HEALTH INSURANCE ELIGIBILITY, UTILIZATION OF MEDICAL CARE, AND CHILD HEALTH*

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We study the effect of public insurance for children on their utilization of medical care and health outcomes by exploiting recent expansions of the Medicaid program to low-income children. These expansions doubled the fraction of children eligible for Medicaid between 1984 and 1992. Take-up of these expansions was much less than full, however, even among otherwise uninsured children. Despite this take-up problem, eligibility for Medicaid significantly increased the utilization of medical care, particularly care delivered in physicians' offices. Increased eligibility was also associated with a sizable and significant reduction in child mortality.

Public concern about the health of children is one of the driving forces behind efforts to reduce the rate of uninsurance in the United States. High rates of child mortality and morbidity suggest that American children do not receive the same quantity or quality of health care as children in other developed countries. For example, at 8 per 1000, the infant mortality rate in the United States is the highest in the developed world. Compared with Canadian children, children one to four years of age in the United States also have a 14 percent higher mortality rate. And American children under fifteen years of age have 28 percent more disability days and 44 percent more bed days than Canadian children [Kozak and McCarthy 1984]. These problems are particularly acute among African-American children, whose child mortality rates are 63 percent higher than those of whites.

Since the uninsured are known to receive less health care than the insured, a potential explanation for this shocking state of affairs is that as many as 30 percent of poor children are without health insurance of any kind [Bloom 1990]. Hence, the debate over reforming the American health care system has increasingly

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emphasized health insurance for children [*Boston Globe,* September 12, 1994]. Absent from this debate, however, is convincing evidence that increased eligibility for public insurance will actually improve the health of children.

Increased eligibility may not translate into health improvements for two reasons. First, increases in eligibility do not automatically increase the utilization of medical care or the efficiency with which medical care is delivered. A large literature documents the fact that individuals do not always take up public assistance benefits for which they are eligible (e.g., Blank and Card [1991] and Blank and Ruggles [1993]. This problem may be more severe among groups who have not traditionally received public assistance. These groups may be both less well informed about public programs and more reticent to avail themselves of public "handouts." Furthermore, even if eligibles take up their benefits, many physicians do not treat publicly insured patients, possibly because public insurance programs generally reimburse at rates far below private fee levels.[1] This problem is exacerbated by the fact that many of the patients who are made eligible for public insurance are concentrated in areas that are underserved by physicians [Fossett and Peterson 1989; Fossett et al. 1992]. As a result, care may be delivered inefficiently to Medicaid patients, with excessive (expensive) visits to hospital outpatient departments and emergency rooms [Long, Settle, and Stuart 1986].

Second, increases in the utilization of care will not necessarily improve child health: for example, a number of studies suggest that much of the acute care received by children is inappropriate and may have little health benefit[2] Lurie et al. [1984] and Bindman et al. [1992] document positive effects of insurance on adult health, but a randomized trial [Newhouse 1993] suggested that increasing the generosity of insurance was of little benefit. However, the first two studies did not look at children, and there were too few children involved in the randomized trial for any firm conclusion regarding child health to be drawn [Valdez 1985].

Several recent papers have shown that children who are uninsured have lower utilization levels, a less efficient distribution of utilization across sites of care, and worse health outcomes

1. See Decker [1995] for evidence on the effect of Medicaid fee policy on physician treatment patterns, and Currie, Gruber, and Fischer [1995] for evidence on the relationship between infant mortality and Medicaid fee policy.

2. Kemper [1988], for example, finds that 21 percent of pediatric hospitalization days were of "doubtful necessity," and that this fraction is higher for insured than for uninsured children.

[Kasper 1986; Short and Lefkowitz 1992; Mullahy 1994]. But since the uninsured are likely to differ from the insured in both observable and unobservable respects, it is difficult to draw causal inferences from these types of comparisons. Furthermore, insurance coverage itself may be a function of health status, leading to endogeneity bias in estimates of the effects of insurance on health, and on the utilization of medical care.

In this paper we attempt to identify the effects of insurance coverage by drawing on dramatic recent expansions of Medicaid eligibility for low-income children. Medicaid is a federal-state matching entitlement program that provides health insurance to the poor. Historically, eligibility for Medicaid was tied to the receipt of cash welfare payments under the Aid to Families with Dependent Children (AFDC) program. Hence, eligibility was effectively limited to very low-income women and children in single-parent families. Beginning in 1984, states were first permitted and then required to extend Medicaid coverage to other groups of children. By 1992 states were required to cover children below age six in families with incomes up to 133 percent of the federal poverty line, and children between ages six and nineteen with family incomes up to 100 percent of the poverty line. States also had the option of covering infants up to 185 percent of the poverty line.[3] Since states have taken up these options at different rates, there is substantial variation in Medicaid eligibility thresholds by state, year, and age of child that can be used to identify the effects of the expansions.

This variation is used to address the effects of expanding public insurance eligibility for children during the 1984 to 1992 period. In Section I, data from the Current Population Survey (CPS) are used to measure increases in eligibility that resulted from state Medicaid policy changes over this era, and to examine the extent to which these increases in eligibility were translated into increases in Medicaid coverage.[4]

The main focus of the paper is on measuring the effects of Medicaid eligibility on the utilization of care using the National Health Interview Survey (NHIS), a large, nationally representa-

3. States received federal matching funds for coverage of these groups. However, some states have extended coverage to children above 200 percent of the poverty line, using state only funds.

4. As we emphasize in Currie and Gruber [1994], the effectiveness of eligibility expansions depends on the extent that the newly eligible take up their benefits. In that paper we show that many pregnant women who became eligible for Medicaid failed to take up their coverage.

tive data set. These data are described in more detail in Section II. We also present our identification strategy: demographic and economic information from the NHIS is used to impute Medicaid eligibility for each child, and then the fraction of children in the same state, age, and year who are eligible for Medicaid, calculated from the CPS, is used as an instrument for imputed individual eligibility. In this way, both the omitted variables and endogeneity problems that bias Ordinary Least Square (OLS) estimates are addressed.

In Section III this empirical strategy is used to measure the effect of Medicaid eligibility on medical utilization and on the site of care. In Section IV we extend our analysis to health outcomes, estimating the effect of Medicaid eligibility on state-level measures of child mortality. Section V concludes.

I. The Medicaid Expansions[5]

I.A. Legislative Background

Historically, Medicaid eligibility for children has been tied to participation in the Aid for Families with Dependent Children program (AFDC). This linkage with AFDC restricted access to the program in three ways. First, despite the existence of the AFDC-Unemployed Parents program (AFDC-UP) which provides benefits to households in which the primary earner is unemployed, AFDC benefits are 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 four in South Carolina was only 29 percent of the poverty line. Third, the stigma of applying for cash welfare programs may have prevented eligible families from receiving Medicaid benefits [Moffitt 1992].

In some states, children could also qualify for Medicaid under state Medically Needy or Ribicoff programs. The Medically Needy program relaxed the income criteria for eligibility by cov-

5. There is a rapidly growing literature that examines the effect of this dramatic shift in public insurance policy in the United States. Cole [1994]; Currie and Gruber [1994]; Piper, Riley, and Griffin, [1990]; Singh, Gold, and Frost [1994]; and Torres and Kenney [1989] examine the effect of expanded Medicaid eligibility for coverage of pregnancies on take-up and on infant health outcomes. Yelowitz [1995a, 1995b] examines the effect of expansions for children on the welfare participation and marital status of their mothers. Cutler and Gruber [1996] examine the extent to which increased public insurance crowded out private insurance coverage.

ering people who would have been eligible for AFDC if their incomes were lower, but who had large medical expenditures that brought their "net 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 has been 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 2. The important point to note is that states took up these options at different rates, so that there was a great deal of variation across states in both the income thresholds and the age limits governing Medicaid eligibility. This variation is documented below.

