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STATE EARNED INCOME TAX CREDITS AND THE PRODUCTION OF CHILD HEALTH: INSURANCE COVERAGE, UTILIZATION, AND HEALTH STATUS

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The Earned Income Tax Credit (EITC) has been credited with reductions in poverty and increases in the labor force participation of single mothers. The credit has the potential to affect the health of children in recipient families through three channels: family income, maternal employment, and health insurance coverage patterns. We exploit variation in state-level EITCs to estimate the effects of the credit on health insurance coverage, utilization of medical care, and health status. We find that the EITC is associated with significant changes in health insurance coverage patterns for children age 6–14, increasing rates of private health insurance but producing offsetting decreases in public coverage. State EITCs are also associated with significant improvements in health status for older children, an effect consistent with higher family income.

Keywords: EITC, child health

JEL Codes: I12, I13, I38, H31

I. INTRODUCTION

The Earned Income Tax Credit (EITC) has been the nation's largest federal cash transfer program for the poor for almost two decades, providing earnings subsidies in the form of refundable tax credits to working families earning below $52,427 per year in 2014. As of 2013, 25 states had also created their own EITCs that supplement the federal EITC. A primary goal of the EITC is to raise the income of low-wage workers, and it is credited with lifting as many as 3.2 million children out of poverty in 2013 (Center on Budget and Policy Priorities (CBPP), 2015). As both the value of the credit and the income range for eligibility have increased over the last 20 years, a substantial literature has emerged on income, anti-poverty, and labor supply effects of the program. More recently, new research has emerged focusing on how the program affects measures of

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well-being beyond income (Dahl and Lochner, 2008; Schmeiser, 2009; Strully, Rehkopf, and Xuan, 2010; Evans and Garthwaite, 2014; Hoynes, Miller, and Simon, 2015).[1]

At the same time, the health of children in the United States has received a great deal of attention in both the research literature and in the formation of policy. Until the recent passage of the Affordable Care Act (ACA) in 2010, most of the expansions of public health insurance programs since the late 1980s have been aimed at children. Additionally, policymakers and economists have historically been concerned with a number of specific indicators of child health, including infant mortality, self-reported health, missed days of school, and body weight (Currie, Decker and Lin, 2008; Currie and Gruber, 1995; De la Mata, 2012; Howell and Kenney, 2012). Further motivating the interest in the determinants of child health is a body of research that suggests that health in childhood has long-lasting impacts on a variety of outcomes in adulthood (Whitaker et al., 1997; Case, Fertig, and Paxson, 2005; Smith, 2009; Biro and Wien, 2010).

The way in which the EITC could affect childhood health outcomes is complex because there are multiple potential mechanisms at play. The EITC has been shown to significantly increase family income in both the short (Neumark and Wascher, 2001) and longer (Dahl, DeLeire, and Schwabish, 2009) terms; there is also substantial evidence in the research literature of a positive relationship between income and both child health (Case, Lubotsky, and Paxson, 2002) and child development (Yeung, Linver, and Brooks-Gunn, 2002; Currie, 2009). The EITC has also been shown to significantly increase labor force participation, particularly for single mothers (Eissa and Liebman, 1996; Ellwood, 2000; Meyer and Rosenbaum, 2000; Grogger, 2003; Cancian and Levinson, 2006; Hotz and Scholz, 2006). Studies on maternal employment and child health provide very mixed evidence; some estimates indicate that child health status is lower when a mother works more hours (Anderson, Butcher, and Levine, 2003; Gordon, Kaestner, and Korenman, 2007; Ruhm, 2008; Gennetian et al., 2010; Morrill, 2011). Whether maternal labor force participation is positively or negatively correlated with child health may depend upon a number of factors, including the family's income, the child's age, and the specific measure of health status. Finally, through both income and employment effects, the EITC may also change health insurance coverage patterns (Baughman, 2005).

In this study, we estimate the effect of the credit on factors that may be a part of the production function for child health, including health insurance coverage, utilization of medical care, and maternal employment. We also estimate the reduced form relationship between the EITC and health status. The small body of existing research on the impact of the EITC on health indicates positive effects of the credit on both maternal health (Evans and Garthwaite, 2014) and birth outcomes (Strully, Rehkopf, and Xuan, 2010; Hoynes, Miller, and Simon, 2015). This paper adds to the literature by expanding both the set of outcomes and the range of ages of children studied.

The Child and Young Adult Supplement to the 1979 National Longitudinal Survey of Youth provides a rich set of data on a variety of health-related outcomes for children,

[1] There is one other related study of note. Milligan and Stabile (2009) examine the effect of a similar income-tested child benefit in Canada and find improvements in maternal health, child behavioral and developmental outcomes, as well as parent-reported health among girls.

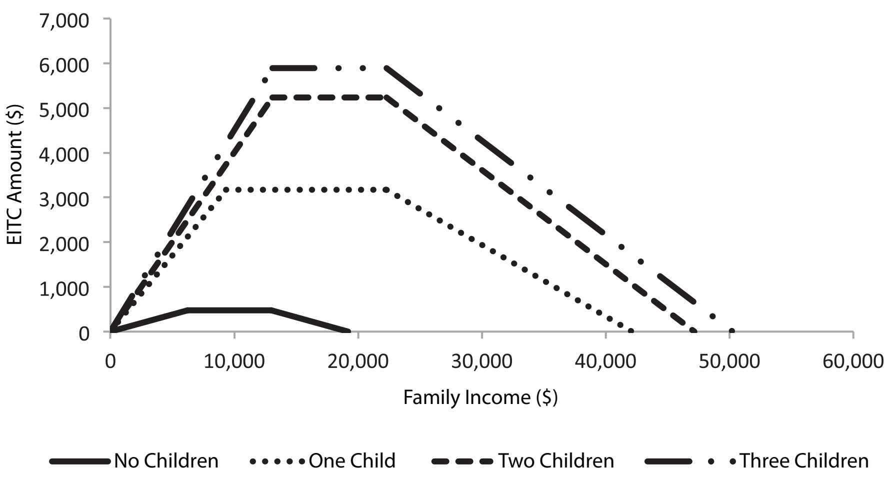
including source of health insurance coverage, utilization of routine medical and dental care, and several indicators of health status. The adoption and expansion of state-level EITC supplements in 23 U.S. states during the 1992–2006 period of our study provide the exogenous variation in policy that we use to identify our empirical models. We do not find any evidence that implementation and expansion of state EITC impacted any health-related outcomes for children under age 6. Among children age 6 to 14, the EITC was associated with an increase in private health insurance coverage and an offsetting decrease in public health insurance coverage, resulting in no significant change in the fraction of children who are uninsured. However, state EITCs were associated with improvements in the health status of children reported by mothers, a result that is most likely consistent with an income (rather than an employment or insurance) effect of the credit.

