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## Public insurance and child hospitalizations: access and efficiency effects

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### Abstract

The 1983–1996 period saw enormous expansions in access to public health insurance for low-income children. We explore the impact of these expansions on child hospitalizations. While greater access to inpatient care may increase hospital utilization, improved efficiency of care for children who are also newly eligible for primary care could lower hospitalization rates. We use a large sample of child discharges from the National Hospital Discharge Survey (NHDS) to assess the net impact of Medicaid expansions on hospitalizations during this period. We find that total hospitalizations increased significantly, with each 10 percentage-point rise in eligibility leading to an 8.4% increase in hospitalizations. Thus, the access effect strongly outweighs any efficiency effect produced by expanded coverage. However, we find some support for an efficiency effect: the increase in hospitalizations for unavoidable conditions is much larger than that for avoidable conditions that are most sensitive to outpatient care. Indeed, the increase in avoidable hospitalizations is less than half that of unavoidable hospitalizations, and it is not statistically significant. We also find that expanded Medicaid eligibility reduced the average length of stay, but increased the utilization of inpatient procedures, so that the net impact on total costs per stay is ambiguous.

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## 1. Introduction

The failure of insurance coverage rates to increase in tandem with economic growth over the past quarter-century has generated substantial political and academic interest. While GDP per capita rose 73% between 1975 and 2000, the share of the nonelderly without any insurance coverage *increased* from 11.2% to 15.8%.<sup>1</sup> Dramatic expansions of the Medicaid program since the mid-1980s have succeeded in reversing this trend for children, yet 12.4% of children under the age of 18 were still uninsured in 2000.

These trends concern policymakers for two reasons. The first is that insurance coverage is generally assumed to lead to more utilization of health care, thereby improving health. The second is that insurance coverage is generally assumed to lead to more efficient provision and utilization of medical care. For example, the newly insured are expected to substitute away from costly emergency rooms and toward personal physicians for their primary care needs. Furthermore, by having a “regular source of care” as well as access to preventative treatments, the newly insured may avoid hospital stays resulting from poorly treated or under-treated ailments.

There is a substantial literature assessing the first of these contentions, both via simple comparisons of individuals across insurance states, and through analyses of exogenous shifts in insurance coverage; see Gruber (1997) for an extensive review. There is much less evidence, however, on the second contention. The purpose of our paper is to address this deficiency. We do so by exploring the impacts of the Medicaid expansions of the late 1980s and early 1990s, the largest change in public insurance policy over the past 30 years, on the nature and incidence of pediatric hospitalizations in the U.S. Our primary interest is the impact of Medicaid on hospitalizations that have been denoted by medical experts as “avoidable.” The effect of Medicaid on these hospitalizations is theoretically ambiguous: while newly covered children may avoid hospital stays by receiving appropriate outpatient care (the “efficiency effect”), they may also be more likely to go to the hospital when ill (the “access effect”). Our empirical strategy allows us to explore the magnitudes of these countervailing effects.

We focus our analysis on children, who have been the primary target of public insurance policy over the past 20 years.<sup>2</sup> Approximately one-quarter of pediatric hospitalizations are classified as avoidable, as compared to 1 in 10 for adults. Despite growing concern about child health, pediatric hospitalizations have received scant attention in the health economics literature, which has focused on the inpatient utilization of adults. As a result, children are a particularly interesting group to study in this context.

Using data from the National Hospital Discharge Survey (NHDS), we investigate the impact of the Medicaid expansions on pediatric hospitalizations. These expansions

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<sup>1</sup> *Statistical Abstract of the United States*, 1979 and 2002. Table 161 (1979) and Tables 137 and 638 (2002).

<sup>2</sup> There have also been substantial expansions in the coverage of pregnancy, as detailed in Gruber (1997). Because virtually every full-term pregnancy results in a hospital admission, the question of interest in this literature is whether insurance coverage increases the utilization of prenatal care and improves birth outcomes. Currie and Gruber (1996a) find support for this efficiency effect.

occurred at a very different pace across states, and across different groups of children within states. This policy heterogeneity provides the exogenous variation in insurance status necessary to conduct our analysis. By matching Medicaid eligibility data to discharge data from the NHDS, the only nationally representative survey of hospital discharges, we can assess the impact of insurance status on the number and type of hospitalizations. The NHDS contains large samples of child discharges for each year, as well as detailed information on admission diagnoses that allows us to assess the “avoidability” of hospitalizations.

We find that expansions in Medicaid eligibility are associated with significant increases in the incidence of total child hospitalization. A 10 percentage-point increase in Medicaid eligibility is associated with an 8.1% increase in the unavoidable hospitalization rate, and a statistically insignificant 3.2% increase in the avoidable hospitalization rate. To the extent that the efficiency effect does operate, reducing the magnitude of the increase in avoidable hospitalizations that would otherwise have resulted from increased access to hospital care, it is strongly outweighed by the access effect of expanded coverage.

Due to the changing composition of patients, diagnoses, and payers, the average treatment intensity for pediatric hospitalizations may also be affected by the Medicaid expansions. We find that increases in eligibility are associated with shorter lengths of stay and a greater number of procedures performed during hospitalization.

Our paper proceeds as follows. Section 2 provides background on child hospitalizations, avoidable hospitalizations, and the Medicaid expansions. Section 3 discusses our data and empirical strategy. Section 4 presents our results, and Section 5 concludes.

## 2. Background

### 2.1. Pediatric hospitalizations

Accounting for 7.2% of all hospital admissions in 1996, pediatric hospitalizations are frequently overlooked in the health economics literature. Yet these hospitalizations were responsible for US\$21.4 billion in charges in 2000, representing 4.5% of hospital charges and over 40% of total expenditures on child health care services.<sup>3</sup> Moreover, recent public health insurance reforms have focused on extending coverage to impoverished children, highlighting the need to understand children’s health and health care utilization patterns.

Fig. 1 presents time trends in hospitalization rates in the U.S. for the under 15, 15 to 64, and 65 plus populations, estimated using the annual National Hospital Discharge Survey (NHDS). The NHDS defines a hospitalization as a “formal admission to the inpatient service of a short-stay hospital for observation, care, diagnosis, or treatment, or by birth.”

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<sup>3</sup> Estimates based on national statistics reported by the Agency for Healthcare Research and Quality, 1999 and 2000. Total charges estimated for children aged 0–17; share of charges is calculated using data on children under 6.