I.B. Effects on Eligibility

A natural first question about these Medicaid policy changes is whether they had a significant effect on the fraction of the population eligible for the program. To answer this question, we use data from the CPS March surveys for 1985 to 1993. The CPS collects information on demographic characteristics as well as income and labor force data for the previous year, so that our analysis will cover the 1984 to 1992 period. Eligibility is imputed to each child in the CPS using the algorithm described in Appendix 1.[6] For this calculation and throughout the paper, children are defined as those less than fifteen years old, in order the avoid complications due to children having children.[7]

The results of our calculation at the national level are sum-

6. The CPS has the advantage of being the largest nationally representative cross section of data over time. The disadvantages relative to a data set like the Survey of Income and Program Participation (SIPP) are annual rather than monthly income, and a lack of information about assets. However, in their study of the Job Training and Partnership Act, Devine and Heckman [1994, p. 56] conclude that despite the use of annual income, "CPS approximations to eligibility are actually very good." Further, in their analysis of take-up welfare programs using the SIPP, Blank and Ruggles [1993] find that applying the asset test in the SIPP does not much affect AFDC take-up rates.

7. Due to the way that the expansions were phased in, pregnant teens could have been eligible for Medicaid coverage of their pregnancies even if family income was too high for the teen to be eligible otherwise.

TABLE I
MEDICAID ELIGIBILITY AND COVERAGE

Year	% of children eligible	% of children eligible— fixed population	% of children covered
1984	16.1	16.1	13.2
1985	18.2	18.4	13.5
1986	19.0	18.9	13.8
1987	19.3	19.7	13.5
1988	18.8	20.3	12.8
1989	20.4	21.6	13.9
1990	25.7	26.2	16.5
1991	28.7	28.1	19.3
1992	31.2	30.3	20.6

Based on data from March 1985–March 1993 CPS. Column 1 shows the percent of children eligible for Medicaid in each year. Column 2 shows the percentage of the 1984 sample that would have been eligible for Medicaid in each subsequent year (holding their characteristics constant and inflating income appropriately). Column 3 gives the percentage of children actually covered in each year. Figures are from the authors' calculations as described in the text and in Appendix 1.

marized in the first column of Table I. There was a dramatic increase in Medicaid eligibility over this period. Eligibility doubled between 1984 and 1992, and almost one-third of all children in the United States were eligible by the end of the period. Some of the increase came from DEFRA and other state law changes between 1984 and 1987, but the bulk came from the expansions to higher income groups between 1987 and 1992. Some of this later increase, of course, resulted from the recession at the end of the period, which lowered family incomes and made more children eligible at each level of legislative generosity. In order to separate business cycle effects from the effects of legislative change, we also estimate the fraction of the 1984 population that would be eligible under each year's laws. These estimates are shown in column 2. The two columns are very similar, suggesting that most of the increase in eligibility was the result of legislative changes.

In addition to the time series variation in eligibility, there is substantial heterogeneity across the states, as shown in Appendix 3.[8] This table lists the fraction eligible in each state in 1984 and 1992, as well as the change over the period. While eligibility rose substantially nationwide, it actually fell in two states, Pennsylvania and Wisconsin. When states are ranked by the fraction

8. Since Arizona did not have a formal Medicaid program for most of this period, we have excluded children in Arizona from the analysis.

covered, the correlation between the rankings in 1984 and 1992 is only 0.16. The median state moved fourteen positions in this ranking. This heterogeneity reflects both changes in state policy and state-specific changes in conditions. Thus, in the last three columns we once again provide eligibility estimates calculated holding population characteristics constant at their 1984 levels. The results are very similar. Nevertheless, we present a methodology below for identifying our models using the state policy changes only, in order to control for state economic and demographic conditions.

I.C. Effects on Coverage

As highlighted in the introduction, increases in Medicaid eligibility do not automatically translate into increases in insurance coverage. In order to examine the take-up of Medicaid coverage by newly eligible children, we use a question about Medicaid coverage from the March CPS.[9]

The second column of Table I shows Medicaid coverage over time. Overall, coverage did rise, but not nearly so steeply as eligibility. The increase in coverage was approximately one-half as large as the increase in eligibility. Coverage was actually flat from 1984 to 1988, and rose sharply thereafter. This time trend follows the business cycle, suggesting that the increase in coverage cannot be automatically attributed to changes in Medicaid policy. In order to control for the business cycle, as well as for observable nonpolicy-related determinants of individual eligibility, we estimate models of Medicaid coverage that include year dummies and state dummies, as well as controls for race, gender, mother's education (high school dropout, high school graduate only, some college), family income ($10,000 intervals up to $50,000), child age, and state.[10]

9. The question asks whether the individual was covered by Medicaid at some point in the previous year: We measure Medicaid eligibility using income information from the previous year. There may be timing problems in our measures of individual eligibility, since income fluctuations during the year can make an individual eligible at one point during the year, even if they are ineligible using average annual income. Approximately 20 percent of those who report Medicaid coverage are deemed ineligible by our imputation procedure. Note also that this information became much more reliable after 1986, since surveys before March 1988 did not gather complete information about insurance coverage obtained through someone other than the head of the household. The results presented below, however, are not sensitive to the restriction of the sample to 1987 onward.

10. The results are very similar if the regressions are estimated as probit or logit models instead. We use mother's education because almost every child in our sample has a mother present but a sizable fraction do not have a father present.

TABLE II
TAKE-UP OF MEDICAID ELIGIBILITY IN THE CPS
LINEAR PROBABILITY MODELS: COEFFICIENTS TIMES 10^2

	(1) OLS	(2) TSLS
Medicaid	29.77	22.73
eligibility	(0.019)	(1.109)
Female	−0.017	−0.003
	(0.098)	(0.098)
Black	9.323	9.555
	(0.162)	(0.166)
Mom is high	−6.218	−6.570
school graduate	(0.131)	(0.142)
Mom has some	−7.405	−7.867
college	(0.166)	(0.180)
Income >	−19.09	−23.15
$10,000	(0.191)	(0.659)
$10,000 < Inc	−22.36	−27.62
< $20,000	(0.217)	(0.845)
$20,000 < Inc	−23.37	−28.72
< $30,000	(0.228)	(0.859)
$30,000 < Inc	−23.28	−28.60
< $40,000	(0.250)	(0.863)
$40,000 < Inc	−23.12	−28.44
< $50,000	(0.250)	(0.862)
R^2	0.395	...
Number of obs.	324,821	324,821

Standard errors are in parentheses. All regressions include a full set of age, state, and year dummies. Eligibility is instrumented using simulated eligibility, matched to individuals by state, year, and age groups, as described in Section II.

The estimation results are presented in Table II. As the first column shows, making a child eligible for Medicaid over the 1984 to 1992 period increased the probability that he or she was covered by insurance by approximately 30 percent. This figure is substantially below the take-up rates estimated for other social programs, but it is not strictly comparable, since we are calculating the marginal effect of a policy change, rather than the average take-up for a given level of eligibility.

These regressions may be subject to some remaining omitted variables bias: For example, a recession in a state will cause both eligibility and coverage to rise simultaneously. This problem is addressed by instrumenting eligibility in the second column of Table II. The instrument that we use measures the legislative generosity of Medicaid policy in a given state in a given year, for each child's age group. This instrument is described at greater length in Section II. The key point is that instrumental variables methods purge the regression of bias due to unobserved individual-level characteristics as well as omitted variables such as economic conditions in a particular state and year. As expected, instrumenting causes our estimate to fall somewhat, yielding a marginal take-up rate of only 23 percent.

One reason for these low take-up rates may be that a number of the children made eligible for Medicaid already had private health insurance. Among those who were ineligible for Medicaid in 1984, but who would have been made eligible by the 1984 to 1992 rule changes, 32 percent were uninsured. Thus, the take-up rate among the otherwise uninsured may be as high as 71 percent (22.7/32). This take-up rate is an upper bound, however, since some of those who take up Medicaid may have previously been privately insured, a point discussed at length by Cutler and Gruber [1996].[11]

Thus, while the Medicaid expansions doubled the fraction of children eligible for public insurance, the increase in the number of children covered was substantially lower. This result suggests that take-up may be an important barrier to the effectiveness of the Medicaid expansions. Still, the coverage response is not trivial, and does leave open the possibility that the expansions had significant effects on utilization and health.

II. NHIS DATA AND EMPIRICAL FRAMEWORK

II.A. NHIS Data

The National Health Interview Survey (NHIS) interviews a large, nationally representative cross section of American families each year. The baseline survey collects information about demographic characteristics, labor force attachment, and family income. There are also a number of questions about the utilization of medical care over the previous two weeks and the previous year. These data cover approximately 30,000 children per year, for a total 1984 to 1992 sample of over 225,000.

