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**Abstract:** In an effort to increase the use of prenatal care by pregnant women and the utilization of medical care by children, eligibility for Medicaid was expanded dramatically for pregnant women and children during the 1980s and early 1990s. By lowering the costs of prenatal care, delivery, and child health care for some individuals, Medicaid expansions may prompt some women to give birth who otherwise would not have children or lead some women to have more children than they otherwise would have. This study uses natality data from 1983 to 1996 to examine the relationship between a state's eligibility threshold for Medicaid and birth rates among various groups. The results suggest that expansions have significant and sizable effects on births. A 10 percentage point increase in the eligibility threshold is associated with a 1.4 percent increase in the birth rate among nonblack women and a 1.0 percent increase among black women. Between 1983 and 1996, the expansions appear to have led to an average increase in the birth rate of about 10 percent.

JEL classification: I18, I38, J13

Key words: births, fertility, Medicaid, health insurance

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# The Effect of Medicaid Eligibility Expansions on Births

## 1. Introduction

Eligibility for Medicaid, the government health insurance program for low-income individuals, expanded dramatically for pregnant women and children during the 1980s and early 1990s. The goal of the expansions was to increase the use of prenatal care by pregnant women and the utilization of medical care by children. The expansions decoupled eligibility for Medicaid from receipt of cash welfare benefits under the Aid to Families with Dependent Children (AFDC) program, the traditional means of obtaining Medicaid coverage that had largely limited eligibility to poor female-headed households with children. As Medicaid eligibility was extended to other groups, the fraction of women and children eligible for Medicaid rose rapidly. The fraction of women of child-bearing age eligible for Medicaid coverage of pregnancy-related services more than doubled between 1987 and 1992, while the fraction of children eligible for full Medicaid coverage rose by at least 50% (Cutler and Gruber, 1996). This dramatic increase in Medicaid eligibility, which varied in extent and timing across states, is used here to measure the effect of increases in program eligibility thresholds on birth rates.

By extending health insurance benefits to pregnant women and children not previously eligible for Medicaid, the eligibility expansions lowered the cost of health care for individuals not previously covered by any health insurance program. In addition, hundreds of thousands of women and children dropped private insurance coverage as Medicaid eligibility expanded, suggesting that the expansions lowered costs for some individuals already covered by health insurance (Cutler and Gruber, 1996). By lowering the costs of prenatal care, delivery, and child health care for some individuals, the expansions may have prompted some women to give birth

who otherwise would not have had children, or led some women to have more children than they otherwise would have had. Alternatively, the expansions could have caused women to alter the timing of their fertility.

Previous research has examined the responsiveness of women's fertility to financial considerations in several ways. Studies have investigated the link between birth rates and health insurance coverage, welfare benefits, and women's earnings. Although the evidence is somewhat mixed, particularly concerning the effects of welfare on fertility, many researchers have found that fertility is responsive to economic incentives.

Previous studies suggest that health insurance coverage affects birth rates. In the Rand Health Insurance Experiment, women randomly assigned to receive free medical care for three to five years had 29% more births than women assigned to cost-sharing medical plans (Leibowitz, 1990). This finding is not clear evidence of increased fertility, however, because the participants knew that the experiment would only run for several years; these women may only have altered the timing of births and not their total number of births. The Medicaid expansions, in contrast, are permanent unless the law is changed. Using data from 15 states between 1987 and 1991, Joyce, Kaestner and Kwan (1998) found that the Medicaid expansions were associated with a 5% increase in the birth rate among unmarried white women aged 19-27 who had not attended college but did not find a significant effect among comparable black women. Yelowitz (1994) used Current Population Survey data from 1989 to 1992 on women aged 15-44 and found that Medicaid eligibility raised the probability that a woman gave birth by about 5%. However, this effect was due entirely to Medicaid coverage of older children, not to coverage of the delivery and medical care for the newborn child.

Researchers have not yet reached a consensus on the relationship between cash welfare benefits and births. Jackson and Klerman (1996) found moderate effects of Aid to Families with Dependent Children (AFDC) benefits on birth rates during 1975-1992.[1] The estimated effects tended to be larger among blacks than among whites and larger for later births than for first births, while there were few clear differences across education groups. Other studies, however, have found little evidence that welfare benefits influence birth rates (Ellwood and Bane, 1985; Schultz, 1994).

Women's fertility also appears to depend on earnings. Among both black and white women aged 15-65, the number of children ever born to a woman is negatively related to her predicted wage and positively related to her potential partner's predicted wage (Schultz, 1994). In addition, nonmarital fertility among young women is negatively associated with current economic opportunities, expected future earnings, and expected future family income (Duncan and Hoffman, 1990; Ribar, 1998). Fertility also appears to be sensitive to tax policy (Whittington, Alm and Peters, 1990).

We add to the previous research on Medicaid eligibility and fertility in several ways. By using natality data stratified by marital status and parity as well as by age and race, we examine whether the estimated effects differ among groups with different eligibility rates for Medicaid. This comparison of the estimated effects among groups with high and low rates of Medicaid eligibility gives a quasi-difference in differences estimate of the effect of increased Medicaid eligibility on fertility. Previous research, in contrast, did not assess whether the estimated effects differ across groups with different eligibility rates for Medicaid except across races. In addition, our data includes the expansions that took place during the mid-1980s and were more narrowly

[1] Among unmarried white women aged 25-29, for example, a 10 percent increase in AFDC benefits was associated with a 2.6 percent increase in the birth rate, compared to a 0.6 percent increase among married women (Jackson and

targeted at poor women as well as the later, broader expansions; previous studies primarily focused on the later, broadly targeted phases of the Medicaid expansions. Unlike previous studies, we control for AFDC maximum benefit levels in the regressions, allowing us to identify the effect of changes in Medicaid eligibility thresholds on birth rates separately from any effect of changes in welfare generosity.

