# Public Health Insurance of Children and Parental Labor Market Outcomes

Konstantin Kunze*

November 12, 2022

**Abstract**

This paper exploits variation resulting from a series of federal and state Medicaid expansions between 1979 and 2014 to estimate the effects of children's access to public health insurance on the labor market outcomes of parents. The results imply that the extended Medicaid eligibility of children leads to positive parental labor supply responses at the extensive and intensive margins. The analysis of mechanisms suggests that Medicaid is less likely to work through marital and educational outcomes and that the effects are driven by the head of the family. The findings also illustrate the importance of siblings' spillovers of Medicaid eligibility for program take-up of children and parental outcomes.

**Keywords:** Labor Supply, Medicaid, Simulated Eligibility, Spillover Effects
**JEL Codes**: I13, I18, I38, J18, J21, J22

*Ph.D. Candidate, Department of Economics, UC Davis; kkunze@ucdavis.edu.

I am deeply indebted to Marianne Bitler and Marianne Page for their invaluable support and guidance. This paper has benefited from the comments and feedback of Gonzalo E. Basante-Pereira, Matthew Bombyk, Colin Cameron, Scott Carrell, Grace Cooper, Sarah Flood, Christian vom Lehn, Santiago Pérez, Monica Singhal, Geoffrey Schnorr, Ivan Strahof, Jenna Stearns, Kari Williams, and seminar participants at the University of California Davis. I thank Sarah Cohodes, Ammar Farooq, Linda Giannarelli, Tal Gross, Xavier Jaravel, Adriana Kugler, Lara Shore-Sheppard, Kendall Swenson, Likun Tian, and Laura Wherry for sharing data used in some analyses of this project or insights about institutional knowledge.

# 1 Introduction

Medicaid, one of the largest government programs in the United States, provided public health insurance to 37.5% of all children ages 0 to 18 in 2019 (CPS, 2022). The literature demonstrates that access to public health insurance during childhood leads to positive short- and long-term effects on their outcomes (Buchmueller et al., 2016). Existing work also documents possible family spillovers of child's access to health insurance on parental outcomes (Aouad, 2021). Hence, it is natural to ask whether and to what extent children's access to Medicaid can affect parental outcomes.

In this paper, I answer this question and study the effects of extended child Medicaid eligibility on parental contemporaneous and long-term labor market outcomes between 1979 and 2014. During the analysis period, Medicaid was expanded, increasing access to health insurance service for many low-income children. However, there was substantial variation in Medicaid eligibility by state, year, and age of children. I exploit this variation using the simulated eligibility strategy first developed by Currie and Gruber (1996a,b) and Cutler and Gruber (1996). This approach uses only legislative variation in public health insurance generosity specific to state, year, and child's age, abstracting from characteristics of the children or family that may be correlated with both Medicaid eligibility and the outcome of interest. To account for the eligibility of each child in the family and address differences in eligibility across race and ethnicity groups, I use a family-level and race-ethnicity-specific simulated eligibility measure.

The main analyses are based on the Annual Social and Economic Supplement to the Current Population Survey from 1980 to 2015 and Decennial Census Samples and the American Community Survey from 1990 to 2010. Using these data, I estimate the effect of children's access to Medicaid on the contemporaneous and long-term labor supply of parents (hours worked per week, weeks worked per year, labor force participation, and earnings). To understand the mechanisms of estimated effects on parental labor market

outcomes, I examine Medicaid take-up of all children in the family, as well as parental marital and educational outcomes. Importantly, in addition to the direct impacts of extended Medicaid eligibility on program take-up, I analyze how siblings' eligibility affects the child's own coverage. Understanding the role a sibling's eligibility plays on a child's own program participation is critical to reconciling the magnitude of parental labor supply responses since sibling spillovers can augment the direct effects of eligible children.

The relationship between children's access to Medicaid and parental labor market outcomes is important from the perspective of the children and the policy maker. Children's well-being crucially depends on material resources and time investments from their parents. Since parents invest financial means and their time into raising their children, parental labor supply is an important factor in the cognitive and physical development of children (Heckman and Mosso, 2014). From the perspective of the policy maker, on the one hand, increased labor supply can recoup some of the costs associated with implementation of Medicaid through increased tax revenue and on the other hand, negative labor supply responses may have implications for program design.

Empirical analysis is necessary since there are many potential mechanisms through which children's Medicaid eligibility can affect the labor market outcomes of their parents. On the one hand, access to Medicaid can improve the health of children and hence lead to increased parental labor supply (Eriksen et al., 2021). On the other hand, extended Medicaid eligibility, which effectively translates into an increase in income, can result in reduced parental labor supply because parents face lower financial burden associated with uninsured children or out-of-pocket expenses for private insurance (Gross and Notowidigdo, 2011). Given that Medicaid is a means-tested program, parents might also only adjust the intensive margin to qualify for Medicaid coverage of their children.[1]

I first provide empirical evidence on first-order effects of children's Medicaid coverage.

[1] Pei (2017), however, does not find evidence for strategic labor supply adjustments of parents in order to gain Medicaid eligibility for their children.

These estimates are important in understanding the mechanism for changes in parental outcomes, because expanded eligibility translates into improved health of children or reduced financial distress through a corresponding increase in program take-up. The results suggest that extended child Medicaid eligibility leads to a family-level marginal take-up rate of 32%. I also find that 30% of a child's own coverage is driven by the extended eligibility to other children in the family emphasizing the importance to account for all eligible children in the family.

I then show that expanding Medicaid eligibility to children positively affects contemporaneous maternal labor market outcomes entirely driven by mothers with non-white and/or Hispanic children. Standard labor supply measures of mothers with non-white and/or Hispanic children increase contemporaneously by 2.9% to 5.3% relative to baseline average labor supply. The elasticities of point estimates range between 0.03 and 0.05. In the long run, the effects persist up to 18 years, and each additional year of eligibility during childhood leads to an increase in labor supply of 0.59% to 0.96% for women. These estimates represent between 15% and 30% of contemporaneous labor supply responses.

The analysis of potential mechanisms indicates that the family structure may be an important factor contributing to race-ethnicity differences in labor supply responses since the effects are concentrated among non-white and/or Hispanic single mothers. These findings suggest that the head of the family is more likely to be affected by extended Medicaid eligibility. Using estimates from existing literature on effects of marital and educational outcomes on labor supply, I show that changes in marriage and education are not likely to be driving maternal labor market outcomes.

This study contributes to three strands of literature. First, it connects to the literature analyzing the effect of child's access to Medicaid on labor supply responses of parents. Earlier work by Ham and Shore-Sheppard (2005a) replicated the study by Yelowitz (1995) by incorporating important institutional features and estimated imprecise effects on parental labor supply using the eligibility of youngest child in the family. A more re-

cent study by Grossman et al. (2022) is the only other paper besides my own that accounts for the Medicaid eligibility of all children in the family and documents negative maternal labor supply responses. In comparison to my paper, they focus on different cohorts (parents born between 1957 and 1964) because they relied in part on data from the National Longitudinal Survey of Youth – 1979 Cohort.

I advance the literature in several ways. Using a larger data set and more cohorts, I focus on the eligibility of all children in the family and empirically quantify the importance of incorporating the eligibility of each child in the family by comparing the estimated effects of one with multiple eligible children in the family. In addition to contemporaneous parental labor supply responses, I analyze if the labor market outcomes of parents are affected up to 18 years after Medicaid eligibility of children was initially expanded, which was not previously explored in the literature. Finally, I use a simulated eligibility measure appropriate for heterogeneous analysis by race and ethnicity. Using a simulated eligibility approach that fails to account for racial and ethnic differences will result in biased estimates since each race-ethnicity group will be assigned an incorrect simulated eligibility.

Second, my findings relate to studies that examine the effects of expanded Medicaid eligibility on enrollment spillovers.[2] The majority of the literature focuses on spillovers from parents to children, except Sommers (2006) who provides suggestive evidence on the relationship between child's actual eligibility and Medicaid disenrollment of siblings and parents. Relative to these studies, I examine if Medicaid take-up of a potentially eligible child is affected by the eligibility of other children in the family using a simulated eligibility approach that accounts for the endogeneity of actual eligibility. Documenting the role of a sibling's eligibility on child's own take-up is important because it captures all indirect effects of access to Medicaid.

[2]See Aizer and Grogger (2003); Dubay and Kenney (2003); Sommers (2006); Sonier et al. (2013); Sommers et al. (2016); Frean et al. (2017); Hudson and Moriya (2017); Hamersma et al. (2019); Sacarny et al. (2022)

Third, this study adds to the strain of literature that employs the simulated eligibility approach.[3] There is a substantial difference across the studies in the method used to construct simulated eligibility, which can be summarized based on two main criteria: type (e.g., fixed or annual simulated eligibility) and structure (e.g., state, year, and age or state, year, age, and race-ethnicity). I show that the results are not sensitive to the choice of the simulated eligibility type, but may be sensitive to the structure of the eligibility measure. Moreover, the analysis demonstrates the importance of the correct simulated eligibility structure when the goal is to capture group-specific effects and eligibility is likely to be different across these groups (e.g., race and ethnicity).

The rest of the paper is structured as follows. Section 2 provides the history and evolution of the Medicaid program. I introduce the simulated eligibility measures and the empirical approach in Section 3. Data sources and sample selection are described in Section 4. Section 5 discusses the results, and Section 6 provides the conclusions.

# 2 History of the Medicaid Program

Medicaid is a joint state and federal program that was signed into law in 1965 as Title XIX of the Social Security Amendments.[4] In the beginning of the analysis period Medicaid eligibility for non-disabled children was originally restricted to single-parent families receiving cash welfare payments under the Aid to Families with Dependent Children (AFDC) program or eligibility under three additional state optional programs.[5] The in-

[3] Among others, Currie and Gruber (1996a,b); Cutler and Gruber (1996); Zavodny and Bitler (2010); Gross and Notowidigdo (2011); Cohodes et al. (2016); East et al. (2017); Brown et al. (2019); Miller and Wherry (2019).

[4] The focus of this review is legislation targeted at the eligibility of children between 1979 and 2014. Table B.1 provides an overview of federal and state legislation for this period. Sources for this section include Gruber (2003) and Buchmueller et al. (2016). See appendix section B for a detailed explanation of legislative changes during the analysis period.

[5] AFDC-Unemployed Parent (AFDC-UP) program extended Medicaid eligibility to families with an unemployed primary earner, the Ribicoff Children program provided Medicaid eligibility to children who did not meet the family structure requirements but met the income and resource requirements for AFDC, and the Medicaid's Medically Needy program extended Medicaid eligibility to those with high medical expenses whose income exceeded the maximum threshold but family structure satisfied the AFDC requirements.

come eligibility thresholds varied by state and family size, most of which were well below the federal poverty line (FPL). These stringent eligibility requirements meant that only a few children of working mothers were eligible for Medicaid and if a women was to leave welfare her child would not be covered by Medicaid. Hence, in order for children to remain eligible for Medicaid, mothers were given the incentive not to participate in the labor force and cut their working hours.

In the mid-1980s legislation started to gradually separate Medicaid and AFDC by expanding eligibility to children not qualifying for AFDC. Initially Medicaid eligibility was extended for children under five years of age who were born after September 30, 1983 and who were living in families that met the financial, but not the family structure requirements for AFDC through the 1984 Deficit Reduction Act. Omnibus Reconciliation Act (OBRA) 1986 and 1987 as well as Medicare Catastrophic Coverage Act and Family Support Act further weakened the link between Medicaid and AFDC by allowing and requiring states to increases the income limits for Medicaid eligibility for children belonging to certain age and birth cohorts.

Additional federal expansions were embedded in OBRA 1989 and 1990 - until then the largest expansions in US history. OBRA 1989 expanded Medicaid eligibility to pregnant women and children up to age six with family incomes below 133% of the federal poverty line and OBRA 1990 required states to cover children born after September 30, 1983 with family incomes below the federal poverty line. These children remained eligible until the age of 18. By introducing Temporary Assistance for Needy Families (TANF) program, Personal Responsibility and Work Opportunity Reconciliation Act of 1996 removed the link between AFDC and Medicaid completely since TANF did not provide Medicaid eligibility automatically. However "Section 1931 eligibility" required states to cover families that would have been eligible under AFDC before the welfare reform. The next milestone in the evolvement of the Medicaid program was the Balanced Budget Act (BBA) in 1997. BBA created the State Children's Health Insurance Program (SCHIP), allowing states to

cover uninsured children in families ineligible for Medicaid and providing continuous coverage for up to twelve months regardless of increases of child's family income.

The policy changes between 1979 and 2014 had a large effect on Medicaid eligibility of children. Figure 1 documents that the fraction of eligible children increased substantially from 0.14 in 1979 to 0.55 in 2014 with the biggest increase in eligibility around the introduction of SCHIP in 1997. The Medicaid expansions affected race-ethnicity groups differently. Throughout the analysis period, the fraction of non-white and/or Hispanic children eligible for Medicaid is higher than the fraction of white non-Hispanic children eligible for Medicaid. Moreover, the increase in eligibility of non-white and/or Hispanic children was stronger after legislations targeted at lower-income families (e.g., OBRA 1990). The difference in eligibility between race-ethnicity groups underlines the importance to use a race-ethnicity-specific eligibility measure in the heterogeneous analysis.

# 3 Methodology

## 3.1 Simulated Eligibility

Sociodemographic characteristics can affect the number of children who are eligible for Medicaid independent of legislative changes as well as outcomes of children and their parents resulting in an endogenous measure of children's actual Medicaid eligibility. For instance, improved economic conditions may increase average income for certain groups of the population and hence reduce the number of children who are income eligible for public health insurance. At the same time, changes in economic environment may also affect outcomes of parents and their children. To address this type of potential endogeneity, I follow the simulated eligibility approach first developed by Currie and Gruber (1996a,b) and Cutler and Gruber (1996). The goal of the simulated eligibility strategy is to create a measure of eligibility abstracting from omitted variables that may be correlated with both children's actual eligibility and parental or child outcomes, so that identification is based only on legislative variation.

I use three simulated eligibility measures in the analysis - contemporaneous child- and family-level as well as long-run family-level eligibility measure.[6] To construct the child's own simulated eligibility measure, I use all children of age 0-18 in each year of the analysis period. Using this national data set, I calculate the child-level simulated eligibility as the fraction of eligible children in each state, year, age, and race-ethnicity group by leaving out children from the state for which the eligibility is being imputed.[7]

Race and ethnicity plays an important role in the analysis of Medicaid expansions.[8] On the one hand, white and non-white children might respond differently to the same level of simulated eligibility, which can be captured by analyzing heterogeneous responses across race-ethnicity groups. On the other hand, white and non-white children are distinguished by different levels of simulated eligibility because of systematic differences in characteristics relevant for eligibility determination resulting in a measurement error in simulated eligibility if not accounted for. Figure 6 explains the importance of a race-ethnicity-specific simulated eligibility measure for the analysis. Subfigure (a) shows that the race-ethnicity-specific simulated eligibility traces out the actual eligibility very well for both race-ethnicity groups. The non-race-ethnicity-specific simulated eligibility, however, is not close to the actual eligibility of any race-ethnicity group shown in subfigure (b). The simulated eligibility measure addresses this concern by allowing the simulated eligibility measure to be race-ethnicity specific.

Since the primary focus is on parental outcomes, one has to account for the eligibility of each child in the family. Using eligibility of only one child in a family might underes-

[6] Appendix B describes in detail the construction of simulated eligibility measures, how eligibility is imputed, and the legislative rules used to impute eligibility. Medicaid eligibility is determined using calculators from Miller and Wherry (2019).

[7] The child-specific simulated eligibility measure is in spirit of earlier literature that examines the effects of child Medicaid eligibility on child's health insurance coverage and health outcomes (e.g., Cutler and Gruber 1996; Currie and Gruber 1996a,b). The main difference to the eligibility measure used in these studies is the race-ethnicity component. Only few studies have used a race-ethnicity-specific simulated eligibility measure (e.g., Dave et al. 2015 and Cohodes et al. 2016).

[8] I define race-ethnicity categories as white non-Hispanic and non-white and/or Hispanic. For simplicity, in the remaining part of the text I will refer to race-ethnicity groups as white and non-white.

timate the effects of extended Medicaid eligibility because having multiple children in the family eligible for Medicaid might affect parental outcomes in a way that is not fully captured by the eligibility of a single child. To construct family's total simulated eligibility, I sum the simulated eligibility fractions across all children in a family which mimics the number of eligible children in a family. The total simulated eligibility measure therefore ranges from 0 to the maximum number of children in a family and is on average 0.65 eligible children per family.

To estimate the long-run effects of children's access to Medicaid and to account for eligibility of each child in the family, I construct a family-level long-run simulated eligibility that captures the number of eligible children in the family during childhood.[9] I obtain the long-run eligibility measure by summing the average total simulated eligibility at each age from birth to the current age of the child for a given birth cohort, age, state, and race-ethnicity group. The long-run simulated eligibility therefore reflects the total child-years of simulated eligibility experienced by a parent and ranges from zero to the number of total child-years of simulated eligibility times the average number of children.

## 3.2 Empirical Approach

I estimate effects of increased Medicaid eligibility on the insurance coverage of children and outcomes of their parents (labor market outcomes, educational attainment, marital outcomes) by running a child-level regression and regressing the outcome of interest on simulated eligibility measures as well as a set of controls. The regression is estimated at the child level because a child-level regression facilitates controlling for child-specific characteristics and allows me to estimate the effects of sibling's eligibility on

[9]The weakness of the long-run simulated eligibility is the lack of information about number of family members between the birth of the fist child and survey date in the publicly available data. Hence it is not possible to predict the exact number of eligible children during childhood and the effects should be interpreted as the average exposure to Medicaid of all children in the family. Moreover, this approach may be problematic if children's Medicaid eligibility affects parental fertility. Existing evidence, however, suggests that access to Medicaid does not affect fertility (e.g., Zavodny and Bitler 2010; DeLeire et al. 2011; East et al. 2017).

child's own Medicaid coverage. In general child-level and parent-level empirical models would provide the same results if the same controls are used at the parent level (including children's ages) and concordant weights are used. Specifically, I begin by estimating the model of the following functional form:

$$y_{ijstr} = \beta_0 + \beta_1 SIMT_{jstr} + X'_{istr}\beta_2 + W'_{jst}\beta_3 + Z'_{st}\beta_4 + \delta_s + \gamma_t + \tau_a + \varepsilon_{ijstr} \quad (1)$$

where the dependent variable is either Medicaid coverage of child *i* or an outcome of parent *j* of child *i* in state *s*, calendar year *t*, and race-ethnicity group *r*. For outcomes of children (e.g. child own Medicaid coverage), equation 1 does not have a *j* component. $SIMT_{jstr}$ is the total simulated Medicaid eligibility - the treatment variable of interest. From the policy maker's perspective, the coefficient $\beta_1$ on total simulated eligibility captures the effect of an additional child in the family becoming eligible for Medicaid.[10]

Equation 1 includes state of residence fixed effects ($\delta_s$) which capture fixed differences in outcomes of parents and their children across states, calendar year fixed effects ($\gamma_t$) to account for potential national changes over time, and child's age fixed effects ($\tau_a$) to account for fixed differences in outcomes of parents and their children across children's ages. Alternative specification add state-by-year, state-by-age, and year-by-age fixed effects to control for omitted variables at the state-year, state-age, and year-age level, respectively. I also include child's sex, race, and ethnicity defined by the vector $X_{istr}$. The vector of parent-level control variables, $W_{jst}$, contains fixed effects for parental age, num-

[10]The parameter of interest using the long-run family-level eligibility measure (SIMC) reflects the effect of making one more child per family eligible for one more year during childhood and contemporaneous child-level simulated eligibility measure (SIM) captures the effect of one child in a family becoming eligible for Medicaid.

ber of children, age of the youngest, and age of the oldest child in the family.[11] In the baseline model, I follow the literature and include a vector of annual state-specific economic and policy characteristics, $Z_{st}$, which incorporates the unemployment rate, the minimum wage, inflation-adjusted maximum welfare benefits for a family of 4, state-level EITC amounts measured as a percentage of the federal EITC, implementation of six types of welfare waivers, and implementation of any waiver or TANF.[12] In specifications for the full sample, all control variables and fixed effects are interacted with a race-ethnicity group indicator.

When the dependent variable is at the child level, I use the corresponding weights of children. For outcomes of parents, the regression is weighted by parental weights divided by number of children in the family because the outcome is at the parent level but the observations are at the child level. Using parental weights divided by number of children makes the results representative of the average parent. Standard errors are clustered at the state level.

The specification shown in equation 1 exploits plausibly exogenous variation that results from the state and federal Medicaid expansions during the analysis period. There are three main sources of variation - across states because of state differences in AFDC eligibility limits prior to the expansions and difference in state's implementation of optional expansions, over time as the expansions were implemented with different pieces of legislation, and across children's age since younger children are more likely to be eligi-

[11] Since child *i* can be the youngest or oldest child in the family, including child's age fixed effects as well age of the youngest, and age of the oldest child in the family fixed effects could result in two duplicate variables in the regression controlling for the age of the child. Hence, to account for child *i* being the youngest or oldest child in the family, the age of youngest child in the family fixed effects are interacted with an indicator for not being the youngest child in the family and the age of the oldest child in the family fixed effects are interacted with an indicator for not being the oldest child in the family. Similarly, to make sure that only one variable capturing age of the child is included in the regression for single-child families, age of the oldest child in the family fixed effects are interacted with an indicator for multiple children in the family.

[12] Parental educational attainment and marital status are not included as controls in main specification because these characteristics may and do respond to simulated eligibility as shown in table A.2 and A.3. However, estimated effects from models including marital and educational outcomes as covariates are very similar to the main specification.

ble.[13] Figures 2 and 3 summarize the underlying variation. Figure 2 shows the difference in total simulated eligibility between 1979 and 2014 for each state for single-child and multiple-child families. While Medicaid eligibility increased over time, there is substantial heterogeneity across states. In some states Medicaid expansions resulted in an average increase in simulated eligibility on the order of two simulated eligible children per family whereas in other states on the order of only 0.2 simulated eligible children per family. This pattern is quite similar across single-child and multiple-child families implying that differences in the number of children per family is not driving the difference in simulated eligibility. Figure 3 shows the national trend in simulated eligibility between 1979 and 2014 by child's age for single-child and multiple-child families. All age groups saw a substantial increase in simulated Medicaid eligibility between 1979 and 2014. While families with younger children were mostly affected during the first half of the analysis period, families with older children were affected during the second half of the analysis period suggesting that the Medicaid program became more generous for older children over time.

Two general difference-in-difference-in-difference identifying assumptions in equation 1 are invoked for the validity of the empirical approach. The first identifying assumption is that no shock differentially affects Medicaid generosity and outcomes of children and their parents in the same state, during the same year, and with the same number of children of the same age. Hence omitted variables specific to parents with the same number of children of the same age and state of residence that change over time and are correlated with both Medicaid legislation and outcomes of children or their parents would invalidate this empirical strategy. To address this potential confounder, I estimate a version of equation 1 with state-by-age linear time trends or region-by-year-by-age fixed effects

[13] Changes in children's characteristics are also contributing to the identifying variation since the simulated eligibility measures are constructed using data from the year for which the eligibility is imputed. Results are, however, robust to alternative eligibility measures that abstract from this possibly non-exogenous part of the identifying variation (see section 5.4.2).

as well as flexible controls for omitted variables at the state-year, state-age, and age-year level.[14]

The second identifying assumption requires that public health insurance eligibility rules are not set based on outcomes of parents and their children. The simulated eligibility approach will therefore fail if states phase in Medicaid expansions because of changing trends in parental or child outcomes. To test the validity of this identifying assumption, I regress the maximum Medicaid eligibility limits for children age 0-18 in a given state and year on contemporaneous and lagged (first and second order) state-level characteristics.[15] I use Medicaid eligibility limits as opposed to state-level annual simulated eligibility since state governments set eligibility levels and do not control simulated eligibility as a policy parameter. The results shown in table 1, suggest that generosity of Medicaid is not affected by outcomes of parents and their children or other state-level policy determinants. Baughman and Milyo (2009) also show that state Medicaid expansions are not driven by percentage of uninsured children in the state and Farooq and Kugler (2020) find no evidence that state demographic and economic characteristics are affecting Medicaid generosity. In comparison to Baughman and Milyo (2009) and Farooq and Kugler (2020), I show that state-level policies discussed in the literature (e.g., Miller and Wherry 2019) have no effect on Medicaid expansions. In addition, the second identifying assumption has been invoked repeatedly in the simulated eligibility literature.[16]

[14] Estimated effects from specifications including state-by-age linear time trends and region-by-year-by-age fixed effects (table A.15 and A.16) and state-by-year, state-by-age, and age-by-year fixed effects (figures A.6-A.10) are very similar in terms of magnitude and precision to the baseline model corroborating the validity of the first identifying assumption.

[15] For the years prior to state expansions (1979-1987), I use the maximum eligibility threshold for Medically Needy Program or AFDC. As an approximation for AFDC eligibility, I use the average ratio of the needs standard to the corresponding poverty guideline across all family sizes. For the years 1988-2014, I use the maximum state-level Medicaid eligibility levels across all ages 0-18. Since states expanded eligibility for different age groups, it is more consistent to use eligibility limits across a broad age group and not focus on narrow defined age groups (e.g., children age 0-5).

[16] See for instance Currie and Gruber (1996a,b); Cutler and Gruber (1996); Gross and Notowidigdo (2011); Cohodes et al. (2016); East et al. (2017); Brown et al. (2019); Miller and Wherry (2019)

# 4 Data

## 4.1 Current Population Survey

To analyze contemporaneous parental labor market responses, child's insurance coverage, and to predict Medicaid eligibility I use data from survey years 1980 to 2015 of the Annual Social and Economic (ASEC) supplement to the Current Population Survey (CPS) obtained from the integrated public use microdata series (Flood et al. 2020). The CPS is a nationally representative survey interviewing approximately 60,000 households per month. The ASEC supplement provides a comprehensive body of data containing information on individuals' demographic characteristics, employment, health insurance coverage, and income. The ASEC supplement provides information on the family composition, educational attainment and demographic characteristics at the interview date, health insurance coverage at any time during the previous calendar year, income during the previous calendar year, and labor supply measures either with reference to last week or previous calendar year.[17] Hence, simulated eligibility is measured contemporaneously to the reference period of the outcome variable of interest. With respect to labor market responses, I analyze usual hours worked per week, labor force participation last week, weeks worked last year, annual earnings last year, and occupational choice last year.[18] Medicaid coverage of children is captured by number of covered children per family or child's own Medicaid coverage. In supplementary analysis, I examine parental educational attainment (no high school, high school, some college, college or more) and marital outcomes (married, never married, ever married, divorced).[19] Since the eligibil-

[17] When answering questions about health insurance coverage, some respondents might ignore the reference period and instead answer based on their status at the time of the interview (Klerman et al. 2009, Ziegenfuss and Davern 2011). The results are, however, very similar when the simulated eligibility is assigned as of last year or as of interview date.

[18] Hours worked per week and weeks worked per year include zeros. I mainly focus on labor supply as of previous year. However, labor force participation is only measured as of last week.

[19] CPS redefined the education variable from years of education to degree receipt in 1992. To attain comparable educational categories across the whole analysis period, I use the method proposed by Jaeger (1997).

ity calculator is able to impute Medicaid eligibility for children age 0-18 and to capture working age parents, the sample is restricted to children age 0-18 with parents age 20-64.[20]

## 4.2 Decennial Census & American Community Survey

To examine long-run effects on parental outcomes, I utilize the 5% sample of the 1990 and 2000 Decennial Census (DCS) and the 2010 American Community Survey (ACS) obtained from the integrated public use microdata series (Ruggles et al. 2020). ACS and DCS are nationwide surveys conducted by the U.S. Census Bureau. The Census Bureau collects information on social, economic, housing, and demographic characteristics. The advantage of ACS and DCS for my analysis is the information about the birth state of the child which allows me to estimate long-run effects of access to Medicaid during childhood on parental outcomes from the birth of the child until the child is observed in the survey. ACS and DCS collect information on family composition and demographic characteristics at the interview date, income during the past 12 months, and labor market outcomes during the previous calendar year with exception of labor force participation which is reported as of last week. Similar to CPS ASEC, the sample is restricted to children age 0-18 with working age parents between 20 and 64 years old.[21]

## 4.3 Supplemental Data

I use supplemental data from various sources. State-level minimum wage, state-level welfare benefits, and state-level Earned Income Tax Credit (EITC) amounts are obtained from U.S. Department of Labor, Urban Institute, and Tax Policy Center respectively. State-level unemployment rate, Consumer Price Index, compensation of employees, and num-

[20] Tables A.11-A.14 show that results are not sensitive to alternative sample selection criteria by dropping children from Arizona (Arizona did not adopt a Medicaid program until 1982), restricting the sample to children with parents in prime working age (25-54), dropping children observed in 2008-2015 to account for Great Recession and the introduction of Affordable Care Act, keeping only children with mothers that gave birth at reproducible age (15-44), and dropping children in families with nine or more children.

[21] Table A.1 shows child-level and parent-level demographic characteristics of ACS, DCS, and CPS ASEC samples.

ber of total non-farm employees come from the U.S. Bureau of Labor Statistics. I obtain federal poverty guidelines and information about implementation of state welfare waivers from the Office of the Assistant Secretary for Planning and Evaluation. Data on Medicaid spending comes from the Medicaid Statistical Information Statistics (MSIS) maintained by Centers for Medicare & Medicaid Services.[22]

# 5 Results

## 5.1 Medicaid Coverage

I first analyze if extended Medicaid eligibility translates into increased Medicaid coverage since program take-up is an important channel in understanding the relationship between child Medicaid eligibility and parental labor market decisions. Table 2 presents results for the estimated effects of race-ethnicity-specific simulated eligibility on public health insurance coverage of children. The first three columns provide the estimated effects of child's own eligibility (SIM) on child's own Medicaid coverage. The estimates indicate that the average marginal take-up rate among children who became eligible over the 1979-2014 period is eight percent. In columns 4-6 in addition to child's own Medicaid coverage, I include siblings' eligibility (SIMS) which is obtained by summing the child's own simulated eligibility across all siblings. The results suggest that the elasticity of the estimated treatment effect of siblings' eligibility on the children's own coverage is roughly 30% of the elasticity of the effect of child's own eligibility emphasizing the importance to account for eligibility of each child in the family. The last three columns show estimates based on a model where the dependent variable is the number of covered children in a family and the independent variable (SIMT) is the total simulated eligibility.

[22] Since children's Medicaid cost is at the state and year level for all children under 21, I disaggregate the cost to obtain a parent-level measure. First, I allocate the aggregated cost to each age based on the fraction of eligible children of a given age and divide by the total population in a given state, year and age. The child-level per capita cost of each child in a family is then summed up to obtain total Medicaid cost per family. MSIS data is only available for the 1980-2012 period. Hence the sample that uses MSIS data is restricted to 1981-2013.

The point estimate implies that the number of covered children per family increases by one-third of a child as a result of one more child per family becoming eligible which is equivalent to a take-up rate of 32% among newly eligible children or an elasticity of 0.47 $(0.32 * 0.65 \div 0.44)$. The family-level take-up rate is higher than the child's own take-up rate because the sibling spillovers of eligibility contribute considerably to Medicaid take-up.

