
Effects of the Earned Income Tax Credit for Childless Adults: A Regression Discontinuity Approach

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Executive Summary

Most antipoverty policy in the United States focuses on families with children, but efforts to assist childless adults have gained traction in recent years. We examine the impact of the earned income tax credit on the labor force outcomes of childless adults using the age-25 eligibility discontinuity. We find no impacts on labor force participation and employment outcomes, which may be due to lack of information about the credit, lack of behavioral response due to its small size, or that childless adults already have very high labor force participation rates.

I. Introduction

The earned income tax credit (EITC) is one of the primary antipoverty tools in the United States, transferring more than \$60 billion to 25 million low-income working households in 2019 (Internal Revenue Service 2019). The EITC is thought to encourage employment for low-income households because people only become eligible if they have positive earnings. Previous research has focused almost exclusively on the impacts of the EITC on households with children; see Hoynes and Rothstein (2016) for a review, and recent work by Neumark and Shirley 2017; Wilson 2020; Kleven 2019; Kuka and Shenhav 2020; Neumark and Williams 2020; and Schanzenbach and Strain 2021, *inter alia*. Most studies find positive impacts on employment among single mothers with lower levels of education,

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and often significant economic and social benefits, such as reductions in recidivism (Agan and Makowsky 2018), improved mother and infant health (Evans and Garthwaite 2014; Hoynes, Miller, and Simon 2015; Markowitz et al. 2017), boosts in educational achievement and attainment (Micheltmore 2013; Bastian and Micheltmore 2018), increases in intergenerational mobility (Jones, Simeonova, and Akee 2020), changes in marriage and fertility (Eissa and Hoynes 2000; Holtzblatt and Rebelein 2000; Baughman and Dickert-Conlin 2009; Maag and Acs 2015; Bastian 2017), and more. Studies looking at the employment effects of married couples find that total labor force participation increases for men and decreases for women. Overall, the combined labor supply of married couples seems to decrease because the increases from married men do not offset declines from married women (Eissa and Hoynes 2004). Research looking at the impacts of the EITC on adults without children, however, is sparse.

In large part, the focus on households with children is because the maximum amount of the federal credit is far lower for those without children, just \$560 in 2022, as compared with \$6,935 for filers with three or more children. Furthermore, the credit for childless adults is fully phased out at \$16,480 for single filers in 2022, far lower than for those with children. Indeed, childless adults are often used as a control group to measure the impacts of the EITC (Meyer and Rosenbaum 2001; Hoynes and Patel 2018; Neumark and Williams 2020), with the justification that “individuals with no children are essentially ineligible for the EITC” (Chetty, Friedman, and Saez 2013) and that “the small credit offered is unlikely to induce a significant behavioral labor supply response” (Neumark and Williams 2020).

However, childless adults do account for a quarter of EITC recipients—nearly 7 million taxpayers—even as they only receive 3% of EITC payments, with an average claimed credit of \$302 (Crandall-Hollick 2021). Recent policy efforts have emphasized expanding the EITC for childless adults: as part of the response to the coronavirus pandemic, the American Rescue Plan Act reduced the minimum age of eligibility from 25 to 19 for the 2021 tax year, increased the phase-in (and phase-out) rate, nearly tripled the size of the credit, and expanded the income range (Crandall-Hollick, Falk, and Boyle 2021). The proposed Build Back Better Act would make this change permanent.¹

States play a large role in the EITC program. More than 30 states have their own EITC program, structured as a percentage of the federal credit. In 2022, all but six of those are fully refundable. Generosity ranges widely, from 3% to 100% of the federal credit (Tax Policy Center 2022). Combined,

the maximum refundable EITC a childless adult could claim for the 2022 tax year ranges from \$560 to \$1,120, depending on the state. Witter (2020) finds that state EITC expansions between 1994 and 2017 led to small but significant increases in employment and labor force participation for younger childless women. Miller et al. (2018) similarly find small increases in employment, tax filing, and child support payments among noncustodial parents in response to the New York City Paycheck Plus program (an EITC-like program). These two studies looking at state expansions of the EITC offer promising but incomplete evidence that expanding the federal EITC can lead to increases in employment and labor force participation, as well as positive impacts on tax revenue and children with low-income noncustodial parents.

Rather than exploit variation in state or local EITC eligibility, we use the federal EITC's age-25 eligibility criterion as a source of identifying variation for the impacts on both laborforce participation and employment. Childless adults become eligible for the EITC in the year in which they turn 25, meaning that observationally similar people born just a few days apart are eligible to receive the credit a full year apart.² Recipients cannot manipulate the running variable, birth date, given that individuals in our sample were born before the EITC was introduced for childless adults in 1994. We use the 2001, 2004, 2008, and 2014 waves of the Survey of Income and Program Participation (SIPP) to examine outcomes on either side of this cutoff. That is, we compare outcomes for those turning 25 at the very end of the calendar year—who are eligible for EITC payments based on their labor income in that year—to those who are born shortly afterward, in the next year, and are thus ineligible until the following year.³

We find no effects on labor force participation or employment around this cutoff, either in the first year of eligibility or the following year. This effect may be driven by several mechanisms. It may be that the amount of the credit is too low to induce entry into the labor force, as hypothesized in Neumark and Williams (2020). It is also possible that the complexity of the eligibility requirements in particular, and of tax filing in general, reduces participation (Miller and Mumford 2015; Benzarti 2020, 2021). Although about four-fifths of eligible households take up the EITC, participation is lowest for the group we examine.⁴ But Kopczuk and Pop-Eleches (2007) document an increase of nearly 4 million EITC recipients in 1994 “mostly due to extending eligibility to childless individuals.” E-filing and assisted preparation also reduces complexity and increases participation (Kopczuk and Pop-Eleches 2007; Goldin 2018). Although knowledge

of the structure of the EITC affects how individuals report their earnings (Chetty et al. 2013), most evidence indicates that providing information to potentially eligible households has little effect on EITC participation (Cranor, Kotb, and Goldin 2019; Linos et al. forthcoming).⁵ Furthermore, labor force participation is very high among this group of young adults. Many of those who are never in the labor force during the sample period report a work-limiting disability. As such, this may simply be a margin of potential workers who are unlikely to enter the labor force, especially in response to such a small incentive.

A short-run analysis indicates that the incidence of such a subsidy will be shared between employers and employees if supply increases meaningfully, muting the impact of the transfer. Because we find no such effects, the incidence of the transfer accrues to employees to the extent that they take up the credit.⁶ And because there is no change in the eligibility of those with children at this age cutoff, there should be little concern about general equilibrium effects that lead to spillovers in this context.

