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 7 percent, representing one million mothers and an elasticity of 0.43. 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, H31, 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). 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) 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 (Figure 1.A). Between 1975 and 1980, the relative employment of mothers rose by about 6 percentage points, closing the employment gap between these two groups by 25 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 nearly identical to the regression-adjusted trend that controls for a rich set of individual- and state-by-year-level covariates (Figure 1.B).

1Cross-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).

2Only 20 percent of married women with infants worked in 1973, compared to 62 percent in 2000 (Goldin 2006).
The EITC also increased labor-force attachment and work intensity, raising average annual work hours by 7.3 percent (43 hours) and earnings by 9.6 percent ($965 in 2013 dollars). Results imply a participation elasticity of 0.43 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 (4.5 and 4.0 percentage points).

My estimates suggest that the 1975 EITC encouraged a 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 af-
ected when a million mothers began working after 1975.\footnote{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 2).}

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 \textit{actual} state EITC response and the \textit{predicted} response (based on preexisting state demographic and occupational 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,\footnote{Subsequent EITC expansions – and their effect on maternal employment – have been studied (see section I).} I find that the EITC encouraged a million mothers to begin working and affected the social attitudes of millions of Americans.
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 and 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

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\(^5\)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 3.A and 3.B). 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 3.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 2017) and employment and earnings (Bastian and Michelmore 2017). 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 3.A shows a budget constraint under the EITC and Figure 3.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. 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. 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. However, the 1975 EITC was large
in at least three ways (Figure A.1): first, about 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 la-
bor 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

9Effects on these margins are generally small: For fertility, Baughman and Dickert-
Conlin (2009) and Bastian (2017) find positive effects. For marriage, Ellwood (2000),
Dickert-Conlin and Houser (2002), Herbst (2011), and Michelmore (2015) find negative
effects, while Bastian (2017) finds positive effects.
10Bastian and Michelmore (2017) is one exception.
11As seen in the following representative quotes: “Between 1975 and the passage of the
Tax Reform Act of 1986, the EITC was small, and the credit amounts did not keep up
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).
II. Conceptual Framework

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

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

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 \(c(\cdot)\) 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.

\[
P(E_{ist}) = f(\beta_1 \text{Mom}_{ist} + \beta_2 \text{Mom} \times \text{Post1975}_{ist} + \beta_3 X_{ist} + \delta_{st} + \epsilon_{ist})
\]

\(^{12}\)I assume working mothers did not displace non-mothers (Neumark and Wascher 2011). However, even if an increase in the supply of 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 Figure 1.A).
$E_{ist}$ is binary for whether a woman is employed. \textsuperscript{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 $\beta_2$ should be positive since the EITC subsidized work. $X_{ist}$ are controls that vary at the individual, state, and year level. $\delta_{st}$ contain state and year fixed effects to control for national trends and state-specific traits associated with female employment. $\epsilon_{ist}$ 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.

\textbf{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 16- to 45-year-old women. The treatment group consists of mothers\textsuperscript{14} and the control group consists of women without children. Table 1 shows summary statistics for all 550,904 women in column 1, while columns 2 and 3 split the sample into treatment and control groups. Women in the sample average 29 years old with 12.1 years of education, 11 and 9 percent are Black and Hispanic, 65 percent work, average individual annual earnings are $12,826 ($19,685 conditional on working), average household earnings are $40,857 (2013 dollars), and 46 percent have household earnings below the EITC limit. On average, mothers are older, less likely to be white, less likely to work, and have less education and higher

\textsuperscript{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).

\textsuperscript{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.
household earnings. See Appendix E for data and sample details.

Figures 1.A and 1.B 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. Although employment levels differed for these groups, employment trends were parallel before 1975 (p-value 0.56).

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 (Figure 1.A).\textsuperscript{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).\textsuperscript{16} Also, trends in marriage, fertility, education, and male earnings were smooth (Figure A.2).\textsuperscript{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): only 1 percent of EITC-eligible tax filers received any CDCTC benefits, compared to 30 percent of EITC-ineligible tax filers with children (Figure A.4), corroborating

\begin{itemize}
  \item \textsuperscript{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.
  \item \textsuperscript{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.
  \item \textsuperscript{17}I cannot rule out a threshold-crossing model (Schelling 1971) where a continuously changing covariate has a discrete impact on an outcome.
  \item \textsuperscript{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.
\end{itemize}
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 1982 (Figure A.4), this pattern bears little resemblance to the maternal employment trends in Figures 1.A and 1.B. 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.\textsuperscript{19} If anything, the expansion of 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.

\textsuperscript{19} 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, Eliehausen 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, Wolters 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.
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 (\textit{Mom}),\textsuperscript{20} whether the observation occurs after 1975 (\textit{Post1975}), and the DD variable of interest (\textit{Mom} × \textit{Post1975}). Column 2 adds state and year fixed effects to account for idiosyncratic state traits and annual shocks affecting all women.\textsuperscript{21} 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 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.

