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Abstract

Infertility currently affects over 6 million individuals in the United States. While most health insurance plans nationwide do not cover infertility diagnoses or treatments, to date fifteen states have enacted some form of infertility insurance mandate. In this paper, I use data from the Vital Statistics Detail Natality Data and Census population estimates to examine whether these state-level mandates were successful in increasing fertility rates. Using a difference-in-differences approach, I exploit variation in the enactment of mandates both across states and over time, and identify treatment and control groups that should have been differentially affected by infertility coverage. My results suggest that the mandates significantly increase first birth rates for women over 35, and these results are robust to a number of specification tests.

Keywords

infertility; impaired fecundity; insurance mandates; fertility

I. Introduction

In 2002, fertility problems affected 7.3 million women in the United States, up 62% from 1982.[1] This increase has occurred across almost all subgroups of women, including along the dimensions of marital status, income, education, race, and ethnicity (Chandra and Stephen, 1998). Medical treatment for fertility problems can be extremely expensive.[2] Despite the high financial costs of infertility treatment, health insurance coverage of such treatments is limited. Nationwide, only 25% of health care plans cover infertility treatment, and coverage varies significantly by state.[3] As a result of the high (and often uninsured) costs associated with

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[1]Chandra and Stephen (2005). The percentage increase in incidence over this time period is smaller, at 44%.

[2] Most instances of impaired fecundity are treated by "conventional" methods such as drug treatment or surgical repair of reproductive organs. Some of the less invasive therapies such as hormone therapy can range from $200–$3,000 per cycle. Tubal surgery can range from $10,000–$15,000, requires a hospital stay and poses a high risk of complication (Resolve, 2003). *In vitro* fertilization (IVF) accounts for approximately five percent of all infertility treatments, and the average cost of an IVF cycle in the United States is $12,400 (ASRM, 2003).

[3] As a comparison, in 2002 78% of covered workers had coverage for oral contraceptives (Kaiser Foundation, 2002), and a study of health plans found that 57% covered colonoscopy (Klabunde et al., 2004).

treatment, medical assistance for infertility is sought primarily by women and couples that are white, college-educated, and affluent.[4]

In response to a perceived need for coverage, legislation was introduced at the federal level in 2005 that would require health plans to provide infertility benefits.[5] As the fraction of the population affected by infertility continues to rise, there are likely to be continued efforts to mandate coverage to increase access to treatment. Understanding the costs and benefits of these policies thus becomes increasingly important.

As of 2003, fifteen states have enacted some form of infertility insurance mandate. I use Vital Statistics Detail Natality Data and Census population estimates to examine whether these state-level mandates were successful in increasing fertility rates. Using a difference-in-differences approach, I exploit variation in the enactment of mandates both across states and over time, and identify treatment and control groups that should have been differentially affected by infertility coverage. My results suggest that the mandates significantly increase first birth rates for women over 35, and these results are robust to a number of specification tests.

II. State Mandates

The first state-level infertility insurance mandate was enacted by West Virginia in 1977. Since that time, fourteen other states have passed mandates, and additional states have ongoing legislative advocacy efforts in this area.[6] The mandates vary along several dimensions. A mandate “to cover” requires that health insurance companies provide coverage of infertility treatment as a benefit included in every policy. A mandate “to offer” requires that health insurance companies make available for purchase a policy which offers coverage of infertility treatment. In addition, some mandates exclude coverage of *in vitro* fertilization (IVF), which is one of the most expensive treatments available for infertility. Finally, some mandates cover all health plans, while others either exclude health maintenance organizations (HMOs) or only cover HMOs. Table 1 provides a list of the states with mandates currently in place, the date the mandates were enacted, whether the provisions are mandates to cover or mandates to offer, whether the mandates cover IVF, and how the mandates treat HMOs.[7] As of 2000, these mandates were in place in only thirteen states. However, as is illustrated in Table 1, these mandates affect a relatively large share of the births in the United States. In 1985, less than one percent of all live births in the US occurred in mandate states, but by 1995, 46% of all births were in mandate states.

Reducing the price of infertility treatment could lead to an increase in utilization of treatments. This could be true if the mandate expands access to individuals who previously could not afford treatment, or if individuals who were previously receiving treatment now choose to consume higher quantities (or a higher quality) of treatment. However, it is also possible that these mandates have no effect on access or on treatment consumed, but simply provide windfall gains to those individuals who would have purchased treatment in the absence of insurance coverage.[8]

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[4] Women with private health insurance coverage were 50% more likely to have received services, as were women with income more than 300% of the poverty line (Stephen and Chandra, 2000).

[5] The Family Building Act of 2005 (HR 735) would require insurance coverage of infertility treatment (including up to four *in vitro* fertilization attempts) by all group health plans that also require obstetrical benefits.

[6] Of the traditional economic justifications for mandated benefits (e.g. Summers, 1989; Gruber, 1994a)), the best efficiency argument is that asymmetric information between the patient and insurer will lead to an adverse selection problem so that benefits will not be provided by the private market (e.g. Rothschild and Stiglitz, 1976).

[7] Detailed information on these mandates, including any further restrictions placed on coverage, can be found in Schmidt (2005).

[8] Mandates may also have dynamic effects on the timing of births. Individuals could seek treatment earlier, which is beneficial from a medical perspective. Alternatively, individuals could further delay childbearing, with the knowledge that they will ultimately be covered.