We begin by presenting a chronology of the Medicaid expansions in Section 2. Section 3 describes the data, which are based on birth certificates over 1983 to 1996, and outlines the estimation methodology. We estimate whether the birth rate among various groups rises as a state's eligibility threshold for Medicaid increases for that group. Section 4 presents the results, which suggest that increases in the Medicaid eligibility threshold raise birth rates among most groups of women. Section 5 concludes and discusses implications of the results.

## 2. Chronology of Medicaid expansions

Until the mid-1980s, Medicaid coverage was generally linked to receipt of AFDC benefits, which effectively limited eligibility to female-headed households with children with incomes low enough to qualify for AFDC. In states that opted to have an AFDC-Unemployed Parent (AFDC-UP) program, two-parent households that met the program income requirements also qualified for Medicaid. Beginning in 1984, a series of laws expanded pregnant women and children's eligibility for Medicaid coverage. Table 1 summarizes the eligibility expansions.

In the early phase of the expansion, states were required to extend Medicaid coverage to several groups that did not meet the family structure requirements of the AFDC program. During this first phase of the expansion, individuals still had to meet the AFDC resource requirements, which required income to be well below the poverty level in most states, to be

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Klerman, 1996).

eligible for Medicaid. We call this first phase "narrowly targeted expansions" because these expansions extended coverage to narrowly defined groups of low-income women.[2] The fraction of women aged 15-44 eligible for Medicaid rose less than 5 percentage points during the narrowly targeted phase of the expansion (Currie and Gruber, 1996).

Beginning in April 1987, states were first given the option and then required to extend Medicaid coverage to pregnant women and children with incomes substantially above the AFDC resource limits in most states, regardless of family structure. States were required to extend coverage by July 1989 to all pregnant women with incomes below 75% of the federal poverty line, and to 133% by April 1990. We call this second phase "broad expansions." Eligibility rates among women aged 15-44 more than doubled during this phase of the expansion (Currie and Gruber, 1996).

States differed in the timing and extent of Medicaid eligibility expansions. In September 1984, for example, 23 states and the District of Columbia had AFDC-UP programs in place and offered Medicaid benefits to pregnant married women who met the program's resource requirements. One of the narrowly targeted expansions required the remaining 27 states to offer Medicaid coverage by October of that year to pregnant women who had a spouse present in the household and met their state's AFDC resource requirements. There was also variation across states in the timing of the broad eligibility expansions; in January 1988, for example, 22 states had opted to expand Medicaid eligibility for pregnant women beyond the AFDC threshold in that state. In February 1996, 28 states had opted for coverage of pregnant women with incomes of

[2] In addition to the expansions listed in Table 1, Medicaid eligibility was extended to teenagers in families with incomes below the AFDC cutoff, regardless of family structure, and to medically needy pregnant women (women with incomes above the AFDC cutoff but with medical expenses that brought net income below the AFDC cutoff) during the first phase of the expansions (Currie and Gruber, 1996).

185% of the federal poverty level or higher, while the remaining states at least met the minimum level of 133%.

The broad phase of the expansions dramatically raised the eligibility threshold beyond the AFDC income threshold in most states. In Texas, for example, the monthly income limit for a 3-person family in 1983 was $148, or 22% of the poverty level for a family of that size. Texas raised its income cutoff for pregnant women to 130% of the poverty level in September 1988, to 133% in July 1990, and to 185% in March 1992. In Vermont, in contrast, the monthly income limit in 1983 was $507, or 74% of the poverty level. Vermont raised its cutoff to 100% in October 1987 and to 200% in July 1988, lowered it to 185% in July 1990, and raised it to 200% in January 1993.[3] Such variation across states in the timing and extent of expansions, as well as in AFDC resource limits, is used to identify the effect of Medicaid eligibility expansions on births, as described below.

Pregnant women covered under the expansions are covered only for services related to the pregnancy, while children receive the same coverage as AFDC recipients. In some states, pregnant women's coverage includes abortion services. States have the option to cover abortions in their Medicaid program but do not receive federal matching funds for most abortions. In almost all states with Medicaid programs that cover abortions, women eligible for Medicaid under the expansions are also eligible for Medicaid-covered abortions.[4]

[3] States had the option of using their own funds to extend coverage beyond groups eligible under the federal law. Details on implementation of the expansions and data on AFDC income thresholds are from National Governors Association (various years) and US House of Representatives (various years).

[4] In 1994, for example, Alaska, California, Connecticut, D.C., Hawaii, Idaho, Illinois, Maryland, Massachusetts, Minnesota, New Jersey, North Carolina, Oregon, Vermont, Washington, and West Virginia had Medicaid programs that covered abortions (Sollom, 1995). The Medicaid program in New York covered abortions for AFDC recipients but not for pregnant women eligible for Medicaid because of the expansions; New York City opted to cover medically necessary abortions for pregnant women covered under the expansions.

## 3. Data and methodology

### *3.1 Birth rates*

The birth data are from national birth certificate tapes produced by the National Center for Health Statistics (NCHS). We use data on all live births between 1983 and 1996 to women aged 15 to 44 residing in the 50 states and the District of Columbia whose age, race, marital status, and number of previous live births were reported. For births in states that did not record marital status on birth certificates, NCHS imputed the mother's marital status.[5] In 1983 and 1984, the tapes included one-half of births for several states; for those states, we counted each record twice when compiling our sample.[6] The data used here are nearly complete even after dropping births with missing data on the mother; they encompass from a low of 96.1% of total births to U.S. female residents aged 15-44 in 1989 to a high of 99.3% in 1987.

