mom\ x\ Post_{i,s}\ x\ SpouseIncome + \beta_4 X_{i,s,t} + \epsilon_{i,s,t})$$

[32] Only about a fifth of married men in the 1971-1980 CPS earn below the EITC kink point, so the average DD estimate for all married women should be smaller than that for single women and theoretically may or may not be statistically larger than zero.

[33] This can be shown by running one regression with an interaction of the DD variable of interest ($Mom\ x\ Post1975_{i,s}$) with an indicator for single. Testing whether this estimate is statistically different than zero yields a chi-squared with one degree of freedom statistic of 275.16, or a p-value of less than 0.001.

[34] This approach follows the intuition provided by Eissa and Hoynes (2004).

which is identical to equation (1) except that I add one variable that interacts the DD variable of interest with real spousal earnings in 1000s of 2013 dollars.[35] Estimates in Table 4 column 3 show that the employment response of married women with non-earning spouses was 5.3-percentage points, and this response declines by an average of 0.1-percentage point for each additional $1,000 (2013 dollars) of spousal earnings.

I show that this heterogeneous response among married women is similar for the 1986 and 1993 EITC expansions. Estimating equation (5) for years surrounding the 1986 and 1993 EITC expansions, Appendix Table A3 columns 1 and 5 show that the average effect of the EITC on married women is insignificant. However when I interact the DD variable of interest with spousal earnings I find a large positive employment response among married women with non-earning spouses that declines by 0.1-percentage point for each additional $1,000 (2013 dollars) of spousal earnings. The same pattern can be seen for weekly work hours in columns 3, 4, 7, and 8. The EITC appears to subsidize maternal employment among the poorest married households. Even if the average effect of the EITC on married women is slightly negative, about a fifth of married women appear to increase their labor supply due to the EITC.[36]

Table 4 column 4 isolates married women with spouses earning below $10,000 (2013 dollars) – corresponding to the bottom fifth of the male income distribution – and finds a treatment effect of 3.8-percentage points. This estimate is nested in Figure 4 which uses the entire spousal-earnings distribution and shows the DD estimate for the sample of married women with spouses earning below various thresholds. As shown in Table 4 column 4, the DD estimate is positive and largest for married women with spouses earning below $10,000; this positive estimate gradually declines as women with spouses earning below $20,000, $30,000, etc., are incrementally added to the sample. These percentage-point estimates are even larger when displayed as a percent effect since women with lower-earning spouses had lower employment rates before the 1975 EITC. Results in Table 4 and Figure 4 provide a more nuanced understanding of how married women respond to the EITC, and show that married women with spouses earning in the bottom fifth of the male income distribution responded positively to the EITC much like single women did.

Since women with spouses earning above the EITC limit are not eligible for EITC benefits, this

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[35] $\beta_1 Mom_{i,s,t}$, $\beta_2 Post_{i,s}$, γ_s, and δ_t are also in the regression but omitted from equation (5) for visual purposes.

[36] Although Eissa and Hoynes (2004) find a negative average response among married women, the authors implicitly anticipate this heterogeneity by acknowledging that the EITC will provide some spouses with an incentive to work and state that "If the spouse is not working, the EITC (as in the case of single parents) encourages the wife to enter the labor force. If the primary earner has income in the subsidy region and the effect of the greater returns to work dominates the income effect, the EITC would encourage employment."

group can be used for another placebo test.[37] Taking spousal income as given (not completely unrealistic given the inelasticity of male labor supply during the 1970s), these mothers could not benefit from the EITC whether or not they worked and faced the same employment incentives before and after 1975. Table 4 column 5 shows that this DD estimate is indeed near zero at -0.1-percentage points, and small employment effects can be ruled out for this group of mothers. This placebo test shows that *something* affected the labor supply of single mothers and married mothers with low-earning spouses after 1975, and did not affect married mothers with higher-earning spouses.

A similar approach to looking at spousal earnings is to look at the amount of EITC benefits a woman was eligible for. Taking spousal earnings as given, and using a small EITC expansion in 1979 along with the changing real value of nominal EITC benefits[38], I calculate the maximum potential real EITC benefits (in 2013 dollars) that a woman could have received ($maxEITC$). For mothers with non-earning spouses and unmarried mothers, the real value of $maxEITC$ varied by year and ranged between about $1,300 and $1,600; for married women with a working spouse earning above the EITC kink point, $maxEITC$ was zero; for married women with a spouse earning below the EITC kink point, $maxEITC$ was equal to 10 percent of the difference between the EITC kink point and her spouse's earnings. For example, a mother with spousal earnings of $10,000 and an EITC kink point of $16,000 would have a $maxEITC$ value of $600.[39] Table 4 column 6 uses the sample of all 16- to 45-year-old women and shows that a $1,000 (2013 dollars) increase in $maxEITC$ is associated with a 4-percentage point increase in maternal employment. This regression also carries out the placebo test from column 5 in a slightly different way since mothers eligible for zero EITC benefits have a statistically insignificant employment response of -0.2-percentage points. This section shows that mothers ineligible for the EITC did not increase their employment after 1975, while the employment response of EITC-eligible single and married women was positive and proportional to the size of potential EITC benefits.

V.B. Heterogeneous Treatment Effects of the EITC: Cost of Living

The 1975 EITC was a 10 percent wage subsidy worth up to about $1,800 (in 2013 dollars) for low-income families. Although these benefits did not vary geographically, mothers living in lower cost of living areas should respond relatively more to the EITC. One reason is that wages are lower in these areas and therefore more likely to be under the EITC limit, and a second reason is that each dollar of

[37] This group of women is also used in Appendix B as a control group for triple differences analyses.
[38] EITC benefits were not automatically adjusted for inflation until 1986
[39] See Appendix Figure A12 for a histogram of $maxEITC$.

EITC benefits has more purchasing power than in more expensive areas. This heterogeneous treatment effect should occur both between states and within states. See Fitzpatrick and Thompson (2010) for further discussion and see Appendix C for comparative statics and conditions required for lower cost of living COL_i to induce an increase in labor supply.

To empirically test this, I estimate the following probit model.[40]

$$(6) \qquad P(E_{i,s,t}) = \Phi(\sum_{r\in\{1,2,3\}}[\beta_{1,r}COL_{i,r} + \beta_{2,r}Mom\ x\ Post\ x\ COL_{i,r}] + \beta_2 X_{i,s,t} + \epsilon_{i,s,t})$$

where $COL_{i,r}$ denotes cost of living for individual i in region r. One simple way to measure cost of living is to divide the sample into women living in large cities, mid-sized cities, and more rural areas.[41] Cost of living is lowest in rural areas and Table 5 verifies that the maternal employment response to the EITC was largest in these locations, second largest in mid-sized cities, and smallest in large cities. These three DD point estimates are 4.7, 3.1, and 2.5 percentage points. Each estimate is statistically larger than zero, and the effect of the EITC is statistically larger than in large and mid-sized cities.

V.C. Heterogeneous Treatment Effects of the EITC: Social Stigma against Working Women

For many women – especially in the 1970s – one potential cost of working was an external social stigma against working women, and a second was the internal identity cost for women with gender norms that precluded mothers from working. Intuition for the first cost comes from the Becker (1957) discrimination model, and intujition for the second cost comes from Akerlof and Kranton (2000). I show that state-level stigma can account for a significant part of the regional variation in the maternal employment response to the EITC. See Appendix C for comparative statics and conditions required for higher stigma S_s to induce a decrease in labor supply in equation (2).

To measure this stigma I use survey questions about the role of women in society in American National Election Surveys (ANES) data and General Social Survey (GSS) data.[42] Appendix Table A3

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[40] Mom and $Post$ are still controlled for but excluded from the equation for visual purposes.

[41] Large cities are defined as one of the 33 metro areas defined in every year of the data, mid-sized cities are "unspecified metro areas," and rural areas are non-metro areas with a population under 50,000 (Nelson 1986). March CPS data in the 1970s is not detailed enough to analyze cost of living at the county-level as is done in Fitzpatrick and Thompson (2010) for the 1993 EITC expansion.

[42] Stigma is created from ANES variable *equalrole*, which asks respondents whether men and women should have equal roles or whether a woman's role is in the home. These responses are scaled between 0 and 6, where 6 represents more stigma. Stigma is also created from the GSS variable *fework*, which asks respondents whether women should work. This binary response is coded so that 1 represents more stigma. Stigma is standardized by subtracting the 1972 mean and dividing by the 1972 standard deviation.

shows correlations between individual stigma against working women and demographics using 1972 to 1988 ANES and GSS data, and show that people with less stigma include younger people, people born more recently, people with more education, and that the trend in stigma has been declining over time. Appendix Figure A13 shows that average annual stigma measured in the GSS and ANES have declined similarly over time, and Appendix Figure A14 shows this trend for each region separately. In general, the regional ranking of stigma has remained somewhat constant over time and the trend over time is fairly parallel.

To empirically test whether stigma can help account for regional variation in the maternal employment response to the EITC, I construct region-level measures of stigma by averaging individual beliefs of males in the ANES data within each region just before 1975. This region-specific variable is merged onto individual-level CPS data and interacted with the DD variable of interest ($Mom\ x\ Post$). I use this measure of stigma to estimate the following probit model.[43]

$$(7) \qquad E_{i,s,t} = \beta_1 Mom\ x\ Post\ x\ Stigma + \sum_{r \in \{1,2,3\}} [\beta_2 Mom\ x\ Post\ x\ COL_{i,r}] + \beta_3 X_{i,s,t} + \epsilon_{i,s,t}$$

Table 5 column 1 shows the DD estimate from Table 3 column 2, and columns 2 and 3 show that region-level stigma had a negative effect on female employment and a negative effect on the maternal employment response to the EITC. Columns 4 and 5 add measures of cost of living and show that even when controlling for cost of living, the effect of stigma is still negative and statistically significant. I conclude that both cost of living and stigma about gender roles can help explain why the response to the EITC varied by region, even though the EITC benefits did not.

VI. Robustness Checks

A number of robustness checks were carried out in previous sections. In Appendix B I show that the results are robust to triple differences analysis, weighting, sample composition, alternate approaches to the default March CPS imputing algorithm, and potentially endogenous fertility responses to the EITC.

VII. Annual Earnings, Work Hours, and Quantile Treatment Effects

Results so far have focused on labor supply at the extensive margin. These results suggest that the

[43] Mom, $Post$, $Stigma$, and COL are also controlled for but excluded from the equation for visual purposes.

EITC should also have led to increases in annual earnings and work hours. Although there may have been some intensive margin responses among women already working,[44] the bulk of these increases due to the EITC are likely driven by women entering into employment on the extensive margin. Results in Table 6 use equation (1) and replaces binary employment as the outcome variable with annual earnings and hours worked. For each outcome I show results for four samples of 16- to 45-year-old women: the "high-impact" sample, women with less than a high school education, all women, and married women with spouses earning above the EITC limit. These four groups were used for analysis in Tables 3 and 4, and the last group serves as a placebo test since the EITC should not have affected their labor supply.

Results in columns 1 and 5 show that the EITC led to increases of 67.7 annual work hours and $1298.6 annual earnings (2013 dollars) for the "high-impact" sample. Columns 2 to 3 and 6 to 7 show that lower educated mothers and the sample of all mothers also increased annual work hours and earnings after 1975, but not by as much as the "high-impact" sample. Columns 4 and 8 show that the work hours and earnings response of EITC-ineligible married women is not statistically different than zero. I can reject small responses on these margins among this placebo group, which corroborates the extensive-margin placebo test in Table 4 column 5.

Another way to examine the impact of the 1975 EITC on female labor supply is to look at the distribution of earnings and work hours. Although this paper has shown that more mothers work after the introduction of the EITC, it is unclear where in the income distribution they enter when they begin working. This also serves as a robustness check because it would raise concerns if women entering the labor force immediately earned above the EITC limit. Figure 5 plots the estimated effect of the EITC on the likelihood of mothers in the "high-impact" sample earning within different ranges of the income distribution. Each point in Figure 5 shows a separate estimate on $Mom\ x\ Post_{i,s}$ from a probit regression resembling equation (3) with the full set of controls, but with a binary outcome variable for earning in a certain income range. Mirroring Table 3 column 2, this figure shows that mothers are 3.9-percentage points less likely to earn zero income after the EITC is in place. Figure 5 also shows that mothers are more likely to earn positive income below about $40,000, with the largest increase occurring between $10,000 and $20,000 (in 2013 dollars). This makes sense since the minimum wage during the late 1970s was between $8 and $9 per hour (in 2013 dollars), and $10,000 to $20,000

[44] There is theoretical justification (but little empirical evidence) that some women will decrease their labor supply on the intensive margin to take advantage of EITC benefits (Meyer 2002; Saez 2002; Eissa and Hoynes 2006). Saez (2010) finds evidence that self-employed workers bunch at the first EITC kink point, and LaLumia (2009) finds that the EITC induces more low-income workers to report self-employed income. These workers are a small fraction of the population.

corresponds to working 25 to 40 hours per week at this rate.

Analysis in Figure 6 shows how the distribution of annual work hours was affected by the EITC. Similar to Figure 5, each point reflects the estimate on $Mom\ x\ Post$ from a probit regression resembling equation (3) with the full set of controls, but with a binary outcome variable for having annual work hours within a particular range. Also similar to Figure 5 (and reflecting the estimate in Appendix Table A1 column 5) mothers are about 3.5-percentage points less likely to work zero hours. More mothers worked part-time and full-time after 1975, with the largest increase being the likelihood of working around 2000 hours a year. It appears that the most common response to the EITC was to begin working full-time full-year (e.g. about 40 hours a week and 50 weeks a year).

Finally, I use quantile regression to show how the EITC affected the full distribution of annual earnings and work hours. Each estimate in Figure 7 (and 8) is from a separate regression using the full set of controls from Table 2 column 5 and the regression behind Table 6 column 5 (and column 1). However, instead of estimating *average* effects, estimates show the effect of $Mom\ x\ Post$ at each centile of the dependent variable.[45] These quantile DD (QDD) estimates in Figures 7 and 8 are shown for the "high-impact" sample and should be interpreted as effects *at* centiles and not *on* centiles, since it is unclear whether rank preservation holds without panel data (which is needed to interpret these as effects *on* deciles; see Bitler, Gelbach, and Hoynes 2003 for a detailed discussion). Figure 7 shows that centiles 45 to 55 saw the largest increase in earnings, and this positive effect declines further up the income distribution. Though there appears to be a positive effect in the top few deciles, the “high-impact” group contains some higher-earning women that were already working before the EITC existed, which likely explains some of this increase in earnings at the top. Figure 8 shows that centiles 50 to 55 saw the largest increase in annual work hours, and this positive effect also declines further up the work-hours distribution. In Figures 7 and 8 the lowest four deciles saw no effect, which reflects the fact that many mothers still did not work after the 1975 EITC was in place. These QDD results show that the EITC had the largest effect on the middle of the income and hours distribution, and these gains drive the average effects shown in Table 6.

VIII. Summary

This paper is the first systemically study of the introduction of the EITC in 1975. I investigate the role

[45] Instead of minimizing the sum of squared residuals like OLS, quantile regression uses potential heteroskedasticity and minimizes a weighted sum of the absolute value of the residuals.

of the 1975 EITC in the rise of working mothers and find that this program can help explain why the U.S. has long had such a high fraction of working mothers despite having little childcare subsidies or maternity leave policies.

Using CPS data spanning 1969 to 1985, a dynamic difference-in-differences approach shows that the EITC led to a 6 percent (or 3.3-percentage-point) increase in the maternal employment of 16- to 45-year-old women. This effect is even larger for mothers more likely to have earnings below the EITC limit, including mothers with relatively low education. Figures 2 and 4 show unadjusted and regression-adjusted evidence that the relative employment of mothers began to increase after 1975. Estimates in this paper suggest that the EITC explains about a third of the 1975 to 1980 rise in maternal employment and a quarter of the overall rise in female employment.

Since the number of EITC-eligible families do not claim the EITC and the number of EITC-ineligible families do claim the credit is about 20 percent (Scholtz 1994), the effects estimated in this paper are intent-to-treat effects. Liebman (1995a; 1995b) finds that 89 and 95 percent of women allocated to the treatment and control groups filed taxes appropriately in the 1980s. Assuming that this misallocation occurs at random, Eissa and Liebman (1996) argue that this contamination should result in scaling up the employment effects of the EITC by 19 percent. This would imply that my DD estimate of 3.3-percentage points should actually be about 4-percentage points, which corresponds to about 1 million mothers.

This paper also shows that the employment response of married women with low-earning spouses is positive and is similar in magnitude to that of single women. The EITC treatment effect on married women is negatively correlated with her spouse’s earnings. This pattern is also evident during the 1986 and 1993 EITC expansions and is not particular to the 1975 program introduction. Regional variation in the maternal employment response to the EITC can be explained by cost of living and by the social stigma against working women.

Using a number of placebo tests, I show that these results cannot be explained by macroeconomic events, other public policies, declining men’s employment and earnings, or evolving demographic traits of mothers. Placebo tests for EITC-ineligible married women with spouses earning above the EITC limit and for higher educated women unlikely to earn in the EITC range yield estimates near zero.

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Figure 1A: Employment Rates for Mothers and Women without Children

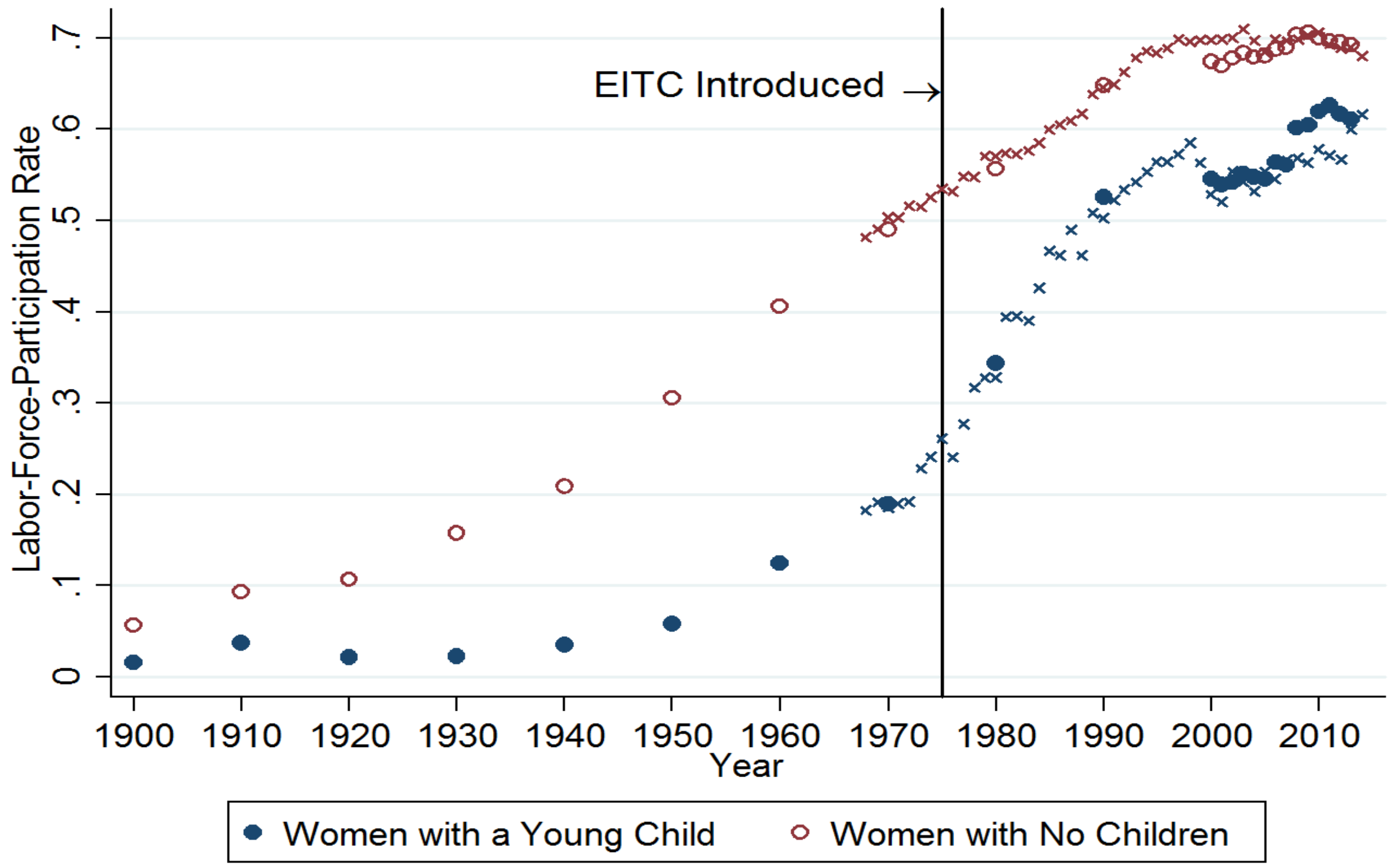
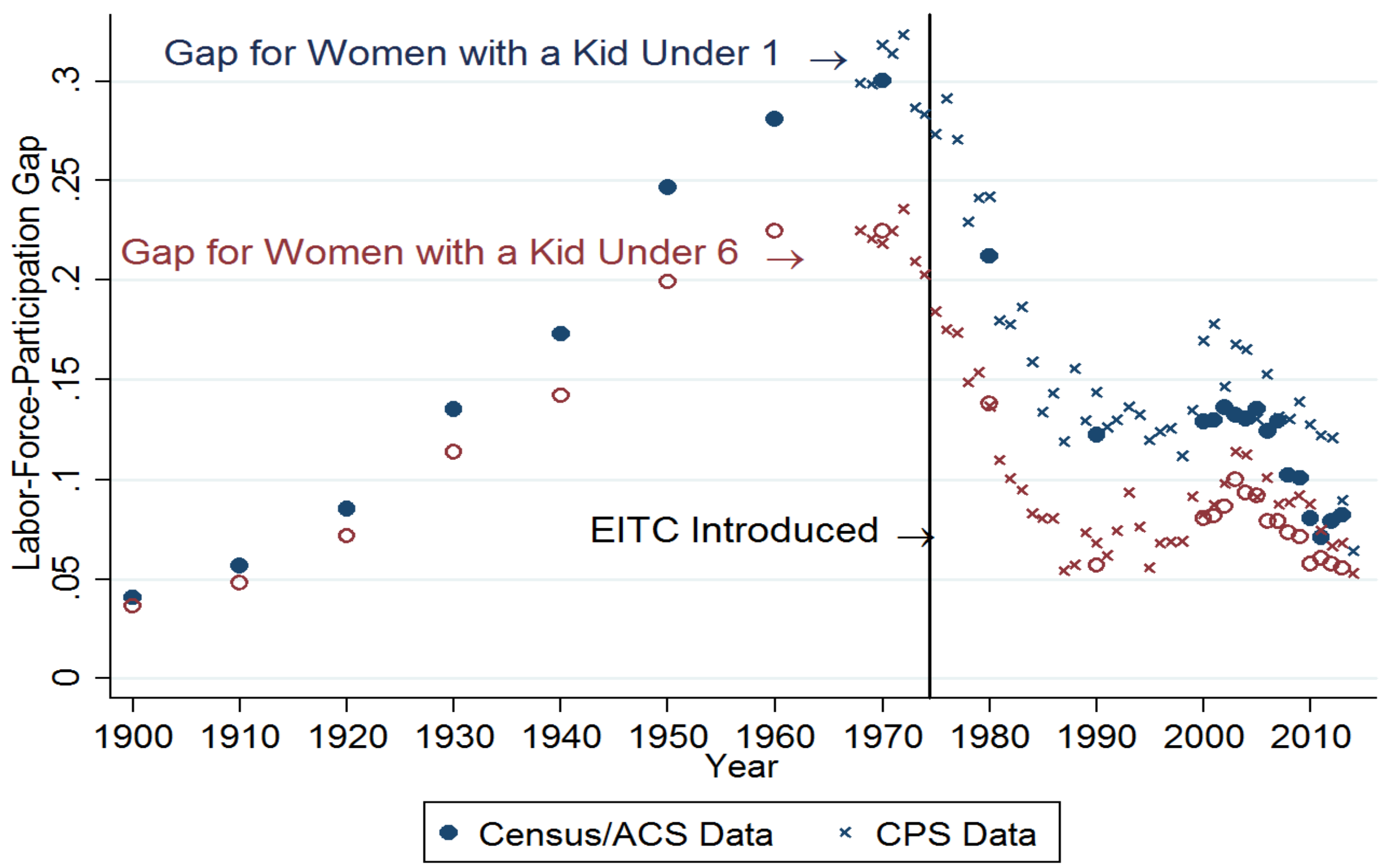


