NBER WORKING PAPER SERIES

CHILDHOOD MEDICAID COVERAGE AND LATER LIFE HEALTH CARE UTILIZATION

Laura R. Wherry
Sarah Miller
Robert Kaestner
Bruce D. Meyer

Working Paper 20929
http://www.nber.org/papers/w20929

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
February 2015

This work was supported by the Robert Wood Johnson Foundation's Health Policy Scholars and Health & Society Scholars Programs at the University of Michigan, Ann Arbor, as well as the California Center for Population Research. We would like to thank participants at seminars at the University of Michigan for excellent comments, as well as Jean Roth, Betty Henderson-Sparks, and Dee Roes for their assistance in accessing the data used in this project. The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

Childhood Medicaid Coverage and Later Life Health Care Utilization
Laura R. Wherry, Sarah Miller, Robert Kaestner, and Bruce D. Meyer
NBER Working Paper No. 20929
February 2015
JEL No. I12,I13,I28

ABSTRACT

Policy-makers have argued that providing public health insurance coverage to the uninsured lowers long-run costs by reducing the need for expensive hospitalizations and emergency department visits later in life. In this paper, we provide evidence for such a phenomenon by exploiting a legislated discontinuity in the cumulative number of years a child is eligible for Medicaid based on date of birth. We find that having more years of Medicaid eligibility in childhood is associated with fewer hospitalizations and emergency department visits in adulthood for blacks. Our effects are particularly pronounced for hospitalizations and emergency department visits related to chronic illnesses and those of patients living in low-income neighborhoods. Furthermore, we find suggestive evidence that these effects are larger in states where the difference in the number of Medicaid-eligible years across the cutoff birth date is greater. We do not find effects on hospitalizations related to appendicitis or injury, two conditions that are unlikely to be affected by medical intervention in childhood. Our calculations suggest that lower rates of hospitalizations and emergency department visits during one year in adulthood offset between 3 and 5 percent of the initial costs of expanding Medicaid. This implies substantial savings if the decline in utilization spans multiple years or grows with age.

Laura R. Wherry
University of California, Los Angeles
David Geffen School of Medicine
911 Broxton Avenue
Los Angeles, CA 90024
lwherry@mednet.ucla.edu

Sarah Miller
University of Michigan
Stephen M. Ross School of Business
Business Economics and Public Policy
Ann Arbor, MI 48109
mille@umich.edu

Robert Kaestner
Institute of Government and Public Affairs
University of Illinois
815 West Van Buren Street, Suite 525
Chicago, IL 60607
and NBER
kaestner.robert@gmail.com

Bruce D. Meyer
Harris School of Public Policy
University of Chicago
1155 E. 60th Street
Chicago, IL 60637
and NBER
bdmeyer@uchicago.edu

A data appendix is available at:
http://www.nber.org/data-appendix/w20929

I. Introduction

One of the goals of publicly-subsidized health insurance is to improve the health of those without insurance. The argument underlying this policy is straightforward: health insurance provides the means to use more, and more timely, medical care, and because of this greater use of care, health is improved. As a result, those who gain coverage may need fewer expensive hospital and emergency department visits later in life due to their improved health, and these long term changes in utilization may partially or completely offset the initial cost of insurance provision. While the intuition behind this argument is strong, the empirical evidence to support it is relatively weak. For example, the Oregon Medicaid Experiment did not find significant health benefits from health insurance during the first two years of coverage and found that the provision of Medicaid increased, rather than decreased, the use of costly hospital and emergency department care.[1]

One limitation of the literature in this area is its relatively short time horizon. Most studies seek to link health insurance to health contemporaneously, or for a few subsequent years. However, the health benefits of insurance may be cumulative and revealed only after a sustained period of insurance and regular use of medical care. Shorter windows of analysis may not be adequate to identify the health benefits insurance.

A second limitation of studies evaluating contemporaneous effects of public health insurance on hospitalizations and emergency department visits is that they are unable to isolate the potential health benefit of insurance. For example, they cannot separate an improvement in health due to insurance that could ultimately result in fewer hospitalizations from the concurrent access effects of insurance that lower out-of-pocket costs and induce greater use of care. Even if an individual's health improves as a result of public insurance coverage, the access effect may dominate in the short term, leading to higher utilization of medical services.

In this paper, we address these issues by examining whether the expansion of Medicaid in the late 1980s and early 1990s improved the health later in life of those affected. Specifically, we exploit plausibly exogenous variation by birthdate in the cumulative number of years an

[1] Study of participants one to two years after expanded access to Medicaid showed significant improvements in self-reported health but no change on physical and clinical health measures (Finkelstein et al. 2012, Baicker et al. 2013).

individual was eligible for public health insurance coverage. To phase in the Medicaid expansions, Congress specified that several eligibility expansions for low-income children applied only to children born after September 30, 1983. As a result, children born before September 30, 1983 experienced lower rates of Medicaid eligibility and fewer Medicaid-eligible years in childhood than children born immediately following the cutoff. This discontinuity in eligibility was first identified and used by Card and Shore-Sheppard (2004) to examine contemporaneous changes in insurance coverage. Wherry and Meyer (2014) later demonstrated that the policy led to cumulative differences in childhood eligibility. They estimated that a child in a family with income just under the Federal Poverty Level (FPL) gained approximately five additional years of Medicaid eligibility during childhood if she were born on October 1, 1983 rather than September 30, 1983. Black children were particularly likely to benefit from the Medicaid expansions, gaining on average more than twice the number of Medicaid-eligible years of white children.

We exploit this policy discontinuity as a source of exogenous variation in Medicaid eligibility in order to evaluate the long-term effects of public insurance. For outcomes we use administrative data on hospital and emergency department (ED) visits from several states. These databases capture the universe of hospitalizations or ED visits in each state for a given year and provide sufficiently large sample sizes to detect changes in utilization among young (e.g., age 25) populations with relatively low usage rates. In addition to birth year and month, the data also provide information on other patient characteristics including race and median income of the zipcode of residence. This information allows us to examine changes in hospitalization and ED use among groups that were especially likely to be affected by the change in Medicaid policy.

An important contribution of our study is that we are able to control for the access effects of health insurance on hospitalizations and ED visits, and by doing so isolate the potential health benefits of insurance. We are able to disentangle the health effect of insurance on utilization from changes in out-of-pocket costs by analyzing the hospitalizations and use of ED care of young adults later in life, when there are no longer policy-driven differences in Medicaid eligibility or out-of-pocket costs between our treatment and control groups that could drive utilization patterns. We examine the effects of coverage one year after the cohorts have experienced the additional coverage (at age 15) and ten years later (at age 25), allowing us to capture both immediate and longer-term effects.

We find no immediate effects of the expansions on health care utilization at age 15. However, we find sizeable effects of Medicaid eligibility in childhood on hospitalizations and emergency department visits at age 25 among black cohorts who gained coverage. Black cohorts born immediately after the cutoff are estimated to experience approximately 8 to 13 percent fewer hospitalizations and 3 to 4 percent fewer emergency department visits at age 25 relative to those born just before the cutoff. Our results are particularly pronounced for hospitalizations and emergency department visits related to chronic illnesses and among patients from low-income zip codes. We do not find reductions in the utilization of non-blacks (who experienced smaller gains in eligibility at the birth date cutoff), nor do we find effects for hospitalizations related to appendicitis or injury, two conditions that are unlikely to be affected by access to care in childhood. Placebo tests using earlier birth cohorts and false “cutoff” points are small relative to the effects we estimate at the true birth date cutoff. Additionally, our analysis suggests that these effects are largest in states where the discontinuity in the cumulative number of Medicaid eligible years is greatest.

Our results provide several insights that are relevant to current policy debates surrounding the provision of public health insurance and the role of government in expanding coverage. First, our estimates indicate that between 3 and 5 percent of the initial cost of the Medicaid expansions for children were “offset” by lower hospitalization and emergency department usage at age 25 alone. If these effects persist, then the size of the cost offset is likely to be even greater. Second, our results highlight the importance of evaluating these programs over a longer time period. Indeed, we find no impact of Medicaid coverage in our analysis of the “immediate” effect at age 15, but do find effects later in life at age 25. These findings suggest that the benefits of insurance may only materialize over a long horizon.

II. Background

High-quality analyses of Medicaid eligibility expansions for children consistently show that Medicaid increases health care utilization, including hospitalizations, in the short term.[2] However, there are fewer studies of the effects of gaining Medicaid on children's health and the evidence from this literature is mixed. A number of studies using parental reports of child health find no evidence of improvement under public insurance, while several papers document significant declines in child mortality.[3] Thus, the effect of gaining health insurance on health remains an important, but unanswered question.

One limitation of studies seeking to assess the effect of insurance on health is that they examine how coverage affects health, for example, as measured by hospital admissions, immediately after or within a few years of the coverage expansion. If the health benefits of insurance are realized later, then the expansion of health insurance coverage may precede the observed health and utilization effects by several years. An emerging literature on the longer-term effects of health insurance coverage in childhood on later life outcomes has begun to address this issue. Boudreaux, Golberstein and McAlpine (2014) use variation in the timing of the introduction of the Medicaid program across states in the 1960s to identify long-term effects among cohorts with different exposure to the program. They find that those who gained access to Medicaid early in childhood were less likely to report having a chronic illness as an adult. Brown, Kowalski, and Lurie (2014) also use state-level variation in the timing of the Medicaid expansions for children in the 1980s to examine long-term effects and find that cohorts who gained coverage have higher wages, receive lower earned income tax credit payouts, have higher graduation rates and lower mortality as adults. Also relying on state-level variation, Cohodes et al. (2014) find that cohorts

---

[2] See evidence of increased hospital use in Dafny and Gruber (2005), Currie and Gruber (1996a) and Boudreaux, Golberstein, and McAlpine (2014). In addition, Currie and Gruber (1996a), Card and Shore-Sheppard (2004), and Currie, Decker and Lin (2008) present evidence indicating an increase in annual doctor visits under expanded public insurance.

[3] For example, Currie, Decker and Lin (2008), Currie and Gruber (1995), De La Mata (2012), and Racine et al. (2001) find no change in subjective measures of child health such as child health status and activity limitations. Meanwhile, Currie and Gruber (1996a, 1996b), Goodman-Bacon (2014), Howell et al. (2010), and Wherry and Meyer (2014) find significant effects on infant or child mortality. Not all studies, however, find mortality improvements; Decker, Almond, and Simon (2015) find no evidence of changes in maternal and child mortality under the rollout of Medicaid. See Howell and Kenney (2012) for additional review of this literature.

who gained coverage in childhood as a result of these Medicaid expansions have higher educational attainment, and Miller and Wherry (2014) find that cohorts whose mothers had higher eligibility rates for prenatal coverage while the cohort was in utero had better health outcomes and fewer hospitalizations in adulthood related to preventable health conditions.

In this paper, we add to this literature by exploiting the discontinuity in Medicaid eligibility and coverage among those born around September 30, 1983. In a paper complementary to this work, Wherry and Meyer (2014) examine the later life mortality of cohorts born after this cutoff date. They provide evidence linking this increase in childhood eligibility to a later decline in teenage mortality for black children who were more likely to gain eligibility under the expansions than white children. All previous papers on the long-term effects of Medicaid coverage (with the exception of Wherry and Meyer 2014) use state and year level variation in Medicaid policy to examine long-term outcomes. Although this empirical approach has been used many times in the literature, some authors have pointed out its limitations (e.g., the estimates tend to be sensitive to the inclusion of state-specific trends; see Dave et al. 2008). The regression discontinuity design we employ allows us to examine the effects of childhood Medicaid coverage in a way that is arguably more credible because it does not rely on using policy changes at the state level as an instrument for eligibility.

III. The Policy Discontinuity

Discontinuity in Eligibility

Prior to the 1980s, eligibility for Medicaid for non-disabled children was primarily limited to children in families receiving cash welfare under the Aid to Dependent Families with Children (AFDC) program. Recipients of AFDC benefits were primarily single-mother families with very low income levels, often well below the poverty line.[4] Beginning in the mid-1980s, Congress took steps to expand eligibility for Medicaid to children not participating in AFDC who would otherwise be ineligible for Medicaid benefits. In a series of legislative acts, eligibility for

[4] Income limits for the AFDC program were established by states and ranged from 14 to 79 percent of the federal poverty line in 1989 (U.S. General Accounting Office 1989).

Medicaid was expanded to all children with family incomes at or below the poverty line, regardless of family structure or participation in the AFDC program.

In an effort to phase in changes in Medicaid eligibility, Congress specified that many of the legislative changes applied only to children born after September 30, 1983. This provision meant that children born just before and after this birthdate cutoff faced very different eligibility criteria for Medicaid during their childhood years. Wherry and Meyer (2014) simulate childhood eligibility for public health insurance for cohorts born on either side of this birthdate cutoff.[5] They show that this unique feature of the expansions led to a large discontinuity in the number of years of Medicaid eligibility during childhood for cohorts born at this birthdate.

Given the nature the expansions, the discontinuity was largest for children with family incomes below the poverty line and above AFDC income levels. Figure 1 displays the average number of years of childhood eligibility for public insurance by birth month cohort for children in families with incomes below 150 percent of the federal poverty line. The magnitude of the discontinuity in childhood eligibility at the September 30, 1983 cutoff is largest for children in families with incomes between 75 and 100 percent of the poverty line. The gain represents an additional 4.6 years of eligibility during childhood. Children with incomes between 50 and 75 percent of poverty, as well as those with incomes between 25 and 50 percent of poverty, also experience sizeable gains with an additional 3.4 and 2.0 years of eligibility, respectively.

Figure 2 reveals that the gain in eligibility was primarily concentrated at ages 8 to 14 for children born immediately after the birthdate cutoff. This graph plots the share of the September versus October 1983 birth cohorts eligible for public health insurance at each age during childhood by race. Eligibility levels are similar for the two cohorts prior to age 8 and again starting at age 15. These cohorts were approximately 8 years of age at the implementation of the Omnibus Budget Reconciliation Act of 1990 (OBRA90), which required all state Medicaid programs to cover

[5] The authors use a random sample of children of ages 0-17 from each year of the 1981-1988 March Supplement to the Current Population Survey (CPS) and estimate eligibility for this pooled sample if born in each month between October 1979 and September 1987. They employ detailed federal and state public health insurance eligibility rules for the years 1979 to 2005 to estimate eligibility status for each month during childhood through age 17. This simulation holds family characteristics, including state of residence, family structure and size, parent employment and family income, constant over the child's lifetime. See Wherry and Meyer (2014) for additional information.

children under age 19 born after September 30, 1983. Later, the State Children's Health Insurance Program (CHIP) authorized state expansions of public health insurance to children in higher income families. The CHIP expansions served to close the gap in public eligibility for cohorts born on either side of the cutoff at around age 15. We examine health care utilization for cohorts born just before and after September 30, 1983 following the differential gain in Medicaid eligibility at age 15 and then 10 years later at age 25.