\textsuperscript{20}Restricting (\textit{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: 0.037 (0.009).

\textsuperscript{21}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-women-group (mother or non-mother) level also yields statistically significant estimates, with slightly larger standard errors of 0.02.
Across each set of controls the DD estimate is stable between 3.9 and 4.9 percentage points (or 7.3 and 9.5 percent from a baseline of 53 percent)\(^{22}\) and significant at the 99-percent level. Results imply that about one million mothers began working because of the 1975 EITC.\(^{23}\) The EITC is responsible for about a third 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 0.034 [0.009]), and additional robustness checks (Appendix B).\(^{24}\)

### 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 the EITC causing this rise in maternal employment. Traits associated with these heterogeneous responses are also used in section VI.F to predict state-

\(^{22}\)53 percent of all mothers in the sample work, compared to 63 percent in Figure 1.A for the “high-impact” sample (section III.B). 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. If this misallocation occurs at random the estimated employment effects of the EITC should be scaled up by 19 percent (Eissa and Liebman 1996).

\(^{23}\)47 percent of the 52.8 million women 15-44 in the 1980 Census are mothers (March CPS). 4 percentage points of these mothers corresponds to about 1 million mothers.

\(^{24}\)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.
level EITC responses and test whether states with larger EITC responses had larger post1975 increases in approval of working women.

**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 3.A). Second, spousal earnings increased the likelihood that the highest feasible indifference curve is achieved with zero labor supply (point A in Figure 3.A).

I verify this heterogeneity in Table 3 column 1, where I add the variable \( \text{Mom} \times \text{Post1975} \times \text{Unmarried} \) to equation (2) and interpret its coefficient (4.2 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 64.3 percent) as the overall effect of the EITC on unmarried mothers.\(^{25}\)

To estimate the average effect of the EITC on married mothers, I carry out two approaches in Table 3. In column 2, I restrict the sample to married women and find that the EITC had a statistically insignificant effect of 1.6 percentage points. In column 1, I estimate an effect of 2.7 percentage points, which is statistically significant but also statistically identical to the estimate in column 2. These estimates align with prior EITC research that has consistently found a larger response among single mothers.\(^{26}\)

\(^{25}\)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).

\(^{26}\)See Eissa and Liebman (1996), Meyer and Rosenbaum (2001), Grogger (2003), and
Although I find a statistically insignificant average response among married mothers to the 1975 EITC (Table 3 column 2), 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 income limit, the EITC increased the employment of this group by 4.9 percentage points, or 10.6 percent. I also test for a negative correlation between spousal earnings and EITC response by adding a variable to equation (2) that interacts $Mom \times Post_{1975}$ with spousal earnings. Column 5 shows that the treatment effect on married women with zero spousal earnings was 6.2 percentage points and declined by 0.9 percentage points for every $10,000 (2013 dollars) in spousal earnings.

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.


A quarter of males earned below the 1975 EITC kink point of $18,000 (2013 dollars).

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; this estimate steadily declines as women with spouses earning below $20,000, $30,000, etc., are incrementally added to the sample.

This 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. Predicting spousal earnings yields similar results. I find similar responses to the 1986, 1993 EITC expansions (Table A.3). Heterogeneous responses among married women also found by (Eissa and Hoynes 2004, Table 8) and Eissa and Hoynes (2006b).
Heterogeneous Treatment Effects: Education

Education is often used as a proxy for EITC eligibility 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), $Mom \times Post \times (<12 \text{ YrsEd})$ and $Mom \times Post \times (12 - 15 \text{ YrsEd})$, 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. 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 5.1, 4.3, and -0.2 percentage points (or 11.3, 8.0, and -0.3 percent).

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. I estimate the effect on this group

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30Sample 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).
31Low-education mothers were more than twice as likely to be EITC-eligible as high-education mothers (42 and 20 percent).
32I also find larger responses among younger mothers, mothers of younger children, and similar responses among white and nonwhite mothers (see Table A.4).
33Disability and unemployment rates were largely exogenous in the 1970s (Autor and Duggan 2003). Figure A.2 shows a smooth increase in education during this period.
by adding a variable to equation (2) that interacts $Mom \times Post_{1975}$ 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 5.9 percentage points (or 9.4 percent).

**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, (0.2 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 columns 4 and 7) extends equation (2) and creates a third difference for triple differences (DDD).  

\begin{equation}
P(E_{ist}) = f(\beta_1 Mom \times Post_{1975} \times Treat_{ist} + \beta_2 X_{ist} + \delta_{st} + \epsilon_{ist})
\end{equation}

The estimate of $\beta_1$ is 4.5 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. When men from Table 3 column 8 are used as a comparison group, I find a similar DDD estimate in Table 4 column 2 (4.4 percentage points).

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34 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).