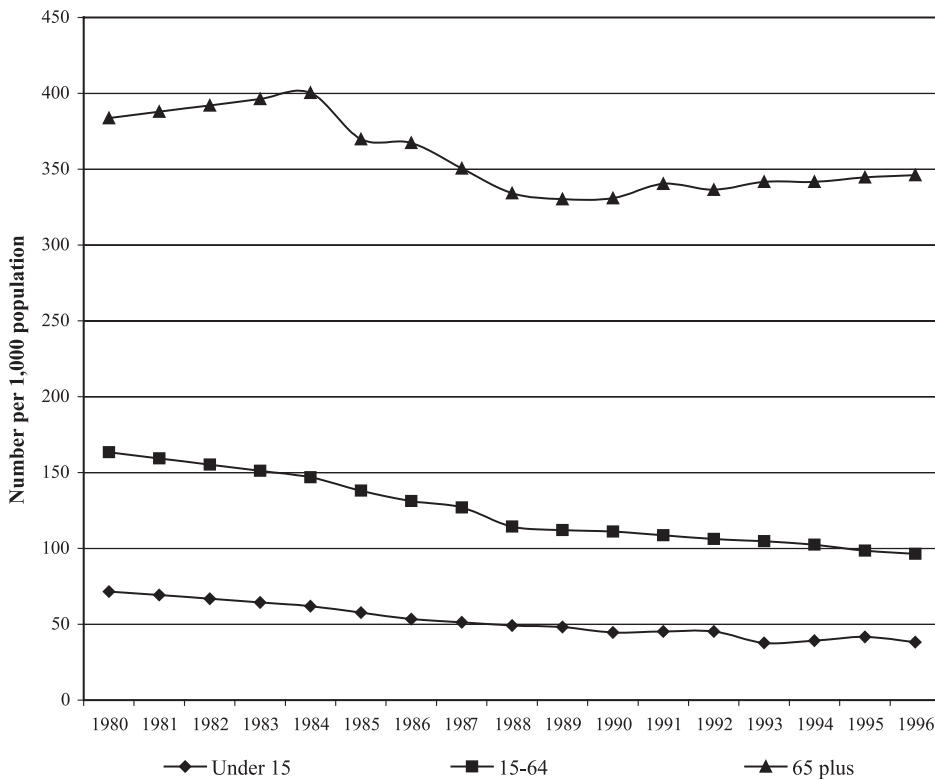


Fig. 1. Hospitalization rates by age group.

It samples discharges from non-federal, short-stay hospitals with six beds or more, whose specialty is general (medical or surgical) or children’s general. All estimates we present exclude hospitalizations of newborns.

The disparity in hospitalization rates across age groups is large, with the over-65 population hospitalized nine times as frequently as children (346 per 1000 elderly vs. 38 per 1000 children in 1996). A marked decline in hospitalization rates during the 1980–1996 period is evident for both the 15–64 and the under 15 groups; the trend in the over-65 category fluctuates during this period, but due to the changing age composition of this group, this trend cannot be meaningfully compared to the trends in the younger groups. Overall, the hospitalization rate declined by 31%, with the relative decline for children (47%) the largest among the three groups.

Table 1 compares the leading causes of pediatric and adult hospitalizations, tabulated using the first-listed diagnosis code in the 1996 NHDS. The table highlights the obvious age-related patterns in hospital needs, with diseases of the respiratory system (asthma, pneumonia, and acute infections) topping the list for children, childbirth ranking first for adults 15–64, and diseases of the circulatory system (largely heart disease) accounting for the plurality of hospitalizations among the elderly. Infectious and parasitic diseases, along

Table 1  
Top hospital diagnoses and prevalence by age category, 1996

	Number (000s)	Percentage of total
<i>Under 15</i>		
Diseases of the respiratory system	653	30
Injury and poisoning	223	10
Diseases of the digestive system	205	9
Endocrine, nutritional and metabolic diseases, and immunity disorders	155	7
Infectious and parasitic diseases	153	7
All diagnoses	2207	100
<i>15–64</i>		
Deliveries	3817	23
Diseases of the circulatory system	2119	13
Mental disorders	1546	9
Diseases of the digestive system	1503	9
Injury and poisoning	1350	8
All diagnoses	16,619	100
<i>65 plus</i>		
Diseases of the circulatory system	3963	34
Diseases of the respiratory system	1550	13
Diseases of the digestive system	1198	10
Injury and poisoning	977	8
Neoplasms	826	7
All diagnoses	11,718	100
All diagnoses, all ages	30,544	

Source: National Center for Health Statistics (1998).

with endocrine, nutritional, metabolic, and immunity disorders, account for the major afflictions specific to children.

Despite this wealth of statistical information, there is little work by health policy analysts on the causal determinants of child hospitalization; the work that exists is largely descriptive in nature. McConnochie et al. (1997a) review the medical literature on pediatric hospitalizations, drawing two conclusions that are important for our purposes. First, there is a substantial amount of “inappropriateness” or “avoidability” in child hospitalization; we discuss this further below. Second, there is a great deal of geographic variation in hospitalization rates of children that is not easily explained by morbidity differences. In their own studies of hospitalization rates for infants of different socioeconomic backgrounds, McConnochie et al. find that nearly 80% of the higher rates for disadvantaged infants is due to “discretionary” as opposed to “mandatory” conditions, suggesting that disease prevalence is not the only determinant of hospitalization. A study of infant hospitalizations for asthma by Homer et al. (1996) controls for morbidity burdens using measures of oxygen saturation in admitted patients, and finds that morbidity does not explain all of the differences in hospitalization rates between Boston and Rochester, New York. Goodman et al. (1994) investigate the relationship between health system characteristics and pediatric discharges, concluding that discharges are positively associated

with bed supply, and negatively associated with distance from the hospital and residence near an academic medical center.

Finally, several studies have noted that hospitalization rates are higher for the uninsured and Medicaid populations, although these studies do not present separate estimates for children (e.g. Weissman et al., 1992; Billings and Teicholz, 1990). Only two studies of which we are aware utilize a framework that addresses causality between insurance status and the probability of hospitalization for children. Manning et al. (1987) exploit randomly assigned variation in copayment rates from the RAND Health Insurance Experiment to show that there are insignificant increases in the probability of admission for children covered fully for inpatient expenses versus those covered by cost-sharing plans. Currie and Gruber (1996b) find the opposite, concluding that becoming eligible for Medicaid increases the probability of hospitalization by 82%. The discrepancy between these findings may arise from the fact that the RAND experiment capped out-of-pocket exposure at a low level, so that in many cases children covered by cost-sharing plans were effectively fully insured for hospitalization.