We also examine differential effects of the expansions by race and by state of residence. Black children were more likely to gain eligibility under the expansions (Table 1) due to their distribution of family income. On average, black children born in October versus September 1983 were 17 percentage points more likely to gain Medicaid eligibility. Among those who were made Medicaid-eligible, the average gain in eligibility throughout childhood was 4.8 years. This is over twice the average years of eligibility gained by non-black children, who would have experienced an 8 percentage point gain in eligibility across the birth date threshold that led, on average, to 4.5 additional Medicaid-eligible years throughout childhood.[6] The gain in Medicaid eligibility for children born after the cutoff also varied by state due primarily to differences in Medicaid policies in place before the expansions.[7] Table 2 presents estimates of the average eligibility gain for each state in our study that resulted from legislative changes in Medicaid policy.[8] In order to estimate changes in eligibility resulting from Medicaid policy rather than state socioeconomic characteristics, a national sample of children is used to estimate the average eligibility gain for each state given its eligibility rules. This follows the simulated eligibility literature and methodology (see Cutler and Gruber 1996; Currie and Gruber 1996a, 1996b). The

[6] Wherry and Meyer (2014) estimate eligibility gains by child race using a similar methodology as for all races but rely on a sample that draws children from the CPS by race and state cells. See their paper for additional information.

[7] Although many of the expansions were first introduced at state option, Wherry and Meyer (2014) estimate that the majority of the variation in eligibility at the September 30, 1983 was the result of a federal requirement for all states to cover children born after this date provided that their family incomes were below the poverty line.

[8] Although slow to implement Medicaid, Arizona provided government-supported health care for families on AFDC both prior to and following the introduction of its Medicaid program in 1982 (Freeman and Kirkman-Liff 1985). In addition, the federal mandate to expand eligibility for children born after September 30, 1983 with family incomes up to the poverty line applied to all states including Arizona (Congressional Research Service 1988).

size of the discontinuity in eligibility varies from 0.02 years of eligibility in California to 1.67 years of eligibility in Maryland.

Discontinuity in Coverage

In addition to a demonstrated discontinuity in childhood eligibility, we also measure any corresponding discontinuity in childhood coverage. It is important to bear in mind, however, that even if they did not take active steps to enroll in the program, all children gaining eligibility for Medicaid had "conditional coverage" in that their expenses were covered in the event of hospitalization or the need for costly medical care (Cutler and Gruber 1996). Not only could eligibility be granted retroactively for a period of up to 3 months prior to the date of application, but many states were giving children the opportunity to sign up for Medicaid at the sites where they received health care (Congressional Research Service 1993). Since the value of Medicaid is highest when children are sick, parents are likely to wait until medical care is needed to sign up for coverage (Marton and Yelowitz 2014).

Card and Shore-Sheppard (2004) first examined changes in Medicaid enrollment for children born after September 30, 1983 following the expansions in Medicaid eligibility. The authors found Medicaid take up rates of between 8 and 11 percent among the newly eligible, with little evidence of substitution of public for private coverage (i.e., crowd out). In an analysis of similar spirit, we explore differences in the discontinuity in coverage by child race. We use the pooled 1992-1996 National Health Insurance Survey (NHIS) Health Insurance Supplements to examine changes in Medicaid coverage for cohorts born after September 30, 1983 at ages 8-13. Our sample includes all cohorts born between the months October 1979 and September 1987. We estimate a simple regression discontinuity model and regress Medicaid coverage on an indicator for birth cohorts October 1983 and later, a quadratic function in birth month cohort interacted with this indicator, and a set of calendar month fixed effects. Standard errors are heteroskedasticity-robust and are allowed to be non-independent within birth month cohort cluster. We also use local linear regression to estimate the discontinuity in Medicaid coverage at the September 30, 1983 cutoff following estimation methods described later in the paper.

Figure 3 plots reported levels of Medicaid coverage for each birth month cohort. The graphs for blacks and, to a lesser extent, for all races, show evidence of an increase in Medicaid coverage at the cutoff. When we look separately at children in households with incomes below the poverty line, we see additional visual evidence of a discontinuity in coverage.

Table 3 presents the corresponding regression estimates. We estimate a significant increase in annual Medicaid coverage for blacks of 5 to 6 percentage points. Given our estimate that 17 percent of black children gained eligibility, this represents a take-up rate of approximately 31 to 34 percent. For non-blacks, we do not find a significant increase in Medicaid coverage and the point estimate is much smaller, indicating less than a one-percentage point change. Examining children with families below the poverty line only, we find an increase in Medicaid coverage of 7 to 8 percentage points, although the estimate is not consistently significant. We find no change among children in families with incomes above the poverty line.

Summary

Changes in Medicaid eligibility and coverage documented in Figures 1 through 3 and Tables 1 through 3 lead to an important empirical implication. There is clear variation in treatment by race, poverty, and state, and the differences range from zero to substantial. Accordingly, if Medicaid coverage has an effect on health and use of medical care such as hospitalization, then it is plausible to expect that effects will vary in a way consistent with the variation in treatment.

IV. Data

To conduct our analysis, we combine discharge-level hospital data from three sources. First, we use hospitalization data from the Healthcare Cost and Utilization Project (HCUP) State Inpatient Databases. These data provide discharge-level information on all inpatient hospitalizations that occurred in 1999 in Arizona, Iowa, New York, Oregon, and Wisconsin, and in 2009, on Arizona, Iowa, Maryland, New Jersey, New York, Oregon, and Wisconsin. We supplement these data with the census of hospital discharges that occurred in Texas and California in 1999 and 2009, obtained from the Texas Department of State Health Services and the California Health and Human Services Agency, resulting in the complete census of hospital discharges for 7 states in 1999 and for 9 states in 2009.

In addition to hospital discharge data, we use data on all outpatient emergency department visits that occurred in Arizona, Iowa, New Jersey, New York, and Wisconsin (obtained from HCUP) and California (obtained from the California Health and Human Services Agency) in 2009. These data cover all visits for which a patient was treated in an emergency department and released the same day, rather than being admitted to the hospital.

Both the hospital discharge and emergency department data contain information on the diagnoses associated with each visit, total charges, and patient demographics including race and birth month and year. In 2009, we observe whether the patient is from a low-income zip code (defined as a zip code with median income below $39,999).[9] We classify primary diagnoses as relating to "chronic" or "non-chronic" conditions using the Chronic Condition Indicator software distributed by HCUP.[10] We exclude hospitalizations and ED visits for diagnoses related to pregnancy and delivery.

Combined, our hospitalization data include 643,342 discharge-level observations for diagnoses not related to pregnancy and delivery and 3,031,928 emergency department visits in 2009 for patients born between 1979 and 1987. Our hospitalization sample covers approximately 29 percent of the national population in 1999 and 34 percent of the national population in 2009, and our emergency department visit sample covers about 20 percent of the US population. These large sample sizes are critical for our analysis because they allow us to detect changes in hospitalizations and emergency department visits even among young populations with low utilization rates and for conditions that are relatively rare.

Table 4 presents descriptive statistics on hospital and emergency department utilization rates from our dataset. The first three columns display hospitalization rates (per 10,000 individuals)

---

[9] Data obtained from HCUP contain a variable indicating that the median income of the patient's zip code is below $39,999. For data from Texas and California, we use the American Community Survey and individual patient zip codes to create this variable following the same criteria.

[10] Downloaded from http://www.hcup-us.ahrq.gov/toolssoftware/chronic/chronic.jsp on 11/11/2014. The HCUP Chronic Conditions Indicator categorizes all diagnosis codes as chronic or not chronic. The definition of a chronic condition requires that it lasts 12 months or longer and that it either (1) places limitations on self-care, independent living, and social interactions; or (2) needs ongoing intervention with medical products, services, and special equipment. The classification was developed based on an existing body of work on the chronicity of conditions and in consultation with a physician panel.

for 15 year-olds in 1999, the first year for which we have data. In 1999, there were approximately 265 hospitalizations (not including hospital visits related to pregnancy and delivery) per 10,000 population for all races. Hospitalization rates at this age were higher for blacks, who experienced approximately 323 hospitalizations per 10,000 individuals, and lower for non-blacks, who experienced approximately 258 hospitalizations per 10,000 individuals. About half of these hospitalizations were for chronic illnesses overall; for blacks, chronic illnesses represented about 60 percent of total hospitalizations. For 15-year-olds, the most common of these chronic illnesses are mental disorders, followed by asthma and diabetes.

The next three columns display hospitalization rates for 25 year-olds in 2009. Hospitalization rates are more common for this age group: among all races, there were 329 hospital visits per 10,000 population; among blacks, there were 548 visits per 10,000 population; among non-blacks, there were 303 visits per 10,000 population. About 57 percent of hospitalizations of black patients were for chronic conditions and about 46 percent of hospitalizations of non-black patients were for chronic conditions. The most common chronic condition for this age group is also mental disorders. The second most common is diabetes and the third most common is asthma.

Emergency department visits are more common than hospitalizations and tend to treat less severe conditions. ED use is described in columns 7 through 9. On average, there are 3167 emergency department visits per 10,000 individuals in 2009, roughly ten times the hospitalization rate. Among blacks, this rate is 5705 per 10,000 individuals; among non-blacks, it is 2892 per 10,000 individuals. ED visits tend to be for acute conditions; only 12 percent of ED visits are for chronic illnesses, relative to 57 percent of hospitalizations for this age group.

These descriptive statistics highlight the importance of using a large dataset to investigate utilization in these age groups. With a per capita hospitalization rate of under 0.03 in 1999, and under 0.04 in 2009, it would be very difficult to detect changes in utilization rates among the relevant cohorts using, for example, survey data. By employing large administrative datasets, we will be able to credibly investigate changes in hospitalizations and ED visits even though overall usage rates in these age groups are low.

V. Empirical Strategy

To estimate the impact of childhood Medicaid eligibility on later life hospitalization and ED visits, we use a regression discontinuity approach and compare outcomes for cohorts born just before and after the September 30, 1983 birthdate cutoff. We rely on both a parametric specification (e.g., polynomial) and a local linear regression to estimate the discontinuity in outcomes at the birth date cutoff point. These complementary methods offer tradeoffs in terms of bias and variance and are presented together to assess the stability of results (Lee and Lemieux 2010). We use the log number of hospitalizations or ED visits as the dependent variable, which assumes that population trends smoothly across birth month cohorts.[11] Estimates of the RD are interpreted as the proportionate change in the rate of hospitalizations or ED visits.

We first estimate a second-order polynomial regression model that uses observations from monthly cohorts born within a 4-year window of the cutoff date. Each cohort born between October 1979 and September 1987 is denoted using the integer values $\in [-48,47]$, where $c=0$ for the first cohort born after the cutoff (October 1983). The regression specification is given by

$$\log(y_c) = \alpha + \beta D_c + \gamma_0 c + \gamma_1 c^2 + \gamma_2 D_c \cdot c + \gamma_3 D_c \cdot c^2 + \delta_m M_c + \varepsilon_c \qquad (1)$$

where y_c represents the number of hospitalizations or ED visits for a given birth cohort and D_c is an indicator for cohorts born after September 30, 1983 ($c \geq 0$). We include a quadratic function in birth month cohort c that is allowed to differ on both sides of the cutoff point by including an interaction term for those cohorts born after the cutoff. In addition, we include calendar month dummies M_c to control for variation in outcomes related to the link between timing of birth and family characteristics (Buckles and Hungerman 2012). The inclusion of these dummies will also net out the effects of policies that may differentially affect cohorts born in certain months (for example, school entry dates). In the Appendix, we explore alternative functional form specifications, as well as window sizes used in the estimation, and find results consistent with

[11] Wherry and Meyer (2014) find evidence of a decrease in mortality at ages 15-18 resulting from the Medicaid expansions for black children born after the cutoff. Without adjusting for the corresponding change in the underlying population count at age 25, this biases us against detecting a decrease in later life hospitalizations or ED visits.

those presented in our main specification. For each outcome, we present visual evidence by plotting the residual for each birth month cohort from a regression on the set of calendar month dummies.

We also use local linear regression to estimate the discontinuity in outcomes at the cutoff point. The estimation is conducted with a triangular kernel and we employ the optimal bandwidth selector procedure proposed by Imbens and Kalyanaraman (2012). We report confidence intervals that were constructed using the variance estimator developed by Calonico, Cattaneo, and Titiunik (2014), which offers robustness to bandwidth choice. In the Appendix, we show that our results are not sensitive to the choice of bandwidth.

These methods estimate the effect of the eligibility expansions averaged across the full sample of children at the cutoff. This is an example of a "fuzzy" RD design because factors other than date of birth determine eligibility for and take-up of public health insurance. Although we do not have information in our data on whether individuals were eligible for or took up public health insurance, we are able to examine outcomes for certain subsamples that were more likely to be affected by the change in Medicaid policy. In particular, we examine outcomes separately by race and income in accord with the variation in treatment documented previously.

We also investigate differences in outcomes by state of residence. States' eligibility criteria in place prior to the expansions, as well as state decisions regarding optional Medicaid and CHIP expansions, led to variation in the size of the gain in Medicaid eligibility for children born after September 30, 1983. We exploit this variation and estimate changes in outcomes associated with the relative size of the discontinuity in childhood eligibility in each state.

We estimate the following specification:

$$\log(y_{cs}) = \alpha + \beta_0 D_{cs} \cdot G_s + \beta_1 D_{cs} + \gamma_{0s} c_s + \gamma_{1s} c^2{}_s + \gamma_{2s} D_{cs} \cdot c_s + \gamma_{3s} D_{cs} \cdot c^2{}_s + \delta_s + \delta_m M_c + \varepsilon_{cs} \tag{2}$$

where we regress the log of a given state-cohort outcome y_{cs} on an indicator for cohorts born after the cutoff D_{cs} and its interaction with a measure of the size of the discontinuity in each state in eligibility-years G_s. In addition to including state and calendar month of birth fixed effects, we also include second order polynomial trends in birth month cohort and allow these trends to vary by state and differ on either side of the discontinuity. Due to the small number of states in our sample, we use the percentile-t bootstrap method with 999 bootstrap repetitions clustered by state for hypothesis testing and constructing confidence intervals. This method has been shown to perform well even when there are relatively few clusters (see Cameron, Gelbach and Miller 2008).