11. Cutler and Gruber estimate that among children, private health insurance coverage dropped by one child for every three children newly covered. If all of these children enrolled in Medicaid, then the take-up rate among the otherwise uninsured would be 47 percent.

Our first step is to assign eligibility to each child in the survey. This is more difficult in the NHIS than in the CPS for two reasons. First, family income is reported in brackets and is missing for a number of households. Missing income data are imputed by estimating yearly regressions of income on household characteristics in the CPS, and then using the regression coefficients to calculate income for NHIS households with similar characteristics in that year.[12] Incomes within a bracket are imputed by randomly choosing a point within the bracket. Second, there is no information on the distribution of income across family members or across income sources. This is problematic because, for example, some portion of earnings, but not other types of income, can be disregarded from total family income in determining AFDC eligibility. We apply those disregards to total income, under the assumption that most family income comes from earnings.[13] These data problems do not seem to lead to any systematic measurement problems. The resulting annual eligibility rate in the NHIS is similar to the CPS in both the level and the time series trend.[14]

Our key variables in the NHIS are measures of health care utilization. A potential problem with utilization measures, however, is that they confound access and morbidity. For example, the Medicaid expansions may have increased access to hospitals, but at the same time they could have increased the use of preventive care, improving health status and reducing the demand for hospital care. One way to surmount this problem is to focus on utilization that is explicitly preventative, and therefore unaffected by morbidity. Pediatric guidelines recommend at least one doctor's visit per year for most children in our sample, so that the absence of a doctor's visit in the previous year is suggestive of a true access problem, regardless of underlying morbidity. This is therefore the measure on which we focus in the analysis.

Two other measures of contacts with the medical system are

12. For most of the missing observations, we know whether income was greater than or less than \$20,000, so we can impute income within those subsamples. These imputation regressions fit fairly well. For those with incomes below \$20,000, the R^2's average 0.32. While for those with incomes above \$20,000, the R^2's average 0.25. Note that the Census bureau uses a similar procedure to impute missing income data in the CPS.

13. In the 1984 CPS, 75 percent of the average child's family income comes from his or her parents' earnings.

14. In the years from 1984 to 1992, the fraction of children eligible for Medicaid in the NHIS data was 14.5, 17.2, 19.1, 19.2, 18.5, 19.3, 25.2, 27.5, and 31.5 percent. These numbers are very close to those shown in Table I.

examined: the probability of having had a doctor's visit in the past two weeks; and the probability of having had a hospitalization in the previous year.[15] Both of these measures suffer from the bias noted above. Nevertheless, the former is useful in assessing the extent that Medicaid affects not only the probability of any contact, but also the frequency of contacts. The latter is useful in assessing whether Medicaid increases the utilization of providers other than physicians. Since the bias to these estimates from improved health status is obviously downward, positive findings are suggestive of true utilization effects.

Concern over the utilization of the uninsured arises not only from their low number of medical contacts, but also from inefficiencies in their patterns of utilization. For example, Aday and Anderson [1984] report that among those with a regular source of care in 1982, 87 percent of the privately insured used a doctor's office as their regular source, and only 8 percent regularly used hospital outpatient departments or emergency rooms. Among the uninsured, however, only 74 percent regularly used a doctor's office, and 18 percent relied on hospitals for their regular source of care.

High hospital utilization rates among the uninsured are of concern for two reasons. First, hospital visits are more expensive. McDevitt and Dutton [1989] found that conditional on the illness, treating AFDC children in hospitals was 66 percent to 107 percent more expensive than treating them in physician's offices: Stuart [1990] et al. find that costs are 50 percent higher in hospital settings, controlling for diagnosis. Second, care may be delivered more sporadically at settings other than the physician's office. As McDevitt and Dutton [1989, p. 44] note, "it is widely accepted that having a regular source of care is an important component of quality care." They find that continuity of care, as measured by the percentage of visits to a single primary provider, is lower for those whose primary provider is a hospital rather than a physician. Furthermore, those with less continuity of care have higher expenditures.

If a child had a visit in the last two weeks, the NHIS reports the site of care. We distinguish between three different sites: the physician's office; a hospital emergency room or outpatient clinic;

15. The NHIS includes telephone contacts as "visits." We exclude telephone contacts in our analysis since it is difficult to know how to interpret them. The results for number of visits in the last two weeks are somewhat stronger if we include telephone contacts.

and other sites, primarily private or public clinics.[16] While it appears clear that care is delivered more efficiently in a physician's office than in a hospital, there is less of a consensus with respect to the relative efficiency of care at other sites. Gold [1981] and Gold and Greenlick [1981] found no differences in outpatient costs between hospital-based and freestanding HMO clinics. The probability of inpatient admission was higher in the former setting, however, so that total costs were higher. Stuart et al. [1990] estimated that total payments per year for Maryland Medicaid patients were approximately twice as high as hospital outpatient departments than at clinics. Clinic costs were approximately equal to office-based physician costs, although the probability of inpatient admission was higher. There is no evidence, however, about continuity of care in clinics versus other sites. Thus, the efficiency implications of shifting visits to and from clinics is less clear than the implications of shifting to physicians and from hospitals.

The means of the NHIS data are presented in Table III. We estimate that, across all years, 22 percent of the NHIS sample were eligible for Medicaid. Not surprisingly, Medicaid eligibles are more disadvantaged than the sample as a whole along a number of dimensions: they have less educated mothers and fathers; they are more likely to be minorities; they are more likely to live in a female-headed household; and they are substantially poorer. Despite these disadvantages, they use medical care at almost exactly the same rate as noneligibles. In fact, they are hospitalized more frequently. This may reflect the fact that utilization confounds morbidity and access, as noted above. Consistent with previous studies, we also see that Medicaid eligibles are less likely to have visited a doctor's office in the past two weeks, and are more likely to have visited a hospital. They are also somewhat more likely to have used another site of care.

II.B. Empirical Strategy

We begin our analysis by estimating linear probability models of the effect of Medicaid eligibility on utilization of the form,[17]

16. We include in the first category a small number of visits that took place in the patient's home. We also include a small number of visits to "doctors' offices in hospitals" in the "other" category since we wish to focus on the distinction between doctors' offices and hospital-based care in our analysis.

17. We use linear probability models for ease of computation and for consistency of our instrumental variables procedure. Heckman and MaCurdy [1985] show that this procedure produces consistent estimates. We correct the standard errors for heteroskedasticity using a White [1984]-type procedure.

TABLE III
NHIS SAMPLE MEANS BY MEDICAID ELIGIBILITY

	All	Medicaid eligible
% Eligible	0.219	1.000
	(0.001)	
# Observations	227,169	49,991
Utilization of care		
No doctor's visits last 12 months	0.194	0.197
	(0.001)	(0.002)
Doctor's visit last 2 weeks	.115	0.118
	(0.001)	(0.001)
Any hospitalization last 12 months	0.036	0.049
	(0.0004)	(0.001)
Visit to doctor's office last 2 weeks	0.087	0.071
	(0.0006)	(0.001)
Visit to ER or hospital clinic last 2 weeks	0.017	0.027
	(0.0003)	(0.001)
Visit to other site of care last 2 weeks	.015	0.024
	(0.0003)	(0.001)
Family & child characteristics		
Male	0.513	0.509
	(0.001)	(0.002)
Black	0.180	0.355
	(0.001)	(0.002)
Hispanic	0.120	0.219
	(0.001)	(0.002)
Age	6.873	5.401
	(0.001)	(0.019)
Female head/spouse is HS dropout	0.240	0.506
	(0.001)	(0.002)
Female head/spouse has some college	0.355	0.133
	(0.001)	(0.002)
Male head has some college	0.342	0.079
	(0.001)	(0.001)
Child is oldest	0.540	0.429
	(0.001)	(0.002)
Number of siblings	1.304	1.800
	(0.002)	(0.006)
No male head	0.221	0.505
	(0.001)	(0.002)
Mom is respondent	0.295	0.588
	(0.001)	(0.002)
Dad is respondent	0.694	0.411
	(0.001)	(0.002)
Other female relatives	0.043	0.064
	(0.0004)	(0.001)