II. DESIGN OF THE EITC

The EITC is designed to increase family incomes and to promote labor force participation of low-income workers by providing wage subsidies to recipients. In 2012, it was available to workers with incomes of up to $19,190 (no children), $42,130 (one child), $47,162 (two children), or $50,270 (three or more children). (All values are for married taxpayers filing jointly.) Figure 1 shows the relationship between income

Figure 1

Design of the Federal EITC, 2012



Note: All values are for single/head of household filers.
Source: IRS, 2012 Earned Income Credit Table

and the EITC: the credit is structured such that recipients' earnings are matched by the credit at a phase in rate (ranging from 7.65 percent for workers without children to 45 percent for workers with 3 or more children) up to a maximum benefit level. After a small earned income range over which the credit stays constant, the credit is gradually phased out to zero at the point where family income equals maximum income for EITC eligibility.

As Figure 1 shows, the credit is primarily targeted at families. In 2012, the maximum credit available to a childless EITC filer was $475, while the maximum credits for families with children were substantially higher (ranging from $3,169 to $5,891). The credit phase-in rate for individuals without children, which is designed to offset payroll tax liability at 7.65 percent, has been unchanged since a credit was first offered to childless workers in 1994. Credits for workers with children, on the other hand, have been significantly expanded since the early 1990s. Between 1993 and 1996, the maximum credit for families nearly doubled. Additionally, a higher credit for families with three or more children was available for the first time in 2009.

At the same time that the federal EITC has grown significantly, a number of states have implemented their own supplemental credits. Although in 1989 only four states offered EITCs, that number had increased to 10 states by 1998, and 25 states and the District of Columbia had EITCs by 2013. The CBPP (2008) estimated that at the start of 2009, almost two out of five federal EITC recipients lived in a state with its own EITC. Table 1 provides a list of states with EITCs covering the 1992 to 2006 period that we study. State EITCs are largely set as a flat percentage of the federal credit, ranging in 2006 from 5 percent for all workers with children in Illinois, Maine, Oklahoma and Oregon to 43 percent for workers with 3 or more children in Wisconsin.[2] This translates to state credits of between $226 and $1,944 per year. By 2013, most states offered refundable EITCs, where any excess of the credit beyond the tax liability is paid as a tax refund. Although a handful of states started out by implementing non-refundable credits, only four states had credits were non-refundable in 2014.

Although in the conceptual framework that follows we refer to the design of the federal credit, changes in state-level credits are used for identifying variation in the empirical analysis. Because states' EITCs are almost always flat percentages of the federal credit, they follow the same phase-in and phase-out schedules as the federal credit. This has two implications that are relevant to our empirical work. First, the existence and generosity of state credit amounts do not change the range (phase-in, flat, and phase-out) of the credit that a recipient family falls into. Second, state supple-

[2] The exceptions are Indiana (from 1999–2002) and Minnesota. According to the Tax Policy Center, "Indiana has enacted a refundable tax credit (described in statute as an "earned income tax credit") for low-income working families with children. [But] the credit is unavailable to a large portion of the recipients of the federal credit (Tax Policy Center, "State Earned Income Tax Credits Based on the Federal EITC Tax Year 2015," http://www.taxpolicycenter.org/taxfacts/Content/PDF/state_eitc.pdf). Minnesota's EITC is not structured as a percentage of the federal credit. Depending on income level, the credit for families with children may range from 25 percent to 45 percent of the federal credit; taxpayers without any children may receive a 25 percent credit (Tax Policy Center, 2010).

Table 1

State EITCs, 1992–2006
(Nominal and Real Values)

Year	States with (Newly Implementing) EITC		Nominal Credit Value for States with an EITC		Real Value ($2009) for States with an EITC	
			Mean	Maximum	Mean	Maximum
1992	Iowa, Maryland (NR),	F	455	1,038	695	1,588
	Rhode Island, Wisconsin,	R	292	1,038	445	1,588
	Vermont, Minnesota					
1993		F	508	1,133	753	1,696
		R	335	1,133	495	1,696
1994	New York	F	655	1,264	950	1,832
		R	406	1,264	589	1,832
1995		F	817	1,555	1,152	2,192
		R	511	1,555	720	2,192
1996		F	950	1,778	1,301	2,435
		R	599	1,529	821	2,094
1997	Massachusetts, Oregon	F	820	1,828	1,100	2,449
		R	520	1,573	697	2,107
1998	Kansas	F	813	1,878	1,073	2,478
		R	575	1,615	758	2,131
1999	Colorado	F	796	1,908	1,027	2,461
		R	577	1,615	745	2,116
2000	District of Columbia, Illinois,	F	722	1,944	902	2,430
	Maine, New Jersey	R	559	1,672	698	2,090
2001		F	801	2,004	970	2,424
		R	637	1,723	771	2,084
2002	Oklahoma	F	788	2,070	937	2,463
		R	630	1,780	750	2,118
2003	Indiana	F	828	2,102	969	2,459
		R	679	1,808	795	2,115
2004		F	848	2,150	966	2,451
		R	705	1,849	804	2,108
2005		F	868	2,200	955	2,420
		R	769	1,892	747	2,081
2006	Delaware, Virginia, Nebraska	F	868	2,260	919	2,395
		R	696	1,944	738	2,060

Notes: R = Refundable portion of credit only; F (Full) = Refundable + non-refundable credit. Maryland's credit was non-refundable until 1998. Indiana had an EITC not based upon the federal credit from 1999–2002. All states with credits in place in 1992 listed for 1992; in years 1993–2006 new adopters are listed in the first year they had an EITC. All states continued their EITCs for all years after adoption except Colorado, which suspended the credit in 2002.
Source: Schmeiser (2009) and CBPP (various publications)

ments that follow the same structure as the federal credit increase both the phase-in and phase-out rates of the credit, intensifying any behavioral incentives generated in those ranges.

III. CONCEPTUAL FRAMEWORK

There are multiple pathways through which the EITC might affect child health. As a starting point, it should be noted that the EITC was designed to encourage employment; because the credit is essentially a wage subsidy in the phase-in range of the credit, the EITC creates strong incentives for parental labor force participation. However, recipients in the phase-out range of the credit face disincentives to work more hours. Furthermore, over the entire range of the credit the income effect of the subsidy discourages labor supply on the margin, all else equal. Therefore, theory does not provide a clear prediction for the direction of the labor supply effect of the credit. However, it is important to note that the extent to which the EITC changes labor supply and other outcomes depends upon the extent to which eligible individuals participate in the program and understand its design. Estimates of EITC participation range from roughly 70 to 85 percent (Scholz, 1994; Blumenthal, Erard, and Ho, 2005; Plueger, 2009; and Jones, 2013). Plueger shows that take-up increased with the number of qualifying children and EITC amounts. Furthermore, taxpayers in the phase-in region of the credit had lower participation rates than those in the maximum benefit or phase-out regions. It is unclear whether those patterns relate to awareness of the credit increasing with income, lower incentives to file a return for people expecting lower amounts, or both. The empirical evidence is much clearer and strongly suggests that the primary labor supply effect of the EITC is to increase labor force participation, particularly among single mothers (Eissa and Liebman, 1996; Ellwood, 2000; Meyer and Rosenbaum, 2000; Grogger, 2003; Cancian and Levinson, 2006; Hotz and Scholz, 2006).[3] The size of these effects is substantial. Grogger (2003) estimates that the approximately $2,000 increase in the maximum federal EITC benefit between 1993 and 1999 resulted in an increase in employment of female heads of household that was just over 10 percent. Meyer and Rosenbaum (2001) conclude 60 percent of the total increase for single women relative to married women in the United States between 1984 and 1996 was due to the EITC and related tax changes.