To compare my estimates to the existing literature and check the robustness to imputation and methodological changes in CPS, I estimate the effect of child- and family-level Medicaid eligibility on Medicaid coverage using a non-race-ethnicity-specific simulated eligibility and harmonized measure of Medicaid coverage constructed by State Health Access Data Assistance Center (SHADAC).[23] The estimates are quantitatively and qualitatively very similar to the baseline model and indicate that the average marginal take-up rate among children who became eligible over the analysis period is ten percent which is in line with the existing literature (Buchmueller et al., 2016).[24]

Next, I examine racial and ethnic differences in Medicaid coverage. The elasticities of the effects of children's own simulated eligibility on children's own coverage are larger for white children. One potential explanation are the greater barriers to Medicaid enrollment (e.g., insufficient knowledge about the programs, confusion about the eligibility, difficulties with the application) for non-white children (Stuber and Bradley 2005).[25] Sibling spillovers of eligibility play a more important role in child's own Medicaid coverage for non-white children. The elasticity of the effect of siblings' eligibility is roughly 60% of elasticity of children's own eligibility for non-white children compared to 12.5% for white

[23] For 1979-1986, CPS imputed health insurance for children age 0-14. In addition, during the analysis period the collection of health insurance coverage in CPS underwent multiple methodological changes (SHADAC 2009).

[24] The results of this exercise are shown in table A.4. The top panel shows estimates from models using non-race-ethnicity-specific simulated eligibility for 1979-2014 and bottom panel shows the estimates from specifications using race-ethnicity-specific simulated eligibility and Medicaid coverage enhanced by SHADAC for 1987-2014.

[25] The average number of children per family as well as the distribution of number of children per family is very similar in white and non-white families. On average white (non-white) parents have 1.8 (1.9) children. Similarly, 56% and 57% of white and non-white families have more than one child.

children. This finding underlines the larger importance of potential knowledge spillovers or reduced costs of application for non-white children. Once the eligibility of each child in the family is taken into account, the take-up of non-white children (elasticity of 0.57) is larger and statistically different ($p<0.05$) than the take-up of white children (elasticity of 0.4).

## 5.2 Labor Market Outcomes

Table 3 presents the estimated treatment effects of race-ethnicity-specific total simulated eligibility on labor market outcomes of mothers with any, white, and non-white children.[26] The estimates for maternal labor force participation and weeks worked per year are positive and significant at one and five percent level. The estimated effects suggest that increasing the number of simulated eligible children in the family by one child leads to an increase of 0.38 weeks worked per year (1.5% relative to the baseline average weeks worked per year of 25.41) and one percentage points increase in labor force participation (1.8% relative to the baseline average labor force participation of 0.57). The estimates for maternal usual hours worked per week are not statistically significant at conventional levels. The results reveal substantial differences between race-ethnicity groups. We see that the estimated effects are entirely driven by mothers with non-white children. Usual hours worked ($p < 0.05$), labor force participation ($p < 0.01$), and weeks worked per year ($p < 0.01$) of mothers with non-white children increase by 2.9-5.3% whereas labor supply responses of mothers with white children are close to zero and not precisely estimated.[27] In addition, the estimates on usual hours worked per week, weeks worked per year, and labor force participation are statistically different between white

[26] Results are very similar when race and ethnicity of parents is used. However, since the race-ethnicity-specific simulated eligibility measure is created using child's race and ethnicity, it is more intuitive to use race and ethnicity of children for heterogeneous analysis. In addition, only 1.5-3% of families are racially and ethnically mixed.

[27] The effects on hours worked are driven by working any hours and increasing full-time employment ($\geq$ 35hours) as shown in table A.5. Weeks worked per year are affected across the whole distribution in response to extended Medicaid eligibility (see figure A.1).

and non-white mothers at conventional levels.

The increased labor supply of mothers with non-white children does not translate into higher average earnings - none of the point estimates are precisely estimated.[28] The earnings in the lower part of the earnings distribution are, however, positively affected by extended Medicaid eligibility. Figure 4 shows the effect of Medicaid eligibility on the distribution of maternal earnings.[29] I find a significant increase in density around the minimum eligibility cutoff where between 30% and 40% of non-white children in the CPS ASEC sample are eligible for Medicaid. Effects above the median eligibility limit are small and statistically insignificant. In line with small and imprecise effects on labor supply of mothers with white children, the effects on earnings are also small and not statistically significant at conventional levels.

Turning to the long-run labor supply responses, table 4 shows the effect of total child-years of simulated eligibility experienced by a parent on usual hours worked per week and labor force participation. As in the the contemporaneous analysis, the effects on maternal labor supply are driven by mothers with non-white children. The results imply that making one more child eligible for one more year leads to an increase of 0.22 ($p < 0.01$) hours worked per week and 0.004 ($p < 0.01$) percentage points in labor force participation. These long-run estimates represent roughly 30% and 13% of contemporaneous effects on usual hours worked per week and labor force participation, respectively.

[28] Shown in table A.17 the average earnings estimates are not driven by nonresonse biases or disclosure avoidance methods. To check for nonresonse biases, I follow Hirsch and Schumacher (2004) and Bollinger and Hirsch (2006) and drop imputed nonrespondents and reweigh the sample with inverse-probability weights to restore population representatives. I apply cell mean replacement topcodes introduced in 1996 from Larrimore et al. (2008) and rank proximity swap topcodes used starting in 2011 from Census Bureau to the earlier period to test if top coding methods are driving the results.

[29] I follow Kuka and Shenhav (2020) and estimate a series of regressions where each dependent variable is an indicator equals to one if earnings are greater than X, where X is (0,5000,...,150000). For reference average minimum, median, and maximum eligibility limits during the analysis period are labeled.

## 5.3 Magnitudes

I show that one additional eligible child in the family leads to a statistically significant increase of 0.90 weeks worked per year of non-white women. The average increase in total simulated eligibility from the beginning to the end of the analysis period represents, however, only 0.70 eligible children per family. To account for the difference of family-level eligibility between 1979 and 2014, I scale the estimated effects on labor supply by the average increase in total simulated eligibility. This corresponds to 0.63 weeks worked per year for non-white mothers. Since not every child enrolls in Medicaid, the results so far should be interpreted as intent-to-treat estimates where treatment is defined as program participation. To convert intent-to-treat (ITT) to treatment-on-treated (TOT) estimates, I divide the estimated effect by the corresponding family-level Medicaid take-up rates of non-white children (0.40). TOT estimates are therefore 1.58 weeks worked per year of non-white women. The scaled estimated effects represent 19.54% of the difference in weeks worked per year between the beginning and end of the analysis period as well as 7.54% of difference in weeks worked per year between non-white mothers with and without some college education. The elasticities of scaled TOT estimates of non-white women range between 0.05 and 0.08.

In order to understand the magnitude of the effect of an additional child per family becoming eligible, I compare the effects of having one eligible child with having multiple eligible children. Table A.6 compares the estimated effects of child- and family-level simulated eligibility on parental labor market outcomes. The point estimates in models using only child's own eligibility measure are not statistically significant at conventional levels (columns 1-3) and once child-level and sibling's total simulated eligibility are included (columns 4-6) the effect of having one additional eligible sibling is twice as big as the effect of child's own eligibility. Therefore, the results of this analysis suggest that labor supply measures are not affected by one eligible child per family emphasizing the

importance to account for each eligible child in the family.

To further put the estimated effects of extended Medicaid eligibility into perspective, I compare the findings to other studies. Ham and Shore-Sheppard (2005a) use the eligibility of the youngest child in the family which provides similar results to the child-level eligibility measure discussed above and find imprecise effects of children's access to Medicaid on parental labor supply. Grossman et al. (2022) account for all eligible children in the family but focus only on parents born between 1957 and 1964 since these birth cohorts are sampled in National Longitudinal Survey of Youth - 1979 Cohort and find that maternal labor force participation decreases as a results of Medicaid expansions. Current work uses a much larger data set with more cohorts that is more suitable to analyze labor supply and shows that the effects are positive once more parental birth cohorts are used.

I also consider literature on early childhood education and childcare since similar to Medicaid, educational services for young children provide incentives for parents to increase labor supply at the extensive and intensive margin. A part of this literature documents positive effects of public schooling for low-income children on maternal employment with an estimated elasticity of 0.34-0.38 (Gelbach 2002; Cascio 2009; Wikle and Wilson 2021).[30] Consistent with these findings, I also find positive effects on maternal labor supply, albeit smaller in magnitude. A potential explanation for the difference in magnitudes is that activities related to child care require a larger time commitment than health related activities (Bastian and Lochner, 2022).

[30] Related literature that examines universal pre-kindergarten programs, finds only weak evidence of employment responses (Fitzpatrick 2010; Cascio and Schanzenbach 2013). This difference could arise since preschool programs for low-income families - a group that is more relevant for my analysis - was available before universal eligibility (Wikle and Wilson, 2021).

## 5.4 Robustness

### 5.4.1 Maternal Eligibility

A small proportion of women were directly affected by extended Medicaid eligibility during the analysis period because some legislations expanded Medicaid to pregnant women and children at the same time. Since Medicaid eligibility for pregnant women was also applicable for their newborns until the first birthday, parents with children age zero might change labor supply as a result of direct effect of extended Medicaid eligibility and not as a result of spillovers from their children. To understand if parental labor market outcomes are driven by direct effects of Medicaid, I test if the estimated effects are sensitive to using maternal eligibility for zero-year old children by using two common measures in the Medicaid literature of maternal eligibility and dropping children of age zero.[31] All estimates across different specifications are very similar in terms of magnitude and significance, suggesting that direct effects of expansions are not driving labor supply responses.

### 5.4.2 Simulated Eligibility Type

Since the simulated eligibility measure used through out the analysis is constructed using all children from the year for which the simulated eligibility is estimated, one might be concerned that characteristics used to determine eligibility (e.g. family structure or family income) may respond to Medicaid expansions.[32] To account for this potential en-

[31] Following Currie and Gruber (1996a,b), the first measure is constructed by using all women of reproducible age (15-44) in each calendar year across the full sample period. Using this national data set, I calculate the fraction of eligible women in each state, year, and race-ethnicity group. The second state-year-race-ethnicity maternal eligibility measure is obtained by using mothers with children of age zero. Similarly to children's simulated eligibility, I leave out women from the state for which the simulated eligibility is being imputed. The maternal eligibility measures are then assigned to zero-year old children based on the state, year, and race-ethnicity group. The results of this analysis are shown in tables A.18 and A.19.

[32] Policy endogeneity can arise because of a response to federal or state-level expansions. For instance, Deficit Reduction Act of 1984 could affect labor market outcomes of parents and hence the family income. Using family income of children observed in years after 1984 to determine eligibility in the same year would result in a biased estimate of the simulated eligibility measure.

dogeneity, I construct alternative simulated eligibility measures that use children from period before the analysis starts. To obtain the simulated fixed eligibility measures, I use all children from 1979 CPS ASEC and inflate the income to the year for which the eligibility is imputed. Changes in the national and regional Consumer Price Index for All Urban Consumers (CPI-U) as well as average wages are used to adjust the income.[33] The demographic characteristics of children in 1979 might however not reflect demographic characteristics of children observed in later years of the analysis period and inflation or wage growth might not fully capture changes in income over time. Using a fixed national data set from pre-analysis period to create simulated eligibility might therefore result in a mismeasured simulated eligibility for later years of the analysis period. Since changes in socio-demographic characteristics are also correlated with changes in the structure of the labor market and hence parental labor market outcomes, using a fixed eligibility measure might as well result in biased estimates.[34] Hence, the annual and fixed eligibility measures have advantages and disadvantages.

Figure 7 documents the differences between the simulated eligibility measures. The changes across the simulated annual eligibility - the measure used throughout the analysis - and actual eligibility track quite well. The trends in simulated fixed eligibility measures, however, deviate from trends in actual eligibility, especially towards the end of the analysis period. The eligibility measures constructed by using CPI-U perform worse than the measure constructed by using average wages. The results are, however, very similar across specification using the different types of the simulated eligibility measure.[35]

[33] Average wages are calculated as the ratio of compensation of employees to total non-farm employees.

[34] Since immigration has a major influence on the size and demographic structure of the US population, immigration can be one factor leading to a change of socio-demographic characteristics of individuals observed in later years of the analysis period.

[35] Estimated effects from models using simulated annual, simulate fixed (CPI), simulated fixed (RCPI), and simulated fixed (WAGE) eligibility are shown in tables A.20 and A.21.

### 5.4.3 Long-Run Analysis

Although Medicaid eligibility can be imputed for children age 0-18, the main analysis uses a balanced long-run eligibility measure from birth to the eleventh birthday since the Medicaid calculator starts in 1979 and the first calendar year in DCS begins in 1990. In addition, to create the long-run simulated eligibility, I use weights to account for potentially endogenous changes in number of children at each age during childhood (see section B.5 for more details.) To understand how sensitive the results are to selective relocation of children and number of childhood years used for the simulated eligibility, I estimate the effects of total childhood eligibility on parental outcomes using a non-weighted and non-balanced long-run eligibility measure. The results are quantitatively and qualitatively very similar to using non-weighted and non-balanced total child-years of simulated eligibility (see table A.23) indicating that the distribution of number of children is not driving the results and that the long-run effects persist up to 18 years.

Since the survey provides only information about the state of birth and residence, it is not possible to know where the children live between the birth and survey date. Given that Medicaid generosity is in part determined based on place of residence, the choice how to assign Medicaid eligibility during childhood may have implications for estimated effects. In the main long-run analysis, I restrict the sample in DCS and ACS to children being born and residing in the same state since these children are more likely to stay in the same state throughout the childhood. On average, however, 16% reside in a different state than birth state. To check if children are more likely to move as a result Medicaid expansions, I estimate the effect of total child years of simulated eligibility on the probability of living in a state different from birth state. Shown in table A.22, access to Medicaid during childhood does not affect the geographic mobility suggesting that endogenous migration is less likely to bias the estimated effects. Moreover, the results are robust to using a non-restricted sample and assigning long-run eligibility measure based on state

of birth or state of residence (see tables A.24 and A.25).

## 5.5 Mechanisms

### 5.5.1 Racial and Ethnic Differences

I first explore if the racial and ethnic differences in labor supply responses of women could arise because of differences in family structure of white and non-white families.[36] To examine if family structure could explain the differences in labor market outcomes between race-ethnicity groups, I analyze heterogeneity in labor supply responses by marital status of mothers. Since children's Medicaid eligibility differs between single- and two-parent families, I construct simulated eligibility separately for children with single and married parents. One potential concern with marital status-specific simulated eligibility is that parents may change marital status as a response to extended children's Medicaid eligibility.[37] Below however, I show that the bias is likely to be small since the effects of marital outcomes on labor market decisions are not big enough to drive labor supply responses. The results of the heterogeneous analysis by maternal marital status provided in table A.7 show that for non-white mothers the effects are concentrated among single mothers. For white mothers, the estimated effects are not statistically different between married and single mothers. The large difference in labor supply responses by marital status between white and non-white mothers suggests that the tendency of non-white mothers for being single is not the only factor contributing to the racial and ethnic difference. A potential explanation could be the difference in income between single white and single non-white mothers. For instance, earnings of single non-white mothers are

[36] Extensive literature demonstrates that the proportion of non-white children living in two-parent families is not only lower than proportion of white children but also decreasing at a faster rate (Hernandez, 1993). For instance in the CPS ASEC analysis sample 65% of non-white mothers and 95% of white fathers are married.

[37] To reduce the endogeneity concern, one could link individuals in CPS over time and use baseline characteristics. Given that CPS is a place-based survey the probability of linking a child over time may be affected by extended Medicaid eligibility resulting in a selected sample. In fact, I find that children are more likely to be linked over time as a result of increased access to Medicaid and hence refrain from using the linked sample (A.9). The children are more likely to be linked over time because the geographic mobility within state but not across states is lower.

40% lower than earnings of single white mothers.

### 5.5.2 Educational Attainment

Since children's access to Medicaid affects educational outcomes of mothers, I check whether extended Medicaid eligibility may affect labor supply of mothers through changes in maternal educational attainment. As shown in table A.2, non-white mothers are less likely to drop out of high school (5.1%) and to graduate from college (25%) as well as more likely to graduate from high school (7.9%) and attend some college (13.3%). The effects on some college attendance are not considered for this calculation since the estimates are not robust to inclusion of additional controls as shown in figure A.3. Back of the envelope calculation suggests that children's Medicaid is less likely to affect maternal labor market outcomes through educational attainment since Medicaid expansions result in larger effects on college non-completion than on high school graduation and earnings premium is higher for college than for high school (Goldin and Katz, 2007). For white mothers only effects on college completion are robust to alternative specifications (see figure A.2). Relative to the baseline mean, mothers of white children are 6.3% more likely to attain college or more in response to extended Medicaid eligibility of children. The positive effects on college graduation of white mothers may explain the zero effects on labor supply since white mothers are more likely to invest the time into education and not the labor market.

### 5.5.3 Marital Outcomes

Next I examine if marital outcomes may be another potential pathway for labor supply responses as I find that children's Medicaid eligibility affects maternal marital outcomes. I mainly focus on the probability of marriage of non-white mothers since the remaining outcomes are not precisely estimated.[38] Given that non-white mothers are 4.7% more

[38] I find that while Medicaid expansions affect the probability of being married of non-white women, the probability of being never married, ever married, and divorced does not respond to extended Medicaid eligibility (see table A.3).

likely to get married as a result of extended Medicaid eligibility and maternal labor supply responses range between 2.9-5.3% relative to the baseline mean, for Medicaid to work through marital status, marriage has to change labor supply by at least 61.7% $(2.9 \div 4.7)$ considering the lower bound of parental labor supply responses. Since, most of the literature documents negative marriage earnings gaps for women, the back of the envelope calculation shows that Medicaid is less likely to work through marital outcomes.[39] The effects on marital outcomes of white mothers show no clear pattern in any direction. While the probability of being never married decreases, the probability of being ever married and divorced increases.

## 5.6 Return on Investment

To examine program's return on investment, I compare the estimated effect of extended Medicaid eligibility on maternal tax payments with the estimated effect on children's Medicaid cost.[40] The results of this analysis provided in table 5 show that the total cost for Medicaid increases by \$941 ($p < 0.01$) as a result of making one additional child per family eligible for Medicaid. At the same time, extended Medicaid eligibility leads to an increase of federal and state tax liabilities of \$432, the point estimate is, however, not significant at conventional levels. Once 19% of FICA is considered as social benefit, the estimates of tax liabilities are slightly higher, but still not precisely estimated.[41] The imprecise effects on net tax liabilities are partially driven by increased tax benefits as a result of labor supply responses (e.g., benefits from EITC). The results of this analysis suggest that government is not able to recover cost of expenditures on the Medicaid

[39] See for instance Waldfogel (1997, 1998). An exception is the recent work by Juhn and McCue (2016) that does not finds lower earnings among married women from 1966-1975 birth cohort. However, even in the 1966-1975 birth cohort married women with children that are most relevant for my analysis have 35 percent lower earnings than their single counterparts.

[40] I calculate net tax liabilities under US federal and state income tax laws using a simulation program (TAXSIM 32 available at http://www.nber.org/taxsim) provided by the National Bureau of Economic Research (Feenberg and Coutts, 1993). In two-parent families, I assume that the father is the primary earner.

[41] Since individuals can recover some of the FICA contributions in form of Medicare and Social Security benefits, Heller and Mumma (forthcoming) suggest treating 19% of FICA as social benefit. For reference, I also report estimates treating 100% FICA as social benefit.

program just through additional taxes collected.[42] Hendren and Sprung-Keyser (2020), however, calculate the marginal value of public funds associated with Medicaid expansion to pregnant women and infants between 1979 and 1992 and find that Medicaid has paid for itself when all benefits (e.g., improved health of children) are accounted for.

## 5.7 Paternal Outcomes

Next, I examine the effects of children's Medicaid eligibility on paternal labor market outcomes. Fathers considered in this analysis are not necessarily spouses of mothers studied so far. Table 6 provides estimated effects on paternal labor market outcomes by race and ethnicity of the child. Except for labor force participation, paternal labor supply responds positively to extended Medicaid eligibility of children. Usual hours worked per week increase by 1.3% and weeks worked per year increase by 1.0% relative to the baseline mean. In contrast to maternal labor market outcomes, effects on paternal labor supply are entirely driven by fathers with white children.[43] While the labor supply responses range between 1.3% and 1.9% for hours and weeks worked, extensive margin response is only marginally significant. The impacts of total simulated eligibility on labor supply of fathers with white children are quantitatively larger and more precisely estimated from impacts on fathers with non-white children. The point estimates are, however, only statistically different between the race and ethnicity groups in models for usual hours worked per week.

The estimated effects on earnings of fathers with non-white children are negative and marginally significant, but not robust to inclusion of additional control variable as shown in figure A.10. The earnings of fathers with white children respond positively to extended

[42] I also analyze if children's access to Medicaid affects family-level government transfers which is another source of public expenditures (see table A.10). While benefits from welfare, disability, educational assistance, and energy subsidy decrease, contributions from housing subsidy and school lunch increase. Program receipt is however often mismeasured in surveys due to nonresponse, imputation, and measurement error potentially resulting in biased estimates (Meyer et al., 2015).

[43] Shown in table A.8, heterogeneous analysis by paternal marital status suggests that married fathers are mainly affected by extended Medicaid eligibility of children supporting the head of the family hypothesis since 95% of white fathers are married in CPS ASEC sample.

eligibility. The results suggest that one additional eligible child per family leads to an increase of $10,318 which represents 15.4% relative to baseline average earnings. The extremely large average earnings effects for white men are puzzling as they are not in line with moderate labor supply responses (1.9% in usual hours worked per week and 1.3% in weeks worked per year) and fraction of white children covered by Medicaid (10.7%). In addition, paternal earnings respond in areas of the earnings distribution above the maximum eligibility limit (see figure 5) which is not expected given that Medicaid is targeted at low-income population. There are two potential explanations which I explore in appendix section C in more detail. First, the race-ethnicity-specific simulated eligibility measure may be correlated with unobservable characteristics which affect the earnings of high-income white men. Second, men may be pushed into higher earnings occupations or working schedules as a result of children's access to Medicaid. Since maternal earnings never exceed the maximum Medicaid eligibility limits (see figures in appendix section C), the effects on maternal labor supply are not driven by high-income mothers and can be interpreted as causal.

# 6 Conclusion

United States has witnessed a substantial increase in public health insurance coverage of children between 1979 and 2014. Despite the extensive literature studying the consequences of expanded Medicaid coverage, spillover effects on other family members have been under-studied. This paper presents new evidence on the effects of children's Medicaid eligibility on parental labor market outcomes. To identify the effects of Medicaid eligibility the empirical strategy exploits legislative variation at the state, year, and age of the child level which resulted from Medicaid expansions between 1979 and 2014. To address endogeneity of actual eligibility I use the simulated eligibility strategy by estimating reduced form impacts of simulated Medicaid eligibility of children on labor market outcomes of their parents.

I first show that Medicaid take-up of children increases as a response to extending Medicaid eligibility and is partially driven by eligibility of siblings. I then demonstrate that extended Medicaid eligibility of children leads to increased maternal labor supply at the extensive and intensive margin and that the labor supply responses are mainly driven by non-white families. These race-ethnicity differences can be potentially explained by family structure as the estimated effects are concentrated among single non-white mothers suggesting that the head of the household is more likely to change labor market activity. Analysis of mechanisms shows that children's access to Medicaid is less likely to work though marital and educational outcomes of parents.

This work may emphasize at least three policy implications. First, the findings of this study may have implications for the overall generosity of Medicaid eligibility since the general equilibrium effects may exceed the direct benefits of the public health insurance coverage. Second, focusing on disadvantaged population may provide guidance about targeting Medicaid to certain groups, for example about making Medicaid more generous for racial and ethnic minorities. Third, the positive sibling spillovers of Medicaid eligibility on Medicaid coverage may suggest that the program take-up can be increased by reducing inefficiencies such as the cost of application.

# References

**Aaronson, Daniel and Eric French**, "The Effect of Part-Time Work on Wages: Evidence from the Social Security Rules," *Journal of Labor Economics*, 2004, *22* (2), 329–252.

**Aizer, Anna and Jeffrey Grogger**, "Parental Medicaid Expansions and Health Insurance Coverage," Working Paper 9907, National Bureau of Economic Research August 2003.

**Aouad, Marion**, "An Examination of the Intracorrelation of Family Health Insurance," IZA Discussion Papers 14541, Institute of Labor Economics (IZA) July 2021.

**Autor, David H. and David Dorn**, "The Growth of Low-Skill Service Jobs and the Polarization of the US Labor Market," *American Economic Review*, August 2013, *103* (5), 1553–97.

**Bastian, Jacob and Lance Lochner**, "The Earned Income Tax Credit and Maternal Time Use: More Time Working and Less Time with Kids?," *Journal of Labor Economics*, 2022, *40* (3), 573–611.

**Baughman, Reagan and Jeffrey Milyo**, "How Do States Formulate Medicaid and SCHIP Policy? Economic and Political Determinants of State Eligibility Levels," *Mimeo*, 2009.

**Bollinger, Christopher R. and Barry T. Hirsch**, "Match Bias from Earnings Imputation in the Current Population Survey: The Case of Imperfect Matching," *Journal of Labor Economics*, 2006, *24* (3), 483–519.

**Borowczyk-Martins, Daniel and Etienne Lalé**, "The welfare effects of involuntary part-time work," *Oxford Economic Papers*, 08 2017, *70* (1), 183–205.

**Brinkley, Mark A**, "Employer Costs for Employee Compensation to Include Information on Part-time and Full-time Workers," *Compensation and Working Conditions*, 1994, *46*, 1–11.

**Brown, David W, Amanda E Kowalski, and Ithai Z Lurie**, "Long-Term Impacts of Childhood Medicaid Expansions on Outcomes in Adulthood," *The Review of Economic Studies*, 07 2019, *87* (2), 792–821.

**Buchmueller, Thomas, John C. Ham, and Lara D. Shore-Sheppard**, "The Medicaid Program," in "In Economics of Means-Tested Transfer Programs in the United States, Volume 1," University of Chicago Press, 2016, pp. 21–136.

**Cascio, Elizabeth and Diane Whitmore Schanzenbach**, "The Impacts of Expanding Access to High-Quality Preschool Education," *Brookings Papers on Economic Activity*, 2013, *44* (2 (Fall)), 127–192.

**Cascio, Elizabeth U.**, "Maternal Labor Supply and the Introduction of Kindergartens into American Public Schools," *The Journal of Human Resources*, 2009, *44* (1), 140–170.

**Center, State Health Access Data Assistance**, "SHADAC-Enhanced Current Population Survey Health Insurance Coverage Estimates: A Summary of Historical Adjustments," *Minneapolis MN: University of Minnesota*, 2009.

**Cohodes, Sarah R, Daniel S. Grossman, Samuel A. Kleiner, and Michael F. Lovenheim**, "The effect of child health insurance access on schooling: Evidence from public insurance expansions," *Journal of Human Resources*, 2016, *51* (3), 727–759.

**Currie, Janet and Jonathan Gruber**, "Health insurance eligibility, utilization of medical care, and child health," *The Quarterly Journal of Economics*, 1996, *111* (2), 431–466.

_ **and** _ , "Saving babies: the efficacy and cost of recent changes in the Medicaid eligibility of pregnant women," *Journal of political Economy*, 1996, *104* (6), 1263–1296.

**Cutler, David M. and Jonathan Gruber**, "Does public insurance crowd out private insurance?," *The Quarterly Journal of Economics*, 1996, *111* (2), 391–430.

**Dave, Dhaval, Sandra L. Decker, Robert Kaestner, and Kosali I. Simon**, "The Effect of Medicaid Expansions in the Late 1980s and Early 1990s on the Labor Supply of Pregnant Women," *American Journal of Health Economics*, 2015, *1* (2), 165–193.

**DeLeire, Thomas, Leonard M. Lopoo, and Kosali I. Simon**, "Medicaid Expansions and Fertility in the United States," *Demography*, 04 2011, *48* (2), 725–747.

**Dubay, Lisa and Genevieve Kenney**, "Expanding public health insurance to parents: effects on children's coverage under Medicaid," *Health services research*, 2003, *38* (5), 1283–1302.

**East, Chloe N., Sarah Miller, Marianne Page, and Laura R. Wherry**, "Multi-generational Impacts of Childhood Access to the Safety Net: Early Life Exposure to Medicaid and the Next Generation's Health," Working Paper 23810, National Bureau of Economic Research September 2017.

**Eriksen, Tine L. Mundbjerg, Amanda Gaulke, Niels Skipper, and Jannet Svensson**, "The impact of childhood health shocks on parental labor supply," *Journal of Health Economics*, 2021, *78*, 102486.

**Farooq, Ammar and Adriana Kugler**, "Impacts of Public Health Insurance on Occupational Upgrading," *ILR Review*, 2020, *75* (1), 225–250.

**Feenberg, Daniel and Elisabeth Coutts**, "An introduction to the TAXSIM model," *Journal of Policy Analysis and Management*, 1993, *12* (1), 189–194.

**Fitzpatrick, Maria Donovan**, "Preschoolers Enrolled and Mothers at Work? The Effects of Universal Prekindergarten," *Journal of Labor Economics*, 2010, *28* (1), 51–85.

**Flood, Sarah, Miriam King, Renae Rodgers, Steven Ruggles, and J. Robert Warren**, "Integrated Public Use Microdata Series, Current Population Survey: Version 7.0 [dataset]," *Minneapolis, MN: IPUMS*, 2020.

**Frean, Molly, Jonathan Gruber, and Benjamin D. Sommers**, "Premium subsidies, the mandate, and Medicaid expansion: Coverage effects of the Affordable Care Act," *Journal of Health Economics*, 2017, *53*, 72 – 86.

**Gelbach, Jonah B.**, "Public Schooling for Young Children and Maternal Labor Supply," *American Economic Review*, March 2002, *92* (1), 307–322.

**Goldin, Claudia and Lawrence F Katz**, "The Race between Education and Technology: The Evolution of U.S. Educational Wage Differentials, 1890 to 2005," Working Paper 12984, National Bureau of Economic Research March 2007.

**Gross, Tal and Matthew J. Notowidigdo**, "Health insurance and the consumer bankruptcy decision: Evidence from expansions of Medicaid," *Journal of Public Economics*, 2011, *95* (7), 767 – 778.

**Grossman, Daniel S, Sebastian Tello-Trillo, and Barton Willage**, "Health Insurance for Whom? The 'Spill-up' Effects of Children's Health Insurance on Mothers," Working Paper 29661, National Bureau of Economic Research January 2022.

**Gruber, Jonathan**, "Medicaid," in "Means-Tested Transfer Programs in the United States," University of Chicago Press, January 2003, pp. 15–78.

**Ham, John C. and Lara D. Shore-Sheppard**, "Did Expanding Medicaid Affect Welfare Participation," *Industrial and Labor Relations Review*, 2005, *58* (3), 452–470.

**Ham, John C and Lara Shore-Sheppard**, "The effect of Medicaid expansions for low-income children on Medicaid participation and private insurance coverage: evidence from the SIPP," *Journal of Public Economics*, 2005, *89* (1), 57–83.

**Hamersma, Sarah, Matthew Kim, and Brenden Timpe**, "THE EFFECT OF PARENTAL MEDICAID EXPANSIONS ON CHILDREN'S HEALTH INSURANCE COVERAGE," *Contemporary Economic Policy*, 2019, *37* (2), 297–311.

**Heckman, James J. and Stefano Mosso**, "The Economics of Human Development and Social Mobility," *Annual Review of Economics*, 2014, *6* (1), 689–733.

**Heller, Blake H. and Kirsten Slungaard Mumma**, "Immigrant Integration in the United States: The Role of Adult English Language Training," *American Economic Journal: Economic Policy*, forthcoming.

**Hendren, Nathaniel and Ben Sprung-Keyser**, "A Unified Welfare Analysis of Government Policies*," *The Quarterly Journal of Economics*, 03 2020, *135* (3), 1209–1318.

**Hernandez, Donald J.**, *America's children: Resources from Family, Government, and the Economy*, Russell Sage Foundation, 1993.

**Hirsch, Barry T.**, "Why Do Part-Time Workers Earn Less? The Role of Worker and Job Skills," *ILR Review*, 2005, *58* (4), 525–551.

_ **and Edward J. Schumacher**, "Match Bias in Wage Gap Estimates Due to Earnings Imputation," *Journal of Labor Economics*, 2004, *22* (3), 689–722.