35 Equation (3) also controls for $Treat$, $Mom \times Treat$, $Post_{1975} \times Treat$, $Mom \times Post_{1975}$, 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 for $y \in [1970, 1975]$ (p-value 0.56), become increasingly positive for $y \in [1975, 1979]$, and remain positive and flat for $y \in [1979, 1985]$ (statistically identical, p-value 0.16). 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).

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), and

\[\text{36The EITC does not pay until the following tax refund; it could take a year before EITC recipients became aware of the EITC (Liebman 1998). To test whether EITC response required an understanding of the tax code (Chetty, Friedman and Saez 2013, Bhargava and Manoli 2015), I plot the annual response by education subgroup and do not find quicker responses by higher-education mothers (omitted).}\]
and the EITC-ineligible placebo group (from Table 3 column 4). Among the “high-impact” sample, the EITC led to increases of 73.7 annual work hours and $1556.6 in annual earnings (Table 4 columns 1 and 4). Among the sample of all women, the EITC led to smaller increases in work hours (43.1) and earnings ($964.8) (columns 2 and 5). Results capture both intensive and extensive margins, but primarily reflect participation responses.\cite{37} Among the placebo group, columns 3 and 6 show that the EITC had a statistically insignificant effect on work hours (-5.6) and earnings (450.9), which corroborates the placebo test in Table 3 column 4.

\textbf{B. The EITC and the Distribution of Hours and Earnings}

Where in the annual hours and earnings 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 work hours or earnings in a particular range. Figures 4 and 5 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.

Figure 4 shows that the most common response to the EITC was to work full-time, full-year (about 2000 annual hours).\cite{38} The EITC may also have led to small increases in part-time work, although estimates on annual hours below 2000 are not statistically significant.

\cite{37}See Figures 4 and 5 for evidence. As a percent, these four estimates are 9.6, 13.1, 7.3, and 9.6. 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 2006a); although see Kline and Tartari (2016).

\cite{38}Annual hours combines the categorical \textit{weeks worked last year} variable (continuous variable not available until 1976 CPS) and \textit{hours worked last week}, in an attempt to reduce measurement error (Bound, Brown and Mathiowetz 2001).
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 4 shows that many mothers began working full time, this maps to about $14,000 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. Consistent with previous results, mothers were about 4 percentage points less likely to have zero work hours or earnings (Figures 4 and 5).39

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, 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, 39To isolate intensive-margin responses, I re-run the analysis in Figures 4 and 5 on working women, and find (noisy) evidence of more mothers working over 1000 hours and earning between $10,000 and $20,000.

20
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 44th centile, with a positive but decreasing effect higher up the earnings distribution. 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 0.43 (0.11) and 0.49 (0.14). See Table C.1 for complete details. Accounting for public assistance take-up rates yields a slightly larger elasticity between 0.48 (0.12) and 0.51 (0.15). Finally, I estimate the total intensive plus extensive margin elasticity from the annual work hours and earnings estimates in Table 4 to be 0.44 (0.09) and 0.59 (0.10). These elasticity estimates are larger than those for more recent decades, but are consistent with elasticity estimates for this period.41

Work hours yield a similar pattern. Oddly, this analysis also suggests that the EITC had a small positive effect on the top of the earnings distribution and could reflect unobservables not available in the CPS, such as work experience, which is known to be a major reason that the gender wage gap narrowed in the 1980s (O’Neill and Polachek 1993, Blau and Kahn 2000). Although most mothers earned in the EITC-eligible range: even 90th-centile earnings were at the end of the EITC phase-out region (Figure 3.B).

Female labor-supply elasticity has steadily declined since World War II (Goldin 1990): Bowen and Finegan (1969) finds 0.67 in 1960; Fields (1970) finds 0.52 in 1970;
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 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 2). 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 al-

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.

42Exposure 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 on gender-role attitudes that largely focuses on describing the time trends and correlates of these attitudes (Thornton and Freedman 1979, Thornt on, Alwin and Camburn 1983, Plutzer 1988, Lottes and Kuriloff 1992).

43Media has been shown to affect teen pregnancy (Kearney and Levine 2015), divorce (Chong and Ferrara 2009), and fertility (La Ferrara, Chong and Dur yea 2012). See DellaVigna and Ferrara (2015) for a recent literature review.
tered 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. 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 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. Table A.5 shows GSS sample summary statistics and Table A.6 shows gender-

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44 Additional evidence that various attitudes can be altered via exposure has also been 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 the role of 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 and whether these individual-level attitude changes scale at the macro-level.

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.6).
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 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).