Figure 1B: Employment Gap between Mothers and Women without Children



Notes: Author's calculation from 1900-2000 Census, 2000-2013 American Community Survey, and 1968 to 2013 March Current Population Survey data. Sample includes all white, married women 16-64. Young children are under 1 and no children means no kids under 18.

Figure 2A: Unadjusted Employment Trends for Mothers and Women without Kids

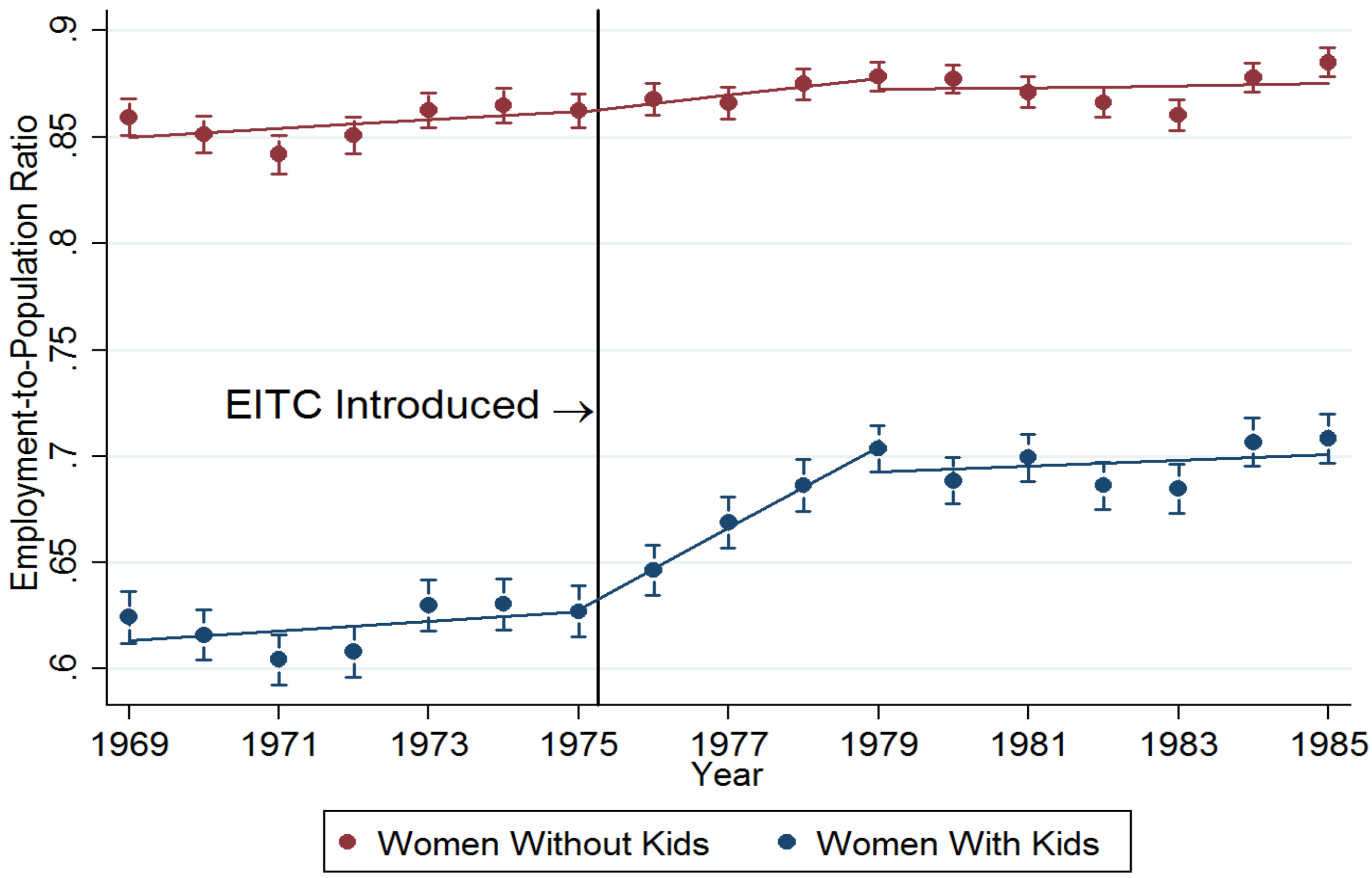
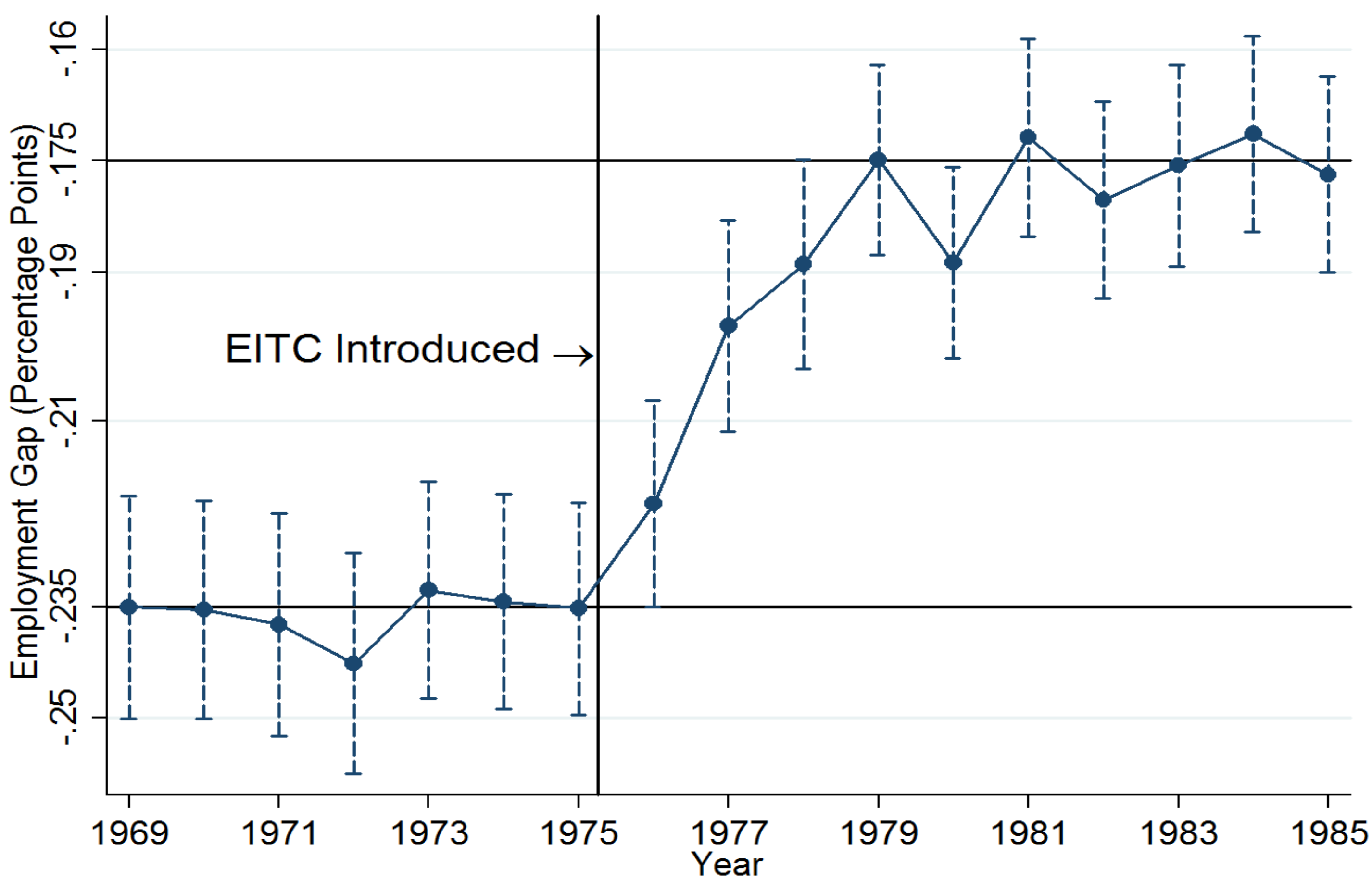
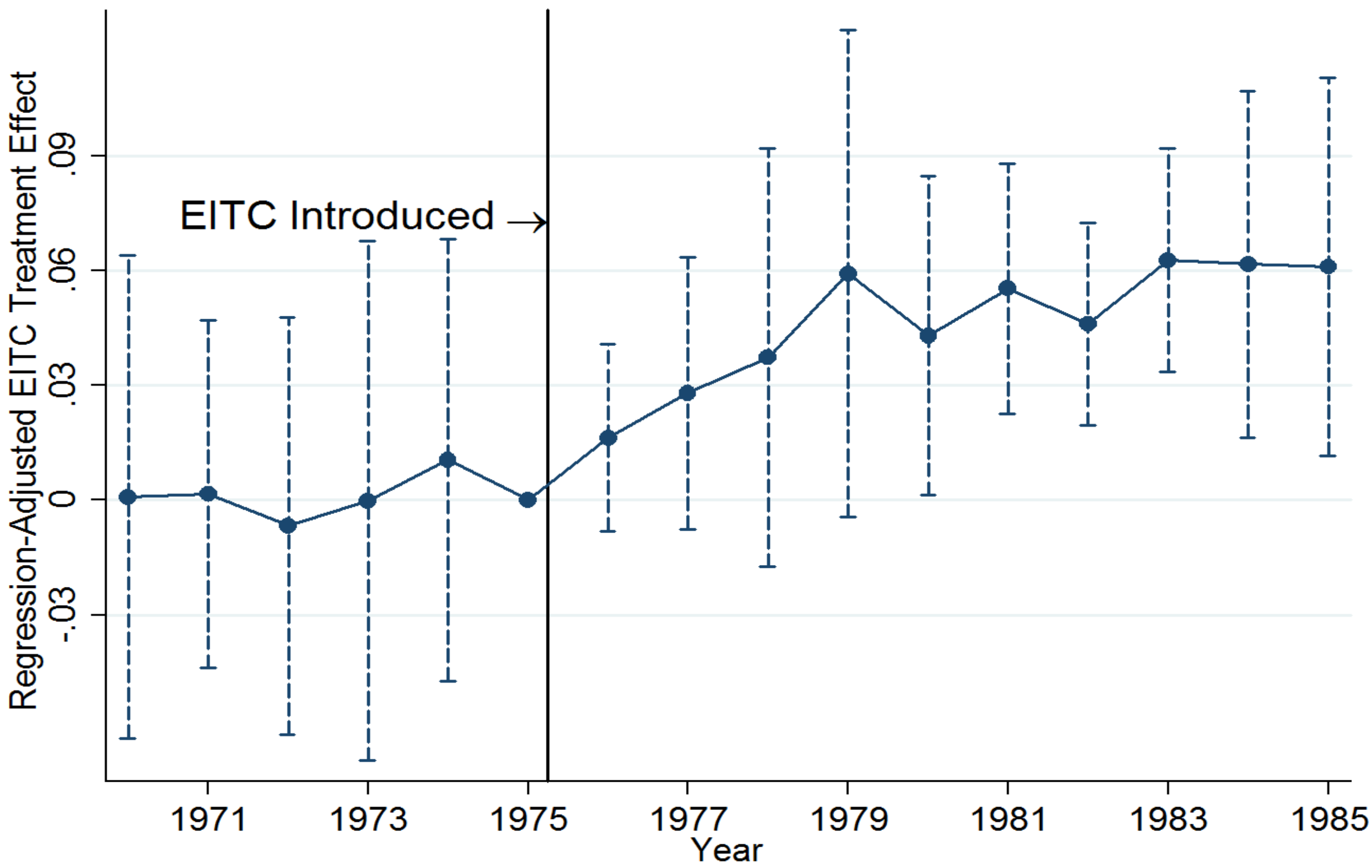


Figure 2B: Unadjusted-Employment Gap between Mothers and Women without Kids



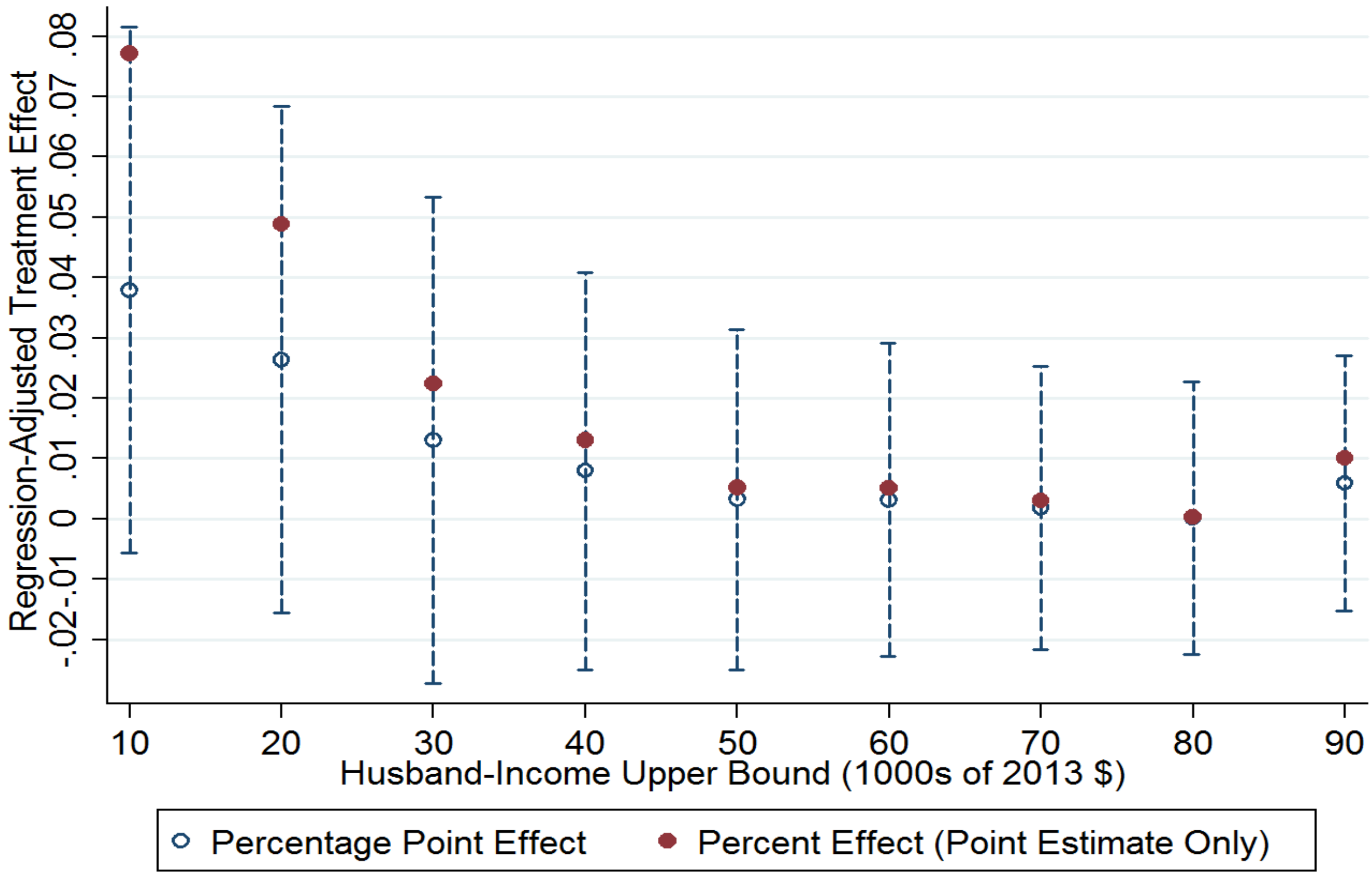
Notes: Employment rates in Panel A calculated from March CPS data by regressing binary employment on a constant for each group-year. Best fit lines are shown for 1969-75, 1975-79, 1979-85. In Panel B, Employment gaps calculated from March CPS data by regressing binary employment on binary variable kids and a constant for each year. Horizontal lines are approximate averages for 1969-75 and 1979-85. Sample includes all women 16-45 that are single or married to a spouse earning below $36,000 (2013 $). Kids are defined as 0-18 years old.