**Hudson, Julie L. and Asako S. Moriya**, "Medicaid Expansion For Adults Had Measurable 'Welcome Mat' Effects On Their Children," *Health Affairs*, 2017, *36* (9), 1643–1651. PMID: 28874493.

**Jaeger, David A.**, "Reconciling the Old and New Census Bureau Education Questions: Recommendations for Researchers," *Journal of Business & Economic Statistics*, 1997, *15* (3), 300–309.

**Juhn, Chinhui and Kristin McCue**, "Evolution of the Marriage Earnings Gap for Women," *American Economic Review*, May 2016, *106* (5), 252–56.

**Klerman, Jacob A, Michael Davern, Kathleen Thiede Call, Victoria Lynch, and Jeanne D Ringel**, "Understanding The Current Population Survey's Insurance Estimates And The Medicaid 'Undercount'," *Health Affairs*, 2009, *28* (Suppl1), w991–w1001.

**Kuka, Elira and Na'ama Shenhav**, "Long-Run Effects of Incentivizing Work After Childbirth," Working Paper 27444, National Bureau of Economic Research June 2020.

**Larrimore, Jeff, Richard V. Burkhauser, Shuaizhang Feng, and Laura Zayatz**, "Consistent Cell Means for Topcoded Incomes in the Public Use March CPS (1976-2007)," *Journal of Economic and Social Measurement*, 2008, *33* (2/3), 89–128.

**Lettau, Michael K**, "Compensation in part-time jobs versus full-time jobs What if the job is the same?," *Economics Letters*, 1997, *56* (1), 101–106.

_ **and Thomas C Buchmueller**, "Comparing benefit costs for full-and part-time workers," *Monthly Lab. Rev.*, 1999, *122*, 30.

**Meyer, Bruce D., Wallace K. C. Mok, and James X. Sullivan**, "Household Surveys in Crisis," *Journal of Economic Perspectives*, November 2015, *29* (4), 199–226.

**Miller, Sarah and Laura R Wherry**, "The long-term effects of early life Medicaid coverage," *Journal of Human Resources*, 2019, *54* (3), 785–824.

**Nightingale, Madeline**, "Looking beyond Average Earnings: Why Are Male and Female Part-Time Employees in the UK More Likely to Be Low Paid Than Their Full-Time Counterparts?," *Work, Employment and Society*, 2019, *33* (1), 131–148.

**Pei, Zhuan**, "Eligibility Recertification and Dynamic Opt-In Incentives in Income-Tested Social Programs: Evidence from Medicaid/CHIP," *American Economic Journal: Economic Policy*, February 2017, *9* (1), 241–76.

**Pongrace, David M and Alan P Zilberman**, "A Comparison of Hourly Wage Rates for Full- and Part-Time Workers by Occupation, 2007," *Monthly Labor Review*, 2009.

**Ruggles, Steven, Sarah Flood, Ronald Goeken, Josiah Grover, Erin Meyer, Jose Pacas, and Matthew Sobek**, "IPUMS USA: Version 10.0 [dataset]," *Minneapolis, MN: IPUMS*, 2020.

**Sacarny, Adam, Katherine Baicker, and Amy Finkelstein**, "Out of the Woodwork: Enrollment Spillovers in the Oregon Health Insurance Experiment," *American Economic Journal: Economic Policy*, August 2022, *14* (3), 273–95.

**Sommers, Benjamin D.**, "Insuring children or insuring families: Do parental and sibling coverage lead to improved retention of children in Medicaid and CHIP?," *Journal of Health Economics*, 2006, *25* (6), 1154–1169.

_ **, Kao-Ping Chua, Genevieve M. Kenney, Sharon K. Long, and Stacey McMorrow**, "California's Early Coverage Expansion under the Affordable Care Act: A County-Level Analysis," *Health Services Research*, 2016, *51* (3), 825–845.

**Sonier, Julie, Michel H. Boudreaux, and Lynn A. Blewett**, "Medicaid 'Welcome-Mat' Effect Of Affordable Care Act Implementation Could Be Substantial," *Health Affairs*, 2013, *32* (7), 1319–1325. PMID: 23804584.

**Stuber, Jennifer and Elizabeth Bradley**, "Barriers to Medicaid Enrollment: Who Is at Risk?," *American Journal of Public Health*, 2005, *95* (2), 292–298. PMID: 15671467.

**vom Lehn, Christian, Cache Ellsworth, and Zachary Kroff**, "Reconciling Occupational Mobility in the Current Population Survey," *Journal of Labor Economics*, forthcoming.

**Waldfogel, Jane**, "The Effect of Children on Women's Wages," *American Sociological Review*, 1997, *62* (2), 209–217.

_ , "Understanding the "Family Gap" in Pay for Women with Children," *Journal of Economic Perspectives*, March 1998, *12* (1), 137–156.

**Wikle, Jocelyn and Riley Wilson**, "Access to Head Start and Maternal Labor Supply: Experimental and Quasi-Experimental Evidence," IZA Discussion Papers 14880, Institute of Labor Economics (IZA) Nov 2021.

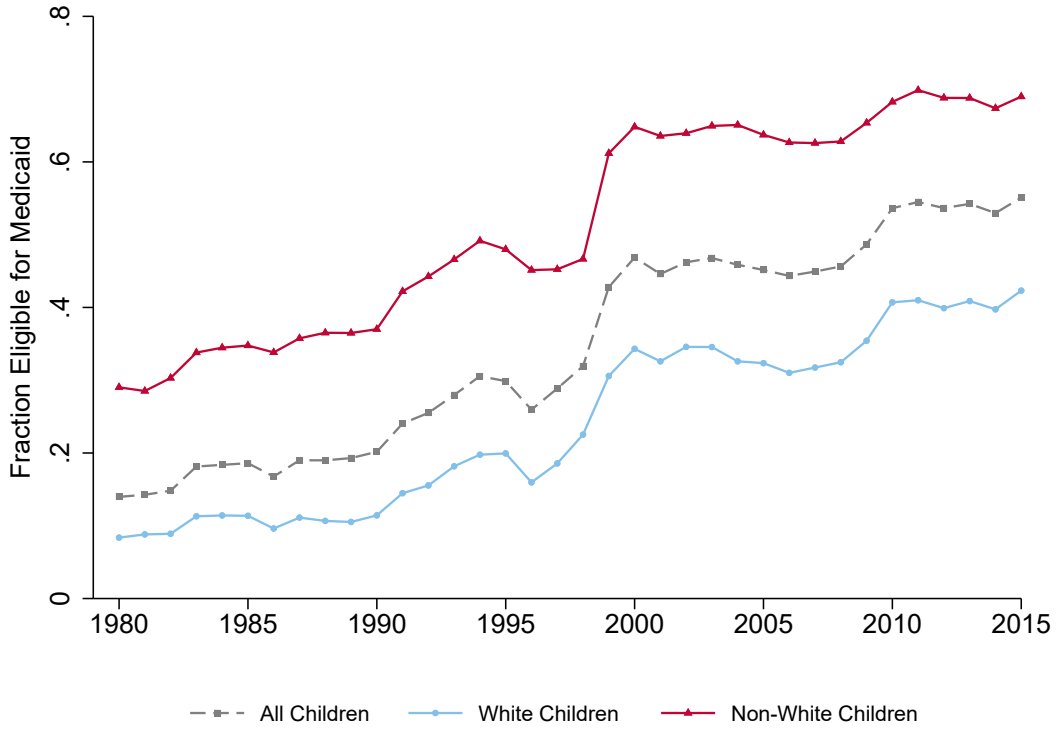
**Yelowitz, Aaron S.**, "The Medicaid Notch , Labor Supply , and Welfare Participation: Evidence from Eligibility Expansions," *Quarterly Journal of Economics*, 1995, *110* (4), 909–939.

**Zavodny, Madeline and Marianne P. Bitler**, "The effect of Medicaid eligibility expansions on fertility," *Social Science & Medicine*, 2010, *71* (5), 918–924.

**Ziegenfuss, Jeanette Y. and Michael E. Davern**, "Twenty Years of Coverage: An Enhanced Current Population Survey—1989–2008," *Health Services Research*, 2011, *46* (1p1), 199–209.

# Figures and Tables

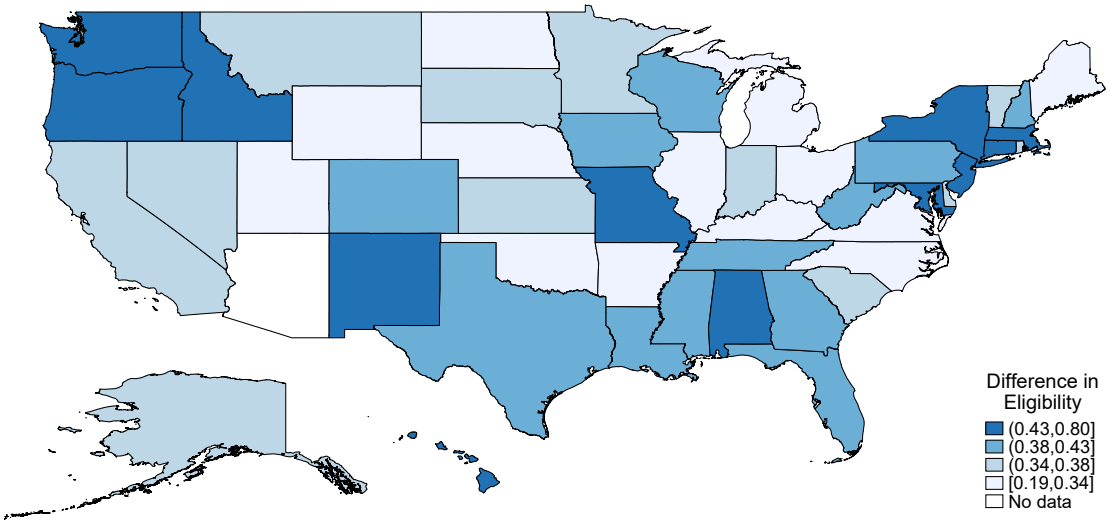
Figure 1:
National Variation in Medicaid Eligibility by Child's Race & Ethnicity



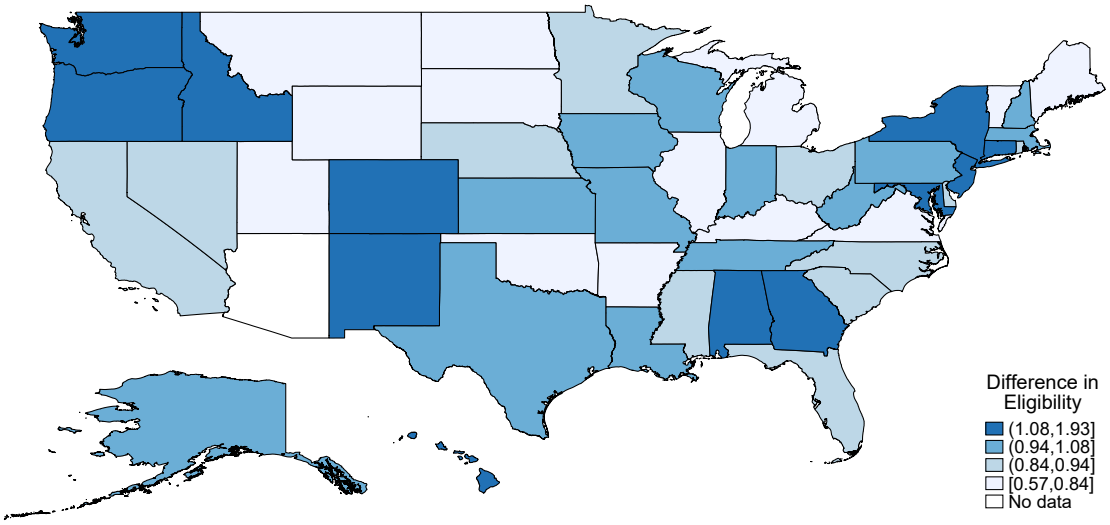

*Notes:* This figure shows the fraction of eligible white and non-white children between 1979 and 2014. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. Arizona is not included because the state did not adopt a Medicaid program until 1982.

Figure 2:
State Variation in Race-Ethnicity-Specific Total Simulated Eligibility

(a) Single-Child Families



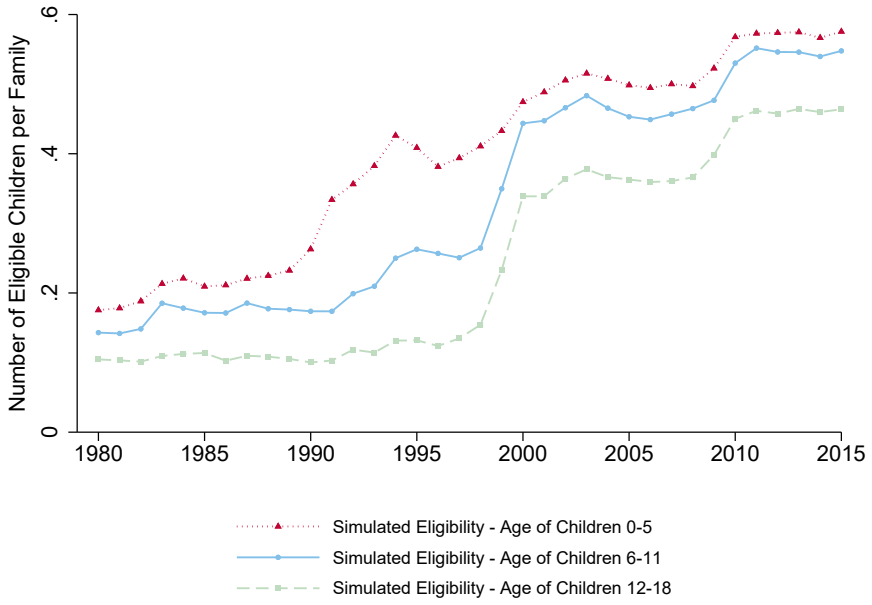

(b) Multiple-Child Families



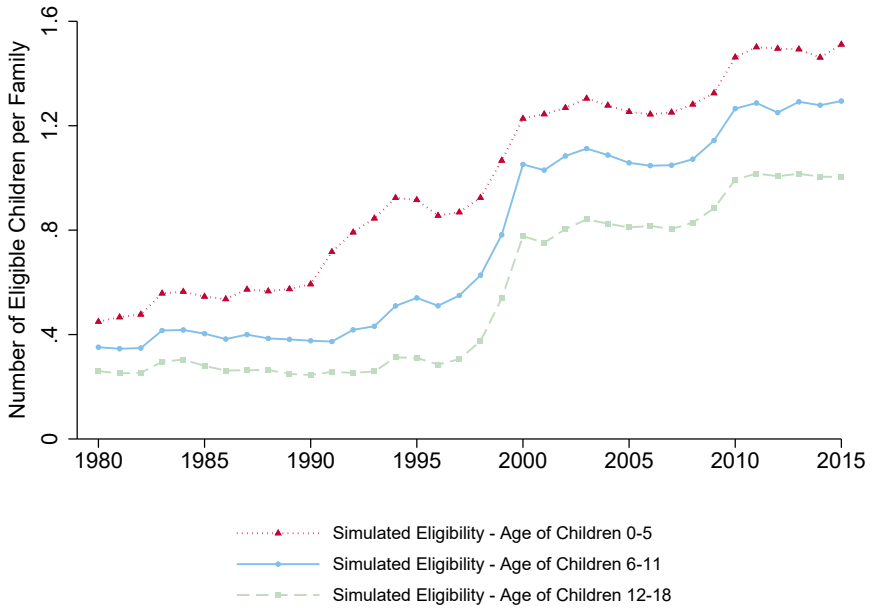

*Notes:* These figures show the difference in race-ethnicity-specific total simulated eligibility between 1979 and 2014 for (a) single- and (b) multiple-child families in each state. The quartiles represent the difference in total simulated eligibility between 1979 and 2014. These years are the start and end of the analysis period. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. Arizona is not included because the state did not adopt a Medicaid program until 1982.

Figure 3:
National Variation in Race-Ethnicity-Specific Total Simulated Eligibility

(a) Single-Child Families



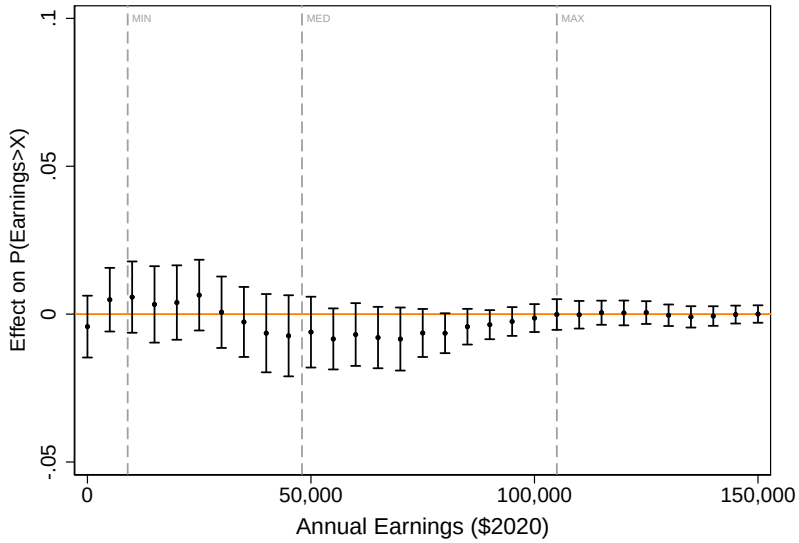

(b) Multiple-Child Families



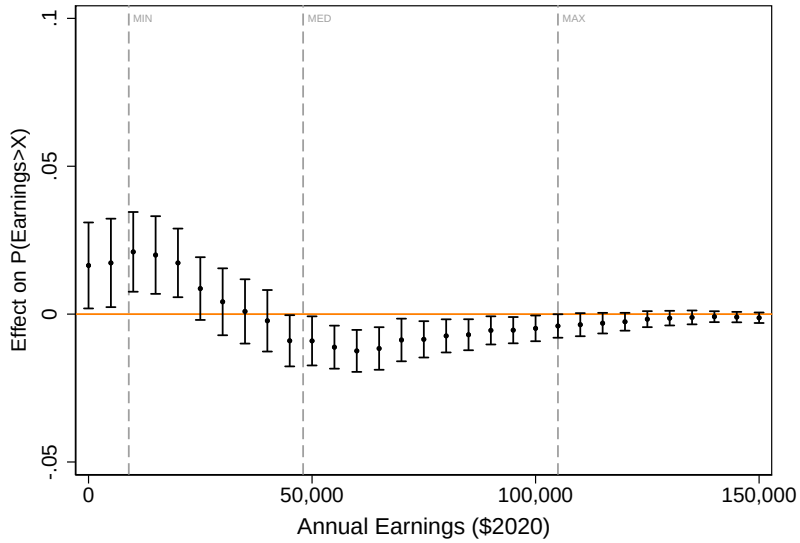

*Notes:* These figures show the average race-ethnicity-specific total simulated eligibility between 1979 and 2014 for (a) single- and (b) multiple-child families by child's age. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. Arizona is only included after 1982 because the state did not adopt a Medicaid program until 1982.

Figure 4:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Maternal Annual Earnings ($2020)

(a) Mothers with White Children




(b) Mothers with Non-White Children

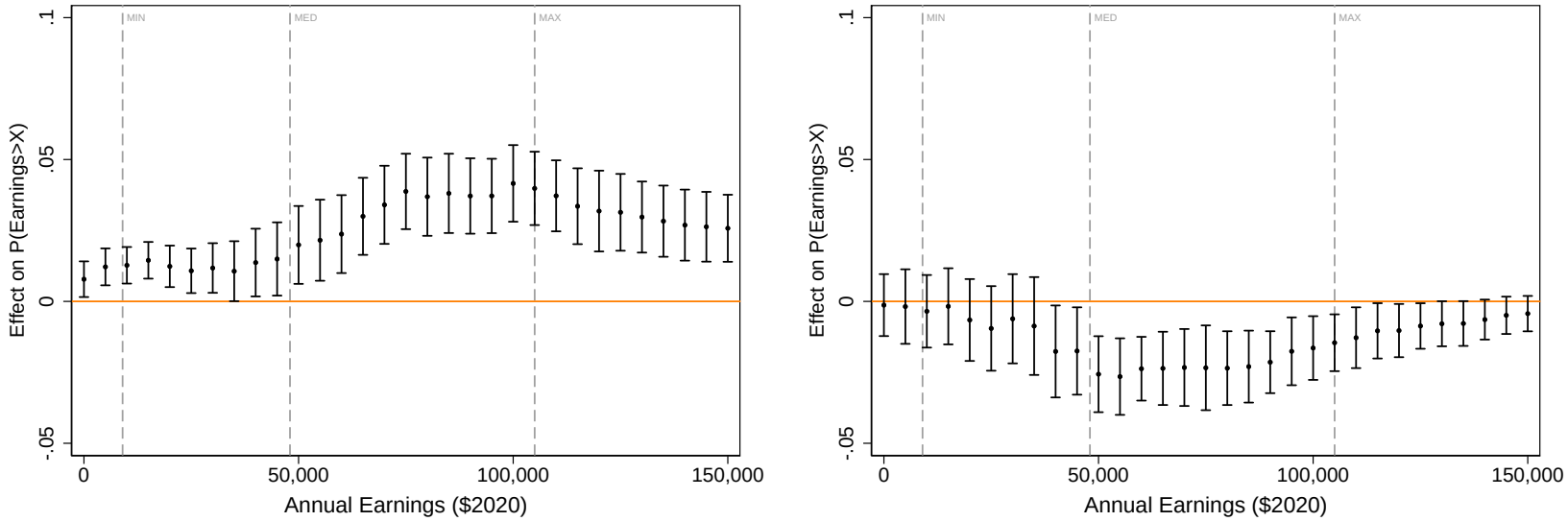



*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on annual earnings ($2020) last year of mothers with (a) white and (b) non-white children. Each point estimate and confidence interval is obtained from a different regression where the dependent variable is an indicator equals to one if maternal annual earnings ($2020) were at least as great as X (0,5000,...,150000) last year. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors are clustered at the state level. The first, second, and third dashed vertical line represent average minimum, median, and maximum Medicaid eligibility cutoffs during the analysis period, respectively. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure 5:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Paternal Annual Earnings ($2020)

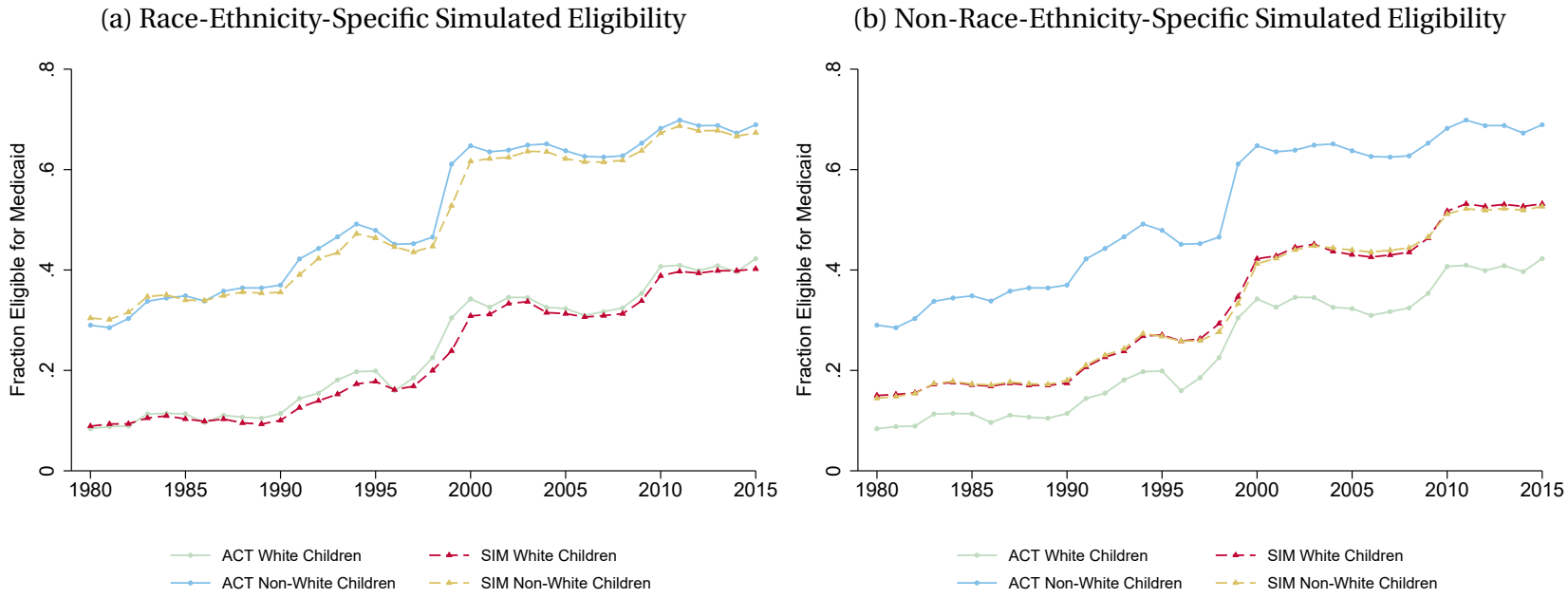
(a) Fathers with White Children

(b) Fathers with Non-White Children



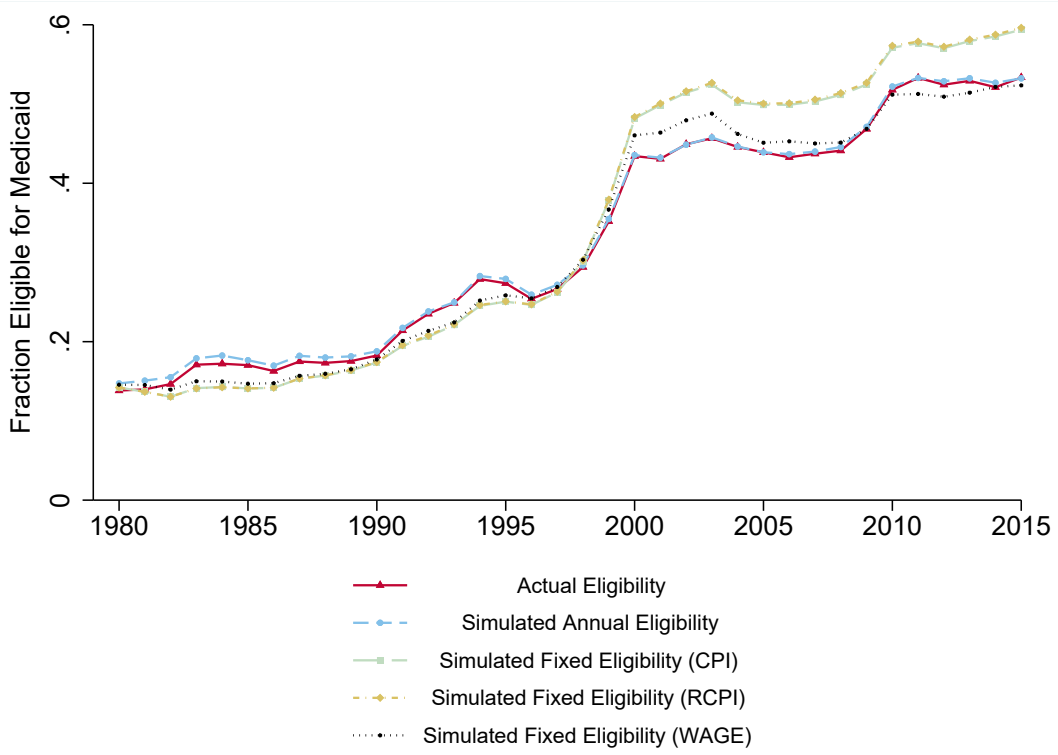

*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on annual earnings ($2020) last year of fathers with (a) white and (b) non-white children. Each point estimate and confidence interval is obtained from a different regression where the dependent variable is an indicator equals to one if paternal annual earnings ($2020) were greater as X (0,5000,...,150000) last year. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors are clustered at the state level. The first, second, and third dashed vertical line represent average minimum, median, and maximum Medicaid eligibility cutoffs during the analysis period, respectively. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure 6:
Race-Ethnicity-Specific vs. Non-Race-Ethnicity-Specific Simulated Eligibility Measure




*Notes:*This figure shows race-ethnicity-specific simulated eligibility measure (a) and non-race-ethnicity-specific simulated eligibility measure (b) for white and non-white children between 1979 and 2014. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. Arizona is not included because the state did not adopt a Medicaid program until 1982.

Figure 7:
National Variation in Alternative Simulated Eligibility Measures




*Notes:* This figure shows different total simulated eligibility measures between 1979 and 2014. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. Arizona is not included because the state did not adopt a Medicaid program until 1982.

