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The effect of Medicaid eligibility expansions on fertility

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ABSTRACT

In the United States, pregnant women and children's eligibility for Medicaid was expanded dramatically during the 1980s and early 1990s. By lowering pregnancy and child health care costs, the Medicaid expansions may have increased the incentives for women to have children. To investigate this possibility, we examine whether state-level birth and abortion rates are related to the extent of states' Medicaid eligibility expansions and the fraction of women eligible for Medicaid, controlling for economic and demographic factors, during the period 1982 to 1996. We examine birth rates by race, marital status and education as well as overall abortion rates. We find little evidence that the Medicaid expansions led to changes in birth rates or abortion rates. However, some results do suggest that the Medicaid expansions boosted the birth rate among white women who have not completed high school. We find that restrictions on Medicaid funding of abortions decrease abortion rates and increase birth rates. The results thus do not provide definitive evidence that expansions in public health insurance eligibility have sizable effects on women's fertility.

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 pregnant women's use of prenatal care and children's access to medical care. The proportion of women of childbearing age eligible for Medicaid coverage of pregnancy-related services more than doubled between 1987 and 1992, and the proportion of children eligible for full Medicaid coverage rose by at least 50% (Card & Shore-Sheppard, 2004; Cutler & Gruber, 1996). Several studies have found that the expansions succeeded at increasing both prenatal care and children's medical care, resulting in better birth outcomes and lower child mortality (Currie & Grogger, 2002; Currie & Gruber, 1996a, 1996b). This article examines whether the Medicaid eligibility expansions had another, likely unintended effect: changing women's fertility.

Understanding the effect of public health insurance on fertility is important for several reasons. Medicaid covers a sizable proportion of the population and involves substantial outlays. In 2007, for example, over 13% of the total population was covered by Medicaid, including over one in four children (U.S. Census Bureau, 2008). Even more of the population is eligible but not currently enrolled in Medicaid. Total program outlays for fiscal year 2007 were over $190 billion. Medicaid has covered over one-third of births in years since the 1990s expansions were completed (National Governors' Association, 2005). The effect of the Medicaid expansions on fertility has implications for welfare reform, which further decoupled eligibility for government-funded health insurance from receipt of cash benefits. In addition, the expansion of government-financed health insurance for children through the State Children's Health Insurance Program (SCHIP) may have affected some women's fertility by lowering the cost of health insurance and care for current and future children. Finally, the 2010 health insurance reform has the potential to increase public (or publicly subsidized) health insurance coverage, making the question investigated here timely.

Economic theory provides a basis for expecting the Medicaid expansions to have increased birth rates. First, some women experienced a drop in the costs of prenatal care, delivery and child health care. This reduction in health care costs lowers the total cost of a child, which standard economic theory suggests should increase the number of children born. In addition, a reduction in health care costs effectively raises income (net of such costs), which in turn may increase births; most studies have concluded that increases in income that are not due to increases in women's earnings have a positive, albeit small, effect on fertility (Hotz, Klerman, & Willis, 1997; Macunovich, 1996). For similar reasons, the expansions may have reduced abortion rates if more women opted to give birth

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instead of terminating a pregnancy. Alternatively, the expansions could have led to no change in abortion rates if all extra births came from an expansion-induced increase in pregnancies.

It is also theoretically possible that the Medicaid expansions had a negative effect on birth rates. Research has found that, by decoupling eligibility for Medicaid from receipt of cash welfare benefits under the Aid to Families with Dependent Children (AFDC) program, the expansions may have caused an increase in women's labor force participation, although there is debate about this effect (Ham & Shore-Sheppard, 2003; Yelowitz, 1995). Higher labor force participation might have led to lower birth rates and, possibly, higher abortion rates. If the expansions improved child health outcomes, parents might have opted to have fewer births in the classic "quality versus quantity" tradeoff (Becker, 1960). This effect could also have raised abortion rates (Joyce & Kaestner, 1996). In addition, abortion rates might have risen because Medicaid covers abortion in some states, and the expansions reduced the cost of abortion for some women in those states. However, a decline in the cost of abortion also could have increased sexual activity and the number of pregnancies, potentially raising both the birth rate and the abortion rate (Kane & Staiger, 1996). The expansions also may have increased some women's access to family planning, reducing both births and abortions. Finally, it is possible that these changes had no impact on birth or abortion rates if women potentially eligible for Medicaid-funded births or abortions were unaware of or did not respond to the law changes. These various theoretical arguments thus suggest we might see an increase, a decrease or no net effect on birth rates, and similarly for abortion rates, when we turn to the data.