Any increase in parental employment could have three consequences that would affect the production of child health: (1) higher family income through increased earnings; (2) less non-work time available for parental investments in child health; and (3) changes in access to public and private health insurance. Additionally, whether or not the EITC changes a recipient's labor supply, the credit itself produces a lump-sum income effect.

[3] There is some weaker evidence that the phase-out range decreases labor supply on the intensive margin, but only for married women (Eissa and Hoynes, 2004).

The EITC, therefore, has the potential to affect child health in a very complex way. Below, we discuss each of the three mechanisms for health effects — income, parental time, and health insurance.

A. Family Income

The first way in which the EITC may affect child health is through family income. The EITC acts as a wage supplement, raising the average income of all EITC recipients, regardless of where their income falls in the phase-in or phase-out structure of the credit. Additionally, to the extent that the EITC increases labor supply, earnings will also increase. Assuming that child health is a normal good, and all else equal, this income effect is expected to improve child health. Case, Lubotsky, and Paxson (2002) and others have documented a strong positive correlation between family income and child health, a relationship that persists even after controlling for parental education; doubling household income is associated with between a 4 and 7.9 percent increase in the probability of the child being in excellent or very good health. Yeung, Linver, and Brooks-Gunn (2002) point out that there are two perspectives that explain why higher family income might improve child outcomes. The first is the traditional economic investment perspective, in which parents invest both time and money in children, and money is used to purchase health and other human capital inputs (Becker, 1991). In this case, a positive effect of income on child health would be attributable to some combination of higher quantity or quality of medical care, healthier food, better living conditions, or toys and activities that promote physical activity.

An alternative explanation for the positive correlation between family income and child outcomes comes from the family process model (Yeung, Linver, and Brooks-Gunn, 2002). In this model, the primary determinants of child outcomes are parental time-based inputs, both in terms of quantity of hours and quality of parenting skills. Specifically, lower family income increases parental stress and lowers the effectiveness of time as an input to child development. The family process model was developed primarily to explain child behavioral and cognitive outcomes, but is relevant to any outcomes for which parental time inputs are important determinants, including health. For example, Case and Paxson (2002) show that higher-income parents are more likely use seat belts and set regular bedtimes and less likely to smoke; differences in parental stress are a much more likely explanation for these correlations than differences in financial resources. Additionally, within the family process framework, both consistently low income and unstable income over time are factors that contribute to parental stress. The EITC not only increases average income among participants but also acts to stabilize consumption by reducing year-to-year variation in income for many recipients whose labor market earnings fluctuate within the EITC eligibility range (Kniesner and Ziliak, 2002). This is because EITC benefits are higher at lower earnings levels, so individuals who are in the phase out range or have incomes too high to qualify for the EITC before an economic shock may lose hours of work, but this will be partially offset by an increase in EITC. This may be an additional income-related pathway from the EITC to improved child health.

B. Parental Time

Assuming that parents are more likely to work because of the EITC, it is not entirely clear what effect this might have on child health. Theoretically, in a model such as the one presented by Grossman (1972), employment has two opposing effects on health. Greater employment generates higher income, which should improve child health, as discussed above. Additionally, either labor force participation or an increase in hours might make a parent eligible for employer-sponsored health insurance coverage. At the same time, a parent's time is also likely to be an important positive input in the production of child health, and employment raises the opportunity cost of time investments in health. Therefore, if it were possible to separately identify the effects of changes in a parent's available time, apart from other effects of employment (which could be either positive or negative — improved self-esteem or increased stress), then one would expect employment to have a negative impact on health outcomes. However, the evidence on the impact of parental employment on child outcomes, in general, is mixed, with both negative and positive estimates, and with estimates that vary by child age. This suggests that parental employment has effects beyond re-allocation of parental time.

There is a small literature on the impact of employment on child health that could be explained by parental time spent with children. Crepinsek and Burstein (2004) document significantly worse nutritional outcomes among children of full-time working mothers compared to children of mothers who do not work outside the home. Overall, children of full-time working mothers are more likely to drink soda and less likely to have adequate iron and dietary fiber intake. The authors also find that children of full-time workers under age 5 are more likely to consume calories in excess of their energy requirements. This is one mechanism for the finding of a significant, positive relationship between maternal work hours and the probability that a child is overweight, even after controlling for unobservable maternal and family characteristics (Anderson, Butcher, and Levine, 2003). They extrapolate from their baseline results to show that the increase in hours worked by mothers in high-income families between the mid-1970s and mid-1990s explains somewhere between 12 and 35 percent of the increased prevalence of overweight children in those families.

Two other studies find that maternal employment may be associated with significantly worse health outcomes for children. Gennetian et al. (2010) provide weak evidence of an adverse effect on health status; their overall estimate is small and negative, but they also note that the deficiency for children of working mothers appears to be made up by higher income and health insurance coverage rates. Morrill (2011) documents much larger negative effects; however, the measures of health status in this study — overnight hospitalizations, asthma episodes, and injuries — are much narrower measures of status than we use in this study.

Ruhm (2008) shows that cognitive outcomes are negatively related to maternal employment for 10- and 11-year olds in higher socioeconomic groups but *positively* related to maternal employment for 10- and 11-year olds in socioeconomically disadvantaged groups. All of this evidence together suggests that as the EITC increases labor supply and, in turn, reduces time spent with children, it could also affect child health. However, the effect of employment may vary by child age and other factors.

C. Health Insurance

The final potential pathway through which the EITC could affect child health is insurance coverage. The EITC is designed to increase family income, which is positively associated with demand for non-group health insurance. Marquis and Long (1995) estimate a statistically significant income elasticity of 0.15. There is also evidence that income has a positive effect on participation in employer-sponsored insurance plans (Abraham, Vogt, and Gaynor, 2002). Case, Lubotsky, and Paxson (2002) consider health insurance as a mechanism by which income improves child health; although they find a strong and significant effect of insurance coverage, it does not change the estimated effect of income on child health. Additionally, as Baughman (2005) points out, the design of the EITC — specifically, the phase-out range, where the majority of recipients fall — intensifies the tax subsidy for employer-sponsored health insurance. She finds that the federal EITC expansion in the mid-1990s significantly increased private health insurance coverage rates of low-skill workers by 3.8 percent, a result of the combination of employment, income, and tax price effects.