Several papers in the medical literature have examined the effects of state level infertility insurance mandates on a number of outcomes. Griffin and Panak (1998) use insurance data and show that the Massachusetts mandate is associated with increases in the use of IVF. Jain et al. (2002) use clinic data from 1998 and find that states with required coverage for IVF have the highest rates of IVF utilization. They also find that in states with full IVF coverage, fewer embryos per cycle are transferred, leading to a reduced risk of multiple births per cycle. While these studies are an important contribution to our understanding of the effects of these mandates, they have several shortcomings. First, they focus exclusively on IVF, even though IVF comprises only about 5% of all infertility treatments. In addition, the studies are cross-sectional and cannot control for unobservable differences in patients or clinics that may be state-specific. It is therefore impossible to tell if, for example, higher rates of utilization in Massachusetts are caused by the mandate, or if Massachusetts had higher rates of utilization prior to the mandate.

In the past several years, a number of economists have also looked at the effects of these mandates, using differences-in-differences-in-differences approaches where they exploit variation across states, over time, as well as by demographic categories (usually age). Bitler and Schmidt (2006a, 2006b) use data from the National Survey of Family Growth and find that while the mandates have not reduced disparities in use of infertility treatment by race and socioeconomic status, they are associated with an increase in reported use of infertility treatment among highly educated, older women. Buckles (2005) uses Vital Statistics Detail Natality Data and finds that mandates that cover IVF are associated with a higher age at first birth.[9] Bitler (2006) shows that the mandates are associated with higher rates of twinning among older mothers. She also uses the mandates to examine the effects of assisted reproductive technologies on infant health outcomes. She finds small negative effects of these technologies on length of gestation, birth weight, and Apgar scores.

Hamilton and McManus (2005) develop a model of the market for IVF and use data from clinics at the Metropolitan Statistical Area level to test the model’s predictions. Using data from 1995–2000, they confirm the findings of Jain et al. (2002) that a mandate increases IVF utilization rates. They also find that clinics are attracted to areas where women are more educated and wealthier, but find no evidence that clinics are attracted to places where mandates are in effect.

In this paper, I use a difference-in-differences-in-differences approach where I exploit variation in the enactment of mandates both across states and over time, as well as by age group to determine whether these mandates have been successful in increasing first birth rates for those women most likely to demand services. By analyzing birth rates rather than IVF utilization, I can estimate the total effects of the mandates on fertility, which will include increases resulting from all types of infertility treatments. In addition, looking at first birth rates rather than changes in age at first birth or probabilities of multiples allows estimates to be made of the number of new births that occur in a given year as a result of these mandates.

III. Data

Information on births comes from Vital Statistics Detail Natality Data, gathered by the National Center for Health Statistics. This information is based on birth certificate data, and includes specific information about the timing, parity (whether it was a first or subsequent birth), and plurality (whether it was a single, twin, triplet, or higher order birth) of each birth. These data also include demographic information on the mother, including age, race, ethnicity, marital status, and educational attainment. Geographic information about the mother’s state of

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[9] Buckles also looks at impacts of the mandates on economic variables such as labor force participation and wages.

residence is also provided. Beginning in 1985 the data cover every birth in the United States. The counts of births by state, year, race, and five-year age cohort are used to generate birth rates.[10]

The denominators for birth rates must come from another data source, since the birth certificates only provide information on those women who actually give birth. Population estimates are available for black and non-black women by age and state through the Census Bureau, and can be used to calculate birth rates. Ideally, the denominator would include only those women who have not yet had a first birth. However, the Census data do not allow for further breakouts. Birth counts by parity of the mother can be calculated, but denominators with counts of women by the number of children they have already borne cannot be generated from the Census data.[11] Other control variables, collected by state and by year, come from a variety of publicly available sources. Summary statistics for the data set, which covers the years 1981–1999, can be found in the Appendix.

IV. Model Specification

Variation in access to infertility treatment induced by the mandates exists both across states and over time. In addition, certain groups should be more likely to be affected by the mandates. Specifically, the probability that a woman experiences infertility is extremely low for young women, and increases with age. The American Society for Reproductive Medicine identifies 35 as an important turning point in the risk of infertility problems. Their *Patient's Fact Sheet on the Prediction of Fertility Potential in Older Female Patients* states that "approximately one-third of couples in which the female partner is age 35 or older will have problems with fertility." Data from the National Survey of Family Growth shows that while 10% of women 29 or younger have reported experiencing fertility problems, 23% of women aged 35–44 report such problems. Similar differences exist in reports of infertility treatments received − 7.5% of women 29 and younger report receiving treatment, compared with 20% of women 35–44 (Bitler and Schmidt 2006b). Therefore, those women most likely to be affected by the mandates are those who have delayed childbearing, and specifically those women 35 and older.[12,13]

I estimate the following Differences-in-Differences-in-Differences (DDD) model[14]:

$$\ln(fstbthrt)_{ajt} = \alpha + \beta_1 Z_{jt} + \beta_2 \eta_a + \beta_3 \delta_j + \beta_4 \tau_t + \beta_5(\eta_a \times \tau_t) + \beta_6(\eta_a \delta_j) + \beta_7 Mandate_{j(t-2)} + \beta_8(Mandate_{j(t-2)} \times Over35) + \varepsilon_{ajt} \quad (1)$$

where the dependent variable is the log first birth rate for age cohort *a* in state *j* and year *t*. The first birth rate is equal to the number of first births within an age cohort-race-state-year cell, divided by the number of women in that same age cohort-race-state-year cell. I focus on first

[10] Five year cohorts are used up to the age of 49 (15–19, 20–24, 25–29, 30–34, 35–39, 40–44, 45–49). The maximum age reported by the Vital Statistics Data was 49 through 1996, and in 1997 births to women up to the age of 54 were also reported. I omit births to women 50 and older to maintain consistency across years.