We use the log of the birth rate in a given state and quarter as the outcome measure. The birth rate is created by dividing the number of births by the population of women in the relevant age and racial group (in thousands). The population data are from the National Cancer Institute and are available for 5-year age groups by race. The analysis focuses on 5-year age groups and distinguishes between blacks and nonblacks. Stratifying by age and race eliminates the need to control for the age and racial composition of the population; birth rates may differ over time within the total female population because of changes in the age structure or racial makeup of the population, creating the need to stratify the sample by age and race or to control for the age and racial distribution of the population.

[5] Jackson and Klerman (1996) discuss the accuracy of the marital status imputation procedure.
[6] NCHS imputed marital status for California, Connecticut, Michigan, Nevada, New York, and Texas. One-half of births in Arizona, California, Delaware, the District of Columbia, and Georgia are reported in the data in 1983 and 1984.

The first column of Table 2 reports the mean birth rate during the sample period for 5-year age groups by race. Birth rates rise and then decline with age, with fertility peaking at age 25-29 among white women and at age 20-24 among black women. Birth rates are higher among black women under age 25 than among young white women, while the converse tends to hold among older women.

We also examine the relationship between Medicaid expansions and birth rates for different marital status and parity groups. The responsiveness of women's fertility to Medicaid eligibility may differ across marital status and parity groups for several reasons. Incomes, and therefore eligibility for Medicaid, may differ systematically across demographic groups; married women tend to have more financial resources, on average, than do unmarried women. In addition, the potential availability of private insurance coverage through a spouse may cause married women's fertility to be less responsive to Medicaid eligibility than the fertility of unmarried women. Medicaid coverage was expanded for some older children as well as for pregnant women and infants, particularly during the later phases of the eligibility expansion. These expansions may have lowered health care costs for women who already have children and caused their fertility to be more responsive to Medicaid expansions than the fertility of women without previous births. Alternatively, the fertility of women who already have children may be less responsive to financial considerations because they may be more likely than nulliparous women to have already achieved their desired family size.

The denominator used to create the birth rates when the data are stratified by parity and marital status is the number of women of the relevant age and race in the state in a given year. We use this number as the denominator for two reasons. First, the size of the female population by marital status and parity is not available except for the years of the decenniel Census. In

addition, as Jackson and Klerman (1996) note, the denominator of the birth rate should not be stratified by marital status if marital status and fertility are jointly determined. In this analysis, the birth rate among unmarried nonblack women aged 25-29 who do not already have children, for example, is the number of first births to unmarried nonblack women aged 25-29 divided by the total number of nonblack women aged 25-29.

*3.2 Medicaid eligibility*

We use two variables to measure the relationship between the Medicaid expansions and birth rates. First, the regressions include a dummy variable equal to one if a state has expanded eligibility for Medicaid beyond women who qualify for AFDC benefits, and zero otherwise. This variable captures the main effect of expanding Medicaid eligibility to women previously not eligible for the program. In 1983, this variable is equal to zero for all states, and the variable is equal to one for all states by the second quarter of 1990.

The regressions also include a linear variable that measures the eligibility threshold for Medicaid. The eligibility threshold for Medicaid is measured using the income limit for eligibility as a fraction of the federal poverty level. During the narrowly targeted first phase of the Medicaid expansion, this limit was the income limit for eligibility for Medicaid benefits under the AFDC program in a state; in most states, the income limit was the maximum AFDC benefit, although a few states had more generous eligibility rules for Medicaid than for AFDC (Currie and Gruber, 1996: Yelowitz, 1995). After the broadly targeted Medicaid eligibility expansions began in April 1987, the income limit was specified as the maximum percentage of the poverty level at which a pregnant woman was eligible for Medicaid. The variable is equal to zero if a state has not yet expanded Medicaid eligibility. This variable measures the incremental

effect of changes in the eligibility threshold, given that a state has expanded Medicaid eligibility, and is of interest because it measures the marginal effect of Medicaid eligibility thresholds on fertility. By also including a variable that measures the AFDC maximum benefit level, we are able to separate out the effects of Medicaid expansions on fertility from possible effects of welfare on fertility.

Columns 2 and 3 of Table 2 report the mean of the Medicaid expansion dummy variable and the conditional mean of the eligibility threshold variable for each age/racial group. Medicaid eligibility was expanded for over 64% of the sample period among nonblack women aged 15-44, and for 67% of the sample period among black women aged 15-44. The mean threshold for Medicaid eligibility, given that eligibility was expanded, was 157% of the poverty line among nonblack women aged 15-44 and 156% of the poverty line among black woman aged 15-44. The differences across age and racial groups reflect differences in residential patterns and in the fraction of women in the age/racial group who are married or have not had a previous birth and therefore may be affected by the narrowly targeted phase of the expansion. For comparison, mean maximum AFDC benefits as a percentage of the federal poverty level are 48% among nonblack woman and 43% among black women during 1983-1996.

The two variables we use to measure the extent of Medicaid expansions serve as proxies for the fraction of women in a demographic group who are eligible for Medicaid. If Medicaid eligibility affects women's fertility, the fraction of women who are eligible for Medicaid should be associated with the birth rate. In the regressions below, we include variables measuring whether Medicaid eligibility has been expanded and level of the eligibility threshold but do not include a measure of the eligibility rate. We do not directly include in the regressions the

fraction of women in a group who are eligible for Medicaid for several reasons.[7] Currie and Gruber (1996) point out that using the estimated fraction of eligible women in each state and year in cells disaggregated by demographic factors, such as age and race, is infeasible because of small cell sizes. In addition, the Current Population Survey (CPS) data sets used to measure eligibility rates undersample very poor women, whose fertility may be quite responsive to the Medicaid expansions. Using eligibility shares calculated from the CPS data could therefore introduce measurement error.