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

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

\[
(5) \quad \Delta GenderEquality_{s}^{(1976-85)-(1972-75)} = \gamma EITC Response_s + \delta \Delta X_s + \epsilon_s.
\]

\( \gamma \) 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

\(^{47}\)Results are robust to extending the GSS sample to any year between the early 1980s and the 1990s (Figure A.7). 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).
Topel 2002). $X_s$ are controls to account for state-level traits. Regressions are weighted by state population since observations represent grouped data.\textsuperscript{49}

\section*{B. Results}

Using equation (5) and no controls, Figure 7 shows that each percentage-point increase in state EITC response led to a 1.8-percentage-point increase in state-level preferences for gender equality (p-value 0.001).\textsuperscript{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).\textsuperscript{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.\textsuperscript{52} The effect is stable between 1.5 and 2.0 percentage points, even when all 12 controls are included together (Table 6). Changes in gender-equality preferences do not

\textsuperscript{49}Similar results if unweighted or weighted by the standard-error inverse (equation 4). \textsuperscript{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).

\textsuperscript{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 $\gamma$, and iterate 10,000 times. Figure A.8 shows that the actual estimate (0.0177) is in the top 0.07 percent of permutations and unlikely to occur by chance.

\textsuperscript{52}State controls are education, age, marriage, race, employment, earnings, whether mother worked and mother’s education, fraction Democrat and religious, and views on public assistance and racial-equality. Results are similar using the pre1975-post1975 change (Panel A) or the pre1975 level (Panel B). See Table A.5 for summary statistics.
seem to be driven by demographics or general trends in social attitudes.  

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 $\gamma$ in equation (5) is 0.021 (p-value 0.001) and for higher-education adults it is 0.005 (p-value 0.58). 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-

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.
value = 0.12) on racial-equality preferences after 1975. Changes in gender-equality preferences do not seem to be driven by general trends in attitudes.

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 0.95), and interestingly, a negative relationship between state EITC response and the 1974 preference level. This negative estimate suggests that the EITC may have led to an attitude “catch up” among states with lower gender-equality preferences.

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

54 The relationship between EITC response and changes in racial attitudes is even less significant (p-value 0.67) when education is controlled for (Table 7 Panel B column 7).

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.9 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.
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. If mean reversion drove attitude changes after 1975, it should also have driven attitude changes in later decades. Figure A.10 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 0.018 to 0.012, 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 education, a trait associated with both EITC response and social attitudes: EITC response continues to have an independent effect on attitude changes even when pre1975 attitudes and education are controlled for (estimate 0.013 [0.005] in column 9).58

57 I also find a strong positive correlation between state-level gender-equality preferences in the 1970s, 1980s, and 1990s (not shown).
58 The estimate of EITC response is similar and significant at the 10-percent level when pre1975 attitudes and all 12 controls from Table 6 column 9 are used.
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 concerns about the potential endogeneity of gender-equality preferences 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 $EITC \ Response_{s}$ (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 $EITC \ Response_{s}$ should be correlated and interpreted as the effect of an exogenous increase in maternal employment on gender-equality preferences. Fourth, $X_{s}^{pre1975}$ should not be correlated with gender-equality preferences before 1975. Conditions one and four together suggest that $X_{s}^{pre1975}$ only affected preferences indirectly through state EITC response.

Figure 8 shows that these four conditions are (largely) met using female education. Figure 8 Panel A shows that pre1975 female education is negatively correlated with gender-equality 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) and illustrates the best-fit line used to generate predicted state EITC response (p-value 0.002). Panel C

59A complementary approach saves the residuals from the regression of $EITC \ Response_{s}$ on $X_{s}^{pre1975}$, regress attitude changes on these residuals, and show that the correlation is zero (results omitted).
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 preferences before 1975 (p-value 0.14).

In Table A.7, I repeat the exercise in Figure 8 using several other state-level demographic and occupational traits and find that predicted EITC response is again associated with increases in gender-equality preferences, with or 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 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).

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

60 These factors include single mothers, fraction female, male earnings, and fraction of jobs that are teachers or librarians, housekeepers or cleaners, and bakers or food makers.

61 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).
(and large increases in working women) affected social attitudes is feasible since Gallup began asking questions related to gender equality in the 1930s (see Figure A.12 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 7-percent increase in maternal employment, which represents about one million mothers and a participation elasticity of 0.43. 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. Although

62 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).
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 pre-1975 demographic and occupational traits – had larger increases 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.