[11] Parity is available in the June Fertility Supplements to the Current Population Survey, but these data are not collected every year, and the sample size is too small to conduct state-level analyses except for the largest states.

[12] Even though the mandates are more likely to affect women 35 and older, it is possible that they could also have an effect on the group of younger women. This could happen for two reasons: First, younger women might also be receiving infertility treatments. Alternatively, it is possible that younger women may be more likely to delay childbearing if they know infertility treatment is covered by their insurance. This empirical approach allows me to test for effects on the group of younger women as well as for the older women.

[13] There are other demographic dimensions along which demand for services varies, including marital status and education level (bthose individuals who are married and have higher levels of education are much more likely to demand services. Bitler and Schmidt (2006b) use the NSFG to look at reports of infertility treatment received and how they are affected by the mandates, and find effects of both education and marital status. Although the Vital Statistics data do include information on both marital status and education level of the mother, these variables are not reported consistently in the Vital Statistics for all states during the years in my sample. Specifically, these variables are either missing or imputed for a number of the mandate states during the years of my sample.

[14] This approach follows Gruber (1994b) and Meyer (1995).

birth rates, because treatments are more likely to be sought by women who have not already borne children.

The DDD model exploits variation across states (mandate versus nonmandate), over time (pre-mandate versus post-mandate), and across age (women 35 and older versus women younger than 35). The fixed effects control for time-invariant differences in fertility by age ($\beta 2$), time-invariant differences in fertility by state ($\beta 3$), and national time-series changes in fertility ($\beta 4$). The second-level interactions control for age-specific changes over time in fertility ($\beta 5$), state-specific differences in the timing of fertility ($\beta 6$), and changes over time (pre-and post-mandate) in the mandate states ($\beta 7$). The third-level interaction ($\beta 8$) captures variation in fertility specific to the "treatment" group (women 35 and older) in the mandate states (relative to nonmandate states) in the years after the mandate was passed (relative to before the law), and is therefore the DDD estimator of the effects of the mandates on first birth rates.

In the specification, a mandate is not allowed to affect fertility rates in the year it is enacted, but can instead affect fertility rates with a two year lag. This is to account for two factors: 1) infertility treatments often do not result in an immediate conception; and 2) even if a conception occurs immediately, there is still a necessary nine-month waiting period before those new conceptions can affect fertility rates.[15]

The coefficient $\beta 7$ is an estimate of the percentage change in first births associated with the mandates for the entire population of women. However, we would expect the effect of the mandates to differ along several different dimensions. First, as mentioned earlier, age plays an important role in the need for fertility services. In addition, there are likely to be racial differences in the effect of a mandate, since evidence suggests that even though African American women are more likely to experience fertility problems than are whites, they are less likely to seek treatment for those fertility problems. In addition, we would not expect the mandate to affect all individuals in the population. State-level mandates only apply to individuals who have private insurance. In addition, the Employment Retirement Income Security Act of 1974 (ERISA) preempts specific state regulation of self-funded insurance plans provided by private-sector employers, so mandated benefits do not affect individuals in firms that self-insure. In the empirical analysis, I will test the robustness of results along these dimensions.[16]

The *Z* vector controls for variables that will vary across states and over time that might also affect birth rates. These include variables that reflect economic conditions, including the state unemployment rate, log median usual weekly earnings, log tenth percentile weekly usual earnings, and female labor force participation rates.[17] A variable describing whether parental involvement is required for minors to obtain an abortion is also included.[18] The *Z* vector also includes the log maximum level of state welfare benefits available to a family of three.[19] The

[15] If a one-year lag structure is used, results are qualitatively similar but the magnitude of the estimated coefficients is slightly smaller.

[16] In analyzing the effects of mandates along these dimensions, I may be introducing additional measurement error. There is some evidence that there might be biases in birth rates generated from Vital Statistics and Census estimates that primarily affect estimates of first birth timing among nonwhites (e.g. Morgan et al., 1999). Therefore, the constructed first birth rates for whites are likely to be measured with less error than the constructed birth rates for African American women. In addition, some states have small African American populations, which could lead to more noise in measuring birth rates. Finally, state-level measures of firm size (a proxy for self-insurance) and the share of the population with private health insurance are generated from the Current Population Survey, and are therefore measured with less error for large states than small states. These sources of measurement error could lead to downward bias in my estimates, and are likely to mean that my results for African Americans are less reliable than the results for whites.

[17] Dehejia and Lleras-Muney (2004) find that parental characteristics vary systematically depending on the unemployment rate at the time of a baby's conception.

[18] Fertility regressions also often include restrictions on Medicaid funding of abortions. I do not control for these policies here, since few states changed these policies over my sample period (see Levine, 2002).