If our linear variable that measures the eligibility threshold for Medicaid under the expansion proxies for the eligibility rate, it should rise as the fraction of women who are eligible for Medicaid increases. Figure 1 shows the annual average eligibility threshold for Medicaid as a percentage of the federal poverty level and the estimated percentage of women aged 15-44 eligible for Medicaid. The two series clearly move together, indicating that increases in the eligibility threshold for Medicaid are associated with increases in the fraction of women eligible for Medicaid coverage. Our linear eligibility threshold variable appears to be a reasonable proxy for women's eligibility for Medicaid.

The fraction of women who are eligible for Medicaid and the effect of the expansions on eligibility differ greatly across demographic groups. Differences in eligibility rates reflect program rules as well as differences in income and state of residence. Table 3 reports the estimated fraction of women who would have been eligible for Medicaid coverage if they had become pregnant in 1983 and 1996. In 1983, before the expansions began, Medicaid eligibility

[7] Previous studies have used data from the March Current Population Survey (CPS) to estimate the fraction of women who are eligible for Medicaid (e.g., Currie and Gruber, 1996; Cutler and Gruber, 1996). The surveys contain about 30,000 women aged 15-44; estimating the eligibility rate at the state level for five-year age groups by race, marital status and parity would result in few or zero observations in a large number of cells. We use data from the March CPS to calculate the national fraction of women eligible for Medicaid shown in Figure 1 and reported in Table 3. Eligibility was calculated using the algorithm described in Currie and Gruber (1996); details are available on request.

was based solely on AFDC eligibility. About 24% of unmarried nulliparous nonblack women would have been eligible if they had gotten pregnant in 1983, for example, compared with 38% of unmarried nonblack women who already had children. Married women who did not already have children were generally not eligible for Medicaid before the eligibility expansions began in 1984.[8] Between 1983 and 1996, eligibility rates expanded for each demographic group, with the largest increases occurring among unmarried women who do not already have children.

*3.3 Empirical methodology*

We employ panel data techniques to estimate the relationship between Medicaid eligibility thresholds and birth rates. The basic regression model is

$$\ln \text{ birth rate}_{st} = \beta \text{Medicaid expanded}_{st-3} + \gamma \text{Medicaid threshold}_{st-3} + \delta X_{st-3} + \phi A_{st-2} + \varepsilon_{st}, \quad (1)$$

where *s* indicates state and *t* indicates time. The regressions are estimated using quarterly data for the 50 states and the District of Columbia over 1983 to 1996, or 2856 observations per age and racial group. The birth rates are annualized by multiplying them by four before taking the log.

As discussed above, two variables are included to measure the effect of Medicaid expansions: a dummy variable equal to one if a state has expanded its Medicaid program and zero otherwise, and a linear variable measuring the eligibility threshold for Medicaid as a percentage of the federal poverty level. The two Medicaid variables are lagged by three quarters

[8] Currie and Gruber (1996) report that such women may have been eligible on a state-by-state basis but were not covered under federal program rules.

because policy at the time of conception should affect women's fertility behavior more than policy at the time of birth.

The regressions include several other variables to control for economic conditions and other factors that may affect fertility. The unemployment rate, average hourly earnings of manufacturing workers as a percentage of the federal poverty level for a family of three, and the employment-to-population ratio are included to control for economic conditions. The maximum AFDC benefit as a percentage of the federal poverty level is included to control for any effect of welfare benefits on fertility.[9] These variables are lagged by three quarters and are represented by $X_{st-3}$ in equation (1). A dummy variable measuring whether a state allows Medicaid funding of abortions is also included because Medicaid funding of abortions may affect birth rates (Klerman, 1996). This variable, $A_{st-2}$, is lagged by two quarters to reflect Medicaid financing of abortions toward the end of the first trimester of a pregnancy.

The regressions also include state, year and quarter fixed effects. The state fixed effects control for time-invariant differences in fertility across states, the year fixed effects control for year-specific differences in fertility common to all states, such as business cycle conditions, and the quarter fixed effects control for seasonal differences in fertility common to all states and years. The regressions are estimated using ordinary least squares (OLS), and observations are weighted using the population size.[10]

[9] Among women who have not already had a live birth, we use the ratio of AFDC benefits for a 2-person family with 1 adult to the federal poverty level for a family of 2 persons. Among women who report a previous live birth, we use the ratio of AFDC benefits for a 3-person family with 1 adult to the federal poverty level for a family of 3. When the data are not stratified by parity and marital status, the AFDC variable is a weighted average, where the weights are the fraction of first births and the fraction of higher-order births.

[10] The results are generally robust to including state-specific time trends. The coefficients tend to fall slightly in magnitude, but the pattern of the estimated effects and the significance levels are similar.

## 4. Results

The results suggest that the increases in the income eligibility cutoff lead to higher birth rates and that the Medicaid eligibility expansions raised fertility among most groups of women.

### *4.1 Age and racial groups*

Table 4 reports the estimated relationships between the birth rate and the Medicaid eligibility variables for various age and racial groups. The estimated coefficients on the dummy variable for whether Medicaid was expanded (the main effect) shown in the first column do not have an economic interpretation by themselves when the linear eligibility variable is included because the dummy variable always equals one when the linear eligibility variable does not equal zero. The estimated coefficients on the Medicaid threshold variable (column 2) give the incremental effect of further raising the income cutoff, given that the state has already expanded eligibility. Because the dependent variable is the log of the birth rate, the incremental effects are interpreted as the percent change in the birth rate if the eligibility threshold increases by 1 percentage point, conditional on eligibility already being expanded.

