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# RE-EXAMINING THE EFFECTS OF MEDICAID EXPANSIONS FOR PREGNANT WOMEN

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**ABSTRACT**

This paper analyzes the effect of Medicaid eligibility expansions on the health insurance coverage of women giving birth and on the use of prenatal care and infant health, controlling for year and state effects and state-specific trends that may be correlated with expansions in Medicaid eligibility. We combine estimates from the two sets of analyses to construct estimates of the effect of health insurance on use of prenatal care and infant health. We find that the eligibility expansions reduced the proportion of pregnant women who were uninsured by approximately 10 percent, although this decrease in uninsured came with the expense of a substantial reduction in private insurance coverage. Changes in Medicaid eligibility were associated with very small and statistically insignificant changes in prenatal care use, birth weight, and incidence of low-birth weight.

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**Introduction**:

The Medicaid income eligibility expansions that began in the latter half of the 1980s were initially focused on poor, pregnant women and were intended to improve and narrow socioeconomic disparities in infant health (The National Commission to Prevent Infant Mortality, 1988). From this somewhat modest beginning, Medicaid has expanded dramatically to the point where today Medicaid pays for the cost of approximately 40 percent of all births and provides health insurance for nearly 30 percent of all children in the United States.[1] Further expansion of Medicaid and its look-alike programs (e.g., the State Children's Health Insurance Program) form the basis of several state and federal proposals to expand health insurance coverage.

Surprisingly, there is little evidence that the expansion of Medicaid has improved infant health, or narrowed socioeconomic disparities in infant health. For example, simple descriptive statistics on infant mortality and low-birth weight provide little evidence that Medicaid expansions have coincided with an improvement in infant health or decreased socioeconomic disparities in infant health. Figure 1 shows a steady decline in infant mortality from 1980 to 2000 for both white and black infants and there is no evidence of a narrowing of the racial gap or of changes in the rate of improvement that coincide with expansions of Medicaid.[2] Rates of low birth weight have increased between 1980 and 2000 and there is no break in trend here either that coincides with the Medicaid expansions. Likewise, it is more the exception than the rule that more sophisticated research has uncovered a beneficial effect of Medicaid expansions on infant or child health (Howell 2001; Hadley 2003; Levy and Meltzer 2008).

The absence of more substantial and consistent evidence demonstrating the health benefits of Medicaid is notable because an important part of the justification for expanding Medicaid is to improve population health by preventing illness. Specifically, the initial expansions of Medicaid that were focused on pregnant women were partly motivated by the 1985 Institute of Medicine Report (IOM 1985) entitled *Preventing Low Birthweight*, which concluded that more, and more timely, prenatal care could reduce

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[1] These figures are derived from statehealthfacts.org of the Henry J. Kaiser Family Foundation (http://www.statehealthfacts.org). Website accessed March 14, 2008.
[2] These figures are from Centers for Disease Control and Prevention (2002).

infant mortality by preventing low birth weight. The Report of the National Commission to Prevent Infant Mortality (1988) also linked universal access to prenatal care to the prevention of low birth weight and infant mortality, and noted the cost-effectiveness of such spending relative to post-birth intervention to prevent infant mortality.

The jury is still out on the efficacy of Medicaid in improving infant health because even after more than 20 years of experience there is simply too little evidence on which to base a conclusion. Part of the explanation for the absence of more definitive evidence is the absence of suitable data and the difficulty of conducting non-experimental research. There is relatively little data available that has information on both insurance coverage and infant health. For example, vital statistics has information on birth weight and other measures of infant health, but no information on health insurance status. Similarly, while the National Health Interview Survey has information on health insurance status, it has limited information on infant health. The lack of information about infant health and insurance coverage in the same survey makes it difficult to link Medicaid coverage directly to infant health. Because of this, all previous studies that exploit the natural experiment afforded by the Medicaid expansions have examined the relationship between Medicaid eligibility and infant health (i.e., reduced form model) and not the relationship between Medicaid (insurance) participation and infant health (i.e., structural model). While this approach is reasonable, and will be one of the multiple approaches we take, there may be a weak link between Medicaid eligibility and the probability that a pregnant woman has health insurance. This is because most women, even poor women, have private insurance and because of the potential for families to substitute public for private insurance (i.e., crowd out). Indeed, several studies of children and families have reported large crowd out effects (Cutler and Gruber 1996; Dubay and Kenney 1997; Lo Sasso and Buchmueller 2004; Gruber and Simon 2008). However, insurance coverage during pregnancy is the most relevant for infant health, and there is only one study that we are aware of that has examined the effect of the Medicaid expansions on health insurance coverage of women who are pregnant (Dubay and Kenney 1997). Thus, it is difficult to assess the credibility of previous studies of the relationship between Medicaid eligibility and infant health because little is known about the relationship between Medicaid eligibility and health insurance coverage of pregnant women, which is a critical link in the causal chain of reasoning that connects Medicaid eligibility to health.

The second impediment to establishing a credible base of evidence regarding the effect of the Medicaid expansions on infant health is the difficulty of conducting non-experimental research even in the context of the natural experiment provided by the Medicaid expansions. Previous studies and the ongoing debate related to insurance crowd out have demonstrated that there are state-specific trends correlated with the timing and magnitude of the Medicaid expansions that can significantly affect estimates (Card and Shore-Sheppard 2004; Gruber and Simon 2008; Shore-Sheppard 2008). In short, Medicaid expansions are not randomly occurring natural experiments.

In this paper, we begin to address some of these issues. First, we provide an extensive analysis of the effect of Medicaid expansions on the health insurance coverage of pregnant women (infants). As noted, there is only one previous study on this topic even though knowledge of the effects of the Medicaid expansions on health insurance coverage is essential to its evaluation. Second, we assess the effect of the Medicaid expansions on infant health and use of medical services associated with deliveries (e.g., prenatal care). While this is an area with a substantial amount of prior research, we use a new data source, the National Hospital Discharge Survey, and examine some previously unstudied outcomes. We also pay particular attention to research design issues and assessing the threats to validity such as the existence of state-specific trends that may be correlated with expansions in Medicaid eligibility that may explain some of the disparate findings of previous research. Third, we combine the two reduced form analyses to construct implied estimates of the effect of health insurance on infant health and use of medical services associated with deliveries. Here too, we will make a novel contribution as there is no reliable evidence of the effect of insurance status itself on infant health.

**The Effect of Medicaid Expansions on Health Insurance Coverage of Pregnant Women:**

In the sole previous study of the effect of Medicaid expansions on the health insurance coverage of pregnant women, Dubay and Kenney (1997) provided a simple descriptive analysis of changes in health insurance coverage of pregnant women using data from the Current Population Study (CPS) from 1989 to 1993 to study health insurance coverage between 1988 and 1992. The analysis consisted of examining changes in health insurance coverage for pregnant women with incomes less than 185 percent of the federal poverty level, which is approximately the maximum level of income to retain eligibility for Medicaid in 1992. The results indicated that for women in this income group, there was an increase in Medicaid coverage, a decrease in private health insurance coverage and a decrease in the proportion of pregnant women who are uninsured. Results differed markedly by income. For poor women, defined as those with incomes under 100 percent of federal poverty, there was a modest increase in both Medicaid coverage and private insurance coverage and a seven percentage point decrease in the proportion of women uninsured between 1988 and 1992. For near-poor women with incomes between 100 and 185 percent of federal poverty, there was virtually no change in the proportion uninsured during the same period; the proportion with Medicaid increased by approximately 15 percentage points and the proportion with private health insurance decreased by approximately 15 percentage points. Given that private health insurance coverage was falling generally during this period (for example private insurance coverage for men declined by seven percentage points) Dubay and Kenney (1997) concluded that the Medicaid expansions resulted in a net increase in insurance coverage of approximately six percentage points among near-poor women.

There are several limitations of this study. Pregnant women in the CPS are identified by women who have children under the age of one at the time of the survey, and questions about insurance status refer to the year prior to the survey year. Therefore, responses to health insurance coverage questions may reflect coverage at the time of the survey, or coverage before or after the birth of the child during the previous year. In addition, not all children and mothers can be linked in the CPS. Also, Dubay and Kenney (1997) did not produce an estimate of the relationship between Medicaid eligibility and insurance coverage. Finally, the use of men as a comparison group was not justified and there are good reasons to believe that trends in their health insurance coverage would differ from trends in the health insurance coverage of pregnant women given differences in employment, occupation and family structure.

Despite the problems with this study, a potentially important finding is that Medicaid expansions during this period may have affected a relatively small part of the target population. Relatively few poor women were uninsured prior to the Medicaid expansions; even though the expansions may have had relatively large effects on the proportion uninsured, the share of the population that changed from uninsured to insured represents only a small fraction of the total population. For example, if we assume that the Dubay and Kenney (1997) estimates are accurate, only about six percent of the population of pregnant women with incomes between 100 and 185 percent of poverty changed insurance status (from not covered to covered) between 1988 and 1992 even though a substantial majority was made eligible for Medicaid during this period. In addition, Ellwood and Kenney (1995) and Simon and Handler (forthcoming) show that many pregnant women who enroll in Medicaid do so after the first trimester. The small fraction of the population that changed insurance status and the timing of obtaining insurance suggest that the effect of eligibility on health is likely to be small, and the statistical power to detect reliably small effects may be limited.

To summarize, there is only one prior study of the effect of Medicaid expansions on pregnant women's health insurance coverage and it is limited in a number of ways. The likely explanation for this paucity of research is the absence of data on insurance coverage of pregnant women. To address this shortfall in research, we will use data from the National Hospital Discharge Survey (NHDS) from 1985 to 1996. This data source provides information on the health insurance coverage of a large, nationally representative sample of pregnant women who deliver in US hospitals. Thus, we will produce the first comprehensive analysis of the effects of Medicaid expansions on health insurance coverage of pregnant women.

