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

By Jacob Bastian*

The rise of working mothers radically changed the U.S. economy and the role of women in society. In one of the first studies of the 1975 introduction of the Earned Income Tax Credit, I find that this program increased maternal employment by 6 percent, representing one million mothers and an elasticity of 0.49. The EITC may help explain why the U.S. has long had such a high fraction of working mothers despite few childcare subsidies or parental-leave policies. I also find evidence that this influx of working mothers affected social attitudes and led to higher approval of working women. (JEL: H24, I38, J16, J38)

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A surprising difference between the U.S. and other developed countries is the large number of mothers in paid work, especially new mothers. By 2000, 56 percent of mothers with infants worked in the U.S. compared to 25 to 45 percent in other developed countries (OECD 2007).[1] The U.S. was not always an outlier in this regard: the number of working mothers in recent decades is also high by U.S. historical standards (Goldin 1990, Costa 2000)[2] and is puzzling since few child-care subsidies or family-friendly work policies (e.g. paid parental leave) exist in the U.S. (Ruhm 1998). This paper finds that the 1975 introduction of the Earned Income Tax Credit (EITC) may help explain this puzzle. Not only do I find that the EITC played an important role in the rise of working mothers, but also that this program led to more positive social attitudes towards working women.

Time-series evidence shows that the relative employment of mothers – compared to women without children – rapidly increased after 1975 (Figures 1.A and 2.A). Between 1975 and 1980, the relative employment of mothers rose by about 5 percentage points, closing the employment gap between these two groups by 20 percent. Using March Current Population Survey data and a dynamic difference-in-differences (DD) approach, I show that much of the 1975-to-1980 increase in the *relative* employment of mothers can be attributed to the 1975 EITC. Interestingly, the unadjusted trend in maternal employment is quite similar to the regression-adjusted trend that controls for a rich set of individual- and state-by-year-level covariates

[1] Cross-country comparisons of working mothers are not straightforward: many countries count mothers on paid parental leave as employed (OECD 2007). The 2003 employment rates of mothers with kids under 3 in Austria, Finland, and Sweden was 80.1, 52.1, and 72.9 percent, but excluding mothers on paid parental leave yields lower rates of 40.1, 33.8, and 45.1 percent (OECD 2007, p.57).

[2] Only 20 percent of married women with infants worked in 1973, compared to 62 percent in 2000 (Goldin 2006).

(Figures 1.B and 2.B). The EITC also increased labor-force attachment and work intensity, raising average annual work hours by 5.7 percent (35 hours) and earnings by 7.3 percent ($750 in 2013 dollars). Results imply a participation elasticity of 0.41 to 0.49, in line with other estimates of this period (Blau and Kahn 2005, Heim 2007, Chetty et al. 2012).

Consistent with the 1975 EITC causing this rise in employment, I find larger responses from mothers more likely to be EITC-eligible and null responses from placebo groups of women and mothers not eligible for EITC benefits. Responses varied by marital status, spousal earnings, and education in a manner consistent with a simple labor-supply model. I use the placebo group of EITC-ineligible mothers in a triple differences (DDD) specification to net out contemporaneous policies and trends (e.g. birth control, divorce laws, abortion) affecting all mothers: the DDD estimate corroborates the DD result (2.6 and 3.3 percentage points).

My estimates suggest that the 1975 EITC encouraged about one million mothers to begin working. Yet, this is unlikely to capture the full impact of the EITC on society. In section VI, I use General Social Survey data to examine whether this influx of working mothers affected social attitudes towards working women ("gender-equality preferences"). This hypothesis is motivated by recent evidence that such attitudes are malleable and increase with exposure to working women: Fernández, Fogli and Olivetti (2004) and Olivetti, Patacchini and Zenou (2016) find that having a working mother – and having friends with working mothers – leads to stronger gender-equality preferences in adulthood. Additionally, Finseraas et al. (2016) shows that exposure to female colleagues reduces discriminatory attitudes. With these results in mind, the attitudes of millions of Americans may have been affected when a million mothers began working after 1975.[3]

To estimate the impact of the EITC on gender-equality preferences, I use a two-sample two-step process, in which I characterize and exploit geographic heterogeneity in the EITC response and test whether states with larger EITC responses experienced larger attitude changes after 1975. Using both the *actual* state EITC response and the *predicted* response (based on preexisting state demographic traits, to help alleviate concerns about the potential endogeneity of gender-equality preferences and EITC response), I find that states with larger EITC responses had larger increases in preferences for gender equality after 1975. Preference changes occurred among both men and women, within and across regions, and do not appear to be driven by preexisting attitudes, demographics, or general trends in social norms. Subgroup analysis confirms larger preference changes among people more likely to know these newly working women: lower-educated adults. I also use a placebo outcome on racial-equality preferences to test and rule out the possibility that states with higher EITC responses were simply experiencing changes in various types of social attitudes. Regarding external validity and whether working women can affect social attitudes towards women in other contexts, I also find evidence of attitude changes due to the large increase in working women during World War II.

In one of the first studies of the 1975 EITC,[4] I find that the EITC encouraged a million mothers to begin working and affected the social attitudes of millions of Americans.

[3]Google ngrams (Michel et al. 2011) provide descriptive evidence that the rise of working mothers was salient and that references to working mothers became much more common after the mid-1970s (Figure 3).

[4]Subsequent EITC expansions – and their effect on maternal employment – have been studied (see section I).

I. EITC History and Known Effects of the EITC

The EITC came to exist partly as a response to the 1960s War on Poverty, which succeeded in improving health (Almond, Hoynes and Schanzenbach 2011, Hoynes, Page and Stevens 2011, Goodman-Bacon 2013, Bailey and Goodman-Bacon 2015) and decreasing poverty, but also had unintentional work disincentives (Moffitt 1992, Hoynes 1996, Hoynes and Schanzenbach 2012).[5] Welfare dependency came to be seen as a growing social problem and momentum built for a guaranteed annual income with support from economists Milton Friedman (Friedman 1962) and James Tobin (Tobin 1969). The U.S. House of Representatives passed such a plan – the Family Assistance Plan – in 1970 with the backing of President Nixon that would have replaced welfare.[6] However, the U.S. Senate never passed the plan because of disagreement about how generous the program should be and concerns about potential work disincentives. An alternative program called the Work Bonus Plan – with work requirements – was introduced by Louisiana Senator Russell Long in 1972. A version of this bill was eventually passed as the Earned Income Tax Credit (EITC) and signed into law by President Ford on March 29, 1975. See Liebman (1998) and Ventry (2000) for a detailed history of the EITC program and legislation.

The 1975 EITC was a refundable tax credit that provided a 10 percent earnings subsidy to working parents with annual household earnings up to $18,000 in 2013 dollars ($4,000 nominal dollars).[7] The EITC was also

[5]See Bailey and Danziger (2013) for a detailed analysis of War on Poverty programs.

[6]FAP would have guaranteed $3,100 (2013 dollars) for each parent and $1,800 for each child – $9,800 for a family of four (the 1970 poverty line was about $23,000 for a family of four). Benefits would phase out at 50 percent when household earned income surpassed $4,400 (Trattner 2007, p.315). See New York Times April 17, 1970. Rhys-Williams (1943) was among the first to outline this type of program.

[7]To be EITC-eligible, tax filers had to have at least one child living in their home for

available to parents with earnings above $18,000, but benefits decreased at a rate of 10 percent and reached zero for earnings above $36,000 (Figures 4.A and 4.B).[8] At this time, there were no additional EITC benefits for having more than one child and benefits did not vary by state or marital status.

Since 1975, the EITC has been expanded many times (see Figure 4.B for details) and has grown into one of the largest anti-poverty program in the U.S., redistributing $66 billion to 28 million individuals and lifting 6.5 million people – including 3.3 million children – out of poverty in 2013 (Center on Budget and Policy Priorities 2014). The EITC has raised maternal employment (Dickert, Houser and Scholz 1995, Eissa and Liebman 1996, Meyer and Rosenbaum 2001, Hotz and Scholz 2006, Eissa, Kleven and Kreiner 2008), increased earnings (Dahl, DeLeire and Schwabish 2009), improved health (Evans and Garthwaite 2014), decreased poverty (Scholz 1994, Neumark and Wascher 2001, Meyer 2010, Hoynes and Patel 2015, Bitler, Hoynes and Kuka 2016), and helped children of EITC recipients by improving health (Hoynes, Miller and Simon 2015, Averett and Wang 2015), test scores (Chetty, Friedman and Rockoff 2011, Dahl and Lochner 2012), and longer-run outcomes like educational attainment (Manoli and Turner 2014, Bastian and Michelmore 2018) and employment and earnings (Bastian and Michelmore 2018). The EITC's unintended consequences include lower pre-

more than half the year ("residency test"). This child must be under 19, under 24 if a full-time student, or any age if disabled. Before 1987, tax filers did not have to provide Social Security numbers for dependents. Until 1990, tax filers had to demonstrate they provided at least half the costs of maintaining the household ("support test"): cash and in-kind public assistance had to be less than half of the household budget (Holtzblatt 1991, Holtzblatt, McCubbin and Gillette 1994). Married couples had to file taxes jointly. Since I do not observe tax filing, I assume all unmarried women file taxes as household head, married couples file joint taxes, and family members under 19 (or 24 if a student) are dependent children. I treat subfamilies within a household as separate tax-filers.

[8] Figure 4.A shows a budget constraint under the EITC and Figure 4.B illustrates the "phase-in" and "phase-out" portion of the EITC schedule while contrasting the 1975 EITC schedule with the 2013 EITC. Benefits phase out with adjusted gross income.

tax wages of low-skill workers (Leigh 2010, Rothstein 2010) and possible effects on fertility and marriage.[9] See Nichols and Rothstein (2015) and Hoynes and Rothstein (2016) for recent EITC literature reviews.

Although much is known about the EITC, almost nothing is known about the 1975 introduction or how the EITC may affect attitudes towards working women. I show that the 1975 EITC encouraged one million mothers to begin working, which subsequently increased approval of working women.

Almost all studies of the EITC ignore the program's first decade.[10] Although there was little policy variation before 1986, the 1975 introduction was itself a large policy change that has received surprisingly little attention, in part due to the common misconception that the original EITC was too small to have had much of an effect.[11] However, the 1975 EITC was large in at least three ways (Figure A.1): first, almost half of all households had earning below the EITC income limit; second, benefits were quite high, up to $1,800 (2013 dollars); third, a 10 percent earnings subsidy represented a substantial year-over-year increase in potential earnings. Other reasons to expect the 1975 EITC to have had a large impact is that female labor supply was more elastic during this period than in later decades (Blau and Kahn 2005, Heim 2007) and the fraction of mothers on the margin of working declined with subsequent program expansions (Björklund and

[9]Effects on these margins are generally small: For fertility, Baughman and Dickert-Conlin (2009) and Bastian (2018) find positive effects. For marriage, Ellwood (2000), Dickert-Conlin and Houser (2002), Herbst (2011), and Michelmore (2015) find negative effects, while Bastian (2018) finds positive effects.

[10]Bastian and Michelmore (2018) is one exception.

[11]As seen in the following representative quotes: "Between 1975 and [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 [EITC] 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).

Moffitt 1987, Heckman and Vytlacil 1999).

II. Conceptual Framework

The EITC was a wage subsidy for low-income parents and should have increased the employment of mothers.[12] Intuition for this can be formalized in the following framework (where work could be binary or continuous).

$$(1) \qquad U\big(c(.), L, g_{st}(.)\big) = \big[c(l_i, w_i, n_i, h_i, k_i) + L_i^{\alpha} - g_{st}(l_i, k_i)\big]$$

Women, states, and years are denoted by i, s, and t. Women divide one unit of time between labor l_i, leisure L_i, and home production h_i. Consumption $c(.)$ is a function of her labor supply l_i, wage w_i, non-labor income n_i, home production good h_i, and kids k_i. Accounting for the EITC requires an interaction between w_i and k_i since only working parents were eligible for the EITC. The cost of working $g_{st}(l_i, k_i)$ is a function of labor supply l_i and kids k_i. The EITC increased w_i for EITC-eligible mothers, making work a relatively more attractive use of time.

To estimate the EITC's effect on maternal employment, I use difference in differences (DD) and compare the employment rates of women with and without kids (first difference), before and after 1975 (second difference). I approximate equation (1) with the following non-linear model that estimates the probability that each woman works.

$$(2) \quad P(E_{ist}) = f(\beta_1 Mom_{ist} + \beta_2 Mom \times Post1975_{ist} + \beta_3 X_{ist} + \delta_{st} + \epsilon_{ist})$$

[12] I assume working mothers did not displace non-mothers (Neumark and Wascher 2011). However, even if an increase in working mothers led to declines in earnings (Leigh 2010, Rothstein 2010), this apparently did not lead to a general-equilibrium effect where the employment of non-mothers decreased (see Figures 1.A and 2.A).

E_{ist} is binary for whether a woman is employed.[13] Mom and $Post1975$ denote whether a woman is a mother and if the year is after 1975; $Mom \times Post1975$ is the DD variable of interest. The EITC treatment effect β_2 should be positive since the EITC subsidized work. X_{ist} are controls that vary at the individual, state, and year level. δ_{st} contain state and year fixed effects to control for national trends and state-specific traits associated with female employment. ϵ_{ist} is an error term. Coefficients are measured in percentage points. Average marginal effects from a logit model are reported throughout (unless otherwise stated). Standard errors are robust to heteroskedasticity and clustered at the state level.

A. Data and Descriptive Statistics

I estimate equation (2) using 1971 to 1986 March CPS data (Ruggles et al. 2015) and the sample of all 18- to 50-year-old women. The treatment group consists of mothers[14] and the control group consists of women without children. Table 1 shows summary statistics for all 571,170 women in column 1, while columns 2 and 3 split the sample into treatment and control groups, and columns 4 and 5 split the sample by marital status. Women in the sample average 32 years old with 12.3 years of education, 12 and 6 percent are Black and Hispanic, 66 percent work, average individual annual earnings are $14,158 ($21,418 conditional on working), average household earnings are $45,822 (2013 dollars), and 41 percent have household earnings below the EITC limit. Mothers are older, less likely to be white, less likely to work,

[13] I focus on employment since this is where most EITC benefits are and since the participation margin generally manifests greater responsiveness to wage variation than hours of work (Heckman 1993).

[14] To match the definition of EITC-eligible children, I define mothers as having at least one child 18 or under, or having a child between 19 and 23 that is in school full time.

and have less education and higher household earnings. Married women are older, have more children, are less likely to work, and have higher household earnings. See Appendix E for data and sample details.

Figures 1.A and 2.A show unadjusted 1970-to-1985 employment trends for women with and without kids and preview the regression-adjusted results. From 1970 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 increased and the gap narrowed to 18 percentage points, where it remained from 1979 to 1985 (Figure 1.B). Although employment levels differed for these groups, employment trends were parallel before 1975 (p-values 0.42 and 0.38 for Figures 1.B and 2.B).

B. **Ruling Out Contemporaneous Shocks to Employment**

In addition to parallel trends, a causal interpretation of DD requires that no contemporaneous factor affected the relative employment of mothers. Even though the 1970s was a period of inflation, oil and food price shocks, and two recessions, in the following discussion I find little evidence of confounding policies or trends that affected maternal employment.

The first oil shock began in 1973 when the Organization of Arab Petroleum Exporting Countries proclaimed an oil embargo against the West in response to the Yom Kippur War. This led to a quadrupling of oil prices by March 1974, double-digit inflation and food-price increases, and a recession from November 1973 to March 1975. A few years later, the second oil shock began when global oil production decreased due to the Iranian Revolution. This preceded the double-dip recession that occurred between 1980 and 1982. Although a recession ended around the time the EITC began, it is not obvious why this would have affected the relative employment of mothers since no such increase occurred after the 1980-1982 recessions (Figures 1.A and 2.A).[15] To account for these factors, I control for annual inflation, state-by-year employment and manufacturing employment, and allow these variables to vary by family size, marital status, and education.

Two potential identification threats include public-program cuts, which could increase maternal employment via an income effect, or a sudden change in demographic traits associated with employment and unrelated to the EITC. However, public assistance expanded in the 1970s (a period of "welfare explosion" (Moffitt 2003)): AFDC, Food Stamps, WIC, and payroll taxes all increased or were flat (Figure A.3).[16] Also, trends in marriage, fertility, education, and male earnings were smooth (Figure A.2).[17] I control for the impact of welfare and demographics on employment, and allow them to vary by state, year, and race.

Perhaps the most serious potential confounder is the 1976 Child and Dependent Care Tax Credit (CDCTC), a non-refundable tax credit for child care expenses. I investigate whether this policy affects my analysis in three ways: First, I look at the fraction of EITC recipients that received CDCTC benefits (using IRS Statistics of Income [SOI] data):[18] only 1 percent of EITC-eligible tax filers received any CDCTC benefits, compared to 30 per-

[15]Theoretically, a permanent price increase could increase labor supply through an income effect, but the 1970s price shocks were temporary and should not have differentially affected mothers.

[16]AFDC denotes Aid to Families with Dependent Children, a cash assistance welfare program. WIC denotes Women, Infants, and Children, an in-kind food assistance program. See Figure A.3 notes for brief histories of these public programs.

[17]I cannot rule out a threshold-crossing model (Schelling 1971) where a continuously changing covariate has a discrete impact on an outcome.

[18]SOI data are de-identified samples of U.S. Federal Individual Income Tax returns with detailed income information, but little demographic information. SOI sampling weights used. More details in Appendix B.

cent of EITC-ineligible tax filers with children (Figure A.4), corroborating previous evidence that most CDCTC benefits go to upper-middle-class families (Maag, Rennane and Steuerle 2011). Second, restricting the sample to women *ineligible* for the EITC and *eligible* for the CDCTC, I do not detect an increase in working mothers after 1975 (Table 3 column 4). Third, I examine the subsequent 1981 CDCTC expansion (rate increased from 20 to 30 percent) and find that although CDCTC benefits doubled after 1981 (Figure A.4), this pattern bears little resemblance to the maternal employment trends in Figures 1.A and 1.B.[19] Together, this evidence suggests that the CDCTC had a minimal effect on the population affected by the EITC.

In conclusion, I find little evidence of confounding policies or trends that affected the relative employment of mothers.[20] If anything, the expansion of

[19]Figure 2.A suggests that the 1981 CDCTC may have increased employment for mothers with relatively high spousal earnings (the group in Table 3 column 4). Triple-differences analysis in Table 4 nets out any employment effect on this group.

[20]Averett, Peters and Waldman (1997) finds that the CDCTC increased the labor supply of mothers in their twenties with young children in 1987. Other potential confounders include Head Start, the 1972 Equal Employment Opportunity Act mandating equal pay for equal work for women, legalized abortion in 1973, the 1974 Equal Credit Opportunity Act allowing women to take out loans without a male co-signer, the 1978 Pregnancy Discrimination Act requiring employers to treat pregnancy as a temporary disability, and changes in birth-control and divorce laws during the 1960s and 1970s. However, Head Start began in the 1960s; the EEOA applied to most states outside the South before 1972 ((Altonji and Blank 1999, footnote 54); four states legalized abortion in 1970 (AK, HI, NY, CA) and had maternal-employment trends similar to other states (results omitted); the ECOA likely did not affect maternal employment (Smith 1977, Elliehausen and Durkin 1989); the PDA had little effect on maternal labor supply since mothers bore the whole cost of the mandated benefits and the return to work remained the same (Gruber 1994) (although Mukhopadhyay (2012) finds a positive labor-supply effect of the PDA on pregnant women and mothers of young children, however, the PDA did not become law until October 1978 and Figures 1.A and 1.B show that most of the rise in maternal employment had already occurred by then); the birth-control pill first became available in 1960 and was available in most states before the mid-1970s (Goldin and Katz 2000, Goldin and Katz 2002, Bailey 2006); divorce began rising in the 1960s (Johnson and Skinner 1986, Peters 1986, Parkman 1992, Wolfers 2006) and California, the first state to pass no-fault divorce in 1970, had similar maternal employment trends as the other states (results omitted). Choo (2015) finds that no-fault divorce laws decreased the *growth rate* of divorce.

public assistance during the 1970s would have led to slight *decreases* in maternal employment, implying that results in this paper may underestimate the employment effects of the 1975 EITC.

III. The EITC and Extensive-Margin Labor Supply

A. Average Treatment Effects

I estimate the average effect of the EITC on maternal employment using equation (2), March CPS weights, and adding controls cumulatively across columns in Table 2. Column 1 controls for whether each observation is a mother (Mom),[21] whether the observation occurs after 1975 ($Post1975$), and the DD variable of interest ($Mom \times Post1975$). Column 2 adds state and year fixed effects to account for idiosyncratic state traits and annual shocks affecting all women.[22] Column 3 adds demographic controls to account for demographic-led increases in maternal employment and help account for the fact that mothers are on average older, have less education, and more likely to be married and nonwhite (Table 1). Column 4 adds state-by-year unemployment rates (that can vary by marital status and whether women have kids) to control for the effects of economic conditions on employment. Columns 5 and 6 show that results in column 4 are robust to using probit or OLS. Finally, column 7 adds a "kitchen-sink" set of controls that interacts each control (along with annual inflation and state-by-year manufacturing

[21] Restricting (Mom) to those with a child born before 1975 avoids potential fertility responses to the EITC, but affects the composition of mothers over time. This approach yields a similar DD: 0.030 (0.006).

[22] Before 1977, CPS did not uniquely identify all states. I merge states into the 21 smallest possible geographical units to provide a balanced panel (details in Appendix E). So few clusters may bias the standard errors (Angrist and Pischke 2009, Cameron, Gelbach and Miller 2008, Cameron, Gelbach and Miller 2011). Block bootstrap yields similar standard errors and clustering at the year-by-(mother/non-mother) level also yields statistically significant estimates, with slightly larger standard errors of 0.0158.

employment) with year, state, marital status, having kids, and race. These interactions flexibly account for the impact of economic conditions, changing demographics, and general trends in the employment of women.

Across each set of controls the DD estimate is stable between 3.3 and 4.9 percentage points (or 6.2 and 9.2 percent from a baseline of 53 percent)[23] and significant at the 99-percent level. Results imply that about one million mothers began working because of the 1975 EITC.[24] The EITC is responsible for about a quarter of the 12-percentage-point rise in absolute maternal employment and a fifth of the 10-percentage-point rise in overall female employment between 1975 and 1985. I use the more conservative logit model and set of controls in column 4 throughout the rest of the analysis (unless otherwise specified). Results are robust to alternate binary definitions of working based on earnings, weeks worked, or labor-force participation (Table A.1), using alternate age cutoffs (Table A.2), not using CPS weights (estimate is 0.030 [0.007]), and additional robustness checks (Appendix B).[25]

B. Heterogeneous and Subgroup Treatment Effects

Although the average employment effect of the EITC was positive, this effect should have varied by the likelihood of receiving EITC benefits. In Table 3, I test whether the treatment effect varied in a way consistent with

[23] 53 percent baseline seen in Figure 2.A. Results are intent-to-treat effects: about 20 percent of households are EITC-eligible and do not claim the EITC or are EITC-ineligible families and do (Scholz 1994). Liebman (1997) and Liebman (2000) find that 89 and 95 percent of women allocated to the treatment and control groups filed taxes appropriately in the 1980s. Random misallocation implies that the estimates should be scaled up by 19 percent (Eissa and Liebman 1996).