REFERENCES


36


VIII. Tables and Figures
<table>
<thead>
<tr>
<th>Variable</th>
<th>All Women</th>
<th>Mothers</th>
<th>Women without Kids</th>
</tr>
</thead>
<tbody>
<tr>
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<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
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<tr>
<td>Age</td>
<td>29.0</td>
<td>32.9</td>
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</tr>
<tr>
<td></td>
<td>(8.4)</td>
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<td>Years of Education</td>
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<td>12.1</td>
<td>12.3</td>
</tr>
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<td>(2.5)</td>
<td>(2.6)</td>
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<td>Black</td>
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<td>0.12</td>
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</tr>
<tr>
<td></td>
<td>(0.33)</td>
<td>(0.32)</td>
<td>(0.33)</td>
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<tr>
<td>Hispanic</td>
<td>0.06</td>
<td>0.07</td>
<td>0.05</td>
</tr>
<tr>
<td></td>
<td>(0.24)</td>
<td>(0.25)</td>
<td>(0.22)</td>
</tr>
<tr>
<td>Kids Under 5</td>
<td>0.25</td>
<td>0.45</td>
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</tr>
<tr>
<td></td>
<td>(0.43)</td>
<td>(0.50)</td>
<td>(0.00)</td>
</tr>
<tr>
<td>Number of Kids</td>
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<td>2.24</td>
<td>0.00</td>
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<tr>
<td></td>
<td>(1.44)</td>
<td>(1.24)</td>
<td>(0.00)</td>
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<td>Employed</td>
<td>0.65</td>
<td>0.58</td>
<td>0.75</td>
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<tr>
<td></td>
<td>(0.48)</td>
<td>(0.49)</td>
<td>(0.43)</td>
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<tr>
<td>Individual Earnings (2013 $)</td>
<td>$13,028</td>
<td>$11,620</td>
<td>$14,808</td>
</tr>
<tr>
<td></td>
<td>(17,007)</td>
<td>(16,225)</td>
<td>(17,789)</td>
</tr>
<tr>
<td>Individual Earnings (2013 $)</td>
<td>$19,960</td>
<td>$20,177</td>
<td>$19,750</td>
</tr>
<tr>
<td>(Conditional on Earnings &gt; 0)</td>
<td>(17,458)</td>
<td>(16,866)</td>
<td>(18,012)</td>
</tr>
<tr>
<td>Household Earnings (2013 $)</td>
<td>$41,268</td>
<td>$53,474</td>
<td>$25,825</td>
</tr>
<tr>
<td></td>
<td>(39,822)</td>
<td>(40,422)</td>
<td>(33,137)</td>
</tr>
<tr>
<td>Household Earnings (2013 $)</td>
<td>$48,755</td>
<td>$59,926</td>
<td>$32,758</td>
</tr>
<tr>
<td>(Conditional on Earnings &gt; 0)</td>
<td>(38,840)</td>
<td>(38,005)</td>
<td>(34,143)</td>
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<tr>
<td>Household Earnings Below EITC Limit</td>
<td>0.46</td>
<td>0.30</td>
<td>0.66</td>
</tr>
<tr>
<td></td>
<td>(0.50)</td>
<td>(0.46)</td>
<td>(0.47)</td>
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<tr>
<td>Household Earnings Below EITC Limit</td>
<td>0.31</td>
<td>0.20</td>
<td>0.45</td>
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<td>(0.46)</td>
<td>(0.40)</td>
<td>(0.50)</td>
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<td>Annual Weeks Worked</td>
<td>25.9</td>
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<td>37.9</td>
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<tr>
<td>(Conditional on Weeks Worked &gt; 0)</td>
<td>(16.6)</td>
<td>(16.4)</td>
<td>(16.9)</td>
</tr>
<tr>
<td>Weekly Hours Worked</td>
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<td>16.2</td>
<td>20.5</td>
</tr>
<tr>
<td></td>
<td>(19.4)</td>
<td>(19.1)</td>
<td>(19.5)</td>
</tr>
<tr>
<td>Weekly Hours Worked</td>
<td>34.0</td>
<td>33.9</td>
<td>34.1</td>
</tr>
<tr>
<td>(Conditional on Hours Worked &gt; 0)</td>
<td>(12.9)</td>
<td>(12.7)</td>
<td>(13.1)</td>
</tr>
<tr>
<td>Observations</td>
<td>550,904</td>
<td>310,875</td>
<td>240,029</td>
</tr>
</tbody>
</table>

Notes: Data source: 1971-1986 March CPS data. Individual March CPS weights used. Sample contains all women 16 to 45 years old. Standard deviations are in parentheses. Kids under 5 is binary. 358,955 observations have positive earnings, 465,247 have positive household earnings, 375,906 have positive weeks worked last year, and 292,911 have positive hours worked last week.
Table 2. The 1975 EITC Increased Maternal Employment, Robust to Various Sets of Controls

<table>
<thead>
<tr>
<th>Variables</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
<th>(5)</th>
<th>(6)</th>
<th>(7)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>0.049</td>
<td>0.049</td>
<td>0.046</td>
<td>0.04</td>
<td>0.041</td>
<td>0.05</td>
<td>0.039</td>
</tr>
<tr>
<td></td>
<td>(0.007)</td>
<td>(0.006)</td>
<td>(0.007)</td>
<td>(0.009)</td>
<td>(0.008)</td>
<td>(0.008)</td>
<td>(0.008)</td>
</tr>
</tbody>
</table>

**Controls**

- State and Year FE
- Demographic Controls
- Unemployment Rate
- "Kitchen-Sink" Controls

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Model</td>
<td>Logit</td>
<td>Logit</td>
<td>Logit</td>
<td>Logit</td>
<td>Probit</td>
<td>OLS</td>
<td>OLS</td>
</tr>
<tr>
<td>R-squared</td>
<td>0.157</td>
<td>0.177</td>
<td>0.177</td>
<td>0.177</td>
<td>0.177</td>
<td>0.177</td>
<td>0.177</td>
</tr>
</tbody>
</table>