[19] These values are for the Aid to Families with Dependent Children program through 1996, and for the Temporary Assistance for Needy Families program from 1997 through 1999.

error term is represented by ε. Difference-in-differences estimation can lead to artificially low standard error estimates if sthe outcomes and the policy changes of interest are serially correlated (Bertrand et al., 2002). I calculate White robust standard errors clustered by state to correct for this potential problem.[20]

I then estimate a differences-in-differences regression, stratified by age group (under 35 versus 35 and older):

$$\begin{aligned}\ln(fstbthrt)_{ajt} &= \alpha + \beta_1 Z_{jt} + \beta_2 \eta_a + \beta_3 \delta_j + \beta_4 \tau_t \\ &+ \beta_5(\eta_a \times \tau_t) + \beta_6(\eta_a \delta_j) + \beta_7 Mandate_{j(t-2)} + \varepsilon_{ajt}\end{aligned} \tag{2}$$

which retains the age-state and age-year interactions used in the DDD approach. The advantage of this age-stratified approach relative to the DDD approach shown above is that it allows the effects of all control variables in the *Z* vector to vary by age. Since age group is an important source of identification, it is important to fully control for any age-specific effects of the other control variables. This approach avoids incorrect attribution of these effects to the mandates.

V. Results

A. Main Results

Results from the estimation of the DDD model in equation (1) can be found in the first Column of Table 2. Panel A first presents results for the full sample. Having a mandate in effect in a state has a small negative effect on overall birth rates, with an estimated coefficient of −0.05 that is statistically significant at the ten-percent level. For women 35 and older, the presence of a mandate leads to an increase in birth rates that is large and statistically significant. The sum of the two coefficients implies that the presence of a mandate increases birth rates for women over 35 by 8%, and this effect is statistically significant at the five-percent level. However, Columns 2 and 3 present results for the age-stratified DD model, and show that once the effects of all control variables are allowed to vary by age, there are no significant effects of the mandates on either the younger or the older women.[21]

However, there are two reasons why we might expect effects to vary by the race of the mother. First, there exist large differences in fertility timing by race. In addition, the existing literature shows that even though white women are less likely to experience fertility problems than are black and Hispanic women, they are more likely to seek treatment (e.g. Stephen and Chandra, 2000; Bitler and Schmidt, 2006b). I re-estimate the model separately by the race of the mother, and Panels B and C present results for whites and blacks, respectively. The results in Panel B show that for white mothers, there are no significant effects of the mandates on women younger than 35, but there are large and significant effects of the mandates on women 35 and older, ranging from a 19% increase in the age-stratified DD model to a 22% increase in the DDD model. In both models, the effects for older women are statistically larger than the effects for younger women at the five-percent level.[22] The results in Panel C show no significant effects of mandates on black women for any age group in any specification, and among black women, the results show no statistical differences in the effects of mandates by age. Tests of the equality of coefficients across the populations of white and black women show that in the age-stratified DD model for women over 35, the coefficients are statistically different by race (p=0.021).

[20] Results are from unweighted regressions, treating each state equally (e.g. Dickens, 1990). Results are robust to weighting by the population counts in each cell.

[21] The estimated coefficient for the older women is statistically larger than the coefficient for the younger women at the ten-percent level.

[22] In the age-stratified DD results, coefficients in bold statistically differ by age group at the ten-percent level or higher.

Since the age-stratified DD specification provides more conservative estimates than the DDD specification, I focus on results from that specification for the remainder of the paper. In addition, since the age-stratified DD results are statistically different by race of the mother, I estimate the models separately by race in all subsequent tables.

In 1999, Vital Statistics Detail Natality data report 104,386 first births to nonblack women over the age of 35 (with a first birth rate for this group of 0.0029). A 19% increase would imply an additional 19,833 births to this group of women. As a comparison, reports from the Center for Disease Control imply that roughly 9,557 live births to women 35 and older resulted from IVF cycles that began in 1998. Many of these births would have occurred in 1999. We might ask whether this ratio is plausible – that is, how many new births generated by the mandates would we expect to be IVF related? A treatment algorithm in Gleicher (2000) suggests that for every 88 births generated by infertility diagnosis and treatment, 24 would be likely to be IVF related, or a ratio of 27%. The combination of my estimates and the CDC estimates suggests a higher ratio of IVF births to all births of 48%. However, given that IVF is one of the most expensive treatments for infertility, it is likely that individuals are more likely to be budget constrained in the absence of insurance for IVF than for diagnosis and other treatments such as ovulation-inducing drugs, and therefore not surprising that the mandates lead to a higher share of IVF births relative to total births.