[24] 60 percent of the 53.8 million women 18-50 in 1980 are mothers (March CPS). 3.3 percentage points of these mothers corresponds to about 1 million mothers.

[25] Appendix B shows results are robust to model choice, sample period, reweighting to account for group composition and CPS data imputations; I also explain how flat EITC beneficiaries and increases in working mothers are compatible, and why I observe larger responses from women with more than one child.

the EITC causing this rise in maternal employment. Traits associated with these heterogeneous responses are also used in section VI.F to predict state-level EITC responses and test whether states with larger EITC responses had larger post1975 increases in approval of working women.

B. i. **Heterogeneous Treatment Effects: Marital Status**

There are at least two reasons why married mothers should have responded less to the EITC than unmarried mothers. First, since EITC eligibility is determined by household earnings, spousal earnings often pushed the household out of EITC eligibility (point C in Figure 4.A). Second, spousal earnings increased the likelihood that the highest feasible indifference curve is achieved with zero labor supply (point A in Figure 4.A).

I verify this heterogeneity in Table 3 column 1, where I add the variable $Mom \times Post1975 \times Unmarried$ to equation (2) and interpret its coefficient (5.3 percentage points) as the treatment effect of the EITC on unmarried mothers relative to married mothers. I interpret the sum of the two coefficients in column 1 (6.9 percentage points, or 10.7 percent from a base of 64.5 percent) as the overall effect of the EITC on unmarried mothers.[26]

To estimate the effect of the EITC on married mothers, I carry out two approaches. In column 1, I pool all women; in column 2, I restrict the sample to married women. These approaches yield statistically significant estimates of 1.6 and 2.1 percentage points and align with prior EITC research that has consistently found a larger response among single mothers.[27]

[26]For comparison, the 1986 EITC expansion increased the number of unmarried working mothers by 2.8 percentage points (Eissa and Liebman 1996), the 1990s EITC expansion was responsible for a 6.1-percentage-point increase (Hoynes, Miller and Simon 2015), and the combined 1984-1996 EITC expansions increased the employment of unmarried mothers by 7.2 percentage points (Meyer and Rosenbaum 2001).

[27]See Eissa and Liebman (1996), Meyer and Rosenbaum (2001), Grogger (2003), and

Although I find a small positive *average* response among married mothers to the 1975 EITC, there should have been substantial heterogeneity that varied by spousal earnings. Mothers with very low spousal earnings should have responded to the EITC much like unmarried mothers. Restricting the sample to EITC-eligible married women with spouses earning below the EITC kink point,[28] the EITC increased the employment of this group by 4.5 percentage points.[29] I also test for a negative correlation between spousal earnings and EITC response by adding a variable to equation (2) that interacts $Mom \times Post1975$ with spousal earnings. Column 5 shows that the treatment effect on married women with zero spousal earnings was 4.6 percentage points and declined by 0.5 percentage points for every $10,000 (2013 dollars) in spousal earnings.[30]

Married mothers with spouses earning above the 1975 EITC kink point were not eligible for the EITC and faced the same work incentives before and after 1975. If it appears that the EITC increased the employment of this placebo group of mothers, this could indicate that an omitted factor is biasing-up the results. However, Table 3 column 4 shows a null effect on this placebo group and small effects can be statistically ruled out.

Hotz and Scholz (2006) for responses of unmarried mothers, and Ellwood (2000), Eissa and Hoynes (2004), and Bitler, Hoynes and Kuka (2016) for responses of married mothers.

[28]This sample is restricted to married women with spouses earning below $18,000 in 2013 dollars (the 1975 EITC kink point) in each year; the bottom fifth of spousal earnings.

[29]This result is nested in Figure A.5 which uses the entire spousal-earnings distribution and shows the largest EITC responses came from women with the lowest earning spouses.

[30]I verify that this pattern is also evident for the 1986 and 1993 EITC expansions (Table A.3). Results are robust to using actual or predicted spousal earnings. Table 3 treats a married woman's work decision like a second mover in a two-person sequential game, where the primary earner's labor supply does not depend on his spouse's labor supply (Eissa and Hoynes 2004). This assumption may not be completely unrealistic since 1970s-male labor supply was inelastic (Blundell and MaCurdy 1999). Also, the EITC is based on household earnings and no additional EITC benefits should arise from substituting labor supply between spouses. Heterogeneous responses among married women are also found by (Eissa and Hoynes 2004, Table 8) and Eissa and Hoynes (2006*b*).

B. *ii.* Heterogeneous Treatment Effects: Education

Education is often used as a proxy for EITC eligibility[31] and generally considered to be a fixed characteristic unlikely to be endogenous with the EITC. Table 3 column 6 adds two variables to equation (2), $Mom \times Post \times (< 12\ YrsEd)$ and $Mom \times Post \times (12 - 15\ YrsEd)$, so that the coefficient on $Mom \times Post$ denotes the treatment effect for mothers with at least 16 years of education and the other two coefficients denote the treatment effect relative to higher-education mothers. EITC response should be negatively correlated with education and mothers with a college degree are a quasi-placebo group, unlikely to have household earnings below the EITC income limit.[32] In line with this prediction, I find that mothers with less than 12, between 12 and 15, and 16 or more years of education had employment responses to the EITC of 6.1, 4.5, and -1.1 percentage points (or 13.4, 8.3, and -1.8 percent).[33]

B. *iii.* Heterogeneous Treatment Effects: "High-Impact" Group

Another way to verify larger effects from mothers most affected by the EITC is to construct a "high-impact" sample that omits EITC-ineligible married mothers with higher-earning spouses (Table 3 column 4) as well as women less able to respond to the employment incentives of the EITC: disabled, retired, and full-time students.[34] I estimate the effect on this group

[31] Sample women with less than, exactly, and more than 12 years of education have average household earnings of $21,000, $45,000, and $53,000 (2013 dollars).

[32] Low-education mothers were more than twice as likely to be EITC-eligible as high-education mothers (42 and 20 percent).

[33] I also find larger responses among younger mothers, mothers of younger children, and similar responses from white and nonwhite mothers (see Table A.4).

[34] The fraction of women in this sample smoothly increased over time, due to falling marriage rates. In the 1970s, disability rates were were slowly rising (Autor and Duggan 2003) and educational attainment was steadily increasing (Figure A.2).

by adding a variable to equation (2) that interacts $Mom \times Post1975$ with a binary for being in this "high-impact" group. The two estimates in Table 3 column 7 show that these mothers had an EITC response of about 5.1 percentage points (or 8.1 percent).

B. *iv.* **Heterogeneous Treatment Effects: Men**

Since most males were already working in the 1970s (over 90 percent), and their participation elasticity was near zero (Blundell and MaCurdy 1999), it should not be surprising that the EITC had no detectable effect on males, (0.3 percentage points) in Table 3 column 8.

C. **Triple Differences Corroborate DD Estimates**

Splitting the sample of mothers into EITC-eligible and EITC-ineligible (Table 3 column 4) creates a third difference for triple differences (DDD).[35]

$$(3) \qquad P(E_{ist}) = f(\beta_1 Mom \times Post1975 \times Treat_{ist} + \beta_2 X_{ist} + \delta_{st} + \epsilon_{ist})$$

The estimate of β_1 is 2.5 percentage points (Table 4 column 1), similar to DD, and suggests that factors affecting all mothers (e.g. abortion and divorce laws, birth control) may not pose a threat to the DD estimates.[36] When men from Table 3 column 8 are used as a comparison group, I find a similar DDD estimate in Table 4 column 2 (2.6 percentage points).

[35] An omitted factor affecting the employment of all mothers could bias DD (discussed in section II.B), which is why DDD "may generate a more convincing set of results" (Angrist and Pischke 2009, p.182).

[36] Equation (3) also controls for $Treat$, $Mom \times Treat$, $Post1975 \times Treat$, $Mom \times Post1975$, along with interactions of each control with $Treat$ for a more flexible model.

D. Extensive Margin Results: Annual DD Estimates

I estimate annual effects of the EITC and test if the DD results are driven by outliers or general trends by replacing $Mom \times Post$ in equation (2) with $Mom \times Year_y$ for $y \in [1970, 1985]$. I omit $y = 1975$ and estimates measure the annual effect of being a mother on the probability of working relative to 1975. Using the "high-impact" sample, Figure 1.B shows that these estimates closely resemble the unadjusted time-series trend. Relative to 1975, the estimates on $Mom \times Year_y$ are jointly insignificant (p-value 0.42) for $y \in [1970, 1975]$, become increasingly positive for $y \in [1975, 1979]$, and remain positive and relatively stable for $y \in [1979, 1985]$. The 1975-to-1979 increase may suggest it took mothers a few years to learn about the EITC, similar to the response to the 1986 and 1993 EITC expansions (Eissa and Liebman 1996; Meyer and Rosenbaum 2001).[37]

IV. Annual Work Hours and Earnings

A. Average Treatment Effects

Results above show that the EITC increased maternal employment and imply that earnings and work hours should also have been affected. Results in Table 4 use equation (2), an OLS specification, and replace the binary employment outcome with annual work hours and earnings (in 2013 dollars). For each outcome, I show results for three samples of women: the "high-impact" group (from Table 3 column 7), all women (from Table 2),

[37]The EITC does not pay until the following tax refund; it could take a year before EITC recipients became aware of the EITC (Liebman 1998). To test whether EITC response required an understanding of the tax code (Chetty, Friedman and Saez 2013, Bhargava and Manoli 2015), I plot the annual response by education subgroup and do not find quicker responses by higher-education mothers (omitted).

and the EITC-ineligible placebo group (from Table 3 column 4). Among the "high-impact" sample, the EITC led to increases of 63.9 annual work hours and $1249.1 in annual earnings (Table 4 columns 1 and 4). Among the sample of all women, the EITC led to smaller increases in work hours (35.1) and earnings ($750.3) (columns 2 and 5). Results capture both intensive and extensive margins, but primarily reflect participation responses.[38] Among the placebo group, columns 3 and 6 show that the EITC had a statistically insignificant effect on work hours (2.4) and earnings (438), which corroborates the placebo test in Table 3 column 4.

B. The EITC and the Distribution of Hours and Earnings

Where in the earnings and work-hours distribution did these newly working mothers enter? To investigate this, I estimate regressions resembling equation (2) but with a binary outcome variable for having annual earnings or work hours in a particular range. Figures 5 and A.6 show the DD estimates using the "high-impact" sample to focus on mothers most affected by the EITC. These figures also serve as robustness checks since it would raise concerns if these newly working mothers earned above the EITC limit.

For annual earnings (in 2013 dollars), the most common response to the EITC was to earn between $10,000 and $20,000, which encompassed the most generous portion of the EITC schedule (Figure 5) and suggests that many of these newly working mothers received the EITC. The minimum wage during this period was $7 to $9 per hour, and since Figure A.6 shows that many mothers began working full time, this maps to about $14,000

[38]See Figures 5 and A.6 for evidence. As a percent, these four estimates are 8.1, 10.3, 5.7, and 7.3. Although some people in or beyond the EITC phase-out region had an incentive to decrease labor supply to receive the EITC, there is little evidence for this (Meyer 2002, Saez 2002, Eissa and Hoynes 2006*a*); although see Kline and Tartari (2016).

to \$18,000 per year, consistent with Figure 5. Figure 5 also suggests that mothers were slightly more likely to earn between \$20,000 and \$50,000.

Figure A.6 shows that the most common response to the EITC was to work full-time, full-year (about 2000 annual hours)[39] and may also have increased part-time work, although estimates on annual hours below 2000 are not statistically significant. Consistent with previous results, mothers were less likely to have zero work hours or earnings (Figures 5 and A.6).[40]

Using IRS SOI data (see footnote 18), I also find suggestive evidence that the EITC affected the composition of tax filers. Consistent with Table 3 column 1, the fraction of unmarried tax filers increased after 1975 in a pattern similar to Figure 1.B (see Appendix B.6 and Figure B.3).

C. **Quantile Analysis**

I now characterize the effect of the EITC on the distribution of earnings. I use the regression behind Table 4, but instead of average effects, I estimate the effect at each centile of the earnings distribution. Instead of minimizing the sum of squared residuals like OLS, quantile regression uses heteroskedasticity as a feature of the data and minimizes a weighted sum of the absolute value of the residuals (Koenker 2005). These quantile difference in differences (QDD) are effects on quantiles, not on individual mothers, since rank preservation would require strong assumptions or panel data (see Bitler, Gelbach and Hoynes (2003)). Using the "high-impact" sample, Figure 6 shows that the EITC had the largest effect on the annual earnings of the

[39] Annual hours combines the categorical *weeks worked last year* variable (continuous variable not available until 1976 CPS) and *hours worked last week*, in an attempt to reduce measurement error (Bound, Brown and Mathiowetz 2001).

[40] To isolate intensive-margin responses, I re-run the analysis in Figures 5 and A.6 conditional on working, and find (noisy) evidence of more mothers working over 1000 hours and earning between \$10,000 and \$20,000.

43th centile, with a positive but decreasing effect higher up the earnings distribution, eventually becoming statistically insignificant for the top centiles. Work hours yield a similar pattern. The EITC had no effect on the lowest four centiles as these mothers did not work before or after 1975. Together, these QDD estimates drive the average effects in Table 4.

V. Implied Elasticities

I follow (Chetty et al. 2012, Appendix B) and calculate the participation elasticity as the pre1975-post1975 change in log employment rates divided by the pre1975-post1975 change in the log net-of-tax earnings from working. I account for various taxes (EITC, income tax, payroll tax, dependent deduction) and transfers (AFDC, food stamps, WIC). I calculate this elasticity for a representative unmarried mother of one child with the average pre-tax earnings of such a mother in the sample ($19,000, in 2013 dollars).

I estimate an elasticity between 0.41 (0.11) and 0.49 (0.13). See Table C.1 for complete details. Accounting for public assistance take-up rates yields a slightly larger elasticity between 0.45 and 0.54. Finally, I estimate the total intensive plus extensive margin elasticity from the annual work hours and earnings estimates in Table 4 to be 0.37 (0.10) and 0.47 (0.13). These elasticity estimates are larger than those for more recent decades, but are consistent with elasticity estimates for this period.[41]

[41] Female labor-supply elasticity has steadily declined since World War II (Goldin 1990): Bowen and Finegan (1969) finds 0.67 in 1960; Fields (1976) finds 0.52 in 1970; Blundell and MaCurdy (1999) shows that empirical studies using data from the 1970s and 1980s produce an average estimate of about 0.8; Blau and Kahn (2005) and Heim (2007) find an uncompensated elasticity of about 0.6 in 1980. Mroz (1987) discusses many of these early studies. The 1968-1982 negative income tax experiments yielded elasticities of 0.2 to 0.3 (Burtless and Hausman 1978, Robins 1985). Chetty et al. (2012) finds a range of 0.30 to 0.45. Elasticities are a function of the tax code (Saez, Slemrod and Giertz 2012) and vary across populations and time.

VI. Effects on Attitudes Towards Working Women

If the 1975 EITC encouraged a million mothers to begin working, this likely had subsequent effects on the country. Although there is a large literature showing that the EITC has benefited children of EITC recipients (see section I), how this program may have affected social attitudes towards working women has remained understudied.[42]

Google ngrams (Michel et al. 2011) show that in the mid-1970s, the phrases *working mom* and – the previously redundant – *stay at home mom* began to be used much more often (Figure 3). This suggests that the rise of working mothers was a salient phenomenon and reflects changes in language and attitudes towards the role of women in society. After 1975, people were more likely to have working-female family members, friends, and coworkers, while media stories about working mothers also became more common.[43]

An emerging literature shows that gender-equality preferences can be altered via *exposure* to working women. Fernández, Fogli and Olivetti (2004) and Olivetti, Patacchini and Zenou (2016) show that having a working mother – and having friends with working mothers – during childhood leads to stronger gender-equality preferences in adulthood.[44] Finseraas et al.

[42]Exposure to working women could theoretically have increased or decreased approval of working women. Analysis in section VI fits into an economics literature analyzing the role of attitudes and social norms (Becker 1957, Arrow 1971, Akerlof and Dickens 1982, Akerlof and Kranton 2000, Bénabou and Tirole 2006). Gender-role preferences are passed on intergenerationally (Fernandez and Fogli 2009, Alesina, Giuliano and Nunn 2011, Farré and Vella 2013) and affect female labor market outcomes (Fortin 2005, Charles, Guryan and Pan 2009, Bertrand, Kamenica and Pan 2015, Fortin 2015, Pan 2015, Janssen, Sartore and Backes-Gellner 2016). Unlike these studies, my goal is to characterize a determinant – not consequence – of these attitudes. There is also a long-standing sociology literature describing the time trends and correlates of these attitudes (Thornton and Freedman 1979, Thornton, Alwin and Camburn 1983, Plutzer 1988, Lottes and Kuriloff 1992).

[43]Media has been shown to affect teen pregnancy (Kearney and Levine 2015), divorce (Chong and Ferrara 2009), and fertility (La Ferrara, Chong and Duryea 2012). See DellaVigna and Ferrara (2015) for a recent literature review.

[44]Additional evidence that various attitudes can be altered via exposure has also been

(2016) shows that exposure to female colleagues reduces discriminatory attitudes.[45] With these results in mind, the attitudes of millions of Americans may have been affected when the EITC led one million mothers to begin working in the late 1970s.

A. **Empirical Strategy**

I characterize and exploit geographic heterogeneity in EITC responses and use a two-sample two-stage approach to test whether states with larger EITC responses had larger changes in gender-equality preferences. Gender-equality preferences are defined as approving of working women and are created from General Social Survey (GSS) data, an appealing source for measuring these social attitudes since the survey question is consistent over time and begins in 1972, providing a few baseline years before 1975.[46] Table A.5 shows GSS sample summary statistics and Table A.6 shows gender-equality preferences are positively correlated with education, having a working mother, and being younger, female, unmarried, and white.

I aggregate the gender-equality preferences of 8,713 adults, ages 18-60, observed between 1972 and 1985, to the state-by-year level using GSS weights. I then construct a state panel on gender-equality preferences before and

shown by Finseraas and Kotsadam (2015) (ethnic minorities), Beaman et al. (2012) (female aspirations), Stouffer et al. (1949) (race), and experimental evidence (Heilman and Martell 1986, Lowery, Hardin and Sinclair 2001, Dasgupta and Asgari 2004). This concept is related to psychology concept of intergroup contact theory (Allport 1954).

[45] Attitude changes consist of individual and intergenerational changes (Firebaugh 1992). Fernández, Fogli and Olivetti (2004) and Olivetti, Patacchini and Zenou (2016) focus on intergenerational change, Finseraas et al. (2016) on individual change. Fernández, Fogli and Olivetti (2004, footnote 1) acknowledges individual change: "as more women joined the labor force, attitudes towards these women changed in society at large." My approach captures both channels and tests how individual attitude changes aggregate.

[46] The GSS question asks, "Do you approve or disapprove of a married woman earning money in business or industry if she has a husband capable of supporting her?" Such approval rose from 20 to 80 percent between the 1930s and the 1990s (Figure A.7).

after 1975 and create the variable $\Delta GenderEquality_s^{(1976-85)-(1972-75)}$ – the change in the fraction of a state's adults that approve of working women – by subtracting the 1972-1975 state average from the 1976-1985 state average.[47]

I use March CPS data and the full sample and full set of controls from Table 2 column 4 to estimate the state-level, EITC-led increase in working mothers (i.e. state EITC response).

$$(4)\ \ P(E_{ist}) = f(\beta_1 Mom_{ist} + \sum_s \beta_{2s} Mom \times Post1975_{is} + \beta_3 X_{ist} + \delta_{st} + \epsilon_{ist})$$

Equation (4) modifies the national-level DD in equation (2) and estimates β_{2s}, state-level DDs.[48] I rename $\hat{\beta}_{2s}$, $EITC\ Response_s$, and estimate:

$$(5)\ \ \Delta GenderEquality_s^{(1976-85)-(1972-75)} = \gamma EITC\ Response_s + \delta \Delta X_s + \epsilon_s.$$

γ measures the effect of a percentage-point increase in state EITC response on the change in the fraction of a state's population with gender-equality preferences after 1975. Since the treatment variable is a generated regressor, standard errors are bootstrapped (Pagan 1984, Hardin 2002, Murphy and Topel 2002). X_s are controls to account for state-level traits. Regressions are weighted by state population since observations represent grouped data.[49]

[47] Results are robust to extending the GSS sample to any year between the early 1980s and the 1990s (Figure A.8). Ideally, I would construct a state-by-year panel, but since GSS samples are relatively small I pool years to increase statistical power.

[48] Results are robust to estimating equation (4) with various sets of controls (including state-by-year fixed effects), using OLS, probit, or logit, and ending the CPS sample in any year between 1979 and 1985 (Table A.7).

[49] Similar results if unweighted or weighted by the standard-error inverse (equation 4). Equation (5) is a first-difference estimator, which nets out the problem of omitted variables and is unbiased and consistent under the condition $E[u_{it} - u_{it-1} | x_{it} - x_{it-1}] = 0$, which is less restrictive than the assumption of weak exogeneity for unbiasedness when pre1975 and post1975 components are separated in a fixed effects estimator (Wooldridge 2015). This second approach produces similar, but noisier, estimates.

B. Results

Using equation (5) and no controls, Figure 7 shows that each percentage-point increase in state EITC response led to a 2.0-percentage-point increase in state-level preferences for gender equality (p-value 0.001).[50] Results are similar with region fixed effects, reflecting changes within and across regions, and are similar by gender (Table 7 columns 2 and 3).[51] Appendix D shows that less parametric approaches yield similar results.

One threat to my hypothesis would be if changes in gender-equality preferences coincided with changes in demographics or other attitudes unrelated to the EITC or working women, implying that an omitted trend is driving the results in Figure 7. To test this, I re-estimate equation (5) with controls for various demographic, political, and social-attitude variables.[52] The effect is stable between 1.6 and 2.3 percentage points, even when all 13 controls are included together (Table 6). Changes in gender-equality preferences do not seem to be driven by demographics or general trends in social attitudes.[53]

[50]The estimated magnitude appears plausible: the interquartile effect is 5.7 percentage points, comparable to having two more years of education or being a decade younger, but less than having a working wife or having racial-equality preferences (Table A.6).

[51]I test how likely Figure 7 is due to chance with a variant of the permutation test in Buchmueller, DiNardo and Valletta (2011): I randomly reassign a new attitude change to each state (with replacement) from the set of state attitude changes, re-estimate equation (5), record γ, and iterate 10,000 times. Figure A.9 shows that the actual estimate (0.0177) is in the top 0.07 percent of permutations and unlikely to occur by chance.

