When Do Health Insurance Mandates Matter? The Case of Infertility Treatment[⊗]

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Abstract

The literature on mandated health insurance benefits has found little evidence of effects on the utilization of health care services. A number of explanations have been considered for these findings of no effects. In this paper, we examine whether mandated insurance coverage for infertility treatment affects utilization for a specific subgroup in the population: older, highly educated women. These women are both most likely to experience fertility problems and most likely to have insurance plans affected by the mandates. We find robust evidence that while an effect of the mandates on utilization can not be found for the full population of women, the mandates do have a large and significant effect on utilization for exactly this subgroup.

I. Introduction

Over the past 30 years, state-level mandated health insurance benefits have grown in popularity as a means of trying to regulate the health care system. Currently, well over 1,000 state mandated benefits are in effect (EBRI, 2005). These laws require insurers to cover specific health services or to cover services provided by specific providers. Proponents of mandated insurance benefits aim to affect utilization of particular health services and, ultimately, improve related health outcomes.

Since the seminal paper of Summers (1989), economists have written a great deal on the economics of mandated benefits and the conditions under which mandates have different impacts than taxes. Most of the research in this area 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. However, the literature finds few effects of mandates along most of these dimensions (Gruber, 1994a; 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 care outcomes (Bao and Sturm, 2004; Pacula and Sturm, 2000).[1]

Several possible explanations have been considered for these findings of no effects. First, if employers do not expect a mandate to have a large impact on health care utilization and costs, they may be less likely to oppose the legislation (Bao and Sturm, 2004). Second, it has been suggested that state mandate law[1]s may not be binding, or that they might be undone by the firm, perhaps through lower wages for the groups targeted by the mandate (Gruber 1994b). Some evidence suggests that benefits are similar in firms in mandate states relative to firms in states that do not mandate. In addition, within mandate states, there is evidence that benefits in firms exempt from the mandates are similar to benefits in firms that are affected by the mandates (Acs et al. 1996; Gruber 1994a; Jensen et al. 1998). If mandates do not affect the benefits offered by firms, then they should not affect utilization of services or health outcomes.

Finally, state-level mandated benefits may not affect all individuals within a state. These 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. As such, it is possible that legislation may not affect enough individuals to discern an impact if looking at the entire population. For example, Liu et al. (2004) find that the effect of drive-through delivery laws has been blunted by ERISA. Furthermore, many mandates potentially affect only a subgroup of the population (for example, mental health mandates affect those in need of mental health services), and this may not be the same subgroup that has private insurance.

In this paper, we examine another health insurance mandate currently under consideration in many states — mandated insurance coverage for infertility treatment. As of 2006, fifteen states have enacted some form of infertility insurance mandate, and additional states have ongoing legislative advocacy efforts in this area. In previous work (Bitler and Schmidt, 2006), we analyze a nationally representative sample of women from the National Survey of Family Growth (NSFG) and find no effect of the mandates on the likelihood that a woman reported ever seeking infertility treatment. These findings — that mandates have no effect on utilization — are consistent with the existing evidence from the literature on mental health mandates. However, as suggested above, it may be important to define subgroups of the population that are most likely to be affected by mandates, and to analyze them separately. In this paper, we revisit the question to examine whether the mandates affect utilization for a specific subgroup in the population: those women who are both most likely to experience fertility problems and most likely to have insurance plans affected by the mandates. We find robust evidence that the mandates have a large and significant effect on utilization for exactly this subgroup.

II. Methodology and Data

We pool individual-level data from the 1982, 1988, 1995, and 2002 rounds of the NSFG to see whether there is more utilization of infertility treatment in states with infertility insurance mandates. Each wave of the NSFG surveys a nationally representative sample of women aged 15–44 on their fertility and marital histories. The NSFG is the only source of data on use of infertility treatment over the previous few decades. We merge information on state infertility insurance mandates to the NSFG data. Table 1 contains a list of states that have passed mandates, along with the year the mandate passed, and shows that there is considerable variation in both the timing of the mandates (ranging from 1977–2001) and in the types of states that have passed mandates (including both small and large states, as well as states from all U.S. regions).[2]

Previous work (Bitler and Schmidt, 2006) shows no effect of these mandates on utilization of infertility services for the population of women aged 15–44. However, in this paper, we revisit the issue by examining a particular subgroup where we would be most likely to detect an impact — highly educated older women.[3]

There are two reasons to expect effects for older, highly educated women. The first is related to demand for treatment. In order to desire treatment for infertility, one has to desire to become pregnant. Over the last several decades, increases in female labor force participation and educational attainment have been accompanied by delays in childbearing. The average age at first birth has risen from 21 years in 1970 to 25 in 2000 (Mathews and Hamilton, 2002), and differences in age at first birth by educational category have been even more striking. College-educated women are more likely to delay, in part to reduce the motherhood wage penalty associated with childbearing (e.g., Blackburn et al., 1993). As women wait longer before attempting to have children, the age at which fertility problems are discovered will rise.

In addition, age is associated with difficulty conceiving and carrying a pregnancy to term (Weinstein et al. 1990). Older women are significantly more likely to experience fertility problems and to seek help for these problems (Stephen and Chandra, 2000; Wright, Schieve, Reynolds, and Jeng, 2003). In 2002, women 30 and older accounted for almost 89% of all Assisted Reproductive Technology procedures (e.g., IVF) performed in the United States.

The second reason to expect any effects to be concentrated among older, highly educated women is that mandates generally only apply to persons with private health insurance. Older, highly educated women are more likely to have private coverage (either through their own employer or through a spouse's employer) than are other women. During calendar year 2002, 85% of women 30 and older with some college were covered by a private health insurance plan, while only 64% of women with at most a high school degree had such coverage.[4]

In addition, white women are also more likely to have coverage than other women. In the 2003 CPS, 87% of white women with some college 30 and older had private coverage during 2002, as compared to only 77% of non-white women with some college aged 30 and above. We therefore expect to find larger effects of the infertility insