The Effect of EITC Exposure in Childhood on Marriage and Early Childbearing

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Abstract

This study analyzes the effects of the Earned Income Tax Credit (EITC) on marriage and childbearing of men and women exposed to the EITC in childhood. Findings suggest that exposure in childhood leads women to delay marriage and first births in early adulthood (ages 18-25), but not men. These results have implications for the well-being of both individuals exposed to the EITC in childhood as well as their future children. In addition, because childless adults cannot claim the EITC until age 25, our back-of-the-envelope calculations suggest that these delays likely save up to $199 million annually in social welfare costs.

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Most Americans believe that becoming a parent as a teenager is far from optimal. In a 2013 Gallup poll of over 5000 adults,[2] 48 percent of respondents said the ideal age to begin having children was 25; only 9 percent of respondents thought the ideal age was younger than 21. Researchers have shown that teenage childbearing is associated with a number of negative outcomes for the mother and child, e.g., decreased educational attainment, poor labor market outcomes, increased likelihood of social welfare program receipt, and poor health (Haveman, Wolfe, and Peterson 1997; Hoffman 2006; Mirowsky 2005).

The consequences of childbearing at a young age have been hotly debated. Beginning with the research of Geronimus and Korenman (1992) scholars have been trying to determine if the negative outcomes associated with teenage childbearing is a real causal effect or simply due to selection. More recent evidence suggests that teenage childbearing is not as harmful as was initially thought, but still has many adverse consequences (Ashcraft, Fernandez-Val, and Lang 2006; Fletcher and Wolfe 2009). Thus, social programs that delay childbearing until after one's teenage years, whether the actual intention of the program or otherwise, provide an important social benefit.

To this effect, the federal government has implemented several programs explicitly designed to affect the family formation decisions of young women: Title X of the Public Health Services Act, the Family Planning Benefit within the Medicaid program, and Title V, Section 510 Abstinence Only Program, to name a few. Unsurprisingly, there is an extensive research literature investigating the efficacy of these family policies (see e.g., U.S. Department of Health and Human Services 2018; Kearney and Levine 2009; Trenholm, et al. 2008). There is also a long research literature that asks if social welfare policies affect contemporaneous marriage and

[2] https://news.gallup.com/poll/165770/americans-ideal-age-women-first-child.aspx

fertility decisions, even when not the explicit goal of these policies. Theoretically, policies that reduce the cost of children or marriage should increase their likelihood (Becker 1974, 1991). However, most reviews of this literature show very little evidence that these policies affect family formation (see, for example, Lopoo and Raissian 2012, 2014; Moffitt 2003). Lopoo and Raissian (2014) suggest that the family effects of most social welfare programs, such as the Supplemental Nutrition Assistant Program (SNAP) or Temporary Assistance to Needy Families (TANF, also known as welfare), are indiscernible because the decision to become a parent or to marry is usually so fundamental to most people that small increases in benefits do not dominate their familial preferences.[3] They go on to argue that social welfare programs that provide large subsidies for significant portions of the population and that affect human capital accumulation, have a much larger capacity to affect family formation decisions. Programs, such as the Hope Scholarship in Georgia or the Say Yes Program in New York, that provide funding for college tuition might produce such a benefit.

In this paper, we ask if the Earned Income Tax Credit (EITC), one of the largest cash-transfer programs in the United States, has intergenerational benefits, i.e., if it affects the family formation decisions of men and women exposed to the EITC in childhood. Our primary interest is in knowing if EITC investment in one's youth, delays the timing of that individual's first birth.[4] As explained above, delaying childbearing beyond one's teenage years has been shown to have important benefits for the young mother and her child. In addition, a delay has direct

[3] Although, see Groves, Hamersma, and Lopoo (2018) who show that while Medicaid benefits do not affect the extensive fertility margin, they do affect the intensive margin.

[4] Demographers frequently make the distinction between the tempo and quantum components of fertility. The "tempo" refers to the timing of childbearing, while quantum refers to the total number of children a woman has during her lifetime. Due to the extensive data requirements necessary to investigate the quantum component, this paper focusses on the tempo alone. One would need completed fertility histories for all women along with their exposure to the EITC during their childhood. In effect, one would need about 60 years of data. At this point, there are no panels that would allow one to test the quantum component.

implications for program eligibility in adulthood. Childless adults cannot claim the EITC until age 25, while there is no such minimum age for parents.[5] Because of this age restriction for childless individuals, each year of foregone fertility is a year in which an individual is not eligible for the EITC, which will reduce the costs of the program. Because fertility and marriage are intertwined for many individuals, although less so today than in the past (Lundberg, Pollack, and Stearns 2016), we also investigate the relationship between EITC benefits in childhood and the timing of first marriage.

As we explain in greater detail below, the EITC is a refundable tax credit that provides benefits for families in the lower portion of the income distribution. The benefits from the program are quite large. For a family with three or more children, the maximum credit was worth over $6,000 in 2018. In addition, in 2017, 29 states and the District of Columbia offered additional benefits through the state EITC.[6] Over a child's formative years, then, this credit has the potential to provide tens of thousands of dollars of benefits for lower-income families. As such, the benefits provided by the EITC could affect family formation through a number of pathways, including by improving the human capital of the child, which increases the opportunity cost of early childbearing (Wolfe, Wilson, and Haveman 2001; Ellwood and Jencks 2004).

To date, there is no research examining how exposure to the EITC in childhood affects subsequent marriage and family formation patterns in adulthood, but there is a vast literature on how education affects early childbearing. Research on the Social Security Student Benefit Program, which provided large subsidies that covered tuition and other costs of attendance for the children of disabled, retired, or deceased parents during the 1970s and early 1980s, for

[5] Provided that the parents and children cannot be claimed by any other tax filer.
[6] https://www.cbpp.org/research/state-budget-and-tax/policy-basics-state-earned-income-tax-credits

example, showed that it not only increased participation in higher education (Dynarksi 2003), but also delayed marriage and childbearing among women who were eligible (Groves and Lopoo 2018). Exploiting policy changes in school dropout laws, previous research finds that increasing the age at which children can drop out of school reduces teen fertility (Black, Devereux, and Salvanes 2008; Geruso and Royer 2018) by 5 to 30 percent. Similar reductions in teen and early-20s fertility have been found for other education policy changes, such as lengthening the vocational upper secondary track in Sweden by one year (Gronqvist and Hall 2013). On the other hand, a study using school start age as an instrument for years of schooling found no reductions in teen fertility among children born just before the cut-off date, who attained more schooling before the legal dropout age, relative to children born just after the cut-off date (McCrary and Royer 2011). Recent evidence suggests that exposure to the EITC in childhood generated increases in human capital and earnings in adulthood (Bastian and Michelmore 2018). These increases in human capital may also result in delayed marriage and family formation among those exposed to the EITC in childhood.

In addition to the human capital pathway, EITC benefits during childhood could affect early childbearing through other avenues. A long line of research on the EITC shows that it increases labor supply among single mothers (Eissa and Liebman 1996; Meyer and Rosenbaum 2001; Ellwood 2000), lifts families out of poverty (Hoynes and Patel 2018), reduces maternal stress (Evans and Garthwaite 2013) and has positive effects for children, ranging from lower incidence of low infant birthweight (Hoynes, Miller, and Simon 2013), and higher test scores in childhood (Dahl and Lochner 2012, 2017). An employed parent may serve as a role-model for his/her child, which might make the child less inclined to give birth at a young age (Chase-Lansdale et al. 2003; Haveman and Wolfe, 1994, 1995). In addition, the reduction in maternal stress from the additional resources could improve parent-child relationships, which might lead to childbearing delays. (Chase-Lansdale et al. 2003; Mayer, 1997). On the other hand, because the EITC induces single mothers to work, it may lead to less supervision and, therefore, higher rates of teen fertility (Haveman and Wolfe, 1994, 1995; Hogan and Kitagawa 1985; McLanahan 1988; McLanahan and Sandefur 1994).

Using the Panel Study of Income Dynamics (PSID) and a panel of respondents born between 1968 and 1992, we exploit variation in the federal and state EITC to examine how policy-induced increases in exposure to the EITC in childhood affects subsequent marriage and fertility when children reach early adulthood; ages 18 to 25. We find that increases in EITC exposure in childhood delays the marital fertility of white women and the non-marital fertility of African American women. We find some support that the delays for white women are due to human capital increases, particularly a greater likelihood of completing college on time. We do not find evidence that EITC exposure affects the marriage or fertility for men through age 25.

Together, these results have implications for the well-being of children growing up with the EITC, and for the intergenerational transmission of EITC eligibility. Our results suggest that a $1,000 increase in EITC exposure during one's childhood reduces the likelihood of an early birth by around 0.7 percentage points or around 2 percent. We believe this is likely a change in the timing of childbearing rather than a change in total fertility. These EITC-induced delays then benefit parents and their children. In addition, because individuals cannot receive the EITC until age 25 unless they become parents themselves, these delays in fertility should reduce the number of eligible EITC recipients. We conduct a back-of-the-envelope cost savings estimate based on these delays and calculate that these EITC benefits paid during an individual's childhood may reduce EITC costs twenty years later by as much as $63 million annually. If one factors in other social welfare programs, such as Medicaid costs and SNAP benefits, the estimate can be as high as $200 million.

EITC Background

The EITC is a refundable tax credit targeting households with earnings below roughly 225% of the federal poverty line, or about $50,000 in 2018. The federal credit is based on earnings, marital status, and the number of children in the household. With the inclusion of state benefits, which typically piggyback off of the federal credit, the EITC benefit is worth as much as $9,200 in 2017 for a family residing in a state with an EITC worth 40% of the federal EITC (such as the District of Columbia). Together, the federal and state EITC could be worth more than 50% of a family's annual earnings, received as a lump-sum along with the rest of a tax refund.

The EITC structure is made up of three regions, a phase-in region, where benefits increase as earnings increase, a plateau region, where benefits remain unchanged as earnings fluctuate, and a phase-out region, where benefits are reduced as earnings increase. The phase-in rate varies depending on the number of children in the household, from 34 cents per dollar for a household with one child to 45 cents per dollar for a household with three children. Benefits phase-out at a rate of approximately 22 cents on the dollar.

Since its inception in 1975, when the federal credit was worth $500 ($1700 in 2016$) and the phase-in rate was 10 percent, there have been several federal and state policy changes to the benefit. In 1987, the credit was indexed to inflation, and throughout the 1980s and 1990s, the phase-in rate was gradually increased from 14 percent in 1990 for families of all sizes, to 34 percent for a family with one child in 1995, 40 percent for a family with two children starting in 1996, and 45 percent for a family with three children starting in 2009. These policy changes resulted in a change in the maximum federal credit available by roughly $2,000, controlling for inflation. A summary of the federal expansions is presented in Appendix Table 1.

In addition to the federal credit, 29 states have also implemented their own EITCs as of 2017. State benefits are usually a fixed percentage of the federal benefit, ranging from 3.5 percent to 40 percent of the federal benefit. Individuals eligible for the federal EITC are also typically eligible for state EITCs, provided they file a state tax return. States began implementing EITCs in the mid-1980s, but the number of states with EITCs increased more rapidly after welfare reform in 1996, when states were given more leeway with welfare block grants to use for funding state EITCs. While some of our youngest cohorts would have been exposed to state EITCs as children, the vast majority of our variation in EITC exposure in childhood is driven by the several federal expansions to the EITC since 1975.

Data

We use the Panel Study of Income Dynamics (PSID) for all analyses. The PSID originated in 1968 with roughly 4,800 families. Information from PSID respondents was collected annually until 1997 and biennially thereafter until 2017, the latest wave of data currently available. Family-level data is collected in each wave from the head of household, defined as the person who has the greatest financial responsibility for the family unit and who is at least 18 years of age. Most of the family units were husband-wife pairs in 1968, and the male was typically assumed to be the head of household. In addition to the family-level data collected in each wave, the PSID produces an individual data file, which includes specific information on the individuals in the PSID separate from the family-level data collected in each wave. For example, beginning in 1985, the PSID began collecting a complete fertility and marital history record for each individual in the PSID. We use both the family-level and individual-level data in this analysis.

Our analytic sample consists of all individuals born between 1968 and 1992. We stop our data series in 1992 allowing us to generate a marital and fertility history until all individuals turn 25. However, because we are also interested in whether any delays in fertility represent postponement or an overall decline in fertility, in some analyses, we restrict our sample to those born between 1968 and 1987 and assess first-time childbearing and marriage up through age 30.

Assessing fertility up through age 25 is a particularly relevant age for this analysis because childless individuals cannot claim the EITC until they turn 25. However, individuals of any age can claim the EITC if they have at least one qualifying child and no one else can claim the tax filer as a dependent on their tax return. Thus, any births averted before age 25 also imply a reduction in the number of individuals eligible for the EITC. We chose 1968-1992 birth cohorts since we can observe these individuals in the PSID from birth to age 15, and the youngest birth cohorts were exposed to some of the largest expansions to the federal EITC—those that occurred in the early 1990s—and thus provide substantial variation in EITC exposure over the course of childhood.

We follow individuals until they turn 25, the 2017 interview, they attrit from the sample, they attrit from the childbirth and adoption history supplement, or the marital history file, whichever comes first. We estimate the influence of the potential EITC benefits on several family formation variables measured at different points in the individual's life course. As such, the sample sizes for models estimating the different outcomes vary. The largest, however, is 6,622 for models of an indicator for births by age 18 and married by age 18. Given the different life-course trajectories, however, we estimate all models separately for men (n=3,321) and women (n=3,301). Following Bastian and Michelmore (2018), we derive a weight for each observation by taking the mean reported weight from the survey years the individual was born until age 18. All monetary measures are inflated to 2016 dollars using the CPI-U-RS. Table 1 presents descriptive statistics for the full analytic sample and separately for females and males.

EITC Measure

To analyze how exposure to the EITC in childhood affects childbearing and marriage in adulthood, we calculate the maximum federal and state EITC a child's family could receive given the number of children in the household, the state of residence, and the year. We cumulate the maximum federal and state EITC available from each year from a child's birth to age 15, allowing the state and number of children to vary in each year according to family circumstances. Notably, this measure is independent of a family's own income; changes in the maximum EITC available are driven by changes in household size, cross-state moves, and federal and state expansions to the credit over time. Under the assumption that cross-state moves and fertility are not influenced by the EITC, of which there is some evidence (e.g. Baughman and Dickert-Conlin 2009), variation in childhood exposure to the EITC is plausibly exogenous with respect to a family's own characteristics. Since we focus on a sample of individuals born between 1968 and 1992, we capture variation in the federal EITC stemming from its inception in 1975, and each subsequent policy change that occurred up through 2007, the year that our youngest cohort turns 15. The mean EITC benefit available for the PSID children from their birth year to age 15 was a little more than $38,000 (in 2016 dollars).

Outcomes

We observe several different outcomes in our analyses. To capture the timing of childbearing, we use a dichotomous measure asking if the individual had given birth by age 18, age 19, and so on until age 25. Table 1 shows that the timing of marriage and fertility is quite different for men and women, with men having children at older ages. For example, nearly 10 percent of women had given birth by age 18 compared to 3 percent of men. By age 25, over 40 percent of women had given birth, while just over 27 percent of men had children.

We use a similar measure for the timing of first marriage. Consistent with previous demographic research (see, for example, Auginbaugh, Robles, and Sun 2013), we find differences in the likelihood of marriage for men and women. Around 3.6 percent of women have been married by age 18 in the PSID compared with less than one percent of men. By age 25, nearly 40 percent of women have been married, while less than 30 percent of men have at the same age. We also combine these two measures to generate unmarried births by a particular age and a mutually exclusive indicator for births by a particular age when the respondent had been married at some point prior to that age. [7]

To explore some of the mechanisms for the findings, we generate an indicator for completing high school by age 19 and another for completing college by age 23. We also use a continuous measure of the number of years of education completed by age 25. Approximately 63 percent of respondents had completed high school by age 19, with 67 percent of women finishing compared to 59 percent of men. We also see a large gender disparity in the likelihood of college completion by age 23. Over 21 percent of women have finished college by 23, while

[7] Given this definition, marital births here reflects births that occurred after the first marriage. Some women may have married, subsequently divorced, and then had a child, and would still be considered to have had a “marital birth”.

under 13 percent of men have. These differences are evident in the education completed by age 25 measure as well. The mean value for women is 13.8 years. For men, it is 13.3 years.

Controls

For race and ethnicity, we generate a set of mutually exclusive variables based on the reports from the head in 1968: white; African American; Hispanic; and a very small set labeled "other," which includes those of Asian descent and missing race. The vast majority of the PSID is composed of white (nearly 80 percent) and African American (around 18 percent) respondents. To control for the socioeconomic status of the family, we use the number of years of education the head had completed in the year the child was born. This variable ranges between 0 and 17 years with a mean of 12.64 years. We also control for the proportion of time between the child's birth and age 15, that there were one, two, three, four, and five or more children in the home. We also have a measure of the proportion of years the head of household was married between the child's birth and age 15.

In addition to the individual measures, we derive several state level controls measured in the year the individual was 17. We measure the mean maximum Aid to Families with Dependent Children/Temporary Assistance for Needy Families benefit for a family of three to control for the state generosity in social welfare programs, which is about $592 dollars per month. We use the state unemployment rate to capture the business cycle effects (mean = 5.67) and state minimum hourly wage (mean = $6.94) as a measure of the value of low-skilled work in the state. These measures were collected from the University of Kentucky's Center for Poverty Research database.

Empirical Strategy

To analyze how exposure to the EITC in childhood affects childbearing and marriage in adulthood, we calculate the maximum federal and state EITC a child's family could receive given the number of children in the household, the state of residence, and the year. We focus on the maximum federal and state EITC available in each year, rather than a family's own EITC, because the latter is endogenous to family characteristics that are also typically associated with fertility and marriage patterns. To be eligible for the EITC in the 2017 tax year, earnings must be less than $53,930 for a married family with three children, meaning that EITC-eligible families are negatively selected. Children growing up in lower income households also typically have births at earlier ages than children growing up in more affluent households, which suggests that a naïve regression of fertility on own EITC benefits received throughout childhood would produce a positive relationship—children exposed to more EITC benefits would be more likely to have a first child at an early age compared to children exposed to lower benefits. Using the maximum available EITC in a given year removes this negative selection into EITC eligibility and allows us to estimate a plausibly causal relationship between EITC exposure and family formation. We focus on a reduced-form, intent-to-treat analysis for our main specification, and in robustness checks we illustrate that effects are largest among subgroups most likely to be affected by the EITC.

To estimate the effect of EITC exposure on family formation, we estimate models of the following form:

$$Y_{ist} = \beta_0 + \beta_1 maxEITC_{ist} + \beta_2 X_{ist} + \beta_3 Z_{st+17} + \alpha_s + \gamma_t + \varepsilon_{ist}$$

where Y_{ist} is the outcome of interest—whether the individual has had a first birth by a given age, whether the individual ever married by a given age, and whether the individual was unmarried when she had the first birth or had been married, ranging from age 18 to 25. Each outcome is evaluated for a given individual i, born in state s, in year t. X_{ist} is a vector of demographic controls including race, parental education, average number of children residing in the household throughout childhood, and share of childhood spent with married parents. Z_{st} is a vector of state contextual variables evaluated in the year the individual turned 17, including the minimum wage, the unemployment rate, and the maximum welfare benefits available for a family of three. We also include state-of-birth-specific time trends to control for other changes in the state over time that may be correlated with EITC exposure and influence family formation. Additionally, we include state-of-birth and year-of-birth fixed effects. State-of-birth fixed effects control for other state-level factors that may be correlated with EITC exposure and family formation, such as political ideology. Year-of-birth fixed effects control for cohort differences in family formation patterns at the national level. Standard errors are clustered at the state level.

Our coefficient of interest is β_1, which represents the change in the outcome of interest among women exposed to a $1,000 increase in maximum potential EITC benefits throughout childhood (all dollars in 2016$). Between our cohorts born in 1968 and 1992, there is more than an $80,000 difference in maximum potential EITC benefits available between birth and age 15, generating substantial variation in exposure to the EITC throughout childhood (see Appendix Figure 1).

We present results for men and women, as well as separately for white and black women. After presenting the main results, we also show sensitivity analyses for different cohorts within our sample, for different age ranges of EITC exposure in childhood, and when making income restrictions to the sample. We then test for different mechanisms that could potentially explain our findings, examining how the EITC affects educational attainment.

Results

Figure 1 shows results from our main regression analysis—the relationship between EITC exposure in childhood and fertility and marriage in adulthood for women. Each point is estimated from a separate regression and represents the change in the outcome of interest (fertility, marriage) at each age point following a $1,000 increase in EITC exposure between birth and age 15. Ninety-five percent confidence intervals are represented by the vertical lines. Panel A depicts results for fertility, while panel B shows results for marriage patterns. All point estimates and standard errors for these models are available in Appendix Table 2.

Results indicate a negative relationship between EITC exposure in childhood and early childbearing and marriage, particularly between the ages of 21 and 24. Increasing EITC exposure in childhood by $1,000 implies a reduction in the likelihood of giving birth by age 20 of 0.6 percentage points, a 2 percent overall reduction in teenage fertility in this sample. Point estimates remain negative and significant up through age 25, but become slightly smaller in magnitude, which implies that women exposed to larger EITC benefits in childhood may "catch up" to their peers exposed to smaller benefits.

Turning to the results for marriage (Panel B), we also find a slight negative relationship between EITC exposure in childhood and marriage, particularly between ages 22 and 24. A $1,000 increase in EITC exposure in childhood leads to a 0.5 percentage point reduction in the likelihood of marrying by age 22, which represents a 3 percent reduction in marriage at that age.

Together, these results indicate that EITC exposure in childhood leads women to postpone marriage and fertility in early adulthood, particularly between ages 20 and 24.

Our youngest cohort, those born in 1992, turn 25 in the last interview year available from the PSID, so it is not possible to observe trends in birth and marriage after age 25 for our full sample, but in Appendix Figure 2, we show results for the subset of our sample we can observe up to age 30, those born between 1968 and 1987. While estimates are noisy, we do see some suggestive evidence that the fertility and marriage results are indicative of postponing family formation, rather than forgoing marriage and childbearing altogether. Point estimates on childbearing become smaller in magnitude and indistinguishable from zero beginning at age 26, though coefficients remain negative through age 30. For marriage, point estimates are insignificant beginning at age 23 for those born from 1968-1987, and coefficients turn slightly positive beginning at age 29.

Figure 2 presents figures analogous to those presented in Figure 1, but for the men in our sample.[8] We see very little evidence that EITC exposure in childhood affects men's fertility or marriage in early adulthood. Point estimates on the likelihood of having a birth are slightly negative, and become increasingly negative between ages 22 and 25, but never attain statistical significance. For marriage, we find very little evidence of a statistically significant relationship between EITC exposure in childhood and marriage in early adulthood, although there is a steady positive trend in the point estimates between ages 22 and 25. The lack of a relationship between EITC generosity and marriage and fertility among men is not entirely surprising since men tend to underreport births. They also tend to partner with younger women so there may be some

[8] Point estimates and standard errors are presented in Appendix Table 3.

evidence of a change at older ages.[9] Given the lack of findings for men, for the remainder of the analyses we focus on the results for women.

All of the point estimates discussed thus far come from regressions that include demographic and state contextual variables, state of birth and year of birth fixed effects, and state of birth-specific linear time trends. We test the robustness of these estimates to the inclusion of each of these controls, which we present in Table 2. For simplicity, we present results for birth by age 21, one of our most precise estimates. Column 1 presents results of a bivariate regression of birth by age 21 on EITC exposure from birth to age 15, and indicates a slight, negative relationship. A $1,000 increase in EITC exposure in childhood is associated with a 0.2 percentage point decline in the likelihood of having a birth by age 21. Adding demographic and state contextual variables reduces the magnitude of the result to a 0.1 percentage point decline in fertility, but the estimate remains significant at the $p<.05$ level. Year of birth fixed effects have a substantial effect on the point estimate, which is not surprising given the correlation between year of birth, fertility, and EITC exposure. Not controlling for year of birth fixed effects biased our estimates towards zero; with their inclusion we find a 0.4 percentage point reduction in the likelihood of having a birth by age 21 following a $1,000 increase in EITC exposure in childhood. Adding state of birth fixed effects does little to change the point estimate, but state-of-birth-specific time trends increases the magnitude of the effect to a 0.8 percentage point reduction in early childbearing.

[9] For the subset of men whom we can observe to age 30 (those born 1968-1987), we find a positive trend in marriage as a function of EITC exposure in childhood, see Appendix Figure 3. Point estimates, however, are noisy, and only significant at the $p<.05$ level at age 27.

How do results vary by family income in childhood?

While we include all individuals born between 1968 and 1992 in our main analysis, not all of these individuals would have received the EITC in childhood. To test whether our results are driven by individuals likely to have received the EITC in childhood, we conduct a robustness check where we re-estimate our results using different income cut-offs. Figure 3 presents results of progressively reducing the upper income cut off from no upper bound (main estimates), to restricting the sample by reducing the income cut off in $10,000 increments, down to individuals growing up with an average family income of $20,000 or less (2016$). Income in childhood is calculated based on the earnings of the head and spouse (if present), and averaged between an individual's birth and age 15.[10] Each point in Figure 3 is estimated from a separate regression and 95 percent confidence intervals are indicated by the vertical bars. We present results for birth by age 21, as this was one of our most precise estimates.

Figure 3 illustrates that reducing the upper income threshold results in a larger reduction in fertility than when no income restriction is imposed. With no restriction, for instance, we find that a $1,000 increase in EITC exposure in childhood leads to a 0.8 percentage point reduction in the likelihood of having a first birth by age 21. Limiting the income threshold to those whose family income was below $50,000 on average between birth and age 15 reduces the likelihood of having a birth by age 21 by 1.1 percentage point. Restricting the sample to those whose family income averaged less than $20,000 yields a reduction in the likelihood of having a first birth by nearly 2 percentage points. None of the estimates is statistically different from each other, but this exercise provides assurance that the reductions in fertility are stemming from individuals

[10] For individuals lacking family income in every year between birth and age 15, we average family income in the years that the individual was present in the PSID.

likely to have received the EITC at some point throughout childhood, and are not driven by individuals growing up in high-income households.

Sensitivity analyses: variation by birth cohort; varying EITC exposure window

We conduct a number of different sensitivity analyses to test the robustness of our results. First, we limit the sample to those born between 1975 and 1992, as those born after 1974 would have had at least some EITC exposure in every year between birth and age 15; those born before 1975 would have zero EITC exposure in at least one year. We also further limit the sample to those born between 1980 and 1992 to test whether our results are robust to restricting the sample to a more recent birth cohort. Results of this exercise are presented in Figure 4, and suggest somewhat similar patterns in fertility and marriage postponement across the three different sample restrictions. One notable exception is that for the more recent birth cohorts, we find stronger evidence that women exposed to larger EITC benefits in childhood "catch up" to their peers who were exposed to smaller benefits in childhood in terms of their fertility by age 25. For the full sample of women born 1968-1992, we continue to find a marginally-significant decline in fertility up through age 25, of about 0.5 percentage points per $1,000 increase in EITC exposure. Limiting the sample to those born 1975-1992 reduces that coefficient to 0.3 percentage points (not significant), and limiting the sample further to those born between 1980-1992, we find an insignificant 0.1 percentage point reduction in fertility by age 25. Patterns for marriage, on the other hand, are very consistent across the three different samples. For all three samples, we find reductions in marriage on the order of about 0.5 percentage points at its strongest point, around age 22, with little evidence of a reversing of this pattern by age 25.

Finally, we test whether results are robust to using different age ranges for the EITC exposure measure. For simplicity, we focus on how the estimates change for likelihood of having a birth by age 21, testing how estimates vary when we limit the age range of exposure to birth to age 14, birth to age 13, and so on to birth to age 5. Results are presented in Figure 5. Limiting the age range for EITC exposure to a narrower window results in somewhat larger reduction in fertility by age 21, though estimates are less precise. In our main estimates, we find that a $1,000 increase in EITC exposure between birth and age 15 results in a 0.8 percentage point reduction in fertility by age 21. Restricting the range of exposure to birth to age five, we find a 1.8 percentage point reduction in fertility by age 21, more than double the point estimate when including birth to age 15 exposure. A word of caution in interpreting these estimates—when we restrict the age range of exposure to birth to age 5, some individuals in our main sample had zero exposure to the EITC in that age range, since the EITC was not established until 1975. This results in a skewed distribution of EITC exposure from birth to age five, which may account for the much larger effects. Still, this exercise reveals that results are not particularly sensitive to the age range of EITC exposure; we find a very consistent pattern of reductions in fertility by age 21 as a function of EITC exposure in childhood.

Subgroup analyses

We next examine how our results vary by race, to analyze which groups are most affected by changes in EITC exposure. We present results separately for black and white women—Figure 6 presents results for fertility, while Figure 7 presents results for marriage. Exposure to the EITC in childhood leads to lower fertility for both black and white women between the ages of roughly 20 and 22, by between 0.5 and 1 percentage point. Since black women have higher fertility than white women, in percentage terms, these effect sizes are very similar between black and white women.

After age 22, the effects of the EITC on fertility differ for black and white women. White women exposed to larger EITC benefits in childhood remain significantly less likely to have a birth up through age 24, and the point estimate on age 25 is similar to that on age 24, though less precisely estimated. This suggests that, at least until age 25, white women do not appear to “catch up” to their peers exposed to lower EITC benefits in childhood in terms of their fertility. On the other hand, black women exposed to larger EITC benefits in childhood do appear to “catch up” to their peers who were exposed to smaller EITC benefits in childhood. Following a $1,000 increase in EITC exposure in childhood, black women are significantly less likely to have a birth up through age 22, but are no less likely to have a birth by ages 23 through 25. In fact, point estimates are positive, though not statistically significant, for ages 24 and 25.

Turning to the results for the relationship between EITC exposure in childhood and marriage, we find very distinct patterns for black and white women (see figure 7). For white women, the relationship between EITC generosity and marriage parallels that of the fertility effects: estimates are close to zero for ages 18 through 20, but turn negative and significant starting at age 22. A $1,000 increase in EITC exposure in childhood reduces the likelihood of a white woman marrying by age 22 by 0.7 percentage points, which is identical to the point estimate on the likelihood of having a birth by age 22. This pattern persists up through age 24, with some slight evidence of a reversal of the relationship between EITC generosity and marriage starting at age 25. More years of data are required to determine if these women “catch up” to their peers after age 25, or whether this is a permanent reduction in the likelihood of marrying. Black women, on the other hand, appear unaffected by the EITC in terms of their marriage rates. There is no clear relationship between the EITC and propensity to marry between ages 18 and 25, with some negative coefficients and some positive coefficients and no point estimates larger than 0.1 percentage point in absolute value, which is somewhat consistent with the decoupling of marriage and childbearing for Black and low-education women (Gibson-Davis 2011; Lundberg et al. 2016).

We also investigate potential EITC benefits and non-marital and marital births by race for women. In the first panel of Figure 8, we see no relationship between EITC exposure in childhood and non-marital births for white women. From age 18 through age 25, the point estimates are sometimes positive, sometimes negative, but rarely significant. For African American women, on the other hand, we observe some delay during the teenage years and in the early twenties, although the point estimates are only significant at conventional levels at age 21 and age 22. By age 23, the delays in non-marital fertility appear to level off, and we even see some positive, although insignificant point estimates in the mid-twenties.

Figure 9 displays results for births that occurred subsequent to first marriage. Here the findings are quite different. First, we see large and increasing delays in fertility among white women in "marital births." In fact, by age 25, the delays appear to be increasing rather than leveling off. We see almost no delay in marital fertility for African American women.

Mechanisms: Educational Attainment

What explains the reduction in early childbearing associated with EITC exposure in childhood? Previous research has linked the EITC with reduced infant birth weight (Hoynes et al 2013), higher test scores in childhood (Dahl and Lochner 2012; 2017), and higher educational attainment and earnings in adulthood (Bastian and Michelmore 2018). These pathways could explain why the EITC also reduces early childbearing among young adult women. Education tends to have a negative effect on fertility, as women typically avoid having children until after completing school. Conversely, an unexpected birth is also likely to reduce future educational prospects. We test for the human capital hypothesis here, by examining whether EITC exposure from birth to age 15 leads to higher educational attainment in adulthood. Table 3 presents results of this exercise.

We examine three different outcomes related to educational attainment: an indicator for whether the individual completed high school by age 19, an indicator for whether the individual completed college by 23, and a continuous term representing the total number of years of schooling completed by age 25. For all women (panel A), we do not find much evidence that the EITC increases the likelihood of completing high school by age 19, but we do find evidence that the EITC increases the likelihood of completing a college degree by age 23. This is consistent with previous research (Bastian and Michelmore 2018; Manoli and Turner 2018), and implies that a $1,000 increase in EITC exposure in childhood increases the likelihood of completing college by about 1 percentage point. Only about 20 percent of our sample completes a college degree by this age, so this represents a 5 percent increase in college attainment. Commiserate with the increase in the likelihood of completing college, we also find evidence that the EITC increases the total number of years of schooling by age 25 by about 0.03 years. These effects appear to be concentrated mostly among white women; we find no significant increase in educational attainment among black women.

These results are consistent with a human capital theory perspective, which suggests that increases in educational attainment increase the opportunity costs of childbearing, leading to lower fertility rates among young women exposed to the EITC. Since we do not find much evidence of a reduction in fertility before age 19, it seems plausible that the fertility reduction effect is driven by an increase in educational attainment rather than the reverse (that reducing teen fertility improves individuals' prospects for higher education).

Cost savings analysis

To calculate the potential cost savings from social welfare programs due to the reductions in early childbearing associated with EITC exposure in childhood, we use a simple calculation based on the approximate number of births delayed for women between the ages of 20 and 24 and an approximation of the annual cost per recipient from the EITC program, the SNAP program, and the Medicaid program. Obviously, there are other costs savings, but this, admittedly, crude analysis should provide a lower bound estimate.

Appendix Table 2 provides marginal effect estimates for a $1,000 increase in the lifetime EITC benefit for PSID respondents for the birth outcome by age 18 through age 25. In the first panel, among the statistically significant estimates, the magnitude of the effect size ranges between 1.3 (-.006/.447 for all 24-year-old women) percent and 3.0 (-.006/.201 for 21-year-old white women) percent with the median value being 2.2 (-.006/.267 for 20-year-old women) percent. For our calculations, we use the median value for a standard deviation change in EITC benefits (approximately $21,013) and a somewhat smaller change, $10,000. Therefore, a $10,000 increase in EITC benefits would reduce the number of births by 22 percent or (as calculated below) approximately 12,252 births, and a $21,000 increase in EITC benefits would reduce the number of births by 46.2 percent or approximately 25,730 births. As a frame of reference, births to 15-19 year olds declined by approximately 70 percent over this time period (between 1990 and 2017), and by 40 percent among 20-24 year olds (Child Trends n.d.).

Next, we use several sources for the fertility change and benefit savings estimates. According to the National Center for Health Statistics (Martin et al. 2018), 3.69 percent of women between the ages of 20 and 24 had a first birth in 2017. The Internal Revenue Service's website reports that 27 million eligible workers gained around $65 billion in EITC benefits in 2017 or an average of $2,445 per recipient.[11] Data from the U.S. Census Bureau American Fact Finder, shows that there were 10,769,493 women aged 20-24 in 2017.[12] Among all men and women aged 20 through 64, a group we will consider the working population eligible, 5.59 percent are women aged 20 to 24. Using that ratio, we assume that 5.59 percent of the 27 million eligible EITC workers, or 1,509,300, are females in that age range. We next assume that 3.69 percent, or 55,693, had a first birth in 2017. This estimate serves as the upper bound estimate of births delayed by the EITC investment. Twenty-two percent of this number is 12,252 births, and 46.2 percent of this number is 25,730 births.

We will use the EITC average benefit reported earlier as an estimate of the cost savings from that program. The United State Department of Agriculture reports that the average monthly SNAP benefit per person in 2017 was $126.21 or $1,515 annually.[13] To generate a per person Medicaid benefit, the Congressional Budget Office frequently values the benefit at the mean expenditure per recipient. However, we follow the methodology of Finkelstein, Hendren, and Luttmer (forthcoming), which uses the Oregon Health Insurance Experiment asking if the CBO estimate is too high. Their models suggest values that fall between 0.15 and 0.85 of the mean expenditure. We chose the center of this range, half of the mean expenditure, for our calculations. In fiscal year 2016, Medicaid spending was $553 billion and covered about 74

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[11] https://www.eitc.irs.gov/eitc-central/statistics-for-tax-returns-with-eitc/statistics-for-2017-tax-returns-with-eitc
[12] https://factfinder.census.gov/faces/tableservices/jsf/pages/productview.xhtml?src=bkmk
[13] https://fns-prod.azureedge.net/sites/default/files/SNAPsummary.pdf. Accessed 4/8/2019.

million Americans (Rudowitz and Valentine 2017). Given these estimates, we value the cost saving of foregone births at 50 percent of $7,473 or $3,736.

Given these changes, we next calculate the EITC savings by multiplying these birth delay figures by $2,445, the SNAP savings by multiplying the birth delay figures by $1,515, and for the Medicaid savings, by multiplying the birth delays figures by $3,736. We report results in Table 4. Depending on the change in the EITC benefit used, EITC savings range between $30 and $63 million annually, SNAP benefits savings range between $19 and $40 million annually, and Medicaid savings range between $46 and $96 million annually. Collectively, then, from these programs alone, these delayed births reduce social welfare costs somewhere between $94 and $200 million annually.

Conclusion

The EITC has become one of the largest cash transfer programs in the United States, and one of the key social safety net programs for low-income families. Previous work on the EITC suggests that the program has improved the human capital accumulation of children who grow up receiving benefits of the program (Bastian and Michelmore 2018; Manoli and Turner 2018). We hypothesized that the EITC may reduce the likelihood of marrying and having children in early adulthood, as prior work has shown the EITC to have a positive effect on educational attainment (Bastian and Michelmore 2018), which typically raises the opportunity cost of childbearing.

We investigate this question by analyzing whether exposure to the EITC in childhood has an effect on marriage and childbearing in early adulthood, from age 18 to age 25. We measure exposure to the EITC in childhood as the sum of the maximum potential federal and state EITCs one could claim given the state, year, and number of children in the household for all years between birth and age 15.

Since its inception in 1975, the EITC has grown from a 10% wage subsidy to a 45% wage subsidy for families with at least three children, and larger in states that offer their own EITCs. In addition to the family formation question, this project also asks whether receiving the EITC as a child affects one's own likelihood of receiving the EITC in adulthood, as has been shown with programs such as welfare (Gottschalk 1996; Pepper 2000). One route through which individuals become eligible for the EITC is by having children at an early age: the EITC is not available to childless adults until age 25, but those with children can claim the EITC at any age, so long as another individual cannot claim them as a dependent on their tax returns. Therefore, reducing teen and early 20s fertility could have a significant impact on the number of families claiming the EITC each year.

Our results suggest that exposure to the EITC in childhood significantly reduces the likelihood of having a birth and marrying among women between ages 20 and 24. The evidence suggests that it is marital births among white women (point estimates range between -0.6 to -0.7 percentage points) and non-marital births among African American women (point estimates range between -0.9 and 1.3 percentage points) that are delayed. We find no effect of EITC exposure in childhood on men's marriage and fertility in early adulthood.

Coupled with this reduction in fertility, we also corroborate previous research in finding that the EITC increases the likelihood of completing a college degree, providing a plausible mechanism through which the EITC reduces early family formation. These reductions in early childbearing likely represent fertility postponement rather than a decline in overall fertility, as we find evidence that women exposed to larger EITC benefits in childhood are no less likely to have had a birth by age 30 than their peers who were exposed to smaller EITC benefits in childhood. Further analysis is needed to determine whether there are quantum effects as well—that is, whether the EITC reduces completed fertility.

These results have implications for the well-being of young women growing up with the EITC, as early childbearing is associated with a host of negative outcomes for both mother and child, e.g., decreased educational attainment, poor labor market outcomes, increased likelihood of social welfare program receipt, and poor health (Haveman, Wolfe, and Peterson, 1997; Hoffman 2006; Mirowsky 2005). From a budgetary perspective, these results also have implications for federal and state spending on the EITC. Each birth averted before age 25 reduces the number of individuals claiming the EITC, since childless individuals are not eligible for the credit until turning 25. Our back of the envelope cost estimates suggest that a one-standard deviation increase in average EITC generosity between birth and age 15 reduces annual EITC spending by $63 million. Factoring in costs savings from other social welfare programs such as food stamps and Medicaid, and annual costs savings increase to about $200 million annually (2016$). Together, these results further suggest that the EITC is a cost-effective program by improving human capital and reducing the incidence of early childbearing among children exposed to the credit.

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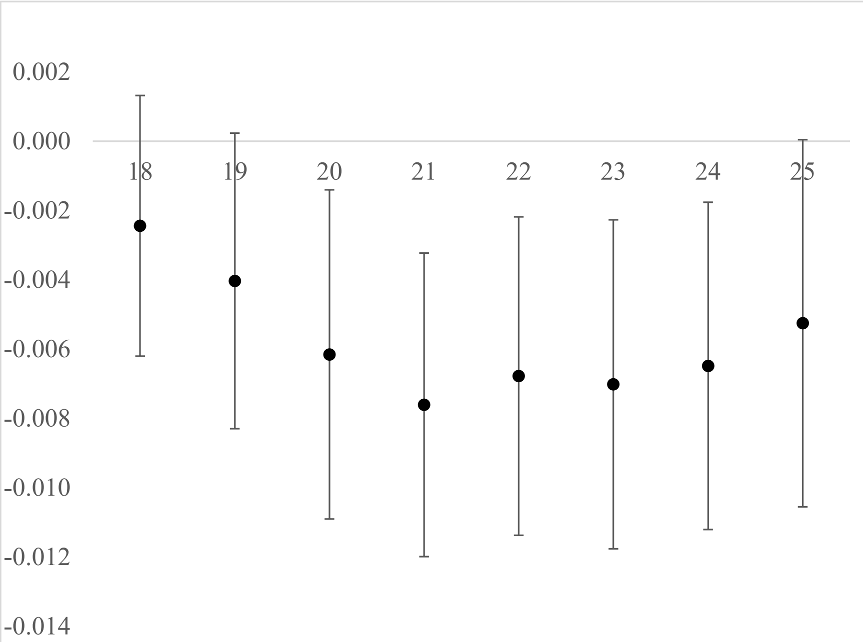
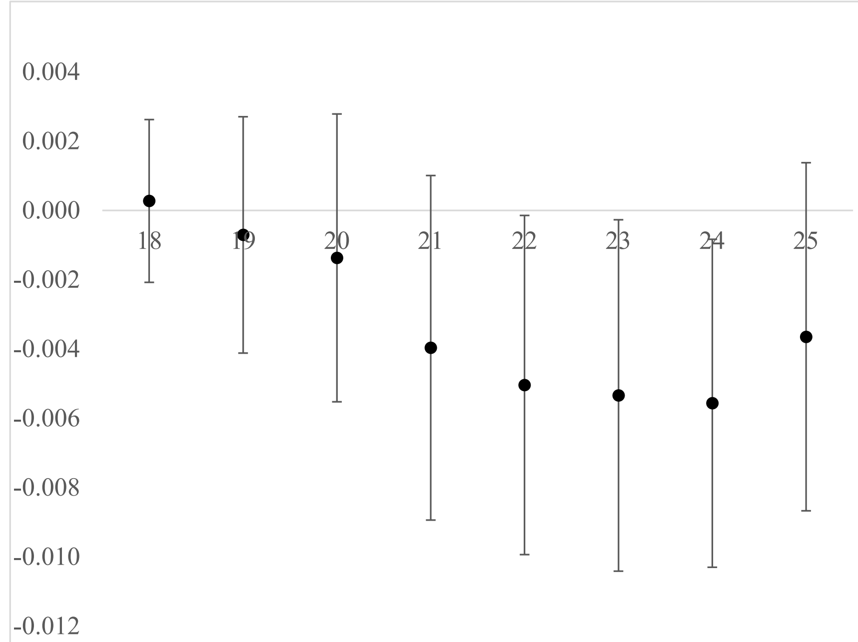
Figure 1. Effect of EITC exposure (in $1,000s of 2016$) from birth to age 15 on marriage and fertility outcomes in early adulthood for women, with 95% confidence intervals



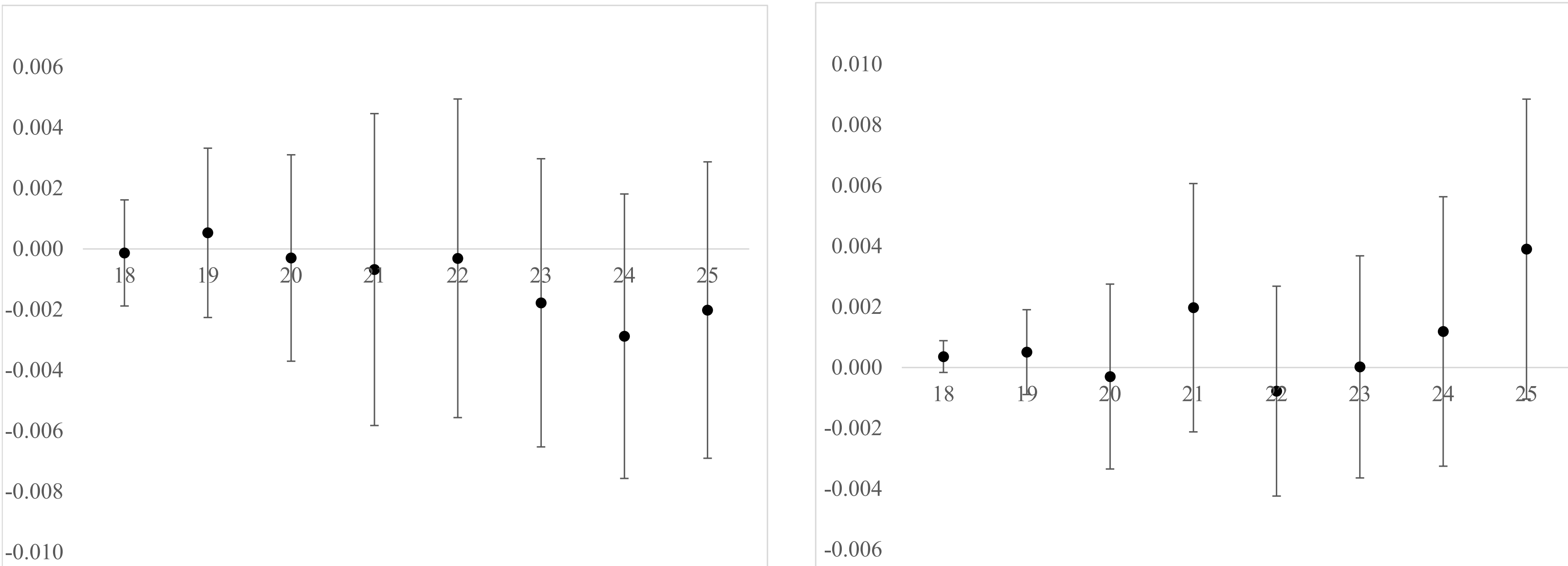
a) Birth by age



b) Marriage by age

Source: Panel Study of Income Dynamics 1968-2017. Women born 1967-1992. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (race, education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends. 95% confidence intervals indicated by vertical bars. Each point represents a separate regression.

Figure 2. Effect of EITC exposure (in $1,000s of 2016$) from birth to age 15 on marriage and fertility outcomes in early adulthood for men, with 95% confidence intervals

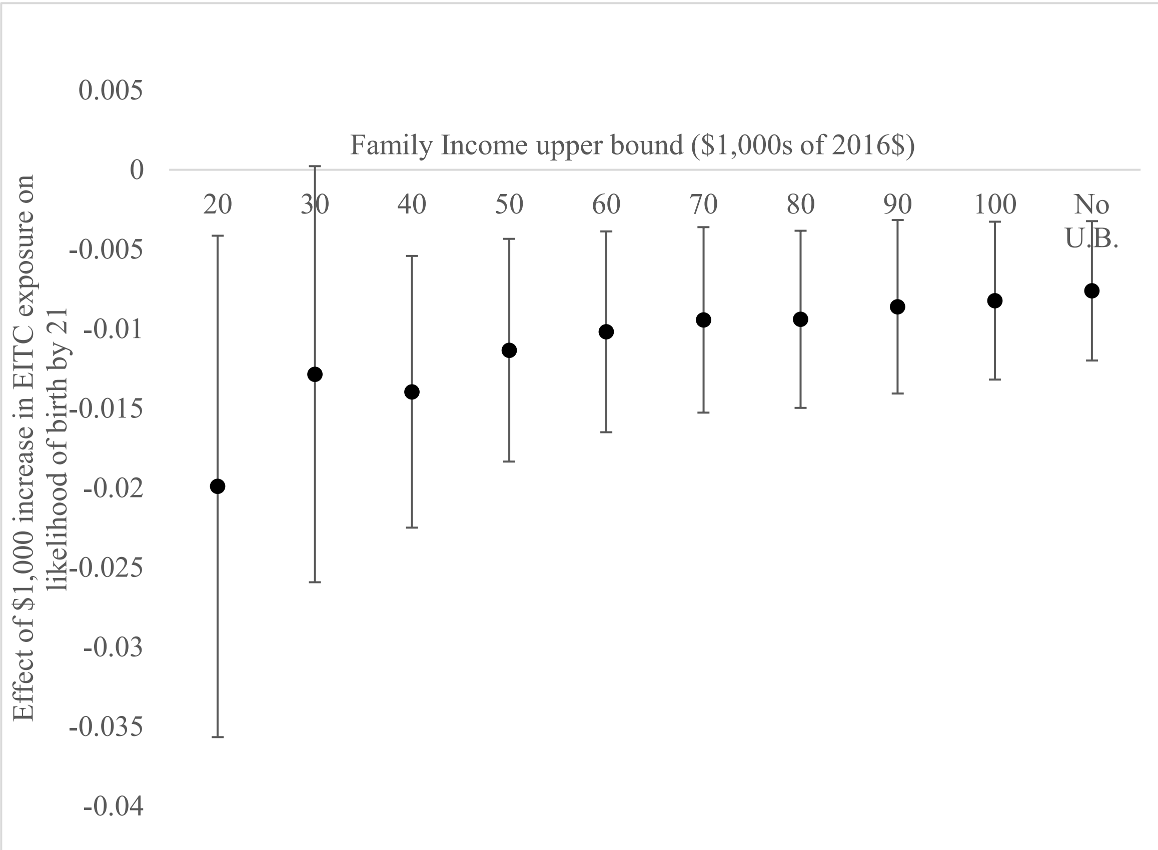


a) Birth by age

b) Marriage by age

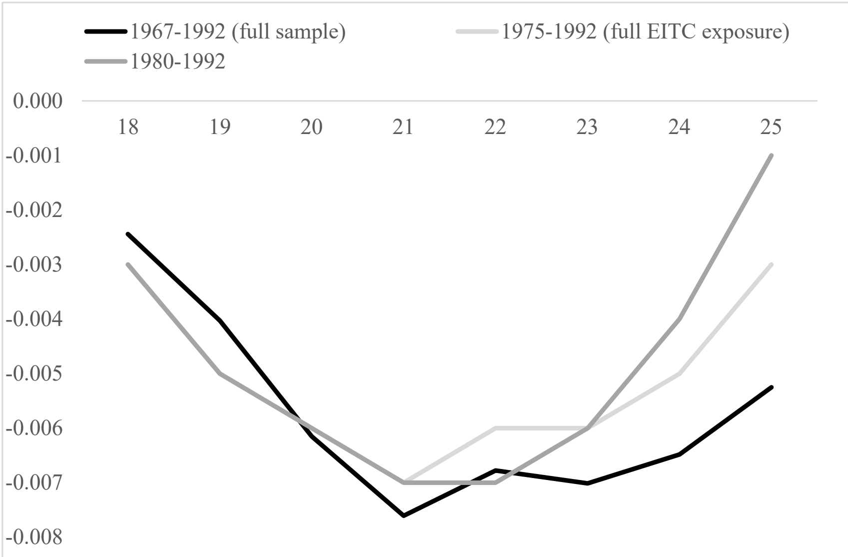
Source: Panel Study of Income Dynamics 1968-2017. Men born 1967-1992. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (race, education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends. 95% confidence intervals indicated by vertical bars. Each point represents a separate regression.

Figure 3. Effect of EITC exposure in childhood on birth by age 21, varying upper income cut-off

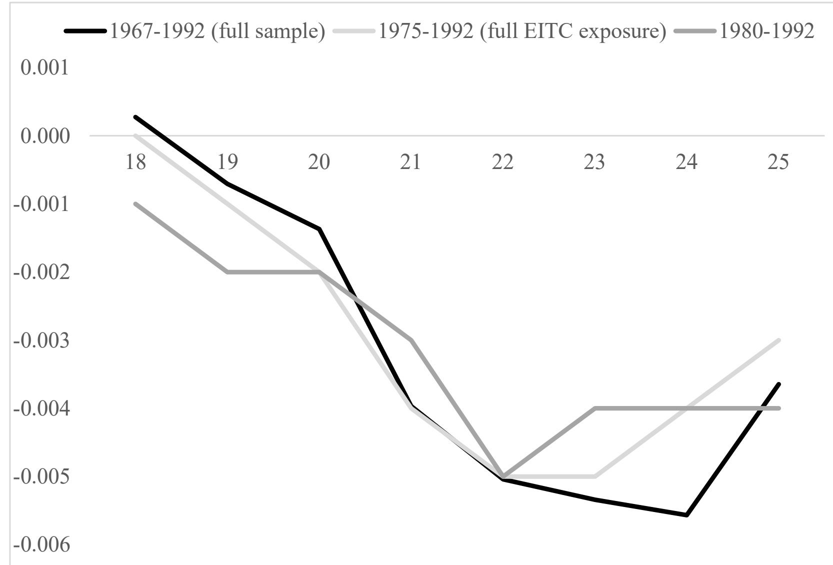


Source: Panel Study of Income Dynamics 1968-2017. Women born 1967-1992. Note: Each point is estimated from a separate regression of birth by age 21 on EITC exposure from birth to age 15, with income upper bound specified along x-axis. All models include controls for demographic characteristics (race, years of schooling of head, number of children residing in household over childhood), state-year controls (welfare generosity, unemployment rate, and minimum wage, all evaluated at age 15), year of birth fixed effects, state fixed effects, and state-year of birth time trends. Income upper bound subjected to average head+wife income observed between birth and age 15.

Figure 4. Effect of EITC exposure in childhood on marriage and fertility; varying birth cohort window



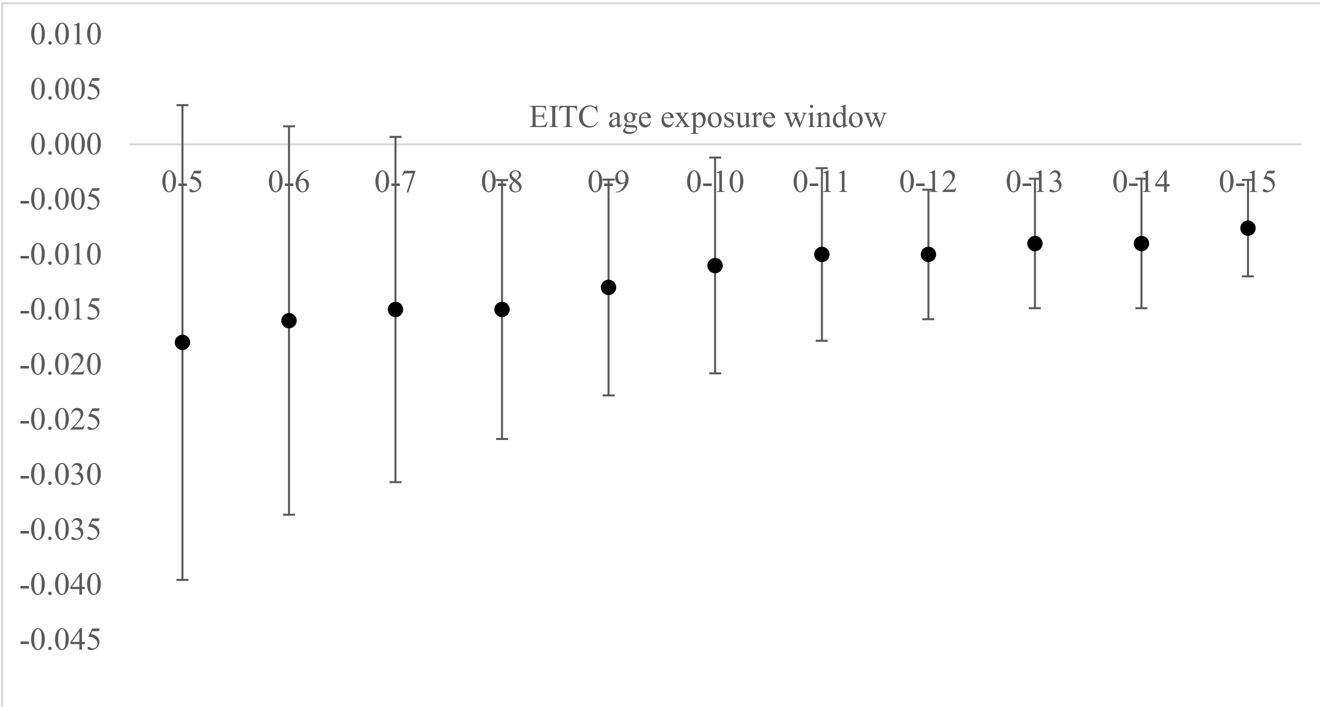
a) Birth by age



b) Marriage by age

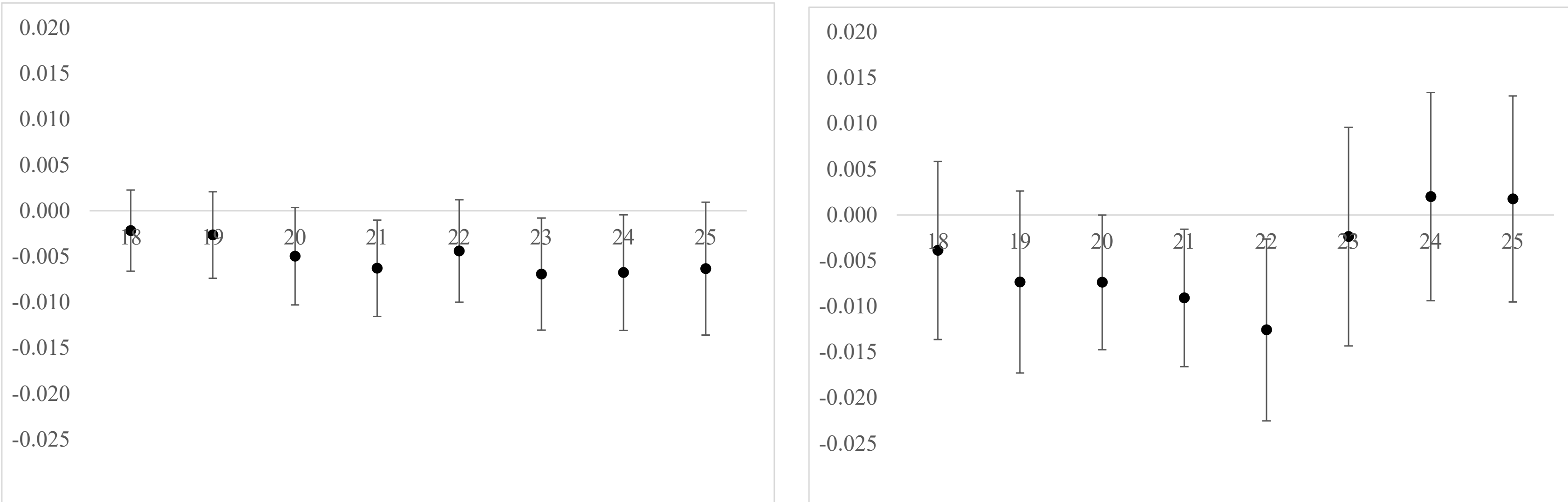
Source: Panel Study of Income Dynamics 1968-2017. Women born 1967-1992. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (race, education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends. 95% confidence intervals indicated by vertical bars. Each point represents a separate regression.

Figure 5. Effect of EITC exposure on likelihood of having a first birth by age 21; varying EITC window of exposure



Source: Panel Study of Income Dynamics 1968-2017. Women born 1967-1992. Note: Each point is estimated from a separate regression of birth by age 21 on EITC exposure, with EITC exposure window specified along x-axis. All models include controls for demographic characteristics (race, years of schooling of head, number of children residing in household over childhood), state-year controls (welfare generosity, unemployment rate, and minimum wage, all evaluated at age 15), year of birth fixed effects, state fixed effects, and state-year of birth time trends. Income upper bound subjected to average head+wife income observed between birth and age 15.

Figure 6. Effect of EITC exposure (in $1,000s of 2016$) from birth to age 15 on fertility outcomes in early adulthood for black and white women, with 95% confidence intervals

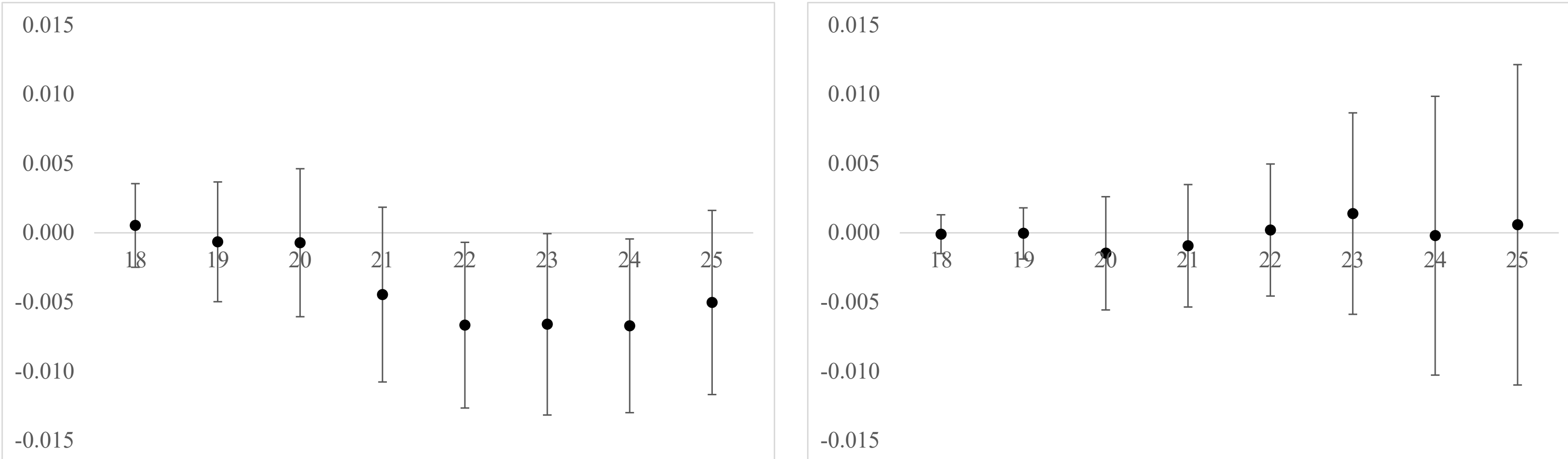


a) White women

b) Black women

Source: Panel Study of Income Dynamics 1968-2017. Women born 1967-1992. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends. 95% confidence intervals indicated by vertical bars. Each point represents a separate regression.

Figure 7. Effect of EITC exposure (in $1,000s of 2016$) from birth to age 15 on marriage outcomes in early adulthood for black and white women, with 95% confidence intervals

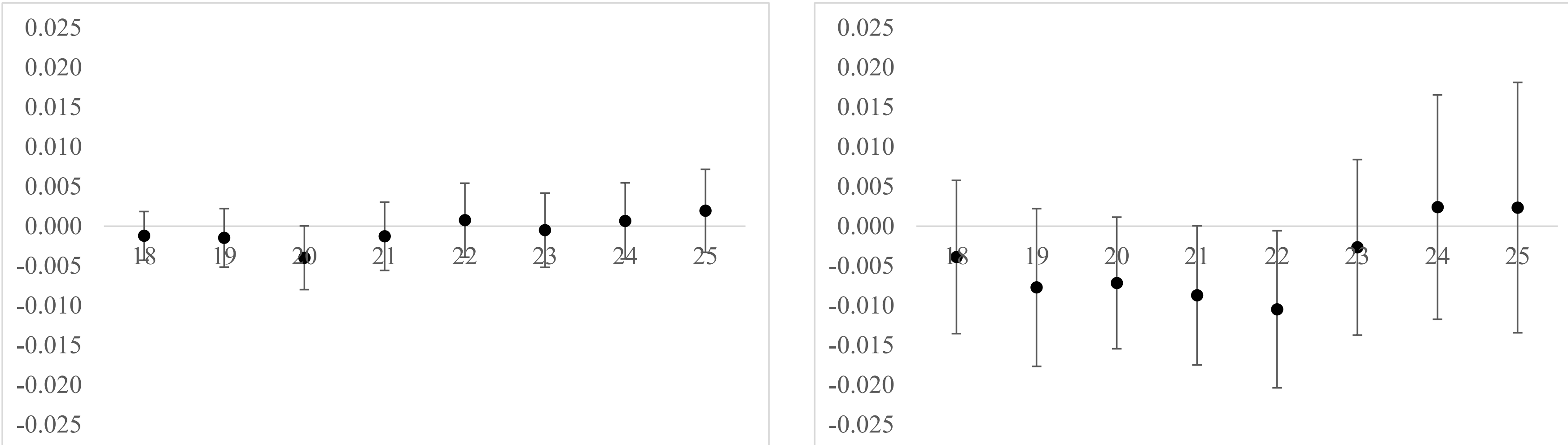


a) White women

b) Black women

Source: Panel Study of Income Dynamics 1968-2017. Women born 1967-1992. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends. 95% confidence intervals indicated by vertical bars. Each point represents a separate regression.

Figure 8. The effect of a $1,000 increase in EITC exposure from birth to age 15 on non-marital births, by race

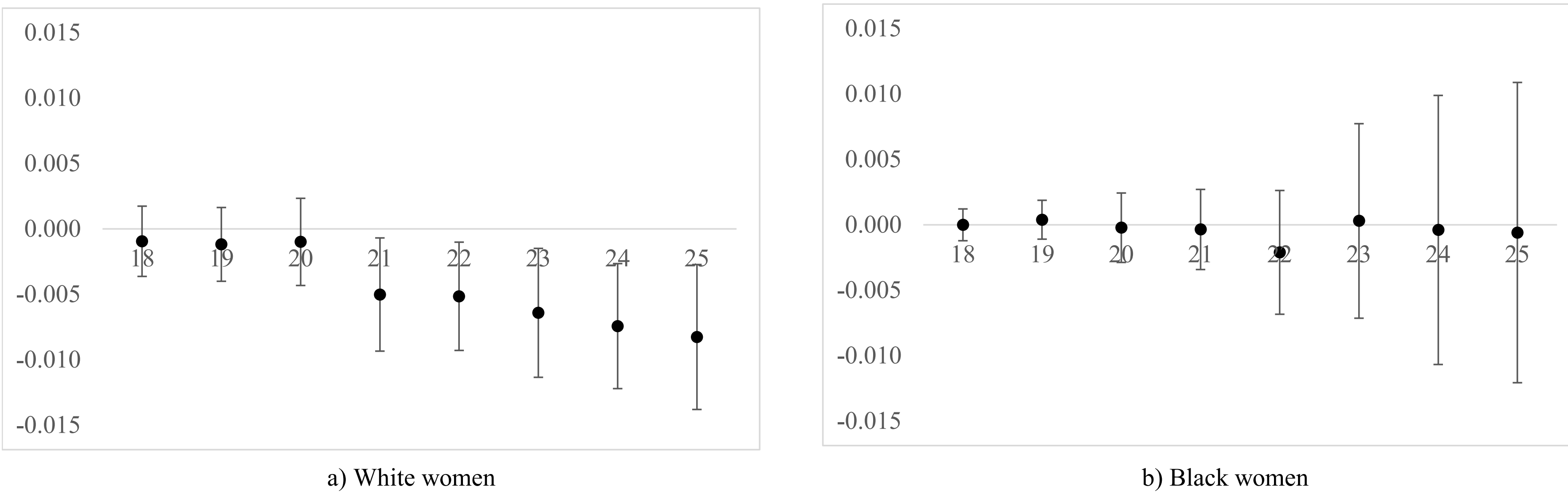


a) White women

b) Black women

Source: Panel Study of Income Dynamics 1968-2017. Women born 1967-1992. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends. 95% confidence intervals indicated by vertical bars. Each point represents a separate regression.

Figure 9. The effect of a $1,000 increase in EITC exposure from birth to age 15 on marital births, by race



Source: Panel Study of Income Dynamics 1968-2017. Women born 1967-1992. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends. 95% confidence intervals indicated by vertical bars. Each point represents a separate regression.

Table 1: Descriptive Statistics

	Full Sample	Women	Men
Max EITC benefit ages 0-15 (in 1000s of 2016 dollars)	38.058	37.994	38.121
	(21.013)	(20.961)	(21.066)
Family background characteristics			
Head's education	12.64	12.618	12.662
	(2.675)	(2.79)	(2.558)
African American	0.181	0.167	0.195
White	0.797	0.811	0.782
Hispanic	0.017	0.015	0.018
Other	0.006	0.007	0.004
Proportion of years 0-15 head married	0.792	0.783	0.802
Proportion of years 0-15 one child in household	0.062	0.061	0.062
Proportion of years 0-15 two children in household	0.423	0.413	0.433
Proportion of years 0-15 three children in household	0.345	0.349	0.342
Proportion of years 0-15 four children in household	0.129	0.132	0.126
Proportion of years 0-15 five or more children in household	0.041	0.045	0.037
State-year contextual variables			
Maximum welfare benefit for a family of three (in 100s of 2016 dollars)	5.916	5.914	5.917
	(2.382)	(2.341)	(2.421)
State unemployment rate	5.674	5.67	5.679
	(1.677)	(1.667)	(1.688)
State minimum hourly wage (in 2016 dollars)	6.94	6.926	6.955
	(1.006)	(1.012)	(1.)
Marriage and fertility			
Birth by age 18	0.063	0.096	0.03
Birth by age 19	0.104	0.149	0.059
Birth by age 20	0.147	0.204	0.09
Birth by age 21	0.188	0.25	0.127
Birth by age 22	0.22	0.285	0.154
Birth by age 23	0.259	0.323	0.194
Birth by age 24	0.305	0.37	0.238
Birth by age 25	0.341	0.407	0.272
Married by age 18	0.022	0.036	0.009
Married by age 19	0.047	0.072	0.023
Married by age 20	0.082	0.114	0.051
Married by age 21	0.123	0.163	0.082
Married by age 22	0.172	0.214	0.129
Married by age 23	0.226	0.273	0.179
Married by age 24	0.282	0.333	0.231
Married by age 25	0.345	0.396	0.292
Education			
High school completion by age 19	0.628	0.672	0.586
College completion by age 23	0.171	0.214	0.127
Years of Education by age 25	13.537	13.753	13.321
	(2.033)	(2.061)	(1.982)
Number of Observations	6,622	3,301	3,321

Source: Panel Study of Income Dynamics 1967-2017. Men and women born 1967-1992. All estimates weighted by average of person weights from birth to age 18.

Table 2. Effect of EITC exposure from birth to age 15 (in 1,000s of 2016 dollars) on childbearing by age 21

	(1)	(2)	(3)	(4)	(5)	(6)
EITC exposure birth to age 15	-0.002	-0.001	-0.001	-0.004	-0.004	-0.008
	(0.0005)	(0.0005)	(0.0005)	(0.002)	(0.002)	(0.002)
Demographic controls		X	X	X	X	X
State-year controls			X	X	X	X
Year of birth fixed effects				X	X	X
State of birth fixed effects					X	X
State specific-year of birth time trends						X
Number of Observations	3063	3063	3063	3063	3063	3063

Source: Panel Study of Income Dynamics 1967-2017. Women born 1967-1992. Each column represents a separate regression. Demographic controls include educational attainment of head of household during childhood, race, number of children in the household, and share of childhood spent with married parents. State-year controls include maximum welfare generosity for family of three, unemployment rate, and minimum wage, all evaluated when individual was 17 years old in state of residence at the time.

Table 3. Effect of the EITC on educational attainment (in 1,000s of 2016 dollars)

	Completed high school by 19	Completed college by 23	Years of schooling at 25
Panel A. All women			
EITC exposure 0-15	0.002	0.009	0.03
	(0.003)	(0.003)	(0.011)
Outcome mean	0.67	0.21	13.75
N	3301	2628	2972
Panel B. White women			
EITC exposure 0-15	-0.001	0.01	0.029
	(0.004)	(0.004)	(0.014)
Outcome mean	0.69	0.24	13.91
N	1777	1403	1601
Panel C. Black women			
EITC exposure 0-15	0.007	0.0002	0.002
	(0.004)	(0.004)	(0.026)
Outcome mean	0.60	0.10	13.01
N	1476	1193	1330

Source: Panel Study of Income Dynamics 1968-2017. Women born between 1967 and 1992. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (race, education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends.

Table 4. Annual Social Welfare Benefit Savings from Reduction in Early Births, 2017 data

	EITC Savings	SNAP Savings	Medicaid Savings	Total
$10,000 increase in EITC benefits	$29,956,000	$18,562,000	$45,773,000	$94,291,000
$ 21,013 (SD) increase in EITC benefits	$62,910,000	$39,981,000	$96,127,000	$199,018,000

Source: Authors calculations based on birth statistics from the National Center for Health Statistics, EITC recipient information from the Internal Revenue Service, and demographic data from the Census Bureau's American Fact Finder. SNAP data come from the United State Department of Agriculture; Medicaid estimates generated using a method developed by Finkelstein, Hendren, and Luttmer (forthcoming).

Appendix Table 1. Selected EITC policy changes since 1975

		One child		Two children	
Year	Event	Income range for maximum benefit (plateau region)	Maximum income (end of phase out range)	Income range for maximum benefit (plateau region)	Maximum income (end of phase out range)
1975	Federal EITC implemented as a temporary credit, worth up to $400	4,000	8,000	4,000	8,000
1978	Federal EITC made a permanent credit, increased maximum benefit to $500	5,000-6,000	10,000	5,000-6,000	10,000
1987	Federal EITC indexed to inflation	6,080-6,920	15,432	6,080-6,920	15,432
1991	Separate EITC established for two-child households	7,140-11,250	21,250	7,140-11,250	21,250
1991-1996	Phase-in rate increased from 16.7 to 34 percent for one-child households, and from 17.3 to 40 percent for two-child households	6,330-11,610	25,078	8,890-11610	28,495
2009	Separate EITC established for three-child households; phase-in rate of 45 percent	8,580-15,740	35,436	12,060-15,740	40,295

Source: Jones and Michelmore (2018); Tax Policy Center: http://www.taxpolicycenter.org/statistics/eitc-parameters

Appendix Table 2. Effect of EITC exposure in childhood on marriage and childbearing in early adulthood (EITC in $1,000s of 2016$)

	18	19	20	21	22	23	24	25
Panel A. Birth								
All women	-0.002	-0.004	-0.006	-0.008	-0.007	-0.007	-0.006	-0.005
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)
Outcome Mean	0.132	0.199	0.267	0.321	0.361	0.403	0.447	0.487
N	3301	3210	3134	3063	3001	2930	2872	2730
White women	-0.002	-0.003	-0.005	-0.006	-0.004	-0.007	-0.007	-0.006
	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)
Outcome Mean	0.072	0.115	0.1597	0.201	0.237	0.271	0.32	0.359
N	1777	1728	1690	1654	1622	1583	1546	1477
Black women	-0.004	-0.007	-0.007	-0.009	-0.013	-0.002	0.002	0.002
	(0.005)	(0.005)	(0.004)	(0.004)	(0.005)	(0.006)	(0.006)	(0.006)
Outcome Mean	0.205	0.303	0.402	0.469	0.507	0.561	0.595	0.64
N	1476	1437	1400	1367	1341	1309	1290	1221
Panel B. Marriage								
All	0.000	-0.001	-0.001	-0.004	-0.005	-0.005	-0.006	-0.004
	(0.001)	(0.002)	(0.002)	(0.003)	(0.002)	(0.003)	(0.002)	(0.003)
Outcome Mean	0.032	0.064	0.097	0.138	0.181	0.229	0.277	0.333
N	3301	3210	3134	3063	3001	2930	2872	2730
White	0.001	-0.001	-0.001	-0.004	-0.007	-0.007	-0.007	-0.005
	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Outcome Mean	0.040	0.080	0.128	0.184	0.245	0.309	0.373	0.456
N	1777	1728	1690	1654	1622	1583	1546	1477
Black	0.000	0.000	-0.001	-0.001	0.000	0.001	0.000	0.001
	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)	(0.004)	(0.005)	(0.006)
Outcome Mean	0.015	0.028	0.044	0.061	0.067	0.107	0.146	0.185
N	1476	1437	1400	1367	1341	1309	1290	1221

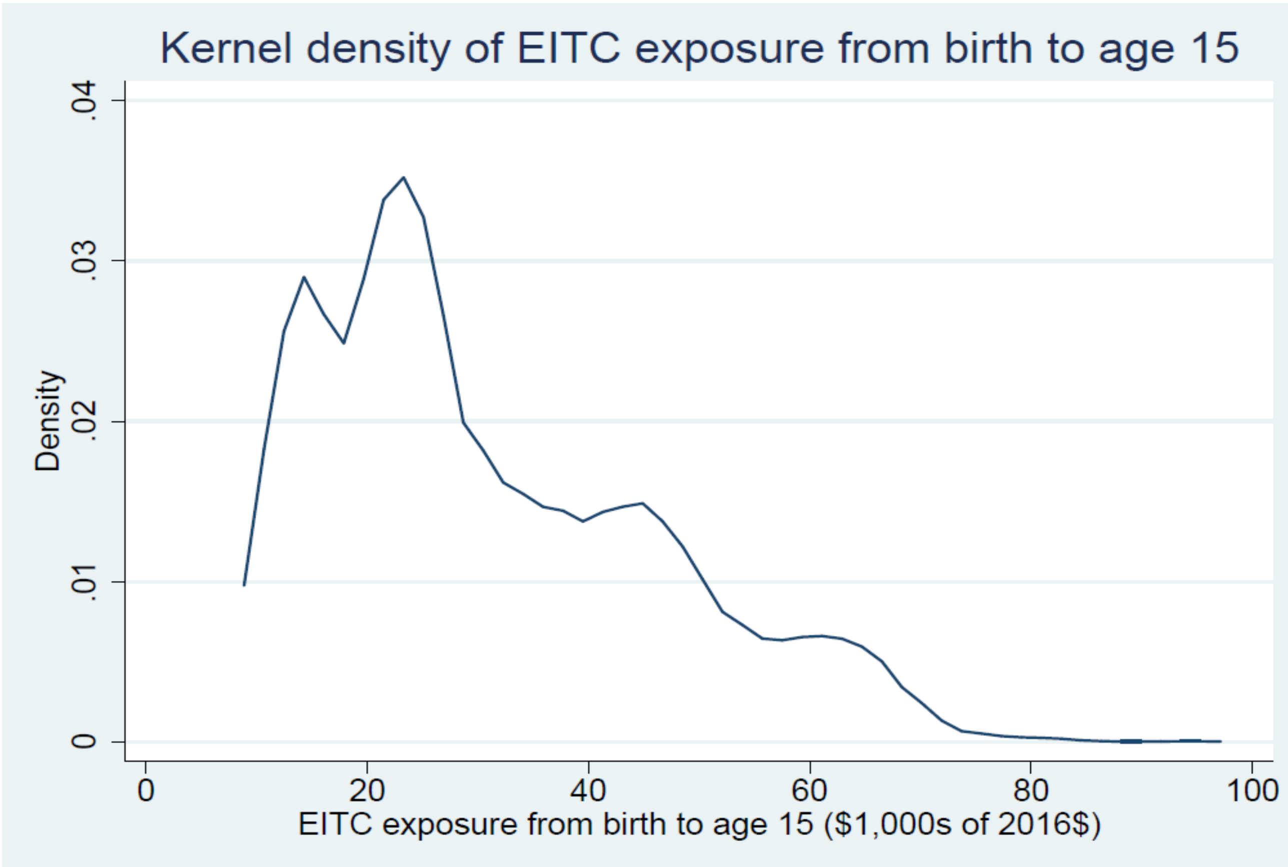
Source: Panel Study of Income Dynamics 1968-2017. Women born between 1967 and 1992. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (race, education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends.

Appendix Table 3. Effect of EITC exposure in childhood on marriage and childbearing in early adulthood (EITC in $1,000s of 2016$)

	18	19	20	21	22	23	24	25
Panel A. Birth								
All men	0.000	0.001	0.000	-0.001	0.000	-0.002	-0.003	-0.002
	(0.001)	(0.001)	(0.002)	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)
Outcome Mean	0.03	0.059	0.09	0.127	0.154	0.194	0.238	0.333
N	3321	3178	3079	2976	2879	2770	2670	2473
White men	0.000	0.002	0.002	0.000	0.000	-0.001	-0.002	0.000
	(0.001)	(0.001)	(0.002)	(0.003)	(0.003)	(0.002)	(0.002)	(0.003)
Outcome Mean	0.014	0.035	0.06	0.088	0.113	0.152	0.193	0.225
N	1737	1668	1627	1590	1555	1512	1468	1377
Black men	-0.001	-0.006	-0.006	-0.002	-0.002	-0.002	-0.004	-0.001
	(0.002)	(0.004)	(0.004)	(0.005)	(0.005)	(0.005)	(0.006)	(0.006)
Outcome Mean	0.089	0.15	0.207	0.27	0.311	0.355	0.416	0.452
N	1532	1461	1409	1344	1285	1223	1170	1068
Panel B. Marriage								
All men	0.000	0.001	0.000	0.002	-0.001	0.000	0.001	0.004
	(0.)	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)
Outcome Mean	0.009	0.023	0.051	0.082	0.129	0.179	0.231	0.250
N	3321	3178	3079	2976	2879	2770	2670	2473
White men	0.000	0.000	-0.001	0.003	-0.001	0.000	0.001	0.004
	(0.)	(0.001)	(0.002)	(0.003)	(0.002)	(0.002)	(0.003)	(0.004)
Outcome Mean	0.010	0.025	0.055	0.092	0.146	0.203	0.258	0.323
N	1737	1668	1627	1590	1555	1512	1468	1377
Black men	0.000	0.000	0.001	0.000	0.000	0.000	-0.001	0.002
	(0.)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.003)	(0.005)
Outcome Mean	0.004	0.015	0.026	0.041	0.051	0.070	0.111	0.157
N	1532	1461	1409	1344	1285	1223	1170	1068

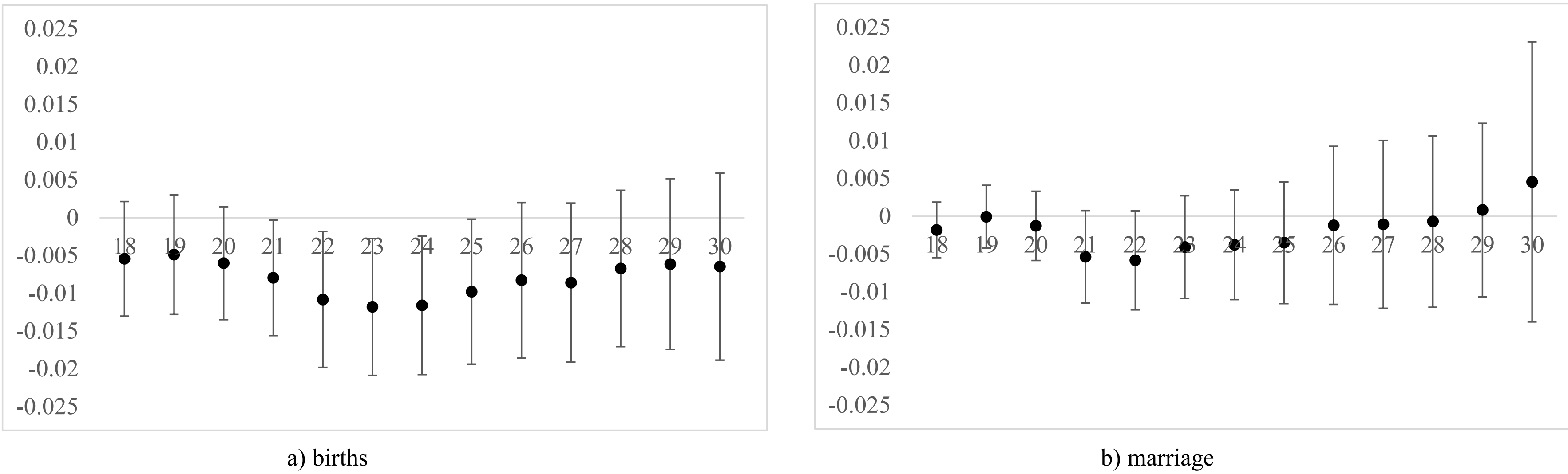
Source: Panel Study of Income Dynamics 1968-2017. Men born between 1967 and 1992. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (race, education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends.

Appendix Figure 1. Kernel density of distribution of EITC exposure from birth to age 15; birth cohorts 1968-1992



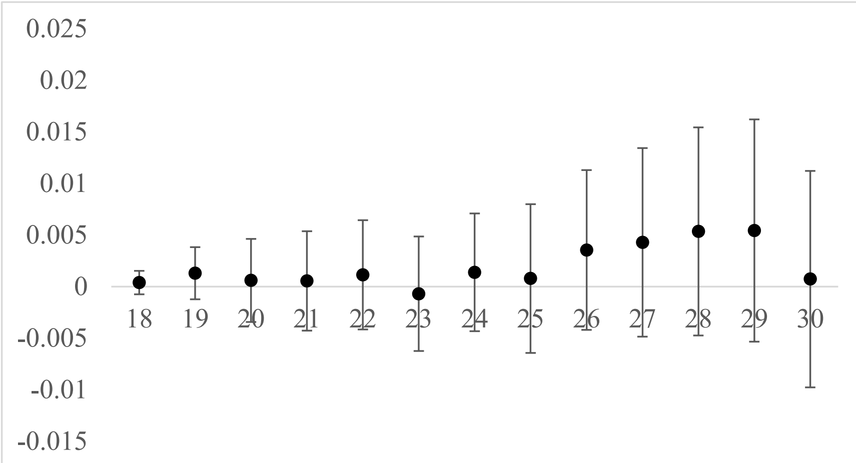
Source: Panel Study of Income Dynamics and authors' calculations. Men and women born 1968-1992. Note: Mean= $38,058. Standard deviation: $21,013.

Appendix Figure 2. Effect of EITC exposure from birth to age 15 on women's marriage and childbearing from 18-30; birth cohorts 1968-1987

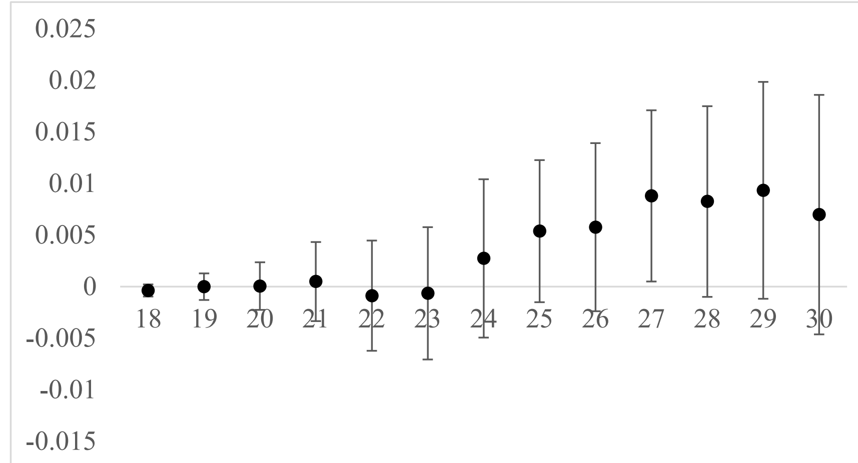


Source: Panel Study of Income Dynamics 1968-2017. Women born 1967-1987. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (race, education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends. 95% confidence intervals indicated by vertical bars. Each point represents a separate regression.

Appendix Figure 3. Effect of EITC exposure from birth to age 15 on men's marriage and childbearing by ages 18 to 30; birth cohorts 1968-1987



a) childbearing



b)marriage

Source: Panel Study of Income Dynamics 1968-2017. Men born 1967-1987. EITC exposure measured in $1,000s of 2016$. All models include demographic controls (race, education of the head, share of childhood spent with married parents, number of children in the household fixed effects) state contextual variables (state welfare generosity, unemployment rate, and minimum wage in the state when individual was 15 years old), state fixed effects, birth year fixed effects, and birth year-specific time trends. 95% confidence intervals indicated by vertical bars. Each point represents a separate regression.