**The Effect of Medicaid Expansions on Infant Health**

Howell (2001) provided an analysis of the effects of the Medicaid expansions on infant health and concluded that there was little evidence that the Medicaid expansions had an effect on low birth weight or infant mortality. She summarized the evidence as follows:

- "Limited evidence that the rate of low birth weight or prematurity declined for some women affected by the expansions. However, a majority of studies that addressed this issue found no effect of the expansions." (Howell 2001, p.20)

- "A single study finds an impact of the expansions on infant mortality. Because this study did not include controls for socioeconomic status and has findings that are inconsistent with state-level studies, it is difficult to draw the conclusion that the Medicaid expansions affected infant mortality rates from this single study." (Howell 2001, pp.20-21)

Hadley (2003) and Levy and Meltzer (2004) provided similar reviews of the literature and concluded that there is little evidence that Medicaid expansions improved infant health.[3] The absence of an effect of Medicaid on infant health is consistent with clinical evidence that there is a limited scope for medical care to improve birth outcomes, such as premature births and low birth weight (Strong 2000; Lu et al. 2003; Behrman and Butler 2006).

Despite limited evidence, many advocacy groups continue to argue that Medicaid has had a positive effect on infant health. For example, a report by the Center on Budget and Policy Priorities states the following:

> "Expansions of Medicaid coverage for low-income pregnant women led to an 8.5 percent reduction in infant mortality and a 7.8 percent reduction in the incidence of low birth weight." (Ku 2005, p. 4)

The findings repeated in this passage are from Currie and Gruber (1996), which is arguably the most influential study in this area. It was the first national study, it was comprehensive in terms of the outcomes it studied, and it used an innovative research design that exploited the legislated changes in Medicaid. Unfortunately, effect sizes obtained in this study seem implausible if considered in light of estimates of the effect of Medicaid expansions on insurance status reported in Dubay and Kenney (1997) and elsewhere (e.g., Cutler and Gruber 1996). According to Dubay and Kenney (1997), it is likely that approximately 50% of the pregnant women who enrolled in Medicaid during this period were previously insured. If we use the 34% rate of take-up for Medicaid reported in Currie and Gruber (1996) and assume that 50% of those were previously uninsured, then a 100% increase in Medicaid eligibility (i.e. making every pregnant woman eligible for Medicaid) would be associated with a decrease in uninsured pregnant

[3] However, Hadley (2003) puts forth several potential explanations for this general finding that are consistent with a true benefit of Medicaid on infant health.

women of 17% (i.e. for every 100 pregnant women, 17 would switch status from uninsured to insured). Therefore, estimates of the effect of Medicaid eligibility on outcomes reported in Currie and Gruber (1996) should be multiplied by approximately five to obtain estimates of the effects on those becoming newly insured, as outcomes are averaged over the entire population while less than 20 percent experienced a change in insurance status. Making this calculation yields estimates that imply that switching a woman's insurance status from uninsured to Medicaid will decrease the probability of infant mortality by between 150% and 190%, which seem implausible.[4] Even if we double the rate of take-up to 68% and reduce the rate of crowd out by half to 25%, estimates of the effects of becoming newly insured on infant mortality would be between 60% and 75%, which still seem too large to be plausible.

The implausible magnitude of the estimates in Currie and Gruber (1996) is likely due to the inadequacy of the underlying natural experiment. The fact that the Medicaid expansions were not a particularly good "natural" experiment is revealed in the sensitivity analyses conducted by Currie and Gruber (1996). When state maternal and child health spending are included in the model, the estimate of the effect of Medicaid eligibility on low birth weight increases by 532%. Similarly, when the number of neonatal intensive care beds are added to the regression the estimate of the effect of Medicaid eligibility on low birth weight increases by 160%. Estimates of the effect of Medicaid eligibility on infant mortality are somewhat less affected, but the addition of the abortion rate to this regression model increases the estimate of the effect of Medicaid eligibility on infant mortality by 25%. The sensitivity of estimates to inclusion of additional variables indicates that the timing and magnitude of the Medicaid expansions were correlated with other changes in states that significantly affected birth weight and infant mortality. It is likely that some of these factors are confounding estimates of the effect of Medicaid eligibility on infant health outcomes.

[4] These figures are calculated using estimates of the effects of targeted Medicaid expansions reported in Table 3 of Currie and Gruber (1996). For example, Currie and Gruber (1996) reported that a 100% increase in Medicaid eligibility was associated with a decrease of 4.088 in infant deaths (per thousand births). To obtain the effect of becoming newly insured, we multiply this estimate by five and divide by the mean of 10.66, which yields 1.9174 (191.74%).

**Summary and Contributions**

In sum, there is little evidence to suggest that the Medicaid expansions had a positive effect on infant health, as measured by birth weight or low birth weight, prematurity, or infant mortality. There are several possible explanations for this fact: low rates of Medicaid take-up, crowd out of private insurance, and limited ability of medical care to improve birth outcomes. The apparent lack of success of the original Medicaid expansions targeted at pregnant women and infants seems inconsistent with the political support for later expansions such as SCHIP and expansion of Medicaid along the lines recently enacted by Massachusetts. These later expansions of Medicaid may have little impact on health because they suffer from the same problems that potentially explain why the earlier expansions targeted at pregnant women and infants had few health benefits. Take-up rates remain low, crowd out of private insurance is a continuing issue, and the link between primary care and improved health may be limited to particular diseases or specific to the health indictors chosen. Therefore, whether or not the initial expansions were a success in terms of improving infant health is an important research question because of the policies the initial expansions have spawned. In addition, the absence of sufficient evidence about the effects of the initial expansions, for example, little study of the extent of take-up and crowd out of pregnant women, suggests the need for additional study.

Our work in this paper begins to fill in the gaps. Specifically we provide virtually the first study of the effects of the initial Medicaid expansions on insurance coverage of pregnant women and infants, and we provide an additional analysis of the effects of the expansions on prenatal care and infant health, as measured by birth weight. We also study additional measures of the use of health services (e.g., cesarean section, length of stay) not previously examined using the National Hospital Discharge Survey. In all analyses, we pay special attention to the empirical problems caused by the absence of a true natural experiment. Finally, we provide estimates of the implied effect of becoming newly insured on infant health (i.e., instrumental variables estimates).

**Research Design**

Our research design initially proceeds as if the Medicaid expansions represented a true natural experiment. We are interested in the effect of the expansions on insurance coverage, prenatal care, and infant health as measured by birth weight. We begin the description of the research design with the effect on insurance coverage. For each type of health insurance category (HI), we estimate the following regression model:

$$
\begin{aligned}
& HI_{ijt} = \alpha_j + \beta_t + \delta\, ELIG_{kjt} + X_i\Gamma + Z_{jt}\lambda + u_{ijt} \\
& \mathrm{i} = 1,\ldots,\mathrm{N} \qquad \text{(persons)} \\
(1)\quad & \mathrm{k} = 1,2,3 \qquad \text{(race)} \\
& \mathrm{j} = 1,\ldots,50 \qquad \text{(states)} \\
& \mathrm{t} = 1985,\ldots,1996 \qquad \text{(years)}
\end{aligned}
$$

In equation (1), HI is a specific health insurance category (private, Medicaid, uninsured) for woman i in state j and year t; ELIG is the fraction of women in group k, which is defined by race, in state j and year t that are eligible for Medicaid; X are individual characteristics of pregnant women such as age and race; and Z are the unemployment rate (lag, contemporaneous, and lead) and fraction of single males with private insurance in state j and year t. Equation (1) also includes state ( $\alpha_j$ ) and year ( $\beta_t$ ) effects. In models in which the sample includes persons of different races, we allow state and year effects to differ by race. We chose to define our group cells by race, in addition to state and year, because of the large racial differences that exist in infant outcomes.

The use of the fraction of women eligible for Medicaid in each state and year to measure Medicaid eligibility follows the innovation of Currie and Gruber (1996) and Cutler and Gruber (1996). The key to this measure is that it consists of the fraction of a fixed sample of women in each state who are Medicaid eligible, avoiding the endogenous relationship between Medicaid eligibility of an individual mother and the health of that mother or her infant.[5] To form the measure, we select a sample of women aged 18 to 39 from the 1986 to 1997 March CPS surveys (covering data from 1985-1996). We adjust all

[5] Currie and Gruber (1996) and Cutler and Gruber (1996) use a national sample instead of state-specific samples. However, given that the measure is based on state variation in timing and magnitude of Medicaid expansions, using state-samples introduces no additional variation that would bias estimates. State samples will produce less measurement error, which is relevant to correlation between eligibility and insurance.

financial variables for price changes so that everyone's incomes are deflated (inflated) from the actual year (e.g., 1985,..,1996) to the policy year (e.g., 1988) for which eligibility will be calculated. Using these inflation-adjusted state-specific samples and the Medicaid eligibility rules in a state in a specific year, we assign eligibility to the (fixed) state sample. So eligibility is calculated using the same sample of women in each year for a given state. We then calculate the proportion of women aged 18 to 39 within a race-state-year cell that would be eligible for Medicaid if they became pregnant.

Data for the analysis, which we describe in more detail below, come from hospital discharge records in the National Hospital Discharge Survey (NHDS). The key aspect of these data is that they provide information on insurance coverage for pregnant women at the time of delivery. We restrict the sample to women aged 18 to 39 years. The instrument—the fraction of women that would be eligible for Medicaid conditional on pregnancy—is merged with the NHDS data by year, state and race.