Some state by birth month cohort cells have zero hospitalizations for blacks; if this is the case for any birth month cohort in a state, we drop that entire state when conducting the analysis for the black and non-black subsamples. This leads us to drop 2 states in our analysis. Sensitivity analyses that run the outcomes in levels and using the full sample of states provide similar results.

VI. Results

Figure 4 presents the profile of log hospitalizations by birth month cohort in 1999, when the cohorts born just on either side of the cutoff are approximately 15 years of age. As seen in the figure, hospitalizations are correlated with age (i.e. birth month), which is the running variable in the RD estimation. As noted earlier, we address the possibility of different trends on either side of the September 30, 1983 cutoff by allowing the polynomial in birth month cohort to have different coefficients on either side of the cutoff. Visually, the figure reveals little evidence of a discontinuity in outcomes at the September 30, 1983 threshold. Estimates of the discontinuity from the regression analysis reported in Table 5 support this conclusion. While the point estimates tend to be negative and suggestive of small declines in hospitalizations among those gaining childhood Medicaid coverage, none are statistically significant. There is no clear evidence of an effect of childhood Medicaid eligibility on rates of hospitalization in this year for any race or visit category; however, we note that the confidence intervals are large and do not rule out that there may have been meaningful changes in utilization.

Figure 5 displays hospitalization outcomes in 2009 when cohorts born around the cutoff were approximately 25 years old. The top panel of Table 6 presents the corresponding discontinuity estimates. Among all races, we find only marginally significant evidence of a reduction in hospitalizations for those born after the birthdate cutoff and only when using local linear regression. However, for blacks, there is a notable drop in hospitalizations visible at the cutoff. The regression analysis indicates a reduction in hospitalizations of either 8 or 13 percent for those cohorts born just after the September 30, 1983 date, depending on the specification, though the 95 percent confidence interval includes values that are much smaller for the global specification. . Furthermore, there is a strongly significant decline in hospitalizations related to chronic illness. Our estimates indicate declines on the order of 13 or 17 percent with confidence intervals ruling out declines smaller than 3 percent. For hospitalizations related to non-chronic illness, the estimated decline is smaller at 2 or 7 percent and only significant when using local linear regression. We do not find evidence of a similar improvement for non-blacks.

The bottom panel of Table 6 presents similar results for emergency department visits, also depicted in Figure 6. We find a significant reduction in rates of ED visits of between 3-4 percent among black cohorts born immediately after the birth date cutoff. When we examine ED visits by their relation to chronic illness, we again find evidence of a sizeable decline in visits related to chronic illness (12 or 16 percent in the two specifications), although the estimate is only marginally significant in the global regression model. There is also some evidence of a smaller decline in ED visits related to non-chronic illness on the order of 1-3 percent. Finally, we find some evidence of a slight increase in ED visits among all races (1%) and non-blacks (2%) that appears to be driven by visits related to non-chronic illness, but the estimates are not consistently significant across specifications and are close to zero under the global regression model.

Assuming similar effect sizes and hospitalization rates across other states, our point estimates imply that there were either 2,703 or 4,359 fewer inpatient hospitalizations among black cohorts born during the first year after the cutoff at age 25.[12] The change in the probability of gaining eligibility across the birth date cutoff was about 17 percentage points for blacks, who gained approximately 4.8 additional years of Medicaid eligibility as a result. If we assume that the reduction in hospitalizations observed in 2009 is entirely a result of the eligibility expansion, the point estimates from our two specifications imply that there were 2.6 or 4.2 fewer hospitalizations in 2009 for every 100 black children who were made eligible for (on average) 4.8 additional years of Medicaid eligibility as a result of the expansions. This reduction for

[12] Using the Census Estimate that in 2009 there were 617,000 blacks age 25, and that the average hospitalization rate at age 25 for blacks was 547.7 per 10,000 (Table 4).

eligibles is large relative to the average rate of hospitalization among all 25-year-old blacks, representing 47 and 76 percent fewer hospitalizations relative to that average. However, because the children that were affected by these expansions were in poor households, and because the poor tend to be in worse health than the general population, it is likely that their baseline hospitalization rates would be higher than that of a typical black 25 year old. Case, Lubotsky and Paxson (2002) find that children from low-income families have worse health in childhood, and that the differences between children raised in low- and high-income families become more pronounced as the children grow older and enter adulthood. Overall, while the point estimates are somewhat large, they are not implausible. In addition, the 95% confidence intervals on our estimates allow for the possibility of smaller effect sizes.

We can further scale these estimates by take-up rates to arrive at the effect of Medicaid coverage, rather than Medicaid eligibility, on hospitalizations later in life. However, because parents tend to enroll children when they become ill or injured (Marton and Yelowitz 2014), this calculation would describe the local effect on (most likely) the sickest children who may benefit the most from medical intervention. This may present an overly optimistic view of what Medicaid coverage expansions can accomplish for the average Medicaid-eligible child. Nonetheless, we perform such a calculation; it implies that for every 100 black children who enrolled in Medicaid there were 7.6 or 12.4 fewer hospitalizations at age 25.[13]

Similarly, our point estimates imply that in 2009 there were either 10,561 or 14,081 fewer emergency department visits experienced by blacks born the first year after the cutoff. Again assuming this reduction is driven entirely by the eligibility expansion, this estimate implies that there were 10 or 13 fewer emergency department visits in 2009 for every 100 black children made eligible as a result of the expansions. Comparing this to average ED use in the population of blacks suggests that gaining an average of 4.8 additional years of Medicaid eligibility in childhood lowers emergency department use by 18 or 23 percent at age 25. However, as we noted previously, baseline ED use among adults who grew up in low-income families is likely higher than average ED use in the population. When scaled by our estimate of take-up, this implies that there were either 29.4 or 38.2 fewer ED visits for every 100 black children newly enrolled in Medicaid.

[13] We use the 34 percent take-up rate estimated in Table 3.

Low-income Zipcodes

Next we examine changes in hospitalizations and ED visits in 2009 for patients from low-income zipcodes (Table 7 and Figures 7-8).[14] In low-income zipcodes, we find a large and statistically significant reduction in total hospitalizations of 15 or 21 percent among black cohorts born just after September 30, 1983. The decline appears to be concentrated among hospitalizations related to chronic illness, where we see a reduction of 22 or 29 percent. There is no significant reduction in hospitalizations for non-chronic illnesses.

Similarly, we find significant evidence of a decrease in ED visits (14 percent) related to chronic illness for blacks, although the visual evidence is less clear. There is some evidence for reductions in total and non-chronic illness related ED visits as well, but the point estimates are not as consistent in size or significance.

Finally, we find no evidence of a reduction in hospitalizations or ED visits from patients of all races or non-blacks. There is some evidence of an increase in total and non-chronic illness related hospitalizations among non-blacks, but the estimates are only statistically significant when parametric methods are used.

Heterogeneity by State

The Federal expansions that affected cohorts around the discontinuity had differential effects across states based on the generosity of each state's Medicaid program prior to the expansions. As a result, the discontinuity in the cumulative number of Medicaid-eligible years at the cutoff birthdate varies by state. For example, in Maryland, being born immediately after the birth date cutoff would result in 1.68 additional expected years of Medicaid eligibility, whereas being born after the birth date cutoff in California would result in only 0.02 additional years of Medicaid eligibility (see Table 2). In this section, we use these differences across states as an additional source of variation in our analysis of the effect of childhood Medicaid coverage on adult utilization. If our observed changes in utilization across the threshold are indeed driven by differences in Medicaid eligibility in childhood, we might expect the discontinuity in utilization

[14] The HCUP hospitalization data includes information on median income of the patient's zip code in 2009 only, so we are unable to conduct this analysis with the 1999 data.

at the cutoff birth date to be larger in states where the change in eligibility across that threshold is greater.

To explore this heterogeneity, we estimate the model described by equation (2). The variable G_s is the estimated size of the discontinuity in eligibility that occurs at the cutoff in state *s* (as reported in Table 2). We report the coefficients on the discontinuity, $D_{cs,}$ and the interaction between D_{cs} and the size of the discontinuity in the number of Medicaid-eligible years in state *s*. The coefficient on the interaction term measures the effect of an additional year of eligibility on the outcome variable. In addition to reporting the coefficients on these terms (β_1 and β_2), we also report the marginal effect of being born after the cutoff birthdate for states with "small" (25^{th} percentile) and "large" (75^{th} percentile) discontinuities in columns 3 and 4. Because the distribution of the discontinuity varies by race, the point at which these marginal effects are measured also varies across the sample stratifications by race; the size of the discontinuity in states with small and large discontinuities is listed in the top row. The confidence intervals of these estimates are also constructed using a clustered percentile-t bootstrap procedure.

Table 8 presents the results using hospitalizations in 1999, when birth cohorts born around the cutoff birth date are 15 years old. As in our original specification, we do not find systematic evidence that those born immediately after the cutoff had fewer hospitalizations at this age. We do find that non-black patients born immediately after the cutoff had fewer hospitalizations if they lived in a state with a large discontinuity, although the effect is small (about 1 percent). We also find a statistically significant decline in chronic illness hospitalizations for patients of all races living in states with small discontinuities.

Using data from 2009, we find that those born immediately after the cutoff had fewer hospitalizations than those born before the cutoff, and that this decrease was most pronounced for individuals living in states with large discontinuities. These results are reported in Table 9. For patients of all races, we find that the effect on total hospitalizations is significantly larger as the size of the discontinuity grows. The coefficient on the interaction term indicates that increasing eligibility by one year in childhood lowers hospitalizations at age 25 by 4.6 percent. We find a statistically significant reduction in total hospitalizations for those born immediately after the cutoff of about 2.5 percent for states with large discontinuities, but no significant effect in states with small discontinuities. The point estimate of the coefficient on the interaction term suggests a similar pattern for chronic and non-chronic illness hospitalizations, although the effect is not statistically significant.

The second panel presents results for black patients. We find a similar pattern for black patients as we find for patients of all races; the point estimate of the interaction term indicates that the effects of being born after the threshold are larger in states with a greater discontinuity, although the interaction term is only statistically significant in the model for non-chronic hospitalizations. In states with large discontinuities, we find that being born immediately after the birth date cutoff is associated with statistically significantly fewer total hospitalizations (a 14.5 percent decline) and hospitalizations for chronic illnesses (a 17 percent decline).[15] We find no statistically significant effects of being born after the cutoff for blacks in states with small discontinuities.

The third panel presents results for non-black patients. We find no statistically significant coefficients on the interaction term for this subsample and the sign of the coefficient on the interaction term varies across outcomes. We do find a significant increase in hospitalizations for non-chronic illnesses in states with small discontinuities. Overall, there is no consistent pattern for this subsample.

In Table 10 we perform a similar analysis using emergency department data. In the top panel we report the results for all races. We find a significant decline in ED visits for cohorts born after the cutoff in states with a large discontinuity in Medicaid eligibility. Being born after the cutoff is associated with 3.9 percent fewer emergency department visits in states with large changes in eligibility across the threshold and this effect is significantly different from zero at the 1 percent level. In states with small discontinuities in eligibility, being born after the birth date cutoff is associated with 1.5 percent fewer emergency department visits and this effect is not statistically significant. Although we do not find statistically significant coefficients on the interaction terms for chronic and non-chronic emergency department visits, our point estimates indicate that the decrease associated with being born after the birth date cutoff is larger in states with greater discontinuities in eligibility. We find a statistically significant reduction in chronic illness related

[15] In this model, the marginal effect at the 75th percentile is statistically different from zero but the interaction term itself is not. This result occurs because the covariance between $\widehat{\beta_0}$ and $\widehat{\beta_1}$ is negative.

ED visits for cohorts born after September 30, 1983 in states with large discontinuities but no significant effect in states with small discontinuities.

Conducting this analysis for blacks, we find a similar pattern of point estimates suggesting larger declines in ED visits associated with a greater average gain in Medicaid eligibility. However, our estimates for the interaction term are not statistically significant. In the third panel we conduct this analysis for non-blacks. Here we find a significant negative effect on the interaction term for all emergency department visits. Our estimates indicate that being born after the cutoff for non-blacks in states with large discontinuities is associated with 3.8 percent fewer emergency department visits.

VII. Sensitivity Analyses

We conduct several sensitivity analyses. First, we estimate the effect of discontinuous Medicaid eligibility on two types of hospitalizations that are unlikely to be affected by medical intervention in childhood: hospitalizations for appendicitis and injury. Second, we estimate placebo effects using birth month cohorts born between January, 1965 and September, 1983 who did not actually experience a discontinuity in Medicaid eligibility. Finally, we explore the sensitivity of our results to the inclusion of cohort-specific characteristics including measures of health at birth.

Hospitalizations for Acute Conditions

We first present estimates using hospitalizations for appendicitis and injuries for all patients and the low-income sample in 2009 by race group. Both appendicitis and injury are acute conditions that should not be sensitive to medical care received in the past. For that reason, we believe it is unlikely that coverage in childhood could plausibly influence hospitalizations for these conditions. If we find effects on these types of hospitalizations, it may indicate that the assumptions of our RD model are incorrect.

The results of these analyses are reported in Table 11. The first panel shows the results for hospitalizations in 2009 for patients from all zip codes stratified by race. The second panel shows similar results for patients from low-income zip codes. In both panels, we find point estimates that are close to zero, none of which are statistically significant for any race group. Although the confidence intervals are large, we note that the point estimates are smaller in magnitude than those reported for all hospitalizations and chronic illness related hospitalizations and that the direction of the estimates is not consistent, with roughly half of the specifications reporting small, statistically insignificant positive effects and half reporting similarly sized negative effects. Overall, this suggests there was little impact of the policy on these types of visits. This result is consistent with our expectation that these types of visits should not be affected by access to medical care in childhood.

Checks for Discontinuities at Non-Discontinuity Points

We conduct a second type of placebo test using data on cohorts born prior to the actual eligibility cutoff. We place an artificial "cut off" date in the center of each four year window (eight year period) beginning with cohorts born in January 1965. Our final placebo "cut off" is placed at September 1979, so that the last month used in the estimation of these placebo effects is September 1983, immediately before the actual change in Medicaid eligibility occurs. This results in 129 "placebo effects" estimated at birth dates where no discontinuity existed.