Most previous research suggested that expansions in Medicaid coverage led to an increase in births and drop in abortions. Studies using samples from natality data or the Current Population Survey found increases in birth rates of 3%–5% as a result of the Medicaid expansions that occurred during the late 1980s and 1990s (Baughman, 2001; Joyce, Kaestner, & Kwan, 1998; Yelowitz, 1994). Medicaid coverage of abortions—which the expansions included in some states—was found to be negatively associated with birth rates, particularly among black, unmarried and less-educated women (Klerman, 1999). Studies that used abortion data from selected states found that the Medicaid expansions reduced abortions among unmarried nonblacks but had no effect among unmarried blacks (Joyce & Kaestner, 1996; Joyce et al., 1998). More generally, restrictions on Medicaid funding of abortions were found to be negatively associated with abortion rates (e.g., Blank, George, & London, 1996; Haas-Wilson, 1993).

However, a recent paper that used methods similar to those here concluded that the expansions did not have a robust, discernable effect on fertility for most groups (DeLeire, Lopoo & Simon, in press). Our findings are generally compatible with that study and thus counter to much of the previous literature; we find little evidence of a significant effect of the expansions on birth rates during a longer time period that covers more phases of the expansions than DeLeire et al.'s analysis.

This article adds to the existing research on Medicaid eligibility and fertility in several ways. First, we estimate whether birth and abortion rates are related to the fraction of women eligible for Medicaid if they became pregnant and to the expansion-related income eligibility threshold for Medicaid. Our measure of the fraction of women eligible uses a national sample of women but each state's eligibility rules to avoid possible endogeneity of state-level eligibility, as advocated by Currie and Gruber (1996a). As discussed below, this approach offers several advantages over previous published research that used other measures of the extent of the expansions. Another contribution of our study is that we control for the impact of several other important factors, such as economic conditions, welfare generosity and whether a state restricted Medicaid funding of abortions.

A further contribution is that this study includes more years and states than most previous research and analyzes the full effect of the expansions. We investigate the time period 1982–1996, which includes several pre-expansion years and the more limited expansions that took place during the mid-1980s as well as the later broader expansions. Most previous studies of the fertility effects of the expansion focused on the broadly targeted phases and did not include any pre-expansion years, thus capturing the impact of moving from the narrow phase of the expansions to the broad phase but not the full effect of the expansions. The early phases of the expansion had considerably larger effects on children's Medicaid coverage than did the later phases (Card & Shore-Sheppard, 2004) and may have also had a large impact on fertility. In addition, whereas some previous research that used natality and abortion data to study the impact of the expansions examined only selected groups of states, this study includes all states (with the exception of 4 states with very incomplete data on maternal education in the education portion of the analysis).

Chronology of Medicaid expansions

Until the mid-1980s, Medicaid coverage was tightly linked to receipt of AFDC benefits, which effectively limited eligibility in most states to female-headed households with children and with income low enough to qualify for AFDC. Beginning in 1984, a series of laws expanded pregnant women and children's eligibility for Medicaid coverage. Table 1 summarizes the eligibility expansions.

During the early, narrowly targeted 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. States also had the option of using their own funds to extend coverage beyond groups eligible under the federal law. During this first phase of the expansion, Medicaid recipients still had to meet the AFDC resource requirements, which required income to be well below the poverty level in most states. The overall fraction of women aged 15–44 eligible for Medicaid rose less than 5 percentage points during the narrowly targeted phase of the expansion (Currie & Gruber, 1996a), although increases were larger among some subgroups. During this period, additional changes included the extension of Medicaid eligibility to older

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.

children in families with incomes below the AFDC cutoff, regardless of family structure ("Ribicoff children"), 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). We do not examine those expansions here.

Beginning in 1987, states were given the option and then later required to extend Medicaid coverage to pregnant women and children in families with incomes substantially above the AFDC resource limits in most states, regardless of family structure. Eligibility rates among women aged 15–44 more than doubled during this broad phase of the expansion (Currie & Gruber, 1996a).