We modify this basic model in a number of ways to address specific issues. First, we allow the eligibility measure (ELIG) to have non-linear effects, which may be likely given that higher levels of eligibility are affecting persons with increasingly higher rates of private insurance and lower rates of uninsured (Card and Shore-Sheppard 2004). We do so by including eligibility and eligibility squared in the regression. Second, in some models we include state-specific trends to adjust for the less than perfect nature of the natural experiment. State-specific trends adjust for the potential endogeneity of the Medicaid expansions. Here too, we allow for separate race-specific state and year effects in models that are estimated using a sample with different races. Third, we restrict the sample to the period 1985 to 1991, which is the period that witnessed the greatest change in Medicaid eligibility. As we show later, while the fraction of Medicaid-eligible births increased from about seven percent in 1985 to almost 30 percent in 1996, most of the expansions had already taken place by 1991. Finally, we also conduct separate analyses by race.

We estimate a similar model related to infant health outcomes: number of prenatal visits, timing of first prenatal care visit, birth weight for gestational age, low birth weight for gestational age, cesarean section, place of birth (public hospital versus non-profit or for-profit hospital) and length of stay in hospital. For these outcomes (represented by BW), the regression model is:

$$
(2)\quad \begin{aligned}
&BW_{ijt} = \tilde{\alpha}_j + \tilde{\beta}_t + \gamma_i + \tilde{\delta}\, ELIG_{ijt} + Z_{jt}\tilde{\lambda} + \tilde{u}_{ijt} \\
&\mathrm{i} = 1,\ldots,\mathrm{N} \qquad \text{(group)} \\
&\mathrm{j} = 1,\ldots,50 \qquad \text{(states)} \\
&\mathrm{t} = 1985,\ldots,1996 \qquad \text{(years)}
\end{aligned}
$$

In equation (2), the symbol ~ is used to differentiate parameters from equation (1). Data for this analysis come from two sources: the National Hospital Discharge Survey for length of stay in hospital, cesarean section, and place of birth, and vital records for measures of prenatal care and birth weight. Analyses of vital statistics data are conducted using individual data collapsed into state-year-group (age by race) cells. The index i in equation (2) reflects this fact and is defined by age and race. Models estimated for birth weight and low birth weight are adjusted for gestational age, but are otherwise the same as equation (1). The fraction of women eligible for Medicaid is merged to these cells by year, state and race. Here too, we alter the model specification to allow for non-linear effects of eligibility and to control for unmeasured, state-specific time trends. All models are estimated via ordinary least squares, with weights as noted.

An implicit estimate of the effect of insurance coverage on measures of infant health, for example birth weight, is provided by the following:

$$
(3)\quad TOT_{IV} = \frac{\tilde{\delta}}{\delta}
$$

We refer to the effect of insurance on birth weight as the treatment on the treated (TOT), or more appropriately as the local average treatment effect (LATE) because it pertains to women and infants affected by the Medicaid expansions, who may not represent the typical Medicaid participant. As equation (3) shows, this estimate can be derived from the ratio of the effect of Medicaid eligibility on birth weight (BW) to the effect of Medicaid eligibility on insurance (i.e., any insurance) coverage.

## Data

We use three primary data sources for the analysis: the National Hospital Discharge Survey (NHDS), the National Detailed Natality Files, and the CPS. The CPS is used to calculate the fraction of women eligible for Medicaid by race, state and year as described in the previous section of the paper. It is also used to calculate the fraction of never-married men aged 18 to 39 who are privately insured by state and year, one control for trends in health insurance coverage by state.

### *National Hospital Discharge Survey*

Our study of the effect of Medicaid expansions on health insurance coverage of mothers and infants at the time of birth relies on the National Hospital Discharge Survey (NHDS), conducted by the National Center for Health Statistics. The data are collected from a sample of inpatient records obtained from a national probability sample of non-Federal, short-stay hospitals (defined as those with average length-of-stay of less than thirty days) within the 50 U.S. states. The NHDS follows a three-stage sample design using geographically defined sampling units to select a sample of hospitals and a sample of records within hospitals. All analyses use sample weights which inflate estimates by reciprocals of the probabilities of sample selection and adjust for non-response among sampled hospitals.[6] The NHDS contains information on patient age, sex, race, up to three expected sources of payment, and up to seven patient diagnoses coded according to the International Classification of Diseases- 9th Revision- Clinical Modification (ICD-9-CM) (Public Health Service, 1998).

We study the effect of Medicaid eligibility on insurance coverage of mothers at the time of birth using weighted linear probability models, controlling for mother's age category (18-24, 25-29, and 30-34, relative to 35-39), race (black and other race relative to white), state-specific unemployment rate (lag, contemporaneous, and lead), state-specific private insurance rate for males, year and state effects, and race-specific year and state effects (in models which pool all races).

[6] Dennison and Pokras (2000) provide additional information about the design of the NHDS.

We examine the effect of Medicaid expansions on the primary expected source of payment for delivering mothers, with expected source of payment categorized as Medicaid, private insurance (including a small number of cases coded as worker's compensation), and uninsured (cases coded as self pay or charity). Our initial sample for 1985 to 1996 consists of 299,634 women aged 18 to 39 delivering a baby (identified in NHDS by primary ICD-9-CM of V27). To put this sample size in perspective, the CPS has about 5,000 women a year who have had a baby in the last year, and it is not possible to know for sure when the insurance data pertain to relative to the pregnancy, whereas here we know when insurance status is coded (at time of delivery). We use the restricted-access version of the NHDS which contains patient zip code of residence and allows us to infer state of residence. We drop 6,918 (about 2%) of mothers with missing, invalid or foreign zip codes of residence or who live in Puerto Rico, U.S. Virgin Islands, and two U.S. states where the number of records is too small to estimate state fixed effects. We have 284,179 observations after dropping 7,547 records with unknown or other primary expected source of payment and 990 records with primary expected source of payment of Medicare.[7]

In addition to insurance status, we also analyze the effect of Medicaid eligibility on four measures of the use of health services from the NHDS. First, we consider whether or not the birth took place via Cesarean section, a relatively rare occurrence for Medicaid patients (Gregory et al., 1999). Second, we consider whether a mother gives birth in a public hospital (as opposed to a non-profit or for-profit hospital). Poor and uninsured women are more likely than others to give birth in public hospitals, and one way that a change in insurance status may affect infant health is by giving mothers access to hospitals that are more often used by insured mothers and may differ in quality of care (Aizer, Lleras-Muney, and Stabile, 2005). Finally, we consider whether or not the length-of-stay in the hospital of mother or infant

[7] We impute race for the 52,469 (nearly 19%) mothers in the sample who are missing this information using data from non-missing observations. For mothers whose race is reported, we use a multinomial logit model to predict race as a function of the fraction of the population of women aged 18 to 39 who are black or other race in the mother's zip code of residence, using information from the 1990 Census. We also control for age and age squared. We then apply these coefficients to data for those missing race in the NHDS, and assign race based on cut offs that match the observed race distribution among those not missing race. We have verified, however, that our results are not substantively different when we drop those missing race from the analysis.

is more than two days, as this may be an indication of relatively poor health at the time of delivery (for infants, see, for example, Phibbs and Schmitt, 2006).[8]

In addition to age and race, our independent variable of interest – Medicaid eligibility – varies by state, year and race. All analyses thus allow for arbitrary correlation of standard errors within state-year-race cells (Moulton 1990; Donald and Lang 2001).[9]

*Natality Files*

In addition to the delivery-related outcomes from the NHDS, we further assess the impact of Medicaid expansions on infant health and the utilization of medical services by pregnant women using data from individual birth records. Detailed information on all individual births occurring in the 50 states and D.C. are submitted by hospitals to state vital registration offices, which is then reported to the National Center for Health Statistics (NCHS). Information on each birth includes date and place of birth along with the demographic characteristics of the mother such as age, race, education, marital status, and parity. We employ data from the Detailed Natality Files for the years 1985 through 1996, enveloping the major Medicaid income eligibility expansions that took place. The sample is limited to women between the ages of 18 to 39 who had non-plural births. Records from two small states are also excluded in order to facilitate comparison with the results based on the NHDS sample and the construction of the implicit IV estimates. This yields a final analysis sample of 39,238,023 births.

We examine the effects of Medicaid expansions on pregnant women's health care utilization, as measured by the frequency and timing of prenatal care visits. In addition to the total number of prenatal

[8] The sample for our analysis of length-of-stay among infants consisted of 294,721 records after dropping 16,687 records for one of the reasons discussed for the sample of mothers. We imputed race for 52,003 infants based on the race distribution of women aged 18 to 39 in the zip code, though again, our results are not sensitive to dropping infants whose race was imputed.

[9] Errors are not clustered by sampling design unit, since no design information is extant for the NHDS prior to 1988. Although weighted linear probability analysis of the effect of Medicaid eligibility (as the only covariate) on Medicaid insurance status using the 1990 NHDS sample of mothers found that standard errors on the eligibility variable are 2.3 times higher when allowing for clustering of errors among sampling design units compared to no clustering, errors are actually slightly higher when allowing for clustering by state-year-race cell compared to sampling design unit. The approach taken in this paper is therefore expected to result in comparable or conservative standard errors compared to analyses that were able to cluster errors by sampling design unit.

visits over the pregnancy, two dichotomous indicators are constructed for whether prenatal care was initiated on time (during the first trimester) or late (during the third trimester), respectively. An additional indicator for whether prenatal care was adequate is also defined. Adequacy is based on the Kessner/Institute of Medicine criteria, which considers both the timing and frequency of care (Kessner et al., 1973). According to these criteria, prenatal care is considered adequate if it is initiated in the first trimester and is composed of a minimum number of visits depending on gestation. Thus, for a normal 36-weeks gestation period, adequate care constitutes first-trimester initiation and more than eight prenatal care visits over the term. We examine standard measures of infant health: birth weight in grams, and a dichotomous indicator for low birth weight (less than 2500 grams). While there are other, more specific measures of infant health that one may consider, low birth weight remains a strong predictor of perinatal health and may also be associated with long-term developmental outcomes (Hack et al., 1995). Focusing on birth weight further allows a comparison of our results with those in the literature. Models estimated for birth weight and low birth weight are adjusted for age of gestation, since the clinical literature suggests that the ability of medical care to affect fetal growth may be greater than its ability to affect prematurity, since relatively little is known about the causes of preterm delivery (Joyce, 1999; Hack and Merkatz, 1995; Li et al., 1993).[10] Specifically, the models control for the percent of births in each cell which are very premature (gestation less than 32 weeks) and premature (32-36 weeks), in reference to normal gestation (37 weeks or above).