## 2.2. Avoidable hospitalizations

Defined as hospitalizations that “might not have occurred had [patients] received effective, timely, and continuous outpatient (ambulatory) medical care for certain chronic disease conditions,” avoidable hospitalizations (AVHs) are commonly used as a measure of access to health care. As such, the list of AVH diagnoses is carefully selected so as to represent conditions more likely to result from inadequate access to ambulatory care than from differences in disease prevalence or provider practices. The list is therefore distinct from so-called “discretionary” admissions, those for which subjective physician judgment is an integral part of the decision to admit.<sup>4</sup> For example, admissions for immunizable conditions are non-discretionary and avoidable, whereas admissions for acute fever are discretionary but unavoidable. Of course, as noted by Weissman, Gatsonis, and Epstein in their oft-cited 1992 JAMA article on AVH rates by insurance status, “being avoidable is a matter of degree.” Healthcare researchers are careful to state that primary care cannot prevent all hospitalizations in this category; disease severity, treatment compliance, and “health-seeking behavior” are also important factors (Gadomski et al., 1998). Nevertheless, the AVH rate is designed to capture the effectiveness of the health care system in providing timely care.

A list of pediatric AVH diagnoses, compiled by Gadomski et al. (1998), is presented in Table 2. This list originates in a 1993 Institute of Medicine report on access to health care, which was modified to reflect pediatric illnesses. Of the 39.4 million pediatric hospitalizations that occurred between 1983 and 1996, 26% were classified as avoidable using this definition.<sup>5</sup> The top 6 avoidable conditions are asthma (24% of AVHs), pneumonia (23%),

<sup>4</sup> McConnochie et al. (1997b).

<sup>5</sup> There are two other definitions used in the literature, but one is not specific to children (Weissman et al., 1992), and the second (Casanova and Starfield, 1995) is a slightly more expansive version of the Gadomski et al. list, classifying 29% of the NHDS hospitalizations as avoidable. Given the minimal differences between the two pediatric definitions available, we choose the more conservative measure.

Table 2  
Pediatric avoidable hospitalization conditions

Condition	ICD-9-CM Code(s)	Qualifiers
Immunization preventable conditions	033, 037, 045, 320, 390, 391	Haemophilus meningitis (320.2) for age 1–5 only
Grand Mal status and other epileptic convulsions	345	
Convulsions “A and B”	780.3	Age >5 years
Severe ENT infections	382, 462, 463, 465, 472.1	Exclude otitis media (382) with myringotomy with insertion of tube (procedure 20.01)
Bacterial pneumonia	481, 482.2, 482.3, 482.9, 483, 485, 486	Exclude cases with secondary diagnosis of sickle cell (282.6) and patients <2 months
Asthma	493	
Tuberculosis	011–018	
Cellulitis	681, 682, 683, 686	Exclude cases with a surgical procedure (01–86.99)
Diabetes “A”	250.1, 250.2, 250.3	
Diabetes “B”	250.8, 250.9	
Diabetes “C”	250	
Hypoglycemia	251.2	
Gastroenteritis	558.9	
Kidney/urinary infection	590, 599, 599.9	
Dehydration-volume depletion	276.5	
Iron deficiency anemia	280.1, 280.8, 280.9	Age 0–5 years
Nutritional deficiencies	260, 261, 262, 268, 268.1	
Failure to thrive	783.4	Age <1 year

Source: Gadomski et al. (1998).

gastroenteritis (14%), ear, nose, and throat (ENT) infections (13%), dehydration (8%), and kidney/urinary tract infections (5%). Previous estimates of the share of hospitalizations that are avoidable range from 7% to 12% for the nonelderly population as a whole and 18% to 28% for children.<sup>6</sup>

Fig. 2 graphs AVH rates per 1000 population under age 16, as well as the share of hospitalizations for this age group that are categorized as avoidable. Between 1983 and 1996, there was a steep decline in the AVH rate of nearly 35%. However, this decline was smaller than the overall decline in the hospitalization of children, leading to a larger share of hospitalizations that are avoidable.

Previous research on AVHs has concentrated on two areas: (1) calculating age and gender-standardized AVH rates for different populations and types of insurance coverage; (2) establishing a causal link between inadequate ambulatory care and subsequent AVHs. Within the first research area, there are a number of papers in the medical literature that document significantly higher AVH rates among low-income populations and blacks (e.g.

<sup>6</sup> Weissman et al. (1992); Pappas et al. (1997); Casanova and Starfield (1995); McConnochie et al. (1997a); Soulen et al. (1994). Note the estimates that include adults exclude psychiatric and obstetrical admissions.

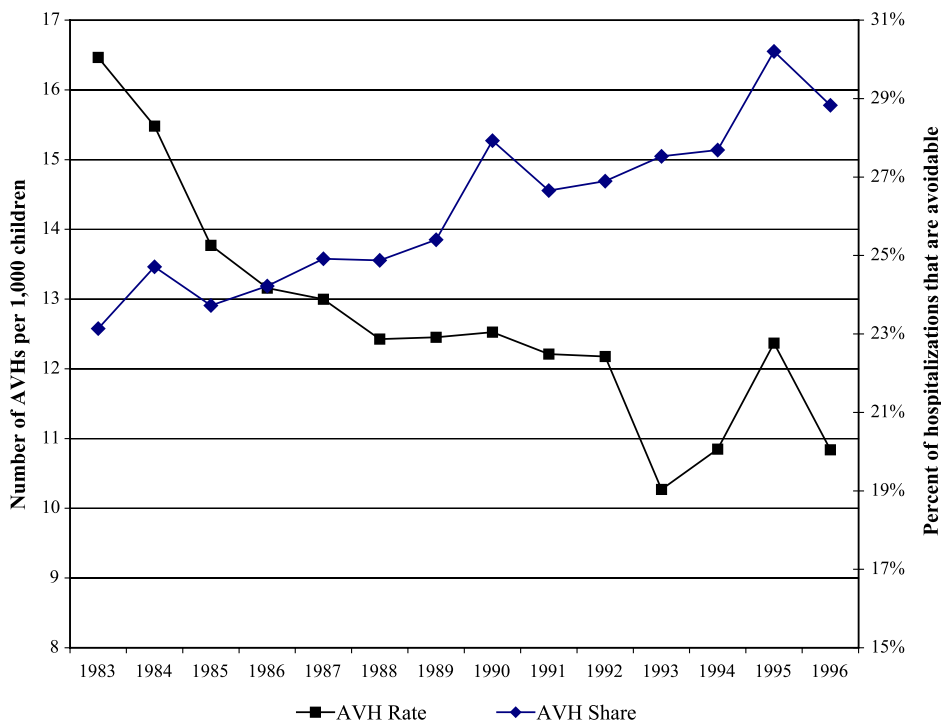


Fig. 2. Avoidable hospitalization of children under 16, 1983–1996.