The impact of the EITC on overall health insurance coverage among children is unclear. As discussed above, the EITC is likely to increase private health insurance coverage through increased labor force participation and higher income. However, because of the wide availability of public insurance for low-income children through Medicaid and the Children's Health Insurance Program (CHIP), the EITC may merely result in a change of the type of health insurance and not necessarily in increases in overall coverage rates. The passage of the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) in 1996 generated similar work incentives as those of the EITC, as it was intended to reduce the receipt of cash assistance and to increase labor force participation among low-income adults. Kaestner and Kaushal (2003) examine changes in the health insurance coverage of low-income single mothers and their children following PRWORA and show a decrease in Medicaid coverage and a smaller increase in employer sponsored and private health insurance, resulting in a higher proportion of uninsured. The authors also show that the effects among children were somewhat smaller.

If the EITC does promote greater health insurance coverage rates among children, it is not clear that it will have an effect on health status. Lower out-of-pocket health care costs associated with any form of insurance coverage (compared to being uninsured) should encourage recipients to use more medical care (Manning et al., 1987). But such increased utilization may or may not translate into measurably improved health. And even with higher utilization, changes in health may not happen immediately.

While the literature suggests that health insurance coverage is associated with improved health outcomes for certain vulnerable populations (Doyle, 2005; Decker, 2005), the effect of insurance on the broader population is not as clear (Levy and Meltzer, 2004). Additionally, all types of health insurance may not produce equivalent utilization and health effects. For example, one might think that a movement from public to private health insurance among newly employed families could result in no net impact on overall health insurance coverage rates, but a positive effect on health utilization and health outcomes if private insurance allows better access to care or higher quality of care. However, the difference in access and quality levels between public and private insurance is not clear in the data. For example, The National Center for Health Statistics (2014) uses having a regular place of medical care as a proxy for access among children, and finds no statistically significant difference between children in public and private health insurance. Similarly, the Medicaid and CHIP Payment and Access Commission (2012) shows that access to care for children in Medicaid and CHIP is similar to access among privately-insured children. On the other hand, there is no cost sharing in Medicaid and only minimal cost sharing (e.g., $5 for many services) in CHIP, so the families of children who transition from public to private insurance may face higher cost sharing (Selden et al., 2009).

IV. DATA

The primary data for the analysis come from the child supplement to the 1979 National Longitudinal Survey of Youth (NLSY79). The NLSY79 sampled approximately 12,000 individuals between ages 14 and 21 in 1979 and has followed them ever since, with annual interviews until 1994 and interviews every other year through 2006. The NLSY79 respondents, who are between ages 41 and 48 in the latest interview wave, have reported data on their families, including births, marriages, divorces, and child outcomes. Detailed data on children born to NLSY79 recipients are available in the matchable Child and Young Adult (CYA) file.

In this analysis, we use observations on children of mothers with less than a college education from the 1992 through 2006 waves of the CYA — a sample likely to be eligible for the EITC.[4] It should be noted that the sample becomes less representative of younger ages as the children of the respondents age. The distribution of children by age in each year is provided in Appendix Table 1. As an example, in 1992, 4- and 12-year-old children each accounted for 7 percent of the sample. In contrast, by 2006, 4-year olds made up 2 percent of the sample while 12-year-old children made up 14 percent of the sample.[5]

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[4] Although adding several years of earlier data would provide a more representative sample of early births to NLSY recipients, all of the health questions used in this analysis were not asked routinely until 1990. In 1990, all of the questions we use were asked but response rates to certain questions, such as mother-rated health status, were very low.

[5] As a result, the sample descriptive statistics shown later (that combine data from 1992 to 2006) may not necessarily match those in cross-sectionally representative surveys.

The data include information on family demographics and income, parental employment, as well as a set of health insurance coverage, medical care utilization, and health status variables for each child. The descriptive statistics for these variables are in Table 2.[6]

Table 2

Sample Descriptive Statistics

	Age 0 to 5	Age 6 to 14
Female (%)	50.42	49.72
Black (%)	23.32	22.67
Hispanic (%)	27.09	30.77
Children in Household	2.58	2.58
Adults in Household	1.92	1.90
Mother Married (%)	68.71	59.60
Mother Divorced (%)	9.39	16.82
Mother has HS Degree (%)	87.58	84.95
Mother Works (%)	64.96	73.17
Real Family Income ($2006)	58,493	50,879
Private Health Insurance (%)	67.54	65.58
Uninsured (%)	8.11	10.21
Doctor Visit (%)	81.02	59.76
Dental Visit (%)	42.49	64.59
Obese (%)	19.45	17.17
Underweight (%)	16.37	8.24
Fair/Poor Health (%)	n/a	3.41
Excellent Health (%)	n/a	67.89
Simulated Real State EITC ($)	201	246
Medicaid/CHIP Instrument	27.56	32.98
Real State AFDC/TANF ($)	540	502

Notes: All descriptive statistics are for the full sample, except for underweight, obese, fair/poor health, excellent health, and dental visits, which in the case of children age 0–5 are only measured for 3- to 5-year olds.
Source: NLSY Child and Young Adult 1992 to 2006 files, restricted to children age 14 or under, with mothers who have less than a college education, and complete health insurance and birth weight measures

[6] All of the descriptive statistics in Table 2 are unweighted. The unweighted statistics for health insurance coverage and medical care utilization match other published estimates for this time period much more closely than weighted statistics. The use of weights does not affect any of the regression results presented later in the paper, however.

The health insurance variables that we use as our first set of health-related outcomes are discrete indicators for whether a child has private health insurance coverage (employer-sponsored insurance or a non-group plan) or no health insurance coverage.[7] Because private and public (Medicaid or CHIP) health insurance coverage rates sum to the overall insurance coverage rate, we do not present results for public coverage separately.

Additionally, we consider two measures of medical care utilization as health-related outcomes that may be affected by the EITC. One reason that we look at utilization is that it may help to explain the mechanisms by which any changes in health status occur. The other reason is that certain types of utilization may be a proxy — although an admittedly imperfect one — for health improvements in the future. We select two measures of utilization that correspond to clinical recommendations for appropriate preventive care: whether or not the child had a "routine medical checkup" during the past 12 months and whether the child visited a dentist in the past 12 months. The American Academy of Pediatrics (2008) recommends well-child doctor visits at least once a year for children of all ages (and twice a year or more for children under age 3). They also recommend that pediatricians begin oral health screenings as early as 6 months and that children start regular dental visits at age 2 (American Academy of Pediatric Dentistry, 2009).

We consider several measures of a child's health status. One is appropriate weight for height based upon body mass index, or BMI. We have calculated discrete indicators for whether a child is underweight (below the 5th percentile in BMI for age) or obese (at or above the 95th percentile in BMI for age) for all children over age 2. Additionally, for the majority of children in the 6- to 14-year old sample, we have the parent's rating of the child's health status on a 1 to 5 scale, which we recode into discrete indicators of fair or poor health status, and excellent health status. Case, Lubotsky, and Paxson (2002) show that parent-reported health is strongly correlated with future hospitalizations and with chronic health conditions, and also that it is correlated with physician-reported health status. There are a large number of missing values (25 percent) for health status in the full 6 to 14 subsample; we explore this issue as a series of robustness checks in Section VI.