Raising the income cutoff for Medicaid eligibility, given that eligibility has already been expanded, appears to increase the birth rate among all age and racial groups. A 10 percentage point increase in the threshold is associated with 1.4% increase in the birth rate among nonblack woman aged 15-44 and a 1.0% increase in the birth rate among black women aged 15-44. The incremental effect rises with age among black women, while there is no clear pattern by age among nonblacks. All of the estimated incremental effects are statistically significant below the 1% confidence level. In addition, the coefficients on the two Medicaid policy variables are

jointly significantly different from zero, as shown by the F-test statistics reported in the third column of Table 4.

The estimated average effect of the expansions on birth rates is also of interest because it suggests the magnitude of the effect of the expansions that occurred through 1996 on birth rates. The average effect of the eligibility expansions that occurred during the sample time period can be calculated by evaluating the estimated incremental effect at the mean eligibility threshold, given that eligibility has been expanded, and adding the estimated main effect.[11] Column 4 of Table 4 shows the estimated mean effects, calculated using the means reported in Table 2 and the estimated coefficients reported in the first two columns of Table 4. The Medicaid expansions that took place between 1983 and 1996 appear to have raised the birth rate among both black and nonblack women aged 15-44 by about 10%. There are no clear differences across age or racial groups in the average effect of the expansions on birth rates.

The estimated relationships between maximum AFDC benefits and birth rates are mixed. In results not reported here, the estimated coefficients on the AFDC variable are positive among nonblacks and negative among blacks but generally not statistically significant. The only subgroups for which the AFDC coefficients are significant at the 5% confidence level indicate a positive association between AFDC benefits and birth rates among nonblack women aged 25-39 and a negative association between AFDC benefits and birth rates among black women aged 35-44.

[11] The standard errors on the mean effects reported in Tables 4-6 take into account the standard errors on the main and incremental effects but do not incorporate uncertainty about the sample mean and therefore should be regarded as lower bounds.

*4.2 Parity and marital status groups*

Table 5 reports the results of estimating equation (1) separately for unmarried and married women who do not already have children, and Table 6 reports results stratified by marital status for women who already have children.[12] As discussed above, the birth rates by parity and marital status are the number of births to women in the particular parity and marital status group divided by the total number of women in the age/racial group. Stratifying the data by parity and marital status allows us to examine whether the effects of Medicaid eligibility expansions are different for unmarried women than for married women and for first births than for later births.

Given that eligibility has been expanded, the estimated effect of further raising the eligibility threshold is similar among unmarried and married women having their first birth and married women having a second or higher birth. Most of the estimates in Tables 5 and 6 indicate that a 10 percentage point increase in the eligibility threshold raises by 0.9 to 1.5% among nulliparous women and among married women who already have children. The effect of raising the income cutoff is considerably different among unmarried women who already have children. The incremental effect of eligibility increases is much smaller among unmarried women having their second or higher birth than among the other groups, and the estimates reported in column 1 of Table 6 are not statistically significant for many of the age/racial groups of unmarried women who already have children.

[12] Tables 5 and 6 do not report the estimated coefficients for the dummy variable measuring whether Medicaid eligibility has been expanded in order to conserve space and because the coefficients do not have an economic interpretation. The results, along with the sample means used to calculate the mean effects in Table 5, are available on request.

The mean effect of the expansions that occurred through 1996, in contrast, is largest among unmarried women who already have children. Evaluating the estimated coefficients at the sample means, the Medicaid expansions appear to have led to a 23% increase in the birth rate among unmarried women who already have children. The estimated mean effects are smaller among unmarried and married women having their first birth and married women having a second or higher birth and suggest that the Medicaid expansions led to a 6 to 15% increase in the birth rate among these groups. The larger effect among unmarried women who already have children is consistent with the results of Yelowitz (1995), who found that most of the effect of the expansions was through extension of benefits to older children.

Among black women, birth rates among married women with children are more sensitive to increases in the Medicaid eligibility threshold than birth rates among unmarried women or among married women who do not already have children. The average effect of the Medicaid expansions was also larger among married black women who already have children than among nulliparous black women.

In results not shown here, the estimated relationship between AFDC benefits and birth rates is not consistent across groups. The results suggest a significant positive relationship between welfare benefits and birth rates to nonblack unmarried women in both parity groups. Some results also indicate a significant negative relationship between cash welfare benefits and birth rates to married black women, particularly among those who already have children.

## 5. Comment

Beginning in 1984, the eligibility of pregnant women and children for Medicaid increased dramatically. Pregnant women made eligible for Medicaid were covered for prenatal

care, delivery and postpartum care, while children received full Medicaid benefits. The expansions lowered the cost of health care for individuals who met the new income limits and were previously uninsured and may have also lowered health care costs for some individuals with private insurance. This reduction in health care costs may have led to an increase in fertility.

Our results indicate a significant positive relationship between birth rates to nonblack and black women aged 15-44 and increases in the Medicaid eligibility threshold as a percentage of the poverty line during 1983 to 1996. The number of births per 1000 women rises as the Medicaid eligibility threshold, and presumably the fraction of women who qualified for Medicaid, increases. The results suggest that a 10 percentage point increase in the income cutoff, given that eligibility had already been expanded, increases the birth rate by about 1% among both nonblack and black women. The expansions that occurred between 1983 and 1996 are associated with an average increase in the birth rate of about 10%.