B. Type of Mandate

The effect of mandates on first birth rates might be expected to vary depending on the type of mandate enacted.[23] In particular, one might expect a "mandate to cover" to affect births differently than a "mandate to offer", or a mandate that includes IVF to affect births differently than one that excludes IVF. In Table 3, I break out the mandates by type. Panel A presents the baseline results from Table 2. Panel B shows results from a specification that separates "mandates to cover" from "mandates to offer". The two mandate variables are jointly statistically significant only for white women 35 and older, and still suggest that the mandates significantly increase first birth rates for this group. However, they provide no evidence that a "mandate to cover" affects first birth rates differently than a "mandate to offer". The estimates imply approximately a 19 percent increase in first birth rates for women over the age of 35, and the coefficients on the cover and the offer mandates are not statistically different from each other.[24]

In Panel C, I break out mandates that cover IVF versus those that exclude IVF. Again, the mandate variables are jointly statistically significant only for white women over the age of 35. The point estimate for mandates that cover IVF suggests a 22% increase in first birth rates, and is roughly 3 times the size of the coefficient for those states with mandates that do not cover IVF. However, these coefficients are not statistically different from each other at conventional confidence levels. The effect of IVF mandates is statistically different for the older women

[23] In addition, the availability of infertility clinics may determine the effectiveness of the mandates. The first reports publishing data on infertility clinics were published in 1995, and in 1995, all states in my sample had at least one infertility clinic, with the exception of Montana. Montana's mandate excludes IVF coverage, so the types of treatment covered in that state would be available even in the absence of a clinic. More generally, to the extent that services are unavailable in some of my mandate states during the sample period, this should bias me against finding an effect of the mandates. However, all results are robust to reclassifying Montana as a non-mandate state.

[24] There are three offer states: California, Connecticut, and Texas. Regressions where each of these states is eliminated in turn show that the effect of an offer mandate on first births of older women is entirely driven by the state of Connecticut. Further analysis using Vital Statistics county level data shows that the increase in first births associated with the offer mandate in Connecticut is entirely driven by Fairfield County, part of the New York City metropolitan area. If a large number of Fairfield County workers commute to work in New York City, it is possible that they would be insured by firms responding to the New York mandate, a "cover" mandate passed only one year after the Connecticut mandate. In fact, nineteen percent of all workers in that region commuted to New York State for employment in 2000 (Southwestern Regional Planning Authority Calculations from the U.S. Census Bureau County Worker Flow Files).

than for the younger women, but the effect of mandates that exclude IVF does not statistically differ by age.[25]

C. Coverage of Population

In addition to differences in the types of mandates, mandates are also likely to have different effects depending on how much of the population they reach. First, the enactment of state mandates will generally only help those individuals who already have access to health insurance. In Panel A of Table 4, I interact the mandate variables with the share of the women 15–44 with private health insurance.[26] This variable has no significant effect on the first birth rates of black women or on white women under 35, but has a positive and statistically significant effect on white women 35 and older. The coefficient for white women over 35 is statistically greater than the coefficient for younger white women at the five percent level.

In addition, even among those firms that provide their employee with health insurance, not all are subject to the mandates. Under the Employer Retirement Income Security Act of 1974 (ERISA), firms that self-insure are exempt from the mandates. It has been argued that the passage of mandates could induce firms to self-insure in order to avoid compliance (e.g. Jensen et al. (1995)). Ideally, I would have information on the share of employees by state and year in firms that self insure. However, these data are not available. It has been shown that the primary determinant of self-insurance by firms is firm size, with large firms being significantly more likely to self-insure (Park, 2000, Gabel et al., 2003). Gabel et al. (2003) report that in 1993, 13% of employees in firms sized 3–199 employees were enrolled in self-insured plans, as compared with 46% of employees in firms sized 200–999 employees, and over 60% for employees in firms with 1000 or more employees.

In panel B of Table 4, I interact the mandate variable with variables that indicate the share of employees who work in firms of a particular size (Categories are less than 99 employees, 100–999 employees, and 1000 or more employees.)[27] For white women under the age of 35, the mandate-firm size interactions are not jointly significantly different from zero, nor are any of the individual coefficients. However, for white women over 35, the effects of mandates are jointly statistically different from zero, and the pattern is suggestive of a larger effect for the midsize firms.[28]

In Panel C of Table 3, I allow mandates covering all health plans to affect births differently than mandates that either exclude HMOs or that only cover HMOs, by replacing the dummy variable for mandate with the share of the population that is covered by the mandate. For states with no mandate, this variable is equal to zero; for states with a mandate covering all health

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[25] While the mandate variables are not jointly statistically significant for African American women of either age group, some specifications result in estimated coefficients that are negative and statistically significant. For example, the results imply that a "mandate to offer" and a non-IVF mandate lead to decreases in the first birth rates of African American women that are statistically significant at the ten-percent level. One possible explanation is that, since only 13 states pass mandates during my sample period, and a much smaller number pass any particular type of mandate (three pass "offer" mandates, while four pass "non-IVF" mandates). That means that these mandate effects are being driven by a very small number of states. Since African American women are not distributed evenly by state, it may be the case that these results are being driven by cells with small numbers of women, and even smaller numbers of births to older, African American women. In addition, as mentioned above, Morgan et al. (1999) find some evidence that the first birth timing of black women may be measured with more error in the Vital Statistics/Census estimates.

[26] These data come from the March Current Population Survey.

[27] These data are generated from the March Current Population Survey. The CPS only began asking this series of questions about firm size in the 1988 survey (for calendar year 1987). A longer series can be constructed using data from the May 1980 CPS, but the firm size categories available group all firms with 100 or more employees together and are therefore unsuitable for looking at the largest firms.