Using these placebo estimates, we construct histograms, which we report in Figures 9-12. The effect estimated at the "true" cutoff is shown on the figure as a black vertical line. The first row of Figure 9 shows the distribution of placebo statistics for the models of the total number of hospitalizations in 2009. Consistent with our results in the previous section, the true estimate for all races and non-blacks is not large relative to the placebo estimates. However, the estimated discontinuity at September 30, 1983 among black patients is large relative to the placebo estimates: it exceeds all but 7 (5.4 percent) of the placebo estimates in absolute value.

The second row of Figure 9 shows the distribution of placebo estimates for chronic illness related hospitalizations. The effect for blacks estimated at the September 30, 1983 cutoff is larger in absolute value than all but 5 (3.9 percent) of the placebo estimates in the chronic illness model. The effect for all races and non-blacks falls near the middle of the placebo distribution for chronic illness hospitalizations.

Figure 10 reports the distribution of placebo effects for hospitalizations of low-income patients in 2009. Among patients of all races, the effect estimated at the true cutoff is near the center of the distribution of placebo effects. For black patients from low-income zip codes, however, the true effect is substantially larger in magnitude than all placebo estimates; this is true for all hospitalizations and chronic illness hospitalizations. For non-black patients, we find that being born after the birth date cutoff is positive and this estimate is larger in absolute value than all but two of the placebo tests conducted for hospitalizations. This result is counter-intuitive because it suggests that non-black cohorts from low-income neighborhoods experienced more hospitalizations in adulthood as a result of being made eligible for Medicaid. However, we note that the effect estimated at the true discontinuity for this group is smaller and not statistically significant when estimated using the local linear regression model.

In Figures 11 and 12 we present similar distributions for placebo tests using emergency department data. Here, our results conform less to the original inference conducted in Tables 6 and 7. While we find that the effect for blacks estimated at the September 30, 1983 cutoff is larger than most placebo estimates, more than 5 percent of placebo estimates exceed the coefficient estimated at the true discontinuity in all models. The effect for blacks exceeds all but 9 percent of placebo estimates in the model of all ED visits and 12 percent of placebo estimates in the model of chronic illness related ED visits. For the low-income sample, the effect for blacks estimated at the September 30, 1983 cutoff exceeds all but 15 percent of placebo estimates in the model of all ED visits and 12 percent of placebo estimates in the model of chronic illness related ED visits.

Overall, the placebo simulations conducted in this section strongly confirm the effect of the discontinuity on hospitalizations among black cohorts and provide particularly convincing evidence supporting our results for the low-income subsample of blacks. The placebo tests are somewhat less convincing when applied to the emergency department results; although the effect of Medicaid on emergency department visits estimated at the true discontinuity exceeds the majority of placebo estimates, a sizeable minority (between 9 and 15 percent) of placebo estimates are larger in absolute value than the "true" effect.

Inclusion of Cohort-Specific Characteristics

Finally, we examine the sensitivity of our estimates to the inclusion of several cohort-specific characteristics drawn from the National Vital Statistics System Birth Data files for 1979 to 1987.

We include controls for the following birth outcomes: the fraction of mothers with at least a high school education, fraction of mothers married, and fraction of mothers receiving any prenatal care; the incidence of low birth weight and very low birth weight births; and the number of births. Reported in the Appendix, the results are robust to the inclusion of these covariates.

VIII. Was the Upfront Cost of the Medicaid Expansions Offset by Lower Utilization Later in Life?

The results presented in this paper provide evidence that expanding public health insurance coverage to children lowers future health care costs by improving health and reducing later life hospital and emergency department use among those who gain eligibility. In this section, we provide "back of the envelope" calculations on the magnitude of these cost savings relative to the upfront cost of expanding Medicaid.

Our point estimates indicate that blacks born immediately after the birth date cutoff had between 8 and 12.9 percent fewer hospitalizations at age 25. Assuming similar effect sizes and hospitalization rates across other states, this would imply that there were between 2,703 and 4,359 fewer hospitalizations at this age among black cohorts born during the first year after the cutoff and who benefited from the Medicaid expansions.[16] Average charges for a hospitalization of a patient this age in 2009 were $18,855. However, while charges indicate the amount billed by hospitals for services, they do not necessarily represent the actual cost of, or payment for, a hospital encounter. HCUP provides hospital-specific "cost to charge" ratios designed to estimate the resource cost of a hospital visit using data from accounting reports collected by the Centers for Medicare and Medicaid Services. We apply these ratios to deflate hospital charges and find that a typical hospitalization for a patient of this age costs the hospital about $5,692. We therefore estimate the total one-year cost savings of these expansions from fewer later life

[16] These calculations use the Census Bureau estimate that there were 617,000 blacks age 25 in 2009.

hospitalizations at between \$5692 x 2703 = \$15.4 million and \$5692 x 4359 = \$24.8 million for these cohorts.[17]

Performing similar calculations for emergency department visits, we estimate that the expansions reduced emergency department visits at age 25 among black cohorts born during the year following the cutoff by between 10,561 and 15,137 visits. Emergency department charges for this age are about \$1,697 per visit; however, HCUP calculates the resource cost of such a visit is approximately \$611. We therefore estimate that the expansions saved between \$6.5 million and \$9.2 million in the cost of emergency department visits for these cohorts by lowering later life utilization.

Average spending per child enrolled in Medicaid in 1991 was \$902 per child (in 1991 dollars) (Congressional Research Services (1993)).[18] Multiplying this amount by the average gain in years enrolled in Medicaid using information from Tables 1 and 3 and assuming a 3 percent discount rate, this implies the total cost of the eligibility expansions for all children born during the year following the September 30, 1983 cutoff was approximately \$652 million dollars in 2009.[19] The cost offsets from childhood Medicaid expansions, totaling between \$15.4 and \$24.8 million at age 25, therefore represent between 3 and 5 percent of the total cost of the expansions. If the reduction in utilization we observe at age 25 persists for several years, the cost offsets associated with these expansions will be even larger.

These calculations indicate that the long-run cost savings from the Medicaid expansions may be quite substantial. Considering the other research on the long run effects of these expansions on other outcomes, the true cost offsets of the Medicaid expansions might be larger still. Specifically, our estimates do not incorporate other benefits to government (such as higher

---

[17] To calculate actual cost-savings under the expansions, we would need to know the marginal cost of the hospitalizations that were prevented. However, without this information, the average cost provides some idea of the potential savings to the Medicaid program. As hospitalizations for chronic illnesses are more costly on average than the average hospitalization, our calculations here likely provide a lower bound on the true cost savings.

[18] Because most (78 percent) of our variation in the discontinuity is a result of the Omnibus Reconciliation Act of 1990 that was implemented in 1991, we use 1991 as our base year in these calculations.

[19] The Census estimate for the total number of 25-year-olds in 2009 is 4,264,000. We multiply this estimate by the 0.42 year average gain in childhood eligibility, the 23.8 percent take-up rate, and the \$902 cost per year of enrollment per child in 1991 to arrive at our estimate.

income tax receipt and lower earned income tax payments, as found in Brown, Kowalski, and Lurie 2014) or to beneficiaries, such as better education outcomes (Cohodes et al. 2014) or lower mortality (Wherry and Meyer 2014, Brown, Kowalski, and Lurie 2014).

IX. Conclusion

Policies that expand public health insurance coverage tend to increase utilization and, thus, the total resources devoted to health care spending in the economy in the short term. However, there may be longer-term costs savings that do not materialize until later in life because of improved health. While these long-term cost savings are often cited in policy discussions and debates, very little credible evidence exists on the magnitude of these effects, or even if they are present at all. This is a crucial gap in our understanding of the role of public health insurance coverage as these cost offsets potentially represent a substantial, but previously unaccounted-for, benefit of such programs.

In this paper, we provide evidence of such effects by exploiting a discontinuity in the number of years a child is eligible for Medicaid based on his or her date of birth. Because several of the early Medicaid coverage expansions to poor children applied only to children born after September 30, 1983, children born immediately before this cutoff received fewer years of Medicaid eligibility throughout childhood. Among blacks, who were most likely to be affected by these expansions, we find that those born immediately after the cutoff had a significant reduction in hospitalizations and emergency department visits at age 25 compared to those born immediately before the cutoff. The effect is particularly pronounced for chronic illness related hospitalizations and ED visits, among patients in low-income neighborhoods, and in states where the size of the eligibility discontinuity was large. A back of the envelope calculation based on our point estimates suggests that these reductions in utilization for the cohort born one year after the birth date cutoff offset between 3 and 5 percent of the total cost of the expansions we study.

References

Baicker, Katherine, Sarah L. Taubman, Heidi L. Allen, Mira Bernstein, Jonathan H. Gruber, Joseph P. Newhouse, Eric C. Schneider, Bill J. Wright, Alan M. Zaslavsky, and Amy N. Finkelstein. 2013. "The Oregon Experiment—Effects of Medicaid on Clinical Outcomes." *The New England Journal of Medicine* 368:1713—1722.

Boudreaux, Michel H., Ezra Golberstein, and Donna McAlpine. 2014. "The Long-Term Impacts of Medicaid Exposure in Early Childhood: Evidence from the Program's Origin." Working paper, University of Minnesota.

Brown, David, Amanda E. Kowalski, and Ithai Z. Lurie. 2014. "Medicaid as an Investment in Children: What is the Long-Term Impact on Tax Receipts?" Working paper, Yale University. Downloaded from http://www.econ.yale.edu/~ak669/medicaid.latest.draft.pdf

Calonico, Sebastian, Matias D. Cattaneo, and Rocio Titiunik. 2014. "Robust Nonparametric Confidence Intervals for Regression-Discontinuity Designs." *Econometrica*, forthcoming.

Card, David and Lara Shore-Sheppard. 2004. "Using Discontinuous Eligibility Rules to Identify the Effects of the Federal Medicaid Expansions." *Review of Economics and Statistics* 86(3): 752-766.

Case, Anne, Darren Lubotsky and Christina Paxson. 2002. "Economic Status and Health in Childhood: The Origins of the Gradient." *American Economic Review* 92 (5):1308-1334.

Cohodes, Sarah, Daniel Grossman, Samuel Kleiner, and Michael F. Lovenheim. 2014. The Effect of Child Health Insurance Access on Schooling: Evidence from Public Health Insurance Expansions. *NBER Working Paper #20178*.

Congressional Research Service. 1988. *Medicaid Source Book: Background Data and Analysis*. Prepared for the use of the Subcommittee on Health and the Environment of the Committee on Energy and Commerce, U.S. House of Representatives. Washington, DC: Government Printing Office.

Congressional Research Service. 1993. *Medicaid Source Book: Background Data and Analysis (A 1993 Update)*. Prepared for the use of the Subcommittee on Health and the Environment of the Committee on Energy and Commerce, U.S. House of Representatives. Washington, DC: Government Printing Office.

Currie, Janet, Sandra Decker and Wanchuan Lin. 2008. "Has Public Health Insurance for Older Children Reduced Disparities in Access to Care and Health Outcomes?" *Journal of Health Economics* 27: 1567-1581.

Currie, Janet and Jonathan Gruber. 1995. "Health Insurance Eligibility, Utilization of Medical Care, and Child Health." *The Quarterly Journal of Economics* 111(2): 431-466.

Currie, Janet and Jonathan Gruber. 1996a. "Health Insurance Eligibility, Utilization of Medical Care, and Child Health." *The Quarterly Journal of Economics* 111(2): 431-466.

Currie, Janet and Jonathan Gruber. 1996b. "Saving Babies: The Efficacy and Cost of Recent Changes in the Medicaid Eligibility of Pregnant Women." *The Journal of Political Economy* 104(6): 1263-1296.

Cutler, David M. and Jonathan Gruber. 1996. "Does Public Insurance Crowd Out Private Insurance." *The Quarterly Journal of Economics* 111(2): 391-430.

Dafny, L. and J. Gruber. 2005. "Public Insurance and Child Hospitalizations: Access and Efficiency Effects." *Journal of Public Economics* 89: 109-129.

Dave, Dhaval M., Sandra Decker, Robert Kaestner, and Kosali I. Simon. 2008. "Re-examining the Effects of the Medicaid Expansions for Pregnant Women." *NBER Working Paper # 14591*.

De la Mata, Dolores. 2012. "The Effect of Medicaid Eligibility on Coverage, Utilization, and Children's Health." *Health Economics* 21(9): 1061-1079.

Decker, Sandra L., Douglas Almond, and Kosali I. Simon. 2015. "The Impact of Medicaid's Introduction on the Use of Health Care and Health Outcomes." Working Paper.

Finkelstein, Amy, Sarah Taubman, Bill Wright, Mira Bernstein, Jonathan Gruber, Joseph P. Newhouse, Heidi Allen, Katherine Baicker and the Oregon Health Study Group. 2012. "The Oregon Health Insurance Experiment: Evidence from the First Year." *Quarterly Journal of Economics*, 127(3): 1057-1106.

Freeman, Howard E. and Bradford L. Kirkman-Liff. 1985. "Health Care Under AHCCCS: An Examination of Arizona's Alternative to Medicaid." *Health Services Research* 20(3): 245-266.

Goodman-Bacon, Andrew. 2014. "Public Insurance and Mortality: Evidence from Medicaid Implementation." Working paper, University of California, Berkeley, ed(October 2014).

Healthcare Cost and Utilization Project (HCUP). 2014. *SID Description of Data Elements*. Accessed on December 1, 2014: http://www.hcup-us.ahrq.gov/db/state/siddist/sid_multivar.jsp

Howell, Embry, Sandra Decker, Sara Hogan, Alshadye Yemane and Jonay Fostser. 2010. "Declining Child Mortality and Continuing Racial Disparities in the Era of the Medicaid and SCHIP Insurance Coverage Expansions." *American Journal of Public Health* 100(12): 2500-2506.

Howell, Embry M. and Genevieve M. 2012. "The Impact of the Medicaid/CHIP Expansions on Children: A Synthesis of the Evidence." *Medical Care Research and Review* 20(10): 1-25.

Lee, David S. and Thomas Lemieux. 2010. "Regression Discontinuity Designs in Economics." *Journal of Economic Literature* 48: 281-355.

Marton, James and Aaron Yelowitz. 2014. "Health Insurance Generosity and Conditional Coverage: Evidence from Medicaid Managed Care in Kentucky." *Working Paper, Andrew Young School of Policy Studies, Georgia State University*.

Miller, Sarah and Laura Wherry. 2014. "The Long-Term Health Effects of Early Life Medicaid Coverage." Working paper, University of Michigan.