V. METHODOLOGY

In our empirical models, we identify the effect of the EITC on health by using variation in state-level EITCs. Specifically, the variation in state EITCs in our analysis is generated by fourteen states enacting EITCs and a number of states with existing EITCs

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7 Health insurance coverage in the NLSY is measured at the date of the interview and does not necessarily reflect coverage during all or part of the preceding year. Approximately 6 percent of children in our estimation sample are reported to have both public and private health insurance coverage. We code these children as having only public coverage for two reasons. First, we believe that enrollment in Medicaid HMOs may explain the double-reporting; we note that in previous specifications before our re-code private health insurance coverage was positively and significant associated with our Medicaid eligibility variable. Second, eliminating the double-counting in this way brings our mean coverage levels in line with other published statistics. Coding private health insurance in this way lowers the mean level of coverage in our sample, but does not change the magnitude of our estimates in Section VI.

increasing the real value of the credit (either overall or for larger families) during our study period.

We capture variation in state EITCs by using a simulated EITC benefit, borrowing from the technique used by Currie and Gruber (1996) in their study of public health insurance. In order to create this variable, we start with a January 2004,[8] nationally representative cross section of the U.S. population from the Survey of Income and Program Participation (SIPP) and restrict the sample to the sub-population most likely to receive EITC benefits — heads of household between ages 19 and 64 who have at least one child under age 18 in the household and less than a college education. All of the 17,719 observations in this dataset are run through the National Bureau of Economic Research's TAXSIM calculator, which estimates (among many other variables) the household's state EITC supplement in any given year. We repeat the TAXSIM calculation using the same sample for each state and each year, with its corresponding state EITC rules, between 1991 and 2006. Our simulated benefit variable equals the median state EITC value in the sample for each state-year combination — calculated separately for households with one and two or more children. The simulated value is matched to individual child observations in the NLSY data by state of residence, year, and number of children in the household.

The advantage of using a simulated benefit to measure variation in state EITCs, rather than actual median benefits, is that it captures variation in benefits within states over time, as well as by number of children in the family, based exclusively on changes in state policy rather than the demographic composition of states at different points in time. Additionally, the simulated variable should measure the dollar amount of EITC that the average recipient receives more accurately than alternative variables — such as maximum benefits for a given family size — and it incorporates more variation than a difference-in-difference specification, where all states with EITCs are treated equally.[9] Alternate specifications are discussed in the second half of Section VI.

We estimate effect of the EITC on child health outcomes by using the simulated benefit (*SIMEITC*) in the following probit model

$$(1) \qquad Y_{ij} = \alpha_s + \alpha_t + \beta SIMEITC_{jst-1} + \gamma_1 X_{it} + \gamma_2 Z_{it} + \delta Policy_{st} + \varepsilon_{ijst} .$$

In these models, Y_{ij} is a discrete variable measuring one of the child health outcomes discussed in Section IV for child *i* in a family with *j* children, where *j*=1, 2, or more. The

[8] This is the first complete month available without attrition in the 2004 Panel of the SIPP, the last one that fits within our study period.

[9] The TAXSIM calculations that we use to create the simulated instrument give the full value of the state EITC credit, but do not allow us to distinguish between refundable and non-refundable credits. However, we have information about whether or not a state credit was refundable in a given year. In results not reported, we estimate models with a "refundable EITC" dummy variable interacted with the simulated instrument. The results for the refundable interaction are very robust to the baseline ones. The only new result is a large, positive, and statistically significant private health insurance effect for 0- to 5-year olds that was not present in the baseline.

EITC variable is lagged by one year because most recipients largely see their credit when taxes are filed and receive any refunds at the start of the next calendar year. Therefore, any income effects would take a year to materialize. Additionally, many recipients are likely to learn about their state's EITC (and may adjust their employment decisions) only after they have learned about it by filing a tax return. We test the robustness of our results to the choice of lag structure in Section VI.B.

The vector X contains child-level control variables including gender, race, and ethnicity. We also control for birth weight to account for the persistence of the effects of low birth weight later in life. This should help to address, at least in part, the heterogeneity in underlying health across children. The vector Z contains controls for parent and family characteristics, including age, marital status, and education of the child's mother, the number of adults and other children in the household, and whether or not the child's biological father lives in the household.

Finally, in the *Policy* vector we include two variables that account for within-state and over-time changes in two other public programs that may have affected child health through changes in family income, health insurance coverage, or both. The first variable is a measure of child eligibility for Medicaid and CHIP, using the same type of technique used to create the EITC variable; in this case the measure is a simulated eligibility level. Specifically, the variable is constructed by calculating the fraction of children under age 19 in a 1990 SIPP national data sample who would be eligible for Medicaid or CHIP, according to each state's rules in a given year. Similar to the simulated EITC variable, the measure varies within and across states over time by legislative rules, but not by differences in the demographic characteristics of an actual state population that might have independent effects on health. The second state-level control variable is the real ($2006) maximum AFDC-TANF benefit for a family of three. Both of these variables are matched to individual child observations in the NLSY by state and year.

All of the models are stratified by child age, with separate models run for children age 5 or younger and over age 5. There are many reasons to believe that the parameters of a production function for child health change as a child ages; for example, time spent with children may be more important at certain ages and monetary inputs at other ages. There is evidence in the literature that the positive income gradient for health gets stronger as children get older (Currie and Stabile, 2003; Case, Lubotsky, and Paxson, 2002). The impact of maternal employment on child health also appears to change, and perhaps even change sign, as children age (Ruhm, 2008).[10]

Each of the models includes state, year, and child age fixed effects. Ideally, because the NLSY-CYA is a panel, we would like to be able to estimate models with child fixed effects. However, because we split the sample by child age, are not able to use every year

[10] The sample stratification between age 0–5 and 6–14 is validated by a series of likelihood ratio tests (based upon the probit specification in Section VI.A) that reject the null hypothesis of pooled age groups. Results are available upon request.

of available data,[11] and have some attrition in the dataset, child fixed effects would tend to absorb most of the variation in the data. In any case, individual fixed effects would be most necessary to deal with unobserved child-level heterogeneity that is somehow correlated with state policy; we do not have reason to believe that this is a major issue in practice, especially after controlling for birth weight.[12] We cluster standard errors in all models by state.