[52]Controls are education, age, marriage, number of mothers, race, employment, earnings, whether mother worked and mother's education, fraction Democrat and religious, and views on public assistance and racial-equality. Table 6 Panels A and B control for the pre1975-post1975 change and the pre1975 level. See Table A.5 for summary statistics.

[53]Although it is impossible to control for every state trait that may be correlated with increases in working mothers and with attitudes towards working women, the GSS has data on a wide range of topics (e.g. racial attitudes, voting behavior, religion, attitudes towards public assistance, mother's work and education). Furthermore, state-level response to the EITC (estimated in equation 4) accounts for changes in demographic traits and economic conditions and isolates the increase in working mothers due to the EITC.

C. Dose Response

If the EITC did affect gender-equality preferences through exposure to working women, then people more likely to know these newly working women should have had larger preference changes. Since the EITC had a larger effect on lower-education mothers (Table 3 column 6), lower-education adults were more likely to know (or even be) these women. Table 7 columns 4 and 5 re-estimate equation (5), but divide the sample into adults with more or less than 12 years of education. For lower-education adults, the estimate of γ in equation (5) is 0.023 (p-value 0.001) and for higher-education adults it is 0.005 (p-value 0.63). These estimates are statistically different at the 99-percent level and confirm that people more likely to know these newly working women did have larger preference changes.

D. Placebo Outcome: Changes in Racial-Equality Preferences

Since attitudes towards gender and race were correlated with the same traits (Table A.6), it is conceivable that an omitted factor – other than the EITC – was driving changes in various types of attitudes. One way to test for this is to use racial attitudes as a control (Table 6 column 7). Another approach is to use racial attitudes as a placebo outcome: Table 7 column 6 shows that state EITC responses had no detectable effect (p-value=0.19) on racial-equality preferences after 1975.[54] Changes in gender-equality preferences do not seem to be driven by general trends in attitudes.

[54]The relationship between EITC response and changes in racial attitudes is even less significant (p-value 0.79) when education is controlled for (Table 7 Panel B column 7).

E. Ruling Out Reverse Causation and Mean Reversion

Perhaps the most obvious threat to the results in Figure 7 is reverse causation: that is, if higher-responding states already had higher approval of working women before 1975. In Table 7 columns 7 and 8, I follow the approach in Acemoglu, Autor and Lyle (2004) and test for a positive relationship between state EITC response and pre1975 gender-equality preferences. I find an insignificant relationship between state EITC response and the 1972-to-1975 preference trend (p-value 0.70), and interestingly, a *negative* relationship between state EITC response and the 1974 preference level.[55] This negative estimate suggests that the EITC may have led to an attitude "catch up" among states with lower gender-equality preferences.[56]

Since states with the *lowest* approval of working women before 1975 had the *largest* increase in approval of working women after 1975, it is possible that Figure 7 simply due to mean reversion. In this context, mean reversion could reflect data limitations and relatively small GSS sample sizes, or real convergence in social norms across states over time. One way to test for mean reversion is to see if states with higher EITC responses (and lower approval of working women) continued to have larger increases in approval of working women in the 1980s and 1990s. As shown by Charles, Guryan and Pan (2009), states with the lowest approval of working women in the 1970s also had the lowest approval of working women in later decades.[57] If

[55]Although the relationship between pre1975 attitudes and EITC response becomes statistically insignificant when education is controlled for (Table 7 Panel B column 7).

[56]If states that voted for the EITC benefited the most from it, perhaps the EITC was the outcome, not the cause, of changing attitudes. To test this, I regress state EITC response on the fraction of a state's Senators and House Representatives that voted for the 1975 EITC legislation. Figure A.10 shows that, in fact, the opposite is true: states voting against the EITC had higher EITC responses and thus preference changes were larger in places less likely to be in favor of a social program like the EITC.

[57]I also find a strong positive correlation between state-level gender-equality prefer-

mean reversion drove attitude changes after 1975, it should also have driven attitude changes in later decades. Figure A.11 re-estimates equation (5), but instead of 1975, measures attitude changes after placebo years in the 1980s and 1990s. I find that EITC response had no apparent relationship with changes in gender-equality preferences after these placebo years, suggesting that mean reversion may not explain post1975 preference changes either.

Another way to investigate whether mean reversion explains Figure 7, is to see if state EITC response is still associated with changes in attitudes when controlling for pre1975 attitudes. Table 7 column 9 shows that while pre1975 attitudes are significantly associated with post1975 attitude changes (corroborating column 7), EITC response continues to have an independent effect on attitude changes; although the estimate falls from 0.020 to 0.013, perhaps suggesting that a third of the estimate in Figure 7 may be due to mean reversion. Panel B takes this approach one step further and re-runs each regression in Panel A with controls for pre1975 education, a trait associated with EITC response and social attitudes: EITC response continues to have an independent effect on attitude changes even when attitudes and education are controlled for (estimate 0.013 [0.006] in column 9).[58]

F. 2SLS and Predicted State EITC Responses

In this section I exploit pre1975 state traits, $X_s^{pre1975}$, and the heterogeneous EITC responses in Table 3 to *predict* state EITC response and test whether *predicted* EITC response is associated with changes in gender-equality preferences. This two-stage-least-squares approach helps alleviate

ences in the 1970s, 1980s, and 1990s (not shown).

[58]The estimate of EITC response remains positive (0.017 [0.005]) when pre1975 attitudes and all 13 controls from Table 6 column 9 are used.

concerns about the potential endogeneity of attitudes and EITC response.

To show that *predicted* state EITC response affected preferences, four conditions should be met. First, the reduced-form version of the two-step regression: $X_s^{pre1975}$ should be correlated with $\Delta GenderEquality_s^{(1976-85)-(1972-75)}$. Second, the 2SLS first stage where $\overline{EITC\ Response}_s$ (predicted state EITC response) is generated: $X_s^{pre1975}$ should be correlated with $EITC\ Response_s$. Third, the 2SLS second stage: regressing $\Delta GenderEquality_s^{(1976-85)-(1972-75)}$ on $\overline{EITC\ Response}_s$ should be correlated and interpreted as the effect of an exogenous increase in maternal employment on gender-equality preferences.[59] Fourth, $X_s^{pre1975}$ should not be correlated with gender-equality preferences before 1975. Conditions one and four together would suggest that $X_s^{pre1975}$ only affected preferences indirectly through state EITC response.

Figure 8 shows that these four conditions are met using female education. Panel A shows that pre1975 female education is negatively correlated with preference changes after 1975 (p-value 0.048). Panel B shows that female education is highly correlated with state EITC response (as expected from Table 3; p-value 0.001) and illustrates the best-fit line used to generate predicted state EITC response. Panel C shows that predicted EITC response is positively correlated with changes in preferences after 1975 (estimate 0.02, p-value 0.048). Finally, Panel D shows an insignificant (although noisy) relationship between female education and pre1975 preferences (p-value 0.14).

In Table A.7, I repeat the exercise in Figure 8 using other state-level demographic traits (e.g. single mothers, male earnings), with and without region fixed effects (Panel A and B). Though not all results are statistically significant, both *actual* and *predicted* state EITC responses suggest that the

[59]A complementary approach regresses attitude changes on the residuals from step two and shows that the correlation is zero (results omitted).

EITC positively affected gender-equality preferences.

G. External Validity: Attitude Changes After WWII

If the EITC-led increase in working women affected attitudes towards working women, then the same pattern should exist during other periods of large increases in female employment. During World War II, more than 7 million women began working – compared to a total of about 14 million women working in 1940 – to make up for the 14 million men that joined the military. More women worked in places with higher mobilization rates (Acemoglu, Autor and Lyle 2004, Goldin and Olivetti 2013).[60]

I follow the approach in equation (5), and construct a state panel on gender-equality preferences before and after WWII, using WWII mobilization rates as the treatment variable. Testing whether mobilization rates (and large increases in working women) affected social attitudes is feasible since Gallup began asking such questions in the 1930s (see Figure A.13 notes for details) and identifies individuals by state. I find that mobilization rates are strongly associated with increases in gender-equality preferences after WWII (p-value 0.003), providing corroborating evidence that increases in working women may affect attitudes about the role of women in society.

VII. Summary

In one of the first systematic studies of the 1975 introduction of the EITC, I find that this program led to a 6-percent increase in maternal employment,

[60]Two-thirds of these rates can be explained by exogenous factors (Goldin and Olivetti 2013). I focus on attitudes and mobilization of white adults, since WWII had a larger effect on white women: "black womens [labor force] participation was high before the war and many were in agricultural occupations" (Goldin and Olivetti 2013). Mobilization rates are not correlated with state responses to the 1975 EITC (p-value 0.38).

which represents about one million mothers and a participation elasticity of 0.49. Regression-adjusted and unadjusted time-series trends show that the relative employment of mothers began to increase after 1975 (Figures 1.A and 1.B). Consistent with the EITC being responsible for this rise in employment, I find larger responses from mothers more likely to be EITC eligible and null responses from placebo groups not eligible for EITC benefits (Table 3). Using the placebo group of EITC-ineligible mothers in a triple-differences specification to net out contemporaneous policies (e.g. birth control, divorce laws, abortion) yields similar estimates.

In hindsight, the employment effect of the 1975 EITC should not be that surprising: female labor-supply elasticity was larger during this period (Blau and Kahn 2005, Heim 2007) and the 10-percent wage subsidy of the EITC represented a large increase in potential earnings.[61] Although much was already known about the rise of working *women* (Killingsworth and Heckman 1986, Goldin 1990, Fernández, Fogli and Olivetti 2004), this study helps explain why so many *mothers* began working in the 1970s.

The 1970s also provide a clean policy environment to evaluate the effects of the EITC. By the 1980s, policymakers were cutting public benefits and nudging low-income women into the labor force, and the 1990s EITC expansion coincided with welfare reductions and the Family Medical Leave Act, which increased maternal employment (Ruhm 1998, Moffitt 1999).

This EITC-led increase in working mothers also appears to have increased approval of working women. States with larger EITC responses – and larger *predicted* responses based on pre1975 demographic traits – had larger in-

[61] This paper may also help resolve an anomaly observed by Smith and Ward (1985): although real wage growth explains most of the increase in the female labor supply between 1950 and 1980, after 1970, the growth rate of female labor supply rose as the real-wage growth rate fell (Parkman 1992).

creases in attitudes approving of women working. Results do not appear to be driven by changes in demographics or general trends in social attitudes, and are larger among people more likely to know these newly working mothers. As for external validity, I find similar attitude changes due to the large increase in working women during World War II. Since social attitudes towards working women and the number of working women are endogenous, I use two episodes of largely exogenous increases in female employment to show that increases in working women affect attitudes towards working women. I conclude that the 1975 EITC played an important role in the rise of U.S. working mothers and in fostering egalitarian social attitudes.

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VIII. Tables and Figures

Table 1. Summary Statistics

	All Women	Mothers	Women without Kids	Married	Not Married
Variable	(1)	(2)	(3)	(4)	(5)
Age	32.1	34.7	28.2	34.3	28.2
	(9.4)	(8.1)	(10.0)	(8.7)	(9.4)
Years of Education	12.3	12.1	12.7	12.3	12.3
	(2.6)	(2.5)	(2.6)	(2.5)	(2.6)
Married	0.64	0.82	0.36	1.00	0.00
	(0.48)	(0.38)	(0.48)	(0.00)	(0.00)
Black	0.12	0.12	0.12	0.07	0.20
	(0.32)	(0.32)	(0.33)	(0.26)	(0.40)
Hispanic	0.06	0.07	0.05	0.06	0.06
	(0.24)	(0.25)	(0.22)	(0.23)	(0.24)
Kids Under 5	0.24	0.39	0.00	0.32	0.09
	(0.43)	(0.49)	(0.00)	(0.47)	(0.29)
Number of Kids	1.35	2.22	0.02	1.76	0.63
	(1.45)	(1.24)	(0.14)	(1.43)	(1.19)
Employed	0.66	0.58	0.79	0.60	0.77
	(0.47)	(0.49)	(0.41)	(0.49)	(0.42)
Individual Earnings (2013 $)	$14,158	$11,854	$17,701	$12,936	$16,341
	(17573)	(16473)	(18592)	(17061)	(18249)
Individual Earnings (2013 $)	$21,418	$20,609	$22,322	$21,500	$21,304
(Conditional on Earnings > 0)	(17653)	(17069)	(18241)	(17310)	(18124)
Household Earnings (2013 $)	$45,822	$54,177	$32,969	$62,320	$16,341
	(40612)	(41378)	(35779)	(40341)	(18249)
Household Earnings (2013 $)	$52,312	$60,963	$38,501	$66,519	$21,304
(Conditional on Earnings > 0)	(39287)	(38896)	(35804)	(38181)	(18124)
Household Earnings Below EITC Limit	0.41	0.30	0.57	0.21	0.76
	(0.49)	(0.46)	(0.50)	(0.41)	(0.43)
Household Earnings Below EITC Limit	0.28	0.19	0.42	0.14	0.53
(Conditional on Earnings > 0)	(0.45)	(0.39)	(0.49)	(0.35)	(0.50)
Annual Weeks Worked	27.4	24.0	32.7	25.5	30.8
	(22.5)	(22.8)	(21.1)	(22.8)	(21.7)
Annual Weeks Worked	39.5	38.8	40.2	39.4	39.5
(Conditional on Weeks Worked > 0)	(16.0)	(16.2)	(15.7)	(15.9)	(16.2)
Weekly Hours Worked	19.3	16.5	23.7	17.5	22.5
	(19.7)	(19.2)	(19.7)	(19.3)	(19.9)
Weekly Hours Worked	34.9	34.0	36.0	34.3	35.8
(Conditional on Hours Worked > 0)	(12.5)	(12.7)	(12.0)	(12.5)	(12.3)
Observations	571,170	350,798	220,372	370,767	200,403

Notes: Data source: 1971-1986 March CPS data. Individual March CPS weights used. Sample contains all women 18 to 50 years old. Standard deviations are in parentheses. Kids under 5 is binary. 376,919 observations have positive earnings, 500,471 have positive household earnings, 397,210 have positive weeks worked last year, and 315,902 have positive hours worked last week.

Table 2. The 1975 EITC Increased Maternal Employment, Robust to Various Sets of Controls

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Mom x Post1975	0.048	0.049	0.040	0.033	0.034	0.042	0.035
	(0.006)	(0.006)	(0.006)	(0.007)	(0.007)	(0.007)	(0.007)
Controls							
State and Year FE		X	X	X	X	X	X
Demographic Controls			X	X	X	X	X
Unemployment Rate				X	X	X	X
"Kitchen-Sink" Controls							X
Observations	571,170	571,170	571,170	571,170	571,170	571,170	571,170
Model	Logit	Logit	Logit	Logit	Probit	OLS	OLS
R-squared						0.146	0.164
Mean Dependent Variable Across Years and Across Treatment and Control Groups=0.66							
Mean Dependent Variable for Treatment Group in 1975=0.53							

Notes: Data source: 1971-1986 March CPS data. Sample includes all women 18 to 50 years old. Dependent variable binary employment for having positive earnings. CPS weights, equation (2) used and average marginal effects from logit, probit, or OLS regression are shown. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level. FE denotes fixed effects. Demographic controls include married, welfare income, number of children, any children under 5, age cubic, years of education quadratic, nonwhite-mom, nonwhite-post1975, age-mom, and married-post1975. Unemployment rate includes state-year employment-to-population ratios and interactions with kid and married. "Kitchen-sink" controls include 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, and state-year-married, as well as annual inflation interacted with low education (<12 years), mom and number of kids, and married, and finally state-year measures of manufacturing employment interacted with low education, mom, and married.

Table 3. Heterogeneous and Subgroup Treatment Effects of the 1975 EITC on Employment

Subgroup:	All	Married				Education	"High-Impact"	Men
Description:	Larger Response Among Unmarried Mothers	All Married Women	Spouse Earning Below EITC Limit	Placebo: Spouse Earning Above EITC Limit	Response Negatively Correlated With Spousal Earnings	Larger Response Among Lower-Education Mothers	Larger Response Among "High-Impact" Sample	No Response Among Married Men
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Mom x Post1975	0.016	0.021	0.045	0.013	0.046	-0.011	0.018	
	(0.008)	(0.009)	(0.013)	(0.010)	(0.014)	(0.009)	(0.007)	
Mom x Post1975 x Unmarried	0.053							
	(0.014)							
Mom x Post1975 x Spousal Earnings					-0.005			
(in 10,000s of 2013 $)					(0.002)			
Mom x Post1975 x (<12 Yrs Ed)						0.050		
						(0.004)		
Mom x Post1975 x (12-15 Yrs Ed)						0.034		
						(0.006)		
Mom x Post1975 x High-Impact							0.033	
							(0.007)	
Dad x Post1975								0.003
								(0.004)
Observations	571,170	370,767	67,277	303,490	370,767	571,170	571,170	343,219
Mean Dependent Variable:	0.66	0.60	0.53	0.62	0.60	0.66	0.75	0.89
Mean Dep Var for 1975 Treat. Group:	0.53	0.51	0.46	0.52	0.51	0.53	0.58	0.89

Notes: Data source: 1971-1986 March CPS data. All samples limited to 18 to 50 year olds. Binary dependent variable employment for positive earnings. CPS weights, equation (2) used and average-marginal effects from logit regression are shown. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level. Each column reflects a separate regression with the full set of controls from Table 2 column 4. The variable *Mom x Post1975* implicitly refers to *Mom x Post1975 x Married* in column 1 and *Mom x Post1975 x (>16 Yrs Ed)* in column 6. In column 1, I avoid potentially endogenous marriage responses to the EITC by using OLS to predict marital status based on the following pre1975 traits: age, education, race, state, and a linear time trend. Results are similar using various sets of covariates used to predict marital status or using actual marital status. In column 5, I avoid potentially endogenous spousal-earning responses to the EITC by using OLS to predict spousal earnings based on the following pre1975 traits: age, education, race, number of kids, kids under 5. Results are similar using various sets of covariates used to predict spousal earnings or using actual spousal earnings. "High-impact" sample excludes EITC-ineligible women with spousal earnings above the EITC limit (column 4) and women not in the labor force due to a disability, health reason, or full-time student in order to capture women most in a position to respond to the employment incentives of the EITC. The mean dependent variable for the "high-impact" sample of moms in 1975 is 0.63 (see Figure 1.A.). The mean dependent variable for 1975 mothers with less than, equal to, or more than 12 years of education in column 6 is 0.45, 0.54, and 0.59. Results for men in column 8 are similar for married or unmarried men. The EITC income limit in columns 3 and 4 was $4,000 nominal dollars in 1975 and increased to $5,000 in 1979 (or about $18,000 in 2013 dollars). For years before 1975, I convert spousal earnings to real 1975 dollars and use a $4,000 cutoff.

Table 4. Triple Differences Corroborates Difference in Differences

Third Difference:	EITC-Eligible Mothers vs Non-EITC-Eligible Mothers	EITC-Eligible Mothers vs Fathers (in Table 3 Column 8)
Variables	(1)	(2)
Mom x Post1975 x EITC Eligible	0.025	
	(0.008)	
Parent x Post1975 x Woman		0.026
		(0.008)
Observations	571,170	610,899

Note: Data source: 1971-1986 March CPS data. Samples limited to 18 to 50 year olds. Binary dependent variable employment equals 1 for positive earnings. EITC-eligible mothers are unmarried mothers and married mothers with low spousal earnings (Table 3 columns 1 and 3). Non-EITC-eligible mothers have higher spousal earnings (Table 3 column 4). Sample of men in column 2 are married men (from Table 3 column 8). Results similar if unmarried men are used. Equation (3), CPS weights, full set of controls from Table 2 column 4 used along with interactions of each control with EITC-eligible mothers, and average-marginal effects from logit regression are shown. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level.

Table 5. The EITC Effect on Annual Work Hours and Earnings (Intensive + Extensive Margins)

Dependent Variable:	Annual Work Hours			Annual Earnings (2013 $)		
Sample:	"High-Impact" Group	All	EITC-Ineligible Placebo Group	"High-Impact" Group	All	EITC-Ineligible Placebo Group
Mean Dependent Variable:	1045	834	755	16,422	13,992	13,295
Mean Dep. Var. 1975 Mothers:	790	611	549	12,137	10,258	9,660
Variables	(1)	(2)	(3)	(4)	(5)	(6)
Mom x Post1975	63.9	35.1	2.4	1,249.1	750.3	438.0
	(15.4)	(11.8)	(17.1)	(219.7)	(219.1)	(394.9)
Observations	236,814	571,170	303,490	236,814	571,170	303,490
R-squared	0.222	0.168	0.140	0.305	0.214	0.170

Notes: Data source: 1971-1986 March CPS data. Each column represents a separate OLS regression with CPS weights and the full set of controls from Table 2 column 4. All samples limited to women 18 to 50 years old. High impact sample from Table 3 column 7. EITC-ineligible placebo group from Table 3 column 4. Annual work hours are constructed by multiplying weeks worked last year and hours worked last week. Weeks worked is given as an interval until 1975, I use this variable for all years to be consistent and assign the midpoint of the interval. Qualitatively similar results using imputed hourly wage (annual earnings divided by annual work hours, with zero assigned if annual work hours equals zero, even if reported annual earnings is positive) as outcome: 0.85 (.19), 0.38 (.015), and 0.07 (0.27), which represent percent increases of 11, 6, and 1 for 1975 mothers. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level.

Table 6. EITC Increased Preferences for Gender Equality, Robust to Various Controls

Controls:	Demographics						Other Social Attitudes			**Using All Controls**
	Educ.	Married	Mothers	Nonwhite	Earnings	Mom Worked	Fraction Democrat	Racial-Equality	Too Much Welfare	
Panel A: Controlling for Pre1975-Post1975 State-Trait Changes										
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
EITC-Led Increase in	0.019	0.019	0.020	0.020	0.019	0.020	0.022	0.017	0.021	0.023
Working Mothers	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.007)
(in Percentage Points)										
Observations	32	32	32	32	32	32	32	32	32	32
R-squared	0.383	0.312	0.292	0.293	0.328	0.325	0.429	0.338	0.387	0.706
Panel B: Controlling for Pre1975 State-Trait Levels										
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
EITC-Led Increase in	0.016	0.020	0.020	0.019	0.018	0.020	0.020	0.017	0.020	0.019
Working Mothers	(0.006)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.006)	(0.004)	(0.008)
(in Percentage Points)										
Observations	32	32	32	32	32	32	32	32	32	32
R-squared	0.326	0.301	0.303	0.298	0.311	0.328	0.316	0.310	0.344	0.590

Notes: 1972-1985 restricted GSS data with state-level identifiers. Gender-equality preferences constructed from the GSS variable *fework* which asks respondents whether married women should work. GSS sample reflects adults 18 to 60 years old in 32 states. State-level EITC response estimated from equation (4). The outcome variable and each control variable in Panel A is constructed by subtracting the pooled 1972-1975 GSS state-average from the 1976-1985 GSS state-average using GSS sample weights. Fernandez et al (2004) shows that mother's employment affects gender-role attitudes. Results are similar for additional controls not shown: average age, employment rate, mother's education, and fraction religious. All controls in column 9 refers to the controls used in columns 1-9 as well as the four additional covariates just mentioned. Education measured in years, mother is fraction of women that are mothers, employment rate denotes labor-force participation (working full-time, working part-time, temporarily laid off, or unemployed), earnings denotes real log earnings. Mom worked and mom education constructed from GSS variables *mawk16* and *maeduc*, democrat from *partyid*, racial equality from *racpres*, religious from *reliten,* and too much welfare from *natfare*. Heteroskedasticity-robust standard errors in parentheses. Regressions weighted by state population, though unweighted results are similar.