Racine, Andrew D., Robert Kaestner, Theodore J. Joyce and Gregory J. Colman. 2001. "Differential Impact of Recent Medicaid Expansions by Race and Ethnicity." *Pediatrics* 108: 1135-1142.

U.S. General Accounting Office. 1989. *Medicaid: States Expand Coverage for Pregnant Women, Infants, and Children*. GAO/HRD-89-90. Washington, DC: U.S. General Accounting Office.

Wherry, Laura R. and Bruce D. Meyer. 2014. "Saving Teens: Using a Policy Discontinuity to Estimate the Effects of Medicaid Eligibility." Working Paper, December. Earlier version: National Bureau of Economics Working Paper No. 18309, August 2012.

Table 1. Childhood Medicaid Eligibility Gain for Children Born in October vs. September 1983 by Race Group

	Percent Gaining Eligibility	Average Gain (in Years) for Children Gaining Eligibility	Average Gain (in Years) for Total Child Population
All Races	9.24	4.53	0.42
Blacks	17.13	4.81	0.82
Non-Blacks	7.98	4.48	0.36

Notes: Weighted averages calculated from simulation of lifetime eligibility if born in September vs. October 1983 for a national sample of children ages 0-17 in the 1981-1988 March CPS. See Wherry and Meyer (2014) for more information on estimation of childhood eligibility at cutoff.

Table 2. Average Childhood Medicaid Eligibility Gain for Children Born in October vs. September 1983 by State

	All Races	Blacks	Non-Blacks
Arizona	0.63	0.78	0.61
California	0.02	0.04	0.02
Iowa	0.38	0.73	0.34
Maryland	1.67	2.38	1.55
New Jersey	0.36	0.67	0.31
New York	0.13	0.23	0.11
Oregon	0.35	0.47	0.31
Texas	0.79	1.42	0.67
Wisconsin	0.38	0.67	0.33

Notes: Weighted averages calculated from simulation of lifetime eligibility if born in September vs. October 1983 for a national sample of children ages 0-17 in the 1981-1988 March CPS. State eligibility gains are estimated using the national sample characteristics. See Wherry and Meyer (2014) for more information on estimation of childhood eligibility at cutoff.

Table 3. Childhood Medicaid Coverage at Ages 8-13, NHIS

	All Races		Blacks		Non-Blacks	
	local	global	local	global	local	global
Medicaid	0.022**	0.012	0.058***	0.053**	0.005	0.002
	[0.005, 0.070]	[-0.004, 0.028]	[0.050, 0.269]	[0.002, 0.104]	[-0.018, 0.039]	[-0.014, 0.018]
	bw=16534		*bw=3747*		*bw=16619*	
Number of Observations:	60119		10059		50060	

	Households in Poverty		Households not in Poverty	
	local	global	local	global
Medicaid	0.076	0.066**	0.001	0.007
	[-0.150, 0.322]	[0.011, 0.120]	[-0.011, 0.038]	[-0.002, 0.015]
	bw=5455		*bw=12432*	
Number of Observations:	11619		43588	

Notes: Data from 1992-1996 National Health Insurance Survey Health Insurance Supplements. All global regression models include birth month fixed effects and a quadratic function in birth month cohort interacted with an indicator that the birth month cohort is October 1983 or later. For local linear regression models, the optimum bandwidth was selected using the Imbens and Kalyanaraman bandwidth selector and is reported in italics. Robust 95% confidence intervals reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1

Table 4. Rates of Hospital and ED Utilization (per 10,000) for 15 and 25-Year-Olds

	Rate for 15-Year-Olds in 1999			**Rates for 25-Year-Olds in 2009**					
	Hospitalizations			Hospitalizations			Emergency Department Visits		
	All Races	Black	Non-Black	All Races	Black	Non-Black	All Races	Black	Non-Black
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Total visits (excluding pregnancy)	264.96	322.96	258.02	329.21	547.73	302.88	3167.11	5705.53	2892.26
By Relation to Chronic Illness									
visits related to chronic illness	139.80	192.81	133.46	156.98	313.84	138.09	386.75	794.82	342.57
visits not related to chronic illness	125.17	130.15	124.57	172.22	233.89	164.79	2780.36	4910.70	2549.70

Notes: Data for inpatient hospitalizations from states: AZ, CA, IA, OR, TX, and WI (1999 and 2009), as well as MD, NJ, and NY (2009 only). Data for emergency department visits from states: AZ, CA, IA, NJ, NY, WI. Rates were calculated using age-specific population estimates by race for these states from the 2009 American Community Survey and the 2000 Census 1% sample downloaded from IPUMS. Hospitalizations and ED visits exclude those related to pregnancy and delivery.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 5. Estimates of Effect of Childhood Medicaid Eligibility on Hospitalizations at Age 15 (1999)

	All Races		Blacks		Non-Blacks	
	local	global	local	global	local	global
Log Total Hospitalizations (excluding pregnancy)	-0.032	-0.004	-0.029	-0.030	-0.037*	0.022
	[-0.071, 0.007]	[-0.042, 0.034]	[-0.236, 0.171]	[-0.141, 0.081]	[-0.100, 0.005]	[-0.029, 0.073]
	bw=27		*bw=30*		*bw=32*	
By Relation to Chronic Illness						
log hospitalizations related to chronic illness	-0.018	-0.026	-0.155	-0.044	-0.037	-0.012
	[-0.068, 0.030]	[-0.083, 0.031]	[-0.409, 0.363]	[-0.210, 0.122]	[-0.156, 0.099]	[-0.094, 0.071]
	bw=24		*bw=31*		*bw=33*	
log hospitalizations related to non-chronic illness	-0.049	0.008	-0.007	-0.001	-0.041	0.045
	[-0.118, 0.010]	[-0.036, 0.051]	[-0.080, 0.100]	[-0.095, 0.093]	[-0.146, 0.030]	[-0.024, 0.113]
	bw=26		*bw=20*		*bw=28*	

Notes: Sample includes 96 birth-month observations and draws data from AZ, CA, IA, OR, NY, TX, WI. All global regression models include birth month fixed effects and a quadratic function in birth month cohort interacted with an indicator that the birth month cohort is October 1983 or later. For local linear regression models, the optimum bandwidth was selected using the Imbens and Kalyanaraman bandwidth selector and is reported in italics. Robust 95% confidence intervals reported in brackets.

*** p<0.01, ** p<0.05, * p<0.1

Table 6. Estimates of Effect of Childhood Medicaid Eligibility on Hospitalizations at Age 25 (2009)

	All Races		**Blacks**		**Non-Blacks**	
	local	global	local	global	local	global
Log Total Hospitalizations (excluding pregnancy)	-0.011*	0.001	-0.129***	-0.080**	0.016	0.019
	[-0.010, 0.003]	[-0.024, 0.026]	[-0.224, -0.082]	[-0.156, -0.004]	[-0.029, 0.049]	[-0.008, 0.047]
	bw=36		*bw=19*		*bw=24*	
By Relation to Chronic Illness						
log hospitalizations related to chronic illness	-0.017*	-0.001	-0.168***	-0.128***	0.019	0.034
	[-0.224, 0.003]	[-0.045, 0.044]	[-0.282, -0.103]	[-0.221, -0.034]	[-0.114, 0.040]	[-0.031, 0.045]
	bw=38		*bw=20*		*bw=26*	
log hospitalizations related to non-chronic illness	-0.006	0.003	-0.071**	-0.018	0.009	0.007
	[-0.049, 0.054]	[-0.027, 0.032]	[-0.196, -0.009]	[-0.114, 0.078]	[-0.020, 0.083]	[-0.031, 0.045]
	bw=36		*bw=20*		*bw=24*	
Log Total ED Visits	0.011*	-0.005	-0.030**	-0.043**	0.024**	0.004
	[-0.003, 0.050]	[-0.024, 0.014]	[-0.060, -0.002]	[-0.077, -0.009]	[0.006, 0.060]	[-0.016, 0.023]
	bw=21		*bw=23*		*bw=16*	
By Relation to Chronic Illness						
log ED visits related to chronic illness	-0.012	-0.016	-0.158***	-0.115*	0.018	0.010
	[-0.097, 0.032]	[-0.052, 0.021]	[-0.330, -0.065]	[-0.232, 0.001]	[-0.025, 0.042]	[-0.021, 0.040]
	bw=42		*bw=20*		*bw=21*	
log ED visits related to non-chronic illness	0.017**	-0.003	-0.012	-0.031**	0.025**	0.003
	[0.004, 0.061]	[-0.023, 0.017]	[-0.031,0.058]	[-0.061, -0.001]	[0.007, 0.070]	[-0.018, 0.024]
	bw=19		*bw=31*		*bw=17*	

Notes: Sample includes 96 birth-month observations and draws hospitalization data from AZ, CA, IA, MD, NJ, NY, OR, TX, WI and ED data from AZ, CA, IA, NJ, NY, WI. All global regression models include birth month fixed effects and a quadratic function in birth month cohort interacted with an indicator that the birth month cohort is October 1983 or later. For local linear regression models, the optimum bandwidth was selected using the Imbens and Kalyanaraman bandwidth selector and is reported in italics. Robust 95% confidence intervals reported in brackets.
*** p<0.01, ** p<0.05, * p<0.1

Table 7. Estimates of Effect of Childhood Medicaid Eligibility on Hospitalizations in Low-Income Zipcodes at Age 25 (2009)

	All Races		**Blacks**		**Non-Blacks**	
	local	global	local	global	local	global
Log Total Hospitalizations (excluding pregnancy)	-0.013	0.009	-0.206***	-0.153**	0.050	0.071**
	[-0.077, 0.059]	[-0.042, 0.061]	[-0.352, -0.105]	[-0.277, -0.029]	[-0.054, 0.201]	[0.008, 0.133]
	bw=26		*bw=20*		*bw=32*	
By Relation to Chronic Illness						
log hospitalizations related to chronic illness	-0.076	-0.037	-0.289***	-0.217***	0.019	0.049
	[-0.200, 0.021]	[-0.116, 0.042]	[-0.491, -0.150]	[-0.365, -0.069]	[-0.168, 0.158]	[-0.047, 0.145]
	bw=23		*bw=20*		*bw=26*	
log hospitalizations related to non-chronic illness	0.037	0.056*	-0.073	-0.057	0.077	0.090**
	[-0.076, 0.269]	[-0.008, 0.120]	[-0.234, 0.030]	[-0.209, 0.095]	[-0.019, 0.274]	[0.009, 0.171]
	bw=42		*bw=24*		*bw=31*	
Log Total ED Visits	-0.017	-0.028	-0.043	-0.069***	-0.008	-0.011
	[-0.053, 0.086]	[-0.063, 0.007]	[-0.107, 0.030]	[-0.118, -0.020]	[-0.040, 0.069]	[-0.054, 0.032]
	bw=37		*bw=31*		*bw=38*	
By Relation to Chronic Illness						
log ED visits related to chronic illness	-0.044	-0.040	-0.140***	-0.142**	-0.003	0.012
	[-0.121, 0.020]	[-0.105, 0.025]	[-0.355, -0.071]	[-0.283, -0.002]	[-0.097, 0.130]	[-0.065, 0.090]
	bw=24		*bw=22*		*bw=40*	
log ED visits related to non-chronic illness	-0.017	-0.026	-0.033	-0.056**	-0.009	-0.014
	[-0.038, 0.138]	[-0.064, 0.011]	[-0.059, 0.130]	[-0.100, -0.013]	[-0.057, 0.034]	[-0.059, 0.031]
	bw=35		*bw=34*		*bw=37*	

Notes: Sample includes 96 birth-month observations and draws hospitalization data from AZ, CA, IA, MD, NJ, NY, OR, TX, WI and ED data from AZ, CA, IA, NJ, NY, WI. All global regression models include birth month fixed effects and a quadratic function in birth month cohort interacted with an indicator that the birth month cohort is October 1983 or later. For local linear regression models, the optimum bandwidth was selected using the Imbens and Kalyanaraman bandwidth selector and is reported in italics. Robust 95% confidence intervals reported in brackets.

Table 8. Estimates of Effect of State Childhood Medicaid Eligibility Gain on Hospitalizations At Age 15 By Race

	Post (1)	Post x Size of Discontinuity (2)	Marginal effect at 25th percentile (3)	Marginal effect at 75th percentile (4)	N
All Races			*Discontinuity Size: 0.35 Years*	*Discontinuity Size: 0.63 Years*	
Log Total Hospitalizations (excluding pregnancy)	-0.012	0.040	0.003	0.014	672
	[-0.043, 0.026]	[-0.654, 0.316]	[-0.051, 0.033]	[-0.227, 0.219]	
By Relation to Chronic Illness					
log hospitalizations related to chronic illness	-0.055	0.088	-0.024**	0.001	672
	[-0.114, 0.060]	[-2.221, 0.387]	[-0.074, -0.0002]	[-0.310, 0.166]	
log hospitalizations related to non-chronic illness	0.027	-0.012	0.023	0.020	672
	[-0.026, 0.051]	[-0.320, 0.724]	[-0.045, 0.085]	[-0.234, 0.535]	
Blacks			*Discontinuity Size: 0.23 Years*	*Discontinuity Size: 1.42 Years*	
Log Total Hospitalizations (excluding pregnancy)	-0.022	-0.045	-0.032	-0.085	384
	[-4.810, 4.059]	[-0.204, 5.998]	[-0.780, 0.125]	[-0.172, 0.029]	
By Relation to Chronic Illness					
log hospitalizations related to chronic illness	-0.105	0.007	-0.103	-0.095	384
	[-4.089, 3.435]	[-2.790, 0.140]	[-0.201, 0.125]	[-7.916, 0.375]	
log hospitalizations related to non-chronic illness	0.118	-0.151	0.083	-0.096	382
	[-67.521, 3.977]	[-3.257, 34.945]	[-1.424, 1.696]	[-28.58, 4.387]	
Non Blacks			*Discontinuity Size: 0.11 Years*	*Discontinuity Size: 0.67 Years*	
Log Total Hospitalizations (excluding pregnancy)	0.003	-0.020	0.001	-0.010**	384
	[-0.640, 0.142]	[-0.055, 0.070]	[-0.026, 0.018]	[-0.017, -0.002]	
By Relation to Chronic Illness					
log hospitalizations related to chronic illness	-0.022*	0.015	-0.020	-0.012	384
	[-0.037, 0.012]	[-0.011, 0.043]	[-0.039, 0.111]	[-0.029, 0.006]	
log hospitalizations related to non-chronic illness	0.019	-0.081	0.010	-0.035	384
	[-2.587, 3.627]	[-0.156, 0.060]	[-0.016, 0.071]	[-0.077, 0.026]	