VI. RESULTS

A. Baseline Models

The results of the baseline models for children ages 5 or younger are presented in Panel A of Table 3. Although the point estimate on private health insurance coverage is positive and the one on uninsured is negative, neither is statistically significant at conventional levels. In fact, there are no significant effects for any health-related outcome in this younger-child subsample for either the state EITC or the other two policy variables.[13]

The lack of an EITC health status effect is consistent with both the lack of an insurance effect and the finding in the child outcomes' literature of a stronger income gradient at older ages. On the other hand, it may be inconsistent with the positive impact of the EITC on maternal employment, particularly since the literature suggests that employment has more negative effects for younger children. However, the health status measures available for the youngest children are limited to body weight variables — and these are only measured for children age 3 or older. It is also possible that changes in health status may not occur immediately following economic investments.

The results of the baseline models for children age 6 to 14 are presented in Panel B of Table 3. In this sub-sample, the EITC is associated with a marginally statistically significant change in health insurance coverage patterns. A $100 increase in the median simulated value of state EITC is associated with a 4.1 percentage point increase in private coverage — a 6 percent increase relative to the group mean. As shown in

[11] There are 1991 and 1993 NLSY waves, but these surveys do not contain all of the necessary health variables for our analysis.

[12] Our baseline models are estimated on the sample of children whose birth weight is reported — 90 percent of NLSY-CYA panel that would otherwise meet our sample selection criteria. Leaving out birth weight does not change the magnitude or significance of any results.

[13] The lack of significant positive Medicaid/CHIP effects on utilization and health status measures from surveys for a nationally representative sample of younger children is consistent with some previous studies (Li and Baughman, 2011; Kaestner, Joyce, and Racine, 1999). The majority of studies of health insurance coverage patterns do find some combination of crowding out of public insurance and reduction in uninsured, although often for a more narrowly defined sample (e.g., mothers with a high school education or less) than ours. However, the lack of result in Table 3 could be driven by model specification. In the models shown in Table 4 with maximum EITC benefit variables and no lags, Medicaid/CHIP is associated with significant decreases in uninsured for both age groups.

Table 3

Baseline Probit Estimates of the Impact of State EITCs by Age Group

	Private Health Insurance	Uninsured	Doctor Visit	Dentist Visit	Fair/Poor Health	Excellent Health	Obese	Underweight
Panel A: Age 0 to 5								
Simulated EITC ($100)	0.0347	–0.0196	0.0021	0.0280	NA	NA	0.0256	0.0072
	(1.34)	(1.25)	(0.11)	(0.90)			(1.08)	(0.28)
Medicaid/CHIP	–0.0007	–0.0007	0.0009	–0.0026	NA	NA	–0.0007	0.0031
	(0.40)	(0.77)	(0.81)	(1.49)			(0.54)	(2.17)
AFDC/TANF ($1,000)	–0.0061	–0.0553	0.0621	0.0449	NA	NA	0.0876	0.0182
	(0.02)	(0.42)	(0.39)	(0.17)			(0.41)	(0.09)
Observations	5,012	4,829	5,019	3,911	NA	NA	2,894	2,894
Panel B: Age 6 to 14								
State EITC ($100)	0.0407*	–0.0053	–0.0261	–0.0102	–0.0120**	0.0339*	–0.0189	–0.0042
	(1.90)	(0.48)	(1.59)	(0.56)	(2.00)	(1.79)	(1.35)	(0.47)
Medicaid/CHIP	0.0014	–0.0008	–0.0001	0.0008	–0.0001	–0.0007	0.0002	0.0003
	(1.41)	(1.45)	(0.07)	(1.00)	(0.23)	(0.88)	(0.45)	(0.58)
AFDC/TANF ($1,000)	0.1300	–0.0935	0.2290	0.1116	–0.0598	0.1681	0.0674	–0.0854
	(0.79)	(1.16)	(1.58)	(0.74)	(1.08)	(0.99)	(0.63)	(1.14)
Observations	14,592	14,592	14,562	14,436	10,587	10,938	13,250	13,120

Notes: Results presented are marginal effects from probit models. Z-statistics calculated using robust standard errors clustered by state are in parentheses. NA = not available. Asterisks denote significance at the 1% (***), 5% (**), and 10% (*) levels. All models contain state, year, and child age fixed effects and controls for birth weight, gender, race/ethnicity, number of children in the household, number of adults in the household, mother's marital status, and mother's education. Self-reported health is not available for age 0 to 5. Body weight is not available for children younger than age 3.

Source: NLSY Child and Young Adult 1992 to 2006 files, restricted to children age 14 or under, with mothers who have less than a college education, and complete health insurance and birth weight measures

Table 2, the average value for the EITC value is just over $200, so this 6 percent increase would be relative to increasing state EITCs by approximately 50 percent, producing an elasticity of 0.12.[14]

However, this increase in private health insurance appears to be offset by a decrease in public coverage, resulting in no significant effect on uninsurance. As noted in Section III, increased private health insurance coverage as a result of the EITC could either be created by higher employment rates that lead to better access to employer-sponsored insurance (ESI), or greater income available to purchase non-group health insurance policies and finance ESI cost sharing.

As shown in the third and fourth columns of Table 3, there are no statistically significant EITC effects on utilization of medical care, nor are there significant changes in measures of underweight and obesity (Columns 7 and 8). However, there are significant and meaningful effects on mother-rated health status. A $100 increase in the median simulated value of the state EITC is associated with a 1.2 percentage point decrease in the probability that a mother reports her child to be in fair or poor health status, translating to 35 percent decrease relative to the group mean. There is also a 3.4 percentage point, or 5 percent, increase in the probability that a mother reports her child to be in excellent health.

B. Specification Tests

First, we test the robustness of our statistically significant results for children ages 6 to 14 from Table 3 to a variety of alternative measures of state EITC values and lag structure. In particular, we compare the baseline results to specifications using either the maximum credit for one, or two, or more children, and no lag in the state EITC. These results are presented in Table 4. The choice of a one-year lag versus no lag (in Panels A and B, respectively) seems to make very little difference in the results. This is consistent with eligible individuals anticipating a future lump sum income increase from the credit and changing their current spending accordingly.

Panels A and B contain comparisons of the results of the baseline models from Table 3 against two alternative EITC measures — the maximum credit for families with two or more children and the maximum credit for a one-child family. The estimates for private health insurance are still positive and become even more significant when using the maximum credit for families with two or more children. In the one-year lag model, the magnitude of the new estimates suggests that a $100 increase in the variable is associated with a 1.1 to 1.4 percentage point (or approximately 2 percent) increase

[14] The simulated benefits levels are not equivalent to measures like maximum or average benefits among recipients. The simulated value captures variation in both whether or not a family lives in a state with an EITC and qualifies for the credit, as well as the conditional amount of the benefit. To give a sense of the dollar increases that would generate a 50 percent increase in the simulated value, it is approximately equal to nominal growth in the maximum annual credit for two children in Minnesota between 1991 and 1996 (from $124 to $467) or in New Jersey between 2000 and 2006 (from $300 to $800).