For computing convenience, specifications are estimated on cells collapsed by mothers' age groups (18-24, 25-29, and 30-34 relative to 35-39), race groups (white, black, other), state of maternal residence and year of birth.[11] Although the NCHS natality files are relatively complete, four states (California, New York, Texas and Washington) had incomplete information on educational attainment

[10] Models estimated for gestation did not find any significant or consistent effects of Medicaid eligibility.
[11] Thus, for the full time-period analysis, there are 7,344 cells (51 states times 12 years times 4 age groups times 3 race groups). All models are weighted by cell frequencies. With this weighting scheme, models estimated on collapsed cells lead to identical parameter estimates as those estimated on the individual records. Standard errors are adjusted for arbitrary correlation within state, year, and race cells.

during various years from 1985 through 1991.[12] Since all models are weighted by cell frequencies, this is not a significant cause for concern. As a sensitivity check, we re-estimated all models omitting these four states for the time periods during which the education records were incomplete. The results are quite similar to those reported below.

## Results

*Effects of Medicaid Eligibility on Insurance Coverage of Pregnant Women*

The first question we examined was the effect of expanding Medicaid eligibility on the health insurance status of pregnant women. Table 1 presents the estimates. For each type of insurance coverage including uninsured or any insurance, we estimated several models that differ in the way Medicaid eligibility is entered into the model, and in the control variables that are included. Medicaid eligibility is first entered using a linear term, and next using both a linear and a quadratic term to allow for non-linear effects. We also used two sets of control variables: one that excluded state-specific trends and one with state-specific trends. In sum, we estimated four separate regression models for each type of insurance.

Estimates in Table 1 indicate that Medicaid eligibility is positively associated with Medicaid coverage and having any type of insurance, and negatively associated with private insurance coverage and being uninsured. Estimates are quite sensitive to the inclusion of state-specific trends, which is consistent with what has been noted in the earlier crowd out literature. Specifically, estimates in the top row (linear specification) indicate that a 20 percentage point increase in Medicaid eligibility, which is the approximate change that occurred over the 1985 to 1996 period (see Figure 2), is associated with 3.5 (with state trends) to 5.6 (without state trends) percentage point increase in Medicaid coverage among pregnant women ages 18 to 39. Given that the mean rate of Medicaid coverage was approximately 0.2 in 1985 (see Figure 3), these estimates suggest that the Medicaid expansions during this period increased Medicaid coverage by between 18 and 28 percent. Estimates in second row (quadratic specification)

[12] Educational information is incomplete for the following years: CA -1985-1988; NY - 1988-1990; TX - 1985-1988; WA - 1985-1991.

suggest a similar conclusion, although there do appear to be non-linear effects. Medicaid take-up (i.e., marginal participation) appears to increase with increases in eligibility. Based on estimates obtained from models that include state-specific trends, the marginal Medicaid take-up rate is 0.122 when eligibility is 0.20 (20 percent of population is eligible) and 0.182 when eligibility is 0.30.

In regards to private insurance coverage, estimates in the top row (linear specification) indicate that a 20 percentage point increase in Medicaid eligibility is associated with a 1.9 (with state trends) to 2.9 (without state trends) percentage point decrease in private insurance coverage. Estimates in the second row (quadratic specification) show that marginal declines in private insurance are greater with increases in eligibility. Estimates from models that include state-specific trends indicate that the marginal decline in private insurance is 0.005 when eligibility is 0.20 and 0.112 when eligibility is 0.30. The decreases in private insurance coverage suggested by estimates in the top row represent approximately 50 percent of the increase in Medicaid coverage, or what has commonly been referred to as crowd out. Moreover, these estimates of crowd out do not depend on model specification even though the addition of state-specific trends had a significant mediating effect on the estimates of the effect of eligibility. Allowing for non-linear effects suggest virtually no crowd out at low levels of eligibility but up to 60 percent crowd out when 30 percent of population is eligible.

Consistent with these estimates of crowd out, expansions in Medicaid eligibility are associated with smaller decreases in the proportion uninsured than increases in the proportion covered by Medicaid. Estimates in the top row indicate that a 20 percentage point increase in Medicaid eligibility is associated with a 1.6 to 2.7 percentage point decrease in the proportion uninsured; these percentage point changes represent relative effects of between 9 to 16 percent given a mean proportion uninsured of approximately 17 percent in 1985 (see Figure 3). The effects of eligibility on uninsured are also non-linear; when eligibility is 0.20, the marginal effect of eligibility on uninsured is -0.126, but when eligibility is 0.30, the marginal effect of eligibility on uninsured is -0.07 (based on estimates controlling for state-specific trends.).

To assess whether the results differ by race or time period, we conducted additional analyses. In Table 2 we present estimates of the effect of Medicaid eligibility on health insurance status of white women. Estimates in Table 2 are similar to those in Table 1, but there are some differences. Estimates in Table 2 suggest that white women have relatively higher take-up rates than black women, having rates that range from 0.17 when 20 percent of population is eligible to 0.25 when 30 percent of population is eligible (based on the quadratic specification with state-specific trends). There also appears to be slightly larger and more pronounced non-linear effects of eligibility on private insurance for white women. Estimates obtained from models that include state-specific trends indicate that the marginal effect of eligibility on private insurance is -0.05 when eligibility is 0.20 and it is -0.21 when eligibility is 0.30. These estimates imply rates of crowd out that are significantly higher than those implied by the estimates in Table 1. For white women, estimates from models that include controls for state-specific trends suggest rates of crowd out of between 32 to 84 percent, as eligibility goes from 20 to 30 percent of the population. Estimates relating to the effect of eligibility on uninsured are consistent with these large estimates of crowd out; the marginal effect of eligibility on uninsured is -0.11 when eligibility is 0.20 and -0.05 when eligibility is 0.30 (figures derived from estimates obtained including controls for state-specific trends). In sum, estimates in Table 2 suggest that expansions in Medicaid eligibility initially decreased the proportion of uninsured white women, but as eligibility increased, the expansions resulted in near complete crowd out of private insurance and relatively little change in the proportion uninsured.

In Table 3, we present estimates of the effect of Medicaid eligibility for years 1985 to 1991, the period that witnessed the largest expansions in Medicaid eligibility (see Figure 2). We only discuss estimates obtained from models that include state-specific trends. Overall estimates in Table 2 indicate less crowd out of private insurance and more pronounced non-linear effects. Medicaid (marginal) take-up rates increase with eligibility; when eligibility is 0.15, as in the beginning of period, (marginal) take-up rates are basically zero, but when eligibility is 0.25 (highest during this period), (marginal) take-up rates are 0.13. Similar nonlinearities characterize estimates of the effect of eligibility on private insurance. However, estimates indicate no decline in private insurance and virtually no crowd out from Medicaid throughout most of the range of Medicaid eligibility observed during the period. It is not until eligibility rises to 25 percent that it is negatively associated with private insurance. Given this pattern of results, the marginal effect of Medicaid eligibility on the proportion uninsured is between -0.12 and -0.14.

Our results of the effect of Medicaid eligibility on the health insurance coverage of pregnant women are consistent with those reported by Dubay and Kenney (1997). The Medicaid expansions of the late 1980s and early 1990s were associated with a significant and relatively large increase in Medicaid participation, but substantially smaller decreases in the proportion uninsured because a large fraction of the newly enrolled Medicaid participants previously had private insurance. The less than perfect "experiment" of the Medicaid expansions warrants appropriate caution in reaching conclusions, but assuming that the inclusion of state-specific trends, state unemployment rates and the fraction of single men with private insurance addresses the endogenous nature of the Medicaid expansions, the results of the current analysis suggest that Medicaid expansions for pregnant women induced a significant substitution of private for public coverage and as a result, increased the cost of expanding public insurance coverage. The problem was particularly severe for white women and for later years of the expansions when eligibility reached relatively large portions of the population. In addition, the relatively small changes in the proportion of the population that was uninsured as a result of the expansions, approximately 10 percent or so, suggests that analyses of the effects of the Medicaid expansions on infant health will have small population effects that may be difficult to detect reliably.

*Effects of Medicaid Eligibility on Prenatal Care and Birth Weight—Natality Files*

Estimates of the effects of Medicaid eligibility on prenatal care and infant health derived from the Natality files are presented in Tables 4 through 7. Statistical analyses of these outcomes are similar to those for insurance. We conduct analyses using the entire sample for the 1985 to 1996 period and then analyses by race and time period. However, because we expect Medicaid eligibility to affect low-educated women the most, in some analyses we limit the sample to those with less than a high school degree. Similar analyses could not be conducted for insurance because the NHDS does not provide information on education. For each outcome, we show estimates from two models: one that excludes state-specific trends and one that includes these controls.