Figure 3: Regression-Adjusted Annual Employment Effect of the EITC



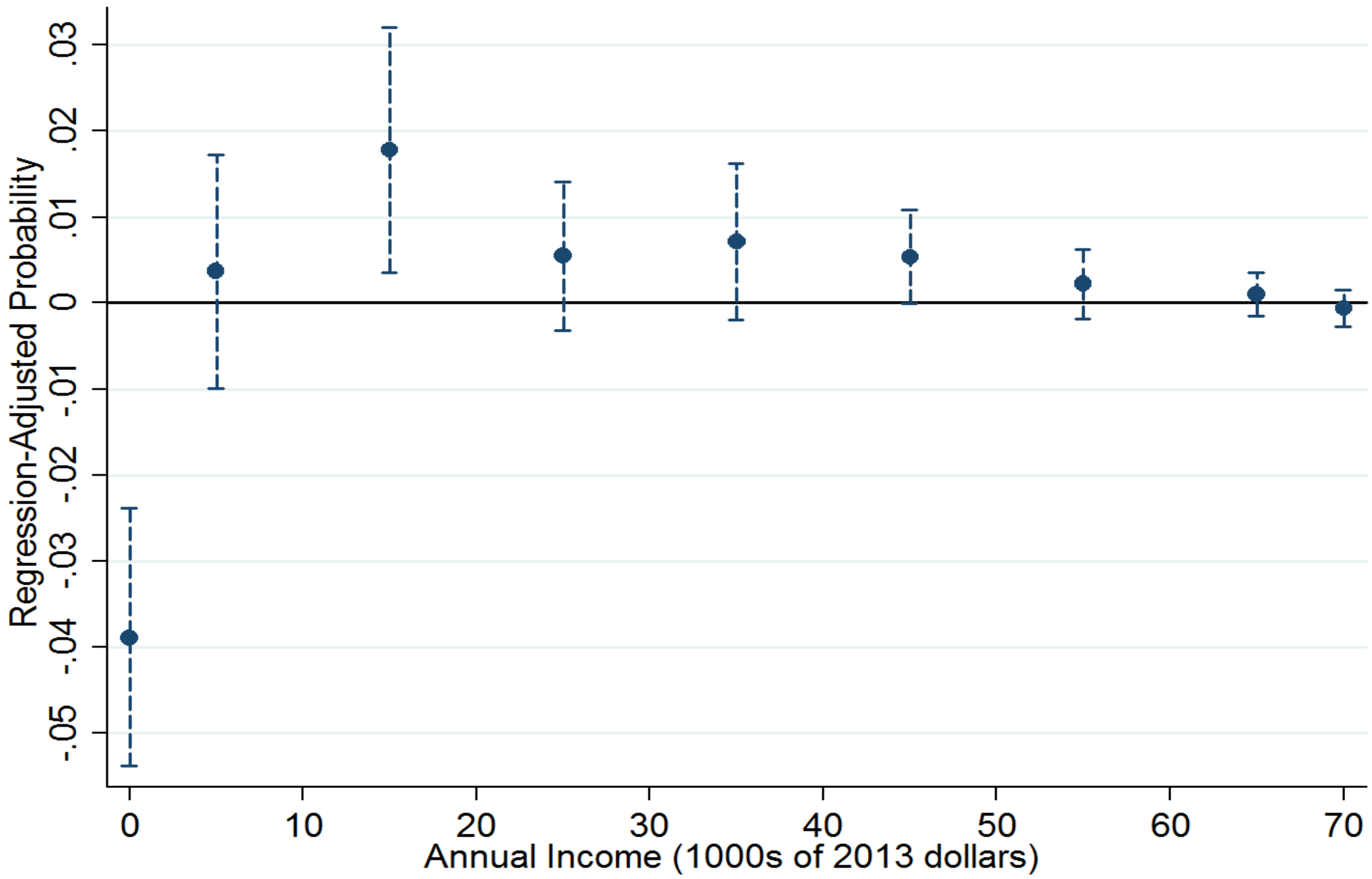
Notes: 1971-1986 March CPS data. Full set of controls used from Table 2 and "high-impact" sample used. Treatment effects are estimates of *Mom x Year* in equation (4). The p-value from the test that the coefficients for 1970-1975 are jointly statistically significant is 0.63.

Figure 4: Married Women with Lower-Earning Spouses Responded to the EITC



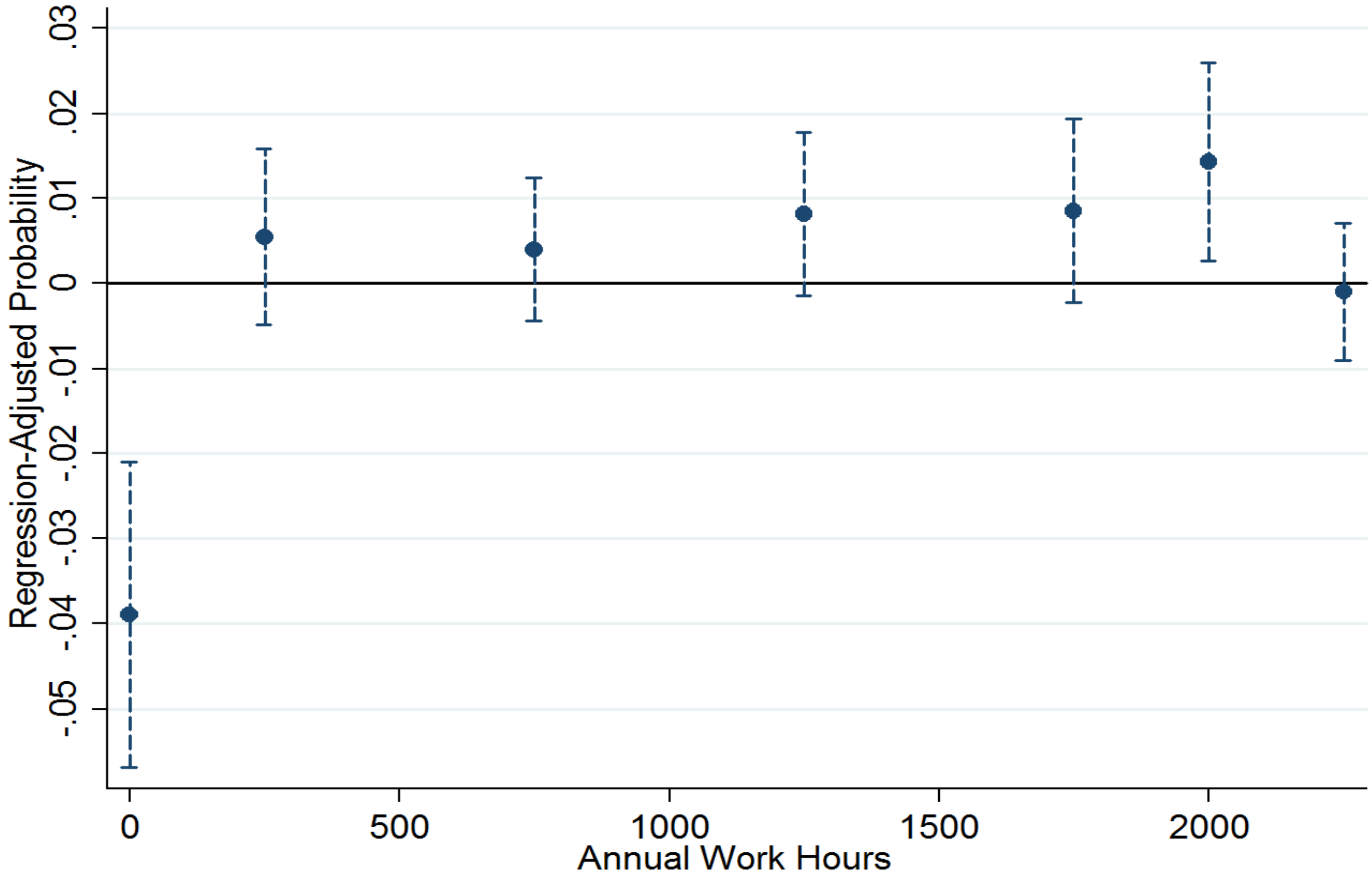
Notes: 1971-1986 March CPS data. Each estimate is from a separate regression that uses the full set of controls from Table 2 column 5 and the sample of all married women with spouses earning below the specified amount. Treatment effects are estimates of *Mom x Post* in equation (3). Sample sizes for these nine regressions are 27625, 37757, 55540, 79267, 106185, 132009, 153904, 168233, and 201212. The mean dependent variable (binary employment) for these nine regressions are 0.49, 0.54, 0.58, 0.61, 0.62, 0.62, 0.62, 0.61, 0.59. The point-estimate response to the EITC from married mothers with spouses earning below $20,000 (2013 dollars) is roughly equal to the response by unmarried mothers (2.6-percentage points).

Figure 5: The Effect of the EITC on the Distribution of Annual Earnings



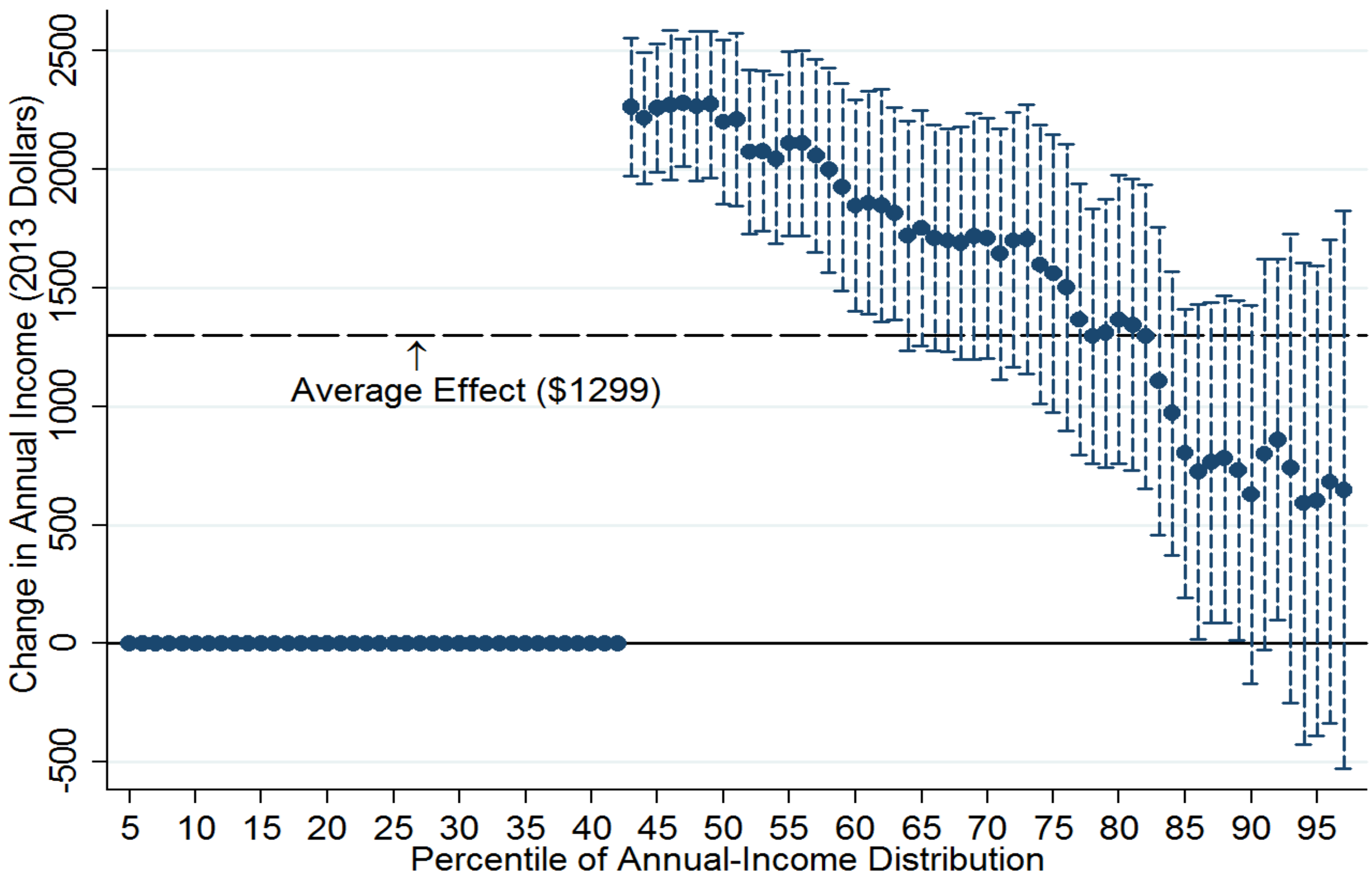
Notes: 1972-1981 March CPS data. Full set of controls used from Table 2 and "high-impact" sample used with 133,805 observations. Each estimate is from a different probit regression where the dependent variable is binary for earning $0, between $1 and $10000, $20000 and $30000, $30000 and $40000, $40000 and $50000, $50000 and $60000, $60000 and $70000, and above $70000. The mean dependent variable for these nine regressions are 0.25, 0.27, 0.15, 0.13, 0.10, 0.05, 0.03, 0.01, 0.01. This figure shows that less mothers earn zero and more mothers earn positive income due to the EITC. It appears that most mothers entering into employment earn between $10,000 and $20,000 (in 2013 dollars).

Figure 6: The Effect of the EITC on the Distribution of Annual Work Hours



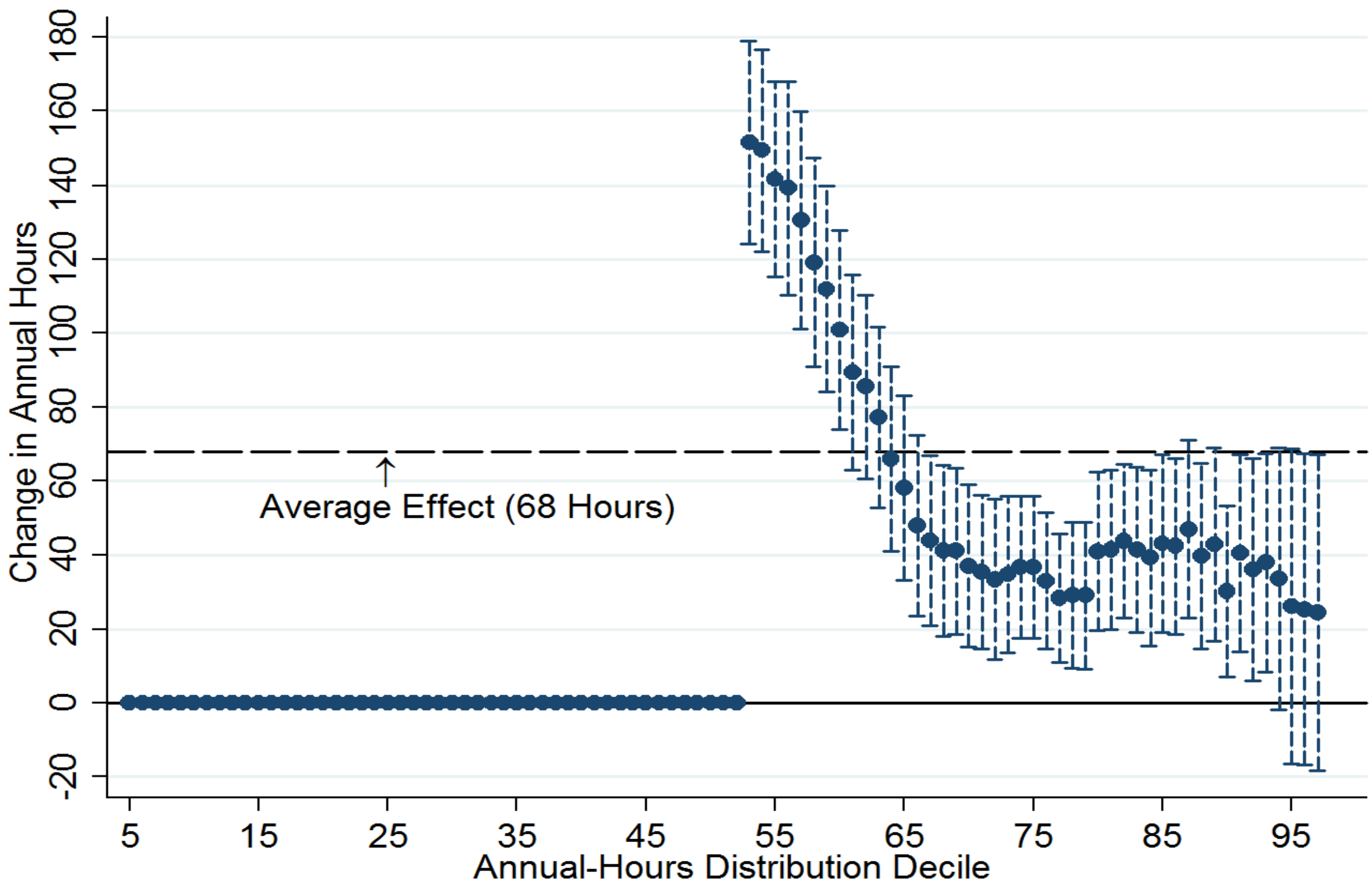
Notes: 1972-1981 March CPS data. Full set of controls used from Table 2 and "high-impact" sample used with 133,805 observations. Each estimate is from a different probit regression where the dependent variable is binary for having annual work hours of 0, or between 1 and 499, 500 and 999, 1000 and 1499, 1500 and 1999, 2000 to 2099, and at least 2100. The mean dependent variable for these seven regressions are 0.36, 0.11, 0.09, 0.08, 0.11, 0.17, and 0.08. This figure shows that less mothers work zero hours and more mothers work a positive number of hours due to the EITC. It appears that the largest change occurs for mothers working around 2000 hours a year (or about 40 hours a week and 50 weeks a year).

Figure 7: Quantile DD on Annual Earnings for the "High-Impact" Group



Notes: 1972-1981 March CPS data. Each estimate is from a separate regression using the full set of controls used from Table 2 column 5 and mimics the regression behind Table 6 column 5 except that instead of estimating average effects, it shows the effect of *mom x post* at each decile. The "high-impact" sample has 133,805 observations. These are effects *at* deciles and not *on* deciles, without panel data it is unclear whether rank preservation holds (which is needed to interpret these as effects *on* deciles). The mean dependent variable at these nine deciles are 0, 0, 1150.9, 4195.1, 8624.0, 14764.0, 22986.3, 29636.9, and 40119.4; for mothers in 1975 the mean dependent variable at these deciles are 0, 0, 0, 0, 1082.0, 6059.3, 13417.1, 21787.7, and 33759.2. This figure shows that the lowest three deciles of women (by annual earnings) in each sample did not work before or after 1975. Mothers in the top three deciles experienced an increase in real earnings, which may reflect positive trends in female earnings (among higher earners) unrelated to the EITC. The sample contains women that were already working before the 1975 EITC and may not be affected by this program. This is especially true for women already earning above the EITC kink point, which varied between $13,000 and $18,000 in 2013 dollars, and corresponds roughly to the threshold for the 70th centile. March CPS data are repeated cross sections and do not allow for this to be directly tested. The EITC had a large effect on the fourth, fifth, and sixth centiles.

Figure 8: Quantile DD on Work Hours for "High-Impact" Group



Notes: 1972-1981 March CPS data. Each estimate is from a separate regression using the full set of controls used from Table 2 column 5 and mimics the regression behind Table 6 column 1 except that instead of estimating average effects, it shows the effect of mom x post at each decile. The "high-impact" sample has 133,805 observations. These are effects at deciles and not on deciles, without panel data it is unclear whether rank preservation holds (which is needed to interpret these as effects on deciles). The mean dependent variable at these nine deciles are 0, 0, 0, 168, 660, 1254, 1785, 2040, and 2040; for mothers in 1975, the mean dependent variable at these deciles are 0, 0, 0, 0, 0, 765, 1530, 2040, and 2040. This figure shows that the lowest four deciles of women (by annual work hours) in each sample did not work before or after 1975. Unlike Figure 7 which showed earnings, women working enough hours to be in the top decile or two with respect to hours, may have low wages and fall in the middle of the income distribution.

Table 1: Summary Statistics

	All Women	Mothers	Women without Kids
Variable	(1)	(2)	(3)
Age	28.73	32.7	23.6
	(8.5)	(7.1)	(7.4)
Years of Education	12.05	12.0	12.2
	(2.5)	(2.5)	(2.5)
Black	0.11	0.11	0.10
	(0.31)	(0.31)	(0.31)
Hispanic	0.08	0.09	0.07
	(0.28)	(0.29)	(0.26)
Kids Under 5	0.25	0.45	0.00
	(0.44)	(0.50)	(0.00)
Number of Kids	1.31	2.30	0.00
	(1.50)	(1.29)	(0.05)
Earned Income (2013 $)	12,262	10,743	14,247
	(16,375)	(15,562)	(17,178)
Earned Income Conditional on Working (2013 $)	19,282	19,453	19,117
	(16,919)	(16,403)	(17,403)
Income Below EITC limit	0.87	0.89	0.84
	(0.34)	(0.32)	(0.37)
Income Below EITC limit Conditional on Working	0.50	0.44	0.58
	(0.50)	(0.50)	(0.49)
Employed	0.64	0.55	0.75
	(0.48)	(0.50)	(0.44)
Annual Weeks Worked	24.7	22.1	28.1
	(22.3)	(22.4)	(21.8)
Annual Weeks Worked Conditional on Working	37.2	37.3	37.1
	(16.9)	(16.7)	(17.1)
Weekly Hours Worked	17.2	15.3	19.8
	(19.2)	(18.9)	(19.4)
Weekly Hours Worked Conditional on Working	33.7	33.7	33.7
	(12.9)	(12.8)	(13.1)
Observations	335,639	191,187	144,452

Note: Data source: 1972-1981 March CPS. Mothers compose the treatment group and women without kids the control group. Sample contains all women 16 to 45 years old. Standard deviations are in parentheses.

Table 2: EITC and the Employment of the "High-Impact" Sample

	Mean Dependent Variable=0.637				
	Mean Dependent Variable for Treatment Group in 1975=0.530				
Variables	(1)	(2)	(3)	(4)	(5)
Mom	-0.207***	-0.207***	-0.053***	-0.053***	-0.056
	(0.016)	(0.016)	(0.008)	(0.008)	(0.101)
Post1975	0.041***	0.083***	0.048***	-0.000	0.001
	(0.005)	(0.006)	(0.006)	(0.013)	(0.014)
Mom x Post1975	0.033***	0.033***	0.035***	0.034***	0.033***
	(0.005)	(0.005)	(0.006)	(0.006)	(0.009)
Controls					
State and Year		X	X	X	X
Demographic Controls			X	X	X
Unemployment Rate				X	X
Interaction Controls					X
Observations	335,639	335,639	335,639	335,639	335,639

Note: *** p<0.01, ** p<0.05, * p<0.1. Data source: 1972-1981 March CPS. Sample includes all women 16 to 45 years old. Binary dependent variable employment. CPS weights used and average marginal effects from probit regression are shown. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. Demographic controls include married, welfare income, number of children, any children under 5, an age cubic, and a years of education quadratic. Unemployment rate includes both annual federal unemployment rates and state-year employment-to-population ratios. Interaction controls include age-kid, nonwhite-kid, nonwhite-post1975, married-post1975, unemployment rate-married, and unemployment rate-kid. Appendix Table A2 shows similar estimates when a logit or OLS is used, as well as a "kitchen-sink" regression where additional double- and triple-interaction controls are included.