Table 4

Probit Estimates of the Impact of State EITCs for Alternative State EITC Variables, Age 6 to 14

	Private Health Insurance	Uninsured	Fair/ Poor Health	Excellent Health
Panel A: One-Year Lag				
Simulated EITC (Baseline)	0.0407*	–0.0053	–0.0120**	0.0339*
($100)	(1.90)	(0.48)	(2.00)	(1.79)
Maximum credit for two + children	0.0109***	–0.0015	–0.0269***	0.0614**
($100)	(3.33)	(0.66)	(2.94)	(2.32)
Maximum credit for one child	0.0139*	0.0007	–0.0060***	0.0112
($100)	(1.90)	(0.20)	(2.95)	(1.60)
Panel B: No Lag				
Simulated EITC ($100)	0.0366*	–0.0094	–0.0131**	0.0317*
	(1.89)	(0.93)	(2.33)	(1.80)
Maximum credit for two + children	0.0136***	–0.0008	–0.0329***	0.0600
($100)	(3.31)	(0.42)	(3.46)	(1.63)
Maximum credit for one child	0.0156**	–0.0006	–0.0066***	0.0098
($100)	(2.16)	(0.18)	(3.27)	(1.36)

Notes: Results presented are marginal effects from probit models. Z-statistics calculated using robust standard errors clustered by state are in parentheses. Asterisks denote significance at the 1% (***), 5% (**) and 10% (*) levels. All models include the same controls as those listed in Table 3.
Source: NLSY Child and Young Adult 1992 to 2006 files, restricted to children age 14 or under, with mothers who have less than a college education, and complete health insurance and birth weight measures

in coverage. This result is comparable to the one estimated without a lag and also to the baseline. Regardless of the specification, we still find no effect on uninsurance.

The reduction in fair/poor health is also highly robust in terms of sign and statistical significance, although the magnitudes of the point estimates vary. For example, using a measure of state EITCs that is lagged by one year (the top panel), the estimated effect on fair/poor health status of a $100 credit increase ranges from 0.6 percentage points (for the maximum credit for one child variable) to 2.7 percentage points (for the maximum credit for two or more children variable). Finally, although not all estimates achieve statistical significance at conventional levels, the estimates for excellent health are robust in sign and, in some cases, are of slightly greater magnitude.

As motivation for our second specification check, we note that although most of the NLSY-CYA child health outcome variables had slightly less than perfect response rates, non-response is much more common for the health index variable, with a full 25 percent of parents of children age 6 to 14 in the sample not reporting a health status for their children. We have analyzed the factors correlated with non-response to the health index question and found that the most important predictors of non-response are the year of the survey and the age of the child. Non-response rates were dramatically higher in 2006 and for children under age 11. Also, the more children in a household, the more likely a parent was to respond to the health status question. None of the other demographic variables in the model were significantly related to non-response probability, nor — more importantly — was the state EITC variable.

As a specification check on our results, we re-estimate the baseline fair/poor health and excellent health models on two different sub-samples: (1) 6- to14-year olds in all years but 2006 and (2) 11- to 14-year olds in all years but 2006. (The non-response rate falls from 25 percent in the baseline sample to 21 and 6 percent, respectively.) The results for these specifications are provided in Table 5. The estimate for fair/poor health for 6- to 14-year olds has almost the exact magnitude as the baseline and is significant at the 5 percent level. And while the estimate for 11- to 14-year olds loses precision

Table 5

Probit Estimates of the Impact of State EITCs on Health Status Models Using Alternative Samples

Outcome Variable	Sample Description	Non-Response Rate for Health Status Question (Percent)	Coefficient (z-statistic)
Fair/Poor	Age 6 to 14 Years 1992–2004	21.0	–0.0129** (2.08)
Fair/Poor	Age 11 to 14 Years 1992–2004	2.6	–0.0133 (1.50)
Excellent	Age 6 to 14 Years 1992–2004	21.0	0.0391** (1.99)
Excellent	Age 11 to 14 Years 1992–2004	2.6	0.0439* (1.78)

Notes: Results presented are marginal effects from probit models. Z-statistics calculated using robust standard errors clustered by state are in parentheses. Asterisks denote significance at the 1% (***), 5% (**), and 10% (*) levels. All models include the same controls as those listed in Table 3.

Source: NLSY Child and Young Adult 1992 to 2006 files, restricted to children age 14 or under, with mothers who have less than a college education, and complete health insurance and birth weight measures (likely due to reduced sample size), it still has a relatively high z-statistic and has a very similar magnitude to the other estimates. Both estimates for excellent health are also highly robust and of a slightly higher magnitude. This gives us confidence that despite the shortcomings in the health status variable, what we are measuring is strongly suggestive of a true health-improving effect of the EITC for older children.

Next we note that in our baseline models we treat the panel NLSY-CYA data as a repeated cross section, and as a consequence we observe many children more than once in each subsample (ages 0 to 5 and 6 to 14). In order to show that this is not driving our significant results for older children, in Table 6 we present results in which the sample is stratified into two-year bands of age.[15] Because the NLSY only surveyed respondents every two years in the period we are studying, this ensures that no child is observed more than once.

Despite the greatly reduced sample sizes in these models, the pattern of results is generally similar to the one in the baseline models (Table 6). The increase in private health insurance coverage is remarkably consistent in terms of point estimate and significance across all age groups. Although, as with the baseline, the effect of state EITCs on uninsurance is not statistically significant, the point estimates are all negative. For fair/poor health, the significant reductions in the baseline model appear to be driven primarily by children ages 7 and 8 and children ages 11 and 12, although all point estimates are negative. There is no obvious explanation — such as an age gradient of the income effect — to explain this pattern; the variation in significance across age groups may simply be driven by smaller sub-sample sizes. None of the excellent health effects are statistically significant at conventional levels, which is not surprising given that the baseline effect was only significant at the 10 percent level. However, three out of four estimates are positive. There is only suggestive evidence of stronger effect with age, as the point estimate for 13- to 14-year olds is largest and closest to being significant.

Finally, we want to make sure that the effects we are estimating are not picking up other time-varying state characteristics that could be correlated with a child's health. We used two approaches to test this aspect of our specification. First, we add state-specific time trends to our model. The results of these models are statistically insignificant, which is likely to be caused by our small sample size. However, some of the point estimates also change in the state trends model, which indicates that there could be some unobservable time-changing state characteristics that we are not controlling for in our baseline model.

Second, we estimate a model in the spirit of a falsification check by estimating our model for a sample of better-educated workers. Education is a strong, if not perfect, proxy for income, so workers with a college education should be much less likely to receive the EITC compared to mothers with lower education levels. Accordingly, we

[15] In these models, we drop 6-year olds in order to create two–year bands. The pattern of results using 6- to 13-year olds is not qualitatively different.