Table 1:
Effect of State-Level Characteristics on Medicaid Eligibility Limits

| | (1) | (2) | (3) |
|---|---|---|---|
| Medicaid Coverage | 43.96 | 23.15 | -2.24 |
| | (36.27) | (39.43) | (40.96) |
| Labor Force Participation | 47.53 | 24.70 | 18.40 |
| | (59.34) | (66.36) | (70.23) |
| Hours Worked per Week | 0.11 | 0.06 | 0.50 |
| | (1.37) | (1.36) | (1.29) |
| State Earned Income Credit | 36.72 | 33.24 | 29.07 |
| | (29.87) | (29.21) | (28.85) |
| State Minimum Wage | 0.26 | 1.13 | 0.83 |
| | (3.23) | (3.18) | (2.97) |
| Welfare Benefit ($2020) | 0.01 | 0.01 | 0.01 |
| | (0.02) | (0.02) | (0.02) |
| Major Waiver or TANF | 6.10 | 7.95 | 8.64 |
| | (10.72) | (11.63) | (9.53) |
| Observations | 1,734 | 1,734 | 1,734 |
| Adjusted $R^2$ | 0.85 | 0.85 | 0.85 |
| Mean Y - Baseline | 65 | 65 | 65 |
| Mean Y - Overall | 163 | 163 | 163 |
| Demographic Controls | X | X | X |
| State Fixed Effects | X | X | X |
| Year Fixed Effects | X | X | X |

*Notes:* This table shows results from regressions estimating the effect of state-level characteristics on the maximum Medicaid eligibility limit for children age 0-18. Column 1, 2, and 3 show models using contemporaneous, first order lagged, and second order lagged state-level characteristics, respectively. Demographic controls include income per capita, fraction of population non-white, married, with high school completion or less, age 0-18, age 25-54, with one child, and with multiple children. Standard errors in parentheses are clustered at the state level. The data is from 1979-2014. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 2:
Effect of Race-Ethnicity-Specific Simulated Eligibility on Child's Medicaid Coverage

| | Child-Level Medicaid Coverage | | | Child-Level Medicaid Coverage | | | Family-Level Medicaid Coverage | | |
|---|---|---|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White | All | White | Non-White |
| SIM | 0.08*** | 0.09*** | 0.06*** | 0.07*** | 0.08*** | 0.05** | | | |
| | (0.02) | (0.03) | (0.02) | (0.02) | (0.02) | (0.02) | | | |
| SIMS | | | | 0.02** | 0.01 | 0.03*** | | | |
| | | | | (0.01) | (0.01) | (0.01) | | | |
| SIMT | | | | | | | 0.32*** | 0.26*** | 0.40*** |
| | | | | | | | (0.06) | (0.06) | (0.06) |
| Observations | 1,418,012 | 889,854 | 528,158 | 1,418,012 | 889,854 | 528,158 | 1,418,012 | 889,854 | 528,158 |
| Adjusted $R^2$ | 0.22 | 0.16 | 0.19 | 0.22 | 0.16 | 0.19 | 0.33 | 0.19 | 0.36 |
| Mean Y - Baseline | 0.10 | 0.06 | 0.21 | 0.10 | 0.06 | 0.21 | 0.23 | 0.13 | 0.52 |
| Mean Y - Overall | 0.22 | 0.14 | 0.34 | 0.22 | 0.14 | 0.34 | 0.44 | 0.27 | 0.73 |
| Mean SIM - Overall | | | | 0.35 | 0.23 | 0.54 | | | |
| Mean SIMS - Overall | | | | 0.30 | 0.19 | 0.49 | | | |
| Mean SIMT - Overall | | | | 0.65 | 0.42 | 1.04 | | | |

*Notes:* This table shows results from regressions estimating the effect of race-specific simulated eligibility on child-level and family-level Medicaid coverage. SIM, SIMS, and SIMT refers to child's own, sibling's total, and family's total simulated eligibility, respectively. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 3:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Maternal Labor Supply

| | All | White | Non-White |
|---|---|---|---|
| | Usual Hours Worked per Week | | |
| SIMT | 0.19 | -0.17 | 0.66** |
| | (0.18) | (0.26) | (0.26) |
| Observations | 1,375,551 | 863,738 | 511,813 |
| Adjusted $R^2$ | 0.07 | 0.08 | 0.06 |
| Mean Y - Baseline | 21.89 | 21.61 | 22.75 |
| Mean Y - Overall | 25.56 | 25.91 | 24.96 |
| | Weeks Worked per Year | | |
| SIMT | 0.38** | -0.02 | 0.90*** |
| | (0.18) | (0.26) | (0.31) |
| Observations | 1,375,551 | 863,738 | 511,813 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.09 |
| Mean Y - Baseline | 25.41 | 25.60 | 24.86 |
| Mean Y - Overall | 31.44 | 32.43 | 29.71 |
| | Labor Force Participation | | |
| SIMT | 0.01*** | 0.00 | 0.03*** |
| | (0.00) | (0.01) | (0.01) |
| Observations | 1,330,378 | 838,593 | 491,785 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.07 |
| Mean Y - Baseline | 0.57 | 0.57 | 0.57 |
| Mean Y - Overall | 0.68 | 0.69 | 0.66 |
| | Annual Total Earnings ($2020) | | |
| SIMT | -269 | -294 | -236 |
| | ( 435) | ( 554) | ( 566) |
| Observations | 1,375,551 | 863,738 | 511,813 |
| Adjusted $R^2$ | 0.08 | 0.08 | 0.07 |
| Mean Y - Baseline | 14,826 | 14,674 | 15,279 |
| Mean Y - Overall | 24,306 | 25,932 | 21,485 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, usual hours worked last year, and labor force participation last week). Usual hours worked per week, weeks worked last year, and usual hours worked last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 4:
Effect of Total Child Years of Simulated Eligibility on Maternal Labor Supply

| | All | White | Non-White |
|---|---|---|---|
| | Usual Hours Worked per Week | | |
| SIMC | 0.09** | -0.02 | 0.22*** |
| | ( 0.03) | ( 0.05) | ( 0.04) |
| Observations | 5,837,237 | 4,029,464 | 1,807,773 |
| Adjusted $R^2$ | 0.06 | 0.06 | 0.05 |
| Mean Y - Baseline | 23.33 | 23.49 | 22.91 |
| Mean Y - Overall | 24.63 | 25.05 | 23.88 |
| | Labor Force Participation | | |
| SIMC | 0.000 | -0.002* | 0.004*** |
| | (0.001) | (0.001) | (0.001) |
| Observations | 5,837,237 | 4,029,464 | 1,807,773 |
| Adjusted $R^2$ | 0.05 | 0.05 | 0.04 |
| Mean Y - Baseline | 0.69 | 0.69 | 0.67 |
| Mean Y - Overall | 0.71 | 0.72 | 0.69 |

*Notes:* This table shows results from regressions estimating the effect of total child years of simulated eligibility on parental labor supply (usual hours worked per week last year and labor force participation last week). Usual hours worked per week last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from 5% census sample 1990 and 2000 and ACS 2010. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 5:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Earnings, Taxes, Government Transfers, and Medicaid Cost

| | Medicaid Cost | Federal & State Tax |
|---|---|---|
| SIMT | 941*** | 432 |
| | ( 113) | ( 655) |
| Observations | 1,189,020 | 1,189,020 |
| Adjusted $R^2$ | 0.81 | 0.11 |
| Mean Outcome - Baseline | 290 | 12,508 |
| Mean Outcome - Overall | 881 | 12,250 |
| | Federal & State Tax (19% FICA) | Federal & State Tax (100% FICA) |
| SIMT | 474 | 653 |
| | ( 677) | ( 775) |
| Observations | 1,189,020 | 1,189,020 |
| Adjusted $R^2$ | 0.12 | 0.14 |
| Mean Outcome - Baseline | 13,792 | 19,263 |
| Mean Outcome - Overall | 14,117 | 22,078 |

*Notes:* This table shows results from regression estimating the effect of race-specific total simulated eligibility on annual federal and state taxes ($2020) and child's total Medicaid cost ($2020). All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1981-2013 and MSIS 1980-2012. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, respectively.

Table 6:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Paternal Labor Supply

| | All | White | Non-White |
|---|---|---|---|
| | Usual Hours Worked per Week | | |
| SIMT | 0.52*** | 0.80*** | 0.02 |
| | (0.17) | (0.21) | (0.24) |
| Observations | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.04 | 0.03 | 0.03 |
| Mean Y - Baseline | 42.07 | 42.93 | 38.74 |
| Mean Y - Overall | 41.63 | 42.93 | 38.73 |
| | Weeks Worked per Year | | |
| SIMT | 0.44** | 0.63*** | 0.10 |
| | (0.17) | (0.18) | (0.32) |
| Observations | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.04 | 0.03 | 0.04 |
| Mean Y - Baseline | 46.53 | 47.32 | 43.45 |
| Mean Y - Overall | 46.08 | 47.04 | 43.93 |
| | Labor Force Participation | | |
| SIMT | 0.00 | 0.01* | 0.00 |
| | (0.00) | (0.00) | (0.00) |
| Observations | 1,058,665 | 724,271 | 334,394 |
| Adjusted $R^2$ | 0.04 | 0.03 | 0.04 |
| Mean Y - Baseline | 0.96 | 0.97 | 0.92 |
| Mean Y - Overall | 0.94 | 0.95 | 0.92 |
| | Annual Total Earnings ($2020) | | |
| SIMT | 5,951*** | 10,318*** | -1,950* |
| | ( 1,340) | ( 1,642) | ( 974) |
| Observations | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.10 | 0.08 | 0.05 |
| Mean Y - Baseline | 62,535 | 66,614 | 46,665 |
| Mean Y - Overall | 65,210 | 72,342 | 49,282 |

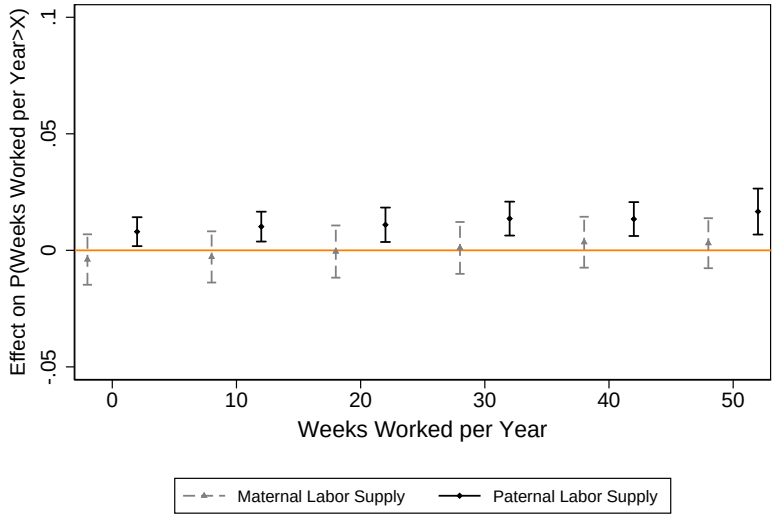
*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, usual hours worked last year, and labor force participation last week). Usual hours worked per week, weeks worked last year, and usual hours worked last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

# Appendix

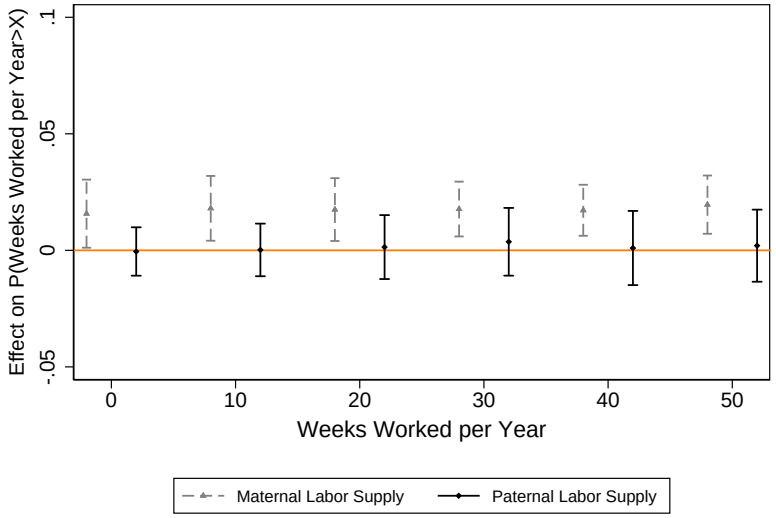
## A Supplemental Figures and Tables

Figure A.1:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Parental Weeks Worked per Year

(a) Parents with White Children

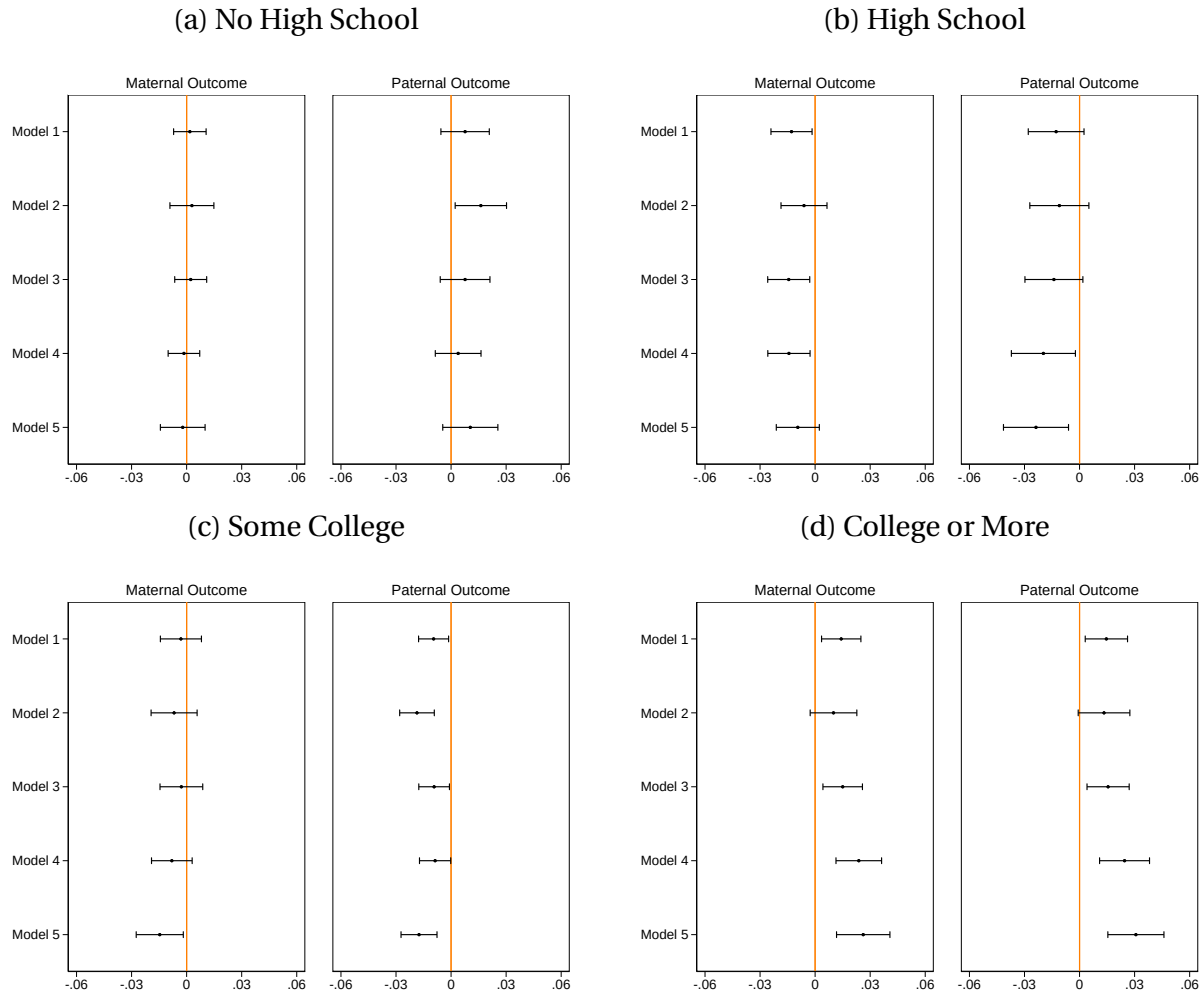



(b) Parents with Non-White Children




*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on parental weeks worked last year. Each point estimate and confidence interval is obtained from a different regression where the dependent variable is an indicator equals to one if a parent worked more than X (0,10,...,50) weeks last year. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure A.2:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility
on Educational Outcomes of Parents with white Children
Robustness to Identifying Assumption

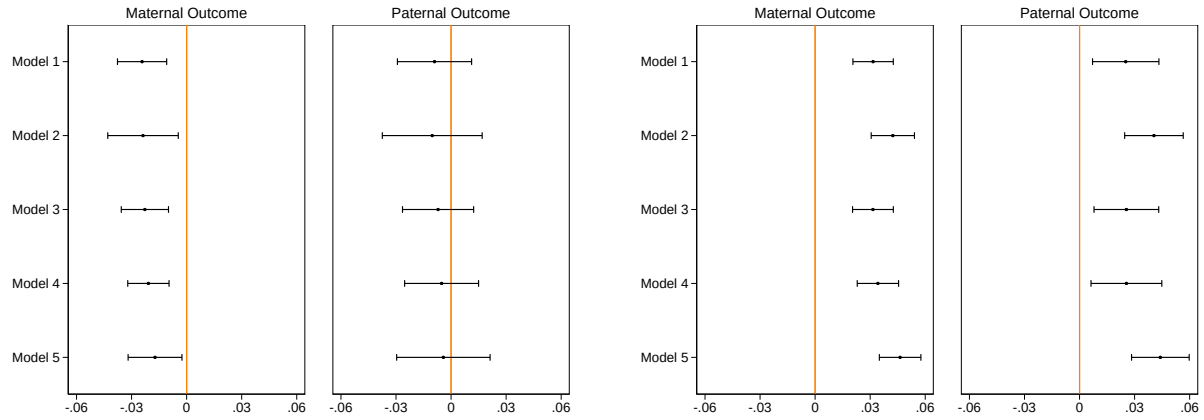



*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on parental marital outcomes (indicator for no high school, high school, some college, college or more). All models include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year) and parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family). All models except model 2 contain state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Model 2 includes state-by-year-by-race-ethnicity fixed effects. Model 3 includes state-by-age-by-race-ethnicity fixed effects. Model 4 includes year-by-age-by-race-ethnicity fixed effects. Model 5 includes state-by-year-by-race-ethnicity, state-by-age-by-race-ethnicity, and year-by-age-by-race-ethnicity fixed effects. Regressions are weighted with parental survey weights divided by number of children per family. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure A.3:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on
Educational Outcomes of Parents with Non-white Children
Robustness to Identifying Assumption

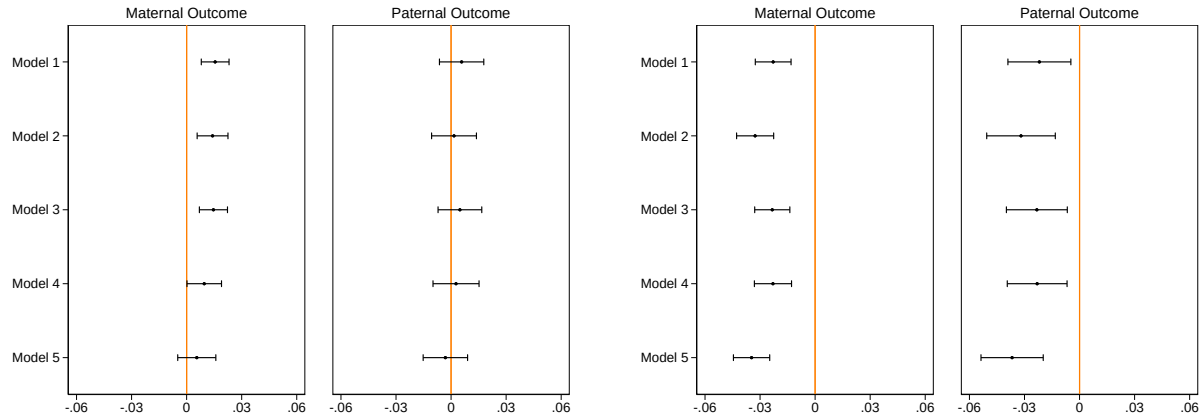
(a) No High School

(b) High School




(c) Some College

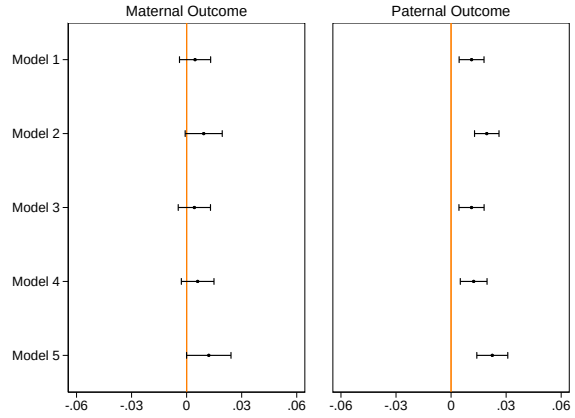
(d) College or More



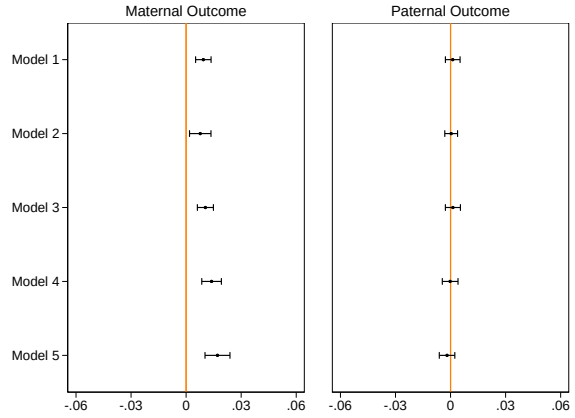

*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on parental marital outcomes (indicator for no high school, high school, some college, college or more). All models include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year) and parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family). All models except model 2 contain state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Model 2 includes state-by-year-by-race-ethnicity fixed effects. Model 3 includes state-by-age-by-race-ethnicity fixed effects. Model 4 includes year-by-age-by-race-ethnicity fixed effects. Model 5 includes state-by-year-by-race-ethnicity, state-by-age-by-race-ethnicity, and year-by-age-by-race-ethnicity fixed effects. Regressions are weighted with parental survey weights divided by number of children per family. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure A.4:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Marital Outcomes of Parents with White Children
Robustness to Identifying Assumption

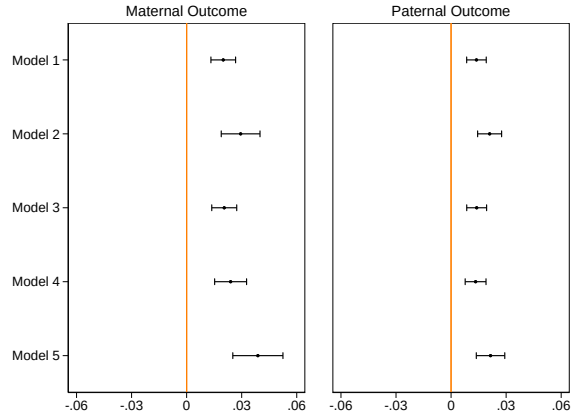
(a) Married



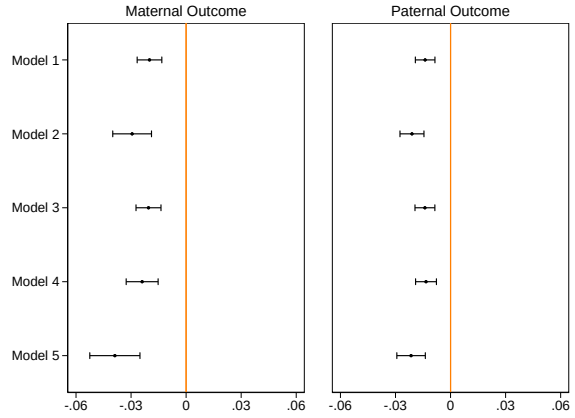

(b) Divorced




(c) Ever Married



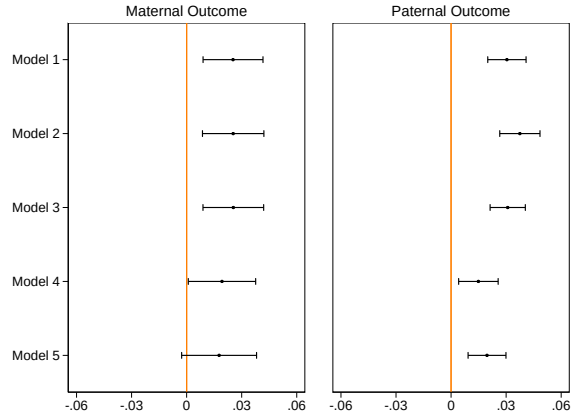

(d) Never Married



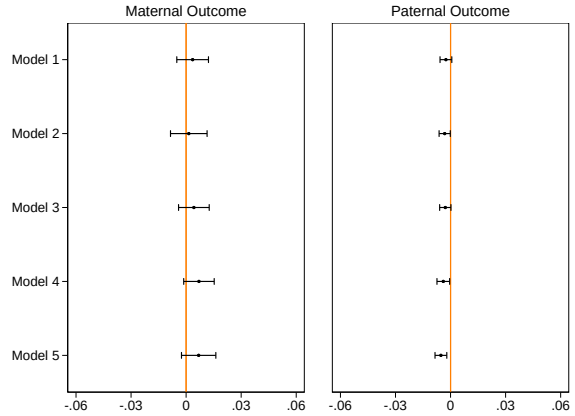

*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on parental marital outcomes (indicator for married, never married, ever married, divorced). All models include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year) and parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family). All models except model 2 contain state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Model 2 includes state-by-year-by-race-ethnicity fixed effects. Model 3 includes state-by-age-by-race-ethnicity fixed effects. Model 4 includes year-by-age-by-race-ethnicity fixed effects. Model 5 includes state-by-year-by-race-ethnicity, state-by-age-by-race-ethnicity, and year-by-age-by-race-ethnicity fixed effects. Regressions are weighted with parental survey weights divided by number of children per family. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure A.5:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility
on Marital Outcomes of Parents with Non-White Children
Robustness to Identifying Assumption

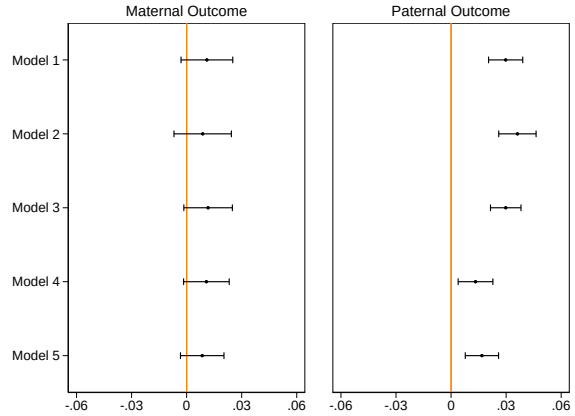
(a) Married



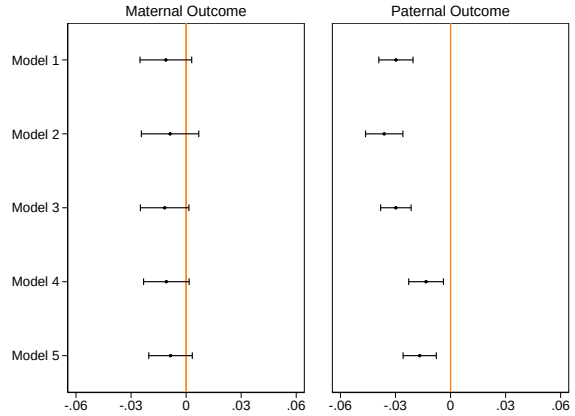

(b) Divorced




(c) Ever Married




(d) Never Married

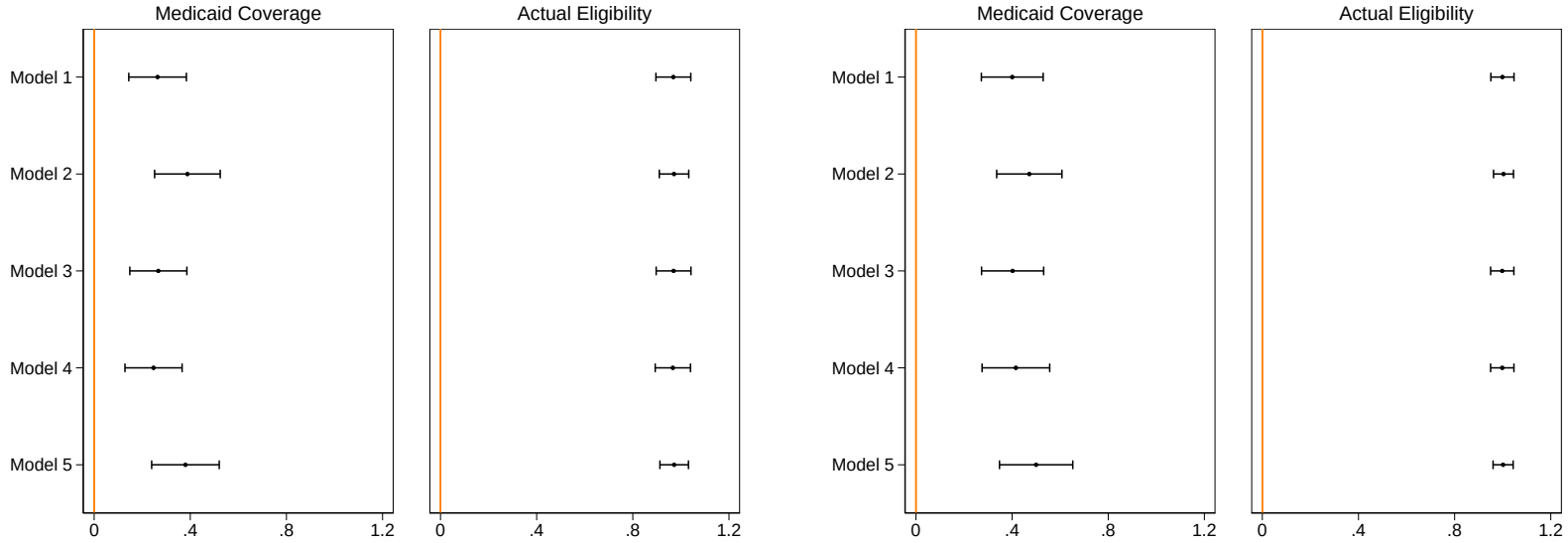



*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on parental marital outcomes (indicator for married, never married, ever married, divorced). All models include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year) and parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family). All models except model 2 contain state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Model 2 includes state-by-year-by-race-ethnicity fixed effects. Model 3 includes state-by-age-by-race-ethnicity fixed effects. Model 4 includes year-by-age-by-race-ethnicity fixed effects. Model 5 includes state-by-year-by-race-ethnicity, state-by-age-by-race-ethnicity, and year-by-age-by-race-ethnicity fixed effects. Regressions are weighted with parental survey weights divided by number of children per family. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure A.6:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Medicaid Coverage and Eligibility
Robustness to Identifying Assumption

(a) White Children

(b) Non-White Children

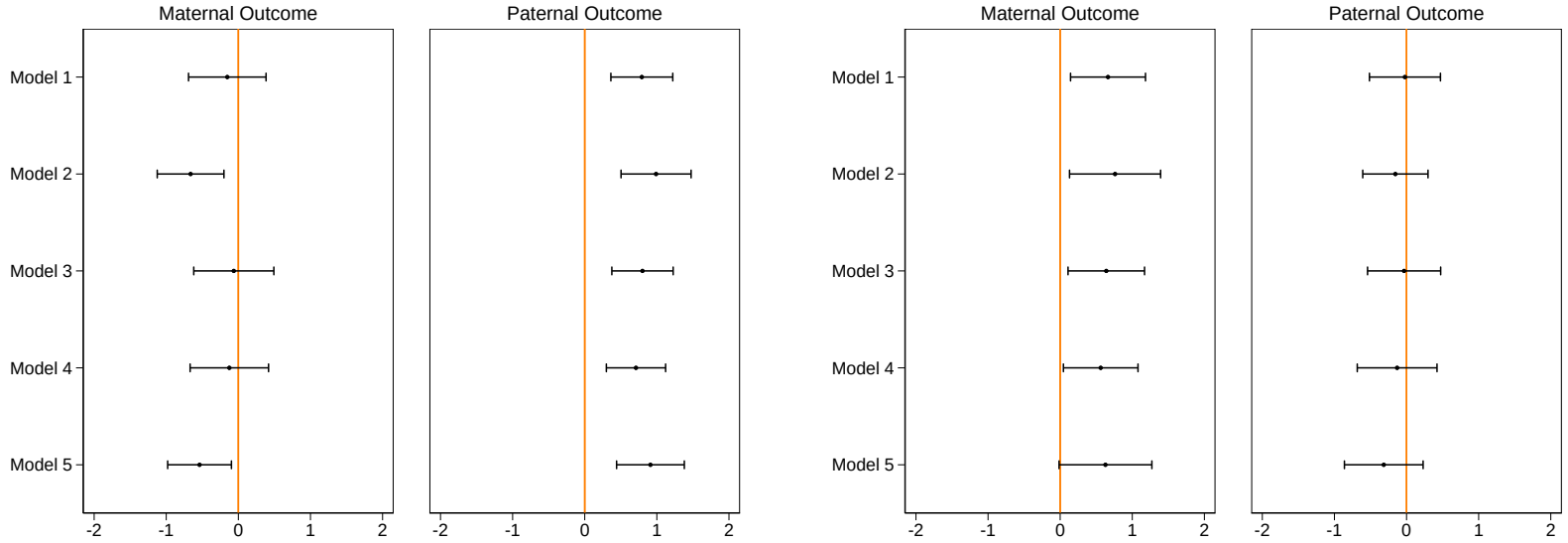



*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on Medicaid coverage and actual eligibility of children. All models include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year) and parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family). All models except model 2 contain state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Model 2 includes state-by-year-by-race-ethnicity fixed effects. Model 3 includes state-by-age-by-race-ethnicity fixed effects. Model 4 includes year-by-age-by-race-ethnicity fixed effects. Model 5 includes state-by-year-by-race-ethnicity, state-by-age-by-race-ethnicity, and year-by-age-by-race-ethnicity fixed effects. Regressions are weighted with parental survey weights divided by number of children per family. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure A.7:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Parental Usual Hours Worked per Week
Robustness to Identifying Assumption