Table 3: EITC and the Employment of Various Groups of Women

Sample of Women:	All	"High-Impact" Group	Less than High-School Education	No More Than High-School Education	Placebo Group: College Education
Mean Dependent Variable:	0.637	0.747	0.500	0.594	0.772
Mean Dependent Variable for Mothers in 1975:	0.530	0.573	0.456	0.511	0.628
Variables	(2)	(1)	(3)	(4)	(5)
Mom	-0.056	-0.008	0.201	-0.049	-0.210***
	(0.101)	(0.077)	(0.149)	(0.119)	(0.072)
Post1975	0.001	-0.023	-0.068***	-0.011	0.030
	(0.014)	(0.016)	(0.020)	(0.017)	(0.021)
Mom x Post1975	0.033***	0.039***	0.041***	0.035***	0.029
	(0.009)	(0.008)	(0.016)	(0.010)	(0.018)
Observations	335,639	133,805	97,683	239,583	39,646

Note: *** p<0.01, ** p<0.05, * p<0.1. Data source: 1972-1981 March CPS. Binary dependent variable employment. CPS weights used and average marginal effects from probit regression are shown. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. Each column represents a separate regression with the full set of controls from Table 2 column 5. All samples are limited to women 16 to 45 years old.

Table 4: Married Women with Lower-Earning Spouses Respond to the EITC Much Like Unmarried Mothers

Sample of Women:	All Single	All Married		Married with Spouse Earning Below $10,000	Married and EITC Ineligible (Due to Spouse Earning Above EITC Kink)	All
Mean Dependent Variable:	0.710	0.589	0.588	0.492	0.601	0.637
Mean Dependent Variable for Mothers in 1975:	0.643	0.509	0.508	0.435	0.518	0.530
Variables	(1)	(2)	(3)	(4)	(5)	(6)
Mom	-0.036	-0.425***	-0.458***	-0.324	-0.459***	-0.043
	(0.046)	(0.110)	(0.110)	(0.204)	(0.093)	(0.091)
Post1975	-0.027	0.058***	0.053***	0.040	0.072***	-0.007
	(0.017)	(0.018)	(0.019)	(0.032)	(0.021)	(0.014)
Mom x Post1975	0.026***	0.006	0.053***	0.038*	-0.001	-0.002
	(0.009)	(0.011)	(0.011)	(0.022)	(0.011)	(0.009)
Mom x Post1975 x Spousal Income (in 1000s of 2013 $)			-0.001***			
			(0.000)			
Mom x Post1975 x MaxEITC (in 1000s of 2013 $)						0.045***
						(0.006)
Observations	134,427	201,212	197,927	27,625	167,949	335,639

Note: *** p<0.01, ** p<0.05, * p<0.1. Data source: 1972-1981 March CPS. Binary dependent variable employment. CPS weights used and average-marginal effects from probit regression are shown. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. Each column represents a separate regression with the full set of controls from Table 2 column 5. All samples are limited to women 16 to 45 years old. EITC kink varied slightly by year and corresponds to roughly $18,000 (2013 dollars). The sample in columns 5 and 6 splits the sample in column 2. Sample sizes differ between columns 2 and 3 due to missing spousal income information. In column 6 *maxEITC* is the maximum potential real EITC benefits (in 2013 dollars) that a woman could have received. For women with non-earning spouses and single women, the real value of *maxEITC* varied by year and ranged between about $1,300 and $1,600. For married women with a working spouse earning above the EITC kink point, *maxEITC* was zero. For married women with a spouse earning below the EITC kink point, *maxEITC* was equal to 10 percent of the difference between the EITC kink point and her spouse's earnings. For example, a mother with a spouse earning $10,000 when the kink point was $16,000 would have a maxEITC of $600. See Appendix Figure A4 for a histogram of *maxEITC*.

Table 5: Effect of the EITC Smaller in Areas with More Stigma Against Working Women

Panel A: Maternal Employment, the EITC, and Region-Level Stigma

	Main DD	Cost of Living	Stigma		Stigma and Cost of Living	
Variables	(1)	(2)	(3)	(4)	(5)	(6)
Mid-Sized City		0.003				0.007
		(0.006)				(0.006)
Rural		-0.020**				-0.012
		(0.008)				(0.009)
Mom x Post1975	0.039***		0.034***	0.033***		
	(0.008)		(0.008)	(0.008)		
Mom x Post1975 x Large City		0.033***			0.034***	0.034***
		(0.008)			(0.010)	(0.008)
Mom x Post1975 x Mid-Sized City		0.035***			0.044***	0.037***
		(0.009)			(0.010)	(0.009)
Mom x Post1975 x Rural		0.048***			0.039***	0.050***
		(0.007)			(0.008)	(0.009)
Region-Level Stigma				-0.006	-0.006	-0.005
				(0.005)	(0.005)	(0.005)
Mom x Post1975 x Stigma			-0.010**	-0.004*	-0.004**	-0.005**
			(0.005)	(0.002)	(0.002)	(0.002)
Observations	133,805	133,805	133,805	133,805	133,805	133,805

Panel B: Bivariate Regression of Region-Level DD Estimate on Region-Level Stigma

Weight Used:	None	Inverse of DD Standard Error	Sum of CPS Weights by Region
Variables	(7)	(8)	(9)
Region-Level Stigma	-0.007*	-0.008***	-0.008**
	(0.003)	(0.002)	(0.003)
	9	9	9
	0.289	0.412	0.295

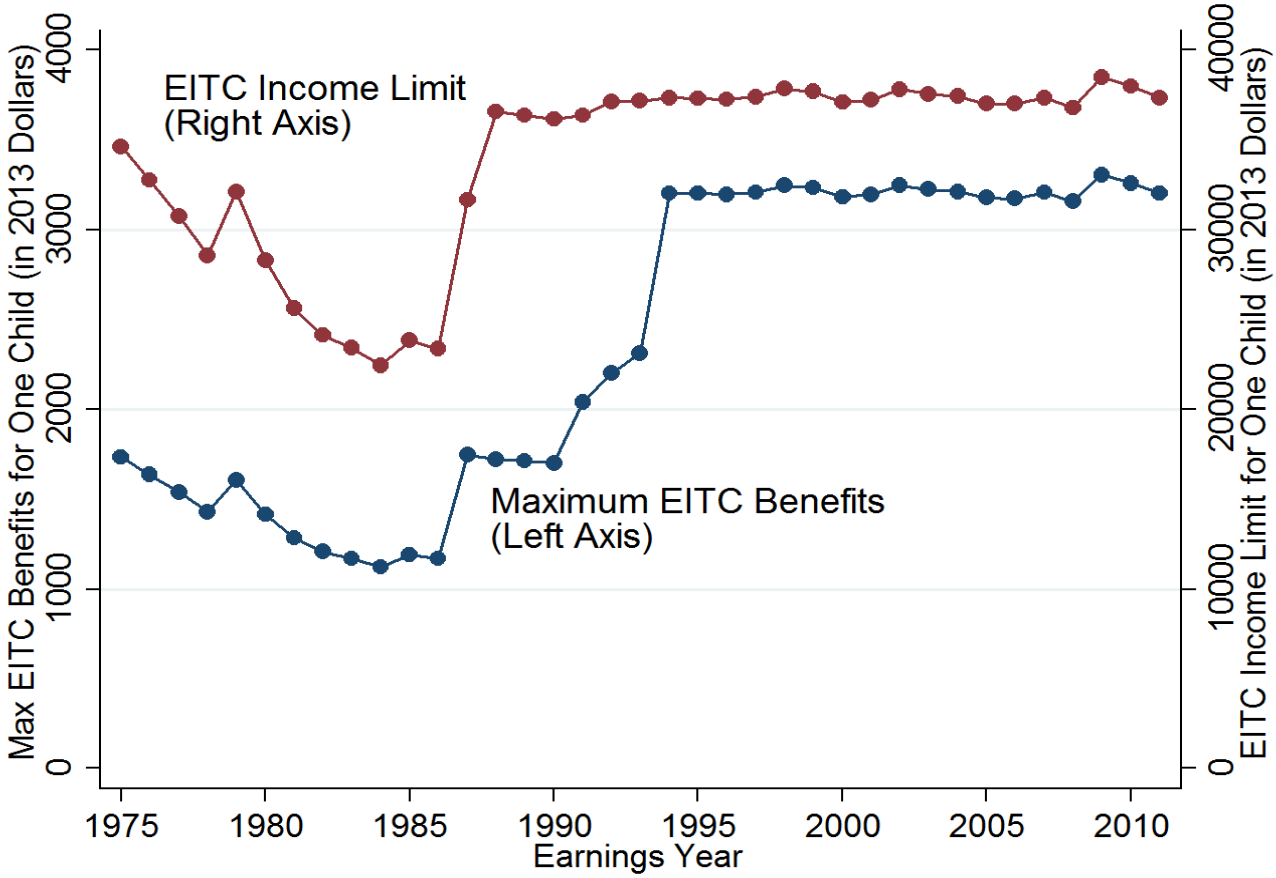
Notes: 1972-1981 March CPS data. "High-impact" sample used. Full set of controls used from Table 2 column 5. Stigma is created from region-level averages of standardized male responses to beliefs about whether women should work (originally scaled 1-7) in American National Election Studies data. As defined, regional sexism took values between -0.3 and 0.1. I multiply this variable by 10 so that stigma units are tenths of a standard deviation. Panel A assigns this region-level variable to all CPS observations and directly interacts it with the DD variable of interest (*mom x post1975)*. Panel B is a simple bivariate regression between the region-level DD estimate and region-level stigma. Both approaches provide evidence that regional stigma had a negative impact on the likelihood that a mother worked in response to the 1975 EITC. Robust standard errors computed by the delta method and clustered at the region level in Panel A only are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6: The EITC Effect on Annual Earnings and Work Hours

Dependent Variable:	Annual Work Hours				Annual Earnings (2013 Dollars)			
Sample of Women:	"High-Impact" Group	Less than High-School Education	All	EITC-Ineligible Placebo Group	"High-Impact" Group	Less than High-School Education	All	EITC-Ineligible Placebo Group
Mean Dependent Variable:	934.3	375.4	718	690.5	15347.6	5516.8	12750.3	12927.0
Mean Dependent Variable for Mothers in 1975:	769.3	451.3	591.3	529.9	12458.0	6506.0	10483.9	9852.5
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Mom	-203.6**	180.3	-267.6*	-865.2***	10,285.3***	3,770.5**	2,549.3	-13,132.3**
	(93.2)	(110.0)	(132.6)	(216.2)	(1,680.1)	(1,580.2)	(2,155.1)	(4,840.0)
Post1975	-69.7*	-68.5**	-0.8	104.0**	-2,481.7**	-1,609.3***	-1,341.1***	79.9
	(35.3)	(25.2)	(23.3)	(39.0)	(922.7)	(343.0)	(421.1)	(488.3)
Mom x Post1975	67.7***	41.3**	36.7***	-12.0	1,298.6***	899.8***	693.1***	-88.7
	(15.5)	(19.4)	(11.8)	(14.7)	(257.8)	(222.1)	(217.9)	(371.8)
	133,805	97,683	335,639	167,949	133,805	97,683	335,639	167,949
Observations	0.265	0.160	0.204	0.149	0.351	0.171	0.249	0.177

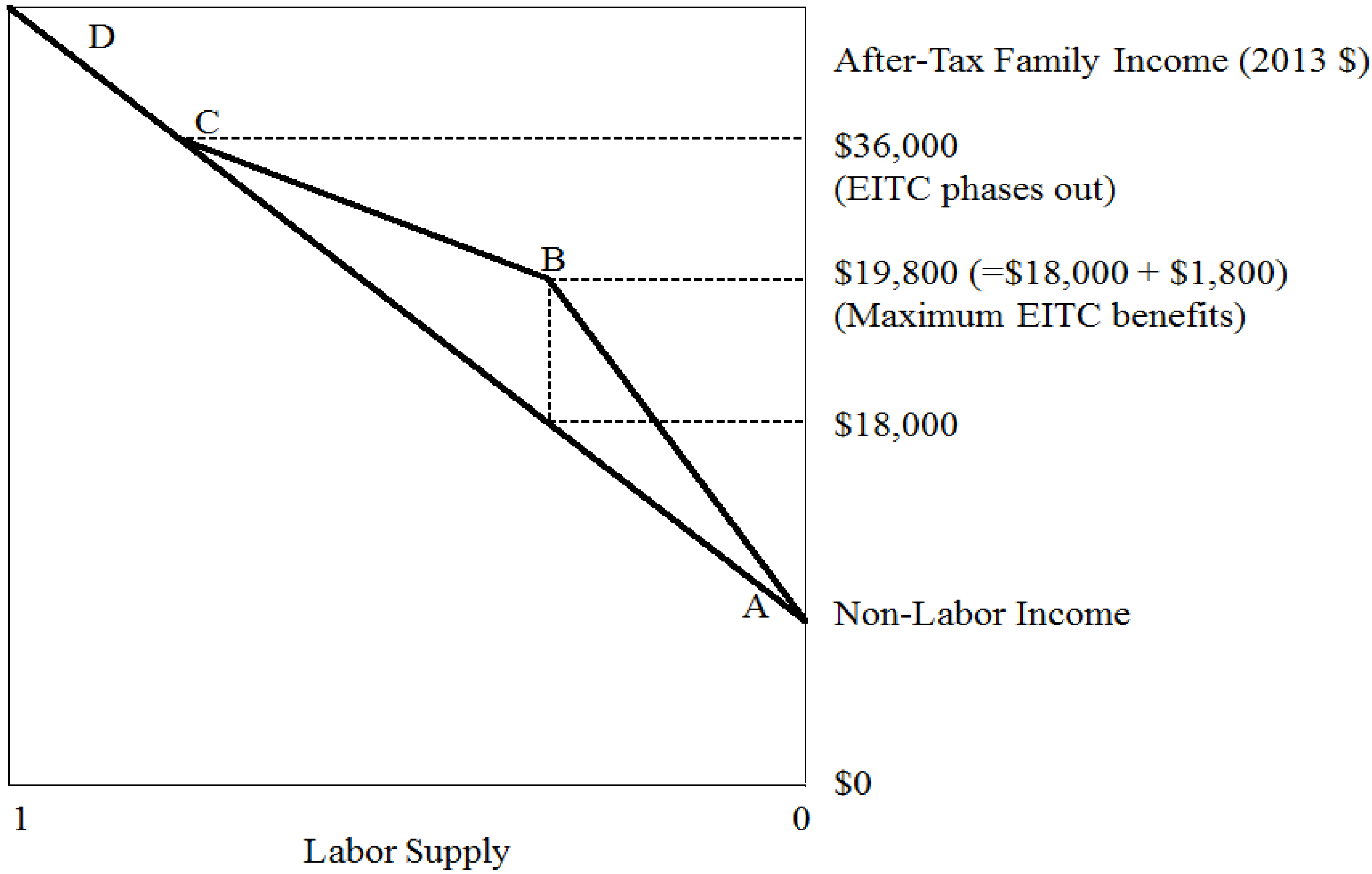
Note: *** p<0.01, ** p<0.05, * p<0.1. Data source: 1972-1981 March CPS. Binary dependent variable employment. CPS weights used and average marginal effects from probit regression are shown. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. Each column represents a separate regression with the full set of controls from Table 2 column 5. All samples are limited to women 16 to 45 years old. Annual work hours are constructed by multiplying weeks worked and hours worked last week. Weeks worked is given as an interval until 1975, so I assign the midpoint of the interval.

Appendix Figure A1: Trends in EITC Income Eligiblity Limit and Max Potential Benefits



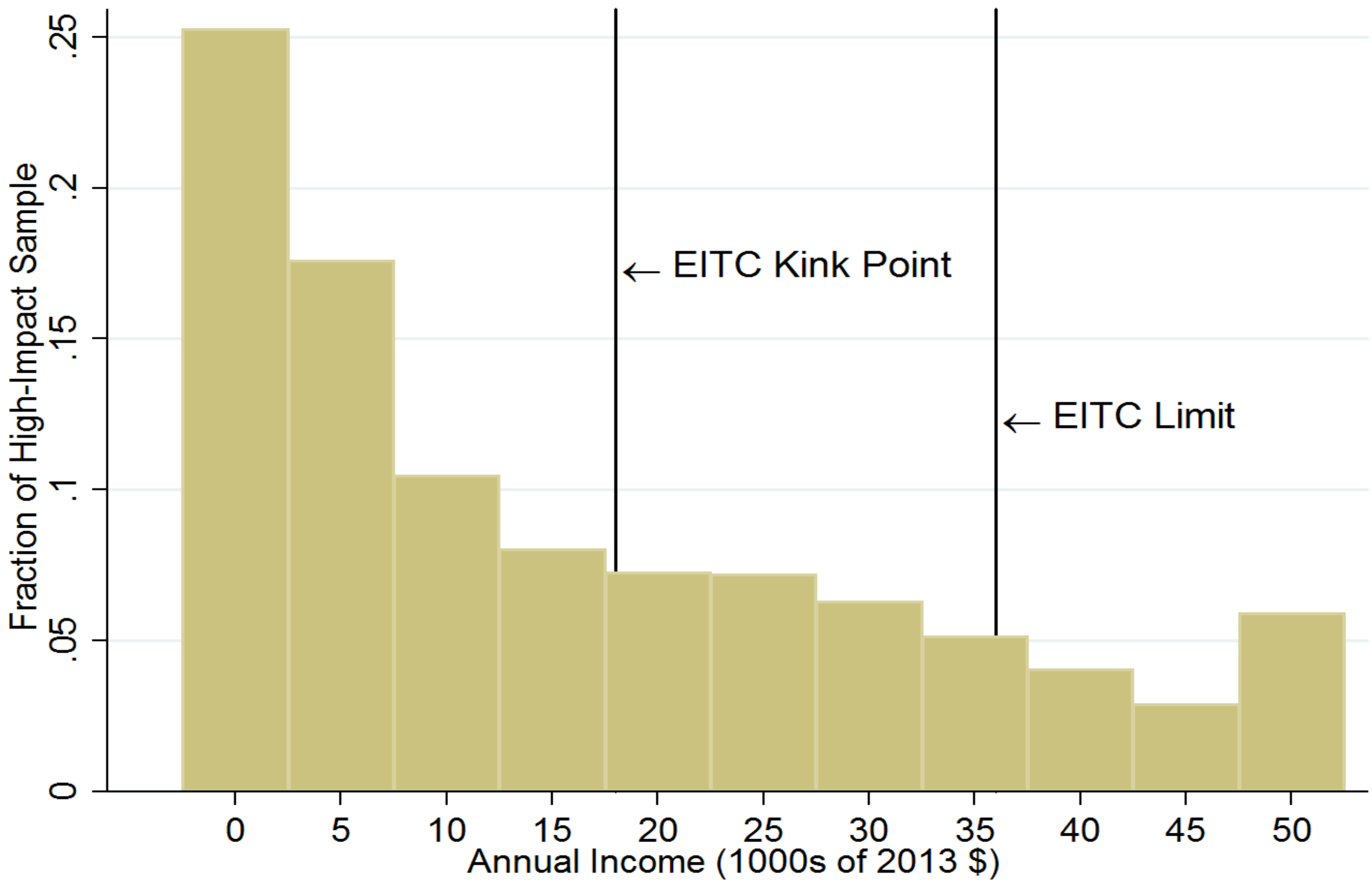
Notes: Author's calculations from IRS data.