[28] For younger African-American women, the coefficient on the interaction between mandate and the share of employees in firms with 99 or fewer employees is large, negative, and statistically significant at the five-percent level. This leads to joint significance of the mandate-firm size interactions for the younger African American women. Regressions where mandate state is eliminated in turn show that this effect is entirely driven by the state of Montana, where non-Hispanic African Americans made up only 0.3 percent of the population in 1995. If Montana is dropped from the regressions, the mandate-firm size interactions are no longer jointly significant for African American women.

plans, it is equal to one; for states with mandates only covering HMOs, it is equal to the HMO penetration rate; and for states with mandates excluding HMOs, and it is equal to one minus the HMO penetration rate.[29] The main results presented above are robust to this alternate specification. The coefficient for white women 35 and older is 0.21, statistically significant at the five-percent level, and is statistically larger than the corresponding coefficient for younger white women (p-value=0.05).

D. Robustness Tests

In Table 5, I re-estimate the main specification in Column 1 with additional controls for state-specific quadratic time trends. When these time trends are included, the point estimate for older, white women falls, but suggests that the mandates increase first birth rates for this group by 9%. This coefficient is less precisely estimated, but approaches statistical significance at the ten percent level (p=0.119).[30] It is roughly eight times the magnitude of the coefficient for the younger white women, but is not statistically different at conventional confidence levels (p=0.191).

In addition, assisted reproductive technologies have changed dramatically between 1981 and 1999. In particular, use of *in vitro* fertilization has increased dramatically over the later years in my sample. The mandates should therefore have a greater impact on first birth rates in the later years of my sample. In Table 6, I interact the mandate variable with dummies for five-year calendar periods 1981–1984, 1985–1989, 1990–1994, and 1995–1999. It is impossible to fully disentangle the effect of technology from the fact that more states were passing mandates over the years in the sample. However, the results in Table 6 are suggestive of a larger effect of the mandates on the first birth rates of older white women in the later years of the sample.

As a final robustness check, I re-estimate the model using log second and higher parity birth rates as the dependent variable. If infertility treatments are more likely to be sought by women who have not already given birth, the effects of mandates on higher parity birth rates should be smaller. Results in Table 7 confirm this. The estimated coefficient on mandates for women over 35 falls from 0.19 for first birth rates to 0.04 for second and higher birth rates, and is no longer statistically different from zero. It is also statistically smaller than the mandate coefficient from the first birth rate regression (p=0.080). For no other group (younger white women, or African American women of age groups, are the coefficients statistically different either from zero or from their counterparts in the baseline regression.

VI. Conclusion

For a woman or couple faced with fertility problems, a conception and birth is the ultimate goal. From this perspective, it appears as if the state-level infertility insurance mandates have been a success, as I find that the mandates significantly increase first birth rates among white women over 35. The estimated effects are robust to a wide variety of specification tests.

The finding of an effect of the mandates on fertility is somewhat surprising. Most of the research on health insurance mandates originally focused on the potential costs of mandating health insurance benefits in terms of reduced wages, reduced employment, or reductions in the probability of insurance being offered, and found few effects of mandates along most of these dimensions (Gruber, 1994b; Kaestner and Simon, 2002). In addition, many supposed "high-cost" mandates, such as mental health mandates, which should reduce the cost of services to consumers, seem to have little effect on utilization of related health care services or on health

[29] HMO penetration rates come from Interstudy, various years.

[30] Use of state linear time trends is not supported by the data, and adding state quadratic time trends requires estimation of over 100 additional parameters.

care outcomes (Bao and Sturm, 2004; Pacula and Sturm, 2000). However, in the case of infertility treatment, those individuals who are most likely to demand services (women who are older and highly educated) are also most likely to be affected by mandates due to their higher probability of having private health insurance. This is not true of all mandates, and could help explain the finding of a utilization impact.[31]

Consistent with other research in the economics and clinical literatures (e.g. Bitler and Schmidt, 2006a; Jain and Hornstein, 2005), I find no evidence that the mandates have reduced racial disparities in access to treatment. My results suggest that the mandates are not associated with increases in the first birth rates of nonwhite women, despite the fact that nonwhite women have a significantly higher likelihood of fertility problems. These findings could also be related in part to differentials in private insurance coverage, but further research is necessary to fully disentangle possible explanations.[32]

As demographic changes and continued trends in delay of childbearing cause infertility to become an increasingly common medical problem, advocacy groups are likely to continue to pressure policymakers to enact mandated benefits at both the state and federal levels. Insurance providers are likely to continue to resist these pressures. Further evaluation of the effects of these mandated benefits is essential to informing this policy debate.

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[31] In work using the National Survey of Family Growth, Bitler and Schmidt (2006b) explore this issue in greater detail.

[32] Other explanations suggested in the clinical literature include access to information, racial discrimination, fewer referrals by physicians, and cultural bias against infertility treatment (Jain and Hornstein, 2005).

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Appendix: Summary Statistics

Variable	Mean
First birth rate	0.0229 (0.0242)
Log median weekly earnings	5.9056 (0.1401)
Log 10th percentile weekly earnings	4.7676 (0.1908)
Unemployment rate	6.2545 (2.1986)
Female labor force participation rate	58.152 (5.1683)
Maximum AFDC/TANF benefit	470.388 (183.166)

Variable	Mean
Parental involvement abortion restrictions	0.4262
Share with private health insurance	0.6540 (0.2017)
Share in firms with less than 25 emp	0.1038 (0.1132)
Share in firms with 25–99 emp	0.0599 (0.0703)
Share in firms with 100+ emp	0.3027 (0.2429)
HMO penetration rate	0.1388 (0.1241)
Number of state-year-age-race cells	13566

Notes: Standard deviations in parentheses. Observations are state/year/5-year age cohort/race cells.