TABLE III
(CONTINUED)

	All	Medicaid eligible
Other male relatives	0.028	0.041
	(0.0004)	(0.001)
Family income < 10,000	0.137	0.532
	(0.001)	(0.002)
10,000 < Inc < 20,000	0.181	0.250
	(0.001)	(0.002)
20,000 < Inc < 30,000	0.174	0.055
	(0.001)	(0.001)
30,000 < Inc < 40,000	0.145	0.006
	(0.001)	(0.0004)
40,000 < Inc < 50,000	0.099	0.0005
	(0.0006)	(0.0001)
Family income > 50,000	0.138	0
	(0.001)	(0)
Central city	0.260	0.413
	(0.001)	(0.002)
Rural	0.288	0.268
	(0.001)	(0.002)

The source is authors' tabulations of NHIS data. Standard errors are in parentheses.

$$(1)\quad UTIL_i = \alpha + \beta_1 X_i + \beta_2 ELIG_i + \beta_3 \delta_j + \beta_4 \tau_t + \beta_5 AGEG_i \times \delta_j + \beta_6 AGEG_i \times \tau_t + \varepsilon_i,$$

where

$UTIL_i$ is a measure of utilization for individual i,
X is a set of control variables,
$ELIG$ is an indicator of the eligibility of individual i for Medicaid,
δ_j and τ_t are a full set of state and year dummies, respectively,
$AGEG$ is a dummy for being in one of five age groups.

These OLS estimates, however, are subject to three sources of bias. The first is omitted variables bias. In all of our regressions we control for observable variables that directly determine eligibility for Medicaid. These include income, the absence of a male head, the number of children in the family, and the age of the child (through single year age of dummies).[18] The income

18. But note that we are unable to control for all of the possible interactions of these variables that might determine eligibility. This is a further argument for IV estimation.

categories reported by the NHIS are used to control for income. In the omitted category are those with missing information on income.[19] We also control for the child's gender, race, and ethnicity, whether he or she is the oldest child, the number of siblings, the education of the mother and (if present) the father, whether the mother or father was the respondent, the presence of other relatives, and whether the family lives in a central city or rural area. Even after conditioning on this detailed set of controls, however, persons who are eligible for Medicaid may have other characteristics that affect both their utilization and their health. For example, they may be more likely to live in areas with limited access to physicians.

A second problem is endogeneity bias. A sick child may cause lower parental income (if a parent is forced to leave work to care for the child, for example), leading to a spurious positive correlation between Medicaid eligibility and utilization. Finally, there may be substantial measurement error in our eligibility indicator, given the limitations of the NHIS income data.

Hence, in addition to the OLS estimates, we present instrumental variables estimates. Our aim is to abstract from characteristics of the child or family that may be correlated with both eligibility and the dependent variables and to achieve identification using only legislative variation in Medicaid policy. One way to do this would be to instrument imputed individual eligibility in the NHIS using the fraction of children in the same state, year, and age who are eligible, calculated from the CPS. This instrument would capture differences in Medicaid eligibility across states, years, and age groups, but would purge the regression of individual-level sources of variation in eligibility.

This approach runs into two problems in practice, however. First, the CPS is simply not large enough to permit reliable estimation of the fraction of children eligible in each state, year, and age category. Second, these estimates could be biased by the omission of characteristics of state, year, and age groups that are correlated with both eligibility and with utilization. For example, if infants in a given state and year were particularly poor, they might have both higher eligibility levels and fewer doctor's visits, resulting in a downward bias to our estimates of the effects of eligibility on utilization.

19. We imputed income for the purposes of determining Medicaid eligibility, as described above. However, we only use reported income categories as control variables in the regressions.

Our strategy, therefore, is to use a "simulated instrument" that varies only with the state's legislative environment and not with its economic or demographic characteristics. In order to construct this instrument, we select a national random sample of 300 children of each age (zero to fourteen), in each year, and calculate the fraction of children in this sample who would be eligible for Medicaid given the rules in each state in that year.[20] This measure can be thought of as a convenient parameterization of legislative differences affecting children in different state, year, and age groups—a natural way to summarize the generosity of state Medicaid policy as it affects each group is in terms of the effect it would have on a given, nationally representative, population.

This instrumental variables strategy addresses the econometric difficulties noted above. First, by using instruments that are arguably exogenous to utilization, we purge the model of endogeneity bias. Second, by using the fraction of children eligible in each child's state, year, and age group, we abstract from any individual-level omitted variables correlated with both eligibility and outcomes. Third, to the extent that the measurement error in our instrument is uncorrelated with the measurement error in our individual eligibility measure, we also surmount the measurement error problem.[21] Finally, by using a national random sample to construct the instrument, we purge it of the effects of state and year-specific economic conditions that might be correlated with both eligibility and utilization. We also overcome the small sample problems that would arise in carrying out his IV strategy using the actual fraction eligible. This instrument is strongly correlated with individual eligibility. The first-stage F

20. The sample size of 300 was chosen due to data and computational constraints. As an alternative, we also tried computing our imputed eligibility measures using samples drawn from each state's region rather than the national sample. By using variation in Medicaid eligibility that is associated with a region's demographic and economic characteristics, we could develop more precise instruments and increase the efficiency of our estimates. The cost is that demographic and economic characteristics of regions may not be exogenous to models of utilization. In fact, we found that this sampling scheme changed our estimates somewhat without appreciably affecting our standard errors. Hence, we retained the somewhat less efficient but consistent national random sample approach.

21. If the measurement error stems mainly from random individual response error, then measurement error in our CPS instrument will be uncorrelated with that in our NHIS data, especially given the fact that the measure we calculate using the CPS is the average eligibility for a large group. However, if measurement error comes primarily from inaccuracies in the program we use to impute Medicaid eligibility, then we will have similar kinds of errors in individual eligibility measures calculated using either the NHIS or the CPS data.

statistic is approximately 10,000. This is the instrument that was used in the two-stage least squares (TSLS) results shown in Table II.

The only remaining potential problem is that of omitted variables that are correlated with changes in state Medicaid policy, and also with differences in utilization or health. For example, rich states may have both more generous Medicaid policy and more medical utilization. Alternatively, expansions may have been phased in first for age groups with the highest exogenous growth in utilization over time. Any such omitted variable will invalidate our IV strategy, which treats the legislative environment for the state, year, and age group as a source of exogenous identification.

In order to control for these effects, we include a full set of dummy variables for states, years, and each single year of age. Interactions between four age groups and calendar year, and between the same four age groups and states are also included.[22] These interactions control for changes in the utilization patterns of different age groups over time, as well as for fixed differences across states in the utilization patterns of different age groups. As a result, in our final model, identification comes from two sources: changes within states over time; and changes within state-age groups over time. Both of these sources of variation are plausibly exogenous to changes in child health care utilization and health outcomes.

In principal, it would be possible to include interactions between states and years in the models, and to achieve identification from changes across age groups within states and years.[23] When state and year interactions are included, the point estimates are similar to those presented below. However, including these variables removes much of the interesting legislative variation in the data. In the early years of the sample, changes in state policy (i.e., through the Ribicoff program) affected children of all ages, so that this variation is lost when state and year interactions are included. And, even though there are "age notches" in

22. The four age groups are zero to one, two to four, five to seven, and eight to ten. Eleven plus is the omitted age group.

23. This would then be a true "difference-in-difference-in-difference" estimator, as employed by Gruber [1994] or Yelowitz [1995a]. The effects of demographic and economic characteristics of state and years have already been purged by using our simulated instrument. Interactions between state and years would only affect our estimates if, conditional on fixed characteristics of states and years, there were characteristics of a particular state and year that were correlated with both the legislative environment and with the utilization of care.

the expansions, much of the change in policy in this period came from secular shifts in the generosity of state policy. As a result, including state and year interactions increases the standard errors by about one-half. They are therefore excluded from the basic analysis.[24]

III. Medicaid Eligibility and Utilization

Models of the utilization of medical care are presented in Table IV. The first three columns show the results from linear probability models estimated by OLS. The coefficients are multiplied by 100 and can therefore be interpreted as percentage point effects. Medicaid is found to significantly decrease (by 2.5 percentage points) the probability of going without a visit in the previous year, but to have no statistically significant effect on the probability of having had a visit in the past two weeks.[25] Compared with the baseline probability of going without a visit for Medicaid eligibles, making a child eligible for Medicaid lowers the probability of going without a visit by 12.8 percent. This estimate is almost certainly biased downward by omitted characteristics correlated with both eligibility and an elevated risk of going without a visit. Hospitalizations are also found to rise significantly, which represents a 14 percent rise relative to the baseline hospitalization rate. As discussed above, this estimate could be biased either upward or downward by omitted variables such as underlying morbidity.