Table 4 presents estimates for the entire sample for the period 1985 to 1996. We focus on estimates obtained from models that include state-specific trends, as the inclusion of these controls appears to affect estimates considerably. Focusing on the first row of Table 4, estimates indicate that Medicaid eligibility is not significantly associated with the timing or number of prenatal care visits, birth weight, or the incidence of low-birth weight. Estimates are also extremely small. For example, a 20 percentage point increase in Medicaid eligibility, which is approximately the change in eligibility that occurred over this period, is associated with 0.02 (0.1 percent) fewer prenatal care visits; a 0.005 percentage point (0.7 percent) lower rate of adequate prenatal care; a 0.3 gram decrease in birth weight; and a 0.0003 percentage point (0.6 percent) increase in the prevalence of low-birth weight. Estimates in row 2 suggest equally small effects. While there is some evidence of nonlinear effects for the measures of prenatal care, the range of marginal effects for the observed variation in eligibility is small and the magnitudes of the marginal effects are also small and similar to those for the linear specification. Most importantly, estimates suggest little evidence that Medicaid eligibility increased use of prenatal care or improved birth weight.

Despite the large sample size underlying estimates in Table 4 and the relatively precise nature of the estimates, our analysis may lack statistical power to detect reliably expected effect sizes. The explanation for this is the fact that Medicaid eligibility is associated with relatively small changes in insurance status. Consider results (row 1) from Table 1 that suggest that a 100 percentage point increase in Medicaid eligibility is associated with an 8 to 13 percentage point increase (decrease) in insurance coverage (uninsured). Assuming that the effects of the eligibility expansions on prenatal care and birth weight work only through increasing the proportion of the population with insurance, we would follow equation (3) and multiply estimates in Table 4 by a factor of 8 to 13 to obtain the (instrumental variables-LATE) effects of gaining insurance coverage on prenatal care and birth weight. Doing so yields a relatively large range of estimates of the effect of insurance coverage. For example, if we construct a 95 percent confidence interval around the estimate of the effect of eligibility on the incidence of low-birth weight and multiply by a factor of 10 (roughly midpoint of 8 and 13), we obtain estimates of the effect of insurance coverage on the prevalence of low-birth weight of between -1.4 to 4.7 percentage points, which are large effects relative to the overall sample mean (5.6 percent), and even relative to the mean among low educated mothers (6.5 percent). While this interval likely indicates no beneficial effect of Medicaid on weight for gestational age, we cannot rule out a clinically important improvement (-1.4 percentage points). An alternative way to illustrate this point is to multiply the standard errors of the estimate (linear specification) of the effect of eligibility on low-birth weight by the same factor (10), which yields an estimate of 0.016. Standard errors of this magnitude indicate that we could not reject an effect of insurance coverage on the incidence of low birth-weight for gestational age of less than 3.1 percentage points, which is a non-trivial effect. In short, the absence of a substantial effect of Medicaid eligibility on insurance coverage raises the possibility that we do not have sufficient power to detect reasonably sized effects of eligibility on prenatal care and birth weight. Only if insurance coverage has very large effects would we be able to detect them reliably.[13] Nevertheless, point estimates indicate that Medicaid eligibility does not increase use and adequacy of prenatal care or improve birth weight.

In Table 5, we present the estimates of the effect of Medicaid eligibility on prenatal care and birth weight of white women. Note that estimates of the effect of Medicaid eligibility on insurance for this group indicated high rates of crowd out and little change in the proportion uninsured. Therefore, we would expect very little effect of the Medicaid expansions on prenatal care and birth weight for this sample (assuming that the primary causal mechanism is a change in insurance). In fact, estimates in Table 5 are very similar to those in Table 4, which is not that surprising given that white women constitute nearly 80 percent of the sample used to obtain estimates in Table 4.

[13] On the other hand, if Medicaid expansions affected prenatal care and birth weight through other avenues (such as increased access to medical care for the uninsured because of the cash infusion to providers from Medicaid expansions) than solely by increasing the proportion of the population with insurance, we may expect larger reduced form effects of eligibility.

One way to improve the statistical power of the analysis is to limit the sample to those most likely affected by the Medicaid eligibility expansions (those with a high school degree or less) which we are able to do only in the vital statistics data (thus not on the insurance outcomes) due to a lack of educational information in the NHDS. Those with relatively less education should have higher Medicaid take-up rates and larger changes in insurance coverage than those with more education because the latter group is unlikely to be affected by the eligibility expansions. Therefore, we expect somewhat larger reduced form estimates. Table 6 presents estimates pertaining to the low-educated sample.

Estimates in Table 6 are slightly more suggestive of a beneficial effect of the Medicaid expansions. The coefficients on Medicaid eligibility indicate that expansions in eligibility were associated with an increase in the number and adequacy of prenatal care visits and an increase in birth weight. Estimates for low-birth weight indicate an increase in prevalence—not the expected decrease. However, estimates remain small in magnitude. For example, the estimate of the effect of eligibility on birth weight indicates that a 20 percentage point change in eligibility is associated with a 1.3 gram increase in birth weight. Further, if we assume a scaling factor of 5 (i.e., insurance coverage for this group goes up by 20 percentage points as a result of a 100 percentage point increase in eligibility) for this group, the instrumental variables (LATE) estimator for birth weight is 33 grams (1%).[14] Magnitudes of other estimates are equally small, although implied instrumental variables (IV) estimates related to the prenatal care variables are non-trivial (using a scaling factor of 5). Implied IV estimates for the number of prenatal care visits is 1.5. Estimates in the second row (quadratic specification) do not indicate any significant nonlinear effects that substantially affect inferences.

The final set of estimates we present are for the years 1985 to 1991, the period that witnessed the largest increases in Medicaid eligibility. This is also the period that had the least amount of crowd out of private insurance coverage and the greatest increase in the proportion of the population covered by insurance, although that proportion remains a small fraction (0.12 to 0.14) of the population. Estimates

[14] The confidence interval of the implied instrumental variables (LATE) estimate of the effect of insurance on the incidence of low-birth weight for gestational age is -0.011 to 0.029, which is a relatively large interval.

for this period are in Table 7. Here we again see some evidence that the Medicaid expansions had beneficial effects: increased number and adequacy of prenatal care, and increased birth weight. However, here too, estimates remain quite small in magnitude. Estimates of the effect of Medicaid eligibility on the number of prenatal care visits and birth weight suggests that a 20 percentage point increase in Medicaid eligibility is associated with a 0.21 increase in the number of prenatal care visits and 0.72 gram increase in birth weight. If we use estimates from Table 3 to construct implied IV estimates, the (instrumental variables) effect of gaining insurance coverage on the number of prenatal care visits is large—10.5 visits (10 times 1.046). However, there is no correspondingly large estimate related to birth weight (or low birth weight); the IV estimate for birth weight (low-birth weight) is small—35.9 grams (-0.002).

*Effects of Medicaid Eligibility on Delivery Length of Stay, Cesarean Section and Location of Delivery*

An advantage of the NHDS is that we have information about additional health related outcomes that are not available on Natality files: a dichotomous indicator of whether the infant or the mother stayed in hospital more than two days, whether the mother had a Cesarean Section (C-section), and whether the delivery occurred in a public hospital. A disadvantage of the NHDS is sample size, which is relatively small for our purposes and results in relatively imprecise parameter estimates. Statistical analyses of these outcomes are similar to those previously presented. Results are presented in Table 8.

Estimates in Table 8 indicate that Medicaid eligibility is not associated with length of stay of the mother or the infant; estimates from models that include state trends are small relative to the mean and not statistically significant. For example, a 20 percentage point increase in Medicaid eligibility is associated with a 0.8 percentage point decrease in the probability that an infant stayed more than two days and 1.4 percentage point increase in the probability that a mother stayed more than two days. Relatively long lengths of stay may be an indicator of poor health, so the absence of any consistent association between Medicaid eligibility and length of stay of the mother or infant accords with evidence from birth certificates that found little evidence of an association between Medicaid eligibility and birth weight. Medicaid eligibility is significantly associated with C- section. Focusing on estimates obtained from models that include state-specific trends, a 20 percentage point increase in Medicaid eligibility is associated with a 2.4 percentage point increase in rate of C- section. For C-sections, there appears to be a significant non-linear effect. The marginal effect of Medicaid eligibility on the probability of having a C-section decreases slightly with eligibility; when eligibility is 20 percent, the marginal effect is 0.149, and when eligibility is 30 percent, the marginal effect is 0.114.[15] Finally, estimates in Table 8 indicate that Medicaid eligibility is positively associated with delivering in a public hospital, but the estimate is imprecise and not statistically significant.

**Conclusions**

The relatively large proportion of the US population that is uninsured is a longstanding issue of concern for researchers and policymakers. To address this problem, state and federal governments have dramatically expanded Medicaid over the last twenty years to provide insurance for an increasingly large proportion of poor and near-poor persons, and further expansions have been proposed as a way to expand health insurance.

Expansion of Medicaid may have several benefits besides providing insurance (e.g., indirect income support), but an important rationale for expansion is that it will decrease the proportion of the population that is uninsured, and as a result increase use of health care services and improve health. This is the rationale that motivated the initial expansions in Medicaid that pertained to pregnant women. The question we sought to answer was whether the initial expansions of Medicaid that were targeted at pregnant women achieved these goals. Did the Medicaid expansions significantly reduce the proportion of pregnant women who were uninsured? And did these expansions result in increased use of health care services by pregnant women and improved birth weight? Surprisingly, there has been only one limited analysis of the former question and while there has been several studies of the effect of Medicaid expansions on use of services and infant health, there is no consensus conclusion.

[15] These marginal effects of eligibility on C-section rate imply implausibly large instrumental variables estimates of the effect of becoming insured on C-section.

Our analysis of the effect of the Medicaid expansions on the health insurance coverage of pregnant women suggests that much of the increase in Medicaid participation that resulted from more generous income eligibility thresholds came at the expense of private insurance coverage. Over the entire period, we estimate that 50% of the increase in Medicaid participation came from private insurance. Crowd out of private insurance was particularly severe among white women and at higher levels of eligibility, which is consistent with the higher rates of private insurance coverage among white women and women with higher incomes. For every newly insured woman, the government paid for at least one additional woman who would have otherwise had private insurance. Indeed, for some demographic groups, estimates indicate that approximately 80% of the increase in Medicaid came from private insurance. Nevertheless, the Medicaid expansions did reduce the proportion of pregnant women who were uninsured by approximately 10 to 15% from the level of uninsured in 1985.