Begley et al., 1994; Billings et al., 1993). Weissman et al. (1992) explore the relative risk of admission, by insurance status, for 12 AVH conditions in the under-65 population residing in Maryland and Massachusetts in 1987. They find that the share of hospitalizations that are avoidable is much greater for the uninsured than for the privately insured, and still higher for Medicaid enrollees.<sup>7</sup> These correlations between insurance or income status and AVHs constitute suggestive but inconclusive evidence of a causal link between the two factors; the possibility of omitted variables bias in such analyses is substantial. For example, those who are uninsured may be in worse underlying health, leading to more avoidable hospitalizations independent of their insurance status. By exploiting exogenous changes in insurance coverage across different age groups and states over time, our approach enables us to surmount these types of biases.

The only previous work of which we are aware that attempts to assess the impact of exogenous insurance shifts on the efficiency of hospitalization is Kaestner et al. (2001), henceforth KJR. Using discharge data from 11 states, KJR compare the change in AVH rate between 1988 and 1992 for children residing in “poor” and “near-poor” zip codes (median family income <US\$25,000 and US\$25,000–30,000, respectively) with the change in AVH rate for children living in more affluent zip codes (median family income

<sup>7</sup> Two other studies that examine the relationship between insurance status and AVH rates are Billings and Teicholz (1990) and Pappas et al. (1997).



>US\$30,000). Because children in the lower-income zip codes were more likely to be affected by the Medicaid expansions during the study period, KJR's estimates can sign the impact of the expansions on the incidence of AVHs, provided that their identifying assumption of similar AVH trends among all income groups in the absence of the expansions is correct.

KJR find that the incidence of AVHs, measured as the number of AVHs divided by the number of births in a hospital-zipcode category-year cell, declined for children aged 2–9, although this decline was more pronounced for the near-poor than for the poor, and was not uniform across all categories of AVHs.<sup>8</sup> In addition to using a more precise instrument for Medicaid eligibility and utilizing data on a larger range of ages, states, and years, our specifications include state×year interaction terms, so our estimates exclude state-level trends that may be affecting KJR's findings.

The second strand of literature provides evidence for the presumption that avoidable hospitalizations are indeed avoidable; that is, that outpatient services do, in fact, prevent AVHs. One approach has been to link aggregate measures of access to medical care with corresponding data on AVH rates. Using hospital discharge data from 26 health service areas (HSAs) in Pennsylvania in 1989, [Parchman and Culler \(1994\)](#) find that higher per-capita rates of family and general-practice physicians are negatively associated with AVH rates, after controlling for the effects of mean per-capita income. Another approach has been to show that avoidable hospitalizations are associated with inadequate pre-hospital care. [Solberg et al. \(1990\)](#) find that 45% of avoidable hospitalizations studied failed explicit quality criteria and 10% were judged by physicians to have received poor-quality care.

A third approach is taken by [Homer et al. \(1996\)](#) and [Halfon and Newacheck \(1993\)](#), who show that places and income groups with lower rates of preventative care for childhood asthma have higher rates of hospitalization, although this association could also be explained by a host of other intermediating factors between location and poverty status and hospitalization. Finally, [Gadomski et al. \(1998\)](#) evaluate the Maryland Access to Care (MAC) Medicaid managed care program, which emphasized improved access to primary care for Medicaid enrollees. They find that the program increased the odds of ambulatory care, and that among those children who did use ambulatory care, the program was associated with decreased probabilities of both hospitalization in general and avoidable hospitalization in particular.

Each of these approaches has limitations, but the weight of the evidence supports the contention of a link between inadequate primary care and avoidable hospitalization. We therefore follow the medical literature in employing the AVH rate as a measure of the efficiency of patient care.

### 2.3. Medicaid expansions

Historically, Medicaid eligibility for children was tied to participation in the Aid for Families with Dependent Children program (AFDC). This linkage with AFDC restricted

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<sup>8</sup> KJR also report slightly smaller declines using the number of unavoidable hospitalizations as the denominator for the AVH rate, but these findings are not directly comparable to our results due to the fact that Medicaid also affects the incidence of unavoidable hospitalizations.

access to the program in three ways. First, despite the existence of the AFDC-Unemployed Parents program (AFDC-UP), which provided benefits to households in which the primary earner was unemployed, AFDC benefits were generally available only to single-parent households. Second, income cutoffs for cash welfare vary across states, and can be very low. For example, in 1984, the cutoff for a family of 4 in South Carolina was only 29% of the poverty line. Third, the stigma of applying for cash welfare programs may have prevented eligible families from receiving Medicaid benefits.

In some states, children could also qualify for Medicaid under state-run Medically Needy or Ribicoff programs. The Medically Needy program relaxed the income criteria for eligibility by covering families who would have been eligible for AFDC if their incomes were lower, but who had large medical expenditures that brought their disposable income below program thresholds. The Ribicoff option allowed states to cover children in two-parent families who met the AFDC income criteria.

Beginning with the Deficit Reduction Act of 1984 (DEFRA '84), the linkage between AFDC coverage and eligibility for Medicaid was gradually weakened. DEFRA '84 eliminated the family structure requirements for Medicaid eligibility of young children by requiring states to cover children born after September 1, 1983 who lived in families that were income-eligible for AFDC. DEFRA was followed by a series of measures that raised the income cutoffs for Medicaid eligibility, first at state option, and then by federal mandate. These options are described in Appendix A. The differential pace with which states pursued these options produced a great deal of variation across states in both the income thresholds and the age limits governing Medicaid eligibility.

We estimate that the fraction of children under 16 who were eligible for Medicaid rose by 16 percentage points between 1983 and 1996. This national trend masks considerable heterogeneity across the states: there was actually a decline in eligibility of 2 percentage points in Alaska during this period, and a rise of 38 percentage points in West Virginia. In addition, there is also heterogeneity within states in the rate at which children of different ages were covered. For example, coverage of infants under 1 rose by 46 percentage points in Texas, while coverage of children ages 11 to 15 rose by less than 5 percentage points. It is this variation across states, within states over time, and even across different age groups in the same state at a given point in time, that we use to identify our models.