Notes: 1999 hospitalization data are from AZ, CA, IA, OR, NY, TX, and WI. .In addition to the indicator for cohorts born after the cutoff and its interaction with the size of the discontinuity, regressions also include state fixed effects and state-specific quadratic functions in birth month cohort that are interacted with the indicator for cohorts born after the cutoff. Percentile-t 95% confidence intervals are reported in brackets and percentile-t p-values were used for hypothesis testing. *** p<0.01, ** p<0.05, * p<0.1

Table 9. Estimates of Effect of State Childhood Medicaid Eligibility Gain on Hospitalizations At Age 25 By Race

	Post	Post x Size of Discontinuity	Marginal effect at 25th percentile	Marginal effect at 75th percentile	N
All Races			*Discontinuity Size: 0.35 Years*	*Discontinuity Size: 0.63 Years*	
Log Total Hospitalizations (excluding pregnancy)	0.004	-0.046**	-0.012	-0.025**	864
	[-0.033, 0.062]	[-0.092, -0.006]	[-0.037, 0.018]	[-0.041, -0.006]	
By Relation to Chronic Illness					
log hospitalizations related to chronic illness	0.030	-0.051	0.012	-0.003	864
	[-0.054, 0.179]	[-0.157, 0.013]	[-0.038, 0.100]	[-0.033, 0.044]	
log hospitalizations related to non-chronic illness	-0.008	-0.048	-0.025	-0.039	864
	[-0.053, 0.217]	[-0.177, -0.002]	[-0.062, 0.132]	[-0.071, 0.057]	
Blacks			*Discontinuity Size: 0.23 Years*	*Discontinuity Size: 1.42 Years*	
Log Total Hospitalizations (excluding pregnancy)	-0.057	-0.062	-0.071	-0.145***	672
	[-0.242, 0.222]	[-0.204, 0.084]	[-0.221, 0.164]	[-0.226, -0.024]	
By Relation to Chronic Illness					
log hospitalizations related to chronic illness	-0.055	-0.082	-0.073	-0.171**	672
	[-0.591, 0.983]	[-0.429, 0.213]	[-0.450, 0.323]	[-0.231, -0.003]	
log hospitalizations related to non-chronic illness	0.026	-0.064**	0.011	-0.065	672
	[-0.033, 0.143]	[-0.116, -0.029]	[-0.046, 0.104]	[-0.160, 0.014]	
Non Blacks			*Discontinuity Size: 0.11 Years*	*Discontinuity Size: 0.67 Years*	
Log Total Hospitalizations (excluding pregnancy)	0.020	-0.010	0.018	0.013	672
	[-0.018, 0.172]	[-0.249, 0.052]	[-0.015, 0.139]	[-0.069, 0.057]	
By Relation to Chronic Illness					
log hospitalizations related to chronic illness	0.011	0.027	0.014	0.029	672
	[-0.090, 0.140]	[-0.126, 0.125]	[-0.076, 0.123]	[-0.100, 0.068]	
log hospitalizations related to non-chronic illness	0.034**	-0.053	0.028**	-0.002	672
	[0.003, 0.101]	[-0.226, 0.043]	[0.001, 0.092]	[-0.138, 0.050]	

Notes: 2009 hospitalization data are from AZ, CA, IA, MD, OR, NJ, NY, TX, and WI. In addition to the indicator for cohorts born after the cutoff and its interaction with the size of the discontinuity, regressions also include state fixed effects and state-specific quadratic functions in birth month cohort that are interacted with the indicator for cohorts born after the cutoff. Percentile-t 95% confidence intervals are reported in brackets and percentile-t p-values were used for hypothesis testing.

*** p<0.01, ** p<0.05, * p<0.1

Table 10. Effect of State Childhood Medicaid Eligibility Gain on ED Visits At Age 25 By Race

	Post	Post x Size of Discontinuity	Marginal effect at 25th percentile	Marginal effect at 75th percentile	N
All Races			*Discontinuity Size: 0.35 Years*	*Discontinuity Size: 0.63 Years*	
Log Total ED Visits in 2009	0.014	-0.084**	-0.015	-0.039***	576
	[-0.027, 0.048]	[-0.177, -0.025]	[-0.033, 0.005]	[-0.070, -0.019]	
By Relation to Chronic Illness					
log ED visits related to chronic illness	-0.0003	-0.128	-0.045	-0.081**	
	[-0.052, 0.074]	[-0.738, 0.005]	[-0.129, 0.017]	[-0.440, -0.004]	576
log ED visits related to non-chronic illness	0.015	-0.079*	-0.012	-0.034	
	[-0.032, 0.037]	[-0.842, 0.142]	[-0.031, 0.036]	[-0.070, 0.053]	
Blacks			*Discontinuity Size: 0.23 Years*	*Discontinuity Size: 1.42 Years*	
Log Total ED Visits in 2009	-0.017	-0.068	-0.032	-0.113	576
	[-0.106, 0.091]	[-0.738, 0.219]	[-0.123, 0.005]	[-0.894, 0.264]	
By Relation to Chronic Illness					
log ED visits related to chronic illness	-0.102	-0.021	-0.107***	-0.133	576
	[-0.308, 0.070]	[-0.802, 0.742]	[-0.171, -0.034]	[-0.931, 0.792]	
log ED visits related to non-chronic illness	-0.003	-0.072	-0.019	-0.105	576
	[-0.180, 0.213]	[-0.462, 0.318]	[-0.208, 0.061]	[-0.558, 0.277]	
Non Blacks			*Discontinuity Size: 0.11 Years*	*Discontinuity Size: 0.67 Years*	
Log Total ED Visits in 2009	0.022*	-0.089**	0.012	-0.038***	576
	[-0.017, 0.038]	[-0.215, -0.031]	[-0.024, 0.028]	[-0.091, -0.023]	
By Relation to Chronic Illness					
log ED visits related to chronic illness	0.023	-0.132	0.008	-0.066	576
	[-0.085, 0.268]	[-1.134, 0.048]	[-0.116, 0.080]	[-0.860, 0.079]	
log ED visits related to non-chronic illness	0.022*	-0.083	0.012*	-0.034	576
	[-0.013, 0.044]	[-0.365, 0.001]	[-0.010, 0.030]	[-0.080, 0.003]	

Notes: 2009 ED data are from AZ, CA, IA, NJ, NY, WI. In addition to the indicator for cohorts born after the cutoff and its interaction with the size of the discontinuity, regressions also include state fixed effects and state-specific quadratic functions in birth month cohort that are interacted with the indicator for cohorts born after the cutoff. Percentile-t 95% confidence intervals are reported in brackets and percentile-t p-values were used for hypothesis testing.
*** p<0.01, ** p<0.05, * p<0.1

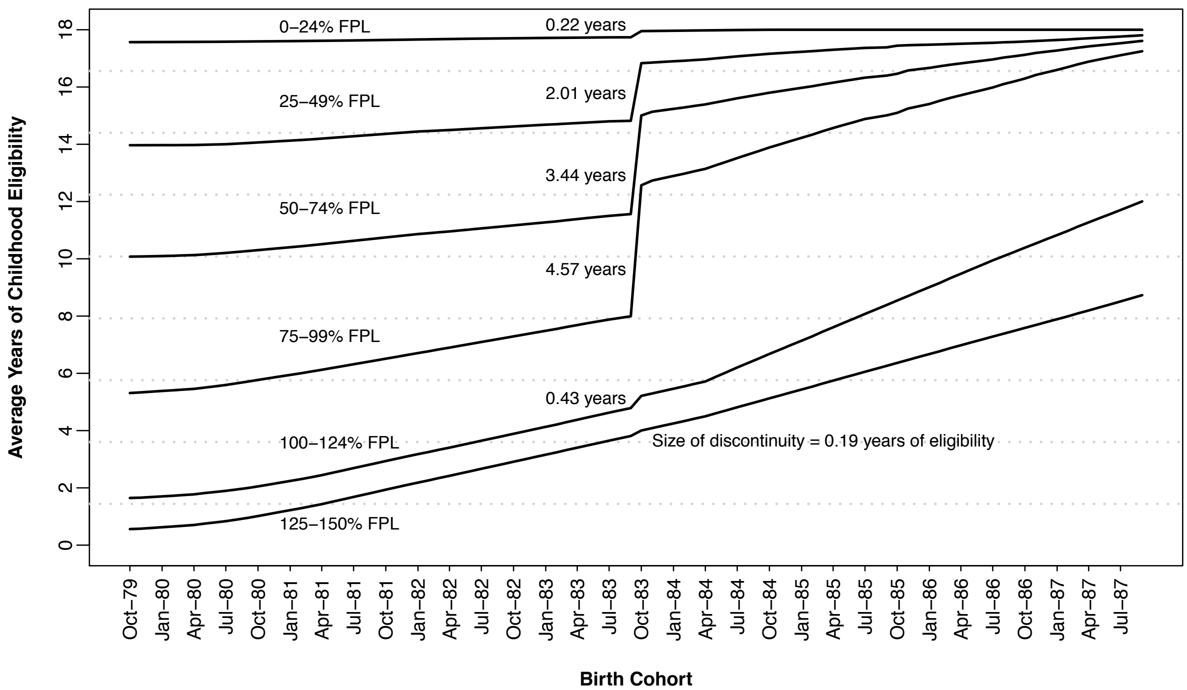
Table 11. "Placebo" tests for 2009 Hospitalizations (Injuries and Appendicitis)

	All Races		Blacks		Non-Blacks	
	local	global	local	global	local	global
Injury	-0.002	-0.015	-0.050	-0.033	0.001	-0.014
	[0.286, 0.314]	[-0.101, 0.075]	[-0.362, 0.128]	[-0.176, 0.111]	[-0.195, 0.332]	[-0.096, 0.069]
	bw=42		*bw=25*		*bw=39*	
Appendicitis	-0.015	-0.013	0.065	0.070	-0.021	-0.013
	[-0.133, 0.095]	[-0.092, 0.063]	[-0.526, 0.575]	[-0.466, 0.607]	[-0.144, 0.112]	[-0.113, 0.087]
	bw=26		*bw=26*		*bw=29*	
Low Income Sample	**All Races**		**Blacks**		**Non-Blacks**	
	local	global	local	global	local	global
Injury	0.001	-0.042	-0.044	-0.054	0.056	-0.030
	[-0.210, 0.321]	[-0.180, 0.097]	[-0.411, 0.265]	[-0.281, 0.172]	[-0.182, 0.378]	[-0.195, 0.134]
	bw=26		*bw=27*		*bw=22*	
Appendicitis	-0.023	0.097	0.012	0.087	-0.030	0.090
	[-0.345, 0.073]	[-0.092, 0.286]	[-0.921, 0.855]	[-.628, 0.802]	[-0.344, 0.094]	[-0.104, 0.283]
	bw=22		*bw=31*		*bw=22*	

States included: AZ, CA, IA, MD, NJ, NY, OR, TX, WI. All global regression models include birth month fixed effects and a quadratic function in birth month cohort interacted with an indicator that the birth month cohort is October 1983 or later. For local linear regression models, the optimum bandwidth was selected using the Imbens and Kalyanaraman bandwidth selector and is reported in italics. Robust 95% confidence intervals reported in brackets.

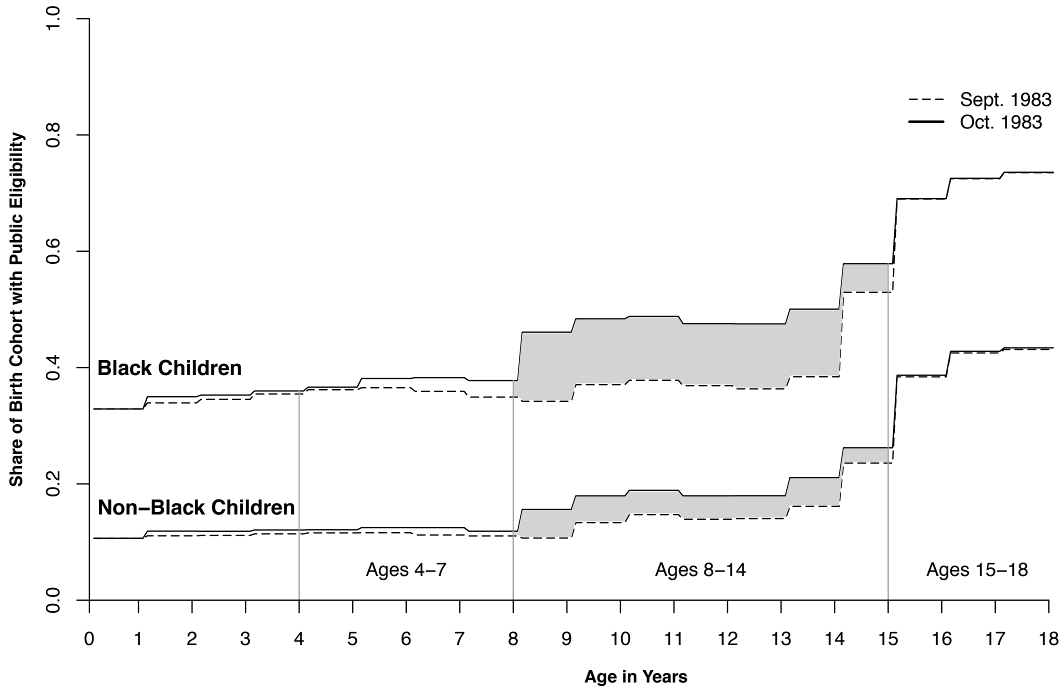
*** p<0.01, ** p<0.05, * p<0.1

Figure 1. Average Years of Childhood Eligibility for Medicaid/SCHIP by Birth Cohort and Family Income (%FPL)



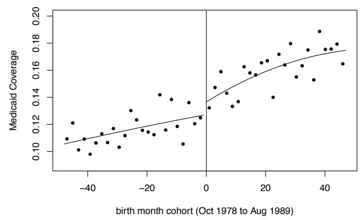
Source: Wherry and Meyer (2014).

Figure 2. Average Public Eligibility at Each Age of Childhood by Birth Month Cohort and Child Race

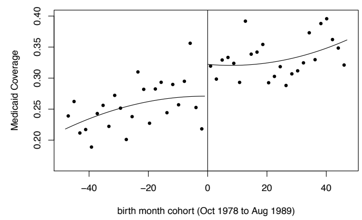


Notes: Weighted average calculated using the characteristics and state of residence of a sample of black or non−black children of ages 0−17 in the 1981−1988 March CPS. See Wherry and Meyer (2014) for additional information.