States differed in the timing and extent of Medicaid eligibility expansions. For example, an early part of the expansions required 27 states to offer Medicaid coverage to pregnant women who had a spouse present in the household and who met their state's AFDC resource requirements; the other 23 states and the District of Columbia already had such programs in place. There also was variation across states in the timing of the broad eligibility expansions; as of February 1996, for example, 28 states had opted to cover pregnant women with incomes of 185% of the federal poverty level or higher, while the remaining states met the minimum level of 133% but were below 185%. This study uses this variation across states in the timing and extent of expansions to estimate the effect of Medicaid eligibility threshold expansions on fertility.

Fig. 1 shows the average income threshold for Medicaid eligibility relative to the federal poverty level during 1982–1996. The figure displays both the expansion-related threshold (after the expansions began in late 1984) and the eligibility threshold for Medicaid coverage via the AFDC program; both are shown relative to the federal poverty level for a family of three. The average AFDC-related threshold declined slightly during the 1990s whereas the average expansion-related threshold rose steadily during the mid-1980s and then plateaued during the 1990s at over three times the AFDC income limit.

The figure also shows the estimated fraction of women eligible for expansion-related Medicaid benefits if they became pregnant. The estimates are constructed from the 1983–1997 March Current Population Surveys based on family income and family structure. Women who were eligible via participation in the AFDC program were not included in these estimates. Separate fractions are shown for white and black women to reflect the large differences in eligibility rates, with blacks much more likely than whites to become eligible for expansion-related Medicaid coverage. The figure also shows that the estimated expansion-related eligibility rates track the rise in the Medicaid expansion threshold.

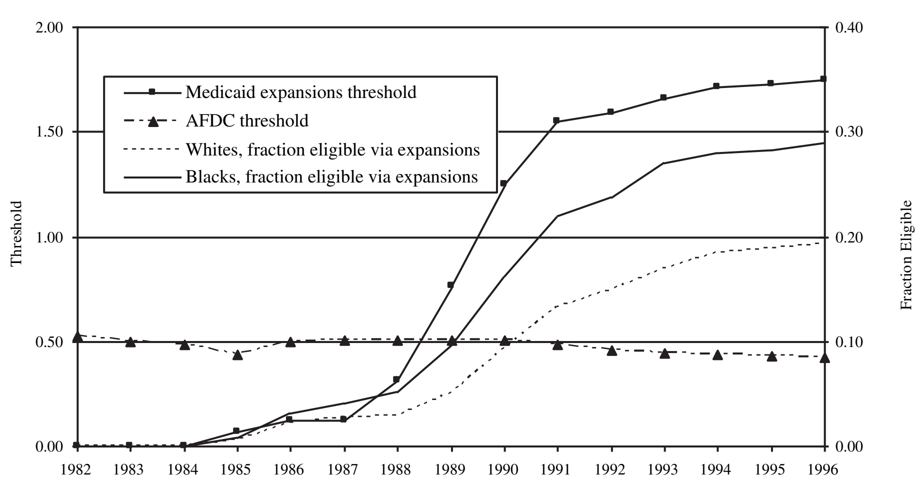
Pregnant women are covered under the expansions only for services related to the pregnancy. In some states, pregnant women's coverage includes abortion services. States have the option to cover abortions in their Medicaid program but do not receive federal matching funds for most abortions. In almost all states with Medicaid programs that cover abortions, women eligible for Medicaid under the expansions are also eligible for Medicaid-covered abortions (Sollom, 1995).

Data

Birth rates

This analysis uses data from national birth certificate data sets created by the National Center for Health Statistics (NCHS). The sample consists of all live births between 1982 and 1996 to women aged 15–44 residing in the 50 states and the District of Columbia whose age, race and marital status were reported; for a small set of states and years, the data are a half-sample of births, so we multiplied those counts by 2. The sample ends in 1996 to minimize confounding the effects of the Medicaid expansions and welfare reform and other policy changes which might impact fertility and abortions, such as the creation of the State Children's Health Insurance Program. In analysis using data on maternal education, 4 states with very incomplete reporting of education during several years are dropped from the sample. We use quarterly birth rates created by dividing the number of births to women in a given quarter by the population (in thousands).