TSLS estimates are shown in the next three columns of Table IV, and suggest that in fact the OLS results are biased downward. Being made eligible for Medicaid is now associated with a 9.6 percentage point drop in the probability of going without a visit last year. This is almost one-half of the baseline probability. That is, our TSLS findings suggest that making children eligible for

24. Relative to the results reported below, including state and year interactions causes the no visit last year coefficient to fall by about one-third and the visit in the past two weeks and visit to a doctor's office coefficients to rise by about one-half. The other coefficients are unchanged. In all cases, however, the standard errors rise by about one-half, so that none of the results differ from those reported below by even one standard error.

25. Information about visits in the past two weeks and about hospitalizations were collected in special supplements to the main questionnaire and a great deal of effort was made in these supplements to avoid missing data. In contrast, the number of visits last year is from the main questionnaire and has more missing values, leading to a somewhat smaller sample size for whether there were any visits in the past year.

Medicaid lowers the probability that they go without a visit over a year period by almost half. The point estimate on the probability of having had a visit in the last two weeks becomes positive and sizable, although it is not statistically significant. Finally, there is a very large and significant four percentage point rise in the probability of a hospitalization in the previous year. This coefficient implies that becoming eligible for Medicaid almost *doubles* the probability of being hospitalized.

The control variables suggest some interesting differences in utilization patterns across demographic groups. Blacks and hispanics use less medical care by all three measures—the differences are particularly large for blacks. There are also clear differences related to parental education. Children whose parents dropped out of high school are less likely to have had any visits in the past twelve months, and less likely to have had a visit in the past two weeks. Conversely, parents with some college education are more likely to take their children to the doctor. The dropout effects are similar for mothers and fathers, while the college effect is stronger for mothers.

Utilization is higher for both first children and those in smaller families: the former may reflect parental diligence with respect to scheduling the first child's checkups that is relaxed for later children, while the latter may reflect the classic Becker [1981] child quality/quantity trade-off. A surprising finding is that, conditional on household income, children in households with no male head have higher utilization levels.[26] Having additional relatives in the house also reduces children's utilization of health care, and the effect is greater for male relatives than for female relatives. Finally, conditional on family structure, fathers report lower utilization rates than mothers, which may reflect inferior recall of the child's actual pattern of use.

The income patterns for visits suggest that, in keeping with the previous literature, visits are normal goods. On the other hand, hospitalizations appears to be an inferior good. Finally, physician utilization is slightly higher in inner cities, but much lower in rural areas. On the other hand, hospitalizations are much higher in rural areas. This last result is consistent with the findings of Goodman et al. [1994], who show that the supply of

26. This result is consistent with evidence from many countries that ceteris paribus resources in the hands of the mother have a greater impact on child well-being than resources in the hands of the father. See Thomas and Strauss [forthcoming] for a review of this literature.

TABLE IV
MEDICAID ELIGIBILITY AND THE UTILIZATION OF MEDICAL CARE
LINEAR PROBABILITY MODELS: COEFFICIENTS × 10^2

Dependent var	(1) OLS No visit last year	(2) OLS Visit last 2 weeks	(3) OLS Hospital last year	(4) TSLS No visit last year	(5) TSLS Visit last 2 weeks	(6) TSLS Hospital last year
Medicaid	−2.510	−0.119	0.681	−9.553	4.853	3.960
eligibility	(0.309)	(0.237)	(0.153)	(3.037)	(2.803)	(1.646)
Male	−0.034	0.691	0.763	−0.033	0.691	0.763
	(0.159)	(0.132)	(0.078)	(0.159)	(0.132)	(0.078)
Black	4.149	−3.354	−0.611	4.362	−3.505	−0.710
	(0.260)	(0.195)	(0.123)	(0.276)	(0.212)	(0.133)
Hispanic	1.738	−0.922	0.019	1.978	−1.093	−0.093
	(0.294)	(0.234)	(0.140)	(0.311)	(0.254)	(0.150)
Mom is HS	2.809	−0.613	0.264	3.255	−0.927	0.057
dropout	(0.246)	(0.180)	(0.118)	(0.316)	(0.252)	(0.157)
Mom has some	−3.098	1.177	−0.263	−3.269	1.298	−0.183
college	(0.197)	(0.175)	(0.098)	(0.210)	(0.188)	(0.064)
Dad is HS	3.069	−0.832	−0.216	3.365	−1.041	−0.354
dropout	(0.296)	(0.212)	(0.137)	(0.323)	(0.243)	(0.154)
Dad has some	−2.392	0.672	−0.252	−2.378	0.662	−0.258
college	(0.223)	(0.191)	(0.111)	(0.223)	(0.192)	(0.109)
Child is oldest	−2.540	0.990	−0.049	−2.372	0.872	−0.127
	(0.197)	(0.157)	(0.092)	(0.210)	(0.171)	(0.099)
Number of	1.610	−0.640	−0.234	2.111	−0.936	−0.430
siblings	(0.095)	(0.066)	(0.040)	(0.204)	(0.175)	(0.105)
No male head	−5.243	2.195	0.618	−4.985	2.012	0.498
	(0.395)	(0.315)	(0.196)	(0.410)	(0.332)	(0.204)
Mom is	−0.214	1.434	−0.445	0.027	1.264	−0.556
respondent	(0.569)	(0.541)	(0.349)	(0.579)	(0.549)	(0.352)

Dad is	1.141	−0.797	−0.826	1.320	0.653	−0.910
respondent	(0.625)	(0.589)	(0.375)	(0.631)	(0.593)	(0.375)
Other female	0.492	−0.211	−0.229	0.675	−0.341	−0.314
relative	(0.410)	(0.032)	(0.193)	(0.417)	(0.331)	(0.197)
Other male	2.548	−0.797	−0.056	2.626	−0.852	−0.093
relative	(0.516)	(0.396)	(0.248)	(0.517)	(0.397)	(0.248)
Income < 10,000	−0.947	2.057	0.705	2.635	−0.472	−0.962
	(0.399)	(0.289)	(0.194)	(1.588)	(1.450)	(0.854)
10,000 < Inc	−0.572	0.813	0.430	−0.412	0.710	0.362
< 20,000	(0.322)	(0.230)	(0.145)	(0.327)	(0.237)	(0.149)
20,000 < Inc	−2.754	1.245	0.547	−3.908	2.060	1.084
< 30,000	(0.319)	(0.239)	(0.145)	(0.590)	(0.515)	(0.305)
30,000 < Inc	−4.623	1.821	0.454	−6.083	2.851	1.134
< 40,000	(0.330)	(0.258)	(0.150)	(0.708)	(0.635)	(0.370)
40,000 < Inc	−6.135	2.444	0.334	−7.609	3.484	1.020
< 50,000	(0.354)	(0.293)	(0.162)	(0.725)	(0.656)	(0.380)
Family income	−6.961	3.232	0.200	−8.497	4.316	0.915
> 50,000	(0.337)	(0.285)	(0.152)	(0.741)	(0.673)	(0.388)
Central city	−1.135	0.222	−0.200	−0.956	0.096	−0.283
	(0.213)	(0.179)	(0.102)	(0.226)	(0.193)	(0.111)
Rural	2.308	−0.207	0.736	2.370	−0.250	0.707
	(0.232)	(0.185)	(0.118)	(0.233)	(0.187)	(0.119)
R^2	0.099	0.044	0.025	...	...	...
Number of obs.	226,603	227,169	227,169	226,603	227,169	227,169

Standard errors are in parentheses. All regressions also include an intercept; dummy variables for each state, calendar year, and year of age; season dummies; interactions between calendar yer and year of age dummies; and interactions between year of age and state dummies. Eligibility is instrumented using simulated eligibility calculated using the CPS, and matched to individuals by state, year, and age. Standard errors are corrected for heteroskedasticity.

outpatient alternatives has a strong effect on pediatric hospitalization rates.