Our findings cannot be extrapolated to measure the impact or incidence of proposals that triple the maximum EITC for childless adults.⁷ But they do shed light on the degree to which the current incarnation of the EITC has affected behavior. In Section II, we discuss the creation of the data set and our empirical approach. Section III presents the results of the regression discontinuity analysis, and Section IV concludes.

II. Data and Empirical Approach

A. Data

The EITC was introduced in 1975 for households with at least one dependent, with a modest increase in benefits introduced in 1986. In 1990, the federal EITC became more generous for families with two or more children relative to one-child families, phased in over several years. A significant increase was introduced in 1993, and in 2009, the EITC expanded for families with three or more children. The federal EITC for childless adults was introduced in 1994, and its generosity has not been expanded since, except for inflation adjustments and the temporary increase during the 2021 tax year.

To qualify for the EITC without a child, tax filers must be between the ages of 25 and 64, have earned income from wages, self-employment, or business, and live in the United States for at least half the year. They cannot be claimed as a dependent on another household's tax return or have

investment income above a threshold (\$3,650 in 2020; temporarily increased to \$10,000 for 2021). Filers with children can still be considered “childless” when filing their taxes if the child lives with them for less than 6 months or is claimed as a dependent on another return.⁸ Indeed, childless adults are diverse in terms of their household structure and family histories. Many are parents without custody of their children or have children who are grown and moved away from home.

In 2022, EITC credit amounts for childless adults are determined as follows. The phase-in rate for the EITC for single childless adults was 7.65% and extends over a range of \$0–\$7,320 in annual earnings. Single childless adults with annual earnings between \$7,320 and \$9,160 are eligible for the maximum EITC of \$560. The phase-out rate is the same as the phase-in rate of 7.65% and extends from \$9,160 to \$16,480, where the EITC becomes \$0. In other words, EITC amounts are equal to $(0.0765 \times \text{Annual Earnings})$ for those with incomes in the phase-in range, \$560 for those eligible for the maximum EITC, and $(\$560 - [0.0765 \times \text{Annual Earnings}])$ for those with incomes in the phase-out range. The credit is structured the same for jointly filing married couples except that it begins phasing out at a higher income threshold.⁹

Our sample is drawn from the 2001, 2004, 2008, and 2014 panels of the SIPP, covering 2000–16. Each SIPP panel surveys a set of households for several years, conducting interviews every 4 months and covering activities since the previous interview, including labor force participation and employment status in each week.¹⁰

We construct our eligibility measure using respondents’ ages on the last day of the calendar year. Eligibility for married couples is determined by the age of the older spouse. We therefore use that age for married respondents; the qualitative conclusions are unaffected by using actual age.¹¹ Because taxes are filed annually, we aggregate the panel to the individual-year level. Our primary outcome measures are the proportion of weeks in the labor force or employed. Our results are similar when we use month-based observations instead. Because the entry and exit of survey participants into the SIPP panels do not line up perfectly with calendar years, many annualized observations are based on less than a full year’s worth of reporting. Our results are unaffected when limiting the sample to observations based on 3, 6, or 9 months of individual data in a given year.

Full-time students under the age of 24 are eligible to be claimed as a dependent by others, so we exclude those with more education than a high school degree. We also exclude all SIPP respondents who had children at

any point during the sample, as well as those under 18 or over 65. Otherwise, we impose no restrictions on the sample. Altogether, our data consist of 106,700 annual observations on 35,321 individuals, of whom 6,942 were, at their oldest, between the ages of 20 and 30 (inclusive) during the sample period. These individuals represent 16,641 person-year observations and form the core of our sample; bandwidth selection reduces the size of the sample, depending on the specification.

Of this sample, 67.0% report being in the labor force in every week about which they were asked in a given year, and 14.5% report not being in the labor force in any week about which they were asked; the percentages for employment status are 57.2% and 21.2%, respectively. The overall proportion of weeks in the labor force is 78.4%, and the percentage is 70.0% for employment. Some 82.8% were employed at any point during the sample. See table 1 for summary statistics.

Table 1
Summary Statistics

	Mean	Standard Error
Observations for persons:		
Female	.333	.471
Ever married	.166	.372
Ever in the labor force	.896	.306
Ever employed	.829	.376
Observations for person-years:		
Share of weeks in the labor force	.785	.373
Share of weeks employed	.700	.417
Married	.124	.330
Age on December 31	24.773	2.917
Hours worked	1,357	1,068
25th percentile		0
Median		1,620
75th percentile		2,085
Total earned income (2019\$)	23,718	30,597
25th percentile		3,128
Median		19,147
75th percentile		32,354

Note: The first set of summary statistics uses 6,942 observations at the individual level, limited to those whose own maximum age or spouse's age in December was between 20 and 30 years during the entire sample. The second set of summary statistics uses 16,641 observations at the individual-year level, limited to those whose age in December was between 20 and 30 years. The analytical sample used for each specification is a subset of these person-year observations that depend on the optimal bandwidth.

B. Empirical Approach

We rely on the assumption that childless adults who turn 25 just before the end of the tax year are similar to those who turn 25 just after the end of the tax year, with the exception that those born before are eligible to claim the EITC in the following year. Our estimating equation is a simple regression discontinuity around the eligibility cutoff:

$$Y_{i,t} = \beta_0 + \beta_1 T_{i,t} + \beta_2 \text{Age}_{i,t} + \beta_3 T_{i,t} \times \text{Age}_{i,t} + \varepsilon_{i,t}.$$

$T_{i,t}$ is an indicator for being under 25 years of age prior to the end of the of the tax year, and β_1 measures the discontinuity in the outcome $Y_{i,t}$ for those who were 25 before the first day of the sample year. Age has a linear effect on the outcome on either side of the discontinuity; results are unchanged when using a quadratic. The bandwidth around the cutoff is selected using mean squared error optimal bandwidths and a triangular kernel (Calonico et al. 2017). Because there are generally multiple observations per individual, we cluster the standard errors at the individual level (Calonico et al. 2017). We use sample weights from the SIPP, though the results are unchanged when not using weights. We also estimate the effects after residualizing the outcome variable for state-year effects and individual effects (Lee and Lemieux 2010); the results are very similar.

Figure 1 shows the distribution of observations by age on December 31. There does not appear to be any manipulation of the running variable based on this figure. We also examine whether there are discontinuities across the cutoff in gender or whether the respondent has ever been married; there are no discontinuities present. Taken together, this suggests that our approach of examining outcomes for those who are just below the age-25 eligibility cutoff at the end of the year to those who are just over it will yield causal estimates of the impacts of the EITC on the labor market outcomes of childless adults.