### **3. Data and empirical strategy**

#### *3.1. Data*

Our study period begins in 1983, nearly a year before the first federally mandated expansions took effect, and continues through 1996, the latest year for which all the data are available. The source of our hospitalization data is the National Hospital Discharge Survey (NHDS), the only continuous nationwide survey of inpatient utilization of non-federal, short-stay hospitals. The NHDS samples approximately 250,000 discharges annually,

collecting data on diagnosis and procedure codes, discharge status, length of stay, and selected hospital and demographic characteristics. Weights provided with the survey data enable estimation of statistics for the universe of annual hospitalizations in the U.S.

Because our primary independent variable of interest, Medicaid eligibility, varies only by state, birth date, and calendar quarter, we group the individual hospital data into cells. The sample size does not permit grouping at such a fine level of detail, so we define cells for four age categories for each state and year. The age categories are children under 1 year old, 1–5-year-olds, 6–10-year-olds, and 11–15-year-olds. Note that the under 1 category does not include the initial hospitalization of newborns admitted upon birth. This age group nevertheless warrants its own category, both because it accounts for 27.6% of total hospitalizations to children under 16 during the study period, and because many federal and state initiatives have expanded Medicaid eligibility specifically for infants under 1 year old.

Of the resulting 2856 cells (4 age categories  $\times$  14 years  $\times$  51 “states” (50 states plus Washington, DC)), we drop 548 because the corresponding state–years are not surveyed at all or several age groups for a given state–year are missing. Sixteen states in total are affected: 3 are dropped entirely from the sample, 11 are dropped in the late 1980s, and 2 are missing data for 2 consecutive years in the middle of the study period but are otherwise included in the sample. Regressions of a dummy for inclusion in the data set on our independent variables reveal no systematic relationship between the probability of inclusion and our variable of interest. Each of the remaining 2308 cells is then matched to the appropriate age group/state/year population estimate from the Census Bureau. We calculate hospitalization rates by dividing the weighted number of hospitalizations in each cell by population.

The main advantage of the NHDS is that it provides a large, nationally representative sample of hospital discharges, with information on diagnosis at hospitalization that can be used to identify avoidable hospitalizations. The main disadvantage is that the NHDS is not designed to yield state-specific estimates of either total or child hospitalizations. Discharges from hospitals with 1000 or more beds are sampled with certainty, while discharges from smaller institutions are selected using a stratified, three-stage design, with selection of primary sampling units (PSUs), hospitals within the selected PSUs, and discharges within the hospitals constituting the first, second, and third stages, respectively.

This sampling design presents two challenges for our empirical approach, which relies on age group/state estimates of annual hospitalization rates. First, this sampling approach can and does leave a number of states entirely (or almost entirely) out of the survey. Discharges in the states that are included in the survey are overweighted in order to produce national estimates; thus, several age group/state cells have large numbers of weighted hospitalizations relative to the underlying age group/state population. Because we cannot include states without survey data in our analysis, the hospitalization statistics for our sample are higher than the true national average. Second, since the PSUs can cross state boundaries, a sample that is representative of a PSU may not be representative of the individual states that comprise it.

Neither of these factors would present a significant problem for the analysis if the sampling rules remained fixed from year to year. By including state fixed effects, we can

capture the extent to which states' estimated hospitalization rates deviate from representative levels. We can also reduce the influence of outliers by censoring the hospitalization rates.<sup>9</sup> However, a 1987 redesign of the NHDS sampling method may have led to a shift in the composition of hospitals across states within a PSU. To control for this possibility, we also include in our models a set of state $\times$ year interactions. These interactions allow for changes within states over time in the sampling frame of hospitals, and also control for other state time trends that may be correlated with Medicaid eligibility policy. Our model is still identified when these are included because the expansions occurred at a differential pace for different age groups.

Descriptive statistics for our data are presented in Table 3. The first set of variables in the table is expressed relative to the underlying population in each age group/state/year cell, i.e. the number of hospitalizations per child. The mean for this hospitalization rate is 9.9%, much higher than the rates graphed in Fig. 1.<sup>10</sup> However, the time trend in the data is very similar to the national trend, suggesting little systematic bias to our estimates as a result. On average, the AVH rate during our study period is 2.3%, or 23% of the total hospitalization rate. The second set of variables is expressed relative to the total number of hospitalizations in each age group/state/year cell, i.e. the share of hospitalizations with at least one procedure. The average length of stay (LOS) per admission is 4.25 days, and a procedure is performed during 46% of hospital stays.

Due to the sensitive nature of the information gathered in the NHDS, geographic identifiers are not released in the public use files. We were able to create the NHDS data cells and match Medicaid policy variables to those cells through an agreement with the Research Data Center at the National Center for Health Statistics. Once the dataset was complete, we were allowed restricted remote access to the data.

### 3.2. Empirical strategy

Our independent variable of interest is the percentage of children in each age group, state, and year who are eligible for Medicaid. We estimate this variable using data from the Current Population Survey (CPS), together with a detailed eligibility model originally developed for Currie and Gruber (1996a,b) and updated through 1996 for this project. This model, described in greater detail in these earlier papers, uses information on family structure, age, income, state and year to impute eligibility for Medicaid using state-specific rules for AFDC and expansion eligibility.

We begin by extracting data for 0- to 15-year-olds from the March Current Population Survey (CPS) for each year, which has sufficient information on income, family structure, and location to determine eligibility for Medicaid. Next, we compute eligibility for each child in the CPS data, calculate weighted averages for the age groups in our NHDS sample (0, 1–5, 6–10, 11–15), and match the eligibility measure onto the NHDS sample by age group, state, and year.

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<sup>9</sup> We censor the individual cell hospitalization rates at 0.5. This affects fewer than 2% of our cells. In addition, we log the hospitalization rates in our analyses, so the impact of censoring on our results is trivial.

<sup>10</sup> It is not possible to correct this rate to reflect multiple discharges for the same child.