Thus, being made eligible for Medicaid has sizable and significant effects on the utilization of medical care. This conclusion emerges both from OLS models and (even more strongly) from TSLS models. These effects are quite large relative to other influences on the utilization of care. For example, the effect of Medicaid eligibility on the probability of going without a visit last year is larger than the effects of moving family income from the \$10,000 to \$20,000 range to the greater than \$50,000 range.

III.A. Site of Care

The above estimates indicated that while eligibility reduced the number of children going without a visit, it also increased hospitalizations. This pattern raises the possibility that care is provided inefficiently to Medicaid patients. One reason for increased hospitalization rates in response to the Medicaid expansions is that hospitals may be better equipped to assist patients in claiming benefits. Potential eligibles for Medicaid must complete lengthy and complex application forms, provide extensive documentation (such as birth certificates, pay stubs, and confirmation of child care costs), and attend several interviews with caseworkers. About one-third to one-half of all Medicaid applications are denied, and half of these denials are because the applicant did not complete all of the necessary steps. In response, many hospitals have established special offices, or contract with private companies, to assist Medicaid eligibles in completing these procedures [U. S. General Accounting Office 1994]. The nontrivial costs of providing these services may be beyond the means of private doctors and clinics, leading them to recommend that potential eligibles seek care in a hospital setting.

In an effort to see whether the eligibility expansions led to an increase in visits to doctors' offices (the most cost-effective site of care), we estimated models with the dependent variable equal to one if the child attended any one of three sites in the last two weeks. These estimates are shown in Table V.[27] Although we were

27. An alternative estimation strategy would involve restricting the sample to those who had a visit in the past two weeks, and then examining the site of care. However, since the probability of having a visit is affected by Medicaid eligibility, restricting the sample in this way could lead to the inference that Medicaid was causing a behavioral shift in the site of care when in fact there was simply a compositional change in the pool of persons with a visit. An additional complica-

TABLE V
MEDICAID ELIGIBILITY AND THE SITE OF CARE
ALL REGRESSIONS RUN AS INSTRUMENTAL VARIABLES
MEDICAID ELIGIBILITY COEFFICIENT AND MEANS ARE × 100

	(1) Doctor's office	(2) ER or hospital outpatient clinic	(3) Other site
Medicaid eligibility	5.073	1.174	−1.217
	(2.479)	(1.117)	(1.100)
Mean of dependent var	8.707	1.666	1.473
Number of obs.	227169	227169	227169

Standard errors are in parentheses. All regressions also include all the variables listed in Table V, as well as an intercept; dummy variables for each state, calendar year, and year of age; season dummies; interactions between calendar year and year of age dummies; and interactions between year of age and state dummies. Eligibility is instrumented using simulated eligibility calculated from the CPS, and matched to individuals by state, year, and age. Standard errors are corrected for heteroskedasticity.

unable to detect a statistically significant effect of eligibility on the probability of a visit in the past two weeks in Table IV, we do find a significant effect on visits to doctors' offices. The probability of a visit to the doctor's office rises five percentage points, which is roughly two-thirds of the baseline probability for eligibles. At the same time, the probability of a visit to a hospital rises by one percentage point, which is about a third of the baseline probability, but this estimate is not statistically significant. There is also a fall at other sites, but it is also not statistically significant.

It is somewhat difficult to draw strong conclusions about efficiency from these estimates because we do not know whether a change in the number of visits to a specific site reflects new visits, or a shifting of patients who would have obtained visits in any case to a new site of care.[28] Nevertheless, the fact that visits to physicians' offices increase more than visits to other sites suggests that the increases in visits are being provided relatively efficiently.

tion is that 839 children had more than one visit in the past two weeks, hence the sum of the visits to different sites may slightly exceed the total number of visits. We use linear probability models instead of multinominal logits here because the latter strategy was infeasible given the sample size, the number of covariates, and the endogeneity of eligibility.

28. For example, it is possible that all new visits are to doctor's offices, while some patients shift from other sites to hospitals. Alternatively, some new visits may be to hospitals rather than to doctors' offices, while other patients shift from other sites to the doctor's office.

IV. Effects on Child Mortality and Equalization

IV.A. Child Mortality

As noted above, even if Medicaid increases the utilization of medical care, it may not improve health status. The NHIS does not contain any objective information on child health, and the subjective indicators that are presented are difficult to interpret.[29] One important objective indicator of child health that is available, however, is child mortality, which is reported for each state and year for children aged one to four and five to fourteen in the publication *Vital Statistics*.[30] In this section we ask whether the fraction eligible in a state, year, and age group, calculated using the CPS, is related to child mortality. We estimate TSLS models of deaths per 10,000 children as a function of the fraction eligible for Medicaid, instrumenting the fraction eligible using the simulated fraction eligible in each state, year, and age group. These models include a full set of age group, state, and year dummies.[31]

These TSLS estimates are presented in the first column of Table VI. Increases in the fraction eligible for Medicaid have a significant negative effect on the child mortality rate. The coefficient implies that for every ten percentage point increase in the fraction of children eligible for Medicaid, mortality drops by 0.128 percentage points, which is 3.4 percent of the baseline mortality rate in the sample. Thus, the 15.1 percentage point rise in eligibility between 1984 and 1992 is estimated to have decreased child mortality by 5.1 percent.

The *Vital Statistics* data also provide information about deaths by cause. If Medicaid eligibility reduces deaths by improving the utilization of care, then we would expect deaths due to

29. In particular, subjective health measures (such as parental reports about the child's health and activity limitations) may capture two effects of increased contacts with the medical system: a "true" health effect; and a "reported" health effect, that arises because contacts with medical practitioners may affect perceptions about the health of the child. The net effect of Medicaid eligibility on subjective health is difficult to evaluate. In Currie and Gruber [1995] we find that Medicaid has insignificant effects on the subjective health measures in the NHIS.

30. These data are available only through 1991.

31. Due to the more limited variation in these grouped data, we do not include age group $\times$ state or age group $\times$ time interactions. Their results are similar if these controls are included, although our standard errors rise by about one-half. All regressions are estimated using White [1984] standard errors, to account for variation in the cell size by state, year, and age group. In our discussion above, we also noted that a problem with using actual eligibility in the CPS is small cell sizes by age/state/year. This is less of a problem in the current context, since we are using two broad age groups. Also, our instrumental variables strategy will remove any measurement error bias arising from the fact that some of our average eligibility measures are computed using small cell sizes.

TABLE VI
EFFECTS OF MEDICAID ELIGIBILITY ON CHILD MORTALITY
DEPENDENT VARIABLE IS DEATHS PER 10,000 CHILDREN

	(1) All causes	(2) Internal causes	(3) External causes
Percent eligible	−1.277	−1.016	−0.261
	(0.482)	(0.359)	(0.363)
Mean of dep var	3.807	1.926	1.881
Number of obs	816	816	816

Standard errors are in parenthese. Dependent variable is death rate per 10,000 children in state/year/race/age group, where age groups are 1–4 years old and 5–14 years old. Regressions are run as instrumental variables, where percent eligible in state/year/age group cell is instrumented using simulated eligibility in that cell. Regressions include state, year, and age group dummies. Standard errors are corrected for heteroskedasticity.

"internal causes" (such as disease) to fall more than deaths due to "external causes" (such as accidents, homicides, suicides, and other external causes).[32] Columns (2) and (3) of Table VI show that this is indeed the case: increases in eligibility are correlated with a significant reduction in deaths due to internal causes, but have no significant effect on deaths due to external causes. In percentage terms, the reduction in internal deaths from the 15.1 percentage point rise in eligibility is fairly large (8 percent). Thus, these data suggest that Medicaid is having a causal effect on child mortality.

One means of interpreting our finding is to calculate the cost to Medicaid per child life saved. The average cost of Medicaid per year per low-income child on the program is $902 [Congressional Research Service 1993]. This estimate excludes the disabled, who cost substantially more. According to the instrumental variables take-up results in Part I, making 10 percent more children eligible for Medicaid raises the number of children covered by the program by 2.27 percent. In 1992 this would have been 1,293,194 children. The cost of covering these children would be $1.17 billion. The regression results imply that 727 lives would be saved (using the coefficient of −1.277 in the first column). This is a cost per life saved of roughly $1.61 million.