Table 6
Probit Estimates of the Impact of State EITCs by 2–Year Age Bands, Children Age 7 to 14

	Private Health Insurance	Uninsured	Fair/Poor Health	Excellent Health
Panel A: Age 7 to 8				
Coefficient	0.0641*	–0.0025	–0.0262**	0.0591
Z-Statistic	(1.82)	(0.13)	(2.26)	(1.05)
Sample Size	*3,028*	*3,006*	*1,134*	*1,565*
Panel B: Age 9 to 10				
Coefficient	0.0595*	–0.0212	–0.0052	–0.0016
Z-Statistic	(1.73)	(0.83)	(0.51)	(0.07)
Sample Size	*3,450*	*3,421*	*2,226*	*2,583*
Panel C: Age 11 to 12				
Coefficient	0.0505*	–0.0118	–0.0202***	0.0340
Z-Statistic	(1.73)	(0.85)	(2.88)	(1.05)
Sample Size	*3,741*	3,725	*3,260*	*3,470*
Panel D: Age 13 to 14				
Coefficient	0.0571**	–0.0260	–0.0138	0.0709
Z-Statistic	(2.11)	(1.39)	(1.05)	(1.54)
Sample Size	*3,054*	*3,042*	*2,564*	*2,646*

Notes: Results presented are marginal effects from probit models. Z-statistics calculated using robust standard errors clustered by state are in parentheses. Number of observations are in italics. Asterisks denote significance at the 1% (***), 5% (**), and 10% (*) levels. All models include the same controls as those listed in Table 3.
Source: NLSY Child and Young Adult 1992 to 2006 files, restricted to children age 7 to 14, with mothers less than a college education, and complete health insurance and birth weight measures

estimate a set of baseline models for children ages 6 to 14 with mothers with college degrees for purposes of comparison to our baseline. These results are presented in Table 7. The only significant effect is a marginally significant decrease in dental visits. The fact that there are no significant increases in private health insurance coverage or reported health status indicate that the estimates in Tables 3 through 6 are probably measuring causal impacts of state EITC expansions rather than simply picking up time-varying state unobservables.

Table 7

Probit Estimates of the Impact of State EITCs,
Children Age 6 to 14 with College-Educated Mothers

Dependent Variable	Coefficient on State EITC Variable
Private Health Insurance	0.0136
	(1.04)
Uninsured	–0.0099
	(1.18)
Doctor Visit	0.0224
	(0.60)
Dental Visit	–0.0371*
	(1.74)
Fair/Poor	–0.0067
	(0.83)
Excellent	0.0227
	(0.60)
Obese	0.0151
	(0.74)
Underweight	0.0013
	(0.07)

Notes: Results presented are marginal effects, with each row being a separate probit model. Z-statistics calculated using robust standard errors clustered by state are in parentheses. Asterisks denote significance at the 1% (***), 5% (**), and 10% (*) levels. All models include the same controls as those listed in Table 3.
Source: NLSY Child and Young Adult 1992 to 2006 files, restricted to children age 6 to 14, with mothers who have a college education, and complete health insurance and birth weight measures

VII. CONCLUSION

In this study, we analyze the impact of changes in state EITC levels on health insurance coverage, utilization of medical care, and health status among children, using variation in the timing and generosity of state EITCs. The EITC is not associated with any significant changes in health insurance coverage, medical care utilization, or health status outcomes for younger children (age 0 to 5); however, it should be noted that the health status outcomes available in the data for younger children are very limited.

State EITCs appear to have led to an increase in private health coverage for children age 6 to 14. Our exploratory analysis (results not reported) suggests that the expansion in private health insurance coverage was made up primarily of an increase in individual (non-group) health insurance policies. This could be a true increase in non-group coverage, an indicator for higher rates of CHIP coverage (based upon the misreporting of non-group for CHIP shown in LoSasso and Buchmueller (2004)), or a combination of the two.

The EITCs are also associated with significantly lower probability that children age 6 to 14 are in fair or poor health, and significantly higher probability that children of the same age are in excellent health. These health status effects are robust to a number of specification checks, despite a high non-response rate to the health status questions in the NLSY-CYA survey. In this analysis, we are not able to determine the precise mechanism for changes in health status. The lack of significant effects for doctor and dental visits suggests that access to medical care is not an important factor; however, the quality of medical care could still have been affected. One possibility is that substitution between public and private insurance improved the quality of providers. On the other hand, quality could have declined if higher cost sharing in private health insurance relative to public plans changes the types health care a family uses – for example, switching from one prescription drug to another. Improved health status is also consistent with a positive income effect of the credit. There is also a small amount of evidence in the literature that maternal employment could be health-improving in older children. However, when we estimate the direct relationship between maternal employment and state EITCs, we do not find a significant relationship.[16] It is also possible that the higher subjective measure of health in our study reflects better mental rather than physical health, a result of lower parental (and therefore child) stress associated with a better financial situation. This type of an effect would be consistent with one discussed in the context of the Oregon Medicaid Experiment, where researchers argue that a substantial portion of the increase in self-reported health status that they observe is associated with a higher level of well-being rather than specific physical health improvements (Finkelstein et al., 2012). The medical and economics literatures show a positive relationship between childhood health and adult health (see, e.g., Case, Fertig, and Paxson, 2005). Thus, to the extent that they improve health status in childhood, state EITCs could potentially result in improved health in years to come.

There are several limitations to this analysis, and the most important one is that our analysis, particularly the study of health outcomes, has necessarily been limited by the

[16] There are no existing positive, significant estimates of the EITC's effect on female labor force participation that are based upon variation in state-level credits across the country. The analysis by Cancian and Levinson (2006) is based on a single-state analysis and finds no effect.

variables available in the survey data. It is possible that there are short-term health status changes that we are unable to capture using only parent-reported health status and body weight measures. It is also likely that certain types of health effects materialize over a much longer time period than the one we study; in this sense, our estimates may understate the true impact of the EITC on health status. Additionally, the estimates from this analysis may not be fully generalizable to all children today because the analysis is based upon a cohort of children that ages over the study period. Specifically, these estimates are most representative of children born to older mothers and least representative of children born to younger mothers.

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APPENDIX TABLE 1

NLSY CYA Sample Size by Child Age and Year

Age	1992	1994	1996	1998	2000	2002	2004	2006
0	221	162	126	102	58	28	12	6
1	298	210	121	101	65	42	22	14
2	306	218	170	126	102	60	34	14
3	320	266	188	136	78	57	40	26
4	354	284	206	177	126	104	65	30
5	348	317	264	202	131	89	67	45
6	363	328	287	210	151	111	101	63
7	425	310	303	285	200	117	98	73
8	401	356	305	281	194	166	119	107
9	399	394	311	317	249	181	123	90
10	412	347	342	327	281	195	178	139
11	366	367	401	308	273	239	200	122
12	345	385	339	330	295	278	208	192
13	248	342	383	371	283	274	241	212
14	209	308	369	352	315	297	308	227
TOTAL	5,015	4,594	4,115	3,625	2,801	2,238	1,816	1,360

Source: NLSY Child and Young Adult 1992 to 2006 files, restricted to children age 14 or under, with mothers who have less than a college education, and complete health insurance and birthweight measures