Table 1

State Mandated Infertility Insurance

State	Year Enacted	Mandate to Cover/ Mandate to Offer	*In Vitro* Fertilization Coverage?	HMO Treatment	Percent of US Births in 1995
Arkansas	1987[a]	Cover	Yes	HMOs excluded	0.86%
California	1989	Offer	No	All plans covered	14.15%
Connecticut	1989	Offer	Yes	HMOs excluded	1.13%
Hawaii	1987	Cover	Yes	All plans covered	0.48%
Illinois	1991	Cover	Yes	All plans covered	4.68%
Maryland	1985	Cover	Yes	All plans covered	1.74%
Massachusetts	1987	Cover	Yes	All plans covered	2.12%
Montana	1987	Cover	No	HMOs only	0.28%
New York	1990[b]	Cover	No	HMOs excluded	6.98%
Ohio	1991[c]	Cover	Yes	HMOs only	3.97%
Rhode Island	1989	Cover	Yes	All plans covered	0.35%
Texas	1987	Offer	Yes	All plans covered	8.37%
West Virginia	1977[d]	Cover	No	HMOs only	0.57%
Total					***45.68%***

Sources: Resolve (www.resolve.org), state laws (see Appendix A ofSchmidt, 2005), and National Center for Health Statistics, Vital Statistics of the United States, Volume I (Natality), various years. Louisiana and New Jersey each enacted mandates in 2001, but since this is out of my sample period, I do not include them here.

[a]Some coverage for IVF was first required of in 1987. The law was revised in 1991 to set maximum and minimum benefit levels and to establish standards for determining whether a policy or certificate must include coverage (see Appendix A ofSchmidt, 2005).

[b]In 2002, New York passed a revised law that clarified the 1990 legislation and appropriated $10 million to a pilot project to help pay for IVF for a small number of individuals.

[c]The original 1991 law did not specifically exclude IVF, but in 1997 the Superintendent of Insurance stated that IVF, GIFT and ZIFT were not essential for the protection of an individual's health and were therefore not subject to mandated insurance coverage. I code Ohio as an IVF state between 1991 and 1997.

[d]In 2001, the law was amended to mandate HMOs to cover infertility treatment only as a "preventative service" benefit (thus excluding IVF).

Table 2

Effects of Mandates on First Birth Rates

	DDD	DD Under 35	DD 35 and older
A. Full sample			
Mandate	−0.0537 (0.0283)*	**−0.0292 (0.0250)**	**0.0521 (0.0421)**
Mandate * 35Plus	0.1385 (0.0537) ***	--	--
B. Whites			
Mandate	−0.0180 (0.0693)	**0.0071 (0.0131)**	**0.1899 (0.0784)** **
Mandate * 35Plus	0.2414 (0.0753)***	--	--
C. Blacks			
Mandate	−0.0895 (0.0561)	−0.0655 (0.0485)	−0.0858 (0.0654)
Mandate * 35Plus	0.0357 (0.0806)	--	--

Notes: Dependent variable is the log of the first birth rate in a state-year-race-age cohort cell. All regressions include state fixed effects, year fixed effects, and a full set of age-year interactions. White robust standard errors clustered by state in parentheses. Regressions are weighted by cell-level population counts. Levels of statistical significance: *** denotes significance at the one-percent level; ** at the five-percent level; and * at the ten-percent level. Coefficients in bold statistically differ by age group at the ten-percent level or higher.

Table 3
Effects of Mandates on First Birth Rates, Type of Mandates

	Whites		Blacks	
	Under 35	**35 and Over**	**Under 35**	**35 and Over**
A. Baseline Results				
Mandate	**0.0071 (0.0131)**	**0.1899 (0.0784)** **	−0.0655 (0.0485)	−0.0858 (0.0654)
B. Cover vs Offer				
Mandate to cover	**0.0049 (0.0147)**	**0.1911 (0.0841)** **	−0.0776 (0.0559)	−0.0541 (0.0707)
Mandate to offer	0.0163 (0.0187)	0.1852 (0.1198)	−0.0163 (0.0409)	−0.2147 (0.1071) *
F test of joint significance of cover & offer coeff.	0.39	2.94 **	0.98	2.05
F-test of equality of cover v offer coefficients	0.28	0.00	1.26	1.99
C. IVF vs non-IVF				
Mandate covers IVF	**0.0067 (0.0145)**	**0.2244 (0.0824)** ***	−0.0341 (0.0479)	−0.0636 (0.0731)
Mandate excludes IVF	0.0084 (0.0219)	0.0760 (0.1102)	−0.1690 (0.0829) *	−0.1590 (0.0960) *
F test of joint significance of IVF & non-IVF coeff	0.15	3.71 **	2.22	1.46
F-test of equality of IVF v non-IVF coefficients	0.01	1.63	2.15	0.86

Notes: Dependent variable is the log of the first birth rate in a state-year-race-age cohort cell. White robust standard errors clustered by state in parentheses. Regressions include state and year fixed effects, as well as age*year interactions. Regressions are weighted by cell-level population counts. Levels of statistical significance: *** denotes significance at the one-percent level; ** at the five-percent level; and * at the ten-percent level. Coefficients in bold statistically differ by age group at the ten-percent level or higher.