This result is considerably larger than the 5% increase in birth rates reported in two previous studies (Yelowitz, 1994; Joyce, Kaestner and Kwan, 1998). This study focuses on the relationship between birth rates and expansions in the Medicaid eligibility threshold as well as changes in the family structure eligibility requirements for Medicaid. Joyce, Kaestner and Kwan (1998) used dummy variables to indicate whether states expanded Medicaid eligibility, while Yelowitz (1994) used an imputed value of health insurance based on family structure and state Medicaid policy to measure the effect of the Medicaid expansions. This difference in the variables used to measure Medicaid may explain the difference in the results. In addition, we control for AFDC benefits in a state, allowing us to separate the effects of Medicaid expansions from those of cash welfare benefits.

The results indicate differences among population subgroups. The group that had the highest eligibility rate before the expansion, unmarried women who already have children, has smaller responses to increases in the Medicaid eligibility threshold than do women who do not already have children and married women who already have children. However, the estimated mean effect of the expansions that occurred from 1983-1996 is largest among unmarried women who already have children.

This finding is consistent with several possibilities. Because they are more likely to qualify for other social insurance programs, unmarried women with children may be more aware of Medicaid program changes than other women and therefore more responsive to program changes. In addition, the Medicaid expansions raised older children's eligibility for Medicaid. Increased coverage of these children may partially underlie the relatively large estimated mean effects among unmarried women with children, as Yelowitz (1994) suggests. These women are also less likely to have private insurance coverage than married women, whose spouses may have coverage through their employers. The greater availability of private coverage among married women with children may cause their fertility to be less responsive to Medicaid expansion than fertility among unmarried women with children.

Our results also indicate that unmarried women with children are less responsive to further increases in the eligibility threshold, given that eligibility has already been expanded, than other groups of women. These women tend to have lower incomes than women in the other parity and marital status groups. Further increases in the income cutoff therefore may have relatively little effect on eligibility rates among unmarried women with children. In addition, the Medicaid expansions coincided with a decline in real welfare benefits in many states. This decline in real welfare benefits caused the insurance value of government programs, or the

"safety net," to fall for many women. Unmarried women with children would have been disproportionately affected by such a decline because they have relatively high rates of eligibility for such programs. The fall in the insurance value of the government programs may have contributed to the relatively small fertility responses to increases in the Medicaid eligibility threshold that we observe among unmarried women with children.

We find that married black women who already had children tend to have larger average increases in birth rates as a result of the expansion than black women who did not already have children. This again suggests that the expansion of Medicaid coverage to existing children may have led some women to have additional births they would not have had absent the expansion.

Our results may affect the interpretation of studies that examined the effect of the Medicaid expansions on prenatal care utilization and neonatal health outcomes (Currie and Gruber, 1996; Piper et al., 1990; Haas et al., 1993). Our findings suggest that expansions prompted some women who otherwise would not have given birth to have a child or to have children sooner than they would have otherwise. If these children had worse health outcomes than children who would have been born regardless of Medicaid policy, the Medicaid expansions would be observed to be associated with smaller improvements in outcomes than would be the case if the composition of women giving birth had not changed. The Medicaid expansions therefore may have led to larger improvements in birth outcomes than previous research suggests. Controlling for changes in the population of women giving birth when estimating the effect of Medicaid eligibility on birth weight and other outcomes is an important area for future research.

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**Table 1. Summary of Medicaid program expansions affecting pregnant women and children**

*Narrowly targeted expansions:*
Effective October 1984, states required to cover first-time pregnant women if they would be eligible for AFDC if their children were already born; states without AFDC-UP programs required to cover married pregnant women if they met AFDC program resource guidelines.

Effective July 1986, states required to cover pregnant women regardless of family structure if they met AFDC program resource guidelines.

*Broad expansions:*
Beginning April 1987, states allowed to cover pregnant women and children under age 2 with incomes up to 100% of federal poverty line.

Beginning July 1988, states allowed to cover pregnant women and children under age 2 up to 185% of poverty line and children under age 5 up to 100%.

Beginning October 1988, states allowed to cover children under age 8 in families with income up to 100% of poverty line.

Effective July 1989, states required to cover pregnant women and children under age 2 with incomes up to 75% of the poverty line.

Effective April 1990, states required to cover pregnant women and children under age 6 up to 133% of the poverty line; given option to extend coverage to 185%.

Effective July 1991, states required to begin phasing in coverage for all children under age 19 up to 100% of the poverty line.

**Table 2. Sample means, by age and race**

| | Birth rate | Medicaid expanded | Medicaid threshold as % fpl, given expansion |
|---|---|---|---|
| Nonblacks | | | |
| All | 63.9 | .64 | 1.57 |
| Age 15-19 | 46.0 | .62 | 1.56 |
| Age 20-24 | 103.1 | .61 | 1.56 |
| Age 25-29 | 113.1 | .62 | 1.56 |
| Age 30-34 | 76.9 | .65 | 1.57 |
| Age 35-39 | 29.5 | .67 | 1.58 |
| Age 40-44 | 5.3 | .69 | 1.58 |
| Blacks | | | |
| All | 77.7 | .67 | 1.56 |
| Age 15-19 | 100.4 | .65 | 1.55 |
| Age 20-24 | 142.6 | .64 | 1.56 |
| Age 25-29 | 103.4 | .65 | 1.56 |
| Age 30-34 | 62.2 | .68 | 1.57 |
| Age 35-39 | 26.6 | .71 | 1.57 |
| Age 40-44 | 5.6 | .72 | 1.58 |

Note: The birth rate is the annualized number of births per 1000 thousand women in the relevant population. The Medicaid expanded variable is equal to one if a state has expanded Medicaid eligibility and zero otherwise. The Medicaid threshold variable is the mean eligibility threshold for Medicaid, as a percentage of the federal poverty line, given that a state has expanded Medicaid eligibility. Observations are weighted using the number of women in the population group. Data are state-level quarterly observations over 1983-1996 for a total of 2856 observations per group.