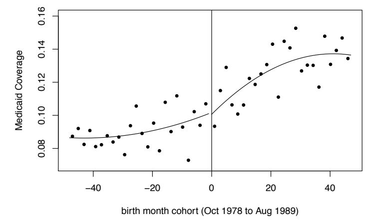
Figure 3: Medicaid Coverage in Childhood, Ages 8 to 13, NHIS



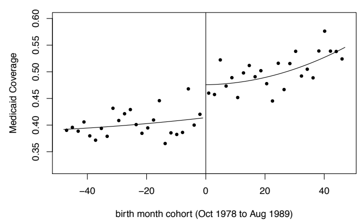
(a) All Races



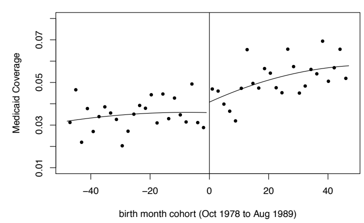
(b) Blacks



(c) Non-Blacks



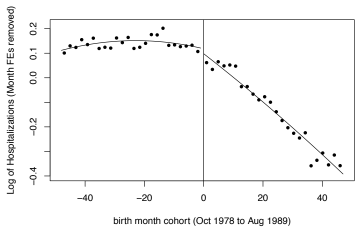
(d) Households below poverty level



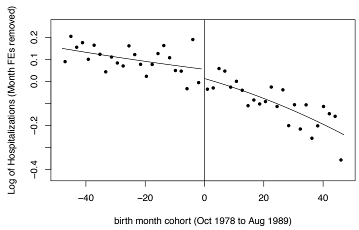
(e) Households above poverty level

Source: Authors' calculations from the National Health Interview Survey, 1992-1996. Cohorts born in 1983 are between the ages of 8 and 13 in these figures. The trend is estimated using children between the ages of 4 and 17.

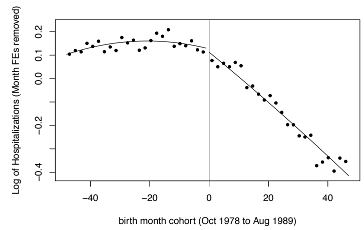
Figure 4: 1999 Hospitalizations, Calendar Month of Birth Fixed Effects Removed



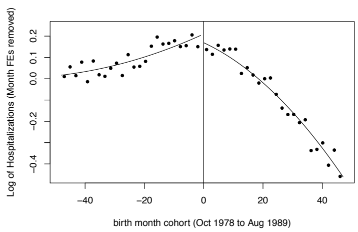
(a) All Hospitalizations, All Races



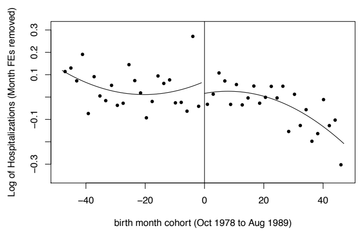
(b) All Hospitalizations, Blacks



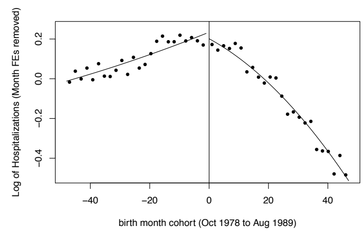
(c) All Hospitalizations, Non-Blacks



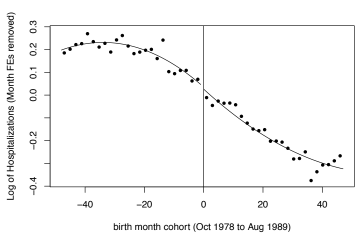
(d) Chronic Illness Hospitalizations, All Races



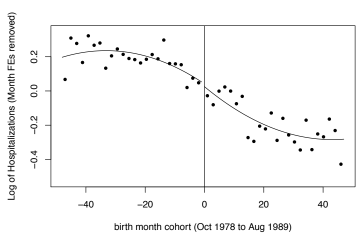
(e) Chronic Illness Hospitalizations, Blacks



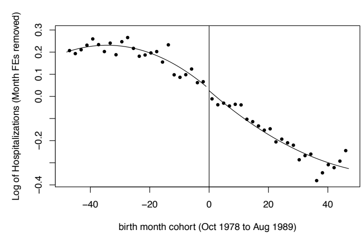
(f) Chronic Illness Hospitalizations, Non-Blacks



(g) Non-chronic Illness Hospitalizations, All Races



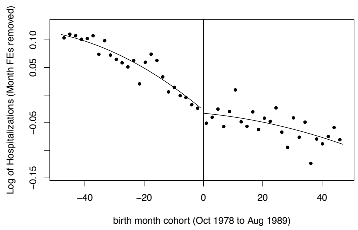
(h) Non-chronic Illness Hospitalizations, Blacks



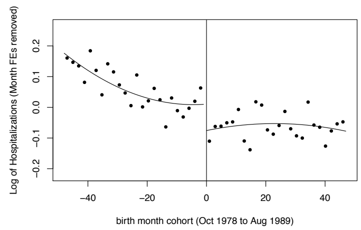
(i) Non-chronic Illness Hospitalizations, Non-Blacks

Figures plot the residuals from a regression of the log of hospitalizations by birth month on calendar month of birth fixed effects. Results presented using two-month bins. These models use data on all hospitalizations that occurred in 1999 in AZ, CA, IA, NY, OR, TX, and WI.

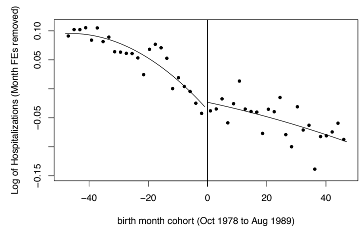
Figure 5: 2009 Hospitalizations, Calendar Month of Birth Fixed Effects Removed



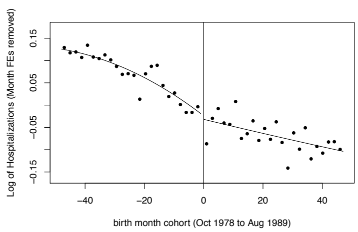
(a) All Hospitalizations, All Races



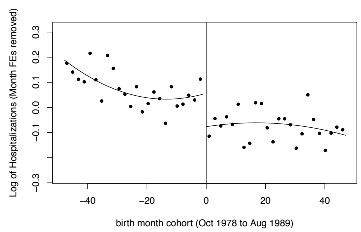
(b) All Hospitalizations, Blacks



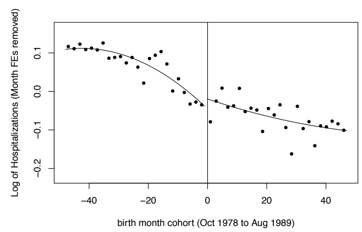
(c) All Hospitalizations, Non-Blacks



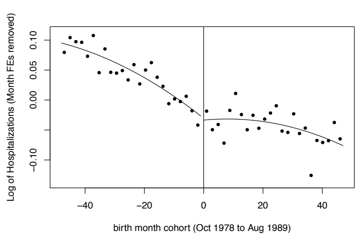
(d) Chronic Illness Hospitalizations, All Races



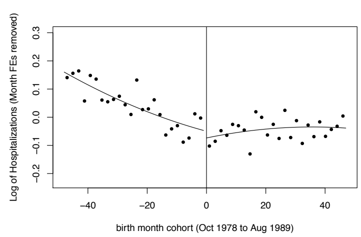
(e) Chronic Illness Hospitalizations, Blacks



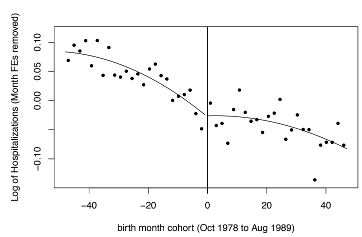
(f) Chronic Illness Hospitalizations, Non-Blacks



(g) Non-chronic Illness Hospitalizations, All Races



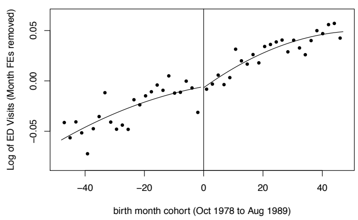
(h) Non-chronic Illness Hospitalizations, Blacks



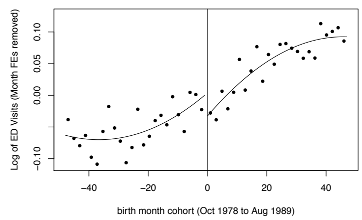
(i) Non-chronic Illness Hospitalizations, Non-Blacks

Figures plot the residuals from a regression of the log of hospitalizations by birth month on calendar month of birth fixed effects. Results presented using two-month bins. These models use data on all hospitalizations that occurred in 2009 in AZ, CA, IA, MD, NJ, NY, OR, TX, and WI.

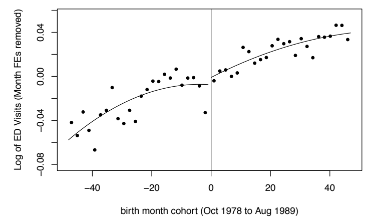
Figure 6: 2009 Emergency Department Visits, Calendar Month of Birth Fixed Effects Removed



(a) All ED Visits, All Races

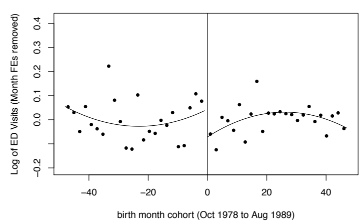


(b) All ED Visits, Blacks

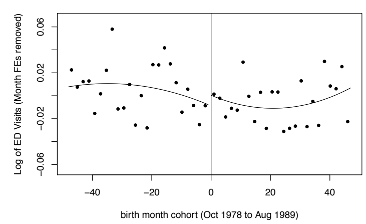


(c) All ED Visits, Non-Blacks

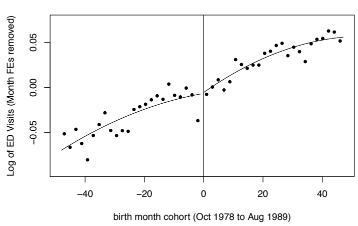
(d) Chronic Illness ED Visits, All Races



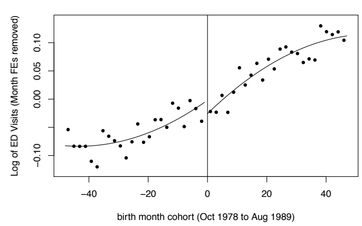
(e) Chronic Illness ED Visits, Blacks



(f) Chronic Illness ED Visits, Non-Blacks



(g) Non-Chronic Illness ED Visits, All Races

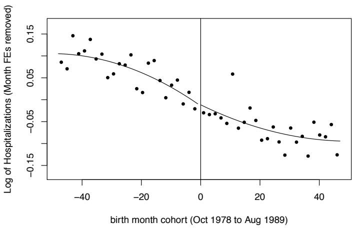


(h) Non-Chronic Illness ED Visits, Blacks

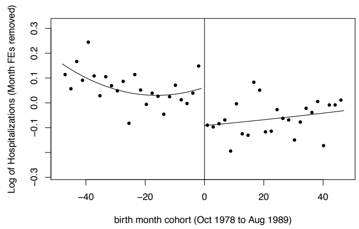
(i) Non-Chronic Illness ED Visits, Non-Blacks

Figures plot the residuals from a regression of the log of hospitalizations by birth month on calendar month of birth fixed effects. Results presented using two-month bins. These models use data on all emergency department visits that occurred in 2009 in AZ, CA, IA, NJ, NY, and WI.

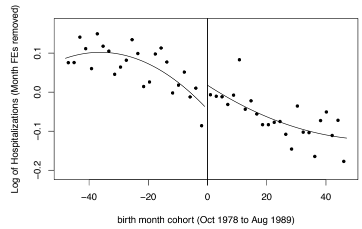
Figure 7: 2009 Hospitalizations, Patients from Low-Income Zipcodes, Calendar Month of Birth Fixed Effects Removed



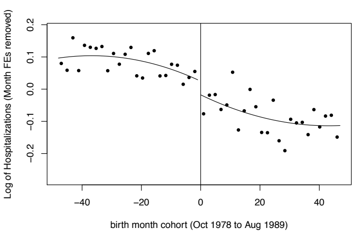
(a) All Hospitalizations, All Races



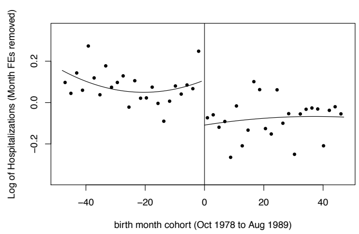
(b) All Hospitalizations, Blacks



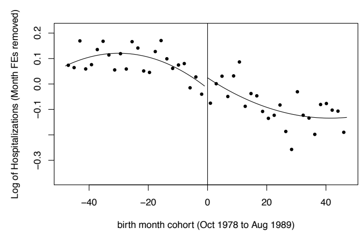
(c) All Hospitalizations, Non-Blacks



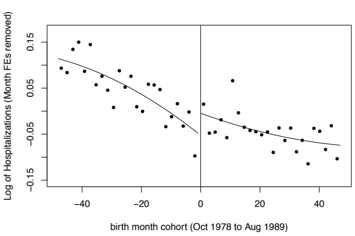
(d) Chronic Illness Hospitalizations, All Races



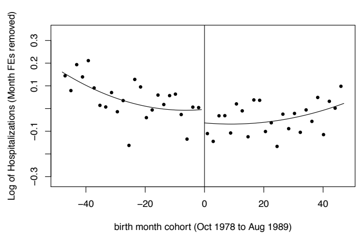
(e) Chronic Illness Hospitalizations in 2009, Blacks



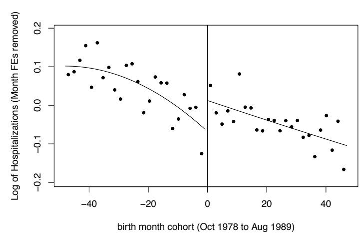
(f) Chronic Illness Hospitalizations in 2009, Non-Blacks



(g) Non-chronic Illness Hospitalizations, All Races



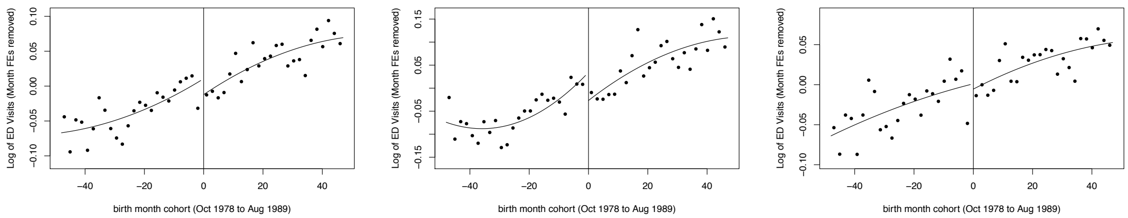
(h) Non-chronic Illness Hospitalizations in 2009, Blacks



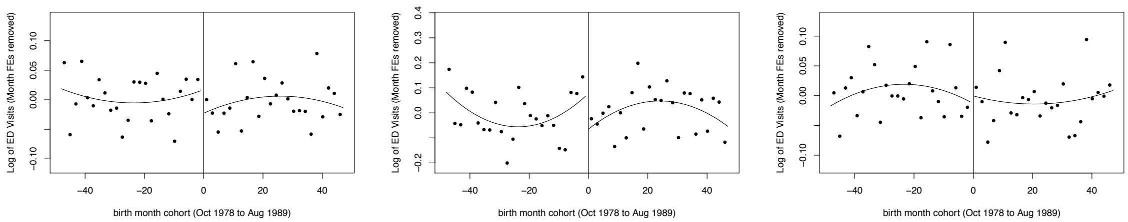
(i) Non-chronic Illness Hospitalizations in 2009, Non-Blacks

Figures plot the residuals from a regression of the log of hospitalizations by birth month on calendar month of birth fixed effects. Results presented using two-month bins. These models use data on all hospitalizations of patients from low-income zipcodes (zipcodes with median income lower than $39,999) that occurred in 2009 in AZ, CA, IA, MD, NJ, NY, OR, TX, and WI.