(a) Parents with White Children

(b) Parents with Non-White Children

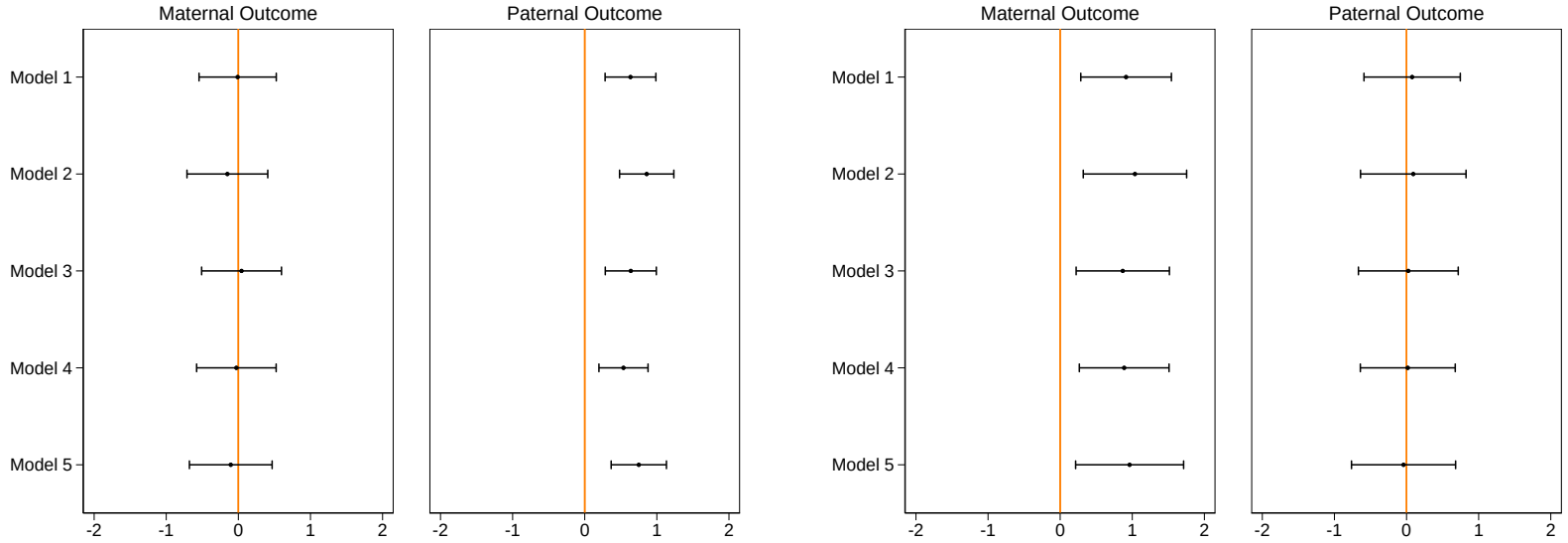



*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on parental usual hours worked per week last year. All models include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year) and parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family). All models except model 2 contain state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Model 2 includes state-by-year-by-race-ethnicity fixed effects. Model 3 includes state-by-age-by-race-ethnicity fixed effects. Model 4 includes year-by-age-by-race-ethnicity fixed effects. Model 5 includes state-by-year-by-race-ethnicity, state-by-age-by-race-ethnicity, and year-by-age-by-race-ethnicity fixed effects. Regressions are weighted with parental survey weights divided by number of children per family. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure A.8:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Parental Weeks Worked per Year
Robustness to Identifying Assumption

(a) Parents with White Children

(b) Parents with Non-White Children

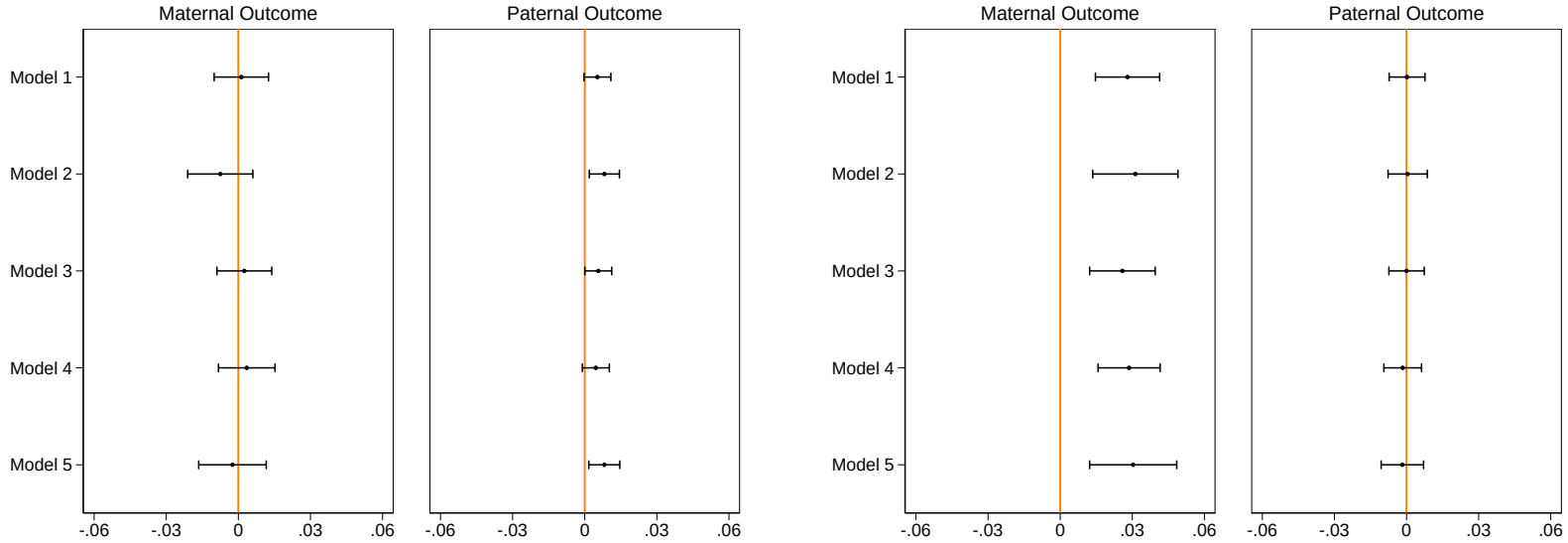



*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on parental weeks worked last year. All models include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year) and parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family). All models except model 2 contain state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Model 2 includes state-by-year-by-race-ethnicity fixed effects. Model 3 includes state-by-age-by-race-ethnicity fixed effects. Model 4 includes year-by-age-by-race-ethnicity fixed effects. Model 5 includes state-by-year-by-race-ethnicity, state-by-age-by-race-ethnicity, and year-by-age-by-race-ethnicity fixed effects. Regressions are weighted with parental survey weights divided by number of children per family. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure A.9:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Parental Labor Force Participation
Robustness to Identifying Assumption

(a) Parents with White Children

(b) Parents with Non-White Children

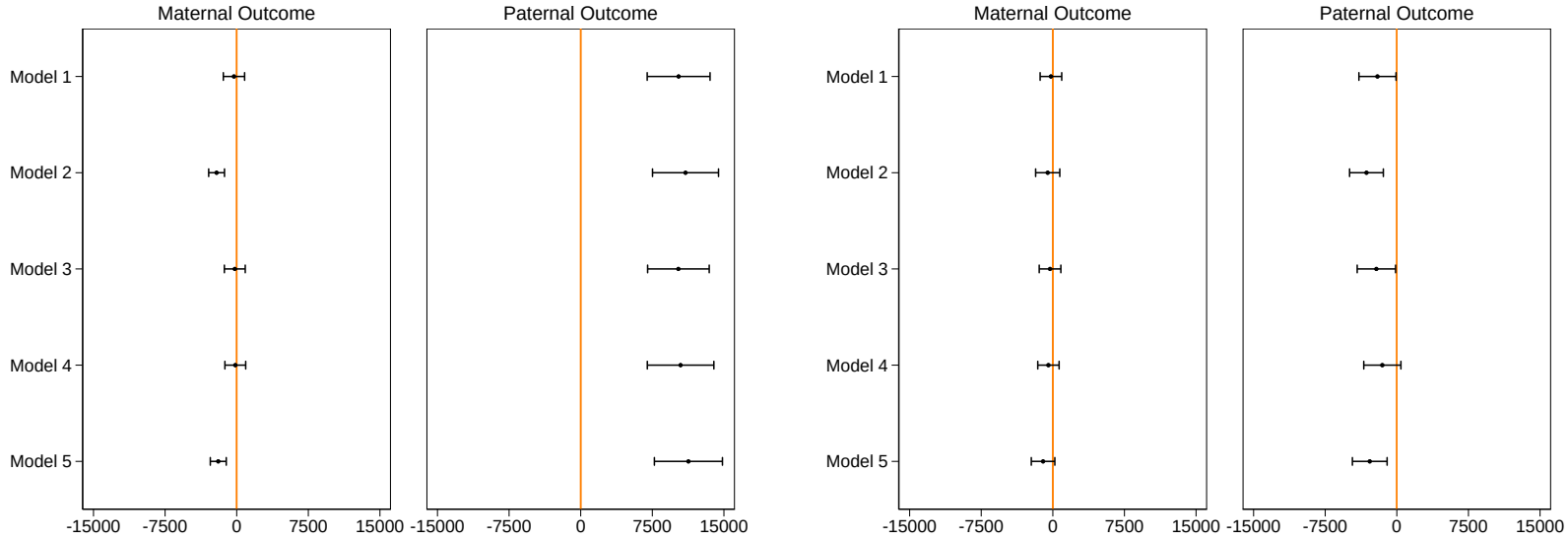



*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on parental labor force participation last week. All models include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year) and parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family). All models except model 2 contain state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Model 2 includes state-by-year-by-race-ethnicity fixed effects. Model 3 includes state-by-age-by-race-ethnicity fixed effects. Model 4 includes year-by-age-by-race-ethnicity fixed effects. Model 5 includes state-by-year-by-race-ethnicity, state-by-age-by-race-ethnicity, and year-by-age-by-race-ethnicity fixed effects. Regressions are weighted with parental survey weights divided by number of children per family. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure A.10:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Parental Annual Earnings ($2020)
Robustness to Identifying Assumption

(a) Parents with White Children

(b) Parents with Non-White Children




*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of race-ethnicity-specific total simulated eligibility on parental annual earnings ($2020) last year. All models include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year) and parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family). All models except model 2 contain state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Model 2 includes state-by-year-by-race-ethnicity fixed effects. Model 3 includes state-by-age-by-race-ethnicity fixed effects. Model 4 includes year-by-age-by-race-ethnicity fixed effects. Model 5 includes state-by-year-by-race-ethnicity, state-by-age-by-race-ethnicity, and year-by-age-by-race-ethnicity fixed effects. Regressions are weighted with parental survey weights divided by number of children per family. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Table A.1:
Child-Level and Parent-Level Demographic Characteristics

| | (1) | (2) |
|---|---|---|
| Fraction Female | 0.49 | 0.49 |
| | (0.50) | (0.50) |
| Fraction White | 0.64 | 0.65 |
| | (0.48) | (0.48) |
| Fraction Black | 0.14 | 0.14 |
| | (0.35) | (0.34) |
| Fraction Other | 0.03 | 0.03 |
| | (0.17) | (0.18) |
| Fraction Hispanic | 0.16 | 0.15 |
| | (0.36) | (0.36) |
| Child's Age | 8.69 | 8.84 |
| | (5.68) | (5.60) |
| Fraction under Age 6 | 0.35 | 0.34 |
| | (0.48) | (0.47) |
| Fraction Age 6-11 | 0.29 | 0.30 |
| | (0.45) | (0.46) |
| Fraction Age 12-18 | 0.36 | 0.36 |
| | (0.48) | (0.48) |
| Fraction under 100% of FPL | 0.17 | 0.15 |
| | (0.37) | (0.36) |
| Fraction 100%-133% of FPL | 0.07 | 0.06 |
| | (0.25) | (0.24) |
| Fraction 133%-185% of FPL | 0.11 | 0.10 |
| | (0.31) | (0.30) |
| Fraction 185%-300% of FPL | 0.22 | 0.22 |
| | (0.42) | (0.41) |
| Fraction above 300% of FPL | 0.43 | 0.47 |
| | (0.50) | (0.50) |
| Fraction with Two Parents | 0.71 | 0.73 |
| | (0.45) | (0.44) |
| Fraction with a Single Mother | 0.24 | 0.23 |
| | (0.43) | (0.42) |
| Fraction with a Single Father | 0.04 | 0.05 |
| | (0.21) | (0.21) |

continued on next page

Table A.1:
Child-Level and Parent-Level Demographic Characteristics (continued)

| | (1) | (2) |
|---|---|---|
| Number of Children per Family | 1.88 | 1.84 |
| | (0.99) | (0.96) |
| Number of Children under 6 per Family | 0.62 | 0.60 |
| | (0.79) | (0.78) |
| Number of Children under 12 per Family | 1.22 | 1.21 |
| | (1.03) | (1.02) |
| Maternal Age | 35.96 | 36.51 |
| | (8.47) | (8.12) |
| Paternal Age | 38.77 | 39.14 |
| | (8.83) | (8.48) |
| Fraction with Mother Attained HS or Less | 0.50 | 0.49 |
| | (0.50) | (0.50) |
| Fraction with Father Attained HS or Less | 0.47 | 0.48 |
| | (0.50) | (0.50) |
| Number of Children with at least a Mother | 1,420,269 | 6,923,107 |
| Number of Children with at least a Father | 1,153,628 | 5,795,794 |

*Notes:* The table shows demographics characteristics of children and their parents observed in CPS ASEC 1980-2015 for the first time (column 1) and census sample 1990, 2000, and ACS 2010 (column 1) The sample is restricted to children age 0-18 with parents age 20-64. Means are weighted with cross-sectional survey weights of parents divided by number of children per family. Standard deviations are shown in parenthesis.

Table A.2:
Effect of Race-Ethnicity-Specific Total Simula−
ted Eligibility on Parental Educational Outcomes

| | Maternal Educational Attainment | | | Paternal Educational Attainment | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | | | No High School | | | |
| SIMT | -0.01** | 0.00 | -0.02*** | 0.00 | 0.01 | -0.01 |
| | (0.00) | (0.00) | (0.01) | (0.01) | (0.01) | (0.01) |
| Observations | 1,331,513 | 839,256 | 492,257 | 1,081,834 | 740,261 | 341,573 |
| Adjusted $R^2$ | 0.13 | 0.07 | 0.09 | 0.11 | 0.05 | 0.07 |
| Mean Y - Baseline | 0.23 | 0.18 | 0.39 | 0.22 | 0.18 | 0.38 |
| Mean Y - Overall | 0.15 | 0.09 | 0.25 | 0.14 | 0.09 | 0.26 |
| | | | High School | | | |
| SIMT | 0.01 | -0.01** | 0.03*** | -0.00 | -0.02** | 0.02*** |
| | (0.00) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) |
| Observations | 1,331,513 | 839,256 | 492,257 | 1,081,834 | 740,261 | 341,573 |
| Adjusted $R^2$ | 0.06 | 0.07 | 0.02 | 0.04 | 0.05 | 0.03 |
| Mean Y - Baseline | 0.45 | 0.47 | 0.38 | 0.35 | 0.35 | 0.33 |
| Mean Y - Overall | 0.35 | 0.36 | 0.34 | 0.33 | 0.33 | 0.33 |
| | | | Some College | | | |
| SIMT | 0.00 | -0.00 | 0.02*** | -0.01* | -0.01*** | 0.00 |
| | (0.00) | (0.01) | (0.00) | (0.00) | (0.00) | (0.01) |
| Observations | 1,331,513 | 839,256 | 492,257 | 1,081,834 | 740,261 | 341,573 |
| Adjusted $R^2$ | 0.02 | 0.02 | 0.02 | 0.02 | 0.02 | 0.01 |
| Mean Y - Baseline | 0.19 | 0.20 | 0.15 | 0.20 | 0.21 | 0.17 |
| Mean Y - Overall | 0.27 | 0.29 | 0.24 | 0.24 | 0.25 | 0.21 |
| | | | College or More | | | |
| SIMT | -0.00 | 0.01** | -0.02*** | 0.00 | 0.02*** | -0.02** |
| | (0.00) | (0.01) | (0.00) | (0.01) | (0.01) | (0.01) |
| Observations | 1,331,513 | 839,256 | 492,257 | 1,081,834 | 740,261 | 341,573 |
| Adjusted $R^2$ | 0.16 | 0.16 | 0.11 | 0.11 | 0.10 | 0.09 |
| Mean Y - Baseline | 0.13 | 0.15 | 0.08 | 0.23 | 0.26 | 0.12 |
| Mean Y - Overall | 0.23 | 0.27 | 0.16 | 0.28 | 0.32 | 0.20 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental educational attainment (indicator for no high school, high school, some college, and college or more). Usual hours worked per week and weeks worked last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.3:
Effect of Race-Ethnicity-Specific Total Simu−
lated Eligibility on Parental Marital Outcomes

| | Maternal Marital Outcomes | | | Paternal Marital Outcomes | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | | | Married | | | |
| SIMT | 0.01*** | 0.00 | 0.03*** | 0.02*** | 0.02*** | 0.02*** |
| | (0.00) | (0.00) | (0.01) | (0.00) | (0.00) | (0.01) |
| Observations | 1,331,513 | 839,256 | 492,257 | 1,081,834 | 740,261 | 341,573 |
| Adjusted $R^2$ | 0.13 | 0.08 | 0.10 | 0.07 | 0.06 | 0.07 |
| Mean Y - Baseline | 0.80 | 0.85 | 0.64 | 0.97 | 0.98 | 0.94 |
| Mean Y - Overall | 0.72 | 0.80 | 0.59 | 0.90 | 0.92 | 0.86 |
| | | | Never Married | | | |
| SIMT | -0.02*** | -0.02*** | -0.01 | -0.02*** | -0.02*** | -0.03*** |
| | (0.00) | (0.00) | (0.01) | (0.00) | (0.00) | (0.01) |
| Observations | 1,331,513 | 839,256 | 492,257 | 1,081,834 | 740,261 | 341,573 |
| Adjusted $R^2$ | 0.21 | 0.13 | 0.17 | 0.11 | 0.08 | 0.10 |
| Mean Y - Baseline | 0.04 | 0.01 | 0.13 | 0.01 | 0.00 | 0.02 |
| Mean Y - Overall | 0.11 | 0.05 | 0.22 | 0.04 | 0.03 | 0.08 |
| | | | Ever Married | | | |
| SIMT | 0.02*** | 0.02*** | 0.01 | 0.02*** | 0.02*** | 0.03*** |
| | (0.00) | (0.00) | (0.01) | (0.00) | (0.00) | (0.01) |
| Observations | 1,331,513 | 839,256 | 492,257 | 1,081,834 | 740,261 | 341,573 |
| Adjusted $R^2$ | 0.21 | 0.13 | 0.17 | 0.11 | 0.08 | 0.10 |
| Mean Y - Baseline | 0.96 | 0.99 | 0.87 | 0.99 | 1.00 | 0.98 |
| Mean Y - Overall | 0.89 | 0.95 | 0.78 | 0.96 | 0.97 | 0.92 |
| | | | Divorced | | | |
| SIMT | 0.01*** | 0.01*** | 0.00 | -0.00 | -0.00 | 0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| Observations | 1,331,513 | 839,256 | 492,257 | 1,081,834 | 740,261 | 341,573 |
| Adjusted $R^2$ | 0.04 | 0.04 | 0.04 | 0.03 | 0.03 | 0.02 |
| Mean Y - Baseline | 0.09 | 0.09 | 0.09 | 0.01 | 0.01 | 0.02 |
| Mean Y - Overall | 0.10 | 0.11 | 0.10 | 0.04 | 0.04 | 0.04 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental marital outcomes (indicator for married, never married, ever married, and divorced). Usual hours worked per week and weeks worked last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.4:
Effect of Simulated Eligibility on Child's Medicaid Coverage

| | Non-Race-Ethnicity-Specific Simulated Eligibility | | |
|---|---|---|---|
| | Child-Level Medicaid Coverage | Child-Level Medicaid Coverage | Family-Level Medicaid Coverage |
| SIM | 0.10*** | 0.08*** | |
| | (0.02) | (0.02) | |
| SIMS | | 0.02** | |
| | | (0.01) | |
| SIMT | | | 0.35*** |
| | | | (0.06) |
| Observations | 1,418,012 | 1,418,012 | 1,418,012 |
| Adjusted $R^2$ | 0.21 | 0.21 | 0.30 |
| Mean Y - Baseline | 0.10 | 0.10 | 0.23 |
| Mean Y - Overall | 0.22 | 0.22 | 0.44 |
| | Race-Ethnicity-Specific Simulated Eligibility (SHADAC) | | |
| | Child-Level Medicaid Coverage | Child-Level Medicaid Coverage | Family-Level Medicaid Coverage |
| SIM | 0.06*** | 0.04*** | |
| | (0.02) | (0.01) | |
| SIMS | | 0.02** | |
| | | (0.01) | |
| SIMT | | | 0.23*** |
| | | | (0.05) |
| Observations | 930,776 | 930,776 | 930,776 |
| Adjusted $R^2$ | 0.22 | 0.22 | 0.34 |
| Mean Y - Baseline | 0.14 | 0.14 | 0.30 |
| Mean Y - Overall | 0.23 | 0.23 | 0.46 |

*Notes:* This table shows results from regressions estimating the effect of simulated eligibility on child-level and family-level Medicaid coverage. SIM, SIMS, and SIMT refers to child's own, sibling's total, and family's total simulated eligibility, respectively. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). The controls in regressions using the full sample and race-specific simulated eligibility are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015 in the top panel and CPS ASEC 1988-2013 in the bottom panel. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.5:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Parental Usual Hours Worked per Week

| | Maternal Labor Supply | | | Paternal Labor Supply | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Positive Hours | | | | | |
| SIMT | 0.004 | -0.004 | 0.015** | 0.005* | 0.008** | 0.000 |
| | (0.004) | (0.005) | (0.007) | (0.003) | (0.003) | (0.005) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.06 | 0.06 | 0.06 | 0.03 | 0.02 | 0.03 |
| Mean Y - Baseline | 0.65 | 0.65 | 0.62 | 0.94 | 0.95 | 0.91 |
| Mean Y - Overall | 0.72 | 0.74 | 0.67 | 0.94 | 0.95 | 0.91 |
| | Part-Time Employment | | | | | |
| SIMT | -0.003 | -0.002 | -0.005 | -0.001 | -0.004** | 0.002 |
| | (0.004) | (0.005) | (0.004) | (0.002) | (0.002) | (0.003) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.03 | 0.02 | 0.01 | 0.01 | 0.01 | 0.01 |
| Mean Y - Baseline | 0.21 | 0.24 | 0.13 | 0.02 | 0.02 | 0.03 |
| Mean Y - Overall | 0.20 | 0.24 | 0.15 | 0.04 | 0.03 | 0.05 |
| | Full-Time Employment | | | | | |
| SIMT | 0.007 | -0.003 | 0.021*** | 0.007* | 0.012*** | -0.002 |
| | (0.005) | (0.008) | (0.006) | (0.004) | (0.004) | (0.006) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.06 | 0.07 | 0.05 | 0.03 | 0.02 | 0.03 |
| Mean Y - Baseline | 0.43 | 0.41 | 0.49 | 0.92 | 0.93 | 0.88 |
| Mean Y - Overall | 0.51 | 0.50 | 0.53 | 0.90 | 0.92 | 0.86 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on likelihood of parents working any hours last year, working part time last year (>0 and <35 hours per week), and working full time last year ($\geq$ 35 hours per week). All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.6:
Effect of Race-Ethnicity-Specific Simulated Eligibility on Maternal Weeks Worked per Year

| | All | White | Non-White | All | White | Non-White | All | White | Non-White |
|---|---|---|---|---|---|---|---|---|---|
| SIM | 0.46 | 0.13 | 1.05 | 0.16 | 0.19 | 0.48 | | | |
| | (0.77) | (0.99) | (0.93) | (0.82) | (1.03) | (0.94) | | | |
| SIMS | | | | 0.43** | -0.08 | 0.98*** | | | |
| | | | | (0.20) | (0.27) | (0.35) | | | |
| SIMT | | | | | | | 0.38** | -0.02 | 0.90*** |
| | | | | | | | (0.18) | (0.26) | (0.31) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,375,551 | 863,738 | 511,813 | 1,375,551 | 863,738 | 511,813 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.09 | 0.09 | 0.09 | 0.09 | 0.09 | 0.09 | 0.09 |
| Mean Y - Baseline | 25.41 | 25.60 | 24.86 | 25.41 | 25.60 | 24.86 | 25.41 | 25.60 | 24.86 |
| Mean Y - Overall | 31.44 | 32.43 | 29.71 | 31.44 | 32.43 | 29.71 | 31.44 | 32.43 | 29.71 |

*Notes:* This table shows results from regressions estimating the effect of race-specific simulated eligibility on maternal weeks worked last year. SIM, SIMS, and SIMT refers to child's own, sibling's total, and family's total simulated eligibility, respectively. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.7:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Maternal Labor Supply by Marital Status

| | All | | | White | | | Non-White | | |
|---|---|---|---|---|---|---|---|---|---|
| | All | Single | Married | All | Single | Married | All | Single | Married |
| | Usual Hours Worked per Week | | | | | | | | |
| SIMT | 0.06 | 1.20*** | -0.27 | -0.18 | 0.18 | -0.26 | 0.36 | 1.98*** | -0.28 |
| | (0.16) | (0.31) | (0.20) | (0.26) | (0.36) | (0.30) | (0.25) | (0.43) | (0.31) |
| Observations | 1,379,087 | 307,333 | 1,071,754 | 865,603 | 131,352 | 734,251 | 513,484 | 175,981 | 337,503 |
| Adjusted $R^2$ | 0.09 | 0.10 | 0.07 | 0.09 | 0.08 | 0.08 | 0.07 | 0.10 | 0.06 |
| Mean Y - Baseline | 21.89 | 26.10 | 20.77 | 21.62 | 28.96 | 20.29 | 22.69 | 22.66 | 22.71 |
| Mean Y - Overall | 25.63 | 27.95 | 24.84 | 25.99 | 30.30 | 25.05 | 25.01 | 26.04 | 24.38 |
| | Weeks Worked per Year | | | | | | | | |
| SIMT | 0.15 | 1.18*** | -0.15 | -0.19 | -0.03 | -0.22 | 0.57* | 2.10*** | -0.04 |
| | (0.15) | (0.42) | (0.18) | (0.24) | (0.51) | (0.27) | (0.30) | (0.60) | (0.31) |
| Observations | 1,379,087 | 307,333 | 1,071,754 | 865,603 | 131,352 | 734,251 | 513,484 | 175,981 | 337,503 |
| Adjusted $R^2$ | 0.10 | 0.13 | 0.09 | 0.10 | 0.10 | 0.09 | 0.09 | 0.13 | 0.07 |
| Mean Y - Baseline | 25.33 | 29.10 | 24.33 | 25.53 | 33.14 | 24.14 | 24.75 | 24.23 | 25.07 |
| Mean Y - Overall | 31.52 | 32.82 | 31.08 | 32.53 | 35.66 | 31.83 | 29.79 | 30.49 | 29.36 |
| | Labor Force Participation | | | | | | | | |
| SIMT | 0.01* | 0.03*** | -0.00 | -0.00 | 0.01 | -0.01 | 0.02*** | 0.04*** | 0.01 |
| | (0.00) | (0.01) | (0.00) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) |
| Observations | 1,333,837 | 297,260 | 1,036,577 | 840,427 | 127,690 | 712,737 | 493,410 | 169,570 | 323,840 |
| Adjusted $R^2$ | 0.08 | 0.09 | 0.07 | 0.08 | 0.07 | 0.08 | 0.08 | 0.09 | 0.06 |
| Mean Y - Baseline | 0.57 | 0.66 | 0.54 | 0.57 | 0.73 | 0.54 | 0.57 | 0.58 | 0.56 |
| Mean Y - Overall | 0.68 | 0.73 | 0.66 | 0.70 | 0.78 | 0.68 | 0.66 | 0.70 | 0.64 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on maternal labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race and marital status indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.8:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Paternal Labor Supply by Marital Status

| | All | | | White | | | Non-White | | |
|---|---|---|---|---|---|---|---|---|---|
| | All | Single | Married | All | Single | Married | All | Single | Married |
| | Usual Hours Worked per Week | | | | | | | | |
| SIMT | 0.47*** | 0.17 | 0.48*** | 0.58*** | -0.24 | 0.61*** | 0.31 | 0.91 | 0.29 |
| | (0.15) | (0.68) | (0.16) | (0.19) | (0.75) | (0.19) | (0.26) | (1.11) | (0.26) |
| Observations | 1,119,770 | 48,255 | 1,071,515 | 763,405 | 29,175 | 734,230 | 356,365 | 19,080 | 337,285 |
| Adjusted $R^2$ | 0.05 | 0.07 | 0.04 | 0.03 | 0.04 | 0.02 | 0.04 | 0.04 | 0.03 |
| Mean Y - Baseline | 42.07 | 38.08 | 42.21 | 42.93 | 40.98 | 42.98 | 38.72 | 33.65 | 39.07 |
| Mean Y - Overall | 41.61 | 36.08 | 41.97 | 42.92 | 38.35 | 43.17 | 38.71 | 32.83 | 39.22 |
| | Weeks Worked per Year | | | | | | | | |
| SIMT | 0.43** | -0.15 | 0.45** | 0.42*** | -0.81 | 0.47*** | 0.45 | 1.03 | 0.43 |
| | (0.18) | (0.86) | (0.17) | (0.16) | (0.87) | (0.16) | (0.34) | (1.50) | (0.34) |
| Observations | 1,119,770 | 48,255 | 1,071,515 | 763,405 | 29,175 | 734,230 | 356,365 | 19,080 | 337,285 |
| Adjusted $R^2$ | 0.06 | 0.07 | 0.04 | 0.04 | 0.04 | 0.03 | 0.06 | 0.05 | 0.04 |
| Mean Y - Baseline | 46.50 | 40.04 | 46.73 | 47.31 | 43.24 | 47.42 | 43.38 | 35.17 | 43.94 |
| Mean Y - Overall | 46.06 | 39.89 | 46.45 | 47.02 | 42.26 | 47.27 | 43.92 | 36.50 | 44.56 |
| | Labor Force Participation | | | | | | | | |
| SIMT | 0.00 | 0.01 | 0.00 | 0.00 | 0.01 | 0.00 | 0.00 | 0.01 | 0.00 |
| | (0.00) | (0.01) | (0.00) | (0.00) | (0.01) | (0.00) | (0.00) | (0.03) | (0.00) |
| Observations | 1,060,726 | 46,085 | 1,014,641 | 725,531 | 27,908 | 697,623 | 335,195 | 18,177 | 317,018 |
| Adjusted $R^2$ | 0.05 | 0.05 | 0.04 | 0.04 | 0.04 | 0.03 | 0.05 | 0.04 | 0.04 |
| Mean Y - Baseline | 0.96 | 0.87 | 0.96 | 0.97 | 0.91 | 0.97 | 0.92 | 0.82 | 0.93 |
| Mean Y - Overall | 0.94 | 0.86 | 0.95 | 0.95 | 0.89 | 0.96 | 0.92 | 0.82 | 0.93 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on paternal labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race and marital status indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.9:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Geographic Mobility

| | All | White | Non-White |
|---|---|---|---|
| | Probability of Linking | | |
| SIMT | 0.006 | 0.009* | 0.003 |
| | (0.005) | (0.006) | (0.008) |
| Observations | 813,767 | 549,484 | 264,283 |
| Adjusted $R^2$ | 0.16 | 0.18 | 0.11 |
| Mean Y - Baseline | 0.567 | 0.586 | 0.499 |
| Mean Y - Overall | 0.500 | 0.526 | 0.452 |
| | Within State Mobility | | |
| SIMT | -0.010*** | -0.011** | -0.008* |
| | (0.003) | (0.005) | (0.004) |
| Observations | 1,370,392 | 861,007 | 509,385 |
| Adjusted $R^2$ | 0.08 | 0.09 | 0.07 |
| Mean Y - Baseline | 0.137 | 0.127 | 0.164 |
| Mean Y - Overall | 0.145 | 0.128 | 0.171 |
| | Across State Mobility | | |
| SIMT | -0.002 | -0.003 | -0.001 |
| | (0.002) | (0.002) | (0.002) |
| Observations | 1,370,392 | 861,007 | 509,385 |
| Adjusted $R^2$ | 0.02 | 0.02 | 0.02 |
| Mean Y - Baseline | 0.033 | 0.032 | 0.036 |
| Mean Y - Overall | 0.031 | 0.031 | 0.031 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on geographic mobility of children. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.10:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Family-Level Government Transfers

| | Welfare | SNAP | Disability |
|---|---|---|---|
| SIMT | -224.76*** | -52.47 | -43.37*** |
| | (72.39) | (41.11) | (12.12) |
| Observation | 1,189,020 | 1,189,020 | 1,189,020 |
| Adjusted $R^2$ | 0.11 | 0.15 | 0.00 |
| Mean Y - Baseline | 884.34 | 503.98 | 0.00 |
| Mean Y - Overall | 505.88 | 514.31 | 113.66 |
| | SSI | Unemployment | School Lunch |
| SIMT | 6.26 | -4.62 | 169.33*** |
| | (20.22) | (21.65) | (31.37) |
| Observation | 1,189,020 | 1,189,020 | 1,189,020 |
| Adjusted $R^2$ | 0.01 | 0.02 | 0.34 |
| Mean Y - Baseline | 75.20 | 967.15 | 0.00 |
| Mean Y - Overall | 180.06 | 521.94 | 166.49 |
| | Education | Housing Subsidy | Energy Subsidy |
| SIMT | -42.62*** | 78.65*** | -12.13** |
| | (13.46) | (16.23) | (4.75) |
| Observation | 1,189,020 | 1,189,020 | 1,189,020 |
| Adjusted $R^2$ | 0.01 | 0.07 | 0.03 |
| Mean Y - Baseline | 0.00 | 0.00 | 0.00 |
| Mean Y - Overall | 261.92 | 130.75 | 39.84 |

*Notes:* This table shows results from regression estimating the effect of income-specific total simulated eligibility on family-level annual income in 2020 dollars from public assistance (welfare), Supplemental Nutrition Assistance Program (SNAP), disability income, Supplemental Security Income (SSI), unemployment compensation, monetary value of school lunch, educational assistance, housing subsidy, and energy subsidy. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1981-2013 and MSIS 1980-2012. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$, respectively.

Table A.11:
Effect of Race-Ethnicity-Specific Total Simulated Eligi−
bility on Labor Supply of Mothers with White Children
Robustness to Sample Selection

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Usual Hours Worked per Week | | | | | |
| SIMT | -0.17 | -0.19 | -0.06 | -0.49 | -0.21 | -0.16 |
| | (0.26) | (0.27) | (0.27) | (0.37) | (0.27) | (0.27) |
| Observations | 863,738 | 854,210 | 853,560 | 656,150 | 855,147 | 862,573 |
| Adjusted $R^2$ | 0.08 | 0.08 | 0.08 | 0.09 | 0.08 | 0.08 |
| Mean Y - Baseline | 21.61 | 21.58 | 21.72 | 21.61 | 21.61 | 21.61 |
| Mean Y - Overall | 25.91 | 25.91 | 25.97 | 25.62 | 25.83 | 25.92 |
| | Weeks Worked per Year | | | | | |
| SIMT | -0.02 | -0.03 | 0.08 | -0.22 | -0.04 | 0.01 |
| | (0.26) | (0.27) | (0.27) | (0.34) | (0.27) | (0.26) |
| Observations | 863,738 | 854,210 | 853,560 | 656,150 | 858,407 | 862,573 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.09 | 0.10 | 0.09 | 0.09 |
| Mean Y - Baseline | 25.60 | 25.60 | 25.71 | 25.60 | 25.60 | 25.60 |
| Mean Y - Overall | 32.43 | 32.44 | 32.48 | 31.78 | 32.37 | 32.44 |
| | Labor Force Participation | | | | | |
| SIMT | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 | 0.00 |
| | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) |
| Observations | 838,593 | 829,299 | 829,042 | 655,688 | 837,299 | 837,498 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.07 | 0.08 | 0.07 | 0.07 |
| Mean Y - Baseline | 0.57 | 0.57 | 0.57 | 0.57 | 0.57 | 0.57 |
| Mean Y - Overall | 0.69 | 0.69 | 0.69 | 0.68 | 0.69 | 0.69 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. Column 1 reports estimates using the baseline model. Column 2 reports estimates using a model that drops children from Arizona. Column 3 reports estimates using a model that restricts the sample to children with parents in prime working age (25-54). Column 4 reports estimates using a model that drops children observed between 2008 and 2015. Column 5 reports estimates using a model that excludes observations with imputed outcome of interest. Column 6 reports estimates using a model that drops children in families with nine or more children. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$,* $p < 0.10$.

Table A.12:
Effect of Race-Ethnicity-Specific Total Simulated Eligibili−
ty on Labor Supply of Mothers with Non-White Children
Robustness to Sample Selection

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Usual Hours Worked per Week | | | | | |
| SIMT | 0.66** | 0.74*** | 0.77*** | 0.62** | 0.63** | 0.64** |
| | (0.26) | (0.25) | (0.26) | (0.30) | (0.26) | (0.25) |
| Observations | 511,813 | 500,482 | 504,672 | 358,176 | 504,473 | 510,846 |
| Adjusted $R^2$ | 0.06 | 0.06 | 0.06 | 0.07 | 0.06 | 0.06 |
| Mean Y - Baseline | 22.75 | 22.81 | 22.94 | 22.75 | 22.75 | 22.77 |
| Mean Y - Overall | 24.96 | 25.01 | 25.05 | 24.87 | 24.78 | 24.96 |
| | Weeks Worked per Year | | | | | |
| SIMT | 0.90*** | 1.01*** | 1.03*** | 1.00** | 0.89*** | 0.90*** |
| | (0.31) | (0.31) | (0.32) | (0.40) | (0.32) | (0.32) |
| Observations | 511,813 | 500,482 | 504,672 | 358,176 | 506,583 | 510,846 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.09 | 0.10 | 0.09 | 0.09 |
| Mean Y - Baseline | 24.86 | 24.94 | 25.04 | 24.86 | 24.86 | 24.89 |
| Mean Y - Overall | 29.71 | 29.78 | 29.79 | 29.03 | 29.56 | 29.72 |
| | Labor Force Participation | | | | | |
| SIMT | 0.03*** | 0.03*** | 0.03*** | 0.03*** | 0.03*** | 0.03*** |
| | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) | (0.01) |
| Observations | 491,785 | 480,990 | 485,024 | 357,877 | 490,258 | 490,846 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.07 | 0.08 | 0.07 | 0.07 |
| Mean Y - Baseline | 0.57 | 0.57 | 0.57 | 0.57 | 0.57 | 0.57 |
| Mean Y - Overall | 0.66 | 0.66 | 0.66 | 0.65 | 0.66 | 0.66 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. Column 1 reports estimates using the baseline model. Column 2 reports estimates using a model that drops children from Arizona. Column 3 reports estimates using a model that restricts the sample to children with parents in prime working age (25-54). Column 4 reports estimates using a model that drops children observed between 2008 and 2015. Column 5 reports estimates using a model that excludes observations with imputed outcome of interest. Column 6 reports estimates using a model that drops children in families with nine or more children. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$,* $p < 0.10$.

Table A.13:
Effect of Race-Ethnicity-Specific Total Simulated Eligi−
bility on Labor Supply of Fathers with White Children
Robustness to Sample Selection

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Usual Hours Worked per Week | | | | | |
| SIMT | 0.80*** | 0.80*** | 0.80*** | 0.59** | 0.80*** | 0.85*** |
| | (0.21) | (0.22) | (0.21) | (0.25) | (0.22) | (0.22) |
| Observations | 762,111 | 753,813 | 736,632 | 577,507 | 754,221 | 761,029 |
| Adjusted $R^2$ | 0.03 | 0.03 | 0.02 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 42.93 | 42.93 | 43.18 | 42.93 | 42.93 | 42.93 |
| Mean Y - Overall | 42.93 | 42.94 | 43.16 | 43.22 | 42.93 | 42.93 |
| | Weeks Worked per Year | | | | | |
| SIMT | 0.63*** | 0.64*** | 0.58*** | 0.51** | 0.65*** | 0.65*** |
| | (0.18) | (0.18) | (0.17) | (0.22) | (0.18) | (0.18) |
| Observations | 762,111 | 753,813 | 736,632 | 577,507 | 757,789 | 761,029 |
| Adjusted $R^2$ | 0.03 | 0.03 | 0.02 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 47.32 | 47.32 | 47.60 | 47.32 | 47.32 | 47.32 |
| Mean Y - Overall | 47.04 | 47.04 | 47.27 | 47.19 | 47.04 | 47.04 |
| | Labor Force Participation | | | | | |
| SIMT | 0.01* | 0.01** | 0.01** | 0.00 | 0.01* | 0.01** |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| Observations | 724,271 | 716,412 | 700,063 | 564,636 | 723,253 | 723,259 |
| Adjusted $R^2$ | 0.03 | 0.03 | 0.01 | 0.04 | 0.03 | 0.03 |
| Mean Y - Baseline | 0.97 | 0.97 | 0.97 | 0.97 | 0.97 | 0.97 |
| Mean Y - Overall | 0.95 | 0.95 | 0.96 | 0.96 | 0.95 | 0.95 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. Column 1 reports estimates using the baseline model. Column 2 reports estimates using a model that drops children from Arizona. Column 3 reports estimates using a model that restricts the sample to children with parents in prime working age (25-54). Column 4 reports estimates using a model that drops children observed between 2008 and 2015. Column 5 reports estimates using a model that excludes observations with imputed outcome of interest. Column 6 reports estimates using a model that drops children in families with nine or more children. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$,* $p < 0.10$.

Table A.14:
Effect of Race-Ethnicity-Specific Total Simulated Eligibili−
ty on Labor Supply of Fathers with Non-White Children
Robustness to Sample Selection

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Usual Hours Worked per Week | | | | | |
| SIMT | 0.02 | -0.02 | 0.07 | -0.06 | 0.01 | 0.03 |
| | (0.24) | (0.24) | (0.23) | (0.25) | (0.24) | (0.23) |
| Observations | 355,534 | 346,744 | 341,267 | 246,425 | 348,855 | 354,900 |
| Adjusted $R^2$ | 0.03 | 0.03 | 0.02 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 38.74 | 38.71 | 39.10 | 38.74 | 38.74 | 38.74 |
| Mean Y - Overall | 38.73 | 38.73 | 39.05 | 38.98 | 38.68 | 38.73 |
| | Weeks Worked per Year | | | | | |
| SIMT | 0.10 | 0.11 | 0.19 | -0.18 | 0.08 | 0.10 |
| | (0.32) | (0.34) | (0.32) | (0.32) | (0.33) | (0.32) |
| Observations | 355,534 | 346,744 | 341,267 | 246,425 | 351,166 | 354,900 |
| Adjusted $R^2$ | 0.04 | 0.04 | 0.03 | 0.04 | 0.04 | 0.04 |
| Mean Y - Baseline | 43.45 | 43.41 | 43.92 | 43.45 | 43.45 | 43.46 |
| Mean Y - Overall | 43.93 | 43.92 | 44.28 | 43.95 | 43.91 | 43.94 |
| | Labor Force Participation | | | | | |
| SIMT | 0.00 | 0.00 | 0.00 | -0.00 | 0.00 | 0.00 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| Observations | 334,394 | 326,208 | 320,956 | 240,791 | 333,407 | 333,797 |
| Adjusted $R^2$ | 0.04 | 0.04 | 0.01 | 0.04 | 0.04 | 0.04 |
| Mean Y - Baseline | 0.92 | 0.92 | 0.94 | 0.92 | 0.92 | 0.92 |
| Mean Y - Overall | 0.92 | 0.92 | 0.93 | 0.92 | 0.92 | 0.92 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. Column 1 reports estimates using the baseline model. Column 2 reports estimates using a model that drops children from Arizona. Column 3 reports estimates using a model that restricts the sample to children with parents in prime working age (25-54). Column 4 reports estimates using a model that drops children observed between 2008 and 2015. Column 5 reports estimates using a model that excludes observations with imputed outcome of interest. Column 6 reports estimates using a model that drops children in families with nine or more children. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$,* $p < 0.10$.

Table A.15:
Effect of Race-Ethnicity-Specific Total Simulated Eligi−
bility on Labor Supply of Parents with White Children
Robustness to Identifying Assumption

| | Maternal Labor Supply | | | Paternal Labor Supply | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (1) | (2) | (3) |
| | Usual Hours Worked per Week | | | | | |
| SIMT | -0.55** | -0.56** | -0.59*** | 0.92*** | 0.90*** | 0.90*** |
| | (0.22) | (0.22) | (0.22) | (0.23) | (0.23) | (0.22) |
| Observations | 863,738 | 863,738 | 863,738 | 762,111 | 762,111 | 762,111 |
| Adjusted $R^2$ | 0.08 | 0.08 | 0.09 | 0.03 | 0.03 | 0.04 |
| Mean Y - Baseline | 21.61 | 21.61 | 21.61 | 42.93 | 42.93 | 42.93 |
| Mean Y - Overall | 25.91 | 25.91 | 25.91 | 42.93 | 42.93 | 42.93 |
| | Weeks Worked per Year | | | | | |
| SIMT | -0.11 | -0.14 | -0.14 | 0.76*** | 0.74*** | 0.73*** |
| | (0.29) | (0.30) | (0.30) | (0.19) | (0.19) | (0.19) |
| Observations | 863,738 | 863,738 | 863,738 | 762,111 | 762,111 | 762,111 |
| Adjusted $R^2$ | 0.09 | 0.10 | 0.10 | 0.04 | 0.04 | 0.04 |
| Mean Y - Baseline | 25.60 | 25.60 | 25.60 | 47.32 | 47.32 | 47.32 |
| Mean Y - Overall | 32.43 | 32.43 | 32.43 | 47.04 | 47.04 | 47.04 |
| | Labor Force Participation | | | | | |
| SIMT | -0.00 | -0.00 | -0.00 | 0.01** | 0.01** | 0.01** |
| | (0.01) | (0.01) | (0.01) | (0.00) | (0.00) | (0.00) |
| Observations | 838,593 | 838,593 | 838,593 | 724,271 | 724,271 | 724,271 |
| Adjusted $R^2$ | 0.08 | 0.08 | 0.08 | 0.03 | 0.04 | 0.04 |
| Mean Y - Baseline | 0.57 | 0.57 | 0.57 | 0.97 | 0.97 | 0.97 |
| Mean Y - Overall | 0.69 | 0.69 | 0.69 | 0.95 | 0.95 | 0.95 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. Column 1 adds state-by-year, state-by-age, and year-by-age fixed effects to the baseline model. Column 2 adds state-by-age linear time trends to model used in column 1. Column 3 adds region-by-year-by-age fixed effects to model used in column 1. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.16:
Effect of Race-Ethnicity-Specific Total Simulated Eligibili−
ty on Labor Supply of Parents with Non-White Children
Robustness to Identifying Assumption

| | Maternal Labor Supply | | | Paternal Labor Supply | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (1) | (2) | (3) |
| | Usual Hours Worked per Week | | | | | |
| SIMT | 0.60* | 0.59* | 0.60* | -0.28 | -0.32 | -0.27 |
| | (0.32) | (0.32) | (0.31) | (0.27) | (0.27) | (0.27) |
| Observations | 511,813 | 511,813 | 511,813 | 355,534 | 355,534 | 355,534 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.07 | 0.04 | 0.04 | 0.05 |
| Mean Y - Baseline | 22.75 | 22.75 | 22.75 | 38.74 | 38.74 | 38.74 |
| Mean Y - Overall | 24.96 | 24.96 | 24.96 | 38.73 | 38.73 | 38.73 |
| | Weeks Worked per Year | | | | | |
| SIMT | 0.93** | 0.94** | 0.92** | -0.02 | -0.06 | -0.00 |
| | (0.37) | (0.37) | (0.36) | (0.35) | (0.35) | (0.36) |
| Observations | 511,813 | 511,813 | 511,813 | 355,534 | 355,534 | 355,534 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.10 | 0.05 | 0.05 | 0.06 |
| Mean Y - Baseline | 24.86 | 24.86 | 24.86 | 43.45 | 43.45 | 43.45 |
| Mean Y - Overall | 29.71 | 29.71 | 29.71 | 43.93 | 43.93 | 43.93 |
| | Labor Force Participation | | | | | |
| SIMT | 0.03*** | 0.03*** | 0.03*** | -0.00 | -0.00 | -0.00 |
| | (0.01) | (0.01) | (0.01) | (0.00) | (0.00) | (0.00) |
| Observations | 491,785 | 491,785 | 491,785 | 334,394 | 334,394 | 334,394 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.08 | 0.05 | 0.05 | 0.06 |
| Mean Y - Baseline | 0.57 | 0.57 | 0.57 | 0.92 | 0.92 | 0.92 |
| Mean Y - Overall | 0.66 | 0.66 | 0.66 | 0.92 | 0.92 | 0.92 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. Column 1 adds state-by-year, state-by-age, and year-by-age fixed effects to the baseline model. Column 2 adds state-by-age linear time trends to model used in column 1. Column 3 adds region-by-year-by-age fixed effects to model used in column 1. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.17:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Parental Annual Earnings
Robustness to Imputation & Topcodes

| | Maternal Annual Earnings | | | Paternal Annual Earnings | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Non-Imputed Observations | | | | | |
| SIMT | -376 | -419 | -310 | 5,209*** | 8,385*** | -1,651* |
| | ( 350) | ( 423) | ( 486) | ( 1,214) | ( 1,457) | ( 832) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.08 | 0.08 | 0.07 | 0.10 | 0.08 | 0.05 |
| Mean Y - Baseline | 15,060 | 14,822 | 15,765 | 63,422 | 67,377 | 48,034 |
| Mean Y - Overall | 24,799 | 26,308 | 22,178 | 66,098 | 73,258 | 50,062 |
| | Rank Proximity Swap Topcodes | | | | | |
| SIMT | -411 | -520 | -240 | 4,782*** | 7,727*** | -1,578** |
| | ( 326) | ( 420) | ( 444) | ( 1,143) | ( 1,372) | ( 740) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.06 | 0.07 | 0.06 | 0.04 |
| Mean Y - Baseline | 14,846 | 14,698 | 15,285 | 64,316 | 68,713 | 47,211 |
| Mean Y - Overall | 24,345 | 25,967 | 21,529 | 66,051 | 73,427 | 49,577 |
| | Cell Means Replacement Topcodes | | | | | |
| SIMT | -367 | -416 | -291 | 4,848*** | 7,820*** | -1,570* |
| | ( 350) | ( 421) | ( 490) | ( 1,163) | ( 1,393) | ( 830) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.08 | 0.07 | 0.07 | 0.09 | 0.08 | 0.05 |
| Mean Y - Baseline | 14,854 | 14,702 | 15,309 | 64,302 | 68,695 | 47,210 |
| Mean Y - Overall | 24,354 | 25,991 | 21,513 | 66,020 | 73,400 | 49,536 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental annual total earnings ($2020) last year. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.18:
Effect of Race-Ethnicity-Specific Total Simulated Eligi−
bility on Labor Supply of Parents with White Children
Robustness to Maternal Eligibility

| | Maternal Labor Supply | | | | Paternal Labor Supply | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (1) | (2) | (3) | (4) |
| | Usual Hours Worked per Week | | | | | | | |
| SIMT | -0.17 | -0.24 | -0.31 | -0.23 | 0.80*** | 0.84*** | 0.80*** | 0.86*** |
| | (0.26) | (0.26) | (0.27) | (0.26) | (0.21) | (0.22) | (0.21 | (0.23) |
| Observations | 863,738 | 863,738 | 812,682 | 863,738 | 762,111 | 762,111 | 713,723 | 762,111 |
| Adjusted $R^2$ | 0.08 | 0.08 | 0.08 | 0.08 | 0.03 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 21.61 | 21.61 | 21.43 | 21.61 | 42.93 | 42.93 | 42.93 | 42.93 |
| Mean Y - Overall | 25.91 | 25.91 | 25.89 | 25.91 | 42.93 | 42.93 | 42.93 | 42.93 |
| | Weeks Worked per Year | | | | | | | |
| SIMT | -0.02 | -0.08 | -0.14 | -0.08 | 0.63*** | 0.65*** | 0.62*** | 0.66*** |
| | (0.26) | (0.25) | (0.26) | (0.25) | (0.18) | (0.18) | (0.17 | (0.19) |
| Observations | 863,738 | 863,738 | 812,682 | 863,738 | 762,111 | 762,111 | 713,723 | 762,111 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.09 | 0.09 | 0.03 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 25.60 | 25.60 | 25.78 | 25.60 | 47.32 | 47.32 | 47.35 | 47.32 |
| Mean Y - Overall | 32.43 | 32.43 | 32.74 | 32.43 | 47.04 | 47.04 | 47.01 | 47.04 |
| | Labor Force Participation | | | | | | | |
| SIMT | 0.00 | 0.00 | -0.00 | 0.00 | 0.01* | 0.01* | 0.00 | 0.01** |
| | (0.01) | (0.01) | (0.01) | (0.01) | (0.00) | (0.00) | (0.00 | (0.00) |
| Observations | 838,593 | 838,593 | 788,756 | 838,593 | 724,271 | 724,271 | 678,446 | 724,271 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.07 | 0.07 | 0.03 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 0.57 | 0.57 | 0.58 | 0.57 | 0.97 | 0.97 | 0.96 | 0.97 |
| Mean Y - Overall | 0.69 | 0.69 | 0.70 | 0.69 | 0.95 | 0.95 | 0.95 | 0.95 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. Column 1 reports estimates using the baseline model. Column 2 reports estimates using maternal eligibility (women age 15-44) for zero-year old children. Column 3 reports estimates dropping children age zero. Column 4 reports estimates using maternal eligibility (mothers with children of age zero) for zero-year old children. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.19:
Effect of Race-Ethnicity-Specific Total Simulated Eligibili−
ty on Labor Supply of Parents with Non-White Children
Robustness to Maternal Eligibility

| | Maternal Labor Supply | | | | Paternal Labor Supply | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (1) | (2) | (3) | (4) |
| | Usual Hours Worked per Week | | | | | | | |
| SIMT | 0.66** | 0.63** | 0.55** | 0.64** | 0.02 | 0.02 | 0.08 | 0.02 |
| | (0.26) | (0.25) | (0.25) | (0.25) | (0.24) | (0.24) | (0.25 | (0.23) |
| Observations | 511,813 | 511,813 | 483,931 | 511,813 | 355,534 | 355,534 | 334,463 | 355,534 |
| Adjusted $R^2$ | 0.06 | 0.06 | 0.06 | 0.06 | 0.03 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 22.75 | 22.75 | 22.89 | 22.75 | 38.74 | 38.74 | 38.93 | 38.74 |
| Mean Y - Overall | 24.96 | 24.96 | 25.19 | 24.96 | 38.73 | 38.73 | 38.67 | 38.73 |
| | Weeks Worked per Year | | | | | | | |
| SIMT | 0.90*** | 0.88*** | 0.77** | 0.90*** | 0.10 | 0.11 | 0.17 | 0.10 |
| | (0.31) | (0.32) | (0.32) | (0.32) | (0.32) | (0.32) | (0.31 | (0.32) |
| Observations | 511,813 | 511,813 | 483,931 | 511,813 | 355,534 | 355,534 | 334,463 | 355,534 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.08 | 0.09 | 0.04 | 0.04 | 0.04 | 0.04 |
| Mean Y - Baseline | 24.86 | 24.86 | 25.36 | 24.86 | 43.45 | 43.45 | 43.63 | 43.45 |
| Mean Y - Overall | 29.71 | 29.71 | 30.23 | 29.71 | 43.93 | 43.93 | 43.88 | 43.93 |
| | Labor Force Participation | | | | | | | |
| SIMT | 0.03*** | 0.03*** | 0.02*** | 0.03*** | 0.00 | -0.00 | 0.00 | 0.00 |
| | (0.01) | (0.01) | (0.01) | (0.01) | (0.00) | (0.00) | (0.00 | (0.00) |
| Observations | 491,785 | 491,785 | 464,819 | 491,785 | 334,394 | 334,394 | 314,575 | 334,394 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.06 | 0.07 | 0.04 | 0.04 | 0.04 | 0.04 |
| Mean Y - Baseline | 0.57 | 0.57 | 0.58 | 0.57 | 0.92 | 0.92 | 0.92 | 0.92 |
| Mean Y - Overall | 0.66 | 0.66 | 0.67 | 0.66 | 0.92 | 0.92 | 0.92 | 0.92 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. Column 1 reports estimates using the baseline model. Column 2 reports estimates using maternal eligibility (women age 15-44) for zero-year old children. Column 3 reports estimates dropping children age zero. Column 4 reports estimates using maternal eligibility (mothers with children of age zero) for zero-year old children. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.20:
Effect of Race-Ethnicity-Specific Total Simulated Eligi−
bility on Labor Supply of Parents with White Children
Robustness to Simulated Eligibility Type

| | Maternal Labor Supply | | | | Paternal Labor Supply | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (1) | (2) | (3) | (4) |
| | Usual Hours Worked per Week | | | | | | | |
| SIMT | -0.17 | -0.05 | -0.06 | -0.07 | 0.80*** | 0.59*** | 0.59*** | 0.61*** |
| | (0.26) | (0.23) | (0.23) | (0.26) | (0.21) | (0.16) | (0.16 | (0.17) |
| Observations | 863,738 | 863,738 | 863,738 | 863,738 | 762,111 | 762,111 | 762,111 | 762,111 |
| Adjusted $R^2$ | 0.08 | 0.08 | 0.08 | 0.08 | 0.03 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 21.61 | 21.61 | 21.61 | 21.61 | 42.93 | 42.93 | 42.93 | 42.93 |
| Mean Y - Overall | 25.91 | 25.91 | 25.91 | 25.91 | 42.93 | 42.93 | 42.93 | 42.93 |
| | Weeks Worked per Year | | | | | | | |
| SIMT | -0.02 | 0.07 | 0.07 | 0.02 | 0.63*** | 0.47*** | 0.47*** | 0.46*** |
| | (0.26) | (0.22) | (0.22) | (0.25) | (0.18) | (0.13) | (0.13 | (0.14) |
| Observations | 863,738 | 863,738 | 863,738 | 863,738 | 762,111 | 762,111 | 762,111 | 762,111 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.09 | 0.09 | 0.03 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 25.60 | 25.60 | 25.60 | 25.60 | 47.32 | 47.32 | 47.32 | 47.32 |
| Mean Y - Overall | 32.43 | 32.43 | 32.43 | 32.43 | 47.04 | 47.04 | 47.04 | 47.04 |
| | Labor Force Participation | | | | | | | |
| SIMT | 0.00 | 0.00 | 0.00 | 0.00 | 0.01* | 0.00* | 0.00* | 0.00 |
| | (0.01) | (0.00) | (0.00) | (0.01) | (0.00) | (0.00) | (0.00 | (0.00) |
| Observations | 838,593 | 838,593 | 838,593 | 838,593 | 724,271 | 724,271 | 724,271 | 724,271 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.07 | 0.07 | 0.03 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 0.57 | 0.57 | 0.57 | 0.57 | 0.97 | 0.97 | 0.97 | 0.97 |
| Mean Y - Overall | 0.69 | 0.69 | 0.69 | 0.69 | 0.95 | 0.95 | 0.95 | 0.95 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. Column 1 reports estimates using simulated annual eligibility. Column 2 reports estimates using total simulated fixed eligibility (CPI). Column 3 reports estimates using total simulated fixed eligibility (RCPI). Column 4 reports estimates using total simulated fixed eligibility (WAGE). All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.21:
Effect of Race-Ethnicity-Specific Total Simulated Eligibili−
ty on Labor Supply of Parents with Non-White Children
Robustness to Simulated Eligibility Type

| | Maternal Labor Supply | | | | Paternal Labor Supply | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (1) | (2) | (3) | (4) |
| | Usual Hours Worked per Week | | | | | | | |
| SIMT | 0.66** | 0.83*** | 0.83*** | 0.79*** | 0.02 | 0.10 | 0.10 | 0.01 |
| | (0.26) | (0.23) | (0.23) | (0.24) | (0.24) | (0.22) | (0.22 | (0.24) |
| Observations | 511,813 | 511,813 | 511,813 | 511,813 | 355,534 | 355,534 | 355,534 | 355,534 |
| Adjusted $R^2$ | 0.06 | 0.06 | 0.06 | 0.06 | 0.03 | 0.03 | 0.03 | 0.03 |
| Mean Y - Baseline | 22.75 | 22.75 | 22.75 | 22.75 | 38.74 | 38.74 | 38.74 | 38.74 |
| Mean Y - Overall | 24.96 | 24.96 | 24.96 | 24.96 | 38.73 | 38.73 | 38.73 | 38.73 |
| | Weeks Worked per Year | | | | | | | |
| SIMT | 0.90*** | 1.12*** | 1.12*** | 1.06*** | 0.10 | 0.23 | 0.23 | 0.14 |
| | (0.31) | (0.27) | (0.27) | (0.29) | (0.32) | (0.31) | (0.31 | (0.31) |
| Observations | 511,813 | 511,813 | 511,813 | 511,813 | 355,534 | 355,534 | 355,534 | 355,534 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.09 | 0.09 | 0.04 | 0.04 | 0.04 | 0.04 |
| Mean Y - Baseline | 24.86 | 24.86 | 24.86 | 24.86 | 43.45 | 43.45 | 43.45 | 43.45 |
| Mean Y - Overall | 29.71 | 29.71 | 29.71 | 29.71 | 43.93 | 43.93 | 43.93 | 43.93 |
| | Labor Force Participation | | | | | | | |
| SIMT | 0.03*** | 0.03*** | 0.03*** | 0.03*** | 0.00 | 0.00 | 0.00 | 0.00 |
| | (0.01) | (0.01) | (0.01) | (0.01) | (0.00) | (0.00) | (0.00 | (0.00) |
| Observations | 491,785 | 491,785 | 491,785 | 491,785 | 334,394 | 334,394 | 334,394 | 334,394 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.07 | 0.07 | 0.04 | 0.04 | 0.04 | 0.04 |
| Mean Y - Baseline | 0.57 | 0.57 | 0.57 | 0.57 | 0.92 | 0.92 | 0.92 | 0.92 |
| Mean Y - Overall | 0.66 | 0.66 | 0.66 | 0.66 | 0.92 | 0.92 | 0.92 | 0.92 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, and labor force participation last week). Usual hours worked per week and weeks worked last year include zeros. Column 1 reports estimates using simulated annual eligibility. Column 2 reports estimates using total simulated fixed eligibility (CPI). Column 3 reports estimates using total simulated fixed eligibility (RCPI). Column 4 reports estimates using total simulated fixed eligibility (WAGE). All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.22:
Effect of Total Child Years of Simulated Eligibility on Geographic Mobility

| | All | White | Non-White |
|---|---|---|---|
| | Total Child Years of Simulated Eligibility up to Period *t* | | |
| SIMTCUM | 0.01 | 0.00 | 0.01 |
| | (0.01) | (0.01) | (0.02) |
| Observations | 7,165,465 | 4,991,673 | 2,173,792 |
| Adjusted $R^2$ | 0.06 | 0.05 | 0.07 |
| | Total Child Years of Simulated Eligibility up to Period *t-1* | | |
| SIMTCUN | 0.01 | 0.00 | 0.01 |
| | (0.01) | (0.01) | (0.02) |
| Observations | 7,165,138 | 4,991,673 | 2,173,465 |
| Adjusted $R^2$ | 0.05 | 0.05 | 0.07 |
| | Total Contemporaneous Simulated Eligibility | | |
| SIMTCON | 0.04 | 0.02 | 0.10 |
| | (0.10) | (0.10) | (0.13) |
| Observations | 7,156,103 | 4,987,829 | 2,168,274 |
| Adjusted $R^2$ | 0.06 | 0.05 | 0.07 |
| Mean Y - Baseline | 0.16 | 0.18 | 0.13 |
| Mean Y - Overall | 0.16 | 0.17 | 0.14 |