**Table 3. Fraction of women eligible for Medicaid, population size and share, by parity, marital status and race**

| | | No children | | Already have children | |
|---|---|---|---|---|---|
| | All (population, 1000s) | Unmarried (share) | Married (share) | Unmarried (share) | Married (share) |
| All races | | | | | |
| 1983 | .17 | .24 | 0 | .43 | .02 |
| | (55,358) | (.22) | (.13) | (.26) | (.39) |
| 1996 | .46 | .62 | .16 | .69 | .26 |
| | (59,561) | (.24) | (.12) | (.28) | (.36) |
| Nonblacks | | | | | |
| 1983 | .15 | .24 | 0 | .38 | .02 |
| | (48,209) | (.22) | (.14) | (.23) | (.41) |
| 1996 | .43 | .61 | .15 | .65 | .26 |
| | (51,145) | (.24) | (.13) | (.24) | (.39) |
| Blacks | | | | | |
| 1983 | .36 | .29 | 0 | .60 | .02 |
| | (7,149) | (.22) | (.06) | (.49) | (.23) |
| 1996 | .66 | .68 | .24 | .82 | .34 |
| | (8,416) | (.27) | (.05) | (.48) | (.19) |

Note: Shown is the estimated fraction of women (within the given demographic group) who would be eligible for Medicaid coverage if they became pregnant. Estimates are based on data from the March 1984 and March 1997 Current Population Survey (CPS). The fraction eligible is calculated by applying AFDC and Medicaid expansion income eligibility thresholds and household structure rules to CPS data on women's family income and family composition. The population size (in thousands) is given in parentheses in column 1, and the distribution of the population across marital status and parity groups is given in parentheses in columns 2-5 (totals may not add up to 1 because of rounding).

**Table 4. Effect of Medicaid eligibility expansion on birth rates, by age and race**

| | Medicaid expanded Main effect | Medicaid threshold Incremental effect | F-test of joint significance | Mean effect |
|---|---|---|---|---|
| Nonblacks | | | | |
| All | -.09 (.03) | .14 (.02) | 28.23 (.00) | .10 (.04) |
| Age 15-19 | -.09 (.03) | .17 (.02) | 37.09 (.00) | .13 (.04) |
| Age 20-24 | -.07 (.03) | .13 (.02) | 27.73 (.00) | .10 (.04) |
| Age 25-29 | -.11 (.03) | .14 (.02) | 25.36 (.00) | .08 (.04) |
| Age 30-34 | -.09 (.03) | .13 (.02) | 22.49 (.00) | .08 (.04) |
| Age 35-39 | -.07 (.03) | .12 (.02) | 21.89 (.00) | .09 (.04) |
| Age 40-44 | -.08 (.03) | .14 (.02) | 24.47 (.00) | .11 (.05) |
| Blacks | | | | |
| All | -.04 (.03) | .10 (.02) | 15.00 (.00) | .10 (.04) |
| Age 15-19 | -.03 (.03) | .08 (.02) | 7.40 (.00) | .07 (.05) |
| Age 20-24 | -.03 (.03) | .10 (.02) | 12.62 (.00) | .10 (.05) |
| Age 25-29 | -.06 (.03) | .12 (.02) | 15.19 (.00) | .09 (.05) |
| Age 30-34 | -.04 (.03) | .12 (.03) | 14.52 (.00) | .11 (.05) |
| Age 35-39 | -.05 (.04) | .14 (.03) | 18.58 (.00) | .14 (.06) |
| Age 40-44 | -.09 (.05) | .16 (.04) | 12.38 (.00) | .13 (.07) |

Note: The dependent variable is the log of the annualized number of births per 1000 women in the relevant population. The Medicaid expanded main effect is the effect of expanding Medicaid eligibility at all beyond AFDC eligibility. The Medicaid threshold incremental effect is the effect of increasing the eligibility threshold, given that Medicaid has been expanded. The F-test tests whether the two estimated coefficients are jointly different from zero. The mean effect is the Medicaid coefficients evaluated at the sample mean of the Medicaid eligibility threshold variable for that group. A separate regression was estimated for each age/racial group. All regressions were estimated by ordinary least squares. See text for list of other controls. Observations are weighted using the number of women in the population group. Standard errors are in parentheses (p-value for the F-test statistics). The standard errors on the mean effect include uncertainty about the estimated coefficients but not uncertainty about the mean. Data are state-level quarterly observations over 1983-1996 for a total of 2856 observations per group.