This analysis distinguishes between whites and blacks; nonwhites/nonblacks ("other race") are not included here because of the relatively small sample sizes in most states. We do not distinguish between Hispanics and non-Hispanics because Hispanic origin is not reported for a substantial fraction of births in early years of the sample. The data are stratified by race because previous research indicated that the determinants of fertility



Note: The Medicaid expansions threshold is the average eligibility threshold across states, relative to the federal poverty level, weighted using the population of women aged 15-44, associated with the eligibility expansions that began in 1984. The AFDC threshold is the average eligibility threshold for obtaining AFDC benefits, which include Medicaid coverage. The fraction of women eligible is the fraction of women aged 16-44 from the 1983-1997 March Current Population Survey who would meet their state's income eligibility rules for the Medicaid expansion (and does not include Medicaid coverage under the AFDC program).

Fig. 1. Medicaid eligibility thresholds and fraction of women eligible via expansions, 1982–1996.

behavior differ for whites and blacks (Levine, Staiger, Kane, & Zimmerman, 1999). In particular, the effect of the expansions may differ because blacks are poorer, on average, and therefore more likely to be below the income threshold (as shown in Fig. 1). However, blacks are more likely to already be eligible for Medicaid before the expansions via the AFDC program. White women's fertility may be more affected by the expansions because they are less likely than blacks to qualify for Medicaid via the AFDC program.

We examine the relationship between Medicaid expansions and fertility by marital status or education within racial groups. The responsiveness of women's fertility to Medicaid eligibility may differ by marital status and education because less educated and single women tend to have fewer financial resources than more educated and married women and therefore are more likely to be eligible for Medicaid via the expansions. The expansions therefore may have had larger effects on less educated and unmarried women's fertility.

We focus here on birth and abortion rates relative to the number of women aged 15–44 in a given group. We use the Census Bureau's annual state-level counts of the number of women aged 15–44 by race. To calculate denominators by marital status or education, we multiply those counts by the distribution of women aged 15–44 across marital status and education groups in that state and year from the March CPS. The birth rate among unmarried white women aged 15–44, for example, is the number of births to unmarried white women aged 15–44 in the natality data divided by the number of unmarried white women aged 15–44, calculated by multiplying the Census Bureau's count of the number of white women aged 15–44 times the fraction of white women aged 15–44 who are married in the CPS.

A primary advantage of using the natality data is that they are a near census of births in the U.S. and thus should not suffer from significant underreporting of births. The 1990 Census, for example, may have missed up to 20% of infants compared with the natality data (Daponte & Haviland, 2001). However, the natality data have one significant disadvantage: because the data only include births, they contain no information on women who are pregnant but do not give birth. This motivates our analysis of abortion rates.

Abortion rates

This analysis uses two sources of annual data on the number of abortions: the Centers for Disease Control (CDC) and the Alan Guttmacher Institute (AGI). The CDC data are based on reports from state public health agencies, but the data are incomplete in states in which not all providers provide reports to the public health agency. The AGI data are based on surveys of all known abortion providers and consistently include more abortions than the CDC data, although the ratio of the two counts varies over time. The AGI data are not available for 1983, 1986, 1989, 1990, 1993 and 1994, so we present estimates using the CDC data for the nine years during the period 1982–1996 when AGI data are available as well as for the period as a whole.

We do not examine the effect of the expansions on abortion rates by race, education or marital status. The CDC reports abortion counts for some demographic groups, such as by race, but the data are only available for a subsample of states and years and their accuracy is uncertain. We use abortion rates calculated by dividing the reported number of abortions in each state and year by the population of women aged 15–44.

Analytic approach

We use OLS panel data techniques to estimate the relationship between Medicaid eligibility and fertility. We regress the natural logarithm of the birth rate or the abortion rate on the expansion-related fraction of women eligible for Medicaid or on the expansion-related income threshold for Medicaid eligibility, expressed as a fraction of the poverty line. All regressions also include economic and demographic variables (described below), state and time fixed effects, and state-specific linear time trends. These other variables help control for other factors that affect fertility besides the Medicaid expansions and are commonly used in studies of fertility determinants. Stratifying on race, marital status, and education also helps control for other factors, such as differences in fertility by women's education. In the discussion below, we focus on several aspects of the estimated coefficients on the Medicaid eligibility measure: what is their magnitude, whether they are significantly different from zero, and whether they are significantly different across groups.