Table 3  
Descriptive statistics

	Mean	Standard deviation
<i>Statistics for total population (N=2308)</i>		
Hospitalization Rate	9.90%	15.3%
AVH rate	2.30%	3.4%
Non-AVH rate	7.60%	12.6%
ELIG	22.2%	15.3%
SIMELIG	22.8%	13.6%
<i>Statistics for hospitalizations (N=2306)</i>		
AVH share	24.9%	15.2%
Average LOS	4.25	1.9%
Any procedure	45.8%	20.2%
Average number of procedures	0.76	0.42

The unit of observation is an age group/state/year cell.

We then estimate models of the following form:

$$\ln(\text{HOSP})_{ajt} = \alpha + \beta_1 \text{ELIG}_{ajt} + \beta_2 \eta_a + \beta_3 \delta_j + \beta_4 \tau_t + \beta_5 \delta_j \tau_t + \varepsilon_{ajt} \quad (1)$$

where  $a$  indexes age groups,  $j$  indexes states, and  $t$  indexes years; HOSP is a hospitalization rate (or one of our other dependent variables), ELIG is the fraction of children eligible for Medicaid in each age group/state/year cell, and  $\eta_a$ ,  $\delta_j$ , and  $\tau_t$  are full sets of dummy variables for age group, state, and year, respectively.

This model relates the rate of hospitalization in a cell to the probability that a child in that cell is eligible for Medicaid. We control for age group, state, and year fixed effects to capture any underlying correlation between Medicaid eligibility and hospitalization within these groups. In addition, as discussed above, we include a full set of state×year interactions to control both for any changes in the NHDS sampling frame that change how a state is represented in our data, and for other state-specific trends that might be correlated with Medicaid eligibility policy. Because these changes and trends are likely to have a proportional effect on each age group, we log the dependent variable.

Even in this rich framework, the estimated coefficient on ELIG may be biased for two reasons. The first is measurement error: given the small sample sizes by age group and state in the CPS, there is likely to be substantial noise in our estimate of true Medicaid eligibility in the underlying population. The second is omitted variables bias: the actual eligibility of these children may be correlated with omitted factors that also determine their hospitalization rates. For example, a recession that hits a given age group/state particularly hard may lead to both rising Medicaid eligibility and rising hospitalization rates.

We therefore instrument for actual eligibility using “simulated eligibility,” an instrument developed by Currie and Gruber (1996a,b). To construct this instrument, we begin by drawing a nationally representative, random sample of 250 children of every age from every year’s March CPS. Then, we take this *same sample* through our simulation programs to calculate the fraction of children of each age who would be

eligible for Medicaid if they lived in each state. That is, we ask how many zero-year-olds would have been eligible had they lived in California, how many would have been eligible had they lived in Texas, etc. We then aggregate the data into the NHDS age categories and match it to the NHDS data by age group, state and year. This simulated eligibility measure provides a convenient index of the generosity of state Medicaid rules that utilizes only variation in the eligibility rules across states, years, and age groups of children. It is independent of factors specific to age group/states that might affect both Medicaid eligibility and hospitalization rates. By using simulated eligibility as an instrument for actual eligibility, we also surmount measurement error problems in our actual eligibility measure, so long as the error does not derive from the miscoding of state rules. As shown in Table 3, 22% of the children for whom hospital data is available are estimated to be eligible for Medicaid over this sample period.

Finally, we estimate our models using a two-step feasible GLS procedure because the accuracy of HOSP depends on the number of observations in the NHDS relative to the volume of hospitalizations in the underlying population (that is, the sampling probability for cell  $ajit$ ). In the first step, we perform 2SLS and calculate the residuals,  $\hat{\varepsilon}_{ajit}$ . We regress the squared residuals on a constant and the estimated variance of each error term, where this variance is calculated as  $((1 - \text{HOSP}_{ajit}) / \text{HOSP}_{ajit}) \times (1 / (\text{population}_{ajit} \times \text{sampling prob}_{jt}))$ .<sup>11</sup> This formula is derived using the Delta method to approximate the variance of the log of a proportion.<sup>12</sup> The coefficients from this regression are used to predict variances,  $\hat{\varepsilon}_{ajit}^2$ . We then re-estimate Eq. (1) by two-stage weighted least squares, with weights  $= \frac{1}{\hat{\varepsilon}_{ajit}^2}$ .

## 4. Results

### 4.1. Total hospitalizations

The first column of Table 4 presents our results for total hospitalizations. As noted above, our dependent variable is the number of hospitalizations per child, and our independent variable of interest is the percentage of children in the corresponding age group/state/year cell who are eligible for Medicaid. We show only the coefficient of interest from models that also include a full set of age group, state, and year dummies, as well as state  $\times$  year interactions.

Our estimate indicates that an increase in Medicaid eligibility of 10 percentage points is associated with an increase in the child hospitalization rate of 8.4%. This estimate is nearly identical to the estimate obtained by Currie and Gruber (1996b) using 1984–1992 micro-data on children from the National Health Interview Survey. Since the expansions

<sup>11</sup> Restrictions on the use of the NHDS data prevented our calculating these sampling probabilities for each age group/state/year cell. We were, however, able to develop sampling probabilities for each state and year using the ratio of total NHDS survey discharges to total discharges reported by the American Hospital Association.

<sup>12</sup> It is based on the variance of a proportion,  $p(1-p)/n$ , where  $n = \text{population} \times \text{sampling probability}$ , and  $p = \text{HOSP}$ .

Table 4  
Basic results for hospitalizations

	Ln(Total hospitalizations/ Population)	Ln(Avoidable hospitalizations/ Population)	Ln(Unavoidable hospitalizations/ Population)
Medicaid eligibility	0.841 (0.129)	0.315 (0.198)	0.806 (0.144)
Main effects (age group, state, year)	Y	Y	Y
State×year interactions	Y	Y	Y
Age group×year interactions	N	N	N
Age group×state interactions	N	N	N
Mean of dependent variable (unlogged)	0.092	0.024	0.072
N	2306	2210	2300

Coefficient is that on Medicaid eligibility from estimating regressions such as (1) in the text, using SIMELIG as an instrument for ELIG and a two-step FGLS procedure. Standard errors are in parentheses.

over our study period increased eligibility by 16 percentage points, we estimate that they increased child hospitalization by 13.4%, or approximately 6.5 additional hospitalizations per 1000 children.<sup>13</sup>

Using Currie and Gruber's estimate of the take-up rate for newly eligible children (30%), our point estimate implies that covering a child with Medicaid increases her odds of hospitalization by 3.8 times. This is a large effect, but it is consistent with evidence on utilization rates of Medicaid enrollees. For example, Weissman et al. (1992) find that the hospitalization rate for Maryland Medicaid enrollees under age 65 is 3.2 times that of uninsured individuals under 65. Our own calculations using the 1990 NHDS and insurance coverage statistics reported by the U.S. Bureau of the Census reveal that Medicaid enrollees under the age of 16 were hospitalized 2.8 times as often as privately insured children. It is not surprising that newly covered enrollees have higher utilization rates than the average enrollee.