Table 7. EITC and Gender-Equality Preferences: Subgroups, Placebo Tests, and Ruling Out Mean Reversion

	Subgroup						Largest EITC Response in States with Lowest Pre1975 Approval of Working Women		Ruling Out Mean Reversion by Controlling for Pre1975 Attitudes
					Placebo Tests				
	All	Men	Women	Low-Educ.	High-Educ.	Racial-Equality Preference	1974 Attitudes	1972-1975 Attitude Trend	
Panel A: No Controls									
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
EITC-Led Increase in Working	0.020	0.016	0.022	0.023	0.005	0.012	-0.023	0.005	0.013
Mothers (in Percentage Points)	(0.005)	(0.007)	(0.006)	(0.006)	(0.010)	(0.009)	(0.008)	(0.013)	(0.005)
1974 Fraction of State Pop. that									-0.298
Approves of Working Women									(0.104)
Observations	32	32	32	32	32	32	32	32	32
R-squared	0.291	0.086	0.327	0.235	0.013	0.071	0.188	0.005	0.450
Panel B: Controlling for Pre1975 State-Level Average Education									
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
EITC-Led Increase in Working	0.016	0.009	0.023	0.022	0.003	0.002	-0.012	-0.004	0.013
Mothers (in Percentage Points)	(0.006)	(0.009)	(0.006)	(0.007)	(0.011)	(0.008)	(0.007)	(0.014)	(0.006)
1974 Fraction of State Pop. that									-0.304
Approves of Working Women									(0.132)
Observations	32	32	32	32	32	32	32	32	32
R-squared	0.326	0.159	0.328	0.236	0.019	0.275	0.388	0.068	0.451

Notes: 1972-1985 restricted GSS data with state-level identifiers. Gender-equality preferences constructed from the GSS variable *fework* which asks respondents whether married women should work. GSS sample reflects adults 18 to 60 years old in 32 states. State-level EITC response estimated from equation (4). The outcome variable is constructed by subtracting the pooled 1972-1975 GSS state-average from the 1976-1985 GSS state-average. Low and high education defined as less than 12 or at least 12 years of education. Although column 7 shows that states with the lowest approval of working women had the largest EITC responses, column 9 shows that even when these pre1975 attitudes are controlled for, state EITC response is still associated with an increase in approval of working women. Regressions are weighted by state population, though unweighted results are similar. Heteroskedasticity-robust standard errors in parentheses.

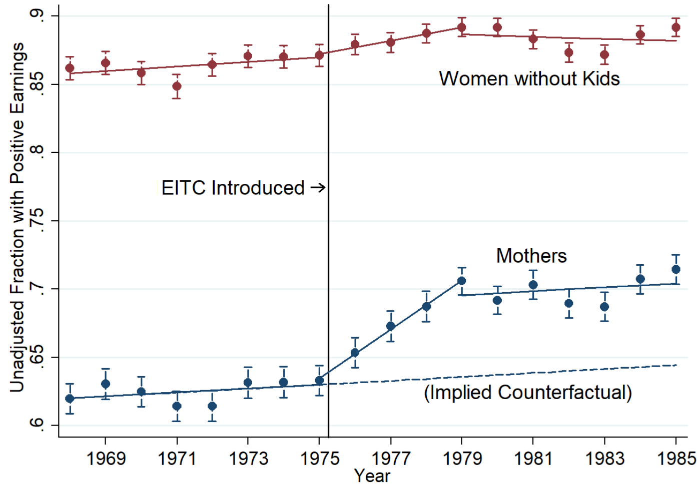

Figure 1.A. Unadjusted Employment Trends (High-Impact Sample)

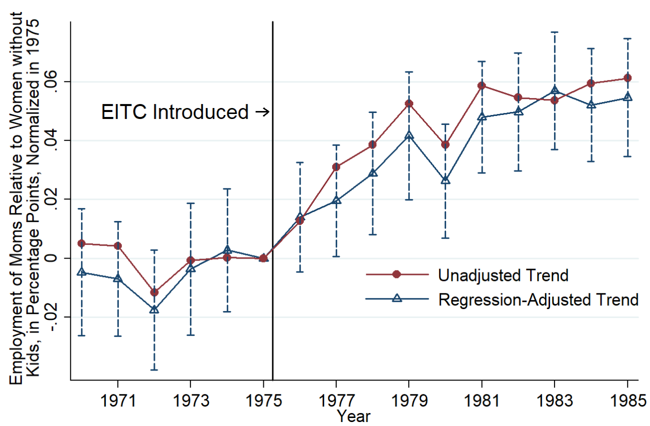

Figure 1.B. Unadjusted and Regression-Adjusted Employment Gap

Notes: 1969-1986 March CPS data. Employment defined as positive income. Best-fit lines shown for 1969-75, 1975-79, 1979-85. Regression adjusted employment gap estimates from a probit and the full set of controls from Table 2 column 4. The estimates are jointly statistically insignificant for all years before 1975 (p-value 0.42). "High-impact" sample includes all women 18-50 and excludes married women with spousal earning above $36,000 in 2013 $ (corresponding to the 1975 EITC income limit), full-time students, disabled, and retired. Kids are 0-18 years old, or 19-23 if a student. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level.

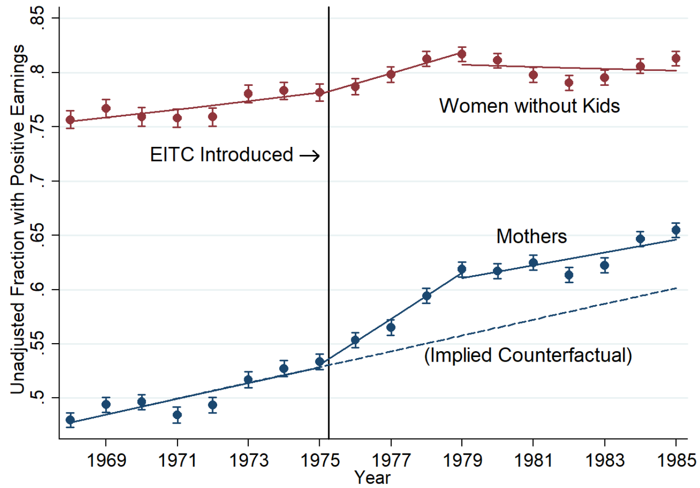

Figure 2.A. Unadjusted Employment Trends (All Women)

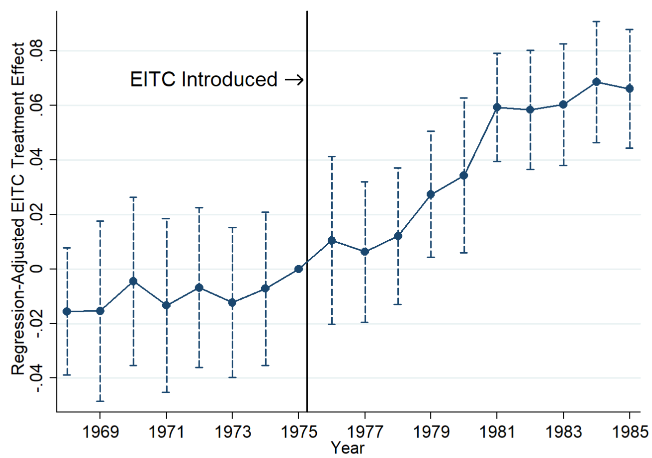

Figure 2.B. Unadjusted Employment Gap: Mothers vs Women without Kids

Notes: Sample includes all women 18-50. Figures are compiled in a manner similar to Figures 1.A and 1.B. Best-fit lines shown for 1969-75, 1975-79, 1979-85. The estimates are jointly statistically insignificant for all years before 1975 (p-value 0.38). The relative rise in maternal employment after 1981 appears to reflect a decline in employment among women without kids.

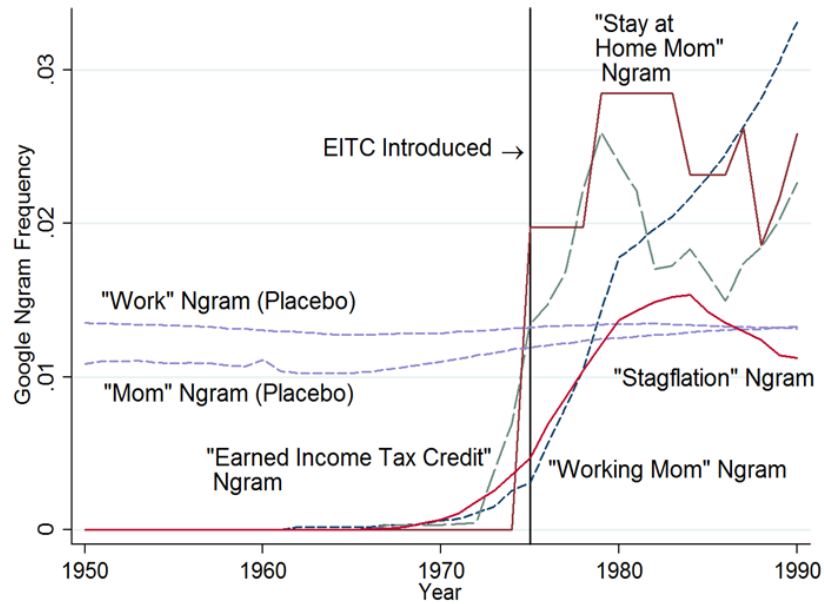

Figure 3. Rise of Working mothers Was Salient: Evidence from Ngrams

Notes: Google Books Ngram Viewer is an online search engine (http://books.google.com/ngrams) that charts frequencies of any set of comma-delimited search strings using a yearly count of ngrams found in over 5 million sources – and over 500 billion words – printed between 1500 and 2008 (Michel et al. 2011). This represents about a 4 percent sample of all possible books and sources. The vertical axis measures the relative frequency that each phrase is used in sources printed between 1950 and 1990. For scaling purposes, earned income tax credit is multiplied by 10,000, working mom is multiplied by 100,000, and stay at home mom is multiplied by 3,800,000. Because of this, the levels within ngrams are comparable over time but levels across ngrams are not. Each ngram includes plural and capitalized variants of these phrases; stay at home mom also uses variants of the word mother. Sources: https://books.google.com/ngrams/graph?content=working+moms&year_start=1950&year_end=1990&corpus=15&smoothing=10&share=&direct_url=t1%3B%2Cworking%20moms%3B%2Cc0, https://books.google.com/ngrams/graph?content=earned+income+tax+credit&year_start=1950&year_end=1990&corpus=15&smoothing=3&share=&direct_url=t1%3B%2Cearned%20income%20tax%20credit%3B%2Cc0, https://books.google.com/ngrams/graph?content=stay+at+home+mom%2Bstay+at+home+moms%2Bstay+at+home+mother&year_start=1950&year_end=1990&corpus=15&smoothing=4&share=&direct_url=t1%3B%2C%28stay%20at%20home%20mom%20%2B%20stay%20at%20home%20moms%20%2B%20stay%20at%20home%20mother%29%3B%2Cc0, https://books.google.com/ngrams/graph?content=working%2Bwork&year_start=1950&year_end=1990&corpus=15&smoothing=4&share=&direct_url=t1%3B%2C%28working%20%2B%20work%29%3B%2Cc0, https://books.google.com/ngrams/graph?content=mom%2Bmother%2Bmoms%2Bmothers&year_start=1950&year_end=1990&corpus=15&smoothing=4&share=&direct_url=t1%3B%2C%28mom%20%2B%20mother%20%2B%20moms%20%2B%20mothers%29%3B%2Cc0. Accessed 9/5/16.

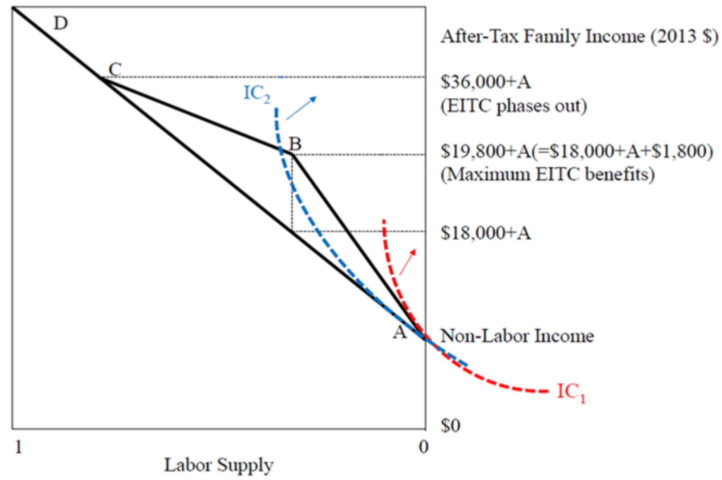

Figure 4.A. Budget Constraint Under the 1975 EITC

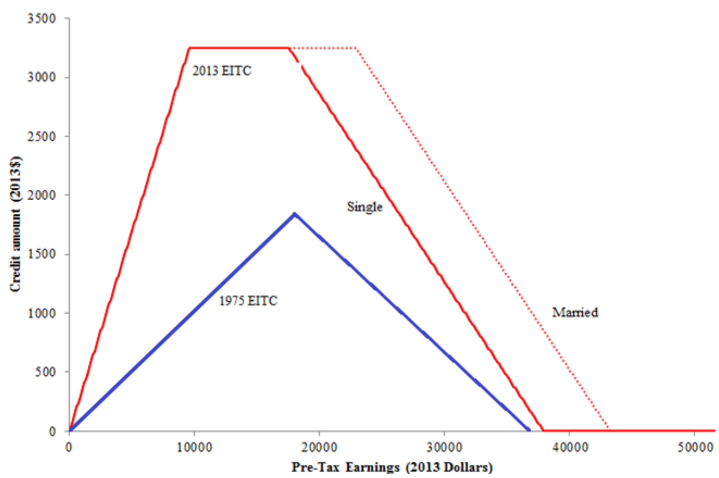

Figure 4.B. Comparing 1975 and 2013 EITC, Households with One Child

Notes: Author's calculation from 1975 and 2013 EITC parameters. 1975 EITC phased in and out at 10 percent. EITC benefits actually phase out with adjusted gross income. 2013 EITC for one child phased in and out at 34 and 15.98 percent. An abbreviated history of 1975-2013 changes to the EITC schedule is: the EITC began as a temporary program and was made permanent in 1978; 1979, a plateau region was added; 1986, the phase-in rate was increased to 14 percent and the EITC parameters were indexed to inflation; 1990, additional benefits available to parents with two children; 1993, benefits were extended to adults without children (at a rate of 7.65 percent); 1993 to 1996, the phase-in rate increased to 34 percent and 40 percent for households with one and two or more children; 2002, the plateau region was extended to married couples to decrease the marriage penalty; 2009, additional benefits available to parents with three children.

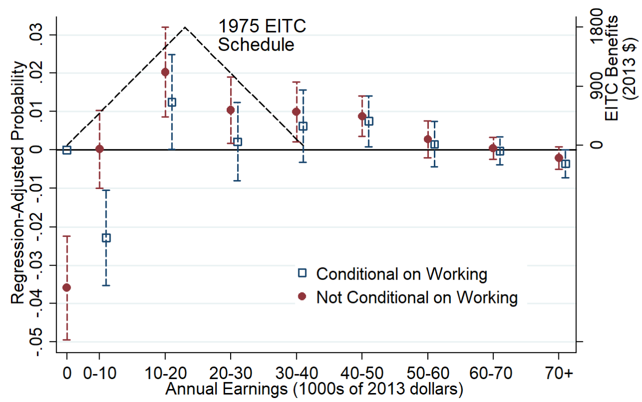

Figure 5. Effect of the EITC on the Distribution of Annual Earnings

Notes: 1971-1986 March CPS data. Full set of controls from Table 2 column 4 and "high-impact" sample used. Each estimate is from a different logit regression of having annual earnings in the specified range. Not conditional on working, the mean dependent variables are 0.25, 0.22, 0.16, 0.16, 0.10, 0.06, 0.03, 0.02, 0.01; and conditional on working are 0.0, 0.29, 0.21, 0.19, 0.14, 0.08, 0.04, 0.02, 0.02. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level.

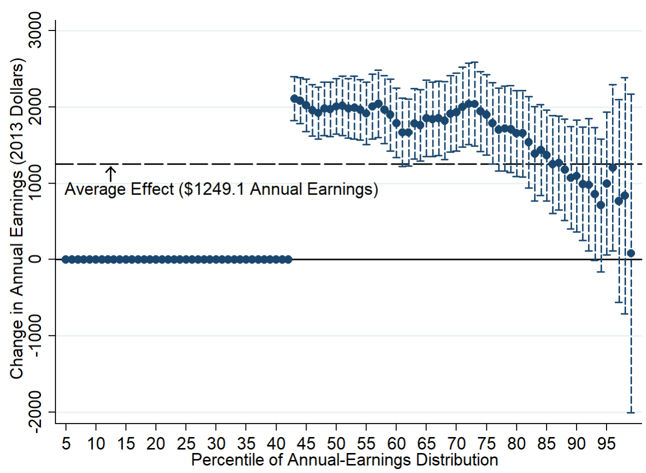

Figure 6. Effect of the EITC on Annual Earnings (Quantile Dif in Dif)

Notes: Same data, sample, controls, and standard errors as Figure 5. Mimics the regression behind Table 4 except instead of average effects, results shown are the effect of $Mom \times Post$ at each centile. The mean dependent variable at deciles 1 to 9 for mothers in 1975 are 0, 0, 0, 0, 2066, 8503, 17229, 25895, 36707.

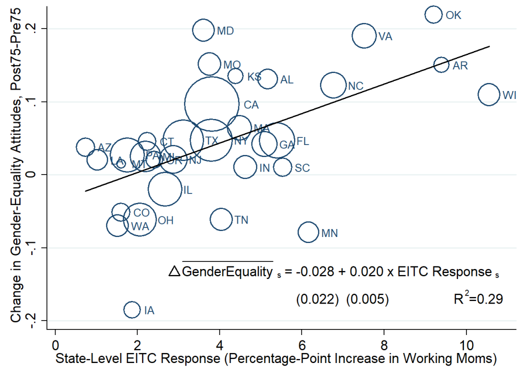

Figure 7. EITC Response and Increased Approval of Working Women

Notes: 1972-1985 restricted GSS data. Sample contains adults 18 to 60 years old. Changes in gender-equality attitudes is calculated by subtracting the pooled 1972-1975 state-average from the 1976-1985 average using GSS weights. Years are pooled to increase power. State EITC response is estimated from equation (5). Heteroskedasticity-robust standard errors shown. Regression weighted by state population.

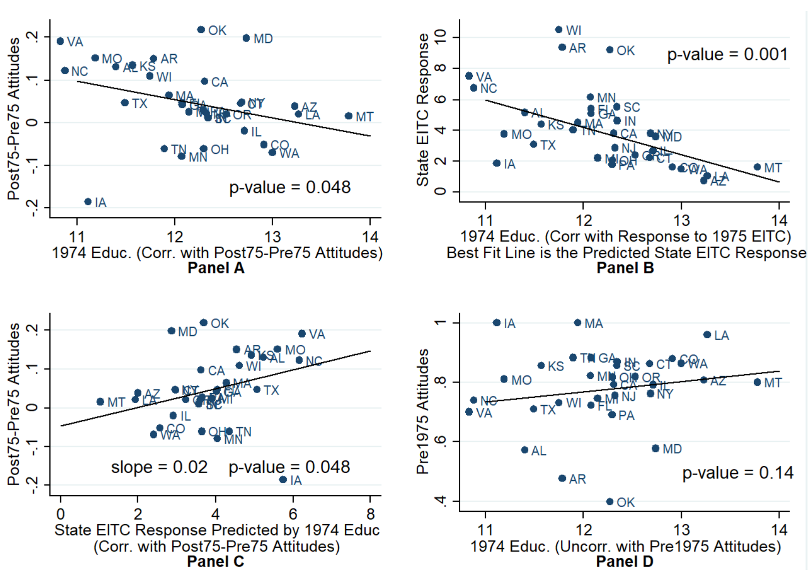

Figure 8. 2SLS: Predicted EITC Response and Attitude Change

Notes: Data, sample, standard errors described in Figure 7. Average state female education correlated with attitude changes (Panel A) and EITC response (Panel B), but not with pre1975 attitude changes (Panel D). Panel C shows that *predicted* EITC response (from Panel B) is associated with changes in attitudes. In Table A.7, I repeat this analysis using other pre1975 traits, with and without region fixed effects.

For Online Publication

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

Jacob Bastian

Appendix A: Additional Tables and Figures

Table A.1. Treatment Effect Robust to Alternate Definitions of Working

Definition of Working:	Earnings >$0 (2013 $)	Earnings >$1000 (2013 $)	Earnings >$5000 (2013 $)	Work Weeks >0	Work Weeks >25	Labor-Force Partic-ipation	Unemp-loyed
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Mom x Post1975	0.033***	0.030***	0.031***	0.029***	0.028***	0.029***	0.007**
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.003)
Observations	571,170	571,170	571,170	571,170	571,170	571,170	571,170
Mean Dependent Variable:	0.66	0.63	0.54	0.70	0.53	0.62	0.05
Mean Dependent Variable for Mothers in 1975:	0.53	0.50	0.42	0.56	0.41	0.50	0.04

Note: Data source: 1971-1986 March CPS data. Sample includes all women 18 to 51 years old. Binary dependent variable. CPS weights, equation (2), and the set of controls from Table 2 column 4 are used. Column 7 may suggest that the increase in female labor supply outpaced labor demand since the unemployment rate also appears to have increased because of the EITC. Average marginal effects from logit regression are shown. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level.

Table A.2. Results Robust to Alternate Sample Age Ranges

Age Lower Bound (Top Row) and Age Upper Bound (Left Column)	16	18	21	25	30
35	0.046	0.04	0.039	0.03	0.017
	(0.009)	(0.008)	(0.007)	(0.010)	(0.014)
	407,261	361,199	296,011	210,668	108,059
45	0.04	0.035	0.034	0.028	0.023
	(0.009)	(0.008)	(0.007)	(0.007)	(0.007)
	550,904	504,842	439,654	354,311	251,702
55	0.038	0.035	0.037	0.035	0.033
	(0.007)	(0.006)	(0.005)	(0.006)	(0.006)
	683,053	636,991	571,803	486,460	383,851

Note: Data source: 1971-1986 March CPS data. Regression identical to Table 2 column 4 regression except for the sample age range. Results larger for younger mothers but results are consistently positive and statistically significant for various age ranges. Sample used for main analysis is 18-51.