Given that the Medicaid expansions had a relatively small effect on the proportion of women uninsured, we would not expect these expansions to have much of an impact on population health. In fact, this is what we found in our analyses of the effect of the expansions on prenatal care, birth weight and hospital length of stay. Changes in Medicaid eligibility were associated with very small changes in prenatal care use, birth weight for gestational age and incidence of low-birth weight for gestational age. However, we also attempted to assess whether gaining health insurance mattered by constructing instrumental variables (LATE) estimates of the effect of gaining insurance on prenatal care and birth weight. Unfortunately, we lacked the statistical power to achieve this goal with certainty. Instrumental variables estimates suggest that the Medicaid expansions may have had some significant effects on prenatal care and birth weight, particularly for low-educated women. We did find some evidence that the Medicaid expansions and the increase in health insurance coverage that resulted from them were associated with increases in the use and adequacy of prenatal care and an increase in birth weight of approximately 35 grams (1%). There was very little evidence that the Medicaid expansions affected the incidence of low-birth weight even for the low-educated group that seemed to increase use of prenatal care services the most. Similarly, we found little evidence that Medicaid eligibility was associated with longer delivery length of stay for either the infant or the mother.

Based on the evidence presented, we believe that the Medicaid expansions for pregnant women had limited effects perhaps because of the significant crowd out of private insurance that occurs, or because they increase the use of services that may have limited impact on birth weight. Future health policy should take these findings into account when developing policies and programs to improve infant health. The difficult part is to identify programs that are effective.

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Table 1
Effects of Medicaid Eligibility on Health Insurance Coverage
National Hospital Discharge Survey 1985 to 1996, All Races

| | Medicaid | | Private Insurance | | Any Insurance | | Uninsured | |
|---|---|---|---|---|---|---|---|---|
| Eligibility | 0.278<br>[0.054]*** | 0.173<br>[0.047]*** | -0.145<br>[0.052]*** | -0.093<br>[0.056]* | 0.133<br>[0.037]*** | 0.08<br>[0.044]* | -0.133<br>[0.037]*** | -0.08<br>[0.044]* |
| Eligibility | 0.254<br>[0.135]* | 0.001<br>[0.107] | 0.086<br>[0.137] | 0.239<br>[0.133]* | 0.34<br>[0.099]*** | 0.239<br>[0.105]** | -0.34<br>[0.099]*** | -0.239<br>[0.105]** |
| Eligibility Squared | 0.041<br>[0.199] | 0.303<br>[0.169]* | -0.401<br>[0.208]* | -0.585<br>[0.205]*** | -0.36<br>[0.155]** | -0.282<br>[0.171]* | 0.36<br>[0.155]** | 0.282<br>[0.171]* |
| State Trends | No | Yes | No | Yes | No | Yes | No | Yes |
| Mean of Dep. Var. | 0.296 | 0.296 | 0.623 | 0.623 | 0.919 | 0.919 | 0.081 | 0.081 |

The table reports coefficients from weighted linear probability models controlling for race, age categories, unemployment rate by state-year with one lead and lag, the fraction of never-married men aged 18 to 39 privately insured by state and year, and race-specific year and state effects. Standard errors are correlated by state-year-race cell and are in brackets. The sample size is 284,179 . ***significant at 1%, * * significant at 5%, * significant at 10%

Table 2
Effects of Medicaid Eligibility on Health Insurance Coverage
National Hospital Discharge Survey 1985 to 1996, White Only

| | Medicaid | | Private Insurance | | Any Insurance | | Uninsured | |
|---|---|---|---|---|---|---|---|---|
| Eligibility | 0.333<br>[0.071]*** | 0.185<br>[0.053]*** | -0.176<br>[0.070]** | -0.083<br>[0.070] | 0.157<br>[0.044]*** | 0.101<br>[0.049]** | -0.157<br>[0.044]*** | -0.101<br>[0.049]** |
| Eligibility | 0.153<br>[0.233] | -0.004<br>[0.159] | 0.14<br>[0.230] | 0.253<br>[0.198] | 0.293<br>[0.152]* | 0.249<br>[0.177] | -0.293<br>[0.152]* | -0.249<br>[0.177] |
| Eligibility Squared | 0.378<br>[0.456] | 0.43<br>[0.338] | -0.664<br>[0.454] | -0.766<br>[0.405]* | -0.286<br>[0.275] | -0.336<br>[0.360] | 0.286<br>[0.275] | 0.336<br>[0.360] |
| State Trends | No | Yes | No | Yes | No | Yes | No | Yes |
| Mean of Dep. Var. | 0.224 | 0.224 | 0.703 | 0.703 | 0.927 | 0.927 | 0.073 | 0.073 |

The table reports coefficients from weighted linear probability models controlling for race, age categories, unemployment rate by state-year with one lead and lag, the fraction of never-married men aged 18 to 39 privately insured by state and year, and race-specific year and state effects. Standard errors are correlated by state-year-race cell and are in brackets. The sample size is 200,217 . ***significant at 1%, * * significant at 5%, * significant at 10%

Table 3
Effects of Medicaid Eligibility on Health Insurance Coverage
National Hospital Discharge Survey 1985 to 1991, All Races

| | Medicaid | | Private Insurance | | Any Insurance | | Uninsured | |
|---|---|---|---|---|---|---|---|---|
| Eligibility | 0.295<br>[0.060]*** | 0.149<br>[0.058]** | -0.132<br>[0.067]** | -0.032<br>[0.064] | 0.163<br>[0.051]*** | 0.117<br>[0.049]** | -0.163<br>[0.051]*** | -0.117<br>[0.049]** |
| Eligibility | 0.076<br>[0.130] | -0.18<br>[0.109]* | 0.183<br>[0.147] | 0.352<br>[0.131]*** | 0.259<br>[0.113]** | 0.172<br>[0.104]* | -0.259<br>[0.113]** | -0.172<br>[0.104]* |
| Eligibility Squared | 0.415<br>[0.208]** | 0.627<br>[0.191]*** | -0.598<br>[0.236]** | -0.732<br>[0.219]*** | -0.183<br>[0.181] | -0.105<br>[0.178] | 0.183<br>[0.181] | 0.105<br>[0.178] |
| State Trends | No | Yes | No | Yes | No | Yes | No | Yes |
| Mean of Dep. Var. | 0.25 | 0.25 | 0.65 | 0.65 | 0.90 | 0.90 | 0.10 | 0.10 |

The table reports coefficients from weighted linear probability models controlling for race, age categories, unemployment rate by state-year with one lead and lag, the fraction of never-married men aged 18 to 39 privately insured by state and year, and race-specific year and state effects. Standard errors are correlated by state-year-race cell and are in brackets. The sample size is 151,232 . ***significant at 1%, * * significant at 5%, * significant at 10%

Table 4
Effects of Medicaid Eligibility on Prenatal Care and Birth Weight
National Natality Files 1985 to 1996, All Races

| | Prenatal Visits | | 1st Trimester Care | | 3rd Trimester Care | | Adequate Prenatal Care | | Birth Weight | | Low Birth Weight | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Eligibility | 0.24642 | -0.07873 | 0.05850*** | -0.00174 | -0.02473*** | -0.00221 | 0.03006 | -0.02520 | -0.99059 | -1.46687 | 0.00139 | 0.00165 |
| | (0.23598) | (0.28266) | (0.01801) | (0.02195) | (0.00576) | (0.00288) | (0.02293) | (0.02044) | (5.19775) | (5.11257) | (0.00136) | (0.00156) |
| Eligibility | -1.22301** | -1.02029** | -0.03759 | -0.07554** | 0.00503 | 0.03339*** | -0.11199** | -0.12953*** | -15.13683 | -3.13465 | 0.00061 | -0.00070 |
| | (0.51456) | (0.47058) | (0.04048) | (0.03591) | (0.01417) | (0.00600) | (0.05056) | (0.03475) | (11.60811) | (10.57654) | (0.00313) | (0.00349) |
| Eligibility Square | 2.58313*** | 1.67696*** | 0.16891*** | 0.13143*** | -0.05232*** | -0.06341*** | 0.24971*** | 0.18581*** | 24.87004 | 2.97065 | 0.00136 | 0.00419 |
| | (0.72976) | (0.55627) | (0.05858) | (0.04239) | (0.01925) | (0.00948) | (0.06994) | (0.04368) | (17.76657) | (15.44569) | (0.00561) | (0.00598) |
| State Trends | No | Yes | No | Yes | No | Yes | No | Yes | No | Yes | No | Yes |
| Mean of Dep. Var. | 11.18 | 11.18 | 0.791 | 0.791 | 0.034 | 0.034 | 0.714 | 0.714 | 3375.6 | 3375.6 | 0.056 | 0.056 |

The table reports coefficients from Ordinary Least Squares models. Standard errors are clustered by state, year and race cells, and reported in parentheses. All models contain state fixed effects, year fixed effects, state by race, and year by race fixed effects. Models control for age (18-24, 25-29, 30-34, relative to 35-39), race (black, other race, relative to white), state unemployment rate (contemporaneous, one-year lag, one-year lead) and state private insurance rate for never married males. Models for birth weight and low birth weight also control for gestation (very premature: <32 weeks, premature: 32-36 weeks, relative to normal gestation: ≥37 weeks). Sample size is 39,238,023 observations. Significance is denoted as follows: *** p-value ≤ 0.01, ** 0.01 < p-value ≤ 0.05, * 0.05 < p-value ≤ 0.10.