Table 4
Effects of Mandates on First Birth Rates, by Coverage of the Population

	Whites		Blacks	
	Under 35	**35 and Over**	**Under 35**	**35 and Over**
A. Private Health Insurance				
Mandate * share with private health insurance	**0.0083 (0.0170)**	**0.2552 (0.0955)** ***	−0.0008 (0.0421)	−0.1087 (0.1074)
B. Firm size				
Mandate * share in firms with less than 99 emp	−0.0294 (0.0518)	−0.6446 (0.5689)	**−1.4028 (0.5498)** **	**0.1759 (0.5506)**
Mandate * share in firms with 100–999 emp	0.1449 (0.1635)	1.1101 (1.4008)	0.3468 (0.9482)	−0.1524 (1.5887)
Mandate * share in firms with 1000 + emp	−0.0338 (0.0707)	0.4450 (0.6970)	1.1796 (1.0573)	−0.3685 (1.0287)
F test of joint significance of firm size*mandate interactions	0.28	3.43 **	6.82 ***	0.56
F-test of equality of firm size* mandate interactions	0.40	0.99	10.13 ***	0.14
C. Share of population affected by Mandate (due to variation in treatment of HMOs)				
	0.0408 (0.0115) ***	**0.2064 (0.0807)** **	−0.0495 (0.0511)	−0.1772 (0.0912) *

Notes: Dependent variable is the log of the first birth rate in a state-year-race-age cohort cell. White robust standard errors clustered by state in parentheses. Regressions include state and year fixed effects, as well as age*year interactions. Regressions are weighted by cell-level population counts. Levels of statistical significance: *** denotes significance at the one-percent level; ** at the five-percent level; and * at the ten-percent level. Coefficients in bold statistically differ by age group at the ten-percent level or higher.

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Table 5
Robustness Test: Adding State-Specific Time Trends

	Whites		Blacks	
	Under 35	**35 and Over**	**Under 35**	**35 and Over**
A. Baseline results				
Mandate	**0.0071 (0.0131)**	**0.1899 (0.0784)** **	−0.0655 (0.0485)	−0.0858 (0.0654)
B. With state-specific time trends				
Mandate	0.0120 (0.0089)	0.0927 (0.0594)	−0.1055 (0.0478) **	−0.1432 (0.1214)

Notes: Dependent variable is the log of the first birth rate in a state-year-race-age cohort cell. White robust standard errors clustered by state in parentheses. Regressions include state and year fixed effects, as well as age*year interactions. Regressions are weighted by cell-level population counts. Levels of statistical significance: *** denotes significance at the one-percent level; ** at the five-percent level; and * at the ten-percent level. Coefficients in bold statistically differ by age group at the ten-percent level or higher.

Table 6

Robustness Test: Calendar Year Effects

	Whites		Blacks	
	Under 35	**35 and Over**	**Under 35**	**35 and Over**
Mandate * (1980–1984)	−0.0074 (0.0189)	0.1634 (0.1180)	−0.1689 (0.0563) ***	−0.1868 (0.1119) *
Mandate * (1985–1989)	0.0024 (0.0163)	−0.0231 (0.1048)	−0.0617 (0.0371) *	−0.1212 (0.1054)
Mandate * (1990–1994)	**0.0114 (0.0135)**	**0.1847 (0.0706)** ***	−0.0651 (0.0368) *	−0.0853 (0.0870)
Mandate * (1995–1999)	**0.0043 (0.0171)**	**0.2318 (0.0973)** **	−0.0664 (0.0663)	−0.0798 (0.0658)
F test of joint significance	0.29	2.25 *	3.01 **	0.80

Notes: Dependent variable is the log of the first birth rate in a state-year-race-age cohort cell. White robust standard errors clustered by state in parentheses. Regressions include state and year fixed effects, as well as age*year interactions. Regressions are weighted by cell-level population counts. Levels of statistical significance: *** denotes significance at the one-percent level; ** at the five-percent level; and * at the ten-percent level. Coefficients in bold statistically differ by age group at the ten-percent level or higher.

Table 7

Second and Higher Birth Rates as dependent variable

	Whites		Blacks	
	Under 35	**35 and Over**	**Under 35**	**35 and Over**
A. Baseline results				
Mandate	**0.0071 (0.0131)**	**0.1899 (0.0784)** **	−0.0655 (0.0485)	−0.0858 (0.0654)
B. Dependent var = Second and Higher birth rates				
Mandate	0.0089 (0.0159)	0.0481 (0.0395)	−0.0675 (0.0869)	0.0499 (0.0658)

Notes: Dependent variable is the log of the higher order birth rate (second births and higher) in a state-year-race-age cohort cell. All regressions include state fixed effects, year fixed effects, and a full set of age-year interactions. White robust standard errors clustered by state in parentheses. Regressions are weighted by cell-level population counts. Levels of statistical significance: *** denotes significance at the one-percent level; ** at the five-percent level; and * at the ten-percent level. Coefficients in bold statistically differ by age group at the ten-percent level or higher.