Table A.3. 1980s and 1990s Married Women with Lower-Earning Spouses Responded to the EITC (Similar to Table 3 Columns 2 and 5)

Sample:	1986 EITC Expansion				1993 EITC Expansion			
Dependent Variable:	Employed		Annual Work Hours		Employed		Annual Work Hours	
Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Mom x Post EITC Expansion	0.017	0.056	24.4	130.8	-0.002	0.033	4.6	92.8
	(0.012)	(0.011)	(17.2)	(18.9)	(0.007)	(0.007)	(14.8)	(18.2)
Kid x Post EITC Expansion x		-0.009		-24.5		-0.007		-17.6
Spousal Income (1000s of 2013 $)		(0.002)		(2.830)		(0.001)		(1.8)
Observations	115,194				173,241			
R-squared	--	--	0.135	0.137	--	--	0.095	0.098

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 (pre EITC expansion) and 1989-1991 (post EITC expansion) 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 1989-1992 (pre EITC expansion) and 1993-1996 (post EITC expansion) March CPS data. One difference required for this analysis is these two papers use unmarried women and I use all women. Binary employment is defined as having positive hours of work (to match the definition in these two papers). Annual work hours equals weekly work hours -- which refers to the week prior to the March CPS interview -- times weeks worked last year. Regressions for binary employment reflect average marginal effects from a logit regression and weekly work hours use OLS. In each regression the set of controls is used from Table 2 column 4 and CPS weights are used. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level. Women with missing spousal earnings were dropped.

Table A.4. Additional Heterogeneous Effects of the 1975 EITC on Employment

Subgroup:	Age	Age of Child	Race
Description:	Larger Response Among Younger Mothers	Larger Response Among Moms with Younger Kids	Similar Response for White and Nonwhite Mothers
Variables	(1)	(2)	(3)
Mom x Post1975	0.0418	0.0483	0.036
	(0.0080)	(0.0078)	(0.010)
Mom x Post1975 x Age	-0.0005	-0.0020	
	(0.0002)	(0.0003)	
Mom x Post1975 x White			-0.005
			(0.010)
Observations	571,170	571,170	571,170
Mean Dependent Variable = 0.66			
Mean Dep Var for 1975 Treat. Group = 0.53			

Notes: Data, sample, and approach is identical to Table 2. Each column reflects a separate regression with the full set of controls from Table 2 column 4. Column 1 uses equation (2) and adds *Mom x Post1975 x (Age-16)*. There are at least two reasons to expect younger mothers to be more responsive to the EITC. First, younger women are more flexible, with smaller adjustment costs of choosing to work. Second, since earnings increase with age, younger workers are more likely to earn below the EITC limit and be eligible for EITC benefits. (Although increased earnings over the life cycle is largely attributed to increased experience, among two women with no experience (one younger, one older), a younger woman should still be more likely to respond to the EITC because even if each began earning the same amount, the younger woman could expect a higher return to lifetime earnings. Column 2 uses equation (2) and adds *Mom x Post1975 x (Age of Youngest Kid)*. Whether the EITC had a larger effect on mothers with younger or older children is not obvious. Mothers with very young children had lower employment rates than mothers with older children and therefore had more room for growth, however, the opportunity and childcare costs associated with working were higher for mothers with very young children. The treatment effect was 4.8 percentage points for mothers with newborn infants and this effect decreased by 0.2 percentage points for every year older her youngest child was. This result suggests that the EITC may help explain why the U.S. has long had such a high number of new mothers that work despite few childcare subsidies or parental-leave policies. Column 3 uses equation (2) and adds *Mom x Post1975 x White*. Whether white or nonwhite women were more affected by the EITC is not theoretically straightforward. Two reasons to suspect that nonwhite mothers were less affected by the EITC are that nonwhite mothers were more likely to already be working before 1975 (55 percent compared to 49 percent) and more likely to have non-labor welfare income (16 percent compared to 4 percent). However, reasons to suspect that nonwhite mothers were more affected by the EITC are that nonwhite mothers had lower household earnings before 1975 (both unconditional and conditional on working or being married), were less likely to be married, and were more likely to be mothers --- making it more likely that they met both the income and children requirements of the EITC. White and nonwhite mothers had statistically identical responses to the EITC of about 3.6 percentage points. This result may reflect the context of the 1970s and not generalize to other EITC expansions.

Table A.5. General Social Survey Data Summary Statistics (Individual and State Level)

Level of Observations:	1972-1975 Years Pooled				1976-1986 Years Pooled			
	Individual-Level		Aggregated to the State-Level		Individual-Level		Aggregated to the State-Level	
	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Average Age	37.23	12.33	36.91	1.74	37.26	12.11	37.23	1.20
Average Education	12.31	2.78	12.26	0.58	12.57	2.78	12.56	0.55
Fraction Married	0.75	0.43	0.73	0.07	0.67	0.47	0.66	0.05
Fraction Nonwhite	0.06	0.24	0.12	0.09	0.15	0.36	0.17	0.10
Employment Rate	0.68	0.47	0.68	0.06	0.75	0.43	0.75	0.05
Average Earned Income (2013 $)	19126	23466	18583	2869	20344	24606	20255	3501
Fraction Female	0.54	0.03	0.54	0.03	0.57	0.02	0.57	0.02
Average Gender-Equality Attitudes	0.76	0.43	0.76	0.08	0.81	0.39	0.81	0.05
Fraction of Women Single Moms	0.12	0.33	0.26	0.09	0.19	0.39	0.34	0.05
Fraction Democrat	0.54	0.50	0.57	0.09	0.53	0.50	0.54	0.07
Average Racial-Equality Attitudes	0.85	0.35	0.80	0.10	0.87	0.32	0.85	0.06
Fraction Religion Important	0.45	0.50	0.45	0.11	0.44	0.50	0.43	0.07
Preference for Less Welfare	0.04	1.01	-0.03	0.19	0.15	1.00	0.11	0.13
Fraction with a Working Mother at 16	0.67	0.47	0.33	0.06	0.71	0.46	0.27	0.04
Average Education of Mother	11.16	1.63	10.23	0.47	11.43	1.73	10.42	0.37
Individuals Observed	1	0	107.6	69.5	1	0	330.6	213.4
Observations	2092		32		6621		32	

Notes: 1972-1985 restricted GSS data with state-level identifiers. State-level summary statistics are created from 19,262 individual observations and weighted using GSS weight *wtssall*. Summary statistics shown are weighted by state population. State-level averages created by averaging individuals observed in 1972, 1974, and 1975 and individuals observed in 1977, 1978, 1982, 1983, 1985, and 1986. These are the years the GSS provides information on gender-role attitudes before 1986. Age, education, married, nonwhite, employment, earned income, gender-equality attitudes, democrat, racial-equality attitudes, religion, want less welfare, had working mother, education of mother are averaged over men and women age 18 to 60. Fraction of women single moms is the number of working moms divided by the number of women in each state. Democrat defined as 1 if having a political party identification as strong democrat, not-strong democrat, independent near democrat, and a 0 if strong republican, not-strong republican, independent near republican, independent, or other party. Racial-equality attitudes defined as would vote for a black president. Religion important is a 1 if strength of religious affiliation is strong or somewhat strong and is a 0 if not very strong or no religion. Want less welfare is constructed from a variable asking if there is too much, too little, or just about right amount of welfare; answers are standardized at 1974 levels and higher values indicate a belief that welfare is too high. GSS only surveyed 33 states until 1977, 34 states from 1979-1982, 36 in 1983, and 40 from 1985-1986. To be consistent I only keep states observed in each year. One state (West Virginia) is dropped because there are few observations in the GSS and CPS and the state EITC response is an outlier (-10 percentage points). Results are similar if this state is included.

Table A.6. Individual Traits Correlated with Gender- and Racial-Equality Preferences

Panel A: Gender-Equality Preferences

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Age / 10	-0.042								-0.035
	(0.004)								(0.004)
Year / 10		0.089							0.078
		(0.015)							(0.013)
Years of Education			0.033						0.03
			(0.002)						(0.002)
Married				-0.018					0.013
				(0.010)					(0.010)
Female					0.015				0.022
					(0.008)				(0.008)
Non-White						-0.057			-0.047
						(0.018)			(0.017)
Mother Worked							0.014		
							(0.018)		
Racial-Equality Attitudes								0.077	0.044
								(0.017)	(0.016)
Observations	8,713	8,713	8,713	8,713	8,713	8,713	3,624	8,713	8,713
R-squared	0.041	0.021	0.075	0.026	0.026	0.027	0.023	0.029	0.085

Panel B: Placebo Outcome: Racial-Equality Preferences

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Age / 10	-0.016								-0.01
	(0.004)								(0.004)
Year / 10		0.039							0.014
		(0.012)							(0.010)
Years of Education			0.014						0.014
			(0.002)						(0.002)
Married				-0.017					0.005
				(0.008)					(0.008)
Female					0.027				0.028
					(0.008)				(0.008)
Non-White						0.124			0.141
						(0.025)			(0.026)
Mother Worked							0.031		
							(0.014)		
Gender-Equality Attitudes								0.052	0.031
								(0.011)	(0.011)
Observations	8,713	8,713	8,713	8,713	8,713	8,713	3,624	8,713	8,713
R-squared	0.034	0.027	0.044	0.031	0.032	0.045	0.038	0.035	0.063

Notes: 1972-1985 restricted GSS data with state-level identifiers. State FE in each regression. Year FE in each regression except column 2, where it is controlled for linearly. Samples consist of adults ages 18 to 60 with non-missing data on gender-equality attitudes, state, age, year, education, married, gender, race, earnings, and racial attitudes. Regressions use GSS weight *wtssall*. Gender-equality preferences constructed from the GSS variable *fework*, which asks respondents whether married women should work; positive values represent egalitarian attitudes. Racial-equality preferences comes from the GSS variable *racpres*, which asks respondents whether they would vote for a black president. Whether mother worked is often not available. Heteroskedasticity-robust standard errors clustered at the state level in parentheses.

Table A.7. EITC and Gender-Equality Preferences: Robust to Alternate Measures of State EITC Response

Equation (4) Specification Used To Measure State-Level EITC Response	OLS			Logit			Probit	
Post1975 Extends Until:	1979	1982	1985	1979	1982	1985	1979	1982
Panel A: Using Full Set of Controls from Table 2 Column 4 to Estimate Equation (4)								
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
EITC-Led Increase in	0.019	0.020	0.016	0.021	0.022	0.017	0.021	0.022
Working Mothers	(0.004)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
(in Percentage Points)								
Observations	32	32	32	32	32	32	32	32
R-squared	0.273	0.291	0.199	--	--	--	--	--
Panel A: Using Full Set of Controls from Table 2 Column 4 and State-by-Year FE to Estimate Equation (4)								
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
EITC-Led Increase in	0.009	0.010	0.008	0.009	--	--	0.009	--
Working Mothers	(0.003)	(0.003)	(0.003)	(0.003)			(0.003)	
(in Percentage Points)								
Observations	32	32	32	32	32	32	32	32
R-squared	0.165	0.192	0.133	--	--	--	--	--

Notes: Equation (4) is estimated using 1971-1986 March CPS data and estimates the state EITC response as a state-level difference in differences. Main analysis uses Panel A column 2 specification. 1972-1985 restricted GSS data with state-level identifiers. Gender-equality preferences constructed from the GSS variable *fework* which asks respondents whether married women should work. GSS sample reflects adults 18 to 60 years old in 32 states. State-level EITC response estimated from equation (3). The outcome variable is constructed by subtracting the pooled 1972-1975 GSS state-average from the 1976-1985 GSS state-average. Regressions are weighted by state population, though unweighted results are similar. Some probit and logit regressions did not converge under default convergence criterion tolerance. Heteroskedasticity-robust standard errors in parentheses.

Table A.8. 2S2SLS: Predicted State EITC Response and Gender-Equality Preferences

Pre1975 State-Level Variable Used to Predict State EITC Response	Female Educ.	Single Mothers	White Single Mothers	Nonwhite Single Mothers	Male Earnings
Panel A: Predicting EITC Response and Change in Preferences					
Variables	(1)	(2)	(3)	(4)	(5)
Predicted State-Level Increase in Maternal Employment due to EITC (in Percentage Points)	0.024	0.050	0.047	0.051	0.043
	(0.012)	(0.033)	(0.021)	(0.022)	(0.025)
Observations	32	32	32	32	32
R-squared	0.111	0.085	0.068	0.188	0.214
Panel B: Predicting EITC Response and Change in Preferences (with Region Fixed Effects)					
Variables	(1)	(2)	(3)	(4)	(5)
Predicted State-Level Increase in Maternal Employment due to EITC (in Percentage Points)	0.035	0.057	0.060	0.055	0.047
	(0.022)	(0.064)	(0.025)	(0.036)	(0.033)
Observations	32	32	32	32	32
R-squared	0.307	0.256	0.302	0.306	0.387

Notes: 1972-1985 restricted GSS data with state-level identifiers. Gender-equality preferences constructed from the GSS variable *fework* which asks respondents whether women should work. GSS sample reflects adults 18 to 60 years old in 32 states. State-level EITC response estimated from equation (4). Implicit behind each estimate is the four-panel scatterplot in Figure 8 and each estimate corresponds to Figure 8 Panel C. To conserve space, I only show the second stage of the 2SLS estimates. Column 1 is the same analysis as Figure 8. Columns 2 measures the fraction of mothers that are single. Columns 3 and 4 measure the fraction of women that are single mothers. Column 5 measures the real average earnings of males. Heteroskedasticity-robust standard errors in parentheses. Regressions weighted by state population, though unweighted results are similar.

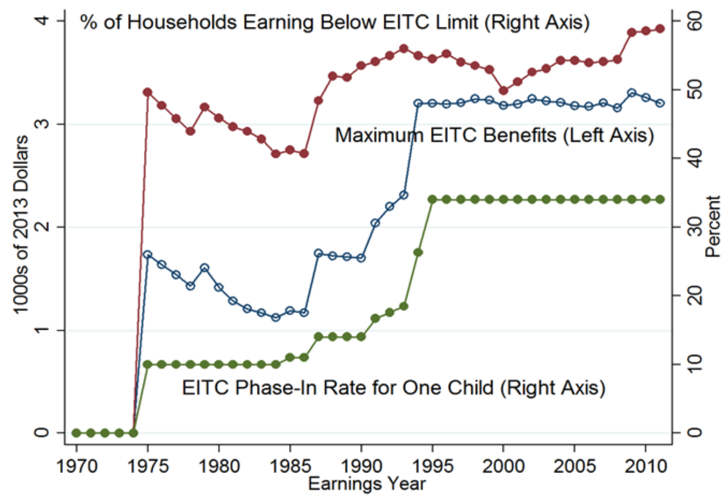

Figure A.1. EITC Trends in Generosity and Eligibility

Notes: Author's calculation from IRS data, March CPS data (using the main sample).

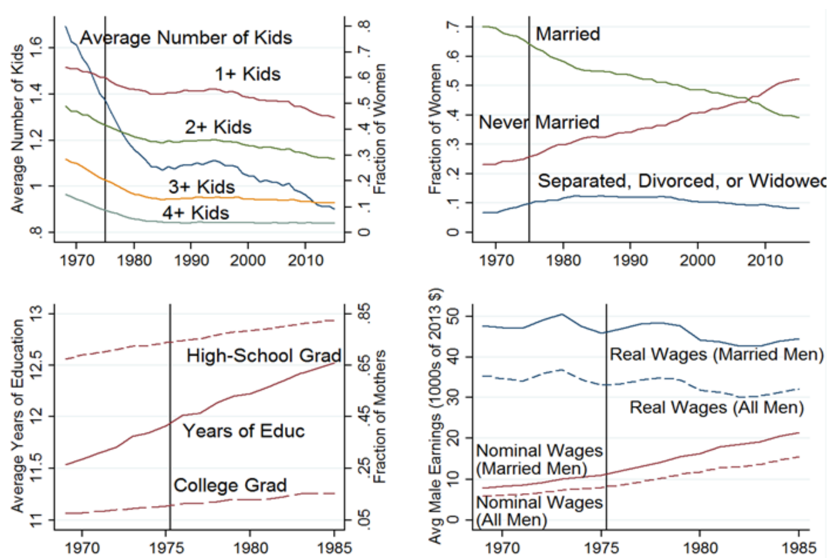

Figure A.2. Ruling Out Confounders: Kids, Marriage, Educ., Male Earn.

Notes: Author's calculation from 1968 to 2015 March CPS (18 to 50 year olds).

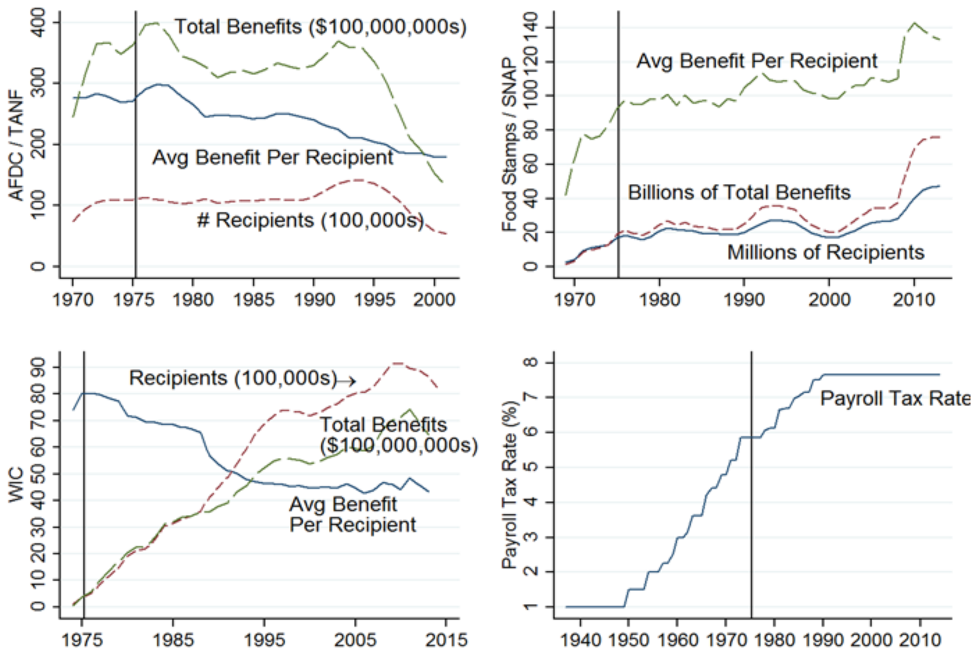

FIGURE A.3. RULING OUT CONFOUNDING POLICIES (TAXES, WIC, AFDC, SNAP)

Notes: Author's calculation from AFDC/TANF data (`https://www.ssa.gov/policy/docs/statcomps/supplement/2005/9g.html#table9.g1`), Food Stamps (SNAP) data (`https://www.fns.usda.gov/pd/supplemental-nutrition-assistance-program-snap`), WIC data (`https://www.fns.usda.gov/pd/wic-program`), and payroll tax data (`http://www.taxpolicycenter.org/statistics/payroll-tax-rates`). Data retrieved 6/25/2017. Food Stamps began rolling out in 1961 and were in all counties by 1975. During the 1970s, families on Food Stamps increased from about 13 to 20 million, which had small negative effects on employment (Hoynes and Schanzenbach 2012). WIC began rolling out in 1972. The percent of counties with WIC rose from 0 in 1973, to 60 in 1975, to 100 in 1979 (Hoynes, Page and Stevens 2011) and had small negative labor-supply effects (Fraker and Moffitt 1988, Hagstrom 1996, Keane and Moffitt 1998, Currie 2003).

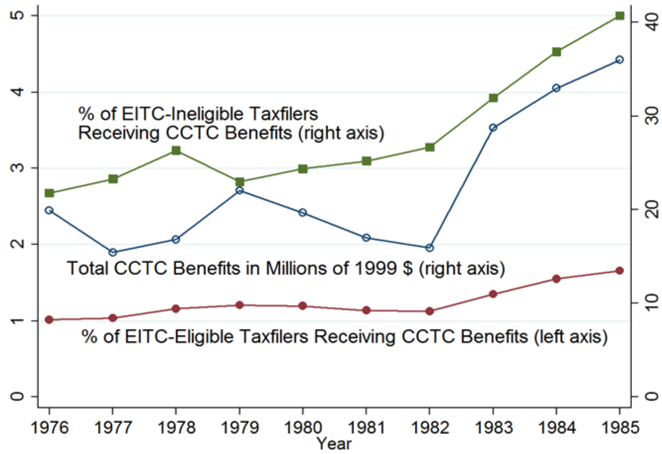

FIGURE A.4. THE 1976 CDCTC AFFECTED FEW EITC-ELIGIBLE TAX FILERS

Notes: 1976-1985 IRS SOI. Sample restricted to tax filers with earned income or business income. EITC eligibility imputed to tax filers with dependents (kids not available in all years) and earnings below the annual EITC income limit. This is imperfect since dependents do not necessarily denote children and I am not able to observe whether tax filers actually claimed the EITC. See Appendix E for SOI sample and variable details.

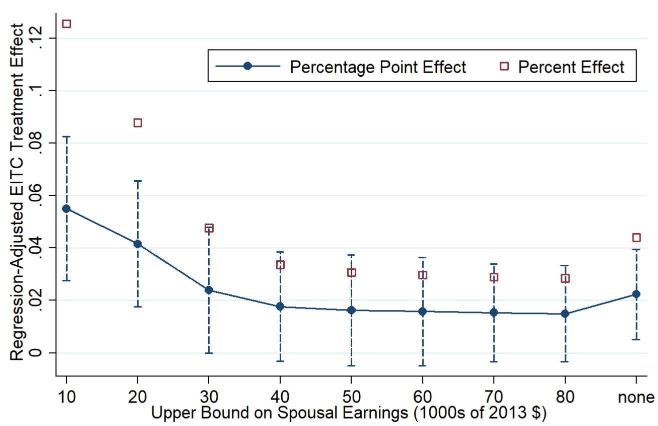

FIGURE A.5. EITC RESPONSE NEGATIVELY CORRELATED WITH SPOUSAL EARNINGS

Notes: 1971-1986 March CPS data. Estimates are from separate logit regressions that use CPS weights, the full set of controls from Table 2 column 4, and the sample of married women with spouses earning below each specified amount. Treatment effects are estimates of $Mom \times Post$ in equation (2). The mean dependent variable for these regressions are 0.51, 0.55, 0.59, 0.61, 0.63, 0.63, 0.63, 0.62, and 0.60, which is why percent-effects are even higher for mothers with low-earning spouses. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level.