Appendix Figure A2: Budget Constraint Under the 1975 EITC



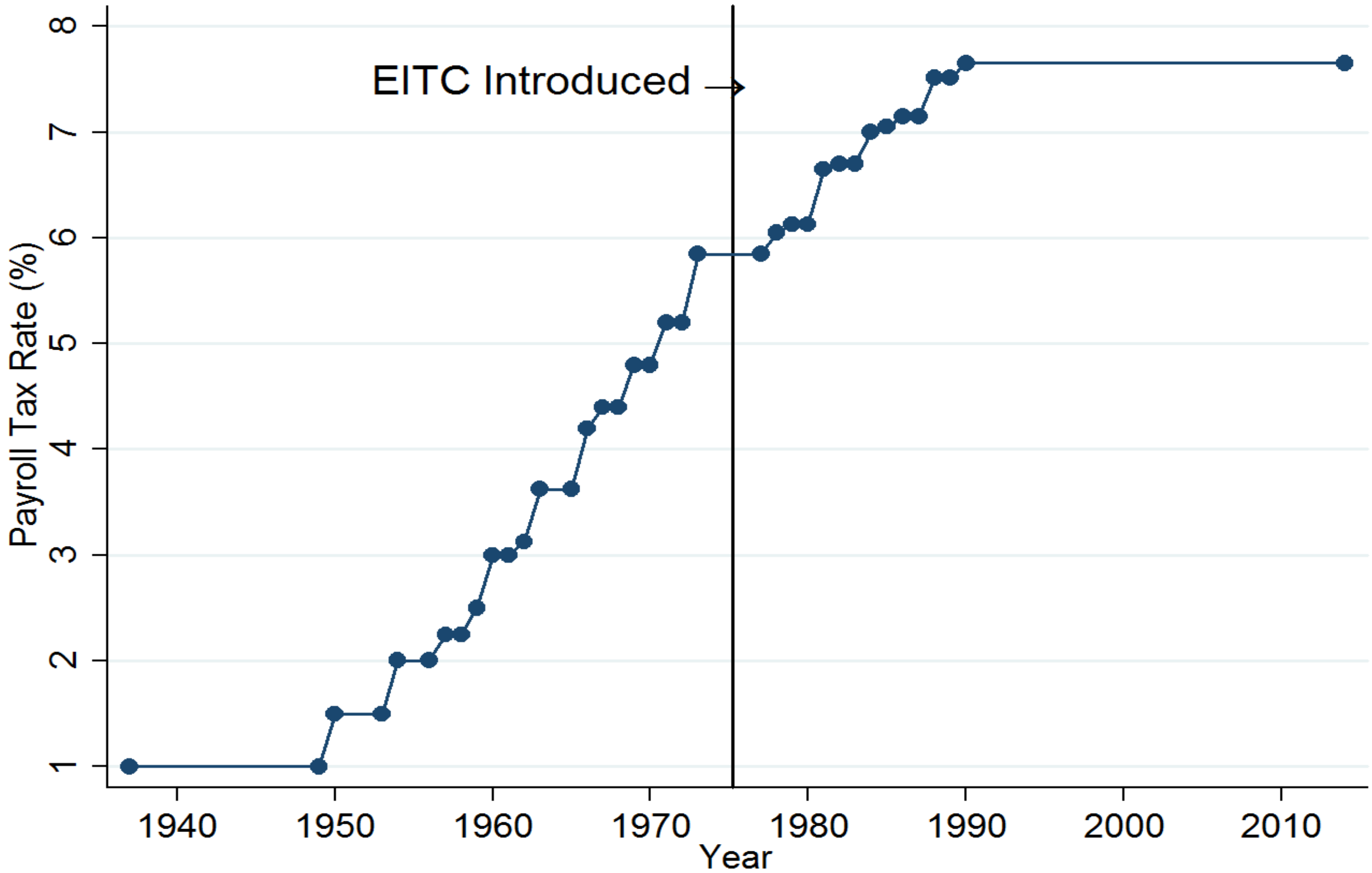
Notes: Author's calculation from 1975 EITC parameters.

Appendix Figure A3: Income Distribution of Females in "High-Impact" Sample



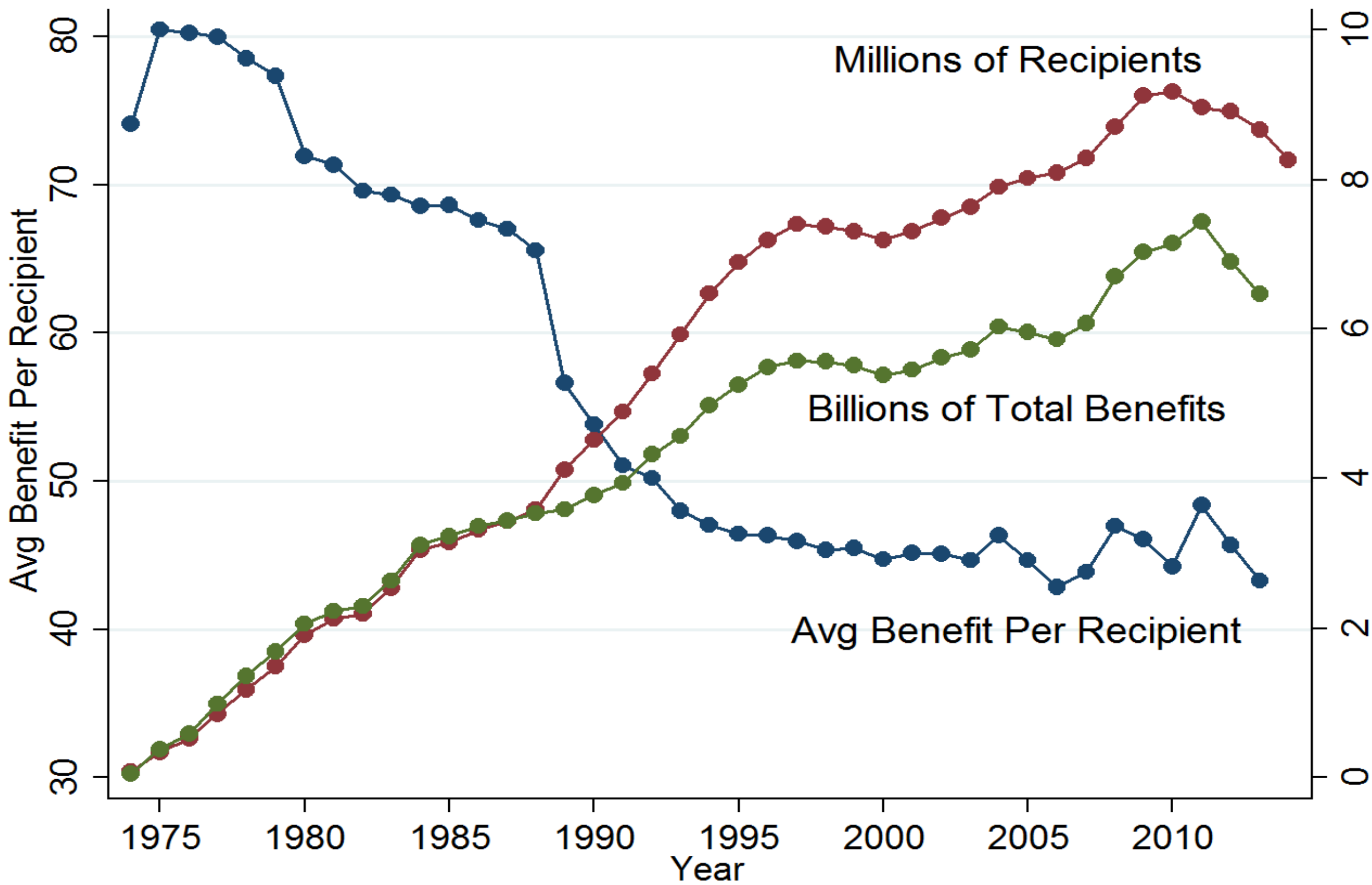
Notes: Author's calculation from 1972-1981 March CPS data. The "high-impact" sample includes women 16-45 excluding full-time students and married women with husbands earning over the EITC-kink point.

Appendix Figure A4: Ruling Out Confounding Policies (Payroll Tax Rate)



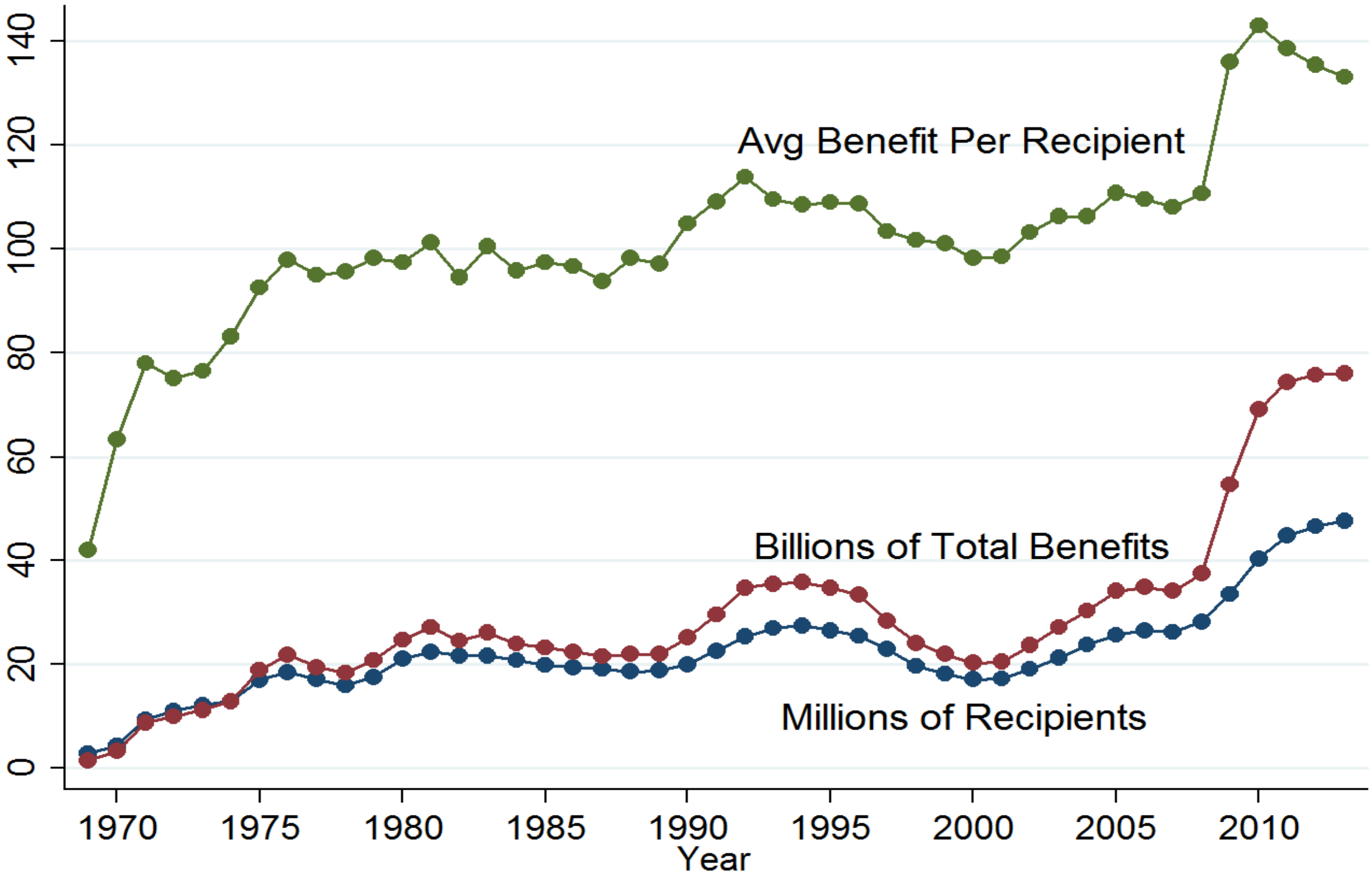
Notes: Author's calculation from payroll tax data.

Appendix Figure A5: Ruling Out Confounding Policies (WIC)



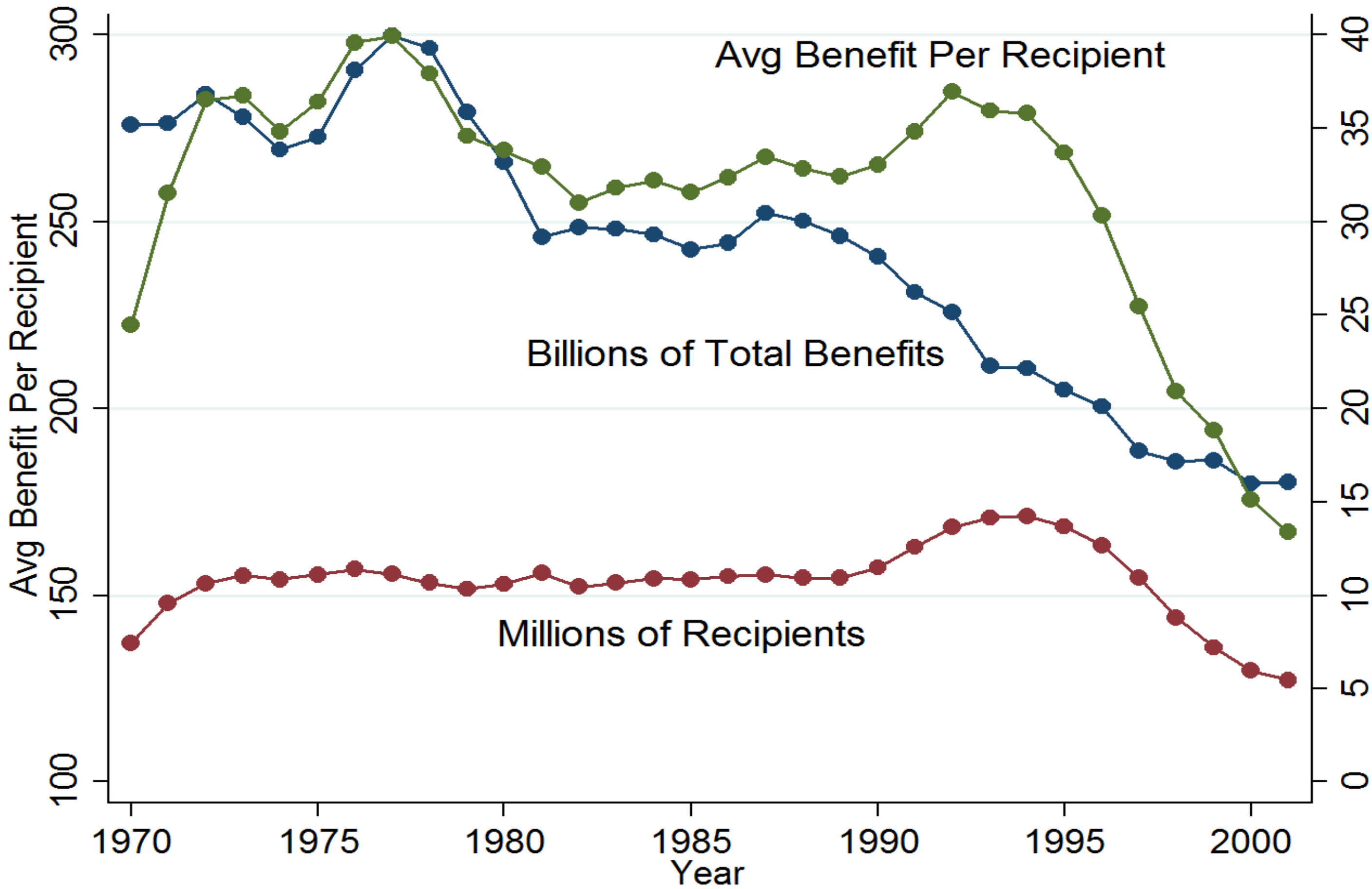
Notes: Author's calculation from WIC data.

Appendix Figure A6: Ruling Out Confounding Policies (Food Stamps)



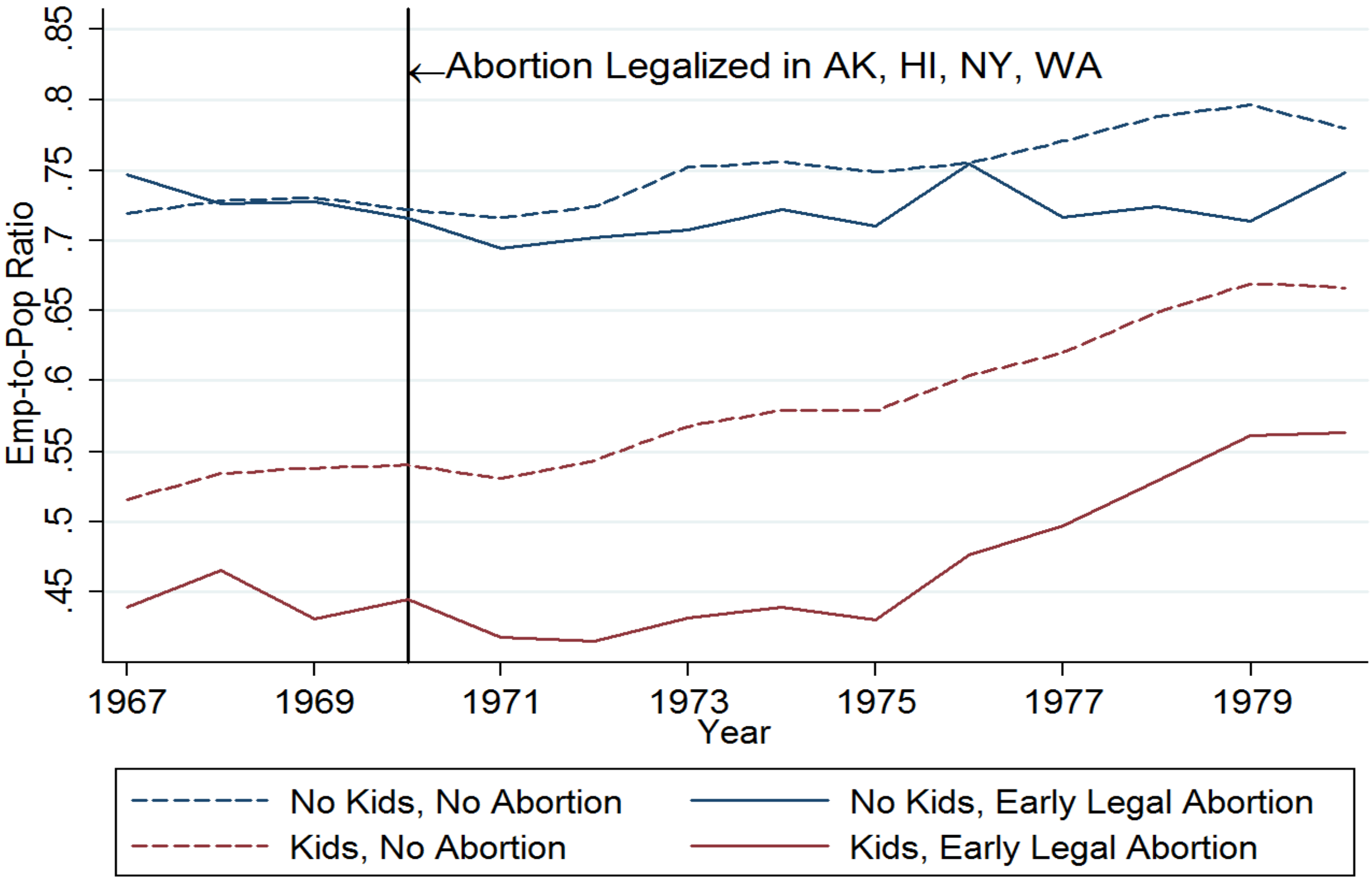
Notes: Author's calculation from Food Stamps data.

Appendix Figure A7: Ruling Out Confounding Policies (AFDC)



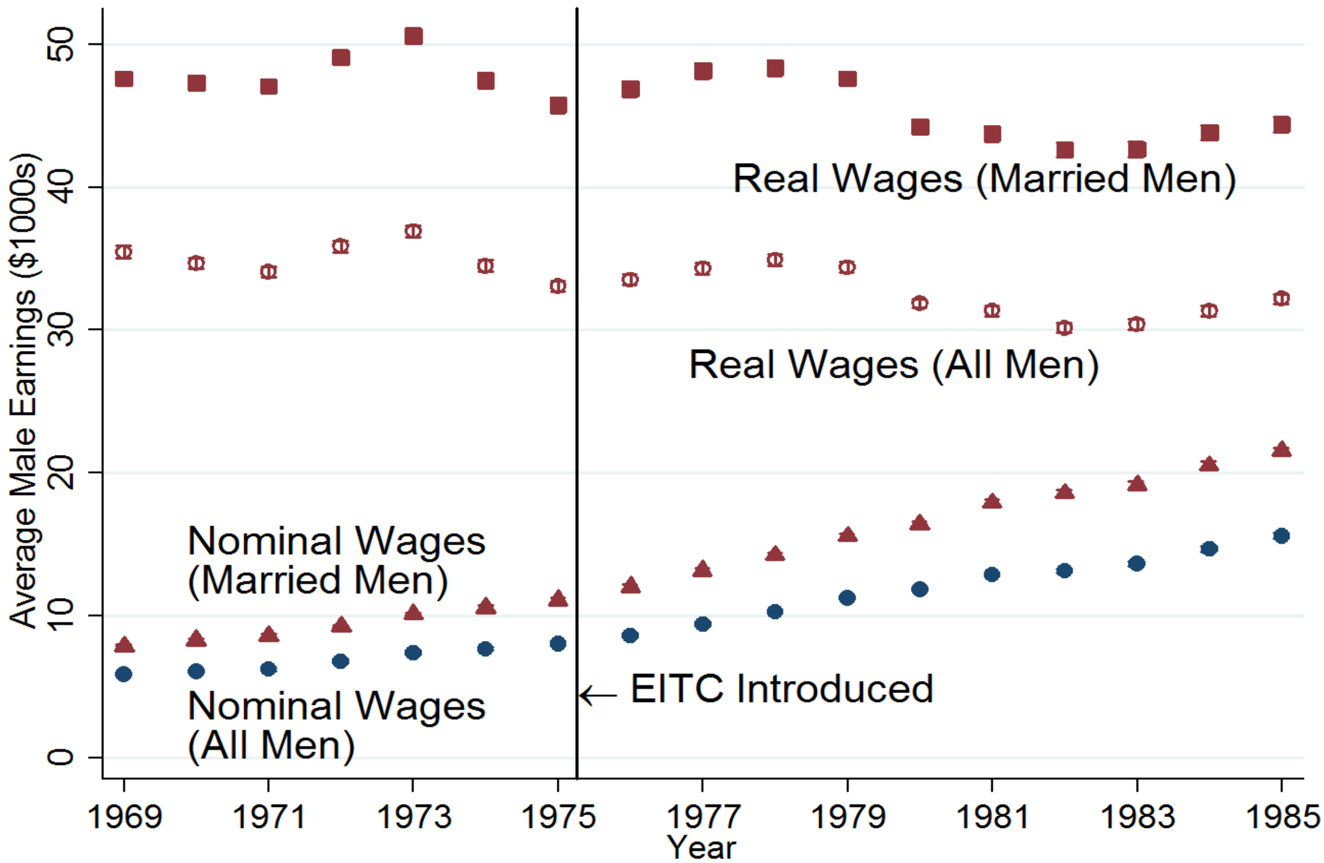
Notes: Author's calculation from AFDC and TANF data.