Table 5
Effects of Medicaid Eligibility on Prenatal Care and Birth Weight
National Natality Files 1985 to 1996, Whites

| | Prenatal Visits | | 1st Trimester Care | | 3rd Trimester Care | | Adequate Prenatal Care | | Birth Weight | | Low Birth Weight | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Eligibility | 0.15841<br>(0.31335) | -0.16058<br>(0.41737) | 0.04805**<br>(0.02301) | -0.01013<br>(0.03266) | -0.02078***<br>(0.00769) | 0.00360<br>(0.00313) | 0.01994<br>(0.03005) | -0.04092<br>(0.03045) | 4.76063<br>(6.20054) | 2.72313<br>(6.30327) | -0.00001<br>(0.00118) | 0.00058<br>(0.00132) |
| Eligibility | -2.93622***<br>(0.89286) | -1.59610**<br>(0.74841) | -0.20306***<br>(0.06416) | -0.15930***<br>(0.05426) | 0.02963<br>(0.02266) | 0.03679***<br>(0.01033) | -0.29180***<br>(0.08427) | -0.20752***<br>(0.05696) | -32.46898*<br>(17.73650) | -25.65803<br>(15.96171) | 0.00094<br>(0.00354) | -0.00126<br>(0.00384) |
| Eligibility Square | 6.44644***<br>(1.81853) | 3.21223<br>(1.97517) | 0.52309***<br>(0.13173) | 0.33381**<br>(0.15122) | -0.10501**<br>(0.04350) | -0.07427***<br>(0.02187) | 0.64940***<br>(0.16871) | 0.37280**<br>(0.14803) | 77.55607**<br>(34.95605) | 63.51186*<br>(33.15278) | -0.00197<br>(0.00701) | 0.00412<br>(0.00798) |
| State Trends | No | Yes | No | Yes | No | Yes | No | Yes | No | Yes | No | Yes |
| Mean of Dep. Var. | 11.43 | 11.43 | 0.819 | 0.819 | 0.029 | 0.029 | 0.748 | 0.748 | 3422.5 | 3422.5 | 0.045 | 0.045 |

The table reports coefficients from Ordinary Least Squares models. Standard errors are clustered by state, year and race cells, and reported in parentheses. All models contain state fixed effects and year fixed effects. Models control for age indicators (18-24, 25-29, 30-34, relative to 35-39), race (black, other race, relative to white), state unemployment rate (contemporaneous, one-year lag, one-year lead) and state-level private insurance rate for never married males. Models for birth weight and low birth weight also control for gestation (very premature: <32 weeks, premature: 32-36 weeks, relative to normal gestation: ≥37 weeks). Sample size is 31,648,294 observations. Significance is denoted as follows: *** p-value ≤ 0.01, ** 0.01 < p-value ≤ 0.05, * 0.05 < p-value ≤ 0.10.

Table 6
Effects of Medicaid Eligibility on Prenatal Care and Birth Weight
National Natality Files 1985 to 1996, All Races
Low Educated Mothers (High School and below)

| | Prenatal Visits | | 1st Trimester Care | | 3rd Trimester Care | | Adequate Prenatal Care | | Birth Weight | | Low Birth Weight | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Eligibility | 0.67862**<br>(0.26566) | 0.30736<br>(0.24257) | 0.09570***<br>(0.02255) | 0.02650<br>(0.02242) | -0.03324***<br>(0.00656) | -0.01677***<br>(0.00599) | 0.08132***<br>(0.03000) | 0.02865<br>(0.02804) | 8.91898<br>(7.12566) | 6.60549<br>(6.19115) | 0.00149<br>(0.00202) | 0.00181<br>(0.00203) |
| Eligibility | -0.60235<br>(0.62240) | -0.12623<br>(0.59655) | 0.02525<br>(0.05556) | -0.00223<br>(0.05322) | 0.00377<br>(0.01554) | 0.0055<br>(0.01502) | -0.04077<br>(0.07167) | -0.01381<br>(0.06832) | 7.62682<br>(15.83398) | 19.95514<br>(13.53802) | -0.00383<br>(0.00442) | -0.00622<br>(0.00456) |
| Eligibility Square | 2.12028***<br>(0.78218) | 0.72536<br>(0.69373) | 0.11662<br>(0.07159) | 0.04805<br>(0.06152) | -0.06125***<br>(0.02065) | -0.03725**<br>(0.01768) | 0.20210**<br>(0.08674) | 0.07103<br>(0.07503) | 2.13937<br>(22.60920) | -22.33246<br>(19.37143) | 0.00882<br>(0.00730) | 0.01344*<br>(0.00737) |
| State Trends | No | Yes | No | Yes | No | Yes | No | Yes | No | Yes | No | Yes |
| Mean of Dep. Var. | 10.66 | 10.66 | 0.725 | 0.725 | 0.046 | 0.046 | 0.638 | 0.638 | 3329.4 | 3329.4 | 0.065 | 0.065 |

The table reports coefficients from Ordinary Least Squares models. Standard errors are clustered by state, year and race cells, and reported in parentheses. All models contain state fixed effects, year fixed effects, state by race, and year by race fixed effects. Models control for age indicators (18-24, 25-29, 30-34, relative to 35-39), race (black, other race, relative to white), state unemployment rate (contemporaneous, one-year lag, one-year lead) and state-level private insurance rate for never married males. Models for birth weight and low birth weight also control for gestation (very premature: <32 weeks, premature: 32-36 weeks, relative to normal gestation: ≥37 weeks). Sample size is 21,442,593 observations. Significance is denoted as follows: *** p-value ≤ 0.01, ** 0.01 < p-value ≤ 0.05, * 0.05 < p-value ≤ 0.10.

Table 7
Effects of Medicaid Eligibility on Prenatal Care and Birth Weight
National Natality Files 1985 to 1991, All Races

| | Prenatal Visits | | 1st Trimester Care | | 3rd Trimester Care | | Adequate Prenatal Care | | Birth Weight | | Low Birth Weight | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Eligibility | 0.08001<br>(0.31594) | 1.04638*<br>(0.60374) | 0.01966<br>(0.02371) | 0.07672*<br>(0.04322) | -0.00287<br>(0.00318) | 0.00630*<br>(0.00365) | -0.01197<br>(0.02154) | 0.03582<br>(0.03718) | -2.10520<br>(6.85720) | 3.59227<br>(5.62865) | 0.00000<br>(0.00179) | -0.00018<br>(0.00186) |
| Eligibility | -0.5965<br>(0.46209) | 0.23412<br>(0.69377) | -0.02043<br>(0.03425) | 0.02042<br>(0.05000) | 0.00557<br>(0.00677) | 0.02058***<br>(0.00590) | -0.07914**<br>(0.03426) | -0.03814<br>(0.04453) | -12.64051<br>(12.80397) | -2.22474<br>(9.41260) | 0.00007<br>(0.00354) | -0.00035<br>(0.00347) |
| Eligibility Square | 1.28665**<br>(0.63990) | 1.56693**<br>(0.76942) | 0.07625<br>(0.04762) | 0.10859*<br>(0.05599) | -0.01605<br>(0.01115) | -0.02754***<br>(0.00955) | 0.12773**<br>(0.05013) | 0.14268***<br>(0.05093) | 20.03886<br>(21.36659) | 11.23864<br>(15.15429) | -0.00014<br>(0.00721) | 0.00033<br>(0.00688) |
| State Trends | No | Yes | No | Yes | No | Yes | No | Yes | No | Yes | No | Yes |
| Mean of Dep. Var. | 11.00 | 11.00 | 0.775 | 0.775 | 0.038 | 0.038 | 0.696 | 0.696 | 3379.0 | 3379.0 | 0.055 | 0.055 |

The table reports coefficients from Ordinary Least Squares models. Standard errors are clustered by state, year and race cells, and reported in parentheses. All models contain state fixed effects, year fixed effects, state by race, and year by race fixed effects. Models control for age indicators (18-24, 25-29, 30-34, relative to 35-39), race (black, other race, relative to white), state unemployment rate (contemporaneous, one-year lag, one-year lead) and state-level private insurance rate for never married males. Models for birth weight and low birth weight also control for gestation (very premature: $<32$ weeks, premature: 32-36 weeks, relative to normal gestation: $\geq 37$ weeks). Sample size is 22,144,278 observations. Significance is denoted as follows: *** p-value $\leq 0.01$, ** $0.01 <$ p-value $\leq 0.05$, * $0.05 <$ p-value $\leq 0.10$.