Figure A.6. Effect of the EITC on the Distribution of Annual Work Hours

Notes: Same data, sample, and approach as Figure 5. Each estimate is from a different logit regression of having annual work hours in the specified range. The mean dependent variable for the nine unconditional-on-working regressions: 0.35, 0.08, 0.08, 0.08, 0.12, 0.20, and 0.09, and for the seven conditional-on-working regressions are 0.17, 0.10, 0.10, 0.10, 0.16, 0.26, and 0.11. Sample sizes are 236,814 and 176,858.

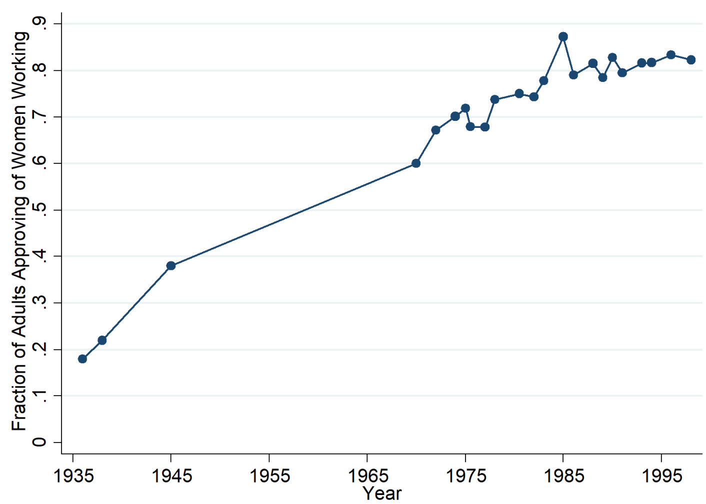

Figure A.7. Gender-Equality Preferences Increasing Over Time

Notes: Attitudes constructed from the binary survey question, "Do you approve or disapprove of a married woman earning money in business or industry if she has a husband capable of supporting her?" Data sources: 1972-1998 GSS data and datasets from the Roper Center (details in Appendix E). GSS weights used to construct annual averages. Other datasets are unweighted (as most do not have weights). Male and female adults.

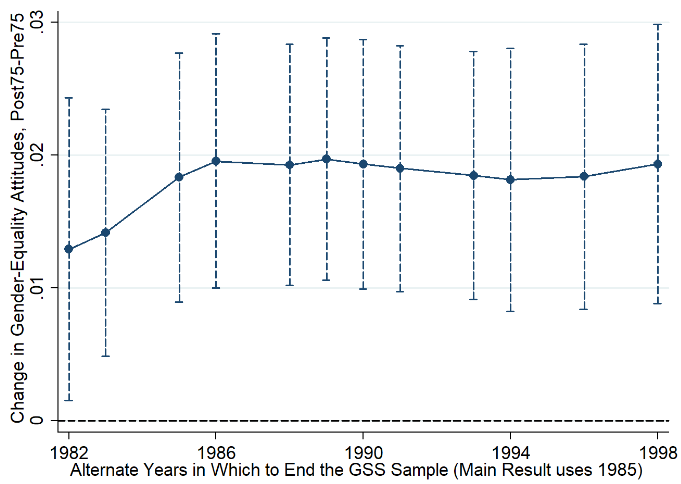

FIGURE A.8. RESULTS ROBUST TO GSS SAMPLE YEARS

Notes: Data and approach resembles Figure 7, except instead of ending the sample in 1985, the post1975 years extend until the specified x-axis year.

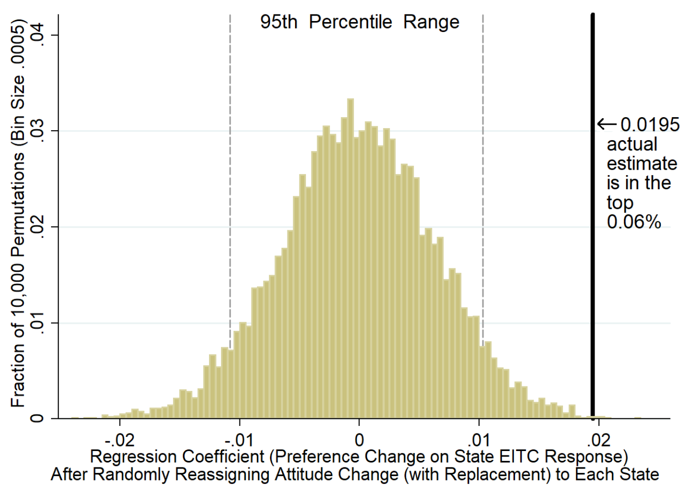

FIGURE A.9. PERMUTATION TEST: RANDOMLY REASSIGN STATE-ATTITUDE CHANGE

Notes: I randomly reassign (with replacement) state attitude changes and re-regress equation (5). 10,000 iterations. Similar to modified Fisher permutation in Buchmueller, DiNardo and Valletta (2011). The actual estimate in Table 7 column 1 is 0.0195 and is in the top 0.06 percent of these permutations, and thus unlikely to be due to chance.

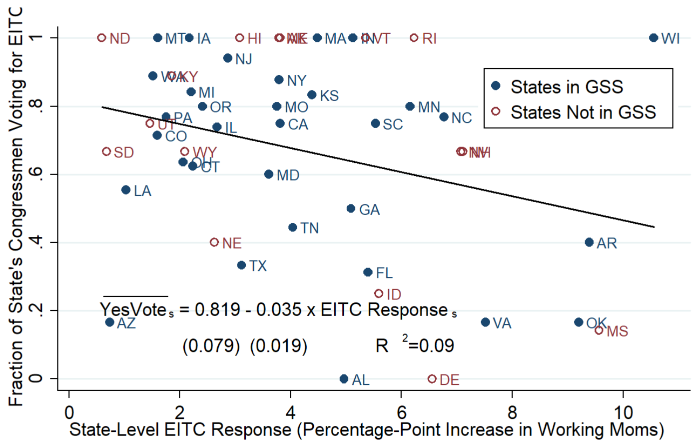

Figure A.10. State EITC Response Neg. Corr. with Voting for EITC Policy

Notes: Congressmen include House of Representatives voting (`https://www.govtrack.us/congress/votes/94-1975/h67`) and Senate voting (`https://www.govtrack.us/congress/votes/94-1975/s112`). State EITC response comes from equation (4). GSS did not interview all 50 states during the 1970s and 1980s. Of course, the Tax Reduction Act of 1975 contained a number of other spending and tax provisions (full bill text: `https://www.gpo.gov/fdsys/pkg/STATUTE-89/pdf/STATUTE-89-Pg26.pdf`).

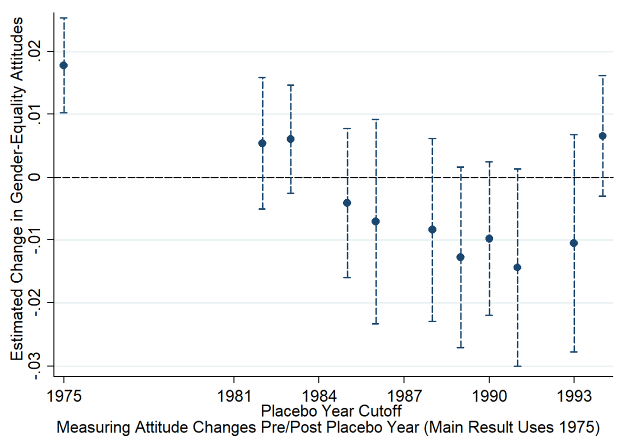

Figure A.11. EITC and (No) Attitude Changes Using Placebo Years

Notes: Data and approach resembles Figure 7, except instead of 1975, I measure the state-level change in attitudes before and after each placebo year. Four years before placebo year and six years after placebo year are used. The identical GSS question about approving of working women is available between 1972 and 1998.

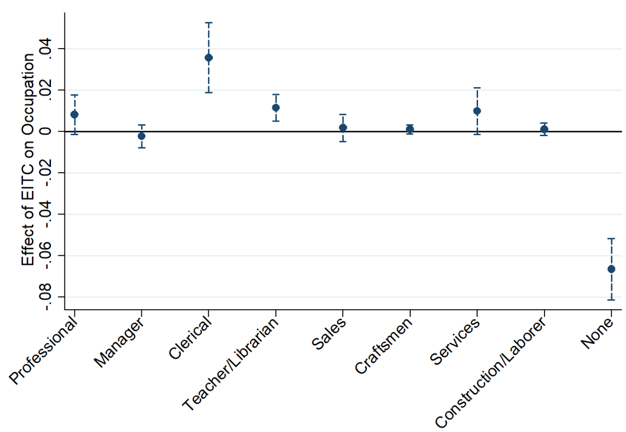

Figure A.12. Which Occupations Did Newly Working Mothers Enter Into?

Notes: 1971-1986 March CPS data. Professions are defined by *occ*1950 codes: professional 0-99, manager 200-290, clerical 300-390, sales 400-490, craftsmen 500-595, services 700-790; and *occ*1990 codes: teacher/librarian 155-165, construction/laborers 558-599, and none 999. Full set of controls and "high-impact" sample used. Each estimate is from a different logit regression of having the specified occupation. Mean dependent variables are 0.13, 0.05, 0.27, 0.04, 0.05, 0.01, 0.16, 0.01, 0.22. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level.

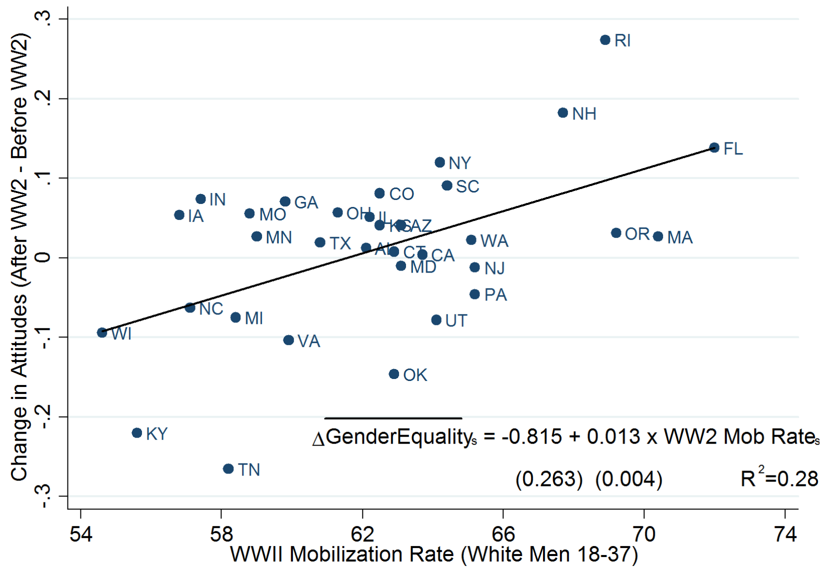

Figure A.13. WWII Working Women Led to Changes in Gender Attitudes

Notes: Mobilization rates from Goldin and Olivetti (2013, Table A1). Attitude data source: Roper Center (`http://ropercenter.cornell.edu/CFIDE/cf/action/ipoll/index.cfm`) and Berinsky and Schickler (2011). The following Gallup datasets and survey questions used: Gallup (1937*c*), Gallup (1937*a*), and Gallup (1937*b*): "Are you in favor of permitting women to serve as jurors in this state?" (Gallup 1937*b*): "Would you vote for a woman for President if she was qualified in every other respect?" (Gallup 1938): "Do you approve of a married woman earning money in business or industry if she has a husband capable of supporting her?" (Gallup 1939): "A bill was introduced in the Illinois State Legislature prohibiting married women from working in business or industry if their husbands earn more than $1,600 a year ($133 a month). Would you favor such a law in this state?" (Gallup 1945): "If the party you most often support nominated a woman for Governor of this state, would you vote for her if she seemed qualified for the job?", "If the party whose candidate you most often support nominated a woman for President of the United States, would you vote for her if she seemed best qualified for the job?", "Would you approve or disapprove of having a capable woman in the President's cabinet?", "A woman leader says not enough of the capable women are holding important jobs in the United States government. Do you agree or disagree with this?", "Would you approve or disapprove of having a capable woman on the Supreme Court?" Change in attitudes (After WWII - Before WWII) created by, first, coding each binary response so that 1 represents gender-equality attitudes; second, averaging each survey question at the state-year level, third averaging the five (November) 1945 questions at the state level to create "After WWII" and averaging the six 1937-1939 questions at the state level to create "Before WWII." Unfortunately, it is not possible to compare exact questions immediately before and after WWII but estimates are very similar if any one or two of the survey questions are omitted: point estimates span 0.017 and 0.007, p-values span 0.001 and 0.065 for these 20+ regressions. Estimates are also positive and statistically significant when the attitudes of men and women are analyzed separately: for men 0.0120 (0.0057) and for women 0.0106 (0.0041).

Appendix B: Additional Robustness Checks

1. Model Choice and Sample Period

In Figure B.1, I show that the estimated DD treatment effect is robust to a probit, logit, or OLS model, and when to end the sample after 1975. As would be expected from Figure 1.B, the treatment effect is small if the sample ends soon after 1975, but grows and flattens out as more years after 1975 are included. OLS results are consistently larger.

2. Larger Response from Mothers Eligible for More EITC Benefits

Conditional on year and spousal earnings (if any), I calculate maximum potential EITC benefits in 2013 dollars ($MaxEITC$) and run a regression identical to equation (2) except with the additional variable $Mom \times Post1975 \times MaxEITC$. For mothers with non-earning spouses and unmarried mothers, the value of $MaxEITC$ varied by year and ranged between \$1,100 and \$1,700 since the EITC schedule was not pegged to inflation until 1986; for married mothers with a working spouse earning above the EITC kink point (placebo group from Table 3 column 4), $MaxEITC$ 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. Table B.1 column 1 shows that a \$1,000 (2013 dollars) increase in $MaxEITC$ is associated with a 3.9-percentage-point increase in maternal employment[1] and carries out the placebo test from Table 3 column 4 in a different way: the estimate of $Mom \times Post$ is now statistically insignificant (that is, a mother after 1975 is no more likely to work than before 1975 if she is eligible for zero EITC benefits) and the effect of the EITC is loaded onto $Mom \times Post \times MaxEITC$.

3. Potentially Endogenous Fertility and Group Composition

In addition to using controls, another way to account for endogenous fertility, marital status, and group composition is by reweighting mothers after 1975 to look like mothers before 1975. Although regression controls should

[1] Similar to Hoynes, Miller and Simon (2015) that \$1000 in EITC benefit increased maternal employment by 7.3 percentage points, and Milligan and Stabile (2007) that \$1,000 increase in public benefits increased maternal employment by 4 percentage points.

largely account for the changing composition of mothers over time, reweighting acts as an additional robustness check (DiNardo 2002). I use two sets of weights: one set is constructed from the approach in DiNardo, Fortin and Lemieux (1996) ("DFL" weights) and the other set is inverse propensity weights ("IP" weights). To construct these weights, I first use a logit[2] and a parsimonious set of traits – six age bins, three education bins, state, and dummies for married, nonwhite, and mother – to estimate the probability than each observation in the sample is from a year before 1975.[3]

$$\text{(B1)} \quad P(Pre75) = f(\beta_1 Age + \beta_2 Ed + \beta_3 St + \beta_4 Marr + \beta_5 Race + \beta_6 Mom + \epsilon)$$

Each observation is assigned a probability p of being from a year before 1975; I create DFL and IP weights by assigning each observation a weight of $p/(1-p)$ and $1/p$.[4] Women are weighted less if their observed characteristics are less likely to be from a year before 1975 and weighed more if their characteristics are more likely to be from a year before 1975 (e.g. low education or high fertility). Figure B.2 verifies that the characteristics of women before and after 1975 overlap sufficiently and have common support (Busso, DiNardo and McCrary 2014). Re-estimating equation (2) with these new weights yields estimates of 3.4 and 3.2 percentage points (Table B.1 columns 2 and 3), similar to the baseline estimate of 3.3.

4. March CPS Imputations

In 1975 the Census changed its hot deck procedure[5] for imputing missing earnings (Welch 1979, Bound and Freeman 1992)[6] and could affect the re-

[2]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). Probit and logit produce very similar results.

[3]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.

[4]Weights are multiplied with the CPS sample weights and normalized to add up to 1 (DiNardo 2002).

[5]Where 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. Source IPUMS: `https://usa.ipums.org/usa/voliii/80editall.shtml#note1`.

[6]Welch (1979): "The imputation procedure used in the first eight surveys differs from that of the ninth so that summary statistics for the 1976 survey (i.e., for 1975 earnings) are not comparable to other years" and "individual records for the first eight surveys

sults in Tables 2, 3, and B1 since I define employment as having positive earnings (although Table A.1 shows similar estimates for other binary definitions of working). The percentage of observations with imputed earnings in the sample is zero before 1975, but between 1975 and 1985 is 13.0, 11.1, 12.7, 14.0, 12.6, 13.1, 10.1, 10.1, 10.5, 11.8, and 10.8. In Table B.2 Column 1 shows the baseline DD estimate using the default CPS imputation and column 2 simply drops all imputed observations. In columns 3 and 4 I use equation (B1) and a logit to predict the probability than an observation has missing earnings data (to account for data missing not at random), create DFL and IP weights (in the way described in the previous section), and re-estimate equation (2) with these weights. Columns 5 and 6 reflect estimates from a bounding exercise – similar to Manski bounds (Manski 1990) – where I assign all observations with missing earnings data to be working or not working. Across each regression, the DD estimate is stable between 3.2 and 3.9 percentage points, similar to the baseline estimate of 3.3.

5. *Additional Response from Women with Multiple Children*

Since the EITC did not provide additional benefits for having more than one child until 1991, mothers with multiple children should not have responded to the EITC more than women with only one child. I test this with the following logit model that expands equation (2) and accounts for any differential impact on employment from having at least J kids.
(B2)

$$P(E) = f(\beta_1 Post1975 + \sum_{k=1}^{J} [\beta_{2k} Mom^k + \beta_{3k} Mom^k \times Post1975] + \beta_4 X + \epsilon)$$

Table B.3 columns 1 to 3 show results of this regression for $J = 1, 2, 3$. Column 1 replicates the baseline estimate where $J = 1$, but surprisingly, in columns 2 and 3 where $J = 2$ and $J = 3$, results show that the estimate of $\beta_{3,k=2}$ is positive and significant. This means that women with at least two kids were more likely to respond to the EITC than women with exactly one child. (Column 3 shows that mothers with at least three children do

contain no flag to identify cases when earnings are imputed. Family records do however identify imputation of total family earnings, which presumably means that earnings for at least one family member are imputed. In contrast, the 1976 survey contains flags for individual amputations but none for families." This issue does not present a problem for my analysis since I focus on the extensive margin, and since I show that results are robust to other definitions of working based on earnings, weeks worked, or labor-force participation (Table A.1).

respond less than women with exactly two children.) Interestingly, Eissa and Liebman (1996) also find an additional response from women with at least two children. They suggest that this may be due to the concurrent increase in the tax exemption for each dependent, which benefited families with multiple children more. During my sample period, the tax exemption for each child also increased from $750 to $1,000 in 1979. However, when I restrict the sample to years before 1979, I still find a positive estimate on $\beta_{3,k=2}$ (Table B.3 column 4) and conclude that increased exemptions is not driving my results.

Another potential explanation is that mothers with multiple children were more likely to have completed their fertility. If mothers that had completed their fertility were more receptive to working – especially when their children reached school age – then with cross-sectional CPS data there could be a mechanical relationship between having multiple children and EITC response. I test this hypothesis in Table B.3 columns 5 to 9 by restricting the sample of mothers in the treatment group to those with a *youngest* child at least 2, 3, 4, 5, and 6 years old. As this youngest-child age restriction increases, the EITC response from mothers with at least two children (relative to mothers with one child) converges to zero, while the estimated response of mothers with exactly one child ($\hat{\beta}_{3,k=2}$) remains positive and grows from 2.3 to 2.8 percentage points. Mothers with multiple children and a youngest child at least 5 years old are statistically no more likely to respond to the EITC than women with just one child. I find the same pattern for the 1986 and 1993 EITC expansion as well (results omitted) and conclude that the additional employment increase for women with at least two children may be explained by mothers that had completed their fertility. (This may also explain why Eissa and Liebman (1996) find the same pattern.)

6. Using IRS Tax Data

Since the CPS shows that the 1975 EITC had a large effect on the employment of mothers, this should be evident in the IRS Statistics of Income (SOI) data as well, however, a few features of the IRS SOI data make it unattractive for detecting the effects of the 1975 EITC. First, many non-working individuals do not file taxes, so detecting an extensive margin response is not easy. Second, *household* income is reported, so it is not possible to determine whether one or two spouses worked. Third, IRS SOI data include few demographic variables so it is not possible to determine the gender, age, race, or education of the tax filer, whether they have children – dependents are not necessarily children – or child's age.[7]

Constrained by the IRS SOI data, I find evidence that the EITC affected the composition of tax filers. Using 1968 to 1985 IRS SOI data, I show that the fraction of unmarried EITC-eligible tax filers (Table 3 shows that single mothers were relatively more affected by the EITC) increased in the years after 1975 (Figure B.3). The pattern closely resembles Figure 1.B: flat before 1975, a quick rise between 1975 and 1980, and relatively flat again after 1980. Without knowing tax filer gender or whether dependents denote children, this is only suggestive evidence that the EITC affected the employment of single mothers.[8]

To corroborate the effect of the EITC with administrative tax records, I first compare the annual number of EITC-eligible households and the amount of EITC benefits implied by CPS data with aggregate IRS EITC statistics. Figure B.4 shows that the number of EITC-eligible households and aggregate EITC benefits – that I calculate from reported household children and earnings – is nearly identical to the published EITC statistics in 1975. However, in the years after 1975, the CPS undercounts EITC recipients and benefits. The ratio of the CPS numbers to the official IRS numbers drops to about 90 percent by 1978, and continues to fall to 70 percent by the mid-1980s. One reason to expect EITC benefits calculated from the CPS to be lower than the actual benefits is that 20 to 25 percent of EITC claims are paid in error[9] due to unintentional tax filer error, divorced parents each claiming the same child, married couples splitting their qualifying children and filing separately as household heads, or lying about having children. Liebman (2000) finds that 11 to 13 percent of EITC recipients had no children.[10] The growing gap between CPS and IRS data in Figure B.4 suggests that tax filer error may have increased between 1975 and 1985.

[7]Marital status is available. Number of children available after 1977, otherwise only in 1970 and 1975. See `http://users.nber.org/~taxsim/taxsim-ndx.txt` for annual available IRS SOI variables.

[8]It is difficult to determine whether the number of tax filers increased, since one million working mothers over a four year period (Figures 1.A and 1.B) corresponds to about 250,000 mothers per year, small in comparison to the 80 million households, 100 million adults in the labor force, and 95 million tax filers in the U.S. by 1980 (source: CPS, BLS, IRS SOI). As a result, I am not able to detect an aggregate rise in tax filers or in the number of working households using IRS SOI or CPS data. Time-series analysis of these data would not detect a newly-working mother that was already a part of a tax-filing household.

[9]See `https://www.eitc.irs.gov/Tax-Preparer-Toolkit/faqs/fraud`.