This estimate is lower than recent compensating differentials estimates of the value of a life. Viscusi [1992] summarizes these studies and concludes that the most reliable range of esti-

32. Medicaid may still affect mortality rates from external causes, as a number of recent studies suggest that insured patients receive more intensive treatment in hospital settings than their uninsured counterparts [Hadley et al. 1991; Wenneker et al. 1990].

mates is $4 to $7 million per life saved. It is also lower than the cost per life saved from a number of current government regulatory policies, according to Breyer [1993].[33] It is, however, somewhat higher than the cost of saving a life through Medicaid expansions to pregnant women estimated in Currie and Gruber [1994].

Moreover, this estimate is likely to be an upper bound, for several reasons. First, the cost for the marginal child brought into the Medicaid program through the expansions of the late 1980s is likely to be lower than the average cost of a low-income child on Medicaid, since the very lowest income children (who are likely to be in the poorest health) were already eligible under the AFDC, Medically Needy, or Ribicoff child programs.[34] Second, this calculation ignores incremental benefits to health that arise from Medicaid coverage but fall short of saving a child's life. Third, we do not consider the savings to both the public and private sectors from reductions in the uncompensated care that is currently being delivered to uninsured children. Finally, we ignore any other savings to the private sector that accrue from individuals dropping private insurance to take up Medicaid.[35]

IV.B. Equalization Effects of the Expansions

One goal of many advocates of health care reform is the equalization of access to medical care between groups. In the United States racial differences in the utilization of health care are particularly glaring—for example, black children are 27 percent more likely than other children to have gone without a visit in the previous year. All else equal, children with less educated parents may also be less able to take advantage of care that is available, to the extent that these parents cannot travel the circuitous route to Medicaid eligibility.

It is not obvious, however, that increasing insurance eligibility will equalize utilization. A number of the problems noted

33. On the other hand, it is difficult to know how to value a child's life relative to the adult lives valued in Viscusi and Breyer's work.

34. In fact, in the NHIS, the parentally reported health of those children made eligible under the expansions is better than that of children who were eligible under the old program rules. For those eligible under AFDC or the Medically Needy or Ribicoff child programs, 5.5 percent report fair/poor health and 6.7 percent report on activity limit. For those made eligible under the 1987 to 1992 income expansions, 3.2 percent report fair/poor health and 3.6 percent report an activity limit.

35. At the same time, however, we do not consider the increased deadweight loss from government revenue raised to finance these Medicaid expenditures.

in the introduction, such as the alleged segregation of Medicaid providers from the population in greatest need, suggest that expanding eligibility could actually make utilization even more unequal by increasing the demands on a limited number of suppliers of care. Furthermore, important disparities in utilization and health remain even in countries with universal insurance coverage (see the *The Black Report* in Townsend et al. [1988]).

We investigated potential differences in the effects of Medicaid eligibility, by estimating separate models for blacks and nonblacks, and for children whose mothers were high school dropouts and other children. Although we were unable to find statistically significant differences between these groups, the point estimates (reported in Currie and Gruber [1995]) were suggestive. It appeared that eligibility decreased the probability of going without a visit more for disadvantaged groups (both blacks and high school dropouts) than for others. On the other hand, disadvantaged groups were more likely to experience increases in the use of hospital clinics and emergency rooms rather than doctor's offices. Hence, the results suggested that Medicaid eligibility could reduce disparities in utilization of care while increasing disparities in the site of care.

We also found that increases in the fraction eligible at the state level had a larger impact on mortality among black children than among white children. This difference, although once again not significant, implies that the equalization in utilization arising from expanding Medicaid was also translated into an equalization of health outcomes. Overall, these findings suggest that further research into the equalizing effects of the Medicaid expansions, perhaps using more precise measures of utilization, could prove useful.[36]

V. Conclusions

Expansion of health insurance for children remains a popular public policy, despite the controversy about health care reform. The strong support for insuring children reflects the assumption that lack of insurance is responsible for the poor health of American children. Furthermore, there is some public

36. Currie and Thomas [1995] examine Medicaid coverage and find that relative to either no insurance or private insurance, Medicaid coverage increases the probability of a routine checkup among both black and white children. However, Medicaid increased the number of visits for illness only among white children.

support for the concept that health care is a basic human right, and that the sizable disparities in the utilization of care among children violate this right [Castelle 1990; U. S. National Committee on the International Year of the Child 1980]. Unfortunately, little is known about the utilization and health effects of extending eligibility for public health insurance to previously ineligible groups of children.

The goal of this paper has been to address this question using recent expansions of Medicaid eligibility as a source of identifying variation in models of the utilization of medical care and health. The resulting instrumental variables models yield a number of interesting findings. In particular, we find that making a child eligible for Medicaid lowers the probability that he or she goes without a doctor's visit during the year by one-half. Becoming Medicaid-eligible also raises the probability that a child went to a physician's office in the past two weeks by approximately two-thirds. Finally, increases in eligibility at the state level are associated with significant reductions in child mortality.

We focus on Medicaid eligibility, rather than coverage, since this is the margin that is directly affected by Medicaid eligibility policy.[37] However, our estimates are quite consistent with previous work on the effects of Medicaid coverage on utilization. For example, Currie and Thomas [1995] find that relative to no insurance, Medicaid coverage increase the probability that a child had a routine checkup in the past year by 15 percent. Our estimates imply that the probability that a child went without a visit in the past year is reduced by 8 percent, which is consistent given that between 50 and 70 percent of the eligible uninsured take up coverage.

This research raises a number of interesting questions that are crucial to a better understanding of the role of public insurance. First, how might the impact of public insurance differ if take-up among eligibles was increased? Are eligibles who do not take up coverage less needy, or do they simply face larger informational and other barriers? Second, how can public policies deliver care to the publicly insured more efficiently? Our results suggest that the Medicaid expansions increased the utilization of visits in a relatively efficient way (since most of these visits seem

37. That is, the government has control over who is made eligible for the Medicaid program, but not over who chooses to take it up (although there are other policy margins, such as outreach programs, which can affect the take-up margin).

to have taken place at the doctor's office rather than in emergency rooms). However, there were also significant increases in the incidence of hospitalization that may reflect inefficiencies in the way that care is delivered to Medicaid patients.

Currie, Gruber, and Fischer [1995] find that increases in the fees paid to obstetrician/gynecologists by the Medicaid program were associated with declines in infant mortality, which suggests that supply side policies can have important effects on access to physicians and on outcomes. How might other supply side policies impact the efficiency of care and outcomes? Finally, what are the health benefits of public insurance beyond the effect on child mortality? In future work we hope to move beyond this relatively crude measure of health to investigate the effects of Medicaid on other, more detailed, indicators of objective health status. The availability of micro-data on objective health outcomes would be an invaluable tool for such future assessments of the effects of public insurance on health.

Appendix 1: Measuring Medicaid Eligibility

In this appendix we describe our procedure for imputing the Medicaid eligibility of individuals in the CPS and NHIS. Our source for information on state Medicaid options is National Governors' Association [various years] and Congressional Research Service [1988, 1993].

A. Eligibility for AFDC

In order to qualify for AFDC, the child's family must satisfy three tests: (1) gross income must not exceed a given multiple of the state needs standard,[38] (2) the gross income less certain "disregards" must be below the state needs standard, and (3) the gross income less the disregards, less a portion of their earnings, must be below the state's payment standard.

The disregards can be computed as follows. Beginning in October 1981 the allowance for work and child care expenses was $75 per month for work expense and a maximum of $160 per child for child care costs. These allowances were not changed until the Family Support Act of 1988, which raised the allowances to $90 for work expenses and $175 per child for child care expenses, effective October 1, 1989. In addition, a portion of earned income

38. In 1984 this multiple was 1.5. From 1985 onward the multiple was 1.85.

was disregarded. In 1984 women were allowed to keep $30 plus one-third of earned income for four months. Hence, we assumed that women could keep $120 + 1/9 of their earnings for the year. From 1985 onward individuals who would have become ineligible for AFDC (and hence for Medicaid) after the four months were allowed to remain eligible for Medicaid for an additional nine to fifteen months depending on the state. We modeled this by assuming that for Medicaid eligibility purposes, women were allowed to keep the full $30 and one-third. Our aim was to consistently model the maximum amount that a person could have received while remaining eligible for Medicaid coverage under AFDC. Finally, AFDC rules until October 1984 mandated that women be assumed to be using the Earned Income Tax Credit, regardless of whether they actually were doing so. We follow this assumption in our eligibility calculation for 1984 only.