The regressions include the unemployment rate and the natural log of real income per capita to control for economic conditions that may affect fertility. The natural log of the real maximum AFDC benefit for a family of three is included to control for any effect of welfare benefits on fertility. The regressions also include the natural log of the real combined maximum value of federal and state earned income tax credits (EITC) for a taxpayer with 1 qualifying child. (The online Appendix provides further details about these variables.)

Because Medicaid funding of abortions may affect abortion and birth rates (e.g., Klerman, 1996, 1999), the regressions also include a dummy variable that indicates whether Medicaid funding of abortions is restricted. In the birth rate regressions, the abortion-funding variable is measured one quarter after conception to reflect Medicaid financing of abortions toward the end of the first trimester of a pregnancy; in the abortion rate regressions, the variable is the contemporaneous annual average.

The regressions also control for the distribution of women across 5-year age groups. Observations were weighted using the population size. The standard errors are Huber-White corrected for heteroscedasticity and clustered at the state level to allow for arbitrary correlation within state. There are a maximum of 3060 observations in the birth rate regressions (2820 in the ones stratified by education since 4 states are not included) and 765 observations in the full sample CDC abortion rate regressions (459 in the AGI and restricted sample CDC ones). The sample sizes for the black birth rate regressions are somewhat smaller, as some state-year cells have no women in some groups.

Most previous research on the fertility effects of the expansions used indicator variables to measure the extent of the expansions (e.g., Joyce & Kaestner, 1996; Joyce et al., 1998). This study, in contrast, uses continuous variables: the income eligibility threshold and the fraction of women eligible. We believe these continuous variables are a better measure of the extent of the expansions than dummy variables indicating whether a state has expanded eligibility to, for example, 100 or 133% of the federal poverty level, for example. We use two measures because each of our variables is imperfect. The income eligibility threshold does not reflect state heterogeneity in actual income levels and hence may not do a good job of capturing cross-state differences in eligibility. The fraction eligible better captures cross-state differences in eligibility but is measured with error. It is constructed by applying each state's eligibility rules to a national sample of women, as pioneered by Currie and Gruber (1996a) (see the online Appendix for details). We cannot include a measure of actual state-level eligibility. The measurement error in such a measure constructed from the CPS would be tremendous, in part because of small sample sizes. More importantly, such a measure would be endogenous since most women's eligibility under the expansions depends on their becoming pregnant (Currie & Gruber, 1996a).

Results

Birth rate results

The results indicate that the Medicaid expansions had little effect on birth rates among women aged 15–44 overall. The estimate in the first column in Table 2 indicates that a 100 percentage point increase in the expansion-related eligibility threshold (measured as a proportion of the federal poverty guideline) is associated with a .9% increase in the birth rate among white women aged 15–44, and the estimate is significantly different from 0. Evaluated at the mean expansion-related eligibility threshold (conditional on expansion) of 158%, this estimate implies that the expansions led to a 1.4% increase (1.58 × .9%) in the birth rate among white women overall. Given that the average number of births to white women was about 3.1 million per year during our sample period, this would correspond to an extra 43,400 births per year (3.1 million × 1.4%). The estimated coefficient for black women is .3% but is not statistically different from 0 at the 5% level. The coefficients are estimated imprecisely enough that we cannot reject the possibility that the effect (in percentage terms) was the same among whites and blacks, based on F-tests from regressions using pooled data. This is likely in part due to power issues.

If we instead measure the extent of the expansions using the fraction of women eligible for expansion-related Medicaid, we do not find a significant relationship between birth rates and the expansions for either whites or blacks. The estimated effect among whites of .079 shown in Table 3, column 1 should be interpreted as a 10 percentage point increase in the fraction of women eligible—about the mean effect in the sample—leading to a .79% increase in the birth rate. However, the 95% confidence interval ranges from almost +20% to −.04%. The estimate for blacks is similarly imprecise, and we again cannot reject the possibility that the effect was the same across whites and blacks.