[10]This is related to the infamous event where millions of children disappeared when taxpayers had to begin reporting the Social Security number of all dependents in 1987 (LaLumia and Sallee 2013).

Observing less (imputed) EITC benefits in the CPS than the aggregate IRS numbers suggest that my employment estimates from the CPS are not overestimates of the actual working response to the EITC. Although Figure C.1 shows evidence of misreporting self-employed income to take advantage of EITC benefits, this represents a relatively small number of the million mothers that begin working in response to the EITC.

Aggregate IRS data also reveal a puzzle in light of estimates in Table 2: the number of EITC recipients and the aggregate EITC benefits remained roughly constant between 1975 and 1985 (Figure B.5). One way to reconcile the positive maternal employment response to the EITC and flat EITC benefits is by considering that the EITC schedule was not pegged to inflation until 1986 and inflation was high in the years after 1975. About 6.3 million households received EITC benefits in 1975, but most recipients seem to have already been working since Figures 1.A and 1.B suggest that the employment was not affected until 1976. Due to rising prices and nominal wages, within a few years some of these households would earn above the nominal EITC earnings limit and no longer receive EITC benefits, akin to "bracket creep" (Saez 2003).[11] The increase in EITC-eligible working mothers (Table 2) and no-longer-EITC-eligible households may have cancelled out and resulted in a roughly constant number of EITC recipients. The following back of the envelope calculation examines whether this is plausible. Using the 1974 SOI earnings distribution (before any labor supply response to the EITC), I use the CPI to inflate the 1974 earnings distribution into 1975, 1976, 1977, and 1978 dollars, and calculate the number of tax filers that were EITC-*eligible* in 1975 but EITC-*ineligible* in 1976, 1977, or 1978 due to rising nominal income.[12] Figure B.6 illustrates that by 1976, 1977, and 1978, 0.6, 1.0, and 1.6 percent of tax filers eligible for the EITC in 1975 would bracket-creep out of EITC eligibility, corresponding to 700,000, 1,200,000, and 1,800,000 tax filers.[13] Even though the stock of EITC recipients remained roughly constant in the decade after 1975, there was substantial flow in and out of EITC eligibility. This may explain why the number of EITC recipients was flat even as a million mothers entered into employment due to the EITC.

[11] This nominal limit was $8,000 through 1978 and $10,000 through 1984.

[12] Assuming constant real earnings. Rising real wages yields even more bracket creep.

[13] Population growth accounts for at most about half a million of these 1.8 million additional EITC recipients: IRS SOI data shows that about a quarter of tax filers with dependents had positive earnings below the EITC limit and CPS data shows that the number of households with children steadily grew from 34.5 million in 1975 to 35.1 million in 1978. Depending on where in the income distribution these new households fell, population growth led to between 200,000 and 600,000 additional EITC recipients.

Tables and Figures for Appendix B

Table B.1. Robustness Checks: *MaxEITC* and Reweighting

	Larger Response from Moms Eligible for More Potential Max EITC Benefits	Reweighting Post1975 Moms to Look Like Pre1975 Moms	
		DFL Weights	IP Weights
Variables	(1)	(2)	(3)
Mom x Post1975	0.005	0.034	0.032
	(0.007)	(0.008)	(0.007)
Mom x Post1975 x MaxEITC	0.039		
	(0.004)		
Observations	571,170	571,170	571,170

Note: Data source: 1971-1986 March CPS. CPS weights used and average marginal effects from logit regression are shown. Reweighting discussed in Appendix B. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level. Each column represents a separate regression with the full set of controls from Table 2 column 4.

Table B.2. Alternate Ways to Treat Imputed CPS Observations

	Baseline: Using CPS Imputations	Drop Imputed Obs.	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 x Post1975	0.033	0.034	0.033	0.032	0.039	0.033
	(0.007)	(0.007)	(0.006)	(0.009)	(0.007)	(0.007)
Observations	571,170	561,402	571,170	571,170	571,170	571,170

Note: Data source: 1971-1986 March CPS. Binary dependent variable employment for positive earnings. CPS weights used. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level. Full set of controls used from Table 2 column 4. CPS imputations discussed in Appendix B.

Table B.3. Completed Fertility May Explain Larger Response from Mothers with Multiple Kids

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 More Likely to Have Completed Fertility				
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 x Post1975	0.033	0.025	0.025	0.012	0.023	0.023	0.024	0.028	0.028
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
Mom of 2+ Kids x Post1975		0.013	0.015	0.014	0.014	0.013	0.012	0.010	0.004
		(0.004)	(0.004)	(0.005)	(0.005)	(0.006)	(0.006)	(0.006)	(0.007)
Mom of 3+ Kids x Post1975			-0.003						
			(0.004)						
Observations	571,170	571,170	571,170	289,943	504,607	476,782	452,817	432,029	412,992

Note: Data source: 1971-1986 March CPS. Binary dependent variable employment. CPS weights used and average marginal effects from logit regression are shown. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level. Each column represents a separate regression with the full set of controls from Table 2 column 4. Each sample includes all women 18 to 51 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 5 to 9 drop mothers in the treatment group with a youngest child below the specified age. See discussion in Appendix B.

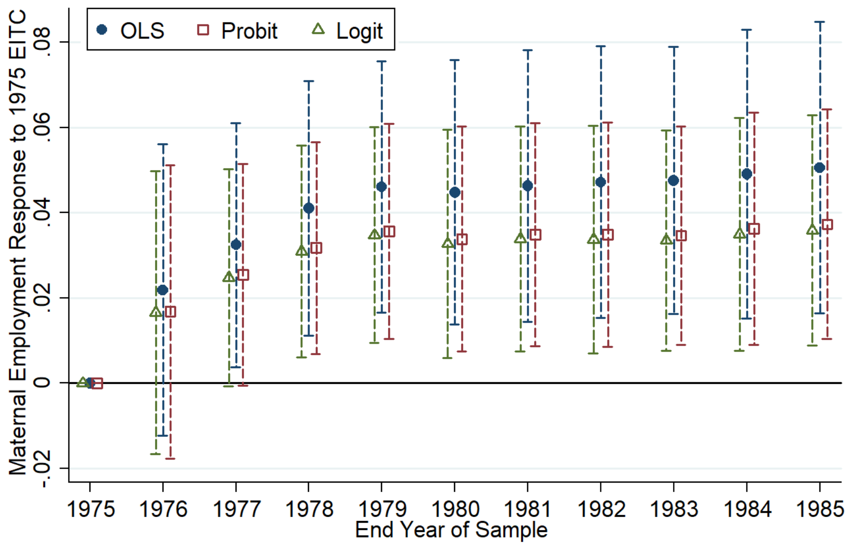

Figure B.1. DD Robust to Model Choice and End of Sample Period

Notes: Data and approach are identical to Table 2 column 4, except that $Post1975$ starts in 1976 and extends through the year specified on the x-axis.

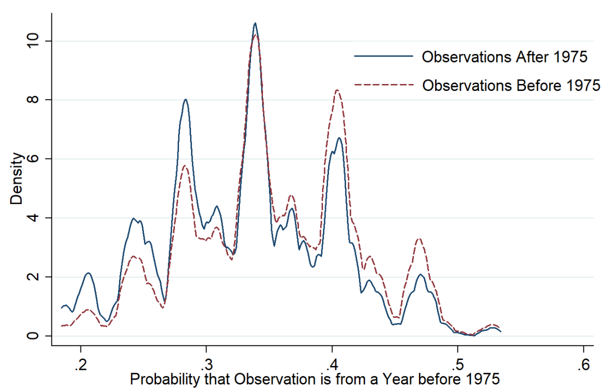

Figure B.2. Kernel Density Plot Shows Common Support for Reweighting

Notes: Data source: 1971-1986 March CPS data. Equation (B1) used and a parsimonious set of controls: six age bins, three education bins, married and nonwhite dummy variables, and 21 state bins. Characteristics of women before and after 1975 overlap and have common support for reweighting (Busso, DiNardo and McCrary 2014).

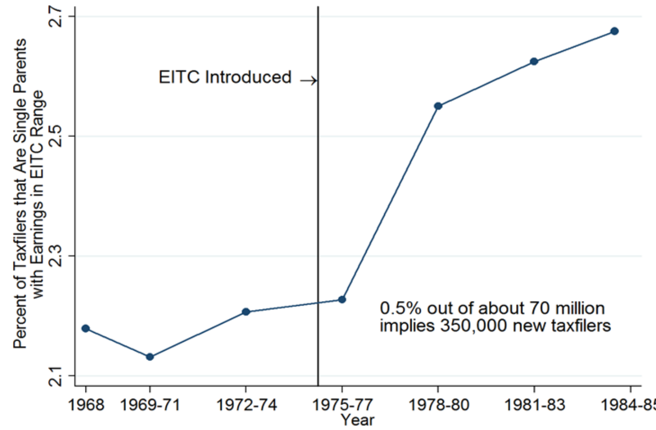

FIGURE B.3. THE 1975 EITC AFFECTED THE COMPOSITION OF TAX FILERS

Notes: Author's calculations from 1968-1985 IRS Statistics of Income Public Use data files. Sample restricted to tax filers with earned income or business income; this eliminates tax filers with only dividend, interest, capital gains, pensions, farm, and alimony income. Refundable portion of the EITC is also provided in the data, but this does not include households who benefit from the EITC through decreased tax liabilities and thus undercounts EITC recipients. Years are grouped into three-year bins to reduce noise.

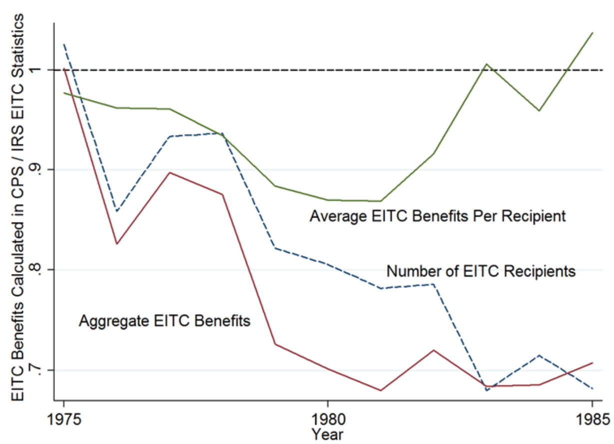

FIGURE B.4. COMPARING EITC RECIPIENTS AND BENEFITS: CPS VS. IRS DATA

Notes: Author's calculation from 1976-1986 March CPS data and published aggregate EITC recipients and benefits (`http://www.taxpolicycenter.org/statistics/eitc-recipients`). EITC recipients and benefits based on household earnings, the annual EITC schedule, and whether the household had any children. The growing gap suggests that tax-filer error may have increased between 1975 and 1985.

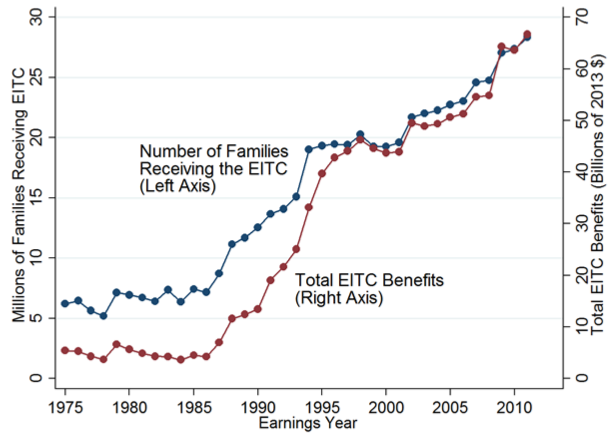

Figure B.5. Trends in EITC Benefits and Recipients

Notes: Author's calculations from IRS data.

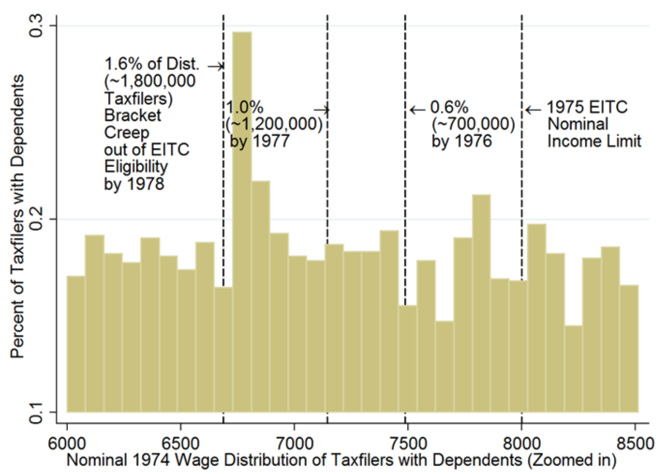

Figure B.6. "Bracket Creep" Reconciles Table 2 and Figure B.5

Notes: Author's calculations from 1974 IRS SOI data and CPI. Sample includes tax filers with wage earnings or business income. EITC schedule not pegged to inflation until 1986; inflation was high in the years after 1975. 6.3 million households received EITC benefits in 1975, but most were already working in 1974. Due to rising prices and nominal wages, within a few years many households would earn above the nominal EITC earnings limit and no longer receive EITC benefits, akin to "bracket creep" (Saez 2003). The rise of working mothers that I find and the no-longer-EITC-eligible households may have cancelled out and resulted in a roughly constant number of EITC recipients.

Appendix C: Calculating Elasticities

1. Extensive-Margin Elasticities

I calculate the extensive-margin labor-supply elasticity as described in Chetty et al. (2012, Appendix B). The numerator of the elasticity is calculated as the pre1975-post1975 change in the log employment rate. The denominator of the elasticity is calculated as the pre1975-post1975 change in the log net-of-tax earnings from working. I allow net-of-tax earnings to account for various taxes (EITC, income tax, payroll tax, dependent deduction) and transfers (AFDC, food stamps, WIC), expanding on the approach in Meyer and Rosenbaum (2001). I calculate net total income for a representative unmarried mother of one child that earns pre-tax $4,427 (in real 1975 dollars)[14] by adding the annual after-tax earnings and public-assistance transfers available to her. I then also add up the transfers available to her if she does not work. If she does not work she is eligible for more transfers, but since the EITC made work more lucrative after 1975, this encouraged many mothers to work and receive less public assistance. The difference in the net total income – from working or not working – measures the financial return to working compared to not working and consists of six pieces: the pre1975 and post1975 after-tax value of $4,427 pre-tax earnings (in real 1975 dollars), the pre1975 and post1975 public assistance available to her if she works, and the pre1975 and post1975 public assistance available to her if she does not work. This is shown in the following equation.

(C1)

$$\epsilon = \frac{log(Emp_{post75}) - log(Emp_{pre75})}{[(log(Earn_{post75} + T^{w}_{post75}) - (T^{\sim w}_{post75})] - [(log(Earn_{pre75} + T^{w}_{pre75}) - (T^{\sim w}_{pre75})]}$$

Values for the numerator can be found in Table 3. Although no estimate in these tables perfectly align with the representative mother described above, two close (and overlapping) matches are unmarried mothers from Table 3 column 1, which experienced a 6.9 percentage point (or 10.7 percent, from a base of 64 percent) increase, and the "high-impact" sample of mothers from Table 3 column 7, which experienced a 5.1-percentage-point (or 8.9-percent, from a base of 58 percent) increase in employment.

In the denominator, *Earn* denotes the real after-tax earnings (in 1975 dol-

[14] This is the average annual earnings of such mothers in the sample. This amount also happens to render her eligible for close to the maximum possible EITC benefits during the sample period (see Figure 4.B). This representative mother is used as an illustration; another type of mother would yield different numerators and denominators of the elasticity calculation.

lars) for the representative mother earning \$4,437 (in constant 1975 dollars) and accounts for the income tax, payroll tax, and dependent deduction. T^w and $T^{\sim w}$ denote the public assistance available to her if she works or does not work.

I transparently show my elasticity calculation in Table C.1 Panel A, which shows the 1970-1985 annual values of the EITC, income tax, payroll tax, dependent exemption, as well as AFDC, Food Stamps, and WIC benefits available if she works and if she does not work. However, calculating public assistance is not straightforward: benefit levels varied by state and did not phase out linearly with earnings. To overcome this, I first calculate AFDC benefits available to this mother if she does not work, using two different sources: one source is the Department of Health and Human Services (DHHS), showing the average benefits for a recipient family; a second source is Fraker, Moffitt and Wolf (1985), which also calculates average family benefits. These two data sources align quite well, and I assign the average of these two values for the case where the mother does not work (Table C.1 Panel A column 17). Second, to calculate the AFDC benefits available to her if she does work, I use the average annual effective tax rates estimated by Fraker, Moffitt and Wolf (1985). This tax rate varies by year, but on average, every dollar of earnings leads to a 25-cent decline in AFDC benefits.[15] Annual AFDC benefits available to this working mother is shown in Table C.1 Panel A column 14. Table C.2 shows details on the DHHS and Fraker, Moffitt and Wolf (1985) data, effective tax rates, and my calculations used to generate the data in Table C.1 Panel A columns 14 and 17.

In Table C.3 I show my calculations for Food Stamps and WIC. It turns out that this representative mother is not eligible for Food Stamps if she works because her earnings are too high. Although the effective tax rate on these benefits are also approximately 25 percent (see Table C.3 notes), Food Stamp benefits are much lower than AFDC benefits. It also turns out that she is eligible for the same amount of WIC whether she works or not, since WIC does not phase out with income and is available for mothers earning below 185 percent of the poverty line (the 1975 poverty line was \$6771, in nominal dollars, for a mother with one child). Table C.3 shows complete details on the Food Stamps and WIC benefits, which are then inputted into Table C.1 Panel A columns 12, 13, 15, and 16.

In Table C.1 Panel B, I aggregate the annual after-tax earnings and transfers (available if she works or not) from Table C.1 Panel A, average them for 1970-1974 and 1975-1985, and plug them into equation (C1) to calcu-

[15]EITC benefits do not count against AFDC eligibility limits (Moffitt 2003).

late the extensive-margin labor-supply elasticity. I find an elasticity of 0.49 using the employment estimate for unmarried women in Table 3 column 1 and 0.41 using the high-impact sample (column 7).[16] As for elasticity standard errors, I assume that the denominator in equation (C1) is measured without error, and approximate the standard error in the numerator by assuming that the T-statistic for the elasticity is identical to the T-statistic for the employment estimates (3.76). This yields standard errors of 0.13 and 0.11 for the elasticity estimates of 0.49 and 0.41. As I explain in the notes to Table C.1, if I account for the various take-up rates of public-assistance programs, this leads to slightly larger elasticities of 0.54 and 0.45. Using the high-impact sample, I also estimate the total intensive plus extensive margin elasticity from the annual work hours and annual earnings estimates in Table 4, I find elasticity estimates of 0.37 (0.10) and 0.47 (0.125).

2. Elasticities from Bunching of Self-Employed Workers

Following Saez (2010), I also use IRS Statistics of Income Public Use Data (SOI) to look for bunching among self-employed tax filers. Figure C.1 shows bunching at the EITC kink point among EITC-eligible tax filers with positive self-employment income (business schedule C), both for the 1975-1978 EITC schedule and the expanded 1979-1984 EITC schedule. Figure C.1 shows no bunching among EITC-ineligible tax filers (claiming zero children) with positive self-employment income. Figure C.2 shows no bunching among wage earners (with no self-employment income), both for EITC-eligible and EITC-ineligible tax filers. There is only evidence of bunching among EITC-eligible tax filers with positive self-employed income and likely reflects income misreporting since there is no third-party reporting for self-employed workers (LaLumia 2009, Saez 2010, Kuka 2014). Following the approach in Saez (2010), I calculate the implied bunching elasticity and find similar results.[17]

Following Saez (2010) and using quasi-linear and iso-elastic utility function, individuals maximize $u(c,z) = c - \frac{n}{1+\frac{1}{e}} \frac{z}{n}^{1+\frac{1}{e}}$, subject to $c = (1-t)z + R$. Where c is consumption, z is the level of earnings, t is the tax rate, n is an

[16]This is similar to the elasticity that Chetty et al. (2012) find when reexamining Meyer and Rosenbaum (2001).

[17]Using a quasi-linear and iso-elastic utility function, the excess bunching density in the earnings distribution, and bandwidths of $1000, $1500, and $2000, I calculate elasticities of taxable income of 0.23, 0.52, and 0.77 in 1975-1978 and 0.58, 1.28, and 2.22 in 1979-1985. The nominal EITC schedule was slightly modified in 1979. Saez (2010) finds bunching at the first EITC kink point in the late 1980s through the 2000s, but does not investigate the first decade of the EITC; my results corroborate Saez (2010).

ability parameter distributed with density $f(n)$, e is the compensated elasticity, and R is non-labor income. The first order condition is $z = n(1-t)^e$ and the bunching elasticity can be estimated by solving the following for e:

$$B = \frac{z^*}{2}\left[\left(\frac{1-t_0}{1-t_1}\right)^e - 1\right]\left[h(z^*)_- + \frac{h(z^*)_+}{\left(\frac{1-t_0}{1-t_1}\right)^e}\right] \tag{C2}$$

Where z^* is the kink threshold, $\left(\frac{1-t_0}{1-t_1}\right)$ is the net of tax ratio at the kink, $h(z^*)_-$ and $h(z^*)_+$ is the density of the distribution just below and above the kink, and B is the amount of bunching at z^*. For a given empirical distribution $h(z)$ and a choice of bandwidth δ, B is equal to the density of tax filers with income is the range $(z^*-\delta, z^*+\delta)-(z^*-2\delta, z^*-\delta)-(z^*+\delta, z^*+2\delta)$. See Saez (2010) Figure 3 for more details and intuition.

I use this formula, the empirical earnings distribution in the SOI tax files for 1975-1978 and 1979-1984 (see Figure C.1 for nominal EITC schedules), and bandwidths of \$1000, \$1500, and \$2000 to calculate the implied bunching elasticity.

For 1975-1978 and δ=\$1000: z^*=\$4000, $\frac{1-t_0}{1-t_1} = 1.2$, $B = .0114$, $h(z^*)_- = 0.0000582$, and $h(z^*)_+ = 0.0000788$. Yielding e=0.23.

For 1975-1978 and δ=\$1500: z^*=\$4000, $\frac{1-t_0}{1-t_1} = 1.2$, $B = .0173$, $h(z^*)_- = 0.0000388$, and $h(z^*)_+ = 0.0000525$. Yielding e=0.52.

For 1975-1978 and δ=\$2000: z^*=\$4000, $\frac{1-t_0}{1-t_1} = 1.2$, $B = .0191$, $h(z^*)_- = 0.0000291$, and $h(z^*)_+ = 0.0000394$. Yielding e=0.77.

For 1979-1984 and δ=\$1000: z^*=\$5000, $\frac{1-t_0}{1-t_1} = 1.1$, $B = .0204$, $h(z^*)_- = 0.0000642$, and $h(z^*)_+ = 0.0000838$. Yielding e=0.58.

For 1979-1984 and δ=\$1500: z^*=\$5000, $\frac{1-t_0}{1-t_1} = 1.1$, $B = .0299$, $h(z^*)_- = 0.0000428$, and $h(z^*)_+ = 0.0000559$. Yielding e=1.28.