III. Results

We begin with the predicted EITC eligibility and amount, in figures 2 and 3. These are calculated using detailed information on family size, geography, and income information from the SIPP panels using the NBER TAXSIM v35 Stata Interface.¹² These estimates include state EITC amounts, for which a small number of childless adults under 25 are eligible. The discontinuity in this intent-to-treat first stage is clear, with those who are just

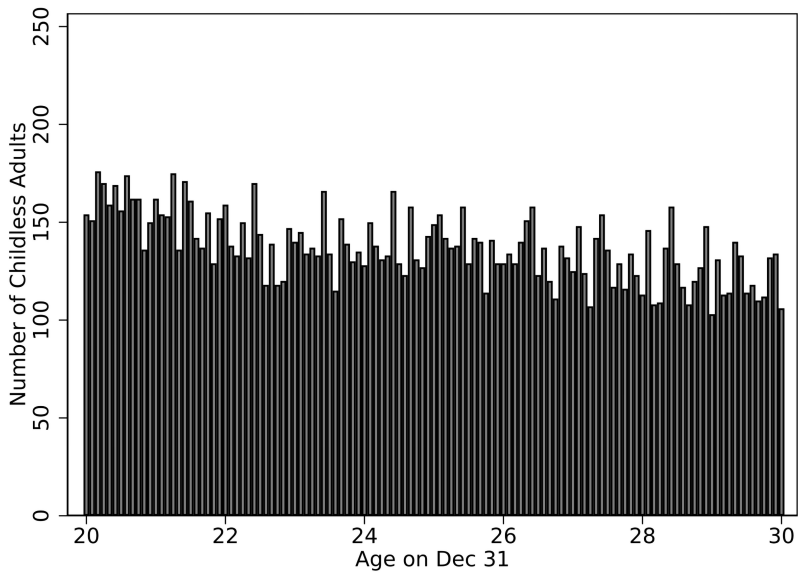


Fig. 1. Frequency of observations by age on December 31.

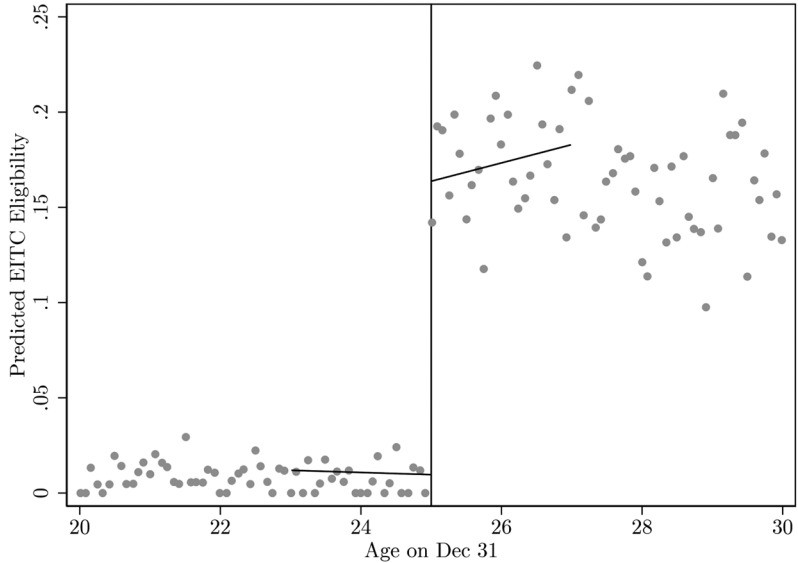


Fig. 2. Predicted earned income tax credit (EITC) eligibility. This figure shows results for predicted EITC eligibility (calculated using NBER TAXSIM), weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 1.99 years, using 7,654 observations.

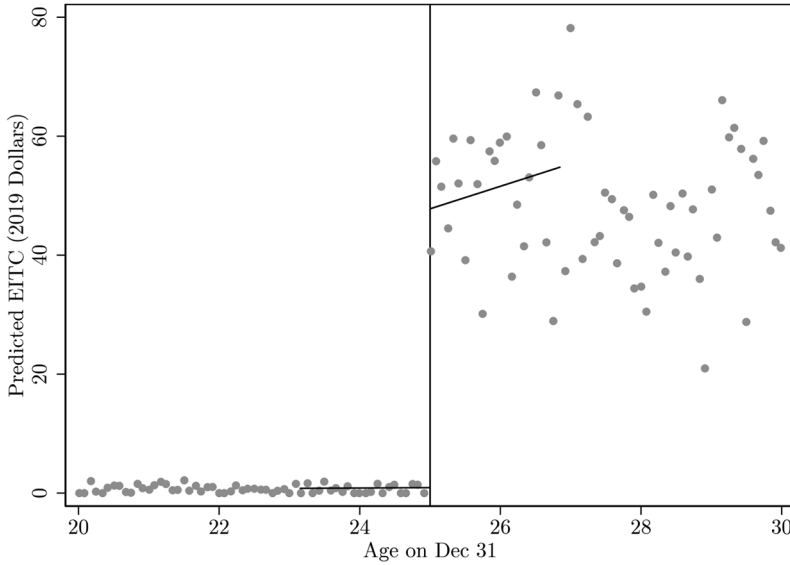


Fig. 3. Predicted earned income tax credit (EITC) amount. This figure shows results for predicted EITC amount received (including zeroes and calculated using NBER TAXSIM), weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 1.86 years, using 7,326 observations.

above age 25 being about 15.4 percentage points ($SE = 1.5$ percentage points) more likely to be eligible for the EITC. Including zero-dollar recipients, the predicted real amount received has a discontinuity of \$46.88 ($SE = \5.00).

Figure 4 shows a simple examination of the effect of age-25 eligibility on the share of the year spent in the labor force by childless adults. There is no discontinuity for those who are just barely eligible for the EITC in a given year, based on their age, as compared with those who are just barely ineligible. The estimated discontinuity is -0.0076 ($SE = 0.020$). Similar results are seen in figure 5, which measures the discontinuity for employment. The measured effect for that outcome is -0.0006 ($SE = 0.020$). Table 2 shows results for labor force participation, and table 3 shows those for employment.¹³

The discontinuity across the age-25 cutoff is fuzzy. We can scale up the results in figures 4 and 5 by a factor of about 6.5, assuming that any estimate is driven entirely by those who are (intended to be) eligible for the EITC. The effect becomes much larger, of course, about -5 percentage points for labor force participation but only -0.4 percentage