**Table 5. Effect of Medicaid eligibility expansion on first birth rates, by marital status, age and race**

| | Unmarried | | | Married | | |
|---|---|---|---|---|---|---|
| | Incremental | F-test | Mean | Incremental | F-test | Mean |
| Nonblacks | | | | | | |
| All | .11 (.02) | 16.31 (.00) | .09 (.04) | .14 (.02) | 28.65 (.00) | .09 (.03) |
| Age 15-19 | .12 (.02) | 16.55 (.00) | .10 (.04) | .20 (.02) | 35.49 (.00) | .11 (.04) |
| Age 20-24 | .12 (.02) | 16.17 (.00) | .08 (.04) | .12 (.02) | 22.39 (.00) | .07 (.03) |
| Age 25-29 | .13 (.02) | 19.86 (.00) | .09 (.04) | .12 (.02) | 20.06 (.00) | .06 (.04) |
| Age 30-34 | .13 (.02) | 14.94 (.00) | .08 (.05) | .12 (.02) | 20.95 (.00) | .07 (.04) |
| Age 35-39 | .05 (.03) | 1.91 (.15) | .03 (.05) | .11 (.02) | 17.91 (.00) | .07 (.03) |
| Age 40-44 | .21 (.04) | 12.78 (.00) | .18 (.08) | .02 (.02) | .76 (.47) | .03 (.05) |
| Blacks | | | | | | |
| All | .09 (.02) | 11.14 (.00) | .06 (.04) | .14 (.02) | 21.77 (.00) | .06 (.04) |
| Age 15-19 | .08 (.02) | 7.14 (.00) | .05 (.04) | .21 (.04) | 13.69 (.00) | .11 (.07) |
| Age 20-24 | .10 (.02) | 9.55 (.00) | .07 (.04) | .13 (.02) | 14.47 (.00) | .07 (.04) |
| Age 25-29 | .10 (.03) | 7.90 (.00) | .05 (.05) | .12 (.02) | 17.15 (.00) | .04 (.04) |
| Age 30-34 | .05 (.03) | 2.58 (.08) | .01 (.05) | .11 (.03) | 11.60 (.00) | .02 (.05) |
| Age 35-39 | .06 (.04) | 1.52 (.22) | -.01 (.08) | .09 (.04) | 4.90 (.01) | .01 (.07) |
| Age 40-44 | -.02 (.06) | 1.08 (.34) | -.08 (.11) | .09 (.06) | 3.71 (.02) | -.07 (.12) |

Note: The dependent variable is the log of the annualized number of births per 1000 women in the relevant age/racial group. The incremental effect measures the effect of further raising the eligibility threshold, given that eligibility has already been expanded. The F-test tests whether the estimated coefficients for the main effect and the incremental effect are jointly different from zero. The mean effect is the Medicaid coefficients evaluated at the mean of the Medicaid eligibility threshold variable for that group. A separate regression was estimated for each group. All regressions were estimated by ordinary least squares. See text for list of other controls. Observations are weighted using the number of women in the age/racial group. Standard errors are in parentheses (p-value for the F-test statistics). The standard errors on the mean effect include uncertainty about the estimated coefficients but not uncertainty about the mean. Data are state-level quarterly observations over 1983-1996 for a total of 2856 observations per group.

**Table 6. Effect of Medicaid eligibility expansion on higher-order birth rates, by marital status, age and race**

| | Unmarried | | | Married | | |
|---|---|---|---|---|---|---|
| | Incremental | F-test | Mean | Incremental | F-test | Mean |
| Nonblacks | | | | | | |
| All | .06 (.03) | 30.41 (.00) | .23 (.06) | .13 (.02) | 28.81 (.00) | .10 (.03) |
| Age 15-19 | .09 (.03) | 29.82 (.00) | .25 (.06) | .17 (.06) | 31.41 (.00) | .13 (.04) |
| Age 20-24 | .06 (.02) | 29.16 (.00) | .21 (.06) | .12 (.02) | 25.55 (.00) | .09 (.03) |
| Age 25-29 | .05 (.03) | 26.10 (.00) | .22 (.06) | .14 (.02) | 27.98 (.00) | .08 (.04) |
| Age 30-34 | .03 (.03) | 26.05 (.00) | .23 (.07) | .12 (.02) | 23.63 (.00) | .08 (.03) |
| Age 35-39 | .07 (.02) | 15.05 (.00) | .14 (.05) | .12 (.02) | 21.25 (.00) | .10 (.03) |
| Age 40-44 | .10 (.04) | 5.78 (.00) | .12 (.08) | .13 (.02) | 23.23 (.00) | .12 (.04) |
| Blacks | | | | | | |
| All | .03 (.03) | 30.50 (.00) | .23 (.06) | .14 (.02) | 26.83 (.00) | .15 (.04) |
| Age 15-19 | .03 (.03) | 25.55 (.00) | .21 (.07) | .20 (.04) | 13.14 (.00) | .17 (.07) |
| Age 20-24 | .03 (.03) | 24.68 (.00) | .23 (.07) | .14 (.02) | 19.06 (.00) | .13 (.04) |
| Age 25-29 | .02 (.03) | 24.24 (.00) | .22 (.07) | .15 (.02) | 23.90 (.00) | .12 (.04) |
| Age 30-34 | .01 (.03) | 22.69 (.00) | .20 (.07) | .14 (.02) | 18.11 (.00) | .13 (.04) |
| Age 35-39 | .08 (.03) | 22.45 (.00) | .22 (.07) | .14 (.03) | 17.83 (.00) | .15 (.05) |
| Age 40-44 | .18 (.05) | 8.34 (.00) | .19 (.12) | .16 (.04) | 8.09 (.00) | .14 (.07) |

Note: The dependent variable is the log of the annualized number of births per 1000 women in the relevant age/racial group. The incremental effect measures the effect of further raising the eligibility threshold, given that eligibility has already been expanded. The F-test tests whether the estimated coefficients for the main effect and the incremental effect are jointly different from zero. The mean effect is the Medicaid coefficients evaluated at the mean of the Medicaid eligibility threshold variable for that group. A separate regression was estimated for each group. All regressions were estimated by ordinary least squares. See text for list of other controls. Observations are weighted using the number of women in the age/racial group. Standard errors are in parentheses (p-value for the F-test statistics). Data are state-level quarterly observations over 1983-1996 for a total of 2856 observations per group.

**Figure 1**
**Medicaid Eligibility Threshold and Eligibility Rate**