When we examine birth rates by marital status, there is little evidence of a significant effect of the expansions on fertility. As column 2 of Table 2 shows, a 100 percentage point increase in the expansion-related eligibility threshold is associated with a 2.7% increase in the birth rate among unmarried white women and a 3.3% increase for unmarried black women, but neither estimate is significantly different from 0 at the 5% level. The estimates also do not indicate a significant difference between white and black unmarried women. The results for the estimated fraction of women eligible for expansion-related Medicaid also indicate no statistically significant effects among unmarried women (Table 3, column 2). For married women, the eligibility threshold is not significantly related to birth rates for either racial group or either specification for the policy change (Table 2, column 3 and Table 3, column 3). We cannot reject the possibility that the expansions had the same effect on unmarried and married white women. However, the results using the eligibility threshold (Table 2) indicate that the expansions did have a significantly larger effect on unmarried black women than on their married counterparts. In addition, the estimated effect of the eligibility threshold is significantly larger among white married women than among black married women, although this difference is driven in part by the negative point estimate for black women.

The effect of the Medicaid eligibility expansions appears to differ somewhat across education groups. The point estimates for women who have not completed high school indicate a sizable positive relationship, although relatively imprecisely estimated; only the result for white women in the eligibility threshold regression is significantly different from 0. That result, shown in column 4 of Table 2, indicates that a 100 percentage point increase in the eligibility threshold would result in a 7.7% increase in the birth rate to white women who do not have 12 years of education. That estimate is significantly larger than the results for white women who have exactly 12 or who have more than 12 years of education (columns 5 and 6 of Table 2). The result for black women who do not have 12 years of education also is not significantly different from the one for black women with either 12 or more than 12 years of education, and the corresponding coefficients in the share eligible specification are not significantly different from 0 or from the corresponding results for more-educated women (Table 3). We do not find any evidence of sizable or statistically significant effects among women with 12 or more years of school in either specification; all of those coefficients are negative but not significantly different from 0 or from each other.

The pattern that emerges from these results is that increases in the eligibility threshold may raise births among white women without a high school diploma, and the effects are sometimes statistically larger among unmarried and low-education women than among their married and more-educated counterparts. The regressions using the fraction eligible measure yield a similar pattern, but none of the differences achieve statistical significance. This is consistent with the fraction eligible measure having more measurement error or being noisier and hence having less power

Table 2
Determinants of birth rates, including Medicaid expansions eligibility threshold, by demographic group, 1982–1996.

	All	Unmarried	Married	<12 years education	12 years education	>12 years education
Whites						
Expanded Medicaid eligibility threshold	.009* (.004)	.027 (.024)	.009 (.010)	.077* (.035)	−.016 (.019)	−.004 (.014)
Ln (real AFDC benefits)	.108 (.056)	.349 (.180)	.047 (.036)	.211 (.127)	.127 (.090)	−.070 (.150)
Ln (real per capita income)	.449** (.093)	.721 (.394)	.363* (.142)	.322 (.345)	.462* (.181)	.409** (.164)
Unemployment rate	.001 (.001)	.009 (.006)	.001 (.002)	−.011 (.007)	.005 (.004)	.001 (.004)
Ln (real EITC)	−.022 (.040)	−.025 (.146)	−.091* (.038)	.123 (.109)	.004 (.059)	.038 (.070)
Medicaid abortion funding restriction	.007 (.011)	.080 (.055)	−.028 (.016)	.032 (.049)	.006 (.027)	.028 (.028)
Blacks						
Expanded Medicaid eligibility threshold	.003 (.011)	.033 (.018)	−.040 (.022)	.059 (.036)	−.027 (.038)	−.040 (.050)
Ln (real AFDC benefits)	.102** (.036)	−.009 (.085)	.339** (.100)	.032 (.234)	.031 (.242)	.162 (.216)
Ln (real per capita income)	.231 (.129)	.451* (.224)	−.102 (.442)	.505 (.587)	.495 (.439)	−.226 (.458)
Unemployment rate	.003 (.003)	.004 (.006)	.005 (.009)	.004 (.014)	.0001 (.008)	.013 (.010)
Ln (real EITC)	.097 (.119)	.055 (.136)	.110 (.181)	.127 (.253)	−.443** (.156)	.295 (.211)
Medicaid abortion funding restriction	.077** (.025)	.137** (.031)	−.056 (.053)	.162 (.100)	.099** (.037)	−.019 (.055)

*$p < .05$; **$p < .01$.

Note: The dependent variable is the log of the annualized number of births per 1000 women in the relevant group each quarter during 1982–1996. Robust standard errors clustered at the state level are in parentheses. Regressions include state and time fixed effects and linear state-specific time trends.