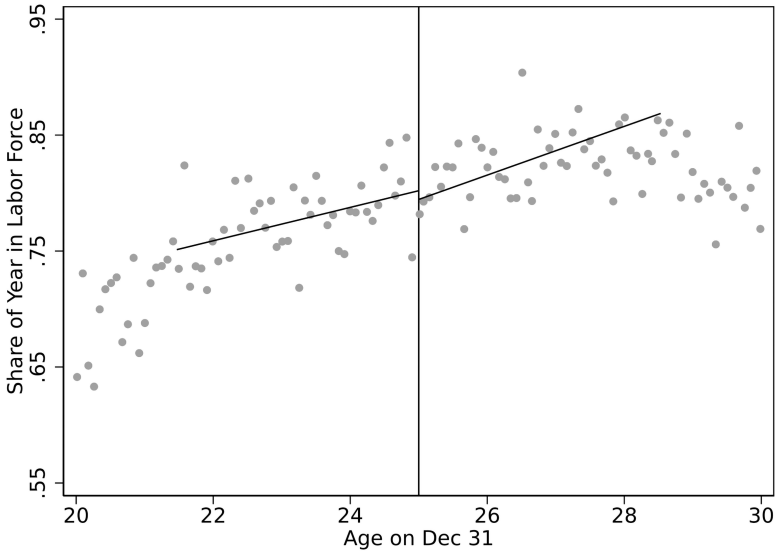


Fig. 4. Labor force participation. This figure shows results for the share of weeks in the labor force, weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 3.53 years, using 11,582 observations.

points for employment. The predicted effect is expected to be positive: the EITC encourages participation and employment. These estimates are not only noisy (and, in the case of employment, still very small) but negative.

As discussed above, it is possible that the population around this cut-off is unaware of the EITC and, as such, we would not expect to see any differences in labor market participation in the year in which they turn 25. Those to the right of the cutoff—that is, those who are old enough to qualify—may discover that they had been eligible and adjust their behavior in the following year. If so, we would expect labor force participation to increase, as has generally been seen in the EITC literature on single mothers. Figures 6 and 7 show the results for labor force participation and employment, respectively, in the following year. The discontinuities are small, statistically insignificant, and negative, running in the opposite direction of what would be expected. As such, we conclude that there is no evidence for dynamic effects of eligibility; again, these may be driven by lack of information or simply lack of response to the relatively small credit.¹⁴

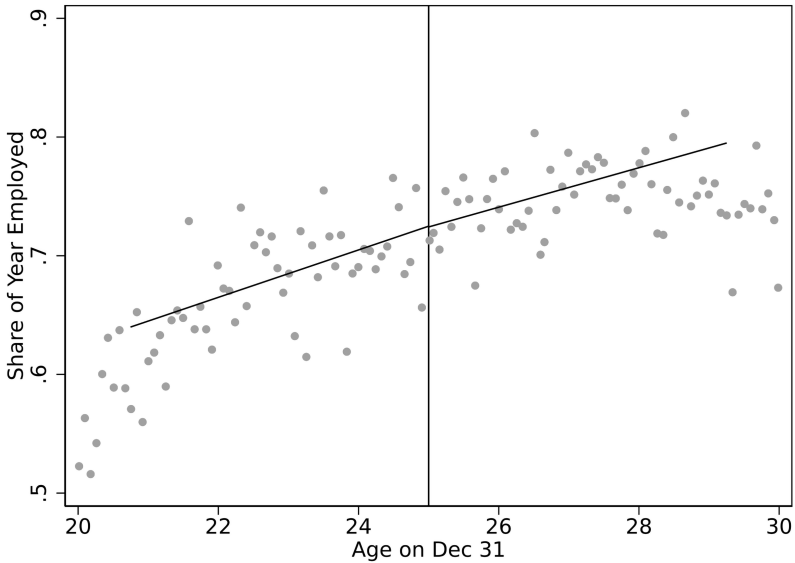


Fig. 5. Employment. This figure shows results for the share of weeks employed, weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual’s age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 4.25 years, using 14,059 observations.

We also estimate effects for those who are more likely be aware of the EITC. People who had a positive tax liability in the previous year are more likely to have filed taxes and become aware of the EITC. The estimated discontinuity for this group continues to be small and statistically insignificant, at -0.015 ($SE = 0.020$) for labor force participation. Another proxy for federal EITC awareness is the existence of a state-level EITC. Neumark and Williams (2020) find that state EITC expansions lead to increases in federal tax filing. However, when we look at discontinuities separately for childless adults in states with EITCs versus those without

Table 2
Labor Force Participation

	(1) Linear	(2) Quadratic	(3) Following Year (Linear)
Estimated discontinuity at age 25	-.008 (.020)	-.014 (.024)	-.006 (.024)
Number of observations used	11,582	15,775	6,549

Note: Each column reports the estimated discontinuity in the share of weeks reporting being in the labor force around the age-25 earned income tax credit eligibility. Estimates are weighted using Survey of Income and Program Participation weights and use a triangular kernel and MSE-optimal bandwidth.

Table 3
Employment

	(1) Linear	(2) Quadratic	(3) Following Year (Linear)
Estimated discontinuity at age 25	-.001 (.0204)	.004 (.025)	-.011 (.026)
Number of observations used	14,059	17,159	6,708

Note: Each column reports the estimated discontinuity in the share of weeks reporting being in the labor force around the age-25 earned income tax credit eligibility. Estimates are weighted using Survey of Income and Program Participation weights and use a triangular kernel and MSE-optimal bandwidth.

EITCs, the differences in labor force participation and employment are negligible in both significance and magnitude.

Because many people in the sample are never in the workforce, it may be that we are unable to measure a labor force participation response for the margin of those who might actually join the workforce. Figure 8 shows the discontinuity in labor force participation for those who were in the labor force at any point during the sample period. The discontinuity is, again, small and statistically insignificant. In the same vein, we limit the

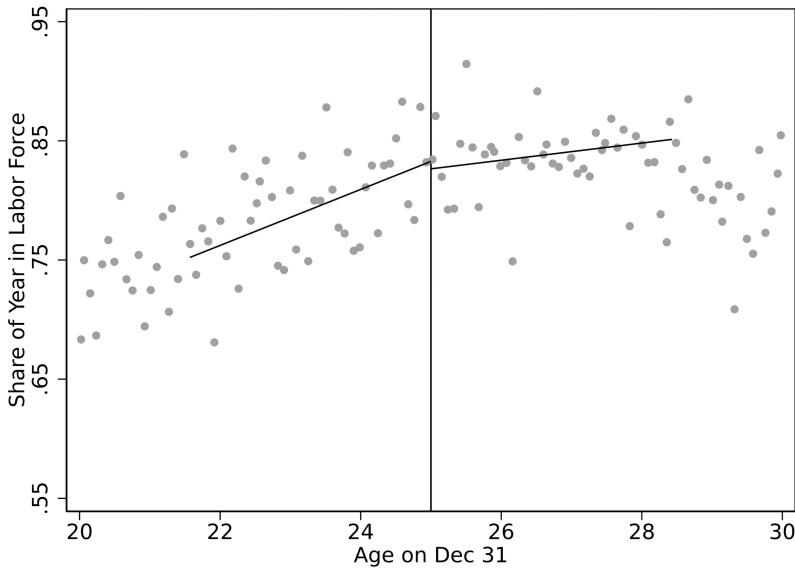


Fig. 6. Following-year labor force participation. This figure shows results for the share of weeks in the labor force in the year following the age used as the running variable, weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The symmetric bandwidth is MSE-optimal and is 3.43 years, using 6,549 observations.