Mean Dependent Variable Across Years and Across Treatment and Control Groups = 0.65

Mean Dependent Variable for Treatment Group in 1975 = 0.53

Notes: Data source: 1971-1986 March CPS data. Sample includes all women 16 to 45 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 data 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-lit, nonwhite-post1975, age-kid, 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), having kids and number of kids, and married, and finally state-year measures of manufacturing employment interacted with low education, having kids, and married.
Table 3. Heterogeneous and Subgroup Treatment Effects of the 1975 EITC on Employment

<table>
<thead>
<tr>
<th>Subgroup:</th>
<th>Larger Response Among Unmarried Mothers</th>
<th>Education</th>
<th>“High-Impact”</th>
<th>Single Men</th>
</tr>
</thead>
<tbody>
<tr>
<td>Description:</td>
<td>All Married</td>
<td>Spouse Earning Below EITC Limit</td>
<td>Spouse Earning Above EITC Limit</td>
<td>Response Negatively Correlated With Spousal Earnings</td>
</tr>
<tr>
<td>Variables</td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(4)</td>
</tr>
<tr>
<td>Mom x Post1975</td>
<td>0.027</td>
<td>0.016</td>
<td>0.049</td>
<td>0.007</td>
</tr>
<tr>
<td>Mom x Post1975 x Unmarried</td>
<td>0.042</td>
<td>(0.012)</td>
<td>-0.009</td>
<td>(0.001)</td>
</tr>
<tr>
<td>Mom x Post1975 x Spousal Earnings (in 10,000s of 2013 $)</td>
<td>0.027</td>
<td>(0.011)</td>
<td>0.049</td>
<td>(0.018)</td>
</tr>
<tr>
<td>Mom x Post1975 x (&gt;12 Yrs Ed)</td>
<td>0.042</td>
<td>(0.012)</td>
<td>0.043</td>
<td>(0.005)</td>
</tr>
<tr>
<td>Mom x Post1975 x High-Impact</td>
<td>0.050</td>
<td>(0.014)</td>
<td>0.065</td>
<td>(0.020)</td>
</tr>
<tr>
<td>Dad x Post1975</td>
<td>0.065</td>
<td>(0.020)</td>
<td>0.065</td>
<td>(0.018)</td>
</tr>
</tbody>
</table>

Observations: 550,904 | 321,147 | 55,313 | 265,834 | 321,147 | 550,904 | 550,904 | 231,087 | 0.002 | (0.020) |

Mean Dependent Variable: 0.65 | 0.61 | 0.54 | 0.62 | 0.61 | 0.65 | 0.65 | 0.78 | 0.002 | (0.020) |

Mean Dep Var for 1975 Treat. Group: 0.53 | 0.51 | 0.46 | 0.52 | 0.51 | 0.53 | 0.53 | 0.85 | 0.002 | (0.020) |

Notes: Data source: 1971-1980 March CPS data. All samples limited to 16 to 45 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 x Unmarried in column 1, Mom x Post1975 x (>12 Yrs Ed) in column 6, and Mom x Post1975 x Non-High-Impact Sample in column 7. In column 1, I predict marital status based on post1975 traits using OLS and marriage, age, years of education, state, and a linear time trend. This avoids potentially endogenous marital responses to the EITC. The choice of covariates used to predict marital status – or using actual marital status – produces qualitatively similar results. The overall effect on unioned women is 65 percentage points (or 10.7 percent from a base of 65 percent). In column 5, predicting spousal earnings (based on age, education, race, number of kids, kids under 5) yields similar results: 0.062 and 0.065. “High-Impact” sample excludes EITC-inseligible 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 means in 1975 is 0.62 (see Figure 1A). The mean dependent variable for 1979 mothers with less than equal to, or more than 12 years of education in column 1 is 0.45, 0.54, and 0.09. The EITC income limit in columns 3 and 4 was 44,000 nominal dollars in 1975 and increased to $50,000 in 1979 (or about $180,000 in 2013 dollars).
Table 4. Triple Differences Corroborates Difference in Differences

<table>
<thead>
<tr>
<th>Variables</th>
<th>Comparing High-Impact and Placebo Women (Table 3 Columns 4 and 7)</th>
<th>Comparing High-Impact Women and Single Men (Table 3 Columns 7 and 8)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mom x Post1975 x EITC Eligible</td>
<td>0.045</td>
<td>0.044</td>
</tr>
<tr>
<td>(0.011)</td>
<td></td>
<td>(0.019)</td>
</tr>
<tr>
<td>Observations</td>
<td>550,904</td>
<td>787,230</td>
</tr>
</tbody>
</table>