Table 3
Determinants of birth rates, including fraction of women eligible for Medicaid because of expansions, by demographic group, 1982–1996.

	All	Unmarried	Married	<12 years education	12 years education	>12 years education
Whites						
Fraction of women eligible	.079 (.060)	.059 (.336)	.141 (.151)	.313 (.258)	−.138 (.197)	−.062 (.300)
Ln (real AFDC benefits)	.109 (.055)	.350 (.177)	.049 (.035)	.227 (.129)	.123 (.089)	−.071 (.151)
Ln (real per capita income)	.454** (.090)	.752 (.396)	.364* (.141)	.365 (.341)	.436* (.182)	.402* (.168)
Unemployment rate	.001 (.001)	.010 (.006)	.001 (.002)	−.011 (.007)	.005 (.004)	.001 (.004)
Ln (real EITC)	−.016 (.040)	−.009 (.143)	−.085* (.038)	.169 (.109)	−.002 (.057)	.037 (.068)
Medicaid abortion funding restriction	.009 (.011)	.083 (.057)	−.027 (.015)	.030 (.049)	.004 (.026)	.027 (.029)
Blacks						
Fraction of women eligible	−.010 (.101)	.063 (.123)	−.269 (.230)	.060 (.318)	−.277 (.226)	−.149 (.490)
Ln (real AFDC benefits)	.102** (.037)	−.005 (.084)	.331** (.099)	.052 (.236)	.023 (.240)	.149 (.226)
Ln (real per capita income)	.238 (.130)	.492* (.238)	−.129 (.467)	.564 (.602)	.462 (.429)	−.258 (.471)
Unemployment rate	.003 (.003)	.005 (.006)	.005 (.009)	.005 (.014)	.0002 (.008)	.012 (.010)
Ln (real EITC)	.099* (.120)	.078 (.129)	.077 (.191)	.169 (.254)	−.448** (.146)	.269 (.199)
Medicaid abortion funding restriction	.077** (.026)	.141** (.035)	−.065 (.052)	.170 (.099)	.095* (.037)	−.022 (.054)

*$p < .05$; **$p < .01$.
Note: The dependent variable is the log of the annualized number of births per 1000 women in the relevant group each quarter during 1982–1996. Robust standard errors clustered at the state level are in parentheses. Regressions include state and time fixed effects and linear state-specific time trends.

than the eligibility threshold measure. Unfortunately, we do not have a good instrument for the fraction eligible measure.

There are several significant relationships between birth rates and the other covariates reported in Tables 2 and 3. In general, birth rates are positively associated with the log of real AFDC benefits for some groups of blacks. Birth rates among whites tend to be positively associated with the log of real income per capita, as are birth rates for unmarried blacks. There is generally no significant relationship between birth rates and the unemployment rate or the log of real EITC benefits. Confirming some existing research, the results indicate a positive relationship between birth rates and state restrictions on Medicaid funding for abortions for some groups, particularly unmarried black women and black women with high school degrees who have not attended college.

Abortion rate results

The Medicaid eligibility expansions do not appear to have affected overall abortion rates. As the results in Table 4 show, abortion rates are not significantly associated with either the eligibility threshold (top panel) or the fraction of women eligible for Medicaid as a result of the expansions (bottom panel). This result holds for the CDC data, the AGI data, and the CDC data restricted to years with AGI data available.

Some other factors do appear to influence abortion rates. In the CDC data, abortion rates are generally positively associated with the log of real average income; these results do not hold in the AGI data, however. Both data sources indicate that abortion rates are significantly lower when states restrict Medicaid funding for abortions. The coefficient estimates are large, but confidence intervals include smaller estimates.

Discussion

Beginning in 1984, the share of pregnant women and children eligible for Medicaid increased dramatically. Eligible pregnant women were covered for prenatal care, delivery and postpartum care, while eligible children received full Medicaid benefits. The expansions lowered the cost of health care for families 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 study examined whether this reduction in health care costs led to changes in fertility behavior.

Like other recent work (DeLeire et al., in press), our results do not provide strong evidence that the Medicaid expansions increased birth rates. We also do not find convincing evidence of an effect on abortion rates. Some results do point to a positive fertility effect among white women who have not completed high school. However, the births rates for all other groups of black and white women are not significantly related to the extent of the expansions, regardless of whether we measure the extent of the expansions using the eligibility threshold or the share of women made eligible via the expansions. Although we cannot rule out small positive effects, our estimates do not encompass overall effects in the 3–5% range suggested by some previous research and certainly are much smaller than the 29% effect found by Leibowitz (1990); she examined a program that

Table 4
Determinants of abortion rates, by data source, 1982–1996.