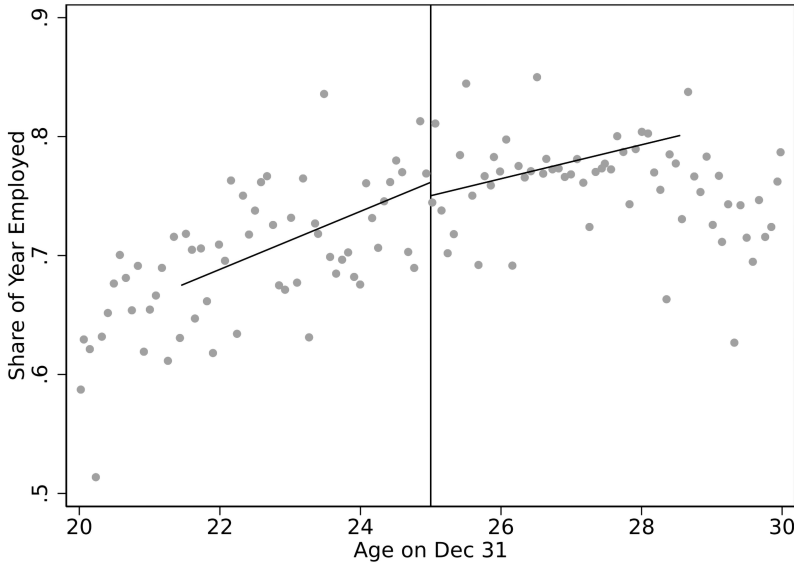


Fig. 7. Following-year employment. This figure shows results for the share of weeks employed in the year following the age used as the running variable, weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The symmetric bandwidth is MSE-optimal and is 3.55 years, using 6,708 observations.

sample to those whose lowest level of annual earned income (both individual and spouse, if applicable) in the sample was less than \$30,000. Based on the income eligibility criteria, these households were far more likely to be exposed to the EITC. The results are unaffected, with the discontinuity for labor force participation estimated at -0.0028 ($SE = 0.021$).

We can look at labor force participation and employment in a slightly different way to ensure we are fully capturing the potential extensive margin effects of the EITC on labor force participation and employment. We do this by running our regression discontinuity specification on binary indicators of whether childless adults were in the labor force or were employed at any time during the year. The first thing of note here is that labor force participation among both eligible and ineligible groups of childless adults is already pretty high. In fact, only about 10% of individuals ages 20–30 are never in the labor force during the sample period. We estimate insignificant discontinuities in labor force participation of about 0.71 percentage points ($SE = 1.35$ percentage points) and 1.03 percentage points ($SE = 1.63$ percentage points) in employment. Of those 10% of young childless adults who are never in the labor force, 39.7% report having a disability that limits their ability to work or prevents them from

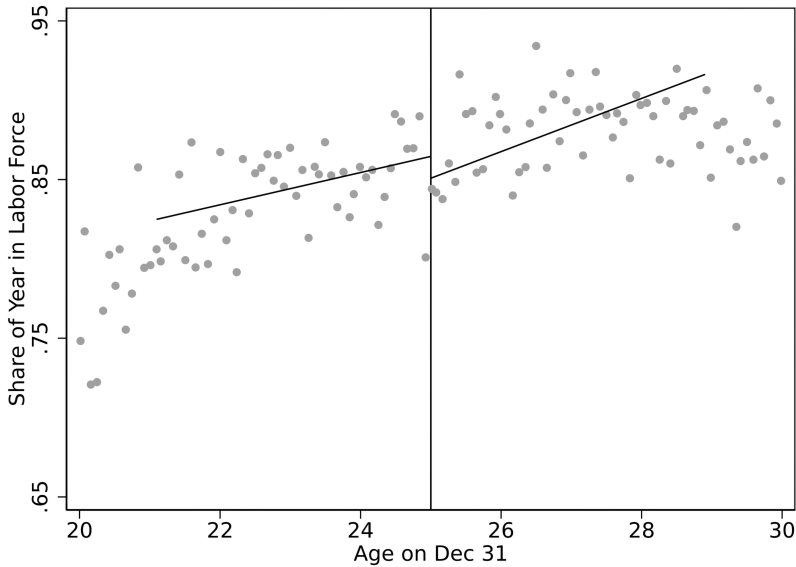


Fig. 8. Labor force participation (for those ever in the labor force). This figure shows results for the share of weeks in the labor force for individuals who were ever in the labor force during the sample period, weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 3.90 years, using 11,762 observations.

working entirely. There may be little scope to induce entry into the labor force in this group, providing an additional explanation for the miniscule effects that we find across this young childless adult population.

We also examine results after residualizing the outcome variable, taking out individual fixed effects as well as state-by-year fixed effects.¹⁵ We are therefore looking at within-individual responses to EITC eligibility based on age at the end of the year. The results, in figure 9, show a statistically insignificant discontinuity of -0.014 ($SE = 0.014$) in labor force participation.

Figure 10 shows the discontinuity for annualized hours of work, and figure 11 shows it for earned income (including spousal income, if applicable). The discontinuities are negligible in magnitude. We also examine results by gender and marital status. Figures 12A and 12B estimate the discontinuity in labor force participation for unmarried and married women, respectively, and figures 13A and 13B show those for men. There are no meaningful patterns or significant differences in effect sizes. Effects for married childless adults are noisy and sensitive to bandwidth selection.