Note: Data source: 1971-1986 March CPS data. Samples limited to 16 to 45 year olds. Binary dependent variable employment equals 1 for positive earnings. High-impact women from Table 3 column 1, placebo women from Table 3 column 4, and single men from Table 3 column 8. Equation (3), CPS weights, full set of controls from Table 2 column 4 used along with interactions of each control with being a "high-impact" women, 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)

<table>
<thead>
<tr>
<th>Dependent Variable:</th>
<th>Annual Work Hours</th>
<th>Annual Earnings (2013 $)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>&quot;High-Impact&quot; Group</td>
<td>All</td>
</tr>
<tr>
<td>Mean Dependent Variable:</td>
<td>972</td>
<td>765</td>
</tr>
<tr>
<td>Mean Dep. Var. 1975</td>
<td>769</td>
<td>592</td>
</tr>
<tr>
<td>Variables</td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>Mom x Post1975</td>
<td>73.7</td>
<td>43.1</td>
</tr>
<tr>
<td></td>
<td>(14.9)</td>
<td>(11.7)</td>
</tr>
<tr>
<td>Observations</td>
<td>230,399</td>
<td>550,904</td>
</tr>
<tr>
<td>R-squared</td>
<td>0.269</td>
<td>0.215</td>
</tr>
</tbody>
</table>

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 16 to 45 years old. EITC-ineligible placebo group have spousal earnings above the 1973 EITC income limit of $18,000 (2013 dollars). 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; zero assigned if annual work hours equals zero, even if reported annual earnings is positive) as outcome. 0.94 (20), 0.91 (0.16), and 0.02 (0.33), which represent percent increases of 13, 6, and 0 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

<table>
<thead>
<tr>
<th>Controls:</th>
<th>Demographics</th>
<th>Other Social Attitudes</th>
<th>Using All Controls</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Educ.</td>
<td>Married</td>
<td>Nonwhite</td>
</tr>
<tr>
<td><strong>Panel A: Controlling for Pre1975-Post1975 State-Trait Changes</strong></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td>EITC-Led Increase in Working Mothers (in Percentage Points)</td>
<td>0.018</td>
<td>0.018</td>
<td>0.018</td>
</tr>
<tr>
<td>(in Percentage Points)</td>
<td>(0.004)</td>
<td>(0.004)</td>
<td>(0.004)</td>
</tr>
<tr>
<td>Observations</td>
<td>32</td>
<td>32</td>
<td>32</td>
</tr>
<tr>
<td>R-squared</td>
<td>0.429</td>
<td>0.324</td>
<td>0.360</td>
</tr>
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</table>

<table>
<thead>
<tr>
<th>Variables</th>
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<th>(6)</th>
<th>(7)</th>
<th>(8)</th>
<th>(9)</th>
</tr>
</thead>
<tbody>
<tr>
<td>EITC-Led Increase in Working Mothers (in Percentage Points)</td>
<td>0.015</td>
<td>0.018</td>
<td>0.017</td>
<td>0.017</td>
<td>0.018</td>
<td>0.018</td>
<td>0.016</td>
<td>0.018</td>
<td>0.017</td>
</tr>
<tr>
<td>(in Percentage Points)</td>
<td>(0.005)</td>
<td>(0.004)</td>
<td>(0.005)</td>
<td>(0.004)</td>
<td>(0.004)</td>
<td>(0.004)</td>
<td>(0.005)</td>
<td>(0.004)</td>
<td>(0.008)</td>
</tr>
<tr>
<td>Observations</td>
<td>32</td>
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<td>32</td>
<td>32</td>
<td>32</td>
<td>32</td>
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</tr>
<tr>
<td>R-squared</td>
<td>0.331</td>
<td>0.315</td>
<td>0.308</td>
<td>0.321</td>
<td>0.336</td>
<td>0.328</td>
<td>0.310</td>
<td>0.352</td>
<td>0.546</td>
</tr>
</tbody>
</table>

Notes: 1972-1985 restricted GSS data with state-level identifiers. Gender-equality preferences constructed from the GSS variable fwork which asks respondents whether married women should work. GSS sample reflects adults 18-60 years old in 32 states. State-level EITC response estimated from equation (1). 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. Fernández et al. (2004) shows that mother's employment effects 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 1-8 as well as the four additional covariates just mentioned. Education measured in years, working 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 mark16 and ineduc, democrat from partyid, racial equality from racpres, religious from religion, and too much welfare from toomw. Heteroskedasticity-robust standard errors in parentheses. Regressions weighted by state population, though unweighted results are similar.
<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>All</td>
<td>Men</td>
<td>Women</td>
<td>Low-Educ.</td>
<td>High-Educ.</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(4)</td>
</tr>
<tr>
<td>EITC-Led Increase in Working Mothers (in Percentage Points)</td>
<td>0.018</td>
<td>0.016</td>
<td>0.019</td>
<td>0.021</td>
<td>0.005</td>
</tr>
<tr>
<td></td>
<td>(0.004)</td>
<td>(0.006)</td>
<td>(0.006)</td>
<td>(0.005)</td>
<td>(0.009)</td>
</tr>
<tr>
<td>1974 Fraction of State Pop. that Approves of Working Women</td>
<td>32</td>
<td>32</td>
<td>32</td>
<td>32</td>
<td>32</td>
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<tr>
<td>Observations</td>
<td>0.309</td>
<td>0.112</td>
<td>0.312</td>
<td>0.271</td>
<td>0.017</td>
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</table>