	CDC data, all years	AGI data	CDC data, AGI years
A. Expansion threshold			
Expanded Medicaid eligibility threshold	.012	.026	.040
	(.040)	(.022)	(.050)
Ln(real AFDC benefits)	.180	−.068	.151
	(.145)	(.152)	(.170)
Ln(real per capita income)	.524*	.634	1.330**
	(.247)	(.423)	(.357)
Unemployment rate	−.011	.008	−.001
	(.009)	(.007)	(.007)
Ln(real EITC)	.019	−.103	.075
	(.120)	(.087)	(.141)
Medicaid abortion funding restriction	−.087*	−.069*	−.119
	(.043)	(.031)	(.064)
B. Fraction of women eligible			
Fraction of women eligible for Medicaid	.055	.187	−.100
	(.255)	(.358)	(.391)
Ln (real AFDC benefits)	.185	−.056	.157
	(.141)	(.149)	(.166)
Ln (real per capita income)	.530*	.643	1.354**
	(.242)	(.431)	(.346)
Unemployment rate	−.011	.008	.0004
	(.009)	(.007)	(.008)
Ln (real EITC)	.027	−.080	.090
	(.128)	(.098)	(.151)
Medicaid abortion funding restriction	−.086	−.067*	−.119
	(.044)	(.031)	(.063)

*$p < .05$; **$p < .01$.
Note: The dependent variable is the log of the annual number of abortions per 1000 women aged 15–44. In column 1, the CDC sample covers all years 1982–1996; in columns 2 and 3, the AGI and CDC samples cover the years 1982–1996 except 1983, 1986, 1989, 1990, 1993 and 1994. Robust standard errors clustered at the state level are in parentheses. Regressions include state and time fixed effects and linear state-specific time trends.

provided free medical care to families for a 3–5 year period and thus likely had large effects on timing but not necessarily on total fertility. Like existing research, we find that Medicaid restrictions on the funding of abortions are related to higher birth rates and lower abortion rates for some groups.

Our failure to find a substantial effect of the Medicaid expansions on fertility accords with much of the literature on the effect of public assistance programs on fertility. Studies of cash welfare, work requirements for welfare recipients, time limits, family caps, and other aspects of welfare and welfare reform tend to find small effects on fertility, although a few studies do find sizable effects (see, for example, the summaries in Grogger & Karoly, 2005, and Moffitt, 1998). These findings do not necessarily indicate that fertility does not respond to public assistance programs. Rather, they indicate that existing programs and changes made thus far to those programs have had little measurable effect. Policymakers could reasonably anticipate that future changes of similar magnitude are unlikely to lead to sizable changes in fertility, but not that fertility is always completely unaffected by the design of public assistance programs. In other words, an expansion of public health insurance similar in magnitude to the Medicaid expansions—which was a fairly large-scale expansion—seems unlikely to have a major effect on fertility, but more radical changes might.

There are two important caveats to our results. First, the Medicaid expansions may not have increased the total number of births among affected women but rather simply shifted some births to younger ages among those women. Because the data used here only cover the period 1982 to 1996, the observed increase in birth rates for some groups could be due to changes in timing but not a long-run increase in the total number of births. These questions cannot be fully examined using the natality data, which only report the interval since the last live birth and last did so in 1994, nor in the Current Population Survey, which last asked about complete fertility histories in 1995. Understanding whether access to public or subsidized health insurance affects the total number of births or simply changes the timing of births is once again a relevant issue given the 2010 health insurance reform. A data source with a detailed fertility history will be needed to completely understand that reform's effect on fertility.

The second caveat is that the natality data we used here do not include women's health insurance status or income. These variables are almost surely endogenous with respect to fertility, but they are also key determinants of fertility. Ideally, researchers would have data not only on women's fertility histories but on their health insurance status and income and plausible instruments for those variables.

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Appendix. Supplementary data

Supplementary data associated with this article can be found, in the online version, at doi:10.1016/j.socscimed.2010.05.046

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