*Notes:* This table shows results from regressions estimating the effect of total child years of simulated eligibility (SIMTCUM), total child years of simulated eligibility up to period *t-1* (SIMTCUN), and total contemporaneous simulated eligibility (SIMTCON) on indicator equals to one if the child resides in a different state than birth state. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from 5% census sample 1990 and 2000 and ACS 2010. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.23:
Effect of Total Child Years of Simulated Eligibility on Parental Labor Supply
Robustness to Non-Weighted & Non-Balanced Simulated Eligibility

| | Maternal Labor Supply | | | Paternal Labor Supply | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Usual Hours Worked per Week | | | | | |
| SIMTCUM | 0.09* | -0.03 | 0.26*** | 0.15*** | 0.14*** | 0.18*** |
| | ( 0.05) | ( 0.06) | ( 0.06) | ( 0.04) | ( 0.04) | ( 0.05) |
| Observations | 5,837,237 | 4,029,464 | 1,807,773 | 4,879,639 | 3,629,847 | 1,249,792 |
| Adjusted $R^2$ | 0.03 | 0.03 | 0.03 | 0.05 | 0.02 | 0.02 |
| Mean Y - Baseline | 23.33 | 23.49 | 22.91 | 41.99 | 43.15 | 37.91 |
| Mean Y - Overall | 24.63 | 25.05 | 23.88 | 40.61 | 42.37 | 36.57 |
| | Labor Force Participation | | | | | |
| SIMTCUM | 0.003** | 0.000 | 0.006*** | 0.003*** | 0.002*** | 0.003*** |
| | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) |
| Observations | 5,837,237 | 4,029,464 | 1,807,773 | 4,879,639 | 3,629,847 | 1,249,792 |
| Adjusted $R^2$ | 0.02 | 0.02 | 0.02 | 0.04 | 0.02 | 0.04 |
| Mean Y - Baseline | 0.69 | 0.69 | 0.67 | 0.95 | 0.96 | 0.92 |
| Mean Y - Overall | 0.71 | 0.72 | 0.69 | 0.93 | 0.95 | 0.89 |

*Notes:* This table shows results from regressions estimating the effect of total child years of simulated eligibility on parental labor supply (usual hours worked per week last year and labor force participation last week). Usual hours worked per week last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from 5% census sample 1990 and 2000 and ACS 2010. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.24:
Effect of Total Child Years of Simulated Eligibility on Parental Labor Supply
Robustness to State of Residence Simulated Eligibility

| | Maternal Labor Supply | | | Paternal Labor Supply | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Usual Hours Worked per Week | | | | | |
| SIMTCUM | 0.01 | -0.14* | 0.22*** | 0.18*** | 0.17*** | 0.22*** |
| | ( 0.05) | ( 0.08) | ( 0.06) | ( 0.04) | ( 0.04) | ( 0.05) |
| Observations | 6,946,198 | 4,852,391 | 2,093,807 | 5,819,311 | 4,359,586 | 1,459,725 |
| Adjusted $R^2$ | 0.02 | 0.02 | 0.03 | 0.05 | 0.02 | 0.02 |
| Mean Y - Baseline | 23.43 | 23.53 | 23.17 | 42.23 | 43.32 | 38.19 |
| Mean Y - Overall | 24.62 | 24.90 | 24.10 | 40.87 | 42.56 | 36.88 |
| | Labor Force Participation | | | | | |
| SIMTCUM | 0.001 | -0.002 | 0.005*** | 0.003*** | 0.002*** | 0.004*** |
| | (0.001) | (0.002) | (0.001) | (0.001) | (0.001) | (0.001) |
| Observations | 6,946,198 | 4,852,391 | 2,093,807 | 5,819,311 | 4,359,586 | 1,459,725 |
| Adjusted $R^2$ | 0.02 | 0.02 | 0.02 | 0.04 | 0.02 | 0.03 |
| Mean Y - Baseline | 0.69 | 0.69 | 0.67 | 0.95 | 0.96 | 0.92 |
| Mean Y - Overall | 0.71 | 0.72 | 0.69 | 0.93 | 0.95 | 0.89 |

*Notes:* This table shows results from regressions estimating the effect of total child years of simulated eligibility on parental labor supply (usual hours worked per week last year and labor force participation last week). Usual hours worked per week last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from 5% census sample 1990 and 2000 and ACS 2010. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.25:
Effect of Total Child Years of Simulated Eligibility on Parental Labor Supply
Robustness to State of Birth Simulated Eligibility

| | Maternal Labor Supply | | | Paternal Labor Supply | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Usual Hours Worked per Week | | | | | |
| SIMTCUM | 0.01 | -0.14* | 0.22*** | 0.17*** | 0.15*** | 0.21*** |
| | ( 0.05) | ( 0.08) | ( 0.06) | ( 0.04) | ( 0.04) | ( 0.06) |
| Observations | 6,946,219 | 4,852,391 | 2,093,828 | 5,819,326 | 4,359,586 | 1,459,740 |
| Adjusted $R^2$ | 0.02 | 0.02 | 0.03 | 0.05 | 0.02 | 0.02 |
| Mean Y - Baseline | 23.43 | 23.53 | 23.17 | 42.23 | 43.32 | 38.19 |
| Mean Y - Overall | 24.62 | 24.90 | 24.10 | 40.87 | 42.56 | 36.88 |
| | Labor Force Participation | | | | | |
| SIMTCUM | 0.001 | -0.002 | 0.005*** | 0.003*** | 0.002*** | 0.004*** |
| | (0.001) | (0.002) | (0.001) | (0.001) | (0.001) | (0.001) |
| Observations | 6,946,219 | 4,852,391 | 2,093,828 | 5,819,326 | 4,359,586 | 1,459,740 |
| Adjusted $R^2$ | 0.02 | 0.02 | 0.02 | 0.04 | 0.02 | 0.03 |
| Mean Y - Baseline | 0.69 | 0.69 | 0.67 | 0.95 | 0.96 | 0.92 |
| Mean Y - Overall | 0.71 | 0.72 | 0.69 | 0.93 | 0.95 | 0.89 |

*Notes:* This table shows results from regressions estimating the effect of total child years of simulated eligibility on parental labor supply (usual hours worked per week last year and labor force participation last week). Usual hours worked per week last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from 5% census sample 1990 and 2000 and ACS 2010. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

# B Medicaid Eligibility

## B.1 Medicaid Legislation

This section describes the underlying legislative rules used to calculate Medicaid eligibility for children in the U.S. for the period 1979-2014. Medicaid eligibility is imputed using the calculator from Miller and Wherry (2019).[44] Eligibility calculations can be broadly categorized into two groups - before and after the Personal Responsibility and Work Opportunity Act (PRWORA). To determine Medicaid eligibility of children, rules under Aid to Families with Dependent Children (AFDC), state-optional programs (AFDC-Unemployed Parents (AFDC-UP), Ribicoff children, Medically Needy), Medicaid "Section 1931", State Children's Health Insurance Program (SCHIP), as well as federal and state Medicaid expansions are used.[45] In general eligibility is imputed using applicable rules based on the date of eligibility determination, the child's age, and the child's birthday, family structure, family income, and information on parental employment.

**Eligibility Calculations before PRWORA (1979-1996)**

***Eligibility under AFDC***

Historically, Medicaid eligibility was restricted to children in families receiving cash welfare benefits. To determine eligibility under AFDC, it is assumed that the child care dedication is not used by eligible families and that the parent has spent one month working. The first set of rules to assess whether the family is eligible for AFDC, are rules regarding family income and earned income disregards. Total family income is calculated by summing all sources of income except public assistance or welfare of each parent. To determine if any earned income disregards are applicable, months spent working are

[44] This section summarizes the most important steps to determine Medicaid eligibility. See appendix and calculator documentation in Miller and Wherry (2019) for more information about legislative information, data sources and methodology used to impute eligibility.

[45] Table B.1 shows the major mandatory and state optional legislations that affected Medicaid eligibility of children during the analysis period.

compared to number of months that disregards are allowed based on state rules. The applicable disregards are then calculated by using state rules. In order to be financially eligible for AFDC, the child's family has to satisfy three tests. First, the family must be eligible for a non-zero AFDC benefit amount based on state rules, monthly total family income, and family size. Second, total family income less applicable disregards must be below the state need's standard. Third, total family income must not exceed a given percentage of the state need's standard. Special rules are implemented for Connecticut and Minnesota. As a second set of rules, the family has to satisfy two family structure requirements. First, eligibility under AFDC requires the child to reside in a single-parent family. Second, the child has to be age 0-17 at date of eligibility determination and either a primary or subfamily member, but not a head or spouse of primary family or subfamily.

***Eligibility under AFDC-UP***

Prior to the federal mandate effective in October 1990, AFDC-UP - a state-optional program - extended eligibility to children in two-parent families where the primary earner was unemployed. To be classified as unemployed, the parent must work less than 100 hours per month. A child is assumed to be eligible under AFDC-UP if AFDC-UP program was effective in state and year, child's family is financially eligible for AFDC, maximum hours worked by any individual in the family do not exceed 1200 per year, and the child resides in a family with married parents.

***Eligibility under Medically Needy Program***

The medically needy program provides states the option to extend Medicaid eligibility to individuals with high medical expenses whose income exceeds the maximum income eligibility threshold, but who satisfy all other eligibility criteria for Medicaid. Income limits could be set no higher than 133% of the state's needs standard for AFDC. However families could use the medical expenditures to reduce the applicable income through spent-down provisions. Since there is no information about medical expenditures in the

CPS, the eligibility limits are set to the Medically Needy levels in states with this program as an approximation. A child is eligible under the medically needy program if the family's income except public assistance or welfare is below the applicable eligibility thresholds.

***Eligibility under the Ribicoff Children Program***

Under the Ribicoff Children Program, states are allowed to cover children who would qualify for cash welfare program given income criteria alone but who do not qualify based on family structure. Hence a child is eligible for Medicaid under the Ribicoff Children Program if the Ribicoff Children Program is present in state and year, child's family is income eligible for AFDC, and the child lives in a family with married parents. In addition, the child is eligible under the Ribicoff Children Program if federally-mandated expansions of Ribicoff Children Program are applicable (Deficit Reduction Act, 1984; Omnibus Budget Reconciliation Act, 1987).

***Poverty-related Eligibility***

Beginning in 1984 states were required or given the option to expand Medicaid eligibility for children living in families with incomes below the eligibility limit. To impute eligibility under federal and state expansions child care deduction is assumed to be zero. All sources of income except public assistance or welfare of each parent are summed up to calculated the total family income. A child is eligible for Medicaid if the total family income minus the work expense deduction used in the net income calculations under AFDC is less than the applicable federal or state eligibility level in the given state, year, and age.

**Eligibility Calculations after PRWORA (1997-2014)**

***Eligibility under Medicaid "Section 1931"***

Medicaid "Section 1931" requires states to provide Medicaid coverage to children in families who meet eligibility requirements under AFDC and AFDC-UP effective on July

16, 1996 in the state of residence. To impute eligibility under "Section 1931", it is assumed that child care deduction is not used by eligible families and that the parent has spent one month working. Eligibility of a child is then determined using the eligibility rules for AFDC programs in effect for the state in July 1996. The procedure to calculate eligibility under AFDC and AFDC-UP is explained above. This calculator does not incorporate state optional "Section 1931" rules under which states have the option to set income and asset standards differently from those in effect under state AFDC program on July 16, 1996. Since eligibility requirements are presumably less restrictive under all other eligibility pathways after the welfare reform, omitting optional "Section 1931" eligibility will not bias the eligibility estimates.

***Eligibility under Separate State Programs under SCHIP***

Balanced Budget Act of 1997 allowed states to create separate state programs. Eligibility under separate state programs is imputed assuming that child care deductions and child support income are zero. To obtain the total family income all sources of income except public assistance or welfare of each parent are summed up. A child is eligible if the total family income minus the state- and SCHIP-specific work expense deduction per worker is less than SCHIP eligibility limit in the given state, year, and age.

***Poverty-related and Targeted Medicaid Eligibility***

After PRWORA two different pathways can determine eligibility under expansion-related rules. The poverty-related pathway is defined by a series of federal and state Medicaid expansions which extended eligibility for certain ages and income levels. The second path is given by targeted Medicaid expansions embedded in the Balanced Budget Act of 1997 which allow states to expand the state Medicaid programs. Technically, different income disregards are applied for the two pathways. However the calculator uses poverty-related disregards for both pathways. To impute eligibility, child care deductions and child support income are assumed to be zero. Total family income is calculated by summing all

sources of income except public assistance or welfare of each parent. If the total family income minus the state- and Medicaid-specific work expense deduction per worker are less than the corresponding cutoffs in a given state, year and age, the child is assumed to be eligible.

## B.2 Eligibility Imputation

This section explains how Medicaid eligibility is imputed based on variables available in CPS ASEC. To impute Medicaid eligibility, I use a calculator that incorporates state and federal legislation based on rules for a given year, state of residence, age of the child, and family characteristics including family income and family structure.[46] A child is considered eligible for public health insurance if the child's family meets eligibility requirements for AFDC, one of the state-optional programs (AFDC-UP, Ribicoff children program, Medicaid's medically needy program), or federal and state-optional Medicaid expansions.

In survey year *t*, CPS ASEC provides information about income from calendar year *t-1*, family structure as of March of calendar year *t*, and age of the child as of March of calendar year *t*. Hence, to calculate eligibility in calendar year *t*, I use data on income from calendar year *t*, family structure from calendar year *t+1*, and adjust the age of the child accordingly. Depending on the birth month and calendar month of eligibility determination, some children are treated as if they were the same age and others are treated as if they were a year younger during the previous calendar year.[47] Child's age *a* is therefore defined as age at calendar month of eligibility determination. I construct two Medicaid eligibility measures - contemporaneous to outcome variables measured as of previous calendar year ("last year eligibility") and contemporaneous to outcome variables measured as of interview month ("last week eligibility"). To obtain Medicaid eligibility

[46] I use a calculator from Miller and Wherry (2019). The calculator allows to impute eligibility for children age 0-18 for years 1979-2014.

[47] I randomly assign birth month because eligibility is imputed using the calendar month of eligibility determination and CPS does not provide the birth month of an individual.

contemporaneous to outcome variables measured as of previous calendar year (e.g., insurance coverage, usual hours worked, weeks worked), I calculate eligibility during each month of the given year and use the average eligibility across all months in that year. Medicaid eligibility contemporaneous to outcome variables measured as of survey date (e.g., marital status, educational attainment) or last week (e.g., labor force participation) is obtained by calculating eligibility during March of the given year.

To obtain the correct family structure and measure of total family income according to rules determining Medicaid eligibility, I construct nuclear families within a household. A family unit is defined as a parent, spouse if present, and children. First spouses within a household are linked and then parents are linked to their children within a household. I obtain family income by combining parental income within the nuclear family. To determine eligibility of a child in calendar year *t*, I follow the legislative rules to calculate Medicaid eligibility and first divide family income (except applicable disregards) by the corresponding poverty guideline for the family size of the nuclear family, state *s*, and calendar year *t*.[48] I then compare this ratio to the eligibility limit for a child of age *a*, in state *s*, and calendar year *t*. Since the eligibility cutoffs depend on the age of the child, the number of children in the family that are eligible for Medicaid may vary for families with the same income and number of children, but with children of different ages.

To check how well the calculator estimates eligibility of children, I examine the percentage of non-eligible children reporting Medicaid coverage or living in families receiving welfare payments since these children should be eligible for Medicaid. For the period 1979-2014, 3.78% of children imputed to be not eligible, report coverage by Medicaid and 0.56% of children live in families where a parent reports receiving cash benefits under the AFDC program, although these children are not eligible for Medicaid based on the calculator.

[48] Poverty guidelines depend on family size, year, and state of residence. All states except Alaska and Hawaii share the same poverty guidelines.

## B.3 Classic Simulated Eligibility Measure

The simulated eligibility is constructed by using all children of each age in each calendar year across the full sample period. Using this national data set, I construct simulated eligibility measures which vary at the state, calendar year, age, and race-ethnicity of the child level.[49] Following Ham and Shore-Sheppard (2005b), the simulated eligibility is obtained by using all children in the national data set except from the state for which the simulated eligibility is calculated. Each state-year-age-race-ethnicity simulated eligibility measure is hence the fraction of children in the national data set in state *s* except children from state *s*, calendar year *t* when outcome of interest is measured, of age *a*, and race-ethnicity *r* who would be eligible for Medicaid given the rules in each state *s*, calendar year *t*, and age *a*. Formally, the simulated eligibility for a given state, calendar year, age, and race-ethnicity of the child is given by the following equation:

$$SIM_{star} = \frac{\sum_{i=1}^{k_{\tilde{s}}} w_{i\tilde{s}tar} \cdot e_{i\tilde{s}tar}}{\sum_{i=1}^{k_{\tilde{s}}} w_{i\tilde{s}tar}} \tag{2}$$

where $k_{\tilde{s}}$ is the number of children in the national data set excluding children from state *s*, of age *a*, race-ethnicity *r*, and in calendar year *t*. $e_{i\tilde{s}tar}$ and $w_{i\tilde{s}tar}$ are individual-level eligibility and CPS ASEC weight of child *i*, not residing in state *s*, in calendar year *t*, of age *a*, and race-ethnicity *r*. Each child-specific simulated eligibility measure (last year and last week simulated eligibility) is then merged to each child based on child's state of residence, calendar year when outcome of interest is measured, age, and race-ethnicity of the child.

[49] Race groups are defined as white non-Hispanic and non-white Hispanic. For simplicity I refer to race-ethnicity groups as white and non-whites.

## B.4 Family-level Simulated Eligibility Measure

To construct family's total simulated eligibility (last year and last week), I sum the simulated eligibility fractions $SIM_{star}$ (last year and last week, respectively) across all children in a family.[50] To facilitate notation, I define a parent type by the number of children of any age between 0 and 18 of that parent - $j := (n_{j_0}, ..., n_{j_{18}})$ where $n_{j_a}$ is the number of children of age *a* of parent *j*.[51] Formally, the total simulated eligibility of a parent type *j*, state *s*, calendar year *t*, and race-ethnicity group *r* is given by the following equation:

$$SIMT_{jstr} = SIMT_{(n_{j_0},...,n_{j_{18}})str} = \sum_{a=0}^{18} SIM_{star} * n_{j_a} \tag{3}$$

where $SIM_{star}$ is the simulated eligibility measure defined in equation 2 and $n_{j_a}$ is the number of children of age *a* of parent *j*. Hence the level of variation of total simulated eligibility within a given state, year, race-ethnicity group, and number of children is the distribution of possible combinations of child's ages. In general, parents in the same state, year, race-ethnicity group, and of the same type (same number of children and same age of each child) are characterized by the same total simulate eligibility measure.

## B.5 Long-Run Simulated Eligibility Measure

I construct the long-run simulated eligibility measure by first calculating the average total simulated eligibility in each state, year, age, and race-ethnicity using CPS ASEC. Second, I create the total child-years of simulated eligibility by summing the average total simulated eligibility at each age from birth to the current age of the child for a given birth cohort, age, state, and race-ethnicity group. To account for different number of children at each age that contribute to the average total simulated eligibility, the sum is created

[50] Consider for example a family with two children age 3 and 5. The child-specific simulated eligibility measure of the first and second child is 0.5 and 0.6 respectively. The family's total simulated eligibility is hence 1.1.

[51] Take for instance a parent with two one-year old, one three-year old, and two five-year old children. The parent type is hence given by the vector (0,2,0,1,0,2,0,...,0).

using relative weights for a given age. The relative weight is obtained by dividing the sum of children's CPS ASEC weights for a given age by the average of the sum of children's CPS ASEC weights across all ages up to the current age of the child. The weighting scheme aims to recover equal distribution of number of children at each year during childhood and allow the long-run eligibility measure to be representative of the average child. Moreover, since the Medicaid calculator starts in 1979 and the first calendar year in DCS begins in 1990, the main analysis uses a balanced long-run eligibility measure from birth to the eleventh birthday. Finally, the total child-years of simulated eligibility experienced by a parent are merged to the child observed in DCS and ACS based on birth year, age, state, and race-ethnicity.[52]

Consider for instance a white three-year old child observed in 2000 born and residing in Alabama. In addition, assume that the sum of CPS ASEC weights of zero-, one-, two-, and three-year old children in Alabama in 1997, 1998, 1999, and 2000 is 40, 30, 20, and 10 respectively. First, I calculate the average total simulated eligibility of zero-year old white children in 1997 in Alabama, one-year old white children in 1998 in Alabama, two-year old white children in 1999 in Alabama, and three-year old white children in 2000 in Alabama. Then, I sum the four average total simulated eligibility measures between age zero and age three using relative weights ($\frac{40}{25}$, $\frac{30}{25}$, $\frac{20}{25}$, and $\frac{10}{25}$ for age zero, one, two, and three respectively) and merge the total child-years of simulated eligibility to the three-year old white child observed in 2000 from Alabama.

[52] Since the long-run eligibility measure captures the average exposure to Medicaid during childhood, I need to control for the number of children a child is exposed to during childhood on average. The total child-years of number of children during childhood is constructed in the exact same way as the total child-years of simulated eligibility except using average number of children per family and not total simulated eligibility in each state, year, age, and race-ethnicity.

Table B.1:
Medicaid and CHIP Legislation Expanding Eligibility of Children 1979-2014

| Year | Legislation | Mandatory Expansion | State Option |
|---|---|---|---|
| 1984 | Deficit Reduction Act | Coverage of children under age 5 born after September 30, 1983 whose families are income and resource eligible for AFDC | |
| 1986 | Omnibus Budget Reconciliation Act | | Increase age level by 1 year each FY for all children under age 5 with incomes below 100% FPL. Infants in families with incomes below 100% FPL |
| 1987 | Omnibus Budget Reconciliation Act | Coverage of children under age 7 born after September 30, 1983 whose families are income and resource eligible for AFDC | Coverage of infants in families with incomes below 185% FPL and children under age 2, 3, 4, or 5 and born after September 30, 1983 in families with incomes below 100% FPL. Coverage of children under age 8 born after September 30, 1983 whose families are income and resource eligible for AFDC and children under age 8 born after September 30, 1983 with incomes below 100% FPL. |
| 1988 | Medicare Catastrophic Coverage Act | Coverage of infants in families with incomes below 75% FPL (1-Jul-89) and infants in families with incomes below 100% FPL (1-Jul-90) | Coverage of children up to eight years of age with family incomes below 75% FPL |
| 1988 | Family Support Act of 1988 | Extension to twelve months transitional Medicaid coverage to families leaving AFDC rolls due to earnings from work. Coverage of two-parent unemployed families meeting state AFDC income and resource standards. | |
| 1989 | Omnibus Budget Reconciliation Act | Coverage of children under age 6 with family incomes below 133% FPL | |
| 1990 | Omnibus Budget Reconciliation Act | Coverage of children under age 19 born after September 30, 1983 with incomes below 100% FPL. | |
| 1996 | Personal Responsibility and Work Opportunity Act | Coverage of families meeting AFDC eligibility standards as of July 16, 1996 (“Section 1931”) | Coverage of higher-income families. |
| 1997 | Balanced Budget Act | | Coverage of children under age 19 in families with incomes below 200% FPL or higher |

*Notes:* Buchmueller et al. (2016) and Miller and Wherry (2019)

# C Paternal Labor Market Outcomes

This section elaborates two potential explanations for the disproportionately large earnings effects of white fathers. First, the race-ethnicity-specific simulated eligibility may be correlated with unobservable characteristics that affect earnings of high-income white men. Second, men may be pushed into higher earnings occupations or working schedules as a result of children's access to Medicaid.

To explore the first point, I create an eligibility measure that aims at capturing differential eligibility across the earnings distribution. This eligibility measure is at the state, year, age, race-ethnicity, and parental education group level since education is positively correlated with income and children with parents in high-education group are less likely to be eligible.[53] Education groups are defined as high school and below or some college and above. Figure C.3 (C.4) shows the distribution of different simulated eligibility measures and actual eligibility across maternal (paternal) earnings.[54] As expected, the actual eligibility is decreasing to almost zero at earnings above the maximum eligibility level. The race-ethnicity-specific simulated eligibility, however, stays fairly constant across the distribution of parental earnings because a child from a low-income family and a child from a high-income family in the same state, year, age, and race-ethnicity group will be by definition assigned the same eligibility measure. The small increase in the race-ethnicity-specific simulated eligibility at higher levels of earnings is due to the higher share of high-income families and more generous Medicaid eligibility in later part of the analysis period. The education-specific simulated eligibility is lower than race-ethnicity-specific simulated eligibility, but does not trace the actual eligibility well at higher earnings implying that parental education does not perfectly capture high- and low-eligibility groups. To show that education is not a good predictor of eligibility, I create an alternative

[53] Since parental educational attainment could respond to children's Medicaid, using education may result in an endogenous eligibility measure. I discuss this point in more detail in section 5.5.2.

[54] I show the average eligibility to account for different number of children between low- and high-income families which would affect the total eligibility measures.

eligibility measure that assigns different simulated eligibility to children with parental earnings relative to federal poverty guidelines above and below the maximum eligibility limit across all states and years. The income-specific simulated eligibility measure at the state, year, age, race-ethnicity, and income group level supports this argument since it converges to actual eligibility around the maximum eligibility limit.[55] I refrain, however, from using the income-specific simulated eligibility for the analysis since an eligibility measure that incorporates a form of income is endogenous.

The estimated effects of education-specific eligibility measure (bottom panel of table C.1) show that the elasticity of the point estimate for earnings of white men is 59% lower than the elasticity of the point estimate from the model using the race-ethnicity-specific eligibility measure emphasizing the importance to correctly capture the simulated eligibility of high earners. Shown in figure C.6, the earnings of white fathers still respond to education-specific simulated eligibility in area outside of maximum eligibility limit consistent with education-specific eligibility measure not being a good predictor for eligibility of high-income individuals.[56]

Another potential explanation for the disproportionately large earnings effects of white men is a switch to better paying jobs or working schedules. More descriptively, I show that fathers with white children are two percentage points more likely to work in managerial and professional occupations as a response to expanding Medicaid eligibility.[57] This is equivalent to 6.1% relative to baseline mean. At the same time, extended eligibility results in reduced probability (4.8%-8.7%) of fathers with white children working

[55] For women the income-specific simulated eligibility measures do not reflect the actual eligibility across the earnings distribution because earnings of mothers relative to federal poverty guidelines almost never exceed the maximum Medicaid eligibility limit which represents the 97th percentile of maternal earnings (see figure C.1).

[56] Shown in table C.2 the effects of education-specific simulated eligibility on usual hours worked per week, weeks worked per year, and labor force participation are very similar to estimates from models using race-ethnicity-specific simulated eligibility.

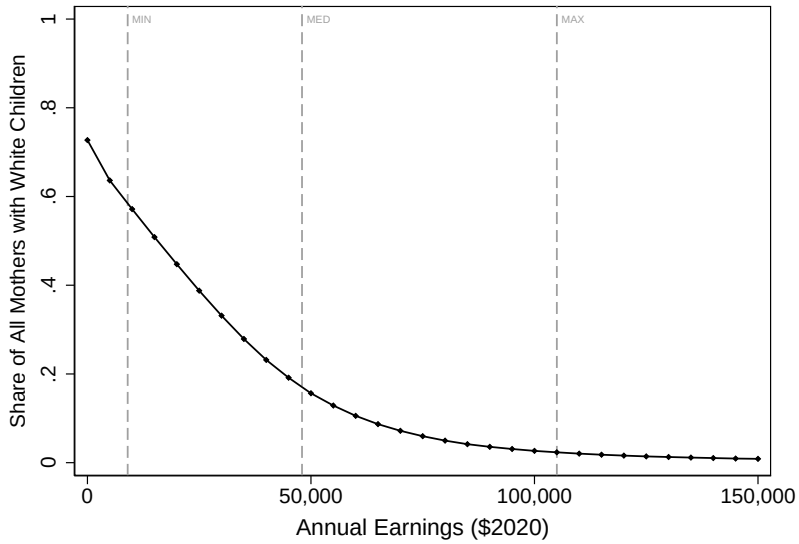
[57] To estimate the effect of simulated eligibility on parental occupational choice (see table C.4), the sample has to be restricted to parents in the labor force resulting in a selected sample. Hence the estimates should be viewed as descriptive. Appendix D provides a detailed description of occupational classification.

in manual occupations (farming, forestry, fishing; precision, craft, repair; operators, fabricators, laborers). Shown in table C.3, the occupational choice of white men translates into 3.4% and 6.3% higher probability of choosing occupations with wages above the 50th and 75th percentile, respectively. The transition from part-time to full-time employment (see table C.6) could also explain the relatively larger earnings effects because the majority of full-time workers earn more per hour than part-time workers.[58]

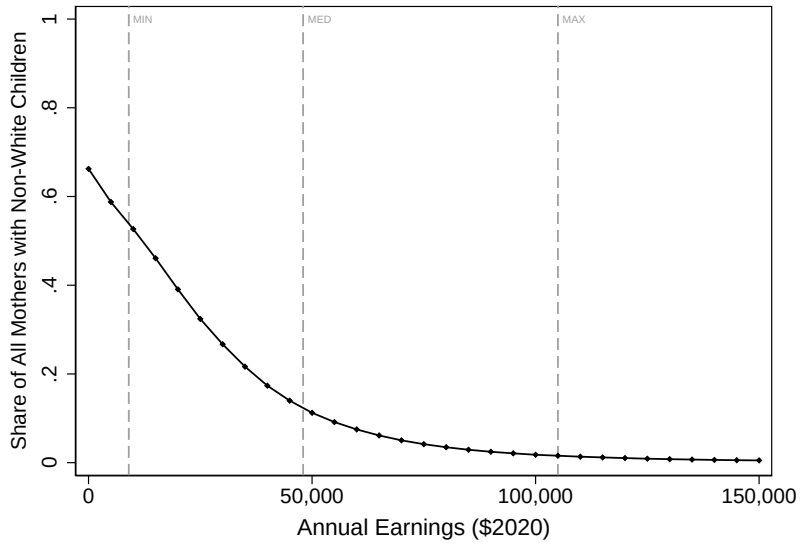
[58]The full-time wage premium has been well documented in the literature for male and female workers across various demographic groups. See for example Brinkley (1994); Lettau (1997); Lettau and Buchmueller (1999); Aaronson and French (2004); Hirsch (2005); Pongrace and Zilberman (2009); Borowczyk-Martins and Lalé (2017); Nightingale (2019).