Panel B: Controlling for Pre1975 State-Level Average Education

<table>
<thead>
<tr>
<th>Variables</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
<th>(5)</th>
<th>(6)</th>
<th>(7)</th>
<th>(8)</th>
<th>(9)</th>
</tr>
</thead>
<tbody>
<tr>
<td>EITC-Led Increase in Working Mothers (in Percentage Points)</td>
<td>0.015</td>
<td>0.010</td>
<td>0.02</td>
<td>0.022</td>
<td>0.004</td>
<td>0.003</td>
<td>-0.008</td>
<td>-0.009</td>
<td>0.013</td>
</tr>
<tr>
<td>1974 Fraction of State Pop. that Approves of Working Women</td>
<td>32</td>
<td>32</td>
<td>32</td>
<td>32</td>
<td>32</td>
<td>32</td>
<td>32</td>
<td>32</td>
<td>32</td>
</tr>
<tr>
<td>Observations</td>
<td>0.333</td>
<td>0.169</td>
<td>0.313</td>
<td>0.271</td>
<td>0.021</td>
<td>0.278</td>
<td>0.370</td>
<td>0.082</td>
<td>0.472</td>
</tr>
</tbody>
</table>

Notes: 1972-1985 restricted GSS data with state-level identifiers. Gender equality preferences constructed from the GSS variable jwork 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. Column 7 shows that states with the lowest approval of working women had the largest EITC responses, however, 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, which suggests that observed increases in working women do not simply reflect mean reversion in state attitudes. Regressions are weighted by state population, though unweighted results are similar. Heteroskedasticity-robust standard errors in parentheses.
Notes: 1970-1986 March CPS data. Employment defined as positive income. Estimates in Figure 1.A calculated by regressing employment on a constant for each group-year. Best-fit lines shown for 1969-75, 1975-79, 1979-85. In Figure 1.B, unadjusted relative employment – relative to women without kids – calculated by regressing employment on whether a woman has a kid and a constant for each year. Regression adjusted annual gap are probit estimates of $Mom \times Year$ with the full set of controls from Table 2 column 4. The estimates are jointly statistically insignificant for all years before 1975 (p-value 0.56); the 1979 to 1985 estimates are statistically identical (p-value 0.16). “High-impact” sample (section III.B) used, which includes women 16-45 and excludes married women with spousal earning above the 1975 EITC limit of $36,000 (in 2013 $), full-time students, disabled, and retired. Kids defined as 0-18 years old. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level.
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%2Cstay%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%2Cworking%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%20stay%20at%20home%20mother%20%2B%28stay%20at%20home%20moms%29%3B%2Cc0. Accessed 9/5/16.
Figure 3.A. Budget Constraint Under the 1975 EITC

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 3.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 4. Effect of the EITC on the Distribution of Annual Work Hours

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 work hours in the specified range. The mean dependent variable for the seven unconditional regressions are 0.35, 0.11, 0.09, 0.08, 0.11, 0.18, and 0.08, and for the seven conditional regressions are 0.18, 0.13, 0.11, 0.10, 0.14, 0.23, and 0.10. Sample sizes are 230,399 and 173,752. Standard errors are computed by the delta method, robust to heteroskedasticity, and clustered at the state level.

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

Notes: Same data, sample, and approach as Figure 4. Each estimate is from a different logit regression of having annual earnings in the specified range. The mean dependent variable for the nine unconditional regressions: 0.25, 0.27, 0.15, 0.13, 0.10, 0.06, 0.03, 0.01, 0.01; and conditional on working: 0.0, 0.35, 0.20, 0.18, 0.13, 0.07, 0.04, 0.02, 0.02.
Figure 6. Effect of the EITC on Annual Earnings (Quantile Dif in Dif)

Notes: Same data, sample, controls, and standard errors as Figure 4. 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, 4814, 12159, 20030, 28132, 38552. The top few centiles are noisy and omitted. More details in footnote 40.

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.

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

Notes: Data, sample, standard errors described in Figure 7. Average state female education correlated with Post1975-Pre1975 attitude changes (Panel A) and state EITC response (Panel B), but not with 1972-1975 attitude changes (Panel D). Panel C shows that predicted EITC response (from Panel B) is associated with changes in gender-equality attitudes. In Table A.7, I repeat this analysis using other pre1975 state-level demographic and occupational traits, with and without region fixed effects. Across traits, predicted EITC response suggests that the EITC positively affected gender-equality preferences.