Appendix Figure A8: Ruling Out Confounding Policies (Abortion Legalized)



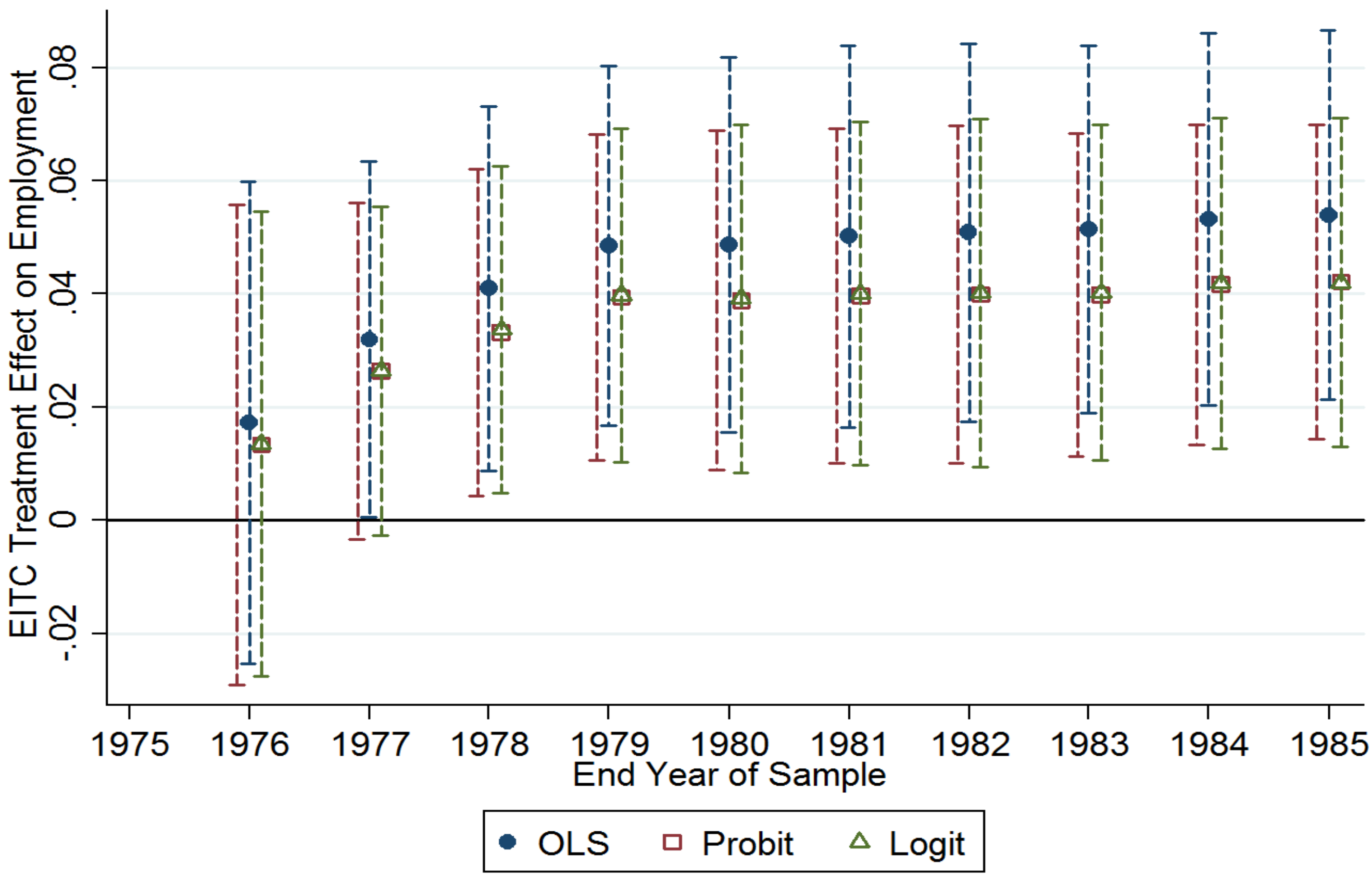
Notes: Author's calculation from 1970-1981 March CPS data. Though there appears to be a level difference in employment rates between the two states, there does not appear to be a detectable change in this gap after abortion is legalized in these four states.

Appendix Figure A9: Ruling Out Confounding Trends (Male Earnings)



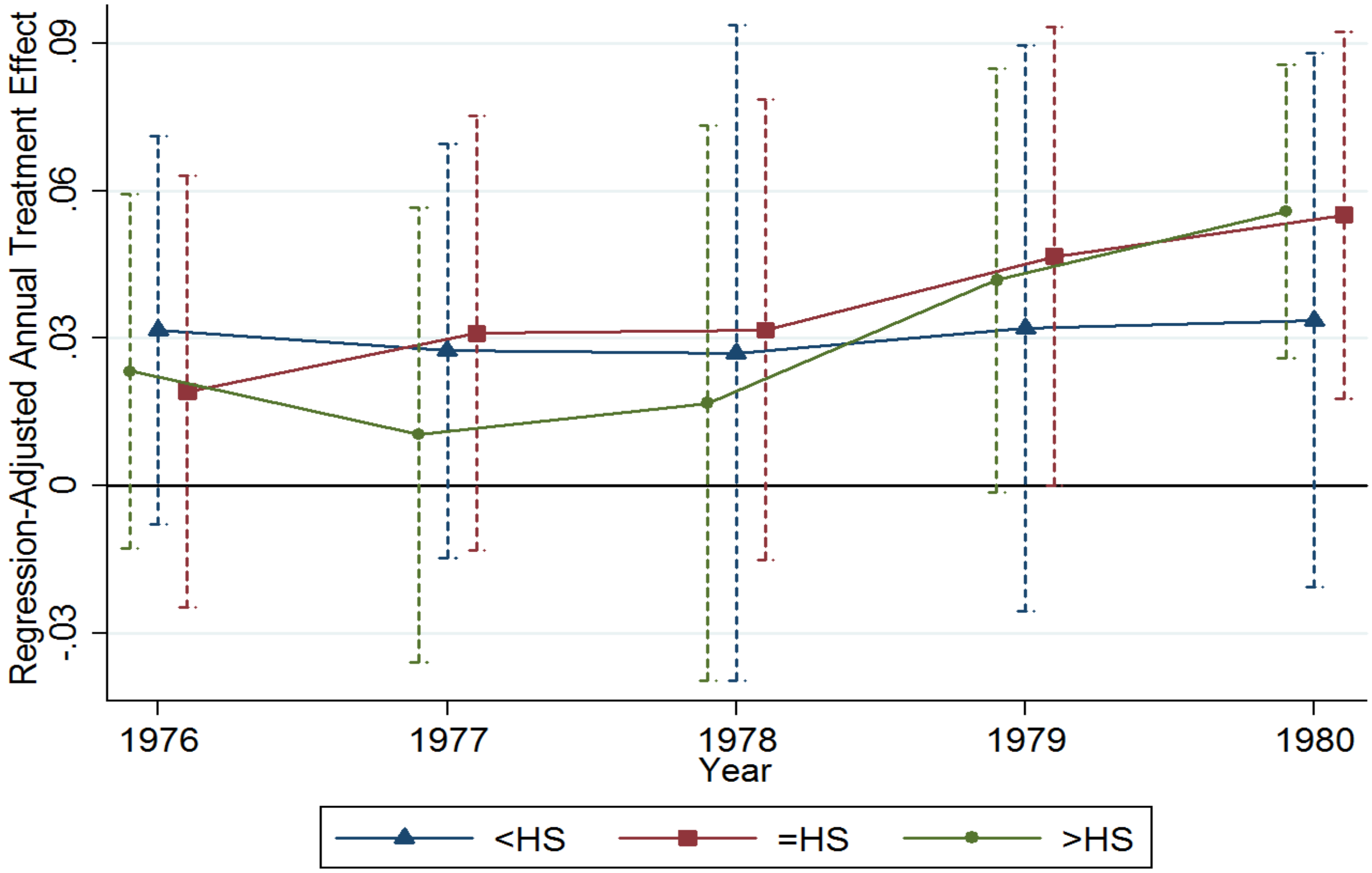
Notes: Author's calculation from 1970-1986 March CPS data.

Appendix Figure A10: DD Estimate Robust to Model Choice and Year that Sample Period Ends



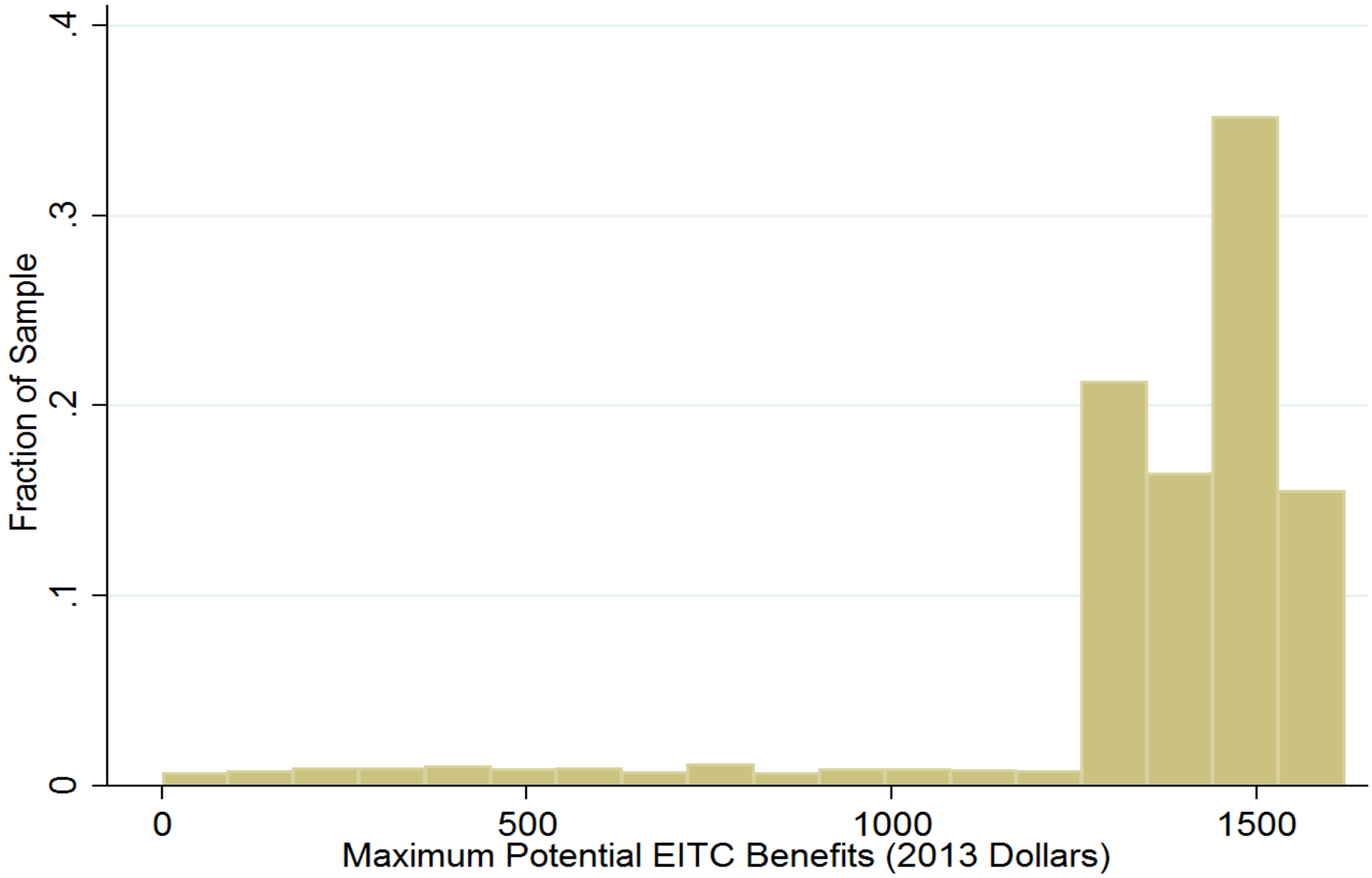
Notes: 1971-1986 March CPS data. Full set of controls used from Table 2 and "high-impact" sample used. Treatment effects are estimates of *Mom x Post1975* in equation (3). *Post1975* starts in 1976 and extends through the year specified on the x-axis. Binary dependent variable employment. CPS weights used and average marginal effects from probit regression are shown. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. 95 percent confidence intervals shown.

Appendix Figure A11: No Evidence that Higher-Educated Women Respond Quicker



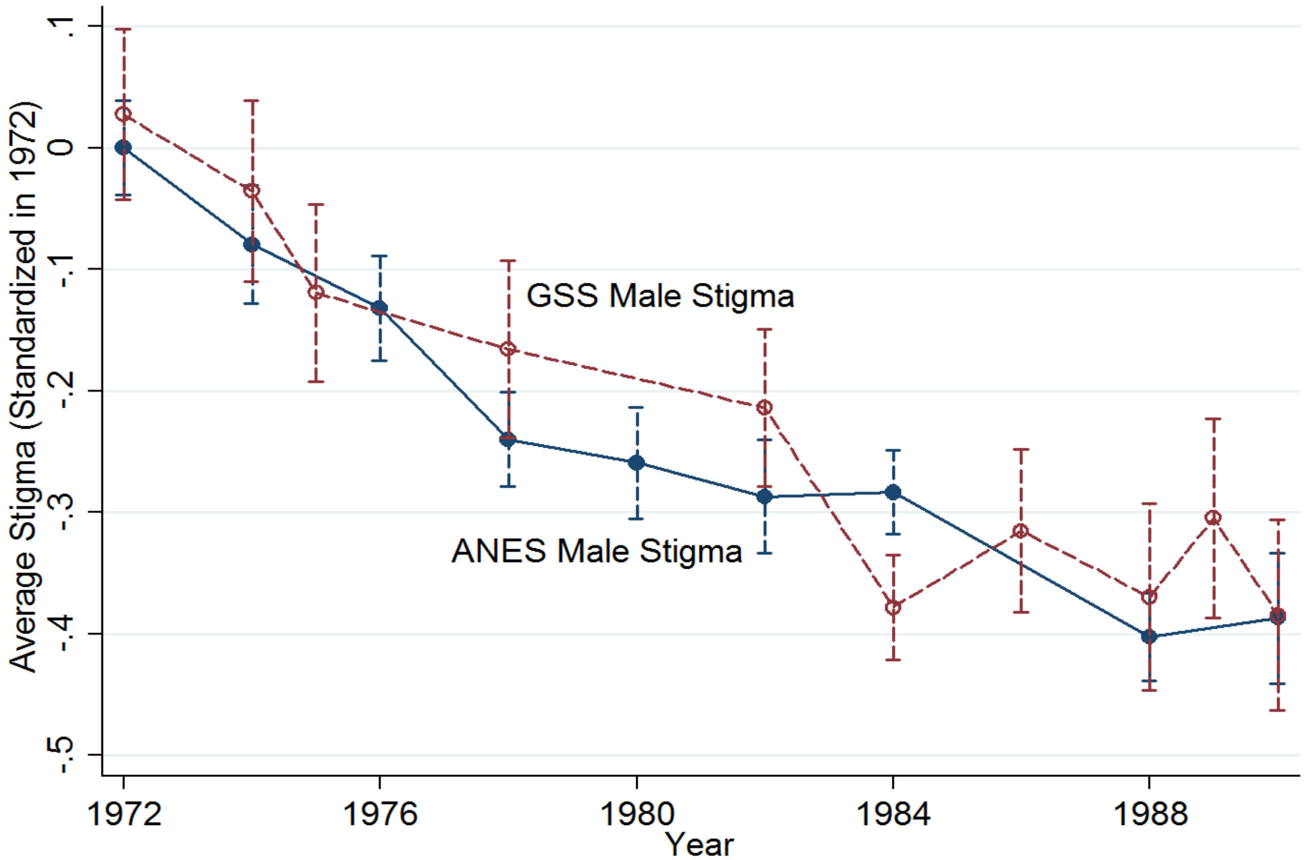
Notes: 1971-1986 March CPS data. Full set of controls used from Table 2 and "high-impact" sample used. Treatment effects are annual estimates of Mom x Year in equation (2) from three separate regressions on each education level. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. 95 percent confidence intervals shown.

Appendix Figure A12: Histogram of Maximum Potential EITC Benefits



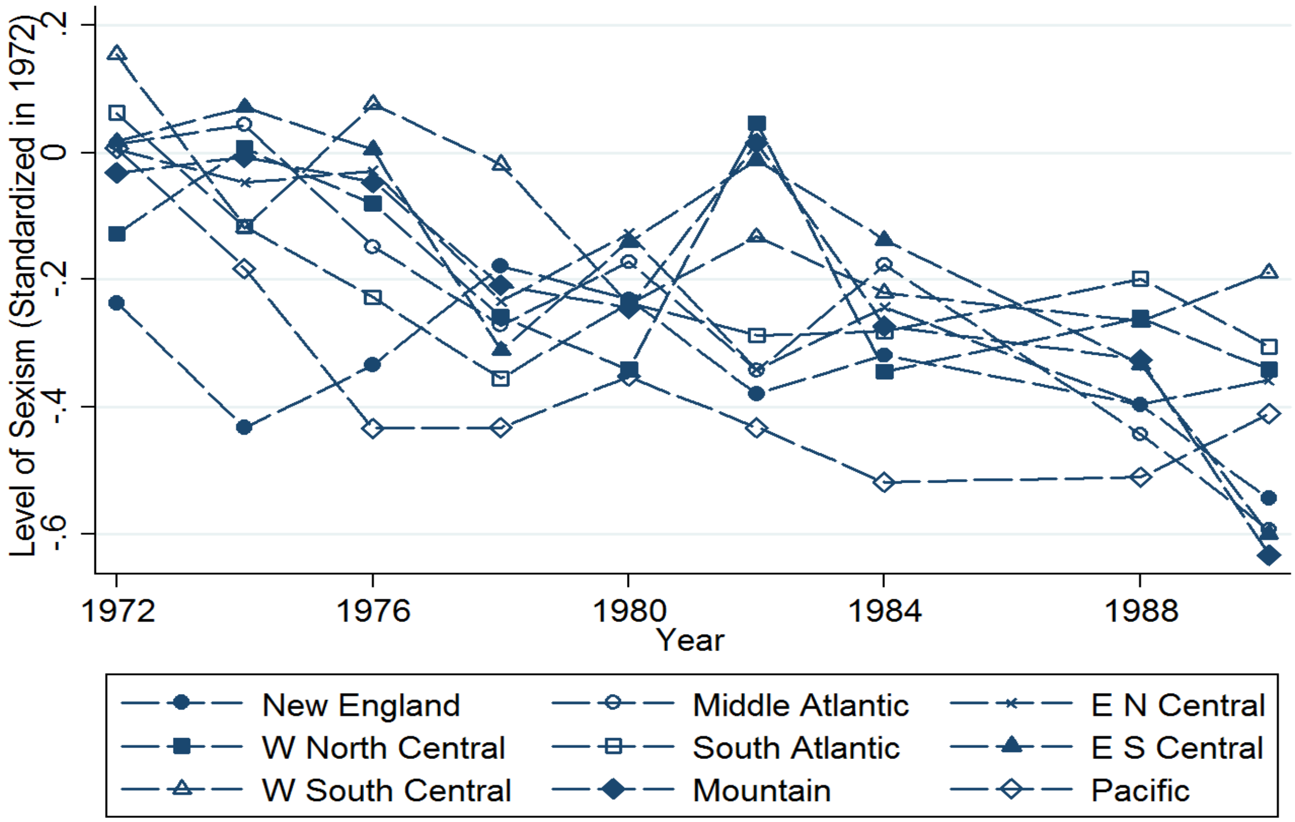
Notes: 1972-1981 March CPS data. Maximum potential EITC benefits (in 2013 dollars) determined by spousal income (zero if unmarried) and year. Women eligible for zero EITC benefits make up about 90 percent of the sample are are excluded from the histogram for scale purposes. In the full sample of 335639 women, *maxEITC* = 0 for 303,120 women, *maxEITC* = 1346 for 6663 women, *maxEITC* = 1361 for 5176 women, *maxEITC* = 1464 for 4935 women, *maxEITC* = 1528 for 6374 women, *maxEITC* = 1560 for 5024 women, and for 4347 married women with positive spousal earnings below the EITC kink point *maxEITC* is somewhat uniformly distributed between 0 and 1560.

Appendix Figure A13: Stigma Against Working Women Decreasing Over Time



Notes: The ANES "stigma" variable *equalrole* asks respondents whether men and women should have equal roles or whether a woman's role is in the home. These responses take values between 0 and 6, where 6 represents more stigma. The GSS "stigma" variable *fework* asks respondents whether women should work. This binary response is coded so that 1 represents more stigma. Both variables are standardized by subtracting the 1972 mean and dividing by the 1972 standard deviation. The sample in each dataset consists of all men and use 1972 through 1990 data. 1984 combines 1983 and 1985 data for visual purposes.