One difficulty in implementing these rules in the NHIS is that the disregards apply only to earned income, and we cannot distinguish between earned income and other income. We therefore assumed that all household income was earned. This assumption yielded AFDC eligibility findings in the NHIS that were similar to those from the CPS, where we have data on individual earnings by source.

The second set of rules that must be evaluated to see whether a child is eligible for AFDC are rules relating to family structure. Eligibility under the traditional program requires that the child reside in a female-headed household. However, children in two-parent households may still have been eligible under the AFDC-UP program. Eligibility for AFDC-UP conditions on both current employment status and work history. Data on AFDC-UP regulations are from Hoynes [1993]. In addition, some states covered families with Medicaid if they had an unemployed head, even if there was no AFDC coverage. These states are identified in the National Governors' Association [various years].

Lacking longitudinal data on work histories, we assume in the CPS that families are eligible if the state has a program, and the spouse had worked less than 40 weeks in the previous year. In the NHIS we are only able to determine whether or not the spouse is currently unemployed. Hence, our estimates of the AFDC-UP caseload is biased upward because we cannot determine that those who are unemployed have been attached to the labor force long enough to qualify for AFDC-UP. Still, our estimates of the size of the AFDC-UP caseload appear to be reason-able as we find that about one in twenty AFDC eligibles qualify through that program, matching the ratio reported in administrative data.

B. Eligibility under State Medically Needy Programs

In some states, children in families with incomes too high for AFDC could qualify for Medicaid under state Medically Needy programs. Income thresholds for these programs could be set no higher than 133 percent of the state's needs standard for AFDC. Families could "spend down" to these thresholds by subtracting their medical expenditures from their gross incomes (less disregards). If they did so, then Medicaid would pay the remainder of their medical expenses. In order to qualify, however, families must have high medical expenditures for several consecutive months (the "spend down period"). We have no way of determining which families have had such high medical spending in the CPS, and we are reluctant to do so in the NHIS since eligibility would then be a direct function of utilization and health. As an approximation, we set the eligibility thresholds to the Medically Needy levels in states with this program. Data on Medically Needy coverage and thresholds is from National Governors' Association [various years].

C. Eligibility for Ribicoff Children

Ribicoff children are those who would qualify for AFDC given income criteria alone, but who do not qualify for reasons of family structure. States may or may not choose to cover children under this optional program. In states that do cover them, we ignore the family structure requirements and screen only on income. Some states cover selected groups of children (such as only those in two-parent families, or only those in institutions). However, we were unable to obtain precise information on the groups of children covered. Hence, we count the state as a "Ribicoff state" only if it covers all categories of children, as reported by the National Governors' Association. We have also tried calling all of the states to obtain information about their Ribicoff children program. The resulting information appeared unreliable, since almost every state said that they had a program, whereas secondary sources report that coverage is much more selective. Using the state self-reported coverage yielded results similar to those reported in the paper.

D. Eligibility under the Medicaid Expansions

See Appendix 2 for a summary of the relevant legislation. If family income and the child's age were less than the cutoffs, we assumed that the child was eligible. One important question is whether states apply AFDC disregards when computing a family's eligibility for the expansions. Discussions with several state and federal Medicaid administrators suggested that such disregards were generally applied, so we used them in our eligibility calculations. Calculating eligibility without the disregards yielded a significantly smaller effect of the expansions, but the regression results were quite similar.

APPENDIX 2: 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 two, three, four, or five who were born after September 30, 1983. Effective October 1, 1988, states could expand coverage to children under age eight born after September 30, 1983. Allows 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 five in fiscal year 1989, and through age six 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 twelve months among 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 six with family incomes up to 133 percent of the federal poverty line.

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

APPENDIX 3: ELIGIBILITY BY STATE

Actual economic conditions				Fixed economic conditions			
State	1984	1992	Diff	State	1984	1992	Diff
AL	0.111	0.252	0.141	AL	0.111	0.348	0.237
AK	0.179	0.208	0.029	AK	0.179	0.213	0.035
AR	0.163	0.292	0.129	AR	0.163	0.356	0.193
CA	0.294	0.406	0.112	CA	0.294	0.343	0.050
CO	0.073	0.211	0.138	CO	0.073	0.191	0.118
CT	0.122	0.273	0.151	CT	0.122	0.170	0.048
DE	0.133	0.198	0.065	DE	0.133	0.277	0.144
DC	0.427	0.474	0.047	DC	0.427	0.523	0.096
FL	0.116	0.337	0.221	FL	0.116	0.311	0.196
GA	0.120	0.327	0.207	GA	0.120	0.314	0.195
HI	0.126	0.278	0.152	HI	0.126	0.287	0.161
ID	0.065	0.292	0.227	ID	0.065	0.299	0.234
IL	0.204	0.287	0.083	IL	0.204	0.291	0.087
IN	0.089	0.272	0.183	IN	0.089	0.243	0.154
IA	0.206	0.266	0.060	IA	0.206	0.300	0.095
KS	0.065	0.190	0.125	KS	0.065	0.181	0.116
KY	0.117	0.330	0.213	KY	0.117	0.298	0.181
LA	0.106	0.357	0.251	LA	0.106	0.329	0.223
ME	0.165	0.335	0.170	ME	0.165	0.338	0.172
MD	0.114	0.317	0.203	MD	0.114	0.185	0.072
MA	0.152	0.259	0.107	MA	0.152	0.242	0.090
MI	0.261	0.283	0.022	MI	0.261	0.334	0.072
MN	0.143	0.325	0.182	MN	0.143	0.284	0.141
MS	0.154	0.380	0.226	MS	0.154	0.390	0.237
MO	0.116	0.329	0.213	MO	0.116	0.241	0.125
MT	0.055	0.233	0.178	MT	0.055	0.230	0.176
NE	0.170	0.221	0.051	NE	0.170	0.270	0.100
NV	0.056	0.261	0.205	NV	0.056	0.220	0.164
NH	0.061	0.167	0.106	NH	0.061	0.205	0.143
NJ	0.164	0.223	0.059	NJ	0.164	0.219	0.056
NM	0.088	0.318	0.230	NM	0.088	0.362	0.275

APPENDIX 3: (CONTINUED)

Actual economic conditions				Fixed economic conditions			
NY	0.286	0.491	0.205	NY	0.286	0.481	0.194
NC	0.068	0.299	0.231	NC	0.068	0.257	0.189
ND	0.168	0.223	0.055	ND	0.168	0.293	0.125
OH	0.173	0.267	0.094	OH	0.173	0.260	0.087
OK	0.153	0.334	0.181	OK	0.153	0.245	0.092
OR	0.060	0.243	0.183	OR	0.060	0.246	0.185
PA	0.223	0.218	−0.005	PA	0.223	0.291	0.068
RI	0.194	0.292	0.098	RI	0.194	0.281	0.088
SC	0.079	0.335	0.256	SC	0.079	0.324	0.245
SD	0.050	0.227	0.177	SD	0.050	0.260	0.210
TN	0.095	0.321	0.226	TN	0.095	0.300	0.206
TX	0.069	0.348	0.279	TX	0.069	0.322	0.253
UT	0.146	0.236	0.090	UT	0.146	0.252	0.106
VT	0.192	0.356	0.164	VT	0.192	0.437	0.244
VA	0.090	0.218	0.128	VA	0.090	0.220	0.130
WA	0.126	0.246	0.120	WA	0.126	0.230	0.104
WV	0.104	0.353	0.249	WV	0.104	0.331	0.228
WI	0.266	0.218	−0.048	WI	0.266	0.320	0.054
WY	0.055	0.224	0.169	WY	0.055	0.224	0.169

Figures are fraction of children eligible for Medicaid in each state year, based on the authors' calculations as described in the text and in Appendix 1.

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