For 1979-1984 and δ=\$2000: z^*=\$5000, $\frac{1-t_0}{1-t_1} = 1.1$, $B = .0388$, $h(z^*)_- = 0.0000321$, and $h(z^*)_+ = 0.0000419$. Yielding e=2.22.

Saez (2010) finds elasticities among self-employed workers in the range of 0.7 to 1.6, depending on bandwidth choice.

Tables and Figures for Appendix C

Table C.1. Calculating Labor Supply Elasticity for a Representative Unmarried Mother of One Kid and Earning $4,427 (1975

Panel A: Nominal Annual Values of Earnings, Taxes, and Transfers Needed to Calculate the Elasticity Denominator

											Transfers					
											Working			Not Working		
Year	CPI	Nominal Earnings	Payroll Tax Rate	Payroll Tax Paid	Dep. Exempt.	Taxable Income	EITC Benefits	Income Taxes Owed	Total Taxes Owed	After-Tax Income	Food Stamps	WIC	AFDC	Food Stamps	WIC	AFDC
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)	(17)
1970	4.54	3195	4.8	153	625	2416	0	515	515	1901	0	0	1432	253	0	2151
1971	4.294	3330	5.2	173	675	2481	0	539	539	1942	0	0	1690	325	0	2222
1972	4.114	3439	5.2	179	750	2510	0	559	559	1951	0	0	1658	324	0	2260
1973	3.986	3654	5.85	214	750	2690	0	598	598	2092	0	0	1605	350	0	2299
1974	3.752	4056	5.85	237	750	3069	0	671	671	2398	0	376	1601	423	376	2473
1975	3.379	4427	5.85	259	750	3418	357	741	384	3034	0	446	1596	514	446	2659
1976	3.097	4683	6.05	283	750	3649	332	790	458	3191	0	470	1707	574	470	2831
1977	2.928	4985	5.85	292	750	3944	301	441	140	3804	0	499	1767	593	499	2963
1978	2.749	5366	6.05	325	750	4291	263	510	246	4045	0	528	1662	642	528	3057
1979	2.555	5970	6.13	366	1000	4604	403	545	142	4462	0	578	1488	734	578	3159
1980	2.295	6776	6.13	415	1000	5360	322	680	357	5003	0	610	1538	827	610	3333
1981	2.022	7479	6.65	497	1000	5982	252	806	554	5427	0	668	1547	948	668	3416
1982	1.833	7939	6.7	532	1000	6407	206	776	570	5837	0	692	1702	940	692	3686
1983	1.726	8190	6.7	549	1000	6642	181	758	577	6065	0	711	1698	1032	711	3745
1984	1.673	8548	7	598	1000	6950	145	770	625	6325	0	734	1684	1026	734	3821
1985	1.603	8850	7.05	624	1000	7226	127	795	669	6557	0	761	1709	1080	761	3921

Table C.1 (continued)

Panel B: After-Tax Earnings + Transfers for a Representative Mother that Either Works or Does Not Work: 1970-74 and 1975-85	1970	1971	1972	1973	1974	1975	1976	1977	1978	1979	1980	1981	1982	1983	1984	1985
$Earn + T^{w}$ (Nominal $):	3334	3632	3609	3697	4376	5076	5368	6070	6235	6527	7151	7642	8231	8474	8743	9027
$T^{\sim w}$ (Nominal $):	2404	2547	2583	2649	3272	3618	3875	4056	4228	4471	4771	5032	5318	5488	5581	5762
$Earn + T^{w}$ (Real 1975 $):	4479	4615	4393	4361	4859	5076	4920	5260	5072	4936	4857	4573	4328	4328	4329	4282
$T^{\sim w}$ (Real 1975 $):	3231	3237	3145	3125	3633	3618	3552	3514	3439	3381	3240	3011	2885	2803	2763	2733

$Earn_{pre75} + T^{w}_{pre75} = 4541$

$Earn_{post75} + T^{w}_{post75} = 4736$

$T^{\sim w}_{pre75} = 3274$

$T^{\sim w}_{post75} = 3176$

$$\epsilon = \frac{\log(Emp_{post75}) - \log(Emp_{pre75})}{[(log(Earn_{post75} + T^{w}_{post75}) - (T^{\sim w}_{post75})] - [(log(Earn_{pre75} + T^{w}_{pre75}) - (T^{\sim w}_{pre75})]} = \frac{0.102}{0.208} = 0.49$$

Notes: Employment estimate based on unmarried women in Table 3 column 1. High-impact sample (Table 3 column 7) yields 0.41 or yields 0.37 and 0.47 when annual hours worked or annual earnings from Table 5 columns 1 and 4 are used. Another approach is to account for take-up rates. According to Currie (2003) reasonable estimates for these are 0.75 for Food Stamps, 0.85 for AFDC (for female headed households), 0.75 for WIC, and 0.85 for EITC. Multiplying these take-up rates with the eligible amounts in Table C.I yields a denominator of 0.189 and an elasticity of 0.54. Nominal earnings in column 3 uses the CPI to put $4427 (1975 dollars) into annual nominal dollars. Annual payrol tax rates from Tax Policy Center (http://www.taxpolicycenter.org/statistics/payroll-tax-rates). Income tax rates from Tax Foundation (https://taxfoundation.org/us-federal-individual-income-tax-rates-history-1913-2013-nominal-and-inflation-adjusted-brackets/). EITC benefits calculated by author from EITC parameters. Dependent exemption data from Tax Policy Center (http://www.taxpolicycenter.org/statistics/historical-individual-income-tax-parameters). For data and details on AFDC, see Table C.II. For data and details on Food Stamps and WIC, see Table C.III. The increase in the return to working largely reflects the 1975 EITC as well as the Revenue Act of 1978 which lowered taxes on low-income earners. Although this tax cut likely increased labor supply, the difference-in-differences empirical strategy in this paper nets out the effect of the Revenue Act of 1978 since this tax cut applied to both women with and without children.

Table C.2. Calculating Annual AFDC Benefits for Representative Mother if She Works or Does Not Work

Year	Calculating AFDC from Data in Fraker et al. (1985) Table 1						Calculating AFDC from DHHS Data			Averaging Two Approaches	
	Average Monthly AFDC for Family with Zero Earnings (Fraker et al. 1985)	1999 CPI	Effective Tax Rate, Averaged Across States (Fraker et al. 1985)	Average Annual AFDC for Non-Working Mother	Nominal Earnings from Table C.I Column 2	Imputed Annual AFDC for Working Mother	Average Monthly AFDC for a Family	Average Annual AFDC for Non-Working Mother	Imputed Annual AFDC for Working Mother	AFDC for Non-Working Mom: Column 5 and 9 Average	AFDC for Working Mom: Column 7 and 10 Average
	1967 Dollars			Nominal Dollars							
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
1967		5.142									
1969	156	4.787	29	2011							
1970		4.54	22.5*	2105*	3195	1386*	183	2198	1479	2151	1432
1971	153	4.294	16	2199	3330	1666	187	2246	1713	2222	1690
1972		4.114	17.5*	2253*	3439	1651*	189	2266	1665	2260	1658
1973	149	3.986	19	2307	3654	1612	191	2291	1597	2299	1605
1974		3.752	21.5*	2495*	4056	1623*	204	2451	1579	2473	1601
1975	147	3.379	24	2684	4427	1622	219	2633	1571	2659	1596
1976		3.097	24*	2828*	4683	1704*	236	2833	1709	2831	1707
1977	141	2.928	24	2971	4985	1775	246	2955	1759	2963	1767
1978		2.749	26*	3068*	5366	1672*	254	3047	1652	3057	1662
1979	131	2.555	28	3164	5970	1492	263	3154	1483	3159	1488
1980		2.295	26.5*	3306*	6776	1510*	280	3360	1565	3333	1538
1981	113	2.022	25	3448	7479	1579	282	3384	1515	3416	1547
1982	111	1.833	25	3737	7939	1752	303	3636	1651	3686	1702
1983		1.726	25*	3737*	8190	1689*	313	3754	1706	3745	1698
1984		1.673	25*	3737*	8548	1600*	325	3905	1768	3821	1684
1985		1.603	25*	3737*	8850	1525*	342	4106	1893	3921	1709

Notes: AFDC data from DHHS (https://www.ssa.gov/policy/docs/statcomps/supplement/2005/9g.html#table9.g1, retrived 6/25/2017) and Fraker et al. (1985) Table 1. Fraker et al. (1985) also source of effective annual tax rate on AFDC from gross earnings, which is used for compute columns 7 and 10. Fraker et al. (1985) Table 1 does not provide estimates for every year, for missing years I impute AFDC benefits and effective tax rates by averaging adjacent years and for 1983-1985 I use 1982 values, these imputations are denoted by *.

Table C.3. Calculating Annual Food Stamps and WIC Benefits for a Representative Mother if She Works or Not

Year	Food Stamps			WIC		
	Average Monthly Food Stamps Per Person	Average Annual Food Stamps for Non-Working Rep. Mother	Average Annual Food Stamps for Working Rep. Mother	Average Monthly WIC Per Person	Average Annual WIC for Non-Working Rep. Mother	Average Annual WIC for Working Rep. Mother
(1)	(2)	(3)	(4)	(5)	(6)	(7)
1970	10.55	253	0	0	0	0
1971	13.55	325	0	0	0	0
1972	13.48	324	0	0	0	0
1973	14.6	350	0	0	0	0
1974	17.61	423	0	15.7	376	376
1975	21.4	514	0	18.6	446	446
1976	23.93	574	0	19.6	470	470
1977	24.71	593	0	20.8	499	499
1978	26.77	642	0	22.0	528	528
1979	30.59	734	0	24.1	578	578
1980	34.47	827	0	25.4	610	610
1981	39.49	948	0	27.8	668	668
1982	39.17	940	0	28.8	692	692
1983	42.98	1032	0	29.6	711	711
1984	42.74	1026	0	30.6	734	734
1985	44.99	1080	0	31.7	761	761

Notes: Food Stamps data from FNS (https://www.fns.usda.gov/pd/supplemental-nutrition-assistance-program-snap, retrived 6/25/2017) and source of effective tax rate on Food Stamps from earnings from CBO (https://www.cbo.gov/publication/43709). WIC data from FNS (https://www.fns.usda.gov/pd/wic-program). WIC eligibility does not decline with income as long as earnings are below 185 percent of the poverty line. For an unmarried female head with one child this was $6771 (nominal dollars) in 1975 (https://www.census.gov/data/tables/time-series/demo/income-poverty/historical-poverty-thresholds.html). Therefore, whether the representative mother works or not, she is WIC-eligible.

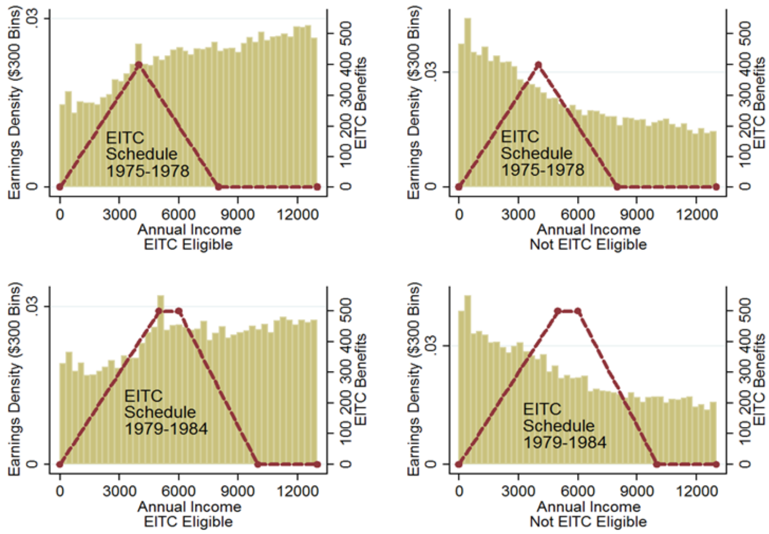

Figure C.1. Bunching Among Self-Employed EITC-Eligible Tax Filers

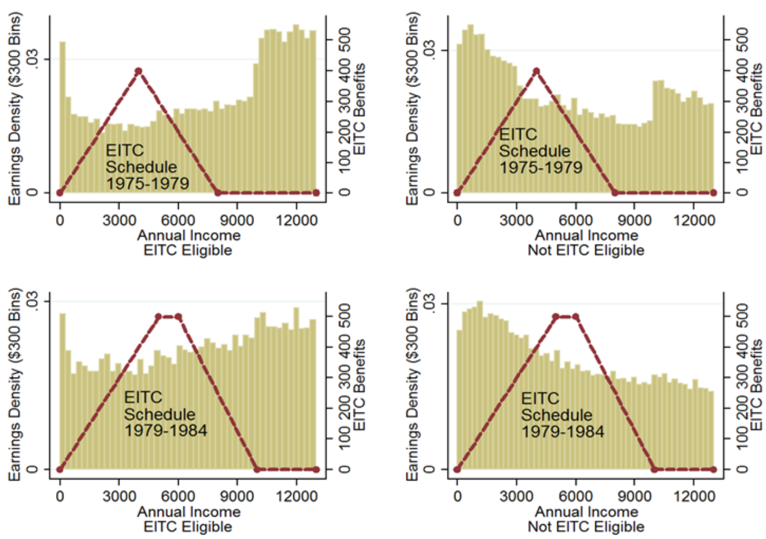

Figure C.2. No Bunching Among Wage Earning EITC-Ineligible Tax Filers

Notes: 1975-1985 IRS Statistics of Income public use files. Sample consists of tax filers with positive self-employment (business schedule C) income. Data on children not available in 1976 so I proxy for EITC-eligible as having at least one dependent.

Appendix D: Less Parametric Approaches

Results in Figure 7 show that each percentage point increase in EITC response led to a 2.0 percentage point increase in positive state gender-equality attitudes. However, if this relationship is not linear – such as with decreasing marginal treatment effects – an OLS specification could be a poor approximation of the true relationship.

One way to test this is to divide up EITC response into a number of categories and regress changes in attitudes on each of these binary categories simultaneously. Results in Figure D.1 show estimates from a regression resembling equation (5), but with three binary variables instead of the continuous variable $EITCResponse_s$. The excluded group represents states with an EITC response less than 2.6 percentage points and the other two categories encompass 2.6-6.7 and 6.7-10.5 percentage points. Figure D.1 shows that state EITC response has an increasingly positive effect on gender-equality attitudes and roughly approximates the predicted effect from a linear OLS specification. This semi-parametric approach shows that OLS closely approximates the effect of the EITC on gender-equality attitudes.

A second approach is to use locally weighted regression (Cleveland 1979). Figure D.2 shows that when the regression behind Figure 7 is locally weighted, the slope is positive and roughly constant, closely resembling a linear OLS estimate.

Figures for Appendix D

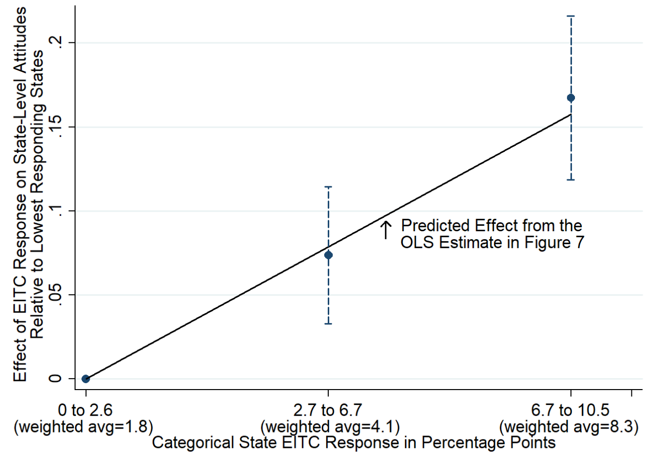

Figure D.1. Categorical EITC Response Corroborates OLS Linear Effect

Notes: Results from regression resembling equation (5) except that $EITCResponse_s$ is replaced with three binary variables for having an EITC response between 0.9 and 2.5, 2.5 and 7.4, or 7.4 and 10. Sample sizes of each group (including the omitted group with $EITCResponse_s$ below 0.9 percentage points) are 5, 7, 15, 5. Heteroskedasticity-robust standard errors are used. Regressions are weighted by state population.

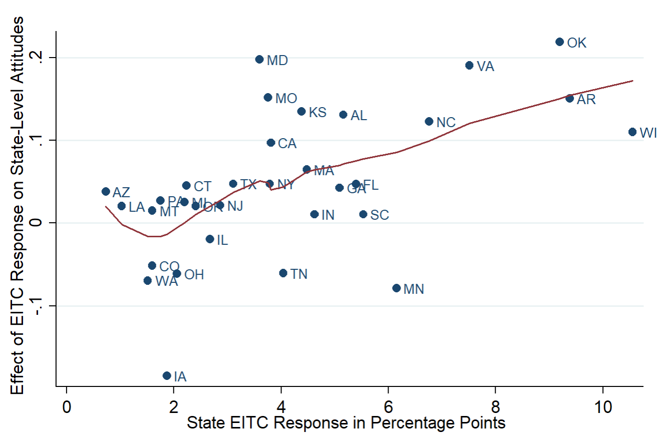

Figure D.2. Locally Weighted Regression

Notes: Locally weighted regression (Cleveland 1979). State EITC response and attitude changes. Stata command *lowess*, default setting: running-line least squares, tricube weighting function, bandwidth 0.8.

Appendix E: Data Appendix

The following information is intended to be detailed enough to replicate my sample.

1. March Current Population Survey Data

I use 1971 to 1986 March CPS (Ruggles et al. 2015) downloaded in December 2014 (2,461,704 observations). I replace *year* with *year*-1 to match the survey year with the work year. I define EITC-eligible households as having at least one child 18 or under, or an adult child between 19 and 23 and in school full time. Households are defined as unique combinations of variables *year* and *serial*. I then drop individuals under 18, observations with a CPS weight (*wtsupp*) of 0, missing education, leaving 1,699,783 observations. Husbands defined as married males. 432,054 individuals live in a household with 0 married males, 1,251,017 individuals live with 1 married male, 16,320 live with two, 384 live with three, and 8 live with four. Each sub-family within a household is assumed to be a separate tax-filing family unit. Dropping women with missing spousal earnings or state, males, and women over 50, yields the 571,170 women used in the main analysis.

The following is a discussion of variables used in employment analysis. Missing *incwage* values of 99999 assigned to be 0 for 574 observations. Weeks worked assigned as the midpoint of the categorical variable *wkswork*. $Post1975$ begins in 1976. Welfare comes from $incwelfr$, married defined as *marstat* equals 1 or 2, and nonwhite created from *race* and *hispan*. Age is rounded to bins of two so that birth year, year, and age can all be controlled for; age squared and cubed are based on actual age. Spousal earnings created from *incwage* and matching a male husband to a female wife; single women assigned zero spousal earnings. States are not identified individually until (working year) 1976. For consistent "states" over time I define 21 "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. National unemployment rates come from BLS: `http://www.bls.gov/cps/cpsaat01.htm`. State-year employment to population ratios created from state-year measures of total employment (found here: `http://www.bea.gov/regional/downloadzip.cfm` under "Local area personal income accounts" file CA25, row 2 in each state file) and state-year measures of population (found at same link under "Local area personal income accounts" file CA25, row 3 in each state file). When state-level measures pertain to these multi-state groups, I weight the variable by annual state population. This data source begins in 1969. Dollars adjusted to real dollars (when specified) using the Consumer Price Index. Occupations detailed in Figure A.12 notes.

2. IRS Statistics of Income Public Use Files

Analysis behind Figures C.1 and C.2 and bunching elasticities calculated in Appendix C use 1975 to 1984 SOI data. Sample restricted to tax filers with positive wages and salaries ($data11$) or positive schedule C business net income ($data17$). EITC-eligible children determined by $data106$, children at home. In 1976 this variable was not available and I instead use $data8$ for number of total dependents. Variable availability in SOI data found here: `http://users.nber.org/~taxsim/taxsim-ndx.txt`.

Analysis behind Figure A.4 use 1976 to 1985 SOI data. EITC-eligible tax filers defined those with wage earnings or business schedule C income below the EITC income limit with a child dependent. Child and Dependent Care Tax Credits given by SOI variable $data64$.

Section .6 uses 1968 to 1985 SOI data. Marital status given by SOI variable $data2$. Number of tax filers in Figure B.3 determined from SOI weight $data1$.

3. General Social Survey Data

I use restricted GSS data with state-level identifiers. Gender-equality attitudes defined from GSS variable $fework$ and racial-equality attitudes from $racpres$. Log income from $conrinc$ and is in real 1000s. Democrat defined as $partyid$ values between 0 and 2, religious defined as $reliten$ values of 1 or 3, too much welfare defined from $natfare$, mom worked and mom education defined from $mawk16$ and $maedyrs$.

In each regression, N=32 since I drop one outlier (West Virginia) that has an EITC response of -10 percentage points and GSS only surveyed 33 states before 1975. Not dropping the outlier has almost no effect on the results. To have a balanced panel and to be consistent over time, I only keep the states that have observations in all years.

Figure A.7 includes adults of all ages (18+) and pools men and women. All other GSS analysis is restricted to adults ages 18-60. This cutoff does not have much of an effect on the results, however when the age cutoff is lowered sufficiently, the sample size and power shrinks, and results become less statistically significant (e.g. age 30 cutoff).

Results define the post1975 period through 1985 and include years 1977, 1978, 1982, 1983, and 1985. The other questions do not have the outcome variable of interest. Results are similar if 1985 (or if 1983 and 1985) is excluded. As would be expected from the employment trends in Figures 1.A and 1.B, the effect on attitudes is larger if 1977 is excluded from the post-1975 period.

4. Gallup Data

Data obtained from Roper Center (`http://ropercenter.cornell.edu/CFIDE/cf/action/ipoll/index.cfm`) and Berinsky and Schickler (2011). The following Gallup datasets and survey questions were used for analysis in Figure A.13. Gallup (1937*c*), Gallup (1937*a*), and Gallup (1937*b*): "Are you in favor of permitting women to serve as jurors in this state?" Gallup (1937*b*): "Would you vote for a woman for President if she was qualified in every other respect?" Gallup (1938): "Do you approve of a married woman earning money in business or industry if she has a husband capable of supporting her?" Gallup (1939): "A bill was introduced in the Illinois State Legislature prohibiting married women from working in business or industry if their husbands earn more than $1,600 a year ($133 a month). Would you favor such a law in this state?" Gallup (1945): "If the party you most often support nominated a woman for Governor of this state, would you vote for her if she seemed qualified for the job?", "If the party whose candidate you most often support nominated a woman for President of the United States, would you vote for her if she seemed best qualified for the job?", "Would you approve or disapprove of having a capable woman in the President's cabinet?", "A woman leader says not enough of the capable women are holding important jobs in the United States government. Do you agree or disagree with this?", "Would you approve or disapprove of having a capable woman on the Supreme Court?"