Appendix Figure A14: Regional Trends in the Stigma Against Working Women



Notes: The ANES "sexism" variable *equalrole* asks respondents whether men and women should have equal roles or whether a woman's role is in the home. These responses take values between 0 and 6, where 6 represents a more "sexist" view. "Sexism" is standardized by subtracting the 1972 mean and dividing by the 1972 standard deviation. The sample consists of all men in the 1972 through 1990 data.

Appendix Table A1: Treatment Effect Robust to Alternate Definitions of Working

Definition of Working:	Earnings >$0 (2013 $)	Earnings >$1000 (2013 $)	Earnings >$5000 (2013 $)	Earnings >$10000 (2013 $)	Work Weeks >0	Work Weeks >25	Labor-Force Participation
Mean Dependent Variable:	0.637	0.594	0.485	0.400	0.667	0.467	0.581
Mean Dependent Variable for Mothers in 1975:	0.530	0.502	0.417	0.345	0.563	0.395	0.495
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Mom	-0.056	-0.078	-0.151*	-0.220***	-0.043	-0.174**	-0.097
	(0.101)	(0.097)	(0.078)	(0.057)	(0.091)	(0.077)	(0.084)
Post1975	0.001	0.001	0.002	-0.007	-0.001	0.005	0.035***
	(0.014)	(0.013)	(0.012)	(0.010)	(0.015)	(0.010)	(0.008)
Mom x Post1975	0.033***	0.029***	0.026***	0.023***	0.031***	0.027***	0.029***
	(0.009)	(0.009)	(0.008)	(0.006)	(0.008)	(0.008)	(0.007)
Observations	335,639	335,639	335,639	335,639	335,639	335,639	335,639

Note: *** p<0.01, ** p<0.05, * p<0.1. Data source: 1972-1981 March CPS. Binary dependent variable employment. CPS weights used. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. Each column represents a seperate regression with the full set of controls from Table 2 column 5. All samples are limited to women 16 to 45 years old.

Appendix Table A2: Probit, Logit, OLS and the "Kitchen-Sink" Set of Controls

	Mean Dependent Variable=0.637			
	Mean Dependent Variable for Treatment Group in 1975=0.530			
Model:	Probit	Logit	Ordinary Least Squares	
Variables	(1)	(2)	(3)	(4)
Kid	-0.056	-0.032	-0.100	-0.203**
	(0.101)	(0.109)	(0.081)	(0.086)
Post1975	0.001	-0.001	0.000	2.503***
	(0.014)	(0.014)	(0.013)	(0.176)
Kid*Post1975	0.033***	0.033***	0.040***	0.037***
	(0.009)	(0.009)	(0.008)	(0.008)
Full Set of Controls	X	X	X	X
"Kitchen-Sink" Controls				X
Observations	335,639	335,639	335,639	335,639
R-squared	--	--	0.153	0.175

Note: *** p<0.01, ** p<0.05, * p<0.1. Data source: 1972-1981 March CPS. OLS regressions and binary dependent variable employment. CPS weights used. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. This "high-impact" sample includes women 16-45 excluding full-time students and married women with husbands earning over the EITC-kink point. This table mirrors and extends Table 2 and columns 1 and 2 here are identical to Table 2 columns 1 and 5 except OLS is used instead of a probit. Column 1 contains no controls; column 2 controls described in Table 2. Column 3 is the "kitchen-sink" regression and in addition to the controls in column 2 also includes: unemployment rate-age, nonwhite-welfare, nonwhite-married, number children-married, child less than 5-married, married-welfare income, education years-married, education-child less than 5, education-nonwhite, a nonwhite-age cubic, unemployment rate-nonwhite, and fixed effects for nonwhite-year, married-year, nonwhite-state, birth-year, state-year, state-married, state-child less than 5, state-year-nonwhite, state-year-married, metropolis status (roughly equivalent to rural, suburban, urban), and metro-nonwhite, metro-year, metro-state, metro-child less than 5, metro-state-year. These additional controls explain about 10 percent of the DD estimate (but see Gelbach 2016 about interpreting this), but the effect of the EITC on maternal employment is still positive and significant.

Appendix Table A3: Married Women with Lower-Earning Spouses Also Responded to the EITC in the 1980s and 1990s

Sample:	1986 EITC Expansion				1993 EITC Expansion			
Dependent Variable:	Employed		Weekly Work Hours		Employed		Weekly Work Hours	
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Mom	-0.260***	-0.279***	-8.799*	-9.619**	-0.203***	-0.206***	-11.274***	-11.691***
	(0.100)	(0.102)	(4.693)	(4.650)	(0.065)	(0.064)	(3.622)	(3.532)
Post1975	0.012	0.016	0.727	0.890*	0.028***	0.030***	1.192***	1.423***
	(0.014)	(0.014)	(0.509)	(0.501)	(0.006)	(0.005)	(0.285)	(0.264)
Mom x Post1975	0.019	0.053***	0.775**	2.701***	-0.002	0.034***	-0.037	1.711***
	(0.011)	(0.011)	(0.373)	(0.413)	(0.007)	(0.007)	(0.297)	(0.362)
Kid x Post1975 x		-0.001***		-0.048***		-0.001***		-0.036***
Spousal Income		(0.000)		(0.006)		(0.000)		(0.003)
(in 1000s of 2013 $)								
Observations	116,914	115,194	116,914	115,194	175,893	173,241	175,893	173,241

Note: Columns 1-4 follow Eissa and Liebman (1996) and examine the effect of the 1986 EITC expansion on the employment of 16-44 year old females using 1985-1987 and 1989-1991 March CPS data. Columns 5-8 follow Eissa and Hoynes (2004) and examine the effect of the 1993 EITC expansion on the employment of 25-54 year old females using 1989-1996 March CPS data. Binary employment is defined as having positive hours of work and weekly work hours refers to the week prior to the March CPS interview. Regressions for binary employment reflect average marginal effects from a probit and weekly work hours reflect estimates from OLS. In each regression the full set of controls is used from Table 2 column 5 and CPS weights are used. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. Some married women did not have information on spousal earnings. *** p<0.01, ** p<0.05, * p<0.1.

Appendix Table A4: Traits Correlated with Social Stigma Against Women Working

Social Stigma Standardized in 1972 (Mean=0, SD=1) (Positive Implies More Stigma)

Panel A: American National Election Surveys Data

VARIABLES	(1)	(2)	(3)	(4)	(5)
Age / 10	0.062***				-0.136
	(0.006)				(0.137)
(Birth Year - 1900) / 10		-0.064***			-0.174
		(0.006)			(0.139)
Education Years			-0.066***		-0.060***
			(0.004)		(0.004)
Time Trend				-0.021***	-0.000
				(0.002)	(0.014)
Observations	6,143	6,143	6,143	6,143	6,143
R-squared	0.041	0.041	0.068	0.024	0.070

Panel B: General Social Survey Data

VARIABLES	(6)	(7)	(8)	(9)	(10)
Age / 10	0.013***				0.011***
	(0.001)				(0.003)
(Birth Year - 1900) / 10		-0.129***			0.019
		(0.008)			(0.029)
Education Years			-0.077***		-0.064***
			(0.003)		(0.003)
Time Trend				-0.029***	-0.025***
				(0.002)	(0.002)
Observations	6,555	6,555	6,558	6,576	6,541
R-squared	0.085	0.083	0.107	0.028	0.127

Notes: The sample in Panel A consists of all men in the 1972 to 1988 ANES data. "Social stigma" in Panel A comes from the ANES variable *equalrole* , which asks respondents whether men and women should have equal roles or whether a woman's role is in the home. These responses are scaled between 0 and 6, where 6 represents more stigma, and then standardized by subtracting the 1972 mean and dividing by the 1972 standard deviation. This results in a mean and standard deviation of -0.20 and 0.89. The sample in Panel B consists of all men in the 1972 to 1988 GSS data. The dependent variable comes from the GSS variable *fework* , which asks respondents whether women should work. This binary response is coded so that 1 represents more stigma and is standardized by subtracting the 1972 mean and dividing by the 1972 standard deviation. The mean and standard deviation of the dependent variable is -0.19 and 0.92. Correlations in each panel show that the following types of people have lower stigmas against working women: younger people, people born more recently, people with more education, and the trend in sexism has been declining over time. Time trend starts in 1972. Robust standard errors clustered at the region level in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Appendix B: Additional Robustness Check

These robustness checks provide further evidence that the employment increase of mothers after 1975 does appear to be due to the EITC and not to general trends or changes in group composition.

I. DDD Corroborates DD Treatment Effects

DD results would be biased if an unaccounted for policy or trend affected the relative employment of all mothers. If, for instance, a law passed that made it cheaper for all mothers to work after 1975, a triple differences (DDD) approach could net out this effect.[1] DDD is akin to estimating equation (2) twice, once for a treatment group eligible for the EITC (e.g. single mothers and married mothers with lower-earning spouses) and again for a control group largely ineligible for the EITC (e.g. married women with higher-earning spouses).[2] The difference between these two DD estimates is the intuition behind the DDD estimate. If the DD estimate for the control group is near zero, then the DDD estimate of the maternal employment response to the EITC should resemble the DD estimate.

Table 4 column 5 shows that married women with spouses earning above the EITC limit did not respond to the EITC. During this time the labor-supply elasticity of men was near zero (especially on the extensive margin), so the EITC should not have affected the employment of men. These two groups contain individuals with and without kids, observed in years before and after 1975, and will be used as DDD control groups. The following DDD probit model uses the full set of controls from Table 2 column 5 and extends equation (3)[3].

(B1) $$E_{i,s,t} = \beta_1 Mom\ x\ Post\ x\ Treat_{i,s,t} + \beta_2 X_{i,s,t} + \gamma_s + \delta_t + \epsilon_{i,s,t}$$

β_1 is the DDD coefficient of interest. When married women with higher-earning spouses are used as the control group, $Treat$ equals zero for these women; when men are used as the control

[1] This is one reason that Angrist and Pischke (2008, p. 182) state that a DDD model "may generate a more convincing set of results" than a simple DD estimator.
[2] Except that covariate estimates in two separate DD regressions can differ, unlike in a single regression.
[3] Each subcomponent and double interaction of $Mom\ x\ Post\ x\ Treat$ are also controlled for and excluded for visual ease.

group, $Treat$ equals zero for all men.

Appendix Table B1 columns 1 to 4 use married women with higher-earning spouses as the control group. Column 1 uses *all* women with spouses earning over the EITC kink point as the control group, and columns 2 to 4 restrict the control group to women with spouses earning above the EITC kink point (about $18,000) and below $60,000, $50,000, and $40,000 (2013 dollars). Married women with spousal earnings just above the EITC limit should serve as a better control group and be more similar to women in the treatment group (which consists of married women with spouses earning below the EITC limit and unmarried women). Columns 1 to 4 show that the four DDD estimates are stable between 2.4- and 3.1-percentage points. Appendix Table B1 columns 5 and 6 use all men and all single men (16 to 45 years old) as the control group. The underlying DD for each group of men (not shown) is near zero and the resulting DDD estimates in these two columns are 3.4- and 3.9-percentage points.

These DDD estimates all fall between 2.4- and 3.9-percentage points and are not statistically different that the DD estimate in Table 3 column 2 (3.3-percentage points) for all 16- to 45-year-old women. This is further evidence that the estimated DD treatment effect is not being driven by a general trend or alternative policy that affected all mothers or all women. These DDD results also confirm that the EITC earnings subsidy did not affect the (largely inelastic) male labor supply.

II. Accounting for Potentially Endogenous Fertility and Changes in Group Composition

The EITC is only available to working adults with children and therefore might impact fertility choices. Although theory would predict a positive effect on the probability of having children[4], existing evidence shows little to no responses on this margin.[5] One simple way to account for potentially endogenous fertility after 1975 is by restricting the sample to women with kids born before 1975 to see if the estimated treatment effect is affected. Doing so barely changes the DD coefficient of interest from 3.3- to 3.2-percentage points (results not shown).

[4] There are no extra EITC benefits to having more than one kid until 1991, so there should not be an incentive on the intensive margin of fertility.

[5] Baughman and Dickert-Conlin (2003) and (2009) find small positive and negative fertility responses during the 1990s. Related, Ellwood (2000) finds no effect on marriage or cohabitation, Dickert-Conlin and Houser (2002) find no impact on single women and a very small negative effect on married women, and Herbst (2011) finds a small reduction in new marriages. Eissa and Hoynes (2004) summarize the empirical evidence as showing very small responses on these margins.

Another way to account for endogenous fertility, marital status, and group composition, is by reweighting mothers observed after 1975 to look like mothers before 1975. Although regression controls should largely account for any changing composition of mothers over time, reweighting is an additional robustness check.[6] Two sets of weights are used, one is detailed in DiNardo, Fortin, and Lemieux (1996) (DFL) and the other is inverse propensity (IP) weights. I generate these weights in the following two step process: First I use a logit[7] to estimate the probability than a given observation in the sample is from a year before 1975.

$$\text{(B2)} \qquad Pre1975_{i,s} = \beta_1 Mom_{i,s,t} + \beta_2 Race_{i,s,t} + \beta_3^j Age_{i,s,t}^j + \beta_4^k Educ_{i,s,t}^k + \beta_5^l State_{i,t}^l + \beta_6 Married_{i,s,t} + \epsilon$$

Using the sample of all women 16 to 45 in the 1971 to 1980 March CPS, equation (B2) predicts the likelihood that a given observation is from a year before the 1975 EITC was enacted. Following DiNardo, Fortin, and Lemieux (1996), I choose a parsimonious set of observable characteristics X, which includes six age bins, three education bins, married and nonwhite dummy variables, and state.[8] Each observation in the sample is assigned a probability of being from a year before 1975. This probability p is then used to create DFL and IP weights by assigning each observation a weight of $p/(1-p)$ and $1/p$. Women are weighted down if they are *less* likely to be observed before 1975 than after 1975 (e.g. a woman with a college education) and weighed up if they are *more* likely to be observed before 1975 (e.g. a woman with low education or high fertility).

In Appendix Table B2 I find that the baseline DD estimate of 3.3 percentage points is largely unchanged at 2.9- and 3.2-percentage points using DFL and IP weights. Since reweighting ensures that the composition of women after 1975 will be “identical” to the composition of women before 1975, this is further evidence that the relative increase in maternal

[6] DiNardo (2002) recommends reweighting as a robustness check. Reweighting analysis requires that the characteristics of the different populations overlap sufficiently and have a common support (Busso, DiNardo, and McCrary 2013), which is shown in Appendix Figure B1, and that any differences between the two groups can be captured by observable characteristics.

[7] The logit has the advantage over a probit in that the sum of predicted values equals the sum of the empirically observed ones (Butcher and DiNardo 2002). Although the results are very similar if a probit is used instead of a logit.

[8] DiNardo, Fortin, and Lemieux (1996) utilize a parsimonious set of controls that contains only 32 education-experience-gender cells. Butcher and DiNardo (2002) utilize several covariates which yields many more cells. My choice results in 1512 cells, although results do not change much with alternate decisions.

employment was due to the EITC and not a change in observable characteristics.

III. March CPS Imputations

In 1975 the Census changed its *hot deck* procedure[9] for imputing missing earnings (Bound and Freeman 1992). This could affect the results since I define employment as having positive income (although Appendix Table A1 shows similar estimates for other definitions of working based on hours and employment status). The percentage of imputed observations in the sample is zero before 1975, but between 1975 and 1980 is 6.6 percent, 6.3 percent, 6.3 percent, 5.3 percent, 1.0 percent, and 1.0 percent.[10] In Appendix Table B3 I show that the DD estimate is robust to various ways of treating imputed observations. Column 1 shows the baseline DD estimate of 3.9-percentage points from Table 3 column 2, column 2 drops all imputed observations and finds that the DD estimate rises to 4.0-percentage points. In columns 3 and 4 I use a logit to predict the probability than an observation is missing and account for selection ("missing not at random") by reweighting each observation with DFL and IP weights (see discussion in part II of this section). These yield DD estimates of 3.4-percentage points and 4.5-percentage points. Columns 5 and 6 are a bounding exercise and assume that none or all observations with imputed earnings are working (similar to Manski bounds), and yield DD estimates of 4.5- and 3.9-percentage points.

IV. Reconciling Constant EITC Recipients with Increasing Employment due to the EITC

Results in this paper show that the EITC encouraged hundreds of thousands of mothers to enter the labor force (and presumably earn EITC benefits). However, Appendix Figure B2 shows that the number of EITC recipients and the aggregate EITC benefits remained roughly constant between 1975 and 1985. At first glance these two trends appear to be contradictory, however,

9 In this process people with missing information are matched with similar people (based on sex, race and ethnicity, household relationship, years of school completed, geographic area, age, disability status, presence of children, veteran status, work experience, occupation, class-of-worker status, earnings, and value of property or monthly rent) and assigned the same value for the missing variable (IPUMS: https://usa.ipums.org/usa/voliii/80editall.shtml#note1).

10 Imputations fall into three categories: where it is unclear whether the individual had positive earnings, or what the actual amount of earnings was, or both. Only in the third case is it unclear whether the individual worked or not. This category is what I refer to as imputations.

they can be reconciled by looking at women already working before the EITC was introduced that had earnings just below the EITC limit. Inflation was high in the 1970s, nominal incomes were quickly rising, and the EITC schedule would not be pegged to inflation until 1986. Some women that were already working in 1974 or 1975 may have initially received EITC benefits, but within a few years had seen their nominal wages rise beyond the EITC limit, akin to "bracket creep" in the tax literature (see Saez 2003 for more details).

To approximate how many working women became ineligible for the EITC due to rising nominal income, I use the 1974 female income distribution from the CPS, and inflate these nominal earnings by the annual inflation rates of 1976 to 1980.[11] This assume constant real earnings, rising real wages would inflate these numbers more and yield even more "bracket creep" out of EITC eligibility. I then calculate the percentage of working women whose nominal income rendered them EITC-*eligible* in 1975 (earnings below nominal $8,000) but would have been EITC-*ineligible* in 1976, or in any other year between 1977 and 1980, due to rising nominal income. Appendix Figure B3 shows a zoomed-in view of this income distribution and the income cutoffs that would "bracket-creep" out of EITC-eligibility in 1976, 1977, and 1978. The percentage of women already working before the introduction of the EITC that would "bracket-creep" out of EITC eligibility was 1.9 percent, 2.9 percent, and 5.2 percent between 1976 and 1978. Since about 66 percent of women were working in 1975, these numbers correspond to percentage points of 1.3, 1.9, and 3.4, which are quite similar to the annual DD estimates in Figure 3. The number of women that would bracket-creep out of EITC eligibility approximates the annual number of mothers entering the labor force due to the EITC. Even though the "stock" of EITC-eligible women may have remained roughly constant in the years after 1975, there was substantial flow in and out of EITC eligibility. I conclude that constant EITC recipients and aggregate benefits can be reconciled with increasing employment due to the EITC.

V. Additional Response from Women with Multiple Children

The EITC did not provide additional benefits to working parents with more than one child until 1991. However the cost of working (e.g. childcare) likely increases with the number of children, and equation (2) accounts for this with a cost of $\gamma_{i,s,t}[l * k_i]$. Therefore mothers with multiple

[11] These numbers are 6.9 percent, 4.9 percent, 6.7 percent, 9.0 percent, and 13.3 percent.

children should work less than women with at most one child, and should not have responded to the EITC any more than women with only one child. I test this prediction with the following probit model – resembling equation (3) – which accounts for any differential impact on employment from having at least *J* kids.

$$\text{(B3)} \quad P(E_{i,s,t}) = \Phi(\beta_1 Post_{i,s} + \sum_{k=1}^{J}[\beta_{2,k} Mom_{i,s,t}^{k} + \beta_{3,k} Mom^{k}\, x\, Post_{i,s}] + \beta_4 X_{i,s,t} + \epsilon_{i,s,t})$$

Appendix Table B4 shows results of this regression for *J* equals 1, 2, or 3. In column 1 *J* equals 1 which repeats the main DD estimate from Table 2 column 2. In columns 2 and 3, *J* equals 2 and 3, and the DD estimate of $\beta_{3,k=2}$ for having at least two kids after 1975 is positive (about 1.2-percentage points) and significant. This means that women with at least two kids were more likely to respond to the employment incentives of the EITC than women with exactly one child.[12] For mothers with at least three children, I do find the expected response of zero, and in columns 2 and 3 the main DD estimate of $\beta_{3,k=1}$ falls from 3.3- to 2.6-percentage points.