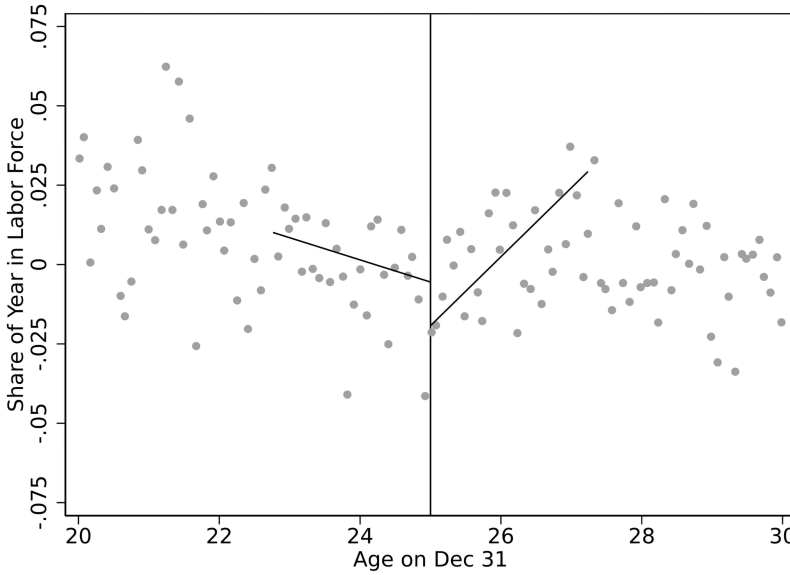


Fig. 9. Labor force participation (residualized). This figure shows results for the share of weeks in the labor force, residualized to remove individual fixed effects as well as state-year fixed effects. The estimation is weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 2.24 years, using 6,369 observations.

There are far fewer young married childless adults than young single childless adults. Those results should be interpreted with caution.

Last, we look at labor force participation and employment results separately for those with earnings above and below the earnings threshold where the EITC eligibility amount is maximized. Labor theory predicts that workers will work fewer hours or weeks when their earnings levels are positioned on the plateau or phase-out region of the EITC schedule, because both the income and substitution effects push toward leisure (Meyer 2002). To investigate this further, we plot the frequency distribution of single and married earners overlaid with their predicted EITC. Figures 14A and 14B show that there are more earners with earnings above the phase-in portion of the EITC schedule (about \$7,000). Next, we estimate the discontinuities in labor force participation and employment for those earning above or below that threshold separately. We find similar insignificant effect sizes as reported above. This suggests that the insignificant negligible effect sizes we find in our main specifications are not the result of an averaging of coefficients from childless adults populations with opposite labor supply incentives.

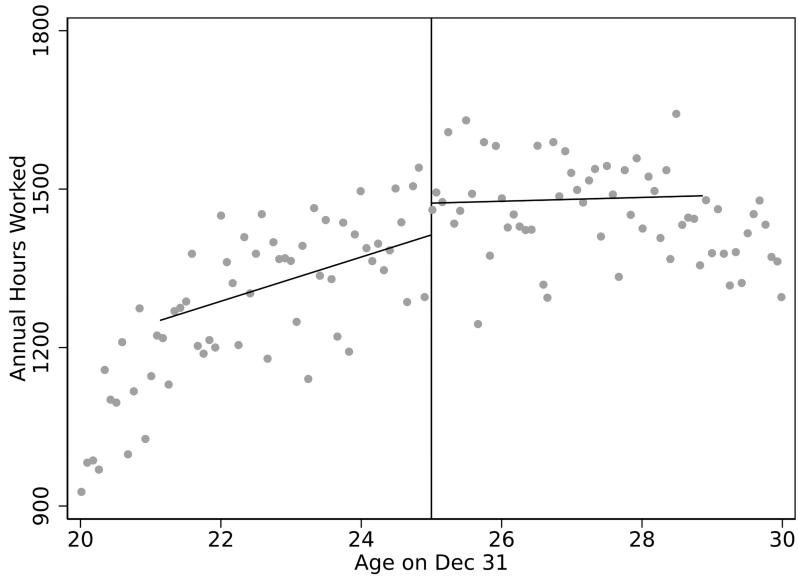


Fig. 10. Hours. This figure shows results for annual hours worked. The estimation is weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 3.87 years, using 12,688 observations.

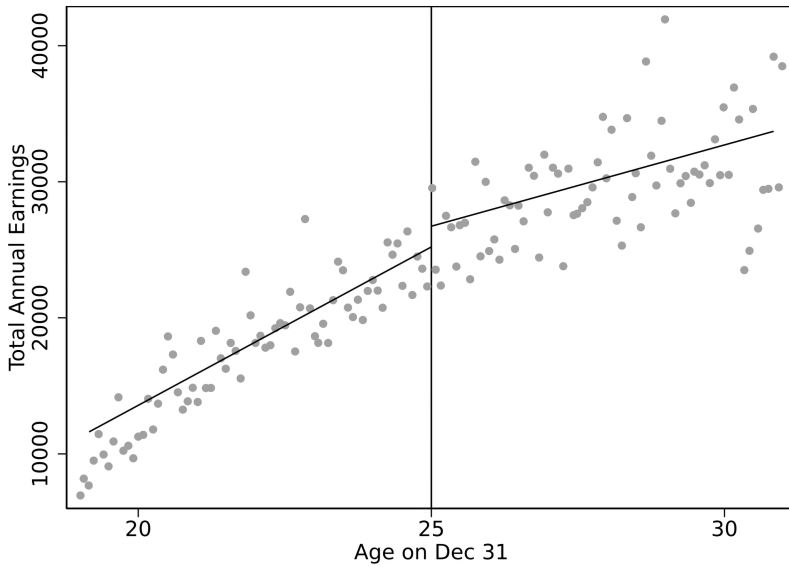


Fig. 11. Earned income. This figure shows results for total earned income in US dollars (includes spousal income, if relevant). The estimation is weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 5.84 years, using 19,500 observations.