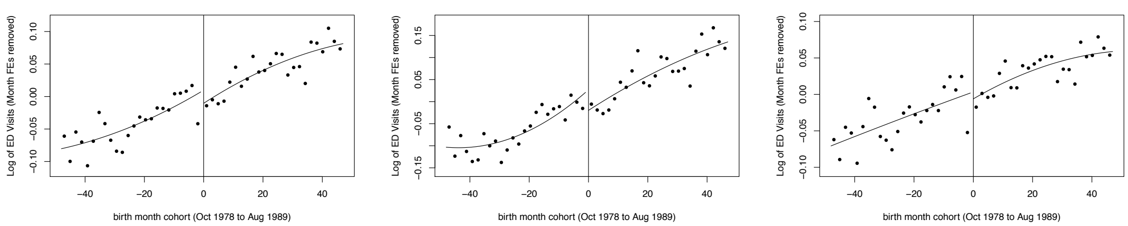
Figure 8: 2009 Emergency Department Visits by Patients from Low-Income Zipcodes, Calendar Month of Birth Fixed Effects Removed



(a) All ED Visits, All Races (b) All ED Visits, Blacks (c) All ED Visits, Non-Blacks



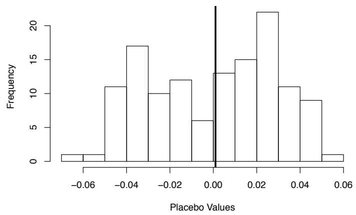
(d) Chronic Illness ED Visits, All Races (e) Chronic Illness ED Visits, Blacks (f) Chronic Illness ED Visits, Non-Blacks



(g) Non-chronic Illness ED Visits, All Races (h) Non-chronic Illness ED Visits, Blacks (i) Non-chronic Illness ED Visits, Non-Blacks

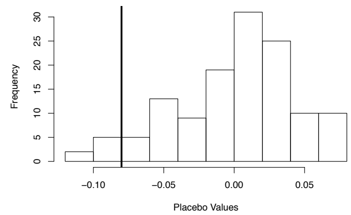
Figures plot the residuals from a regression of the log of hospitalizations by birth month on calendar month of birth fixed effects. Results presented using two-month bins. These models use data on all emergency department visits by patients from low-income zipcodes (zipcodes with median income below $39,999) that occurred in 2009 in AZ, IA, CA, NJ, NY, and WI.

Figure 9: 2009 Hospitalization Placebo Tests (Jan 1965 to Sep 1983 - 129 total placebo tests)



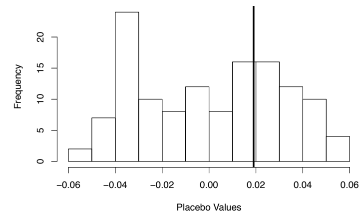
(a) All Hospitalizations, All Races

True effect is larger in magnitude than ≈ 0 % of placebo estimates



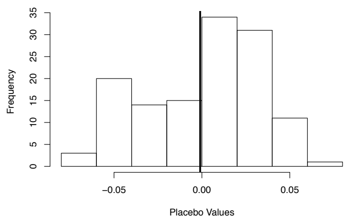
(b) All Hospitalizations, Blacks

True effect is larger in magnitude than ≈ 95 % of placebo estimates



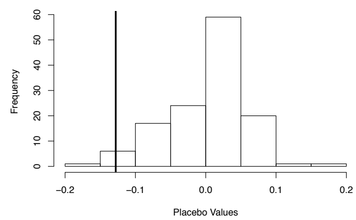
(c) All Hospitalizations, Non-Blacks

True effect is larger in magnitude than ≈ 33 % of placebo estimates



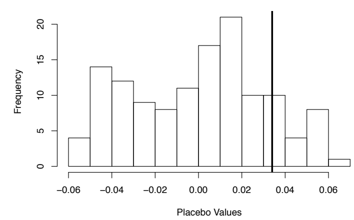
(d) Chronic Illness Hospitalizations, All Races

True effect is larger in magnitude than ≈ 4 % of placebo estimates



(e) Chronic Illness Hospitalizations, Blacks

True effect is larger in magnitude than ≈ 96 % of placebo estimates

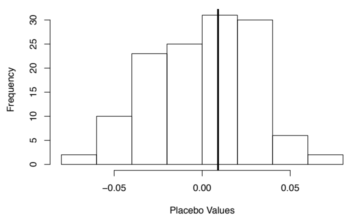


(f) Chronic Illness Hospitalizations, Non-Blacks

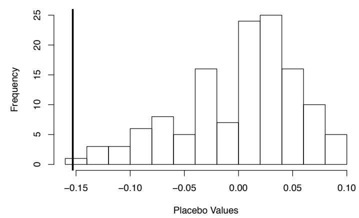
Larger in magnitude than ≈ 44 % of placebo estimates

These figures are histograms of the 129 "placebo" regression discontinuity estimates generated using every 4-year period between January 1965 and September 1983, prior to the actual discontinuity in Medicaid eligibility. The vertical line represents the effect estimated at the true Medicaid eligibility discontinuity.

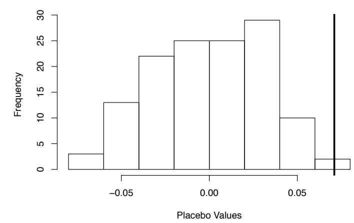
Figure 10: 2009 Hospitalization Placebo Tests (Jan 1965 to Sep 1983 - 129 total placebo tests), Patients from Low Income Zipcodes



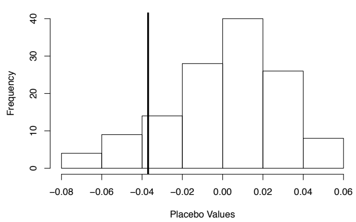
(a) All Hospitalizations, All Races
True effect is larger in magnitude than $\approx$ 22 % of placebo estimates



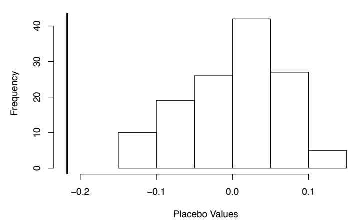
(b) All Hospitalizations, Blacks
True effect is larger in magnitude than $\approx$ 100 % of placebo estimates



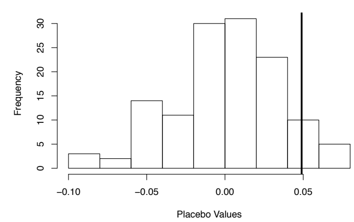
(c) All Hospitalizations, Non-Blacks
True effect is larger in magnitude than $\approx$ 99 % of placebo estimates



(d) Chronic Illness Hospitalizations, All Races
True effect is larger in magnitude than $\approx$ 82 % of placebo estimates



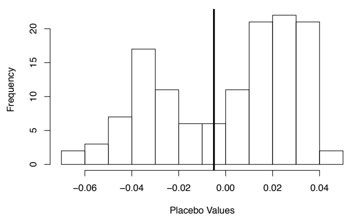
(e) Chronic Illness Hospitalizations in 2009, Blacks
True effect is larger in magnitude than $\approx$ 100 % of placebo estimates



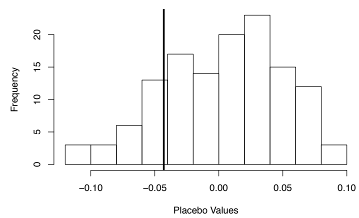
(f) Chronic Illness Hospitalizations in 2009, Non-Blacks
True effect is larger in magnitude than $\approx$ 87 % of placebo estimates

These figures are histograms of the 129 "placebo" regression discontinuity estimates generated using every 4-year period between January 1965 and September 1983, prior to the actual discontinuity in Medicaid eligibility. The vertical line represents the effect estimated at the true Medicaid eligibility discontinuity.

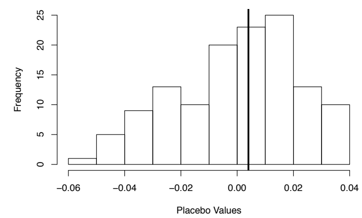
Figure 11: 2009 Emergency Department Visit Placebo Tests (Jan 1965 to Sep 1983 - 129 total placebo tests)



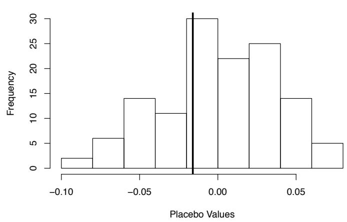
(a) All ED Visits, All Races
True effect is larger in magnitude than ≈ 5 % of placebo estimates



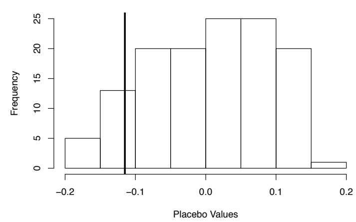
(b) All ED Visits, Blacks
True effect is larger in magnitude than ≈ 91 % of placebo estimates



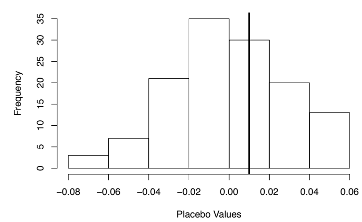
(c) All ED Visits, Non-Blacks
True effect is larger in magnitude than ≈ 5 % of placebo estimates



(d) Chronic Illness ED Visits, All Races
True effect is larger in magnitude than ≈ 24 % of placebo estimates



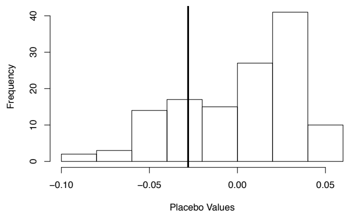
(e) Chronic Illness ED Visits in 2009, Blacks
True effect is larger in magnitude than ≈ 82 % of placebo estimates



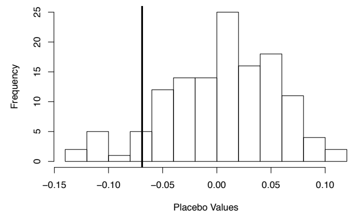
(f) Chronic Illness ED Visits in 2009, Non-Blacks
True effect is larger in magnitude than ≈ 30 % of placebo estimates

These figures are histograms of the 129 “placebo” regression discontinuity estimates generated using every 4-year period between January 1965 and September 1983, prior to the actual discontinuity in Medicaid eligibility. The vertical line represents the effect estimated at the true Medicaid eligibility discontinuity.

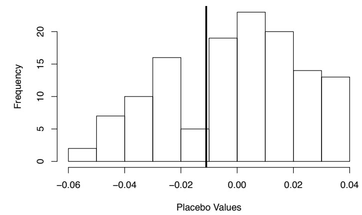
Figure 12: 2009 Emergency Department Visit Placebo Tests (Jan 1965 to Sep 1983 - 129 total placebo tests), Patients from Low Income Zipcodes



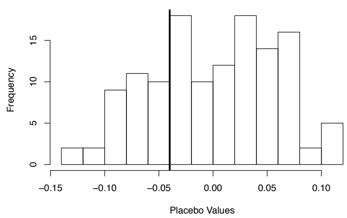
(a) All ED Visits, All Races
True effect is larger in magnitude than $\approx$ 52 % of placebo estimates



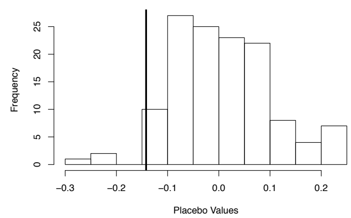
(b) All ED Visits, Blacks
True effect is larger in magnitude than $\approx$ 85 % of placebo estimates



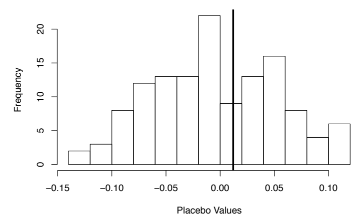
(c) All ED Visits, Non-Blacks
True effect is larger in magnitude than $\approx$ 35 % of placebo estimates



(d) Chronic Illness ED Visits, All Races
True effect is larger in magnitude than $\approx$ 45 % of placebo estimates



(e) Chronic Illness ED Visits in 2009, Blacks
True effect is larger in magnitude than $\approx$ 88 % of placebo estimates



(f) Chronic Illness ED Visits in 2009, Non-Blacks
True effect is larger in magnitude than $\approx$ 16 % of placebo estimates

These figures are histograms of the 129 "placebo" regression discontinuity estimates generated using every 4-year period between January 1965 and September 1983, prior to the actual discontinuity in Medicaid eligibility. The vertical line represents the effect estimated at the true Medicaid eligibility discontinuity.