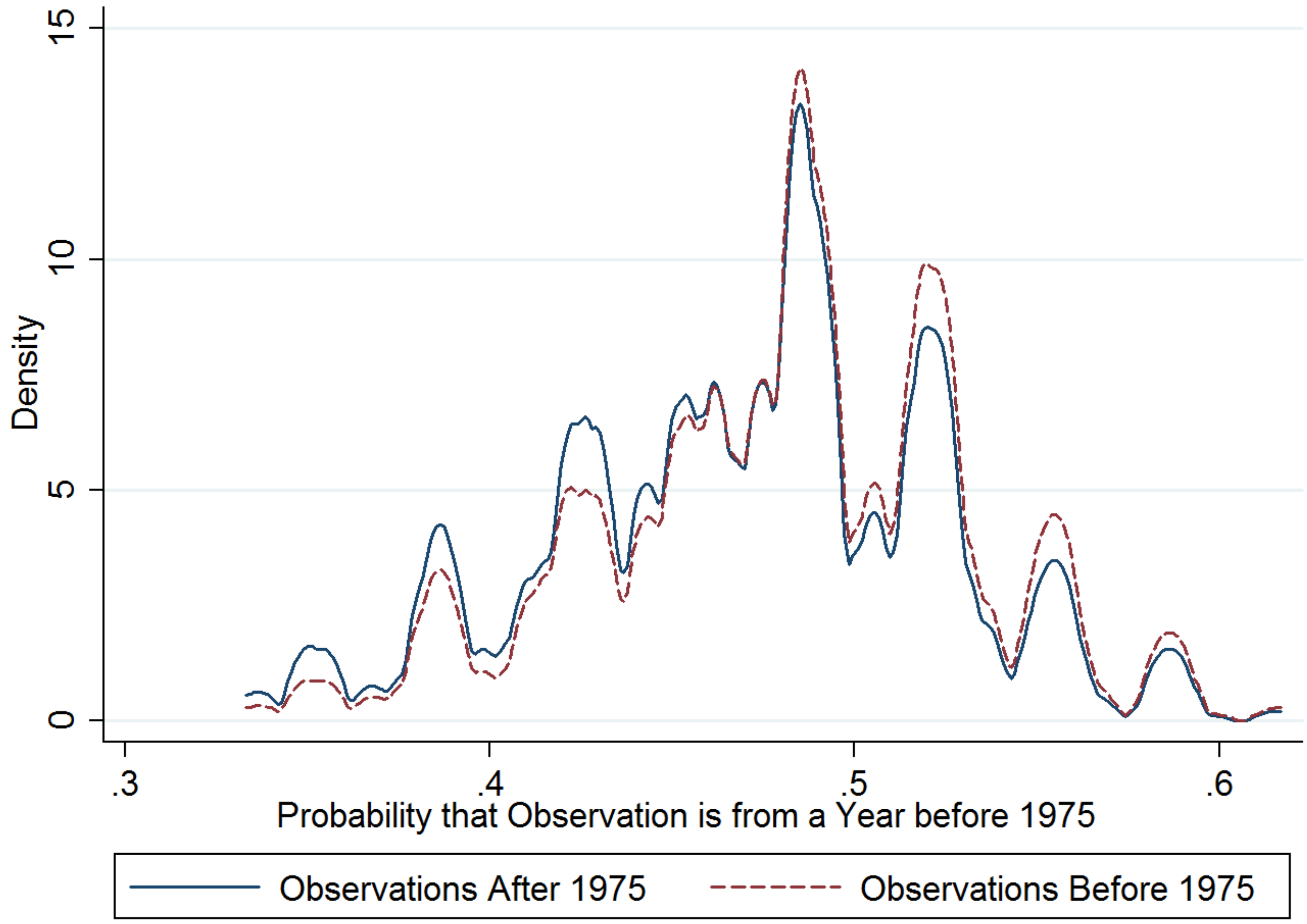
This puzzling finding could indicate the presence of another policy or trend encouraging mothers to work. Interestingly, Eissa and Liebman (1996) also find an additional employment response from women with at least two children after the 1986 EITC expansion, and suggest that this might be due to an increase in tax exemptions for each dependent in the 1986 Tax Reform Act. I consider this explanation in the 1975 EITC context. In 1979 the tax exemption for each child increased from \$750 to \$1,000, however when I restrict to sample to end in 1978 I still find a positive estimate of $\beta_{3,k=2}$ (see Table 6 column 4).

Another potential explanation is that mothers with multiple children are more likely to have completed their fertility. At this time most mothers would not consider working if they already had an infant to take care of and were preparing to have an additional child soon. Rather, mothers that had finished childbearing would be more receptive to the employment incentives of the EITC. This hypothesis can be tested by looking at the age of each mother's youngest child. In Appendix Table B4 columns 5 to 9, I restrict the sample of mothers in the treatment group to

[12] Eissa and Liebman (1996) also find this additional response from women with at least two children. They suggest that this may be due to the new 1986 tax exemptions given for each dependent, which benefited families with multiple children more and therefore may have given an extra incentive to such women to enter employment. The tax exemption for each child was also increased in 1979 from \$750 to \$1,000. However, even when I exclude 1979 and 1980 I still find a positive estimate of $\beta_{3,k=2}$.

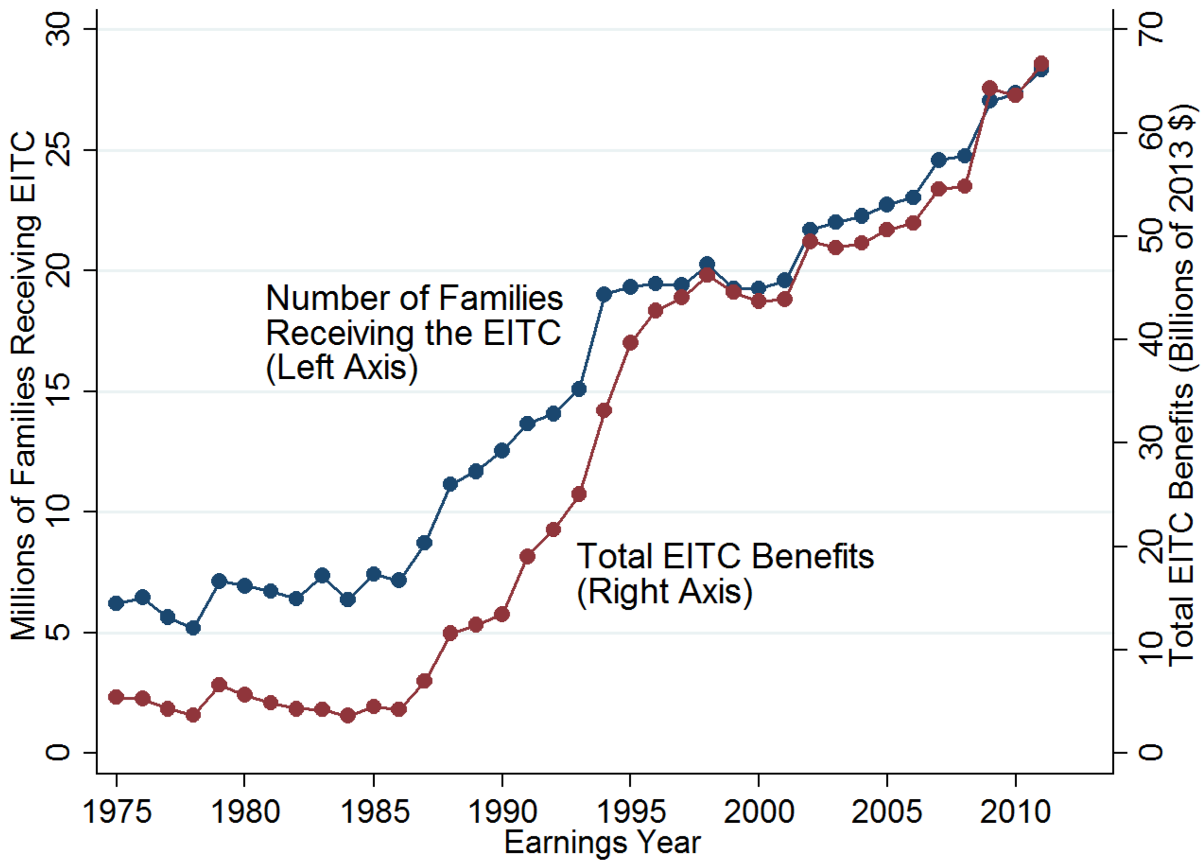
those with a youngest child at least 2, 3, 4, 5, and 6 years old. As this age restriction increases, the DD estimate on having at least two children after 1975 $\beta_{3,k=2}$ converges towards zero (and the baseline DD estimate $\beta_{3,k=1}$ on having at least one child remains positive between 2.5- and 3.5-percentage points). Mothers with a youngest child at least 3 years old are statistically no more likely to respond to the EITC than women with just one child. I also verify that this pattern holds for the 1986 EITC expansion (results not shown), and conclude that the apparent additional employment increase for women with at least two children can be explained by completed fertility.

Appendix Figure B1: Kernel Density Plot for Observations Before and After 1975



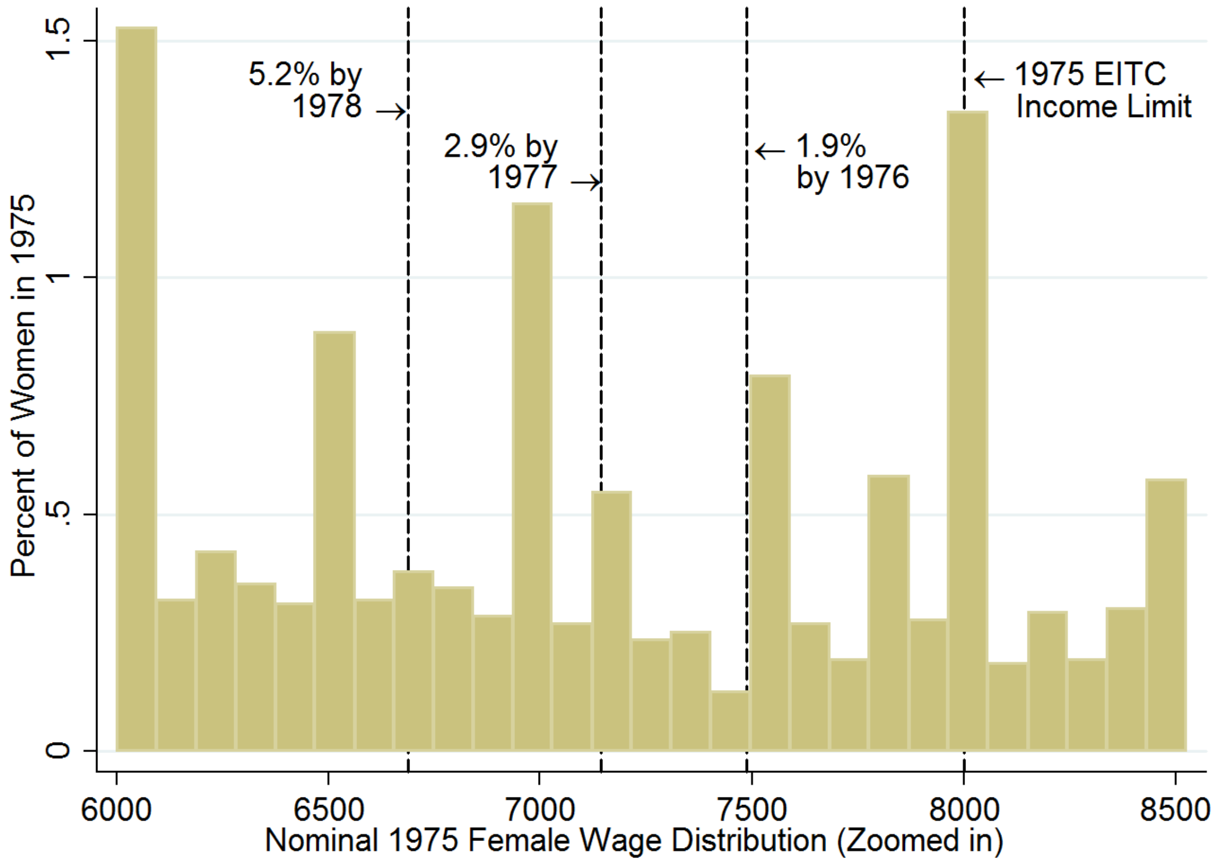
Notes: 1972-1981 March CPS data. Equation (5) used to predict the probability of each observation being observed in a year before 1975. Following DiNardo, Fortin, and Lemieux (1996), I choose a parsimonious set of observable characteristics X, which includes six age bins, three education bins, married and nonwhite dummy variables, and 21 state bins for a total of 1512 cells.

Appendix Figure B2: Trends in EITC Benefits and Recipients



Notes: Author's calculations from IRS data.

Appendix Figure B3: Previously Working Women Bracket-Creep Out of EITC After 1975



Notes: Data source 1974 March CPS. High-impact sample of women. T'his figure shows that many women working in 1974 would have been eligible for EITC benefits just after 1975. However, due to high inflation in the 1970s (up to 13 percent per year) many of these women would "bracket-creep" out of the EITC income range due to rising nominal wages (see Saez 2003 for further "bracket-creep" discussion). This can reconcile the apparent contradiction between my findings that the EITC induced a large number of mothers into the labor force and Appendix Figure A18 that shows a roughly constant number of EITC recipients between 1975 and 1980. Inflating annual earnings in the 1974 income distribution by the Consumer Price Index shows that by 1978 fully 5.2 percent of EITC-eligible women would no longer be eligible for the EITC since the nominal income limit was $8,000. The rough estimates of the number of women that would bracket-creep out of EITC eligibility approximates the annual number of mothers entering the labor force due to the EITC. I conclude that even though the "stock" of EITC-eligible women may have remained roughly constant in the years after 1975, there was substantial flow in and out of EITC eligibility.

Appendix Table B1: Triple Differences Using Various Comparison Groups

Comparison Group:	Married Women with Spouses Earning Above EITC Kink Point				All Men	Single Men
Upper Bound on Spousal Earnings (2013 $):	None	$60,000	$50,000	$40,000		
Variables	(1)	(2)	(3)	(4)	(5)	(6)
Parent x Post1975 x EITC Eligible	0.024**	0.029***	0.028***	0.031***		
	(0.011)	(0.011)	(0.010)	(0.010)		
Parent x Post1975 x Woman					0.034***	0.039***
					(0.006)	(0.013)
Observations	298,885	230,876	205,411	178,883	482,068	303,035

Note: *** p<0.01, ** p<0.05, * p<0.1. Data source: 1972-1981 March CPS. Binary dependent variable employment with mean 0.637 for the whole sample and mean 0.530 for the treatment group in 1975. CPS weights used and average marginal effects from probit regression are shown. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. Each column represents a separate regression with the full set of controls from Table 2 column 5. Columns 1 to 4 also include the following sub-components of the DDD as controls: Parent x Post1975, Parent x EITC Eligible, and Post1975 x EITC Eligible. Likewise columns 5 and 6 include Parent x Post1975, Parent x Woman, and Post1975 x Woman. These additional controls ensures that the DDD model is saturated and the DDD estimate can be appropriately interpreted. The sample in columns 1 to 4 is the "high-impact" sample and married women with a spouse earning above the EITC kink point (these are the two groups of 16- to 45-year-old women in Table 3 columns 1 and 2). The samples in columns 2 to 4 are also restricted to spouses with earnings in the specified range. The sample in columns 5 and 6 includes all 16- to 45-year-old men and women, except column 6 only keep single men.

Appendix Table B2: Reweighting to Account for Changing Group Composition

			Reweighting Mothers After 1975 to Look like Mothers Before 1975	
	Baseline Probit	Baseline Logit	DFL Weights	IP Weights
Variables	(1)	(2)	(3)	(4)
Mom	-0.056	-0.032	-0.037	-0.035
	(0.101)	(0.109)	(0.151)	(0.104)
Post1975	0.001	-0.001	-0.025*	0.001
	(0.014)	(0.014)	(0.015)	(0.013)
Mom*Post1975	0.033***	0.033***	0.029**	0.032***
	(0.009)	(0.009)	(0.012)	(0.009)
Observations	335,639	335,639	335,639	335,639

Note: *** p<0.01, ** p<0.05, * p<0.1. Data source: 1972-1981 March CPS. Binary dependent variable labor force participation. CPS weights used. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. Full set of controls used from Table 2 column 5. Sample includes all women 16-45.

Appendix Table B3: Alternate Ways to Treat Imputed CPS Observations

	Baseline (using default March CPS imputations)	Dropping Imputed Observations	Using DFL Weights	Using IPW	Assigning 0 to all Imputed Obs	Assigning 1 to all Imputed Obs
Variables	(1)	(2)	(3)	(4)	(5)	(6)
Mom	0.002	0.008	-0.054	0.086	-0.049	-0.001
	(0.041)	(0.042)	(0.033)	(0.056)	(0.045)	(0.041)
Post1975	-0.021**	-0.022**	-0.011	-0.036***	-0.039***	-0.021**
	(0.010)	(0.011)	(0.010)	(0.013)	(0.011)	(0.010)
Mom*Post1975	0.039***	0.040***	0.034***	0.045***	0.045***	0.039***
	(0.007)	(0.007)	(0.006)	(0.009)	(0.007)	(0.007)
Observations	132,272	128,577	128,577	128,577	132,272	132,272

Note: *** p<0.01, ** p<0.05, * p<0.1. Data source: 1972-1981 March CPS. Binary dependent variable labor force participation. CPS weights used. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. Full set of controls used from Table 2 column 5. "High-impact" sample used, though results are similar for the sample of all 16 to 45 year old women.

Appendix Table B4: No Additional Maternal Employment Response for Second Child, Conditional on Age of Youngest Child

Specification:	Baseline DD for 1+ Children	Add DD for 2+ Children	Add DD for 3+ Children	End Sample in 1978 to Rule Out Exemption Explanation	DD for 2+ Children Approaches Zero for Mothers Likely to Have Completed Fertility				More
Restricting Mothers by Age of Youngest Child	No	No	No	No	> 1	> 2	> 3	> 4	> 5
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Mom	-0.056	-0.063	-0.069	-0.044	-0.187*	-0.268***	-0.334***	-0.406***	-0.354***
	(0.101)	(0.100)	(0.099)	(0.115)	(0.096)	(0.098)	(0.101)	(0.105)	(0.104)
Post1975	0.001	0.000	0.000	0.005	-0.008	-0.008	-0.008	-0.010	-0.009
	(0.014)	(0.013)	(0.014)	(0.011)	(0.013)	(0.012)	(0.011)	(0.012)	(0.012)
Mom x Post1975	0.033***	0.026***	0.026***	0.015*	0.025***	0.027***	0.029***	0.034***	0.035***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.010)
Mom of 2+ Kids		-0.102***	-0.104***	-0.107***	-0.077***	-0.071***	-0.062***	-0.053***	-0.041***
		(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.006)	(0.006)	(0.006)
Mom of 2+ Kids x Post1975		0.012**	0.014**	0.017***	0.013*	0.011	0.010	0.007	-0.000
		(0.005)	(0.006)	(0.005)	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
Mom of 3+ Kids			-0.015***						
			(0.005)						
Mom of 3+ Kids x Post1975			-0.002						
			(0.005)						
Observations	335,639	335,639	335,639	251,273	294,093	276,845	262,114	249,222	237,503

Note: *** p<0.01, ** p<0.05, * p<0.1. Data source: 1972-1981 March CPS. Binary dependent variable employment. CPS weights used and average marginal effects from probit regression are shown. Standard errors are computed by the delta method and are robust to heteroskedasticity and clustered at the state-year level. Each column represents a separate regression with the full set of controls from Table 2 column 5. Each sample includes all women 16 to 45 years old. Column 2 is identical to column 1 except it adds an indicator for having 2+ children and an interaction between having 2+ children and the year being after 1975. Column 3 includes these two variables as well as two more for having 3+ children. Columns 4 to 8 drop mothers in the treatment group with a youngest child below the specified age.

Appendix C: Comparative Statics from a Simple Labor-Leisure Model

Consider a simply labor-leisure model where a woman's decision to work is a function of her consumption c_i, leisure L_i, and cost of working $g_i(.)$.

(C1) $$V(c_i(.), L_i, g_i(.)) = \max[f(c_i(.), L_i, g_i(.))]$$

$c_i(.)$ is determined by her wage w_i and non-labor income n_i (e.g. spousal wage, welfare benefits), while $g_i(.)$ is determined by childcare and opportunity costs, as well as psychological costs that come from a social stigma against women working (Fortin 2005; Charles, Guryan, and Pan 2009) or from her own beliefs about gender roles and identity (Akerlof and Kranton 2000). A woman will work if the additional consumption benefits of working outweigh the costs of working, where working can be thought of as either a binary or continuous decision.

To fix ideas, consider a less general form of equation (1):

(C2)

$$V(c_i(.), L_i, g_i(.)) = \max_{l,L\in[0,1]} \{\alpha \log[l * w_i(k_i, COL_i) + n_i] + \beta \log[L] - \gamma_{s,t}[l * (k_i + S_s)]\}$$

Where α and β weight the relative importance of consumption and leisure, and $\gamma_{i,s,t}$ weights the cost of working, which can vary by state and year, and increases with work hours l and number of kids k_i. Wages $w_i(k_i)$ are a function of whether she has children and is eligible for the EITC. Normalizing labor and leisure such that $l + L = 1$ results in the following first order condition:

(C3)

$$\frac{\partial}{\partial l} = \frac{\alpha w(.)}{lw(.) + n} - \frac{\beta}{1-l} - \gamma_{i,s,t}k_i = 0$$

Using the implicit function theorem to derive $\frac{\partial l}{\partial w}$ yields:

(C4)

$$\frac{\partial l}{\partial w} = \frac{(\alpha - l)(1-l)^2}{w(.)(1-l)^2 + (lw(.) + n)^2}$$

which is positive if $\alpha > l$.

Using the implicit function theorem to derive $\frac{\partial l}{\partial n}$ yields:

(C5)

$$\frac{\partial l}{\partial n} = \frac{-\alpha w(.)(1-l)^2}{w(.)(1-l)^2 + (lw(.) + n)^2}$$

which is always negative.