This finding also reveals that any efficiency effect generated by the Medicaid expansions was not large enough to produce a decline in total hospitalizations. That is, either efficiency did not increase, or any efficiency gains were offset by the increased access of children to inpatient care. To assess which explanation is most consistent with the data, we next turn to a decomposition of the total hospitalization rate into its avoidable and unavoidable components.

#### 4.2. Avoidable and unavoidable hospitalizations

The next two columns of Table 4 present our findings for avoidable hospitalizations and their complement, unavoidable hospitalizations. A 10 percentage-point increase in Medicaid eligibility is associated with an 8.1% increase in the unavoidable hospitaliza-

<sup>13</sup> This estimate is based on a rate of 48.4 hospitalizations per 1000 children under age 15. This is the average of the annual hospitalization rates from 1983 to 1996 for children under 15, as reported by the NCHS (using the NHDS data) and graphed in Fig. 1.

tion rate, and a statistically insignificant 3.2% increase in the avoidable hospitalization rate.<sup>14</sup> A *t*-test for equality between the two coefficients is easily rejected ( $p < 0.05$ ), implying that unavoidable hospitalizations increased more in response to the Medicaid expansions than did avoidable hospitalizations.

Thus, these findings provide evidence for both access and efficiency effects of expanding insurance coverage. While there is a large increase in unavoidable hospitalizations, there is a much smaller and statistically insignificant increase in avoidable hospitalizations. Assuming access to hospital care increases the likelihood of all hospitalizations equally, these results imply that the increased use of primary care engendered by the Medicaid expansions mitigated the increase in total hospitalizations by reducing the increase in avoidable hospitalizations that would otherwise have occurred.

#### 4.3. Specification checks

The results presented in Table 4 arise from a rigorous empirical specification that controls for fixed differences across states, years, and age groups, and even for time trends by state. It is possible, however, that other omitted factors that determine hospitalizations are not captured by this regression framework. In order to explore this possibility, we further enrich the set of controls in the regression. We first add age group  $\times$  year interactions. This allows for arbitrary time trends in hospitalization rates by age group that may confound our findings. We then add age group  $\times$  state interactions to control for any fixed differences in hospitalization across age group/state cells. With both sets of controls included, this is a full “differences-in-differences-in-differences” setup, with only third-level interactions of age group  $\times$  state  $\times$  year identifying the empirical models.

The results of these robustness checks are presented in Table 5. In each case, the addition of the interaction terms increases the standard errors on the coefficient estimates and reduces their magnitude, so that in the “differences-in-differences-in-differences” model, none of our estimates is significant. At the same time, the basic pattern of findings holds: there is an increase in unavoidable hospitalizations of the same magnitude as total hospitalizations, and a smaller increase in avoidable hospitalizations. Due to the increased imprecision of the estimates, these specification checks are largely uninformative. However, the increase in the standard errors implies that none of the estimates in Table 5 are significantly different from those presented in Table 4.

#### 4.4. Impact of Medicaid on the intensity of inpatient care provided

Our analysis thus far has examined the impact of Medicaid on the incidence of child hospitalization, with the purpose of evaluating the hypothesis that expanding public health insurance increases the efficiency with which medical care is delivered. However, public health insurance may affect not only the volume but also the intensity of care children receive in the hospital.

<sup>14</sup> Observations with a hospitalization rate of zero are dropped. 2SLS estimation of specification (1) using an indicator variable for a positive AVH rate as the dependent variable reveals no relationship to Medicaid eligibility.



Table 5  
Sensitivity tests, hospitalization regressions

	Ln(Total hospitalizations/ Population)		Ln(Avoidable hospitalizations/ Population)		Ln(Unavoidable hospitalizations/ Population)	
Medicaid eligibility	0.520 (0.263)	0.202 (0.268)	0.181 (0.421)	0.095 (0.460)	0.471 (0.284)	0.173 (0.290)
Main effects (age group, state, year)	Y	Y	Y	Y	Y	Y
State×year interactions	Y	Y	Y	Y	Y	Y
Age group×year interactions	Y	Y	Y	Y	Y	Y
Age group×state interactions	N	Y	N	Y	N	Y
N	2306	2306	2210	2210	2300	2300

Coefficient is that on Medicaid eligibility from estimating regressions such as (1) in the text, using SIMELIG as an instrument for ELIG and a two-step FGLS procedure. Standard errors are in parentheses.

The relationship between insurance status and the intensity of inpatient care has received considerable attention in the health literature. Several studies, reviewed in Weissman and Epstein (1990) and Currie and Gruber (2001), compare the length of stay and mean number of procedures for groups with different types of insurance coverage and reach somewhat mixed conclusions: those with private insurance coverage are treated much more intensively than are the uninsured, but those with public insurance coverage do not appear to be consistently treated more intensively.

Currie and Gruber (2001) extend this literature by examining the impact of the Medicaid expansions on the intensity of treatment for childbirth, using a variety of obstetric procedures as intensity measures. They find no aggregate impact on intensity of treatment, but they identify an important compositional impact. For those mothers who were likely to be uninsured prior to becoming Medicaid-eligible, there is a significant increase in treatment intensity. However, for those mothers who were likely to be privately insured, and therefore potentially subject to crowdout, treatment intensity declines. This result is consistent with the disparity in reimbursement rates between Medicaid and private insurers. Medicaid reimburses providers at much lower levels, so that a move from private to Medicaid coverage lowers incentives for intensive treatment.

We examine the impact of Medicaid on the intensity with which children are treated in the hospital. Following prior research (e.g. Cutler, 1991), we consider two dependent variables: length of stay (LOS) in the hospital, in days; and the number of procedures performed on the child. Due to the high volume of discharges without any procedures, we also consider an indicator variable for having an inpatient procedure, denoted “any procedure.” The mean value for each of these variables is calculated by age group/state/year. Descriptive statistics are presented in Table 3.