Figure C.1:
Distribution of Maternal Annual Earnings ($2020)

(a) Mothers with White Children



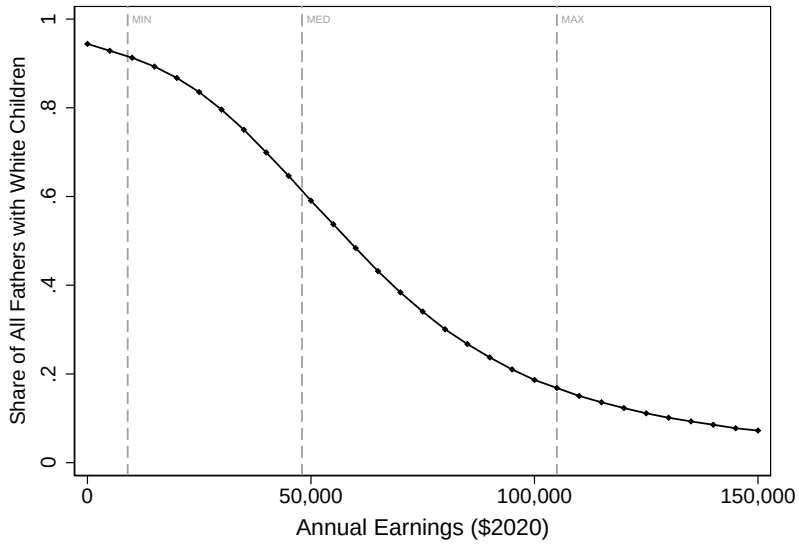

(b) Mothers with Non-White Children



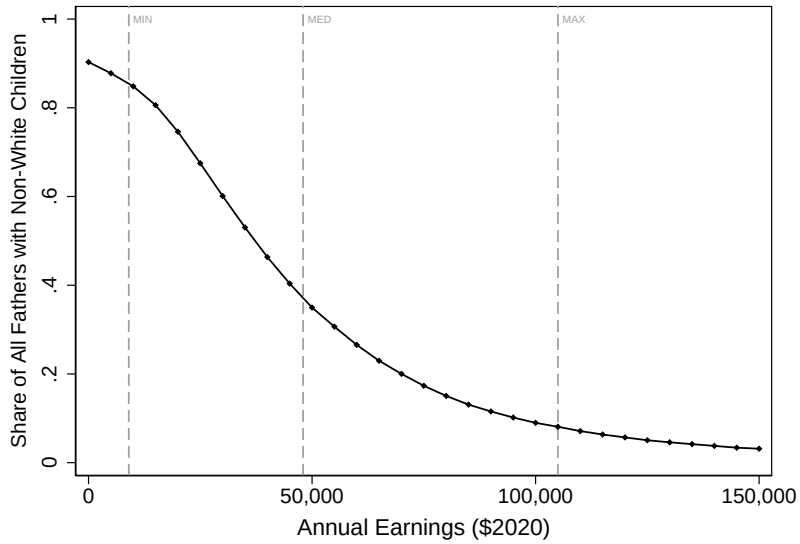

*Notes:* These figures show the truncated distribution of annual earnings ($2020) last year excluding zeros of mothers with (a) white and (b) non-white children.. The first, second, and third dashed vertical line represent average minimum, median, and maximum Medicaid eligibility limits during the analysis period, respectively. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure C.2:
Distribution of Paternal Annual Earnings ($2020)

(a) Fathers with White Children




(b) Fathers with Non-White Children

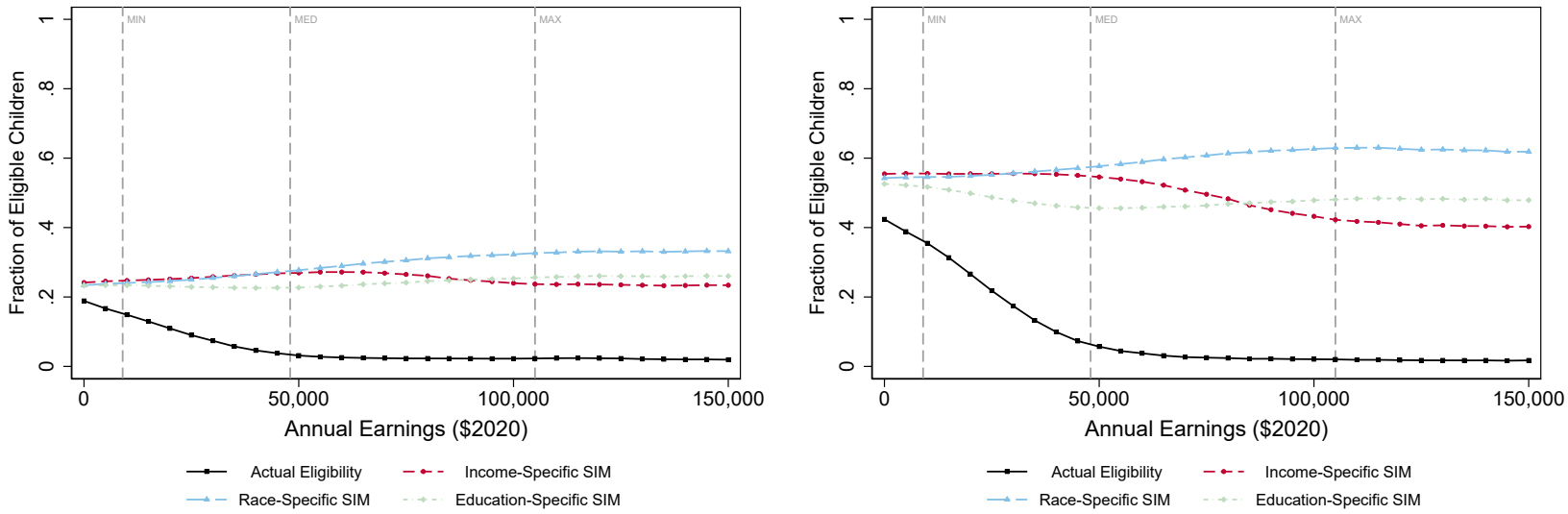



*Notes:* These figures show the truncated distribution of paternal annual earnings ($2020) last year excluding zeros of fathers with (a) white and (b) non-white children. The first, second, and third dashed vertical line represent average minimum, median, and maximum Medicaid eligibility limits during the analysis period, respectively. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure C.3:
Average Eligibility Measures across the Distribution of Maternal Annual Earnings ($2020)

(a) Mothers with White Children

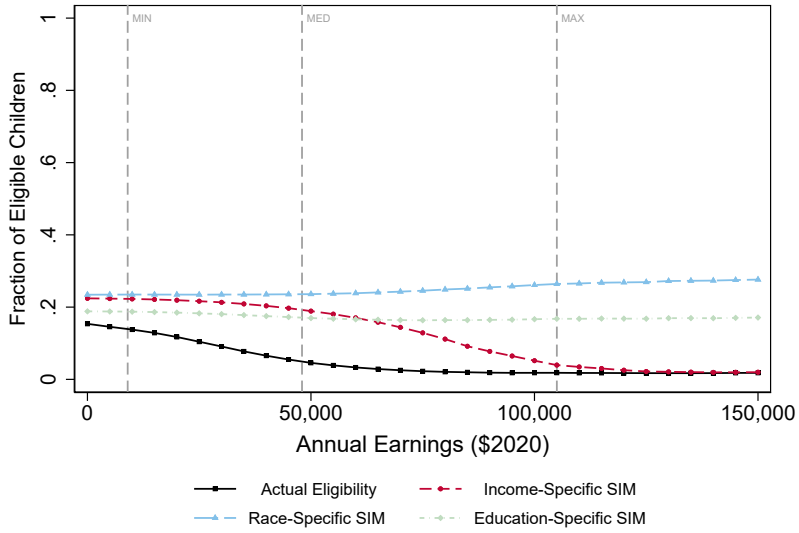
(b) Mothers with Non-White Children



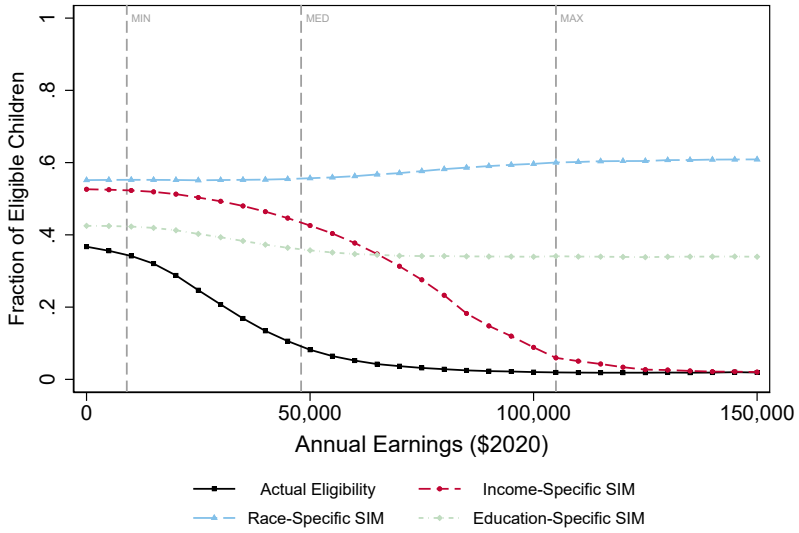

*Notes:* These figures show average eligibility measures across the truncated distribution of annual earnings ($2020) last year excluding zeros of mothers with (a) white and (b) non-white children. The first, second, and third dashed vertical line represent average minimum, median, and maximum Medicaid eligibility limits during the analysis period, respectively. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure C.4:
Average Eligibility Measures across the Distribution of Paternal Annual Earnings ($2020)

(a) Fathers with White Children



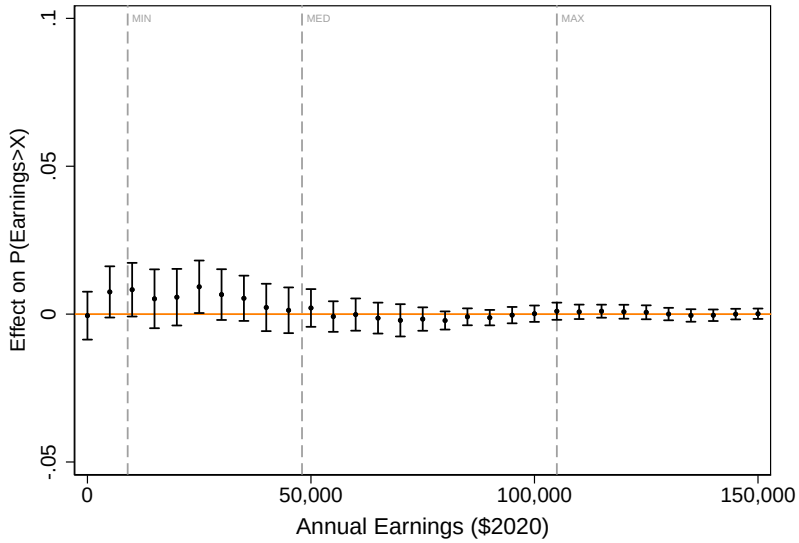

(b) Fathers with Non-White Children



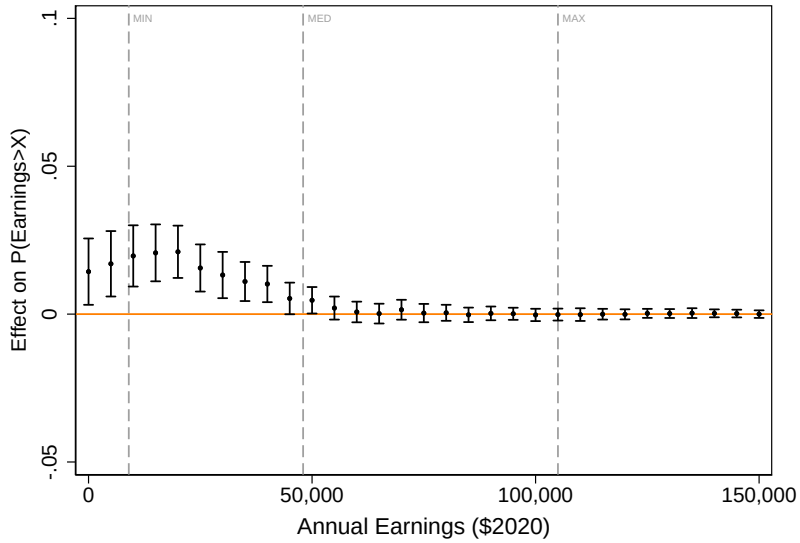

*Notes:* These figures show average eligibility measures across the truncated distribution of annual earnings ($2020) last year excluding zeros of fathers with (a) white and (b) non-white children. The first, second, and third dashed vertical line represent average minimum, median, and maximum Medicaid eligibility limits during the analysis period, respectively. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure C.5:
Effect of Education-Specific Total Simulated Eligibility on Maternal Annual Earnings ($2020)

(a) Mothers with White Children



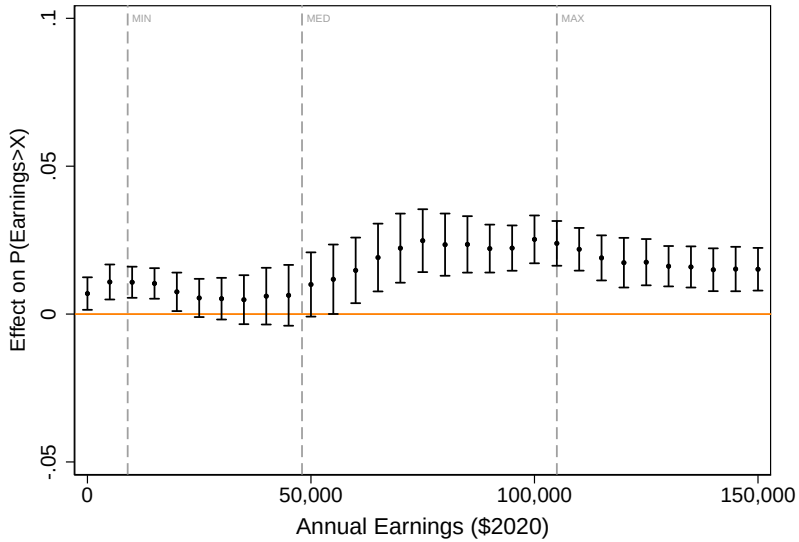

(b) Mothers with Non-White Children



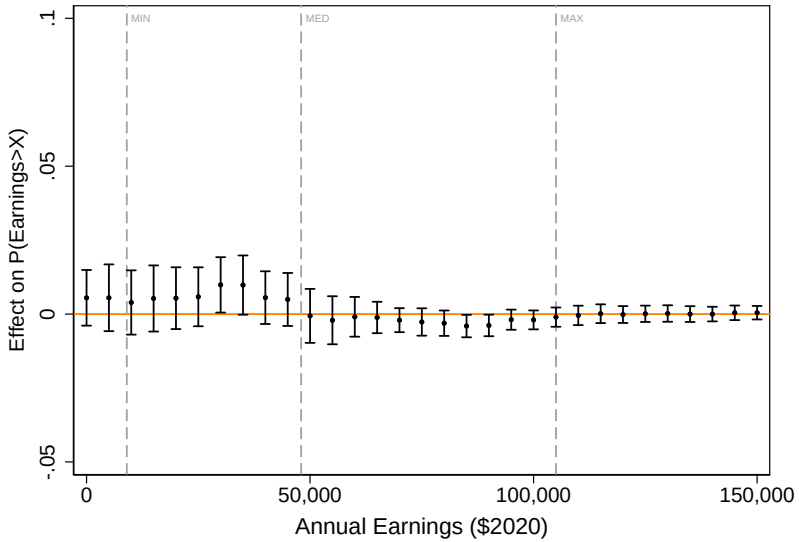

*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of education-specific total simulated eligibility on annual earnings ($2020) last year of mothers with (a) white and (b) non-white children. Each point estimate and confidence interval is obtained from a different regression where the dependent variable is an indicator equals to one if maternal annual earnings ($2020) were at least as great as X (0,5000,...,150000) last year. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors are clustered at the state level. The first, second, and third dashed vertical line represent average minimum, median, and maximum Medicaid eligibility limits during the analysis period, respectively. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Figure C.6:
Effect of Education-Specific Total Simulated Eligibility on Paternal Annual Earnings ($2020)

(a) Fathers with White Children




(b) Fathers with Non-White Children




*Notes:* These figures show the coefficients and 95% confidence intervals from regressions estimating the effect of education-specific total simulated eligibility on annual earnings ($2020) last year of fathers with (a) white and (b) non-white children. Each point estimate and confidence interval is obtained from a different regression where the dependent variable is an indicator equals to one if paternal annual earnings ($2020) were greater as X (0,5000,...,150000) last year. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). Regressions are weighted with parental survey weights divided by number of children per family. Standard errors are clustered at the state level. The first, second, and third dashed vertical line represent average minimum, median, and maximum Medicaid eligibility limits during the analysis period, respectively. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64.

Table C.1:
Effect of Total Simulated Eligibility on Parental Annual Earnings

| | Maternal Annual Earnings | | | Paternal Annual Earnings | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Race-Ethnicity-Specific Simulated Eligibility | | | | | |
| | Annual Total Earnings ($2020) | | | | | |
| SIMT | -269 | -294 | -236 | 5,951*** | 10,318*** | -1,950* |
| | ( 435) | ( 554) | ( 566) | ( 1,340) | ( 1,642) | ( 974) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.08 | 0.08 | 0.07 | 0.10 | 0.08 | 0.05 |
| Mean Y - Baseline | 14,826 | 14,674 | 15,279 | 62,535 | 66,614 | 46,665 |
| Mean Y - Overall | 24,306 | 25,932 | 21,485 | 65,210 | 72,342 | 49,282 |
| | Annual Wage Earnings ($2020) | | | | | |
| SIMT | -377 | -472 | -253 | 5,975*** | 10,413*** | -2,055** |
| | ( 389) | ( 496) | ( 557) | ( 1,376) | ( 1,727) | ( 939) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.07 | 0.07 | 0.06 | 0.08 | 0.07 | 0.05 |
| Mean Y - Baseline | 14,315 | 14,080 | 15,015 | 56,118 | 59,350 | 43,539 |
| Mean Y - Overall | 23,345 | 24,783 | 20,849 | 60,352 | 66,660 | 46,265 |
| | Education-Specific Simulated Eligibility | | | | | |
| | Annual Total Earnings ($2020) | | | | | |
| SIMT | 446** | 221 | 748** | 3,920*** | 6,458*** | 450 |
| | ( 203) | ( 327) | ( 303) | ( 757) | ( 964) | ( 430) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.12 | 0.11 | 0.14 | 0.16 | 0.14 | 0.14 |
| Mean Y - Baseline | 14,826 | 14,674 | 15,279 | 62,535 | 66,614 | 46,665 |
| Mean Y - Overall | 24,306 | 25,932 | 21,485 | 65,210 | 72,342 | 49,282 |
| | Annual Wage Earnings ($2020) | | | | | |
| SIMT | 335* | 48 | 719** | 3,773*** | 6,375*** | 215 |
| | ( 178) | ( 299) | ( 297) | ( 798) | ( 1,057) | ( 458) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.12 | 0.11 | 0.14 | 0.15 | 0.13 | 0.13 |
| Mean Y - Baseline | 14,315 | 14,080 | 15,015 | 56,118 | 59,350 | 43,539 |
| Mean Y - Overall | 23,345 | 24,783 | 20,849 | 60,352 | 66,660 | 46,265 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on parental annual total earnings ($2020) last year. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). The controls in models using full sample and race-specific or education-specific simulated eligibility are interacted with race or race and high school or less indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.2:
Effect of Education-Specific Total Simulated Eligibility on Parental Labor Supply

| | Maternal Labor Supply | | | Paternal Labor Supply | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Usual Hours Worked per Week | | | | | |
| SIMT | 0.36** | 0.15 | 0.65*** | 0.49*** | 0.60*** | 0.34* |
| | (0.14) | (0.20) | (0.20) | (0.13) | (0.16) | (0.19) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.10 | 0.06 | 0.04 | 0.05 |
| Mean Y - Baseline | 21.89 | 21.61 | 22.75 | 42.07 | 42.93 | 38.74 |
| Mean Y - Overall | 25.56 | 25.91 | 24.96 | 41.63 | 42.93 | 38.73 |
| | Weeks Worked per Year | | | | | |
| SIMT | 0.41*** | 0.12 | 0.81*** | 0.59*** | 0.64*** | 0.52** |
| | (0.13) | (0.20) | (0.25) | (0.17) | (0.18) | (0.26) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.11 | 0.10 | 0.12 | 0.06 | 0.05 | 0.05 |
| Mean Y - Baseline | 25.41 | 25.60 | 24.86 | 46.53 | 47.32 | 43.45 |
| Mean Y - Overall | 31.44 | 32.43 | 29.71 | 46.08 | 47.04 | 43.93 |
| | Labor Force Participation | | | | | |
| SIMT | 0.01*** | 0.01 | 0.02*** | 0.01*** | 0.01** | 0.01** |
| | (0.00) | (0.00) | (0.01) | (0.00) | (0.00) | (0.00) |
| Observations | 1,374,344 | 863,041 | 511,303 | 1,093,792 | 745,655 | 348,137 |
| Adjusted $R^2$ | 0.09 | 0.09 | 0.09 | 0.05 | 0.04 | 0.05 |
| Mean Y - Baseline | 0.57 | 0.57 | 0.57 | 0.96 | 0.97 | 0.92 |
| Mean Y - Overall | 0.68 | 0.69 | 0.66 | 0.94 | 0.95 | 0.92 |

*Notes:* This table shows results from regressions estimating the effect of education-specific total simulated eligibility on parental labor supply (usual hours worked per week last year, weeks worked last year, usual hours worked last year, and labor force participation last week). Usual hours worked per week, weeks worked last year, and usual hours worked last year include zeros. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.3:
Effect of Race-Ethnicity-Specific Total Simulated Eligibili−
ty on Parental Occupational Choice by Wage Distribution

| | Maternal Occupation | | | Paternal Occupation | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Occupation with Wages about 25th Percentile | | | | | |
| SIMT | 0.011** | 0.009 | 0.015** | -0.002 | 0.007 | -0.018*** |
| | (0.004) | (0.006) | (0.006) | (0.005) | (0.006) | (0.006) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.078 | 0.062 | 0.050 | 0.066 | 0.031 | 0.032 |
| Mean Y - Baseline | 0.361 | 0.386 | 0.285 | 0.786 | 0.819 | 0.657 |
| Mean Y - Overall | 0.449 | 0.505 | 0.353 | 0.766 | 0.820 | 0.646 |
| | Occupation with Wages about 50th Percentile | | | | | |
| SIMT | -0.002 | -0.001 | -0.004 | -0.001 | 0.016** | -0.033*** |
| | (0.005) | (0.006) | (0.006) | (0.005) | (0.007) | (0.006) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.076 | 0.066 | 0.049 | 0.085 | 0.049 | 0.044 |
| Mean Y - Baseline | 0.157 | 0.174 | 0.106 | 0.534 | 0.583 | 0.344 |
| Mean Y - Overall | 0.269 | 0.311 | 0.196 | 0.506 | 0.573 | 0.356 |
| | Occupation with Wages about 75th Percentile | | | | | |
| SIMT | -0.004 | 0.003 | -0.012** | 0.007* | 0.022*** | -0.021*** |
| | (0.003) | (0.004) | (0.005) | (0.004) | (0.006) | (0.006) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.045 | 0.039 | 0.036 | 0.070 | 0.052 | 0.049 |
| Mean Y - Baseline | 0.077 | 0.085 | 0.052 | 0.283 | 0.316 | 0.158 |
| Mean Y - Overall | 0.139 | 0.161 | 0.099 | 0.289 | 0.333 | 0.192 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on the probability of parents reporting being last year in a one-digit occupation with average wage above 25, 50, and 75 percentile. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.4:
Effect of Race-Ethnicity-Specific Total Simula−
ted Eligibility on Parental Occupational Choice

| | Maternal Occupations | | | Paternal Occupations | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Managerial, Professional | | | | | |
| SIMT | 0.010* | 0.021*** | -0.007 | 0.004 | 0.019*** | -0.024*** |
| | (0.005) | (0.007) | (0.009) | (0.005) | (0.005) | (0.007) |
| Observations | 950,688 | 623,757 | 326,931 | 1,038,667 | 717,856 | 320,811 |
| Adjusted $R^2$ | 0.093 | 0.088 | 0.061 | 0.082 | 0.067 | 0.061 |
| Mean Y - Baseline | 0.202 | 0.223 | 0.138 | 0.296 | 0.327 | 0.167 |
| Mean Y - Overall | 0.301 | 0.341 | 0.226 | 0.299 | 0.339 | 0.206 |
| | Technical, Sales, Administrative | | | | | |
| SIMT | 0.014** | 0.001 | 0.035*** | 0.004 | 0.004 | 0.005 |
| | (0.006) | (0.007) | (0.009) | (0.003) | (0.004) | (0.005) |
| Observations | 950,688 | 623,757 | 326,931 | 1,038,667 | 717,856 | 320,811 |
| Adjusted $R^2$ | 0.020 | 0.017 | 0.021 | 0.006 | 0.005 | 0.007 |
| Mean Y - Baseline | 0.430 | 0.456 | 0.351 | 0.134 | 0.141 | 0.107 |
| Mean Y - Overall | 0.389 | 0.405 | 0.360 | 0.175 | 0.183 | 0.155 |
| | Farming, Forestry, Fishing | | | | | |
| SIMT | -0.003** | -0.004*** | -0.001 | -0.012*** | -0.017*** | -0.004 |
| | (0.001) | (0.001) | (0.002) | (0.003) | (0.004) | (0.005) |
| Observations | 950,688 | 623,757 | 326,931 | 1,038,667 | 717,856 | 320,811 |
| Adjusted $R^2$ | 0.002 | 0.002 | 0.003 | 0.019 | 0.020 | 0.015 |
| Mean Y - Baseline | 0.016 | 0.014 | 0.020 | 0.231 | 0.233 | 0.225 |
| Mean Y - Overall | 0.020 | 0.018 | 0.023 | 0.211 | 0.216 | 0.200 |

continued on next page

Table C.5:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Parental Occupational Choice (continued)

| | Maternal Occupations | | | Paternal Occupations | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Service | | | | | |
| SIMT | -0.014*** | -0.013* | -0.017** | 0.008*** | 0.004** | 0.016*** |
| | (0.005) | (0.007) | (0.008) | (0.002) | (0.002) | (0.005) |
| Observations | 950,688 | 623,757 | 326,931 | 1,038,667 | 717,856 | 320,811 |
| Adjusted $R^2$ | 0.039 | 0.032 | 0.017 | 0.026 | 0.008 | 0.012 |
| Mean Y - Baseline | 0.198 | 0.175 | 0.269 | 0.067 | 0.056 | 0.113 |
| Mean Y - Overall | 0.200 | 0.166 | 0.265 | 0.096 | 0.072 | 0.154 |
| | Precision Production, Craft, Repair | | | | | |
| SIMT | -0.003** | -0.004*** | -0.001 | -0.012*** | -0.017*** | -0.004 |
| | (0.001) | (0.001) | (0.002) | (0.003) | (0.004) | (0.005) |
| Observations | 950,688 | 623,757 | 326,931 | 1,038,667 | 717,856 | 320,811 |
| Adjusted $R^2$ | 0.002 | 0.002 | 0.003 | 0.019 | 0.020 | 0.015 |
| Mean Y - Baseline | 0.016 | 0.014 | 0.020 | 0.231 | 0.233 | 0.225 |
| Mean Y - Overall | 0.020 | 0.018 | 0.023 | 0.211 | 0.216 | 0.200 |
| | Operators, Fabricators, Laborers | | | | | |
| SIMT | -0.006 | -0.004 | -0.008 | -0.001 | -0.008* | 0.011 |
| | (0.004) | (0.003) | (0.007) | (0.004) | (0.004) | (0.007) |
| Observations | 950,688 | 623,757 | 326,931 | 1,038,667 | 717,856 | 320,811 |
| Adjusted $R^2$ | 0.042 | 0.031 | 0.037 | 0.045 | 0.039 | 0.026 |
| Mean Y - Baseline | 0.138 | 0.116 | 0.205 | 0.237 | 0.209 | 0.351 |
| Mean Y - Overall | 0.081 | 0.063 | 0.115 | 0.194 | 0.167 | 0.257 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on the probability of parents reporting being in a one-digit occupation last year. All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table C.6:
Effect of Race-Ethnicity-Specific Total Simulated Eligibility on Parental Usual Hours Worked per Week

| | Maternal Labor Supply | | | Paternal Labor Supply | | |
|---|---|---|---|---|---|---|
| | All | White | Non-White | All | White | Non-White |
| | Positive Hours | | | | | |
| SIMT | 0.004 | -0.004 | 0.015** | 0.005* | 0.008** | 0.000 |
| | (0.004) | (0.005) | (0.007) | (0.003) | (0.003) | (0.005) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.06 | 0.06 | 0.06 | 0.03 | 0.02 | 0.03 |
| Mean Y - Baseline | 0.65 | 0.65 | 0.62 | 0.94 | 0.95 | 0.91 |
| Mean Y - Overall | 0.72 | 0.74 | 0.67 | 0.94 | 0.95 | 0.91 |
| | Part-Time Employment | | | | | |
| SIMT | -0.003 | -0.002 | -0.005 | -0.001 | -0.004** | 0.002 |
| | (0.004) | (0.005) | (0.004) | (0.002) | (0.002) | (0.003) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.03 | 0.02 | 0.01 | 0.01 | 0.01 | 0.01 |
| Mean Y - Baseline | 0.21 | 0.24 | 0.13 | 0.02 | 0.02 | 0.03 |
| Mean Y - Overall | 0.20 | 0.24 | 0.15 | 0.04 | 0.03 | 0.05 |
| | Full-Time Employment | | | | | |
| SIMT | 0.007 | -0.003 | 0.021*** | 0.007* | 0.012*** | -0.002 |
| | (0.005) | (0.008) | (0.006) | (0.004) | (0.004) | (0.006) |
| Observations | 1,375,551 | 863,738 | 511,813 | 1,117,645 | 762,111 | 355,534 |
| Adjusted $R^2$ | 0.06 | 0.07 | 0.05 | 0.03 | 0.02 | 0.03 |
| Mean Y - Baseline | 0.43 | 0.41 | 0.49 | 0.92 | 0.93 | 0.88 |
| Mean Y - Overall | 0.51 | 0.50 | 0.53 | 0.90 | 0.92 | 0.86 |

*Notes:* This table shows results from regressions estimating the effect of race-specific total simulated eligibility on likelihood of parents working any hours last year, working part time last year (>0 and <35 hours per week), and working full time last year ($\geq$ 35 hours per week). All regressions include child-level controls (indicators for sex, race and ethnicity, age, state of residence, calendar year), parental-level controls (indicators for parental age, age of the youngest, age of the oldest child, and number of children in the family), and state-level controls (unemployment rate, minimum wage, inflation-adjusted maximum welfare benefit for a family of 4, state-level EITC, implementation of six types of welfare waivers, implementation of any waiver or TANF). In models using the full sample all controls are interacted with a race indicator. Regressions are weighted with parental survey weights divided by number of children per family. Standard errors in parentheses are clustered at the state level. The data is from CPS ASEC 1980-2015. The sample is restricted to children age 0-18 with parents age 20-64. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

# D Occupation Classification

This section describes the occupational coding scheme. Following Lehn et al. (forthcoming), I use a modified version of the 1990 Census Bureau occupational classification created by Autor and Dorn (2013). This coding scheme offers a consistent and balanced panel of occupations. The categories of one-, two- and three-digit occupations are shown in table D.1.

Table D.1:
Occupational Classification

| One-Digit Occupation | Two-Digit Occupation | Occupational Codes |
|---|---|---|
| Managerial and Professional Specialty Occupations | Executive, Administrative, and Managerial Occupations | 3-22 |
| | Management Related Occupations | 23-37 |
| | Professional Specialty Occupations | 43-199 |
| Technical, Sales and Administrative Support Occupations | Technicians and Related Support Occupations | 203-235 |
| | Sales Occupations | 243-283 |
| | Administrative Support Occupations | 303-389 |
| Service Occupations | Private Household Occupations | 405-408 |
| | Protective Service Occupations | 415-427 |
| | Other Service Occupations | 433-472 |
| Farming, Forestry and Fishing Occupations | Farm Operators and Managers | 473-475 |
| | Other Agricultural and Related Occupations | 479-498 |
| Precision Production, Craft and Repair Occupations | Mechanics and Repairers | 503-549 |
| | Construction Trades | 558-599 |
| | Extractive Occupations | 614-617 |
| | Precision Production Occupations | 628-699 |
| Operators, Fabricators and Laborers | Machine Operators, Assemblers, and Inspectors | 703-799 |
| | Transportation and Material Moving Occupations | 803-889 |