Table 8
Effects of Medicaid Eligibility on Health
National Hospital Discharge Survey 1985 to 1996, All Races

| | Infant LOS > 2 Days | | Maternal LOS > 2 Days | | Cesarean Section | | Delivery in Public Hospital | |
|---|---|---|---|---|---|---|---|---|
| Eligibility | 0.084<br>[0.065] | -0.044<br>[0.058] | 0.205<br>[0.058]*** | 0.068<br>[0.047] | 0.116<br>[0.029]*** | 0.12<br>[0.031]*** | -0.076<br>[0.090] | 0.065<br>[0.087] |
| Eligibility | -0.011<br>[0.157] | 0.009<br>[0.131] | -0.065<br>[0.141] | 0.003<br>[0.110] | 0.172<br>[0.078]** | 0.218<br>[0.077]*** | 0.193<br>[0.250] | 0.132<br>[0.214] |
| Eligibility Squared | 0.164<br>[0.235] | -0.093<br>[0.199] | 0.469<br>[0.218]** | 0.115<br>[0.169] | -0.098<br>[0.120] | -0.173<br>[0.118] | -0.467<br>[0.382] | -0.119<br>[0.309] |
| State Trends | No | Yes | No | Yes | No | Yes | No | Yes |
| Mean of Dep. Var. | 0.387 | 0.387 | 0.425 | 0.425 | 0.235 | 0.235 | 0.296 | 0.296 |

The table reports coefficients from weighted linear probability models controlling for race, age categories (except for infants), unemployment rate by state-year with one lead and lag, the fraction of never-married men aged 18 to 39 privately insured by state and year, and race-specific year and state effects. Standard errors are correlated by state-year-race cell and are in brackets. The sample size is 284,179 for maternal deliveries, and 294,721 for infants. ***significant at 1%, * * significant at 5%, * significant at 10%

Figure 1.
Infant Mortality (per 1000 Births) and Low Birth Weight (Per 100 Births)
by Race, 1980 to 2000

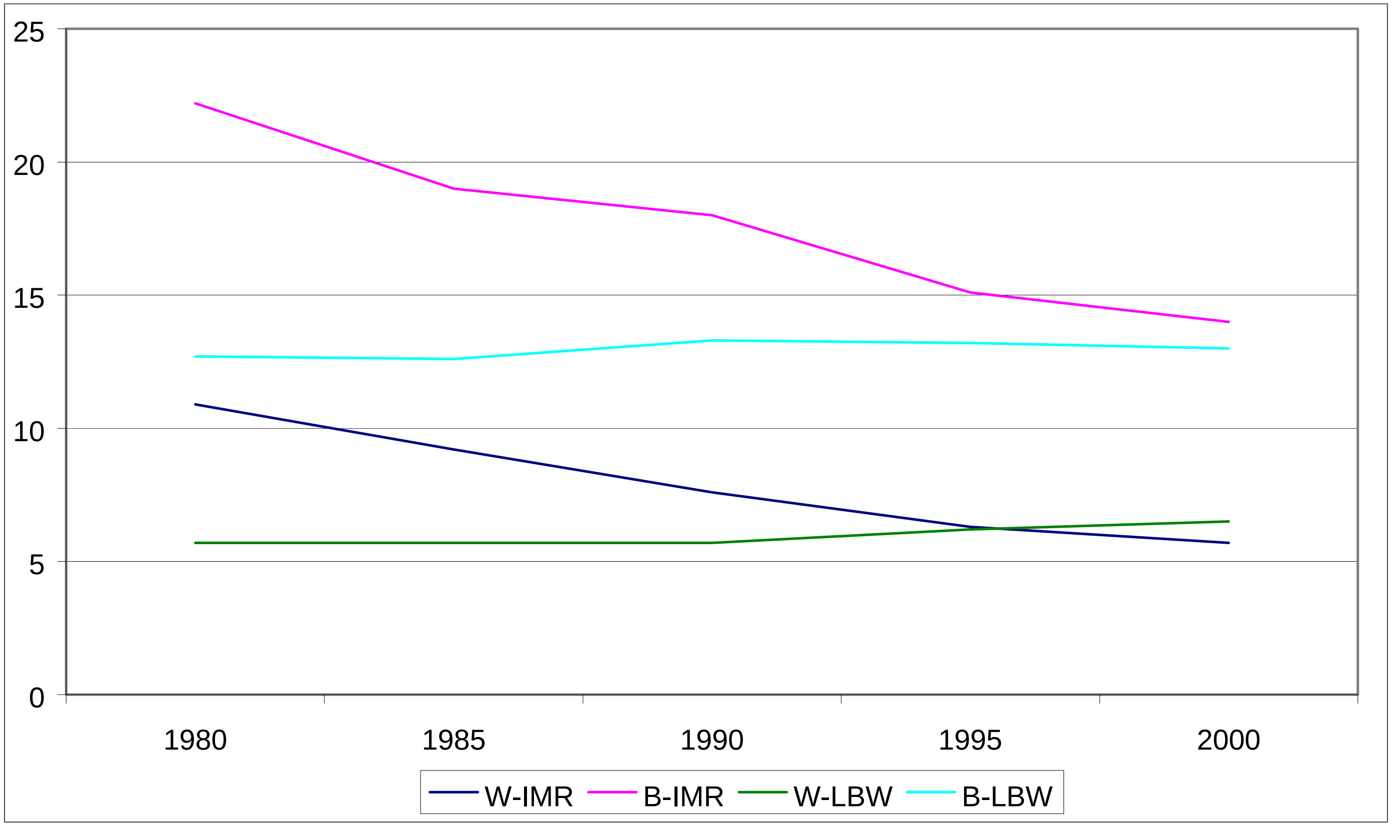


Figure 2

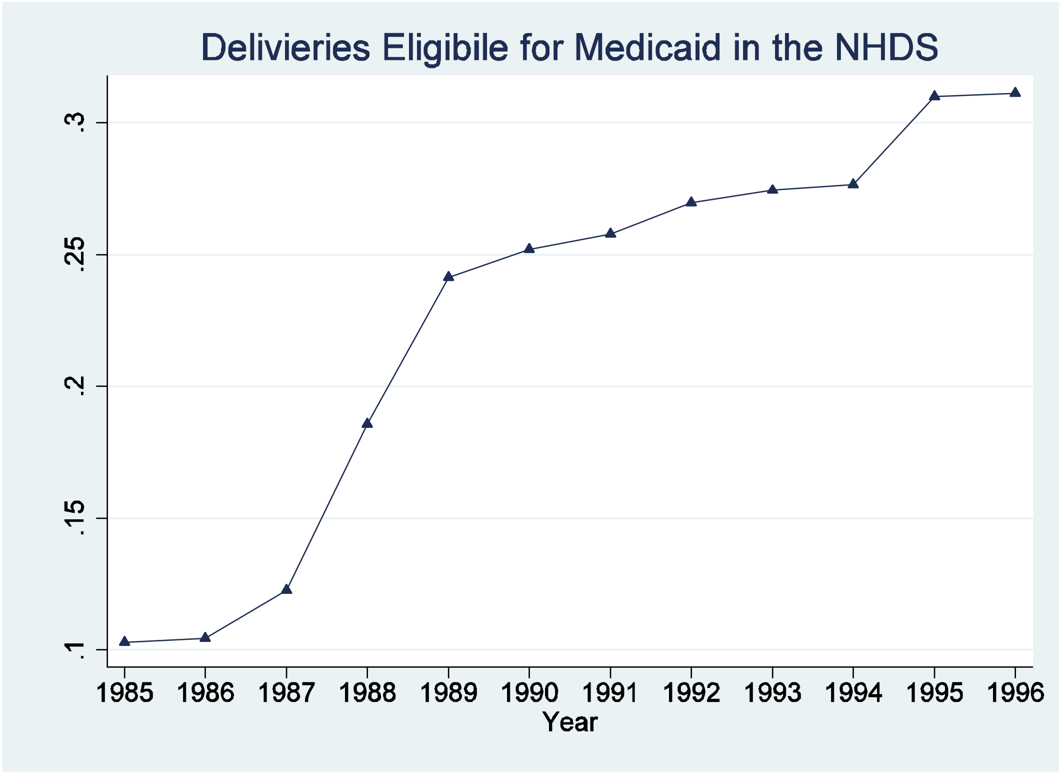
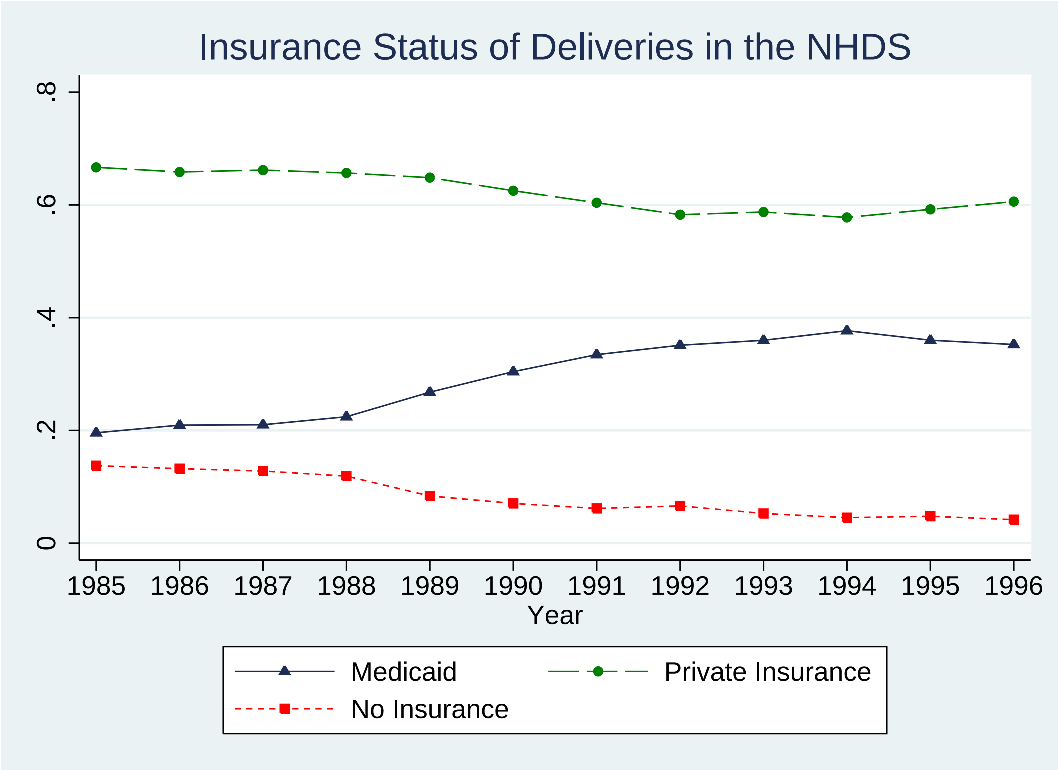

Delivieries Eligibile for Medicaid in the NHDS
.1
.15
.2
.25
.3
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
1995
1996
Year

Figure 3


Insurance Status of Deliveries in the NHDS
0
.2
.4
.6
.8
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
1995
1996
Year
Medicaid
Private Insurance
No Insurance