Table 6 contains results from estimating specification (1) and substituting the intensity measures for hospitalization rates. Simulated eligibility again instruments for actual eligibility of children in each cell. The estimation method for the length of stay and number of procedures regressions is two-stage weighted least squares, with weights equal

Table 6  
Effect of Medicaid on intensity of treatment

	Ln(LOS)	Ln(Number of procedures)	Ln(Any procedure)
Medicaid eligibility	-0.311 (0.086)	0.504 (0.121)	0.659 (0.110)
Main effects (age group, state, year)	Y	Y	Y
State×year interactions	Y	Y	Y
Age group×year interactions	N	N	N
Age group×state interactions	N	N	N
Mean of dependent variable (unlogged)	4.25	0.774	0.468
<i>N</i>	2306	2259	2259

Coefficient is that on Medicaid eligibility from estimating regressions such as (1) in the text, using SIMELIG as an instrument for ELIG and a two-step FGLS procedure. Standard errors are in parentheses.

to the number of observations in each cell. The any procedure regression is estimated using the two-step FGLS method described above, but replacing sampling probability<sub>ast</sub>×population<sub>ast</sub> in the denominator of the estimated variance formula with the number of observations in each cell.

The Medicaid coefficients reported in Table 6 capture both demand-side and supply-side effects of the expansions on intensity of care. On the demand side, Medicaid increases the number of hospital admissions. If these marginal admissions require less (more) intensive treatment than did the average admission prior to the expansions, then the demand-side effect should decrease (increase) the average intensity of care provided. On the supply side, there will be more intensive treatment of those moving from an uninsured state to Medicaid, but less intensive treatment of those moving from private insurance to Medicaid. Thus, the theoretical prediction for treatment intensity is quite ambiguous, and the empirical results will reflect all of these factors.

We find a significant increase in both the probability of having any procedure and the number of procedures. A 10 percentage-point increase in eligibility raises the probability of having a procedure by 6.6% (3.1 percentage points), and the mean number of procedures by 5.0% (.04 procedures per child). At the same time, there is a significant decline in length of stay: 3.1%, or 0.13 days per child. Thus, the Medicaid expansions appear to be leading to shorter stays for children in the hospital on average, but more intensive treatment per day when hospitalized. Absent cost data on procedures and hospital days, we cannot estimate the impact of Medicaid on the average cost of a child hospitalization; however, so long as the cost per procedure is less than 3.25 (=0.13/0.04) times the cost per day, the average cost will not increase.

## 5. Conclusions

Due to the relentless rise in the uninsured population in the U.S., public insurance policy is likely to remain a topic of considerable interest for some time. Improvements in the efficiency of medical care provided could theoretically offset some of the costs

associated with expanding coverage. However, few studies to date have empirically evaluated this proposition. Using a comprehensive, nationwide survey of hospital discharges, together with detailed data on Medicaid eligibility by age, state, and year, we identify the impact of Medicaid expansions on hospital utilization and the intensity of care provided in the hospital.

We find that over the 1983–1996 period, expansions in the Medicaid program increased the rate at which children were hospitalized by 13.4%, accounting for 6.5 additional hospitalizations per 1000 children. This increase is produced by a large increase in unavoidable hospitalizations, and a smaller, statistically insignificant increase in avoidable hospitalizations. These findings suggest that an efficiency effect may be acting to offset the access effect associated with the expansions. Nevertheless, our results imply that the Medicaid expansions are associated with more, rather than fewer, hospitalizations of children.

Given that we do not find strong evidence of cost decreases for the average hospitalization (the length of stay declined, but the number of procedures rose), we conclude that the Medicaid expansions were associated with large increases in hospital expenditures on behalf of children. Holding the average cost per child hospitalization constant at the 1996 rate for Medicaid enrollees, approximately US\$2418, the expansions increased hospital expenditures by nearly US\$1 billion annually.<sup>15</sup>

While we find that the Medicaid expansions are associated with increases in total hospitalization, and possibly even increases in avoidable hospitalization, these hospitalizations may contribute to the Medicaid-induced health improvements documented in earlier studies. Furthermore, we find evidence consistent with reduced avoidable hospitalizations, *holding access to the hospital constant*. Thus, the argument that improved access to outpatient care reduces avoidable hospitalizations may certainly hold. Because the Medicaid expansions simultaneously increase access to inpatient and outpatient care, we cannot evaluate this argument directly. Future research in this area will help policymakers to assess the longer-term effect of expanded coverage on health outcomes, hospitalizations, and expenditures.

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<sup>15</sup> Among Medicaid enrollees in 1996, average hospital expenditures per year per user of hospital services was US\$3627 (U.S. Department of Health and Human Services, 1998). The typical user of hospital services has 1.5 inpatient stays per year, so that the cost per stay is approximately US\$2418.

## Appendix A. The Medicaid expansions

**Deficit Reconciliation Act, 1984:** Effective October 1, 1984. Required states to extend Medicaid coverage to children born after September 30, 1983, if those children lived in families that were income-eligible for AFDC.

**Omnibus Budget Reconciliation Act, 1986:** Effective April 1, 1987. Permitted states to extend Medicaid coverage to children in families with incomes below the federal poverty level. Beginning in fiscal year 1988, states could increase the age cutoff by one year each year, until all children under age five were covered.

**Omnibus Budget Reconciliation Act, 1987:** Effective July 1, 1988. Permitted states to cover children under age 2, 3, 4, or 5, who were born after September 30, 1983. Effective October 1, 1988, states could expand coverage to children under age 8 born after September 30, 1983. Allowed states to extend Medicaid eligibility to infants up to one year of age in families with incomes up to 185 percent of the federal poverty level. States were *required* to cover children through age 5 in fiscal year 1989, and through age 6 in fiscal year 1990, if the families met AFDC income standards.

**Medicare Catastrophic Coverage Act, 1988:** Effective July 1, 1989, states were required to cover infants up to age one in families with incomes less than 75 percent of the federal poverty level. Effective July 1, 1990, the income threshold was raised to 100 percent of poverty.

**Family Support Act, 1988:** Effective April 1, 1990. States were required to continue Medicaid coverage for 12 months for families who had received AFDC in three of the previous six months, but who had become ineligible because of earnings.

**Omnibus Budget Reconciliation Act, 1989:** Effective April 1, 1990. Required states to extend Medicaid eligibility to children up to age 6 with family incomes up to 133 percent of the federal poverty line.

**Omnibus Budget Reconciliation Act, 1990:** Effective July 1, 1991. Required states to cover all children under age 19 who were born after September 30, 1983 and whose family incomes were below 100 percent of the federal poverty level.

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