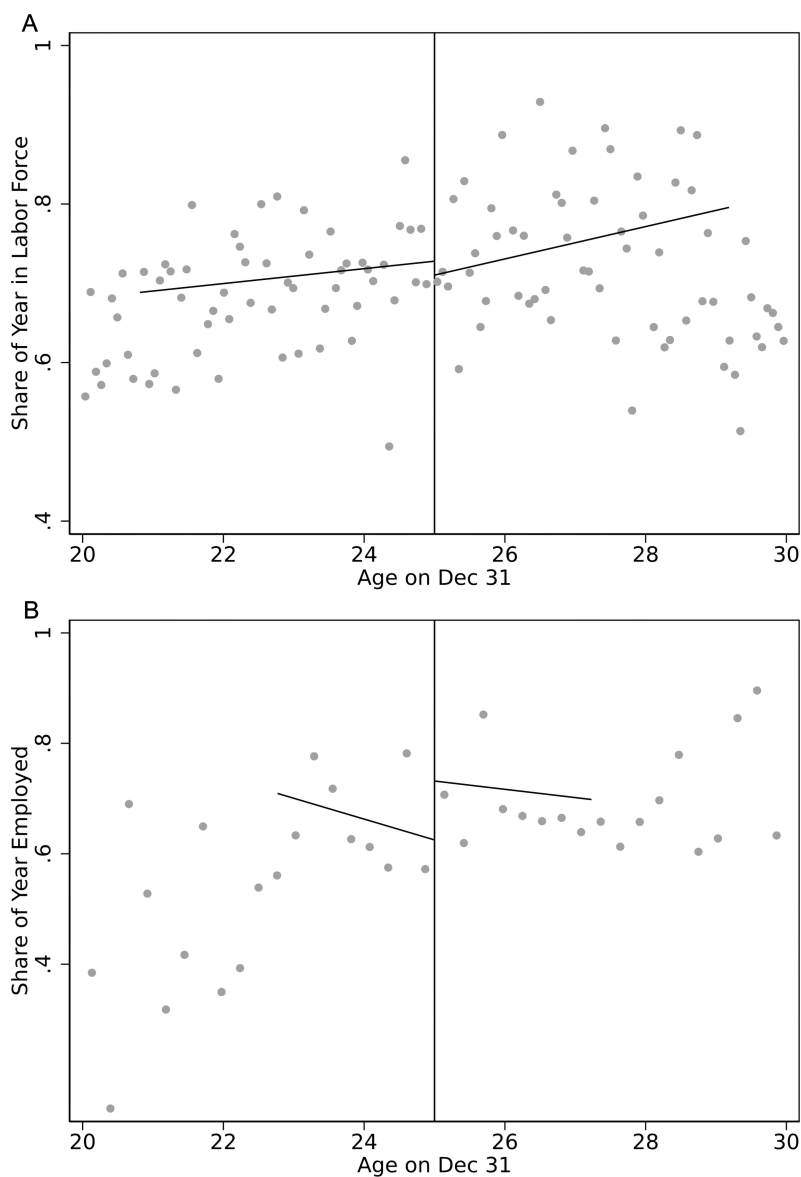


Fig. 12. (A) Labor force participation (unmarried females). This figure shows results for the share of weeks in the labor force for unmarried females, weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 4.19 years, using 3,654 observations. (B) Labor force participation (married females). This figure shows results for the share of weeks in the labor force for married females, weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 2.06 years, using 360 observations.

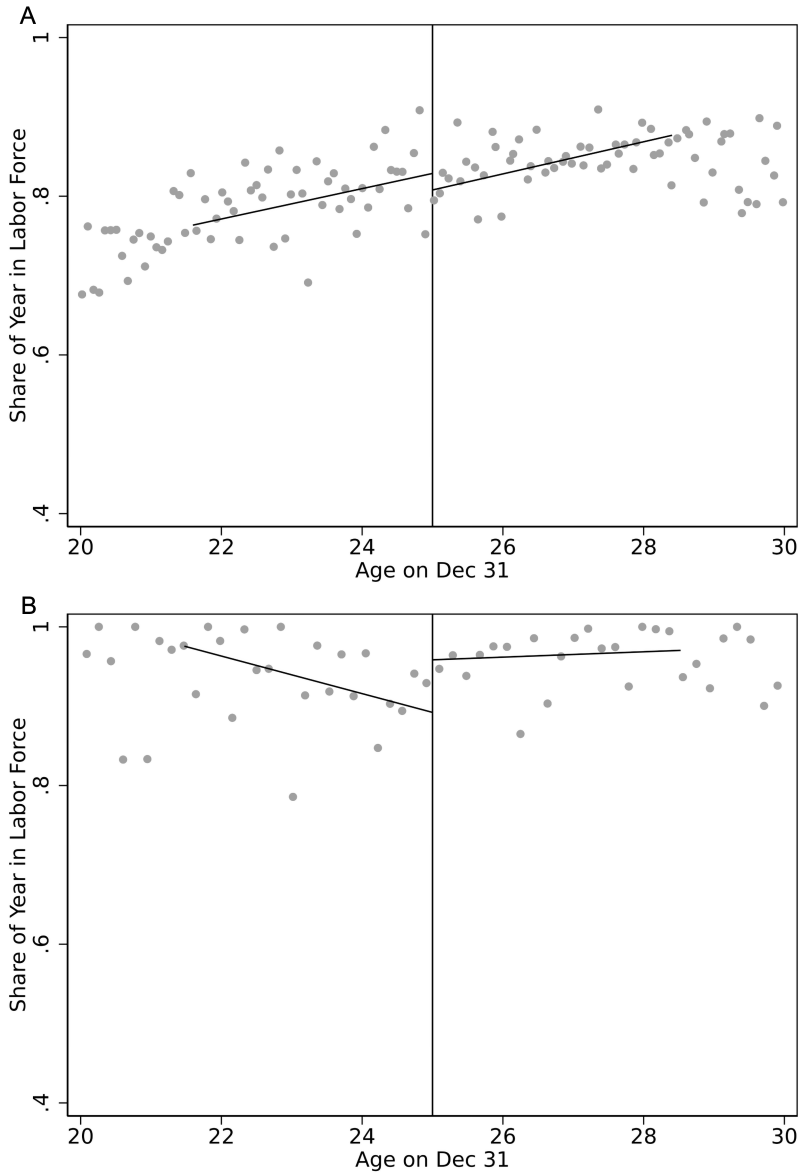


Fig. 13. (A) Labor force participation (unmarried males). This figure shows results for the share of weeks in the labor force for unmarried males, weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 3.40 years, using 6,681 observations. (B) Labor force participation (married males). This figure shows results for the share of weeks in the labor force for males, weighted using Survey of Income and Program Participation weights and estimated using a triangular kernel. The running variable is the individual's age on December 31 of that year. The symmetric bandwidth is MSE-optimal and is 3.53 years, using 958 observations.

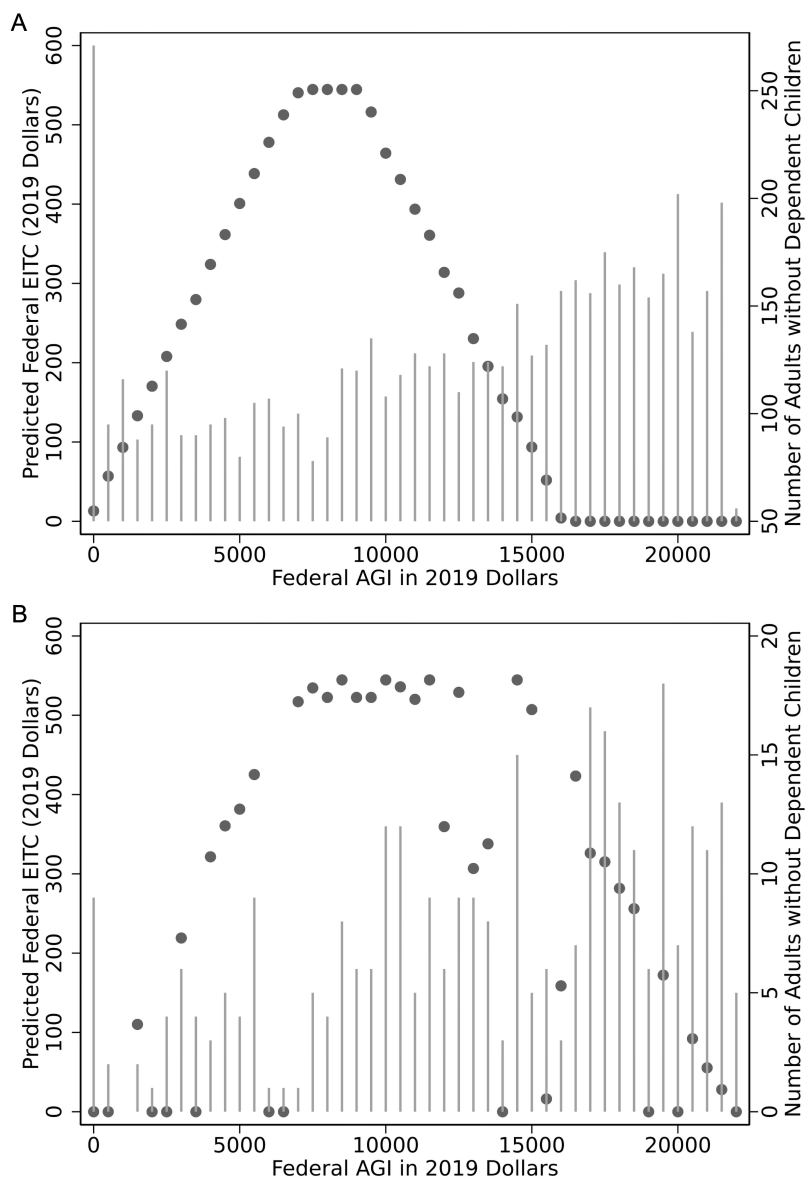


Fig. 14. (A) Predicted earned income tax credit (EITC) and income distribution (unmarried). AGI = adjusted gross income. (B) Predicted earned income tax credit (EITC) and income distribution (married). AGI = adjusted gross income.

IV. Conclusions

We examine the impact of the age-25 EITC eligibility cutoff for childless adults, a group whose economic well-being has been the subject of an increased focus for policy makers. Comparing the labor market behavior of those who became eligible in a given year with those who were just outside the eligibility range, we find no impacts on labor force participation, employment, or hours worked either in the year in which people become eligible or in the following year. These findings could be driven by a lack of information about the credit, a lack of response to it due to its small size, or the fact that only 6% of young childless adults are both out of the labor force and do not have a work-limiting disability. The lack of employment effects contrasts with previous studies looking at the impact of state and local EITC expansions on employment among the childless adult population. These differences likely arise from the fact that this study looks at effects around the age-25 cutoff specifically, whereas the other studies look at wider ranges of childless adults. Differences could also stem from differential baseline employment or labor force participation rates between expansion and nonexpansion states, awareness of own EITC eligibility, EITC generosity, or all of the above. A large expansion of the federal EITC may have very different effects.

Endnotes

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1. In 2016, Speaker Paul Ryan and President Barack Obama proposed nearly identical plans to lower the eligibility age for that group, expand the eligible income range, and increase the maximum credit amount. No legislative action resulted due to disagreements about paying for the expansion.

2. Shirley (2020) uses a similar approach around the timing of a first birth (and thus EITC eligibility) to examine mothers' labor-supply responses. Barr, Eggleston, and Smith (2022) do so to investigate long-run outcomes for those children.

3. There is a robust literature on the relationship between the season of birth and attributes such as family socioeconomic status (Buckles and Hungerman 2013), but these seasonal differences do not show up between December and January births, the discontinuity we exploit in our study (Wingender and LaLumia 2017).

4. US Census documents show that the take-up rate among eligible taxpayers without qualifying children was 64% in tax year 2017 (Jones 2020), a rate that has been fairly consistent over time (Plueger 2009). Guyton et al. (2016) show that about three-quarters of potentially eligible nonfilers have qualifying children.

5. Bhargava and Manoli (2015) find that additional information can increase EITC participation among those who had already been notified that they failed to claim benefits; about 15% of unclaimed credits were taken up due to that intervention. Clemens and Wither (2021) find complementary evidence that low-wage individuals face frictions to adjusting their labor supply in response to moderate changes to their budget constraints.

6. Leigh (2010) analyzes employment and wage changes using staggered state EITC expansions and differences in demographics during the federal EITC expansion and finds wage losses of 2%–5% for both eligible and noneligible workers.

7. Moreover, if the EITC causes wages to fall, workers may choose to use the additional posttax income to consume nonwage job attributes such as schedule flexibility. See Clemens (2021), who shows that models incorporating nonwage attributes of jobs can substantially alter the conclusions of incidence analyses of minimum wages, with similar implications for analyses of a wage subsidy like the EITC.

8. See <https://www.irs.gov/credits-deductions/individuals/earned-income-tax-credit/qualifying-child-rules> for a full explanation of the rules regarding “qualifying children.”

9. The EITC for adults with no dependents and a tax filing status of married filing jointly begins its phase-out at a higher level of earnings, \$15,920. The EITC phases out completely for married filers earning more than \$22,610.

10. The SIPP was redesigned in 2014 as an annual survey.

11. About 4% of the sample has different birth dates listed at different points in the sample. Our results are unaffected by using the youngest or oldest listed ages or dropping those individuals from the data.

12. More information on the TAXSIM can be found at <http://taxsim.nber.org>.

13. Carr, Moffitt, and Wiemers (2020) show that imputed values in the SIPP can lead to problems with inference. We also estimate our specifications excluding those with imputed labor force participation values; the conclusions are unchanged.

14. An alternate approach is to use age 26 as a cutoff rather than following-year outcomes for those who are age 25. The conceptual approach is the same and the samples have a good deal of overlap. Figures 4 and 5 show no obvious discontinuity at age 26. The estimated discontinuity is very small: -0.0027 ($SE = 0.020$) and 0.0022 ($SE = 0.022$) for labor force participation and employment, respectively.

15. To examine whether there are different effects of labor-supply incentives when labor demand changes with the business cycle, we also estimate the discontinuity for the three panels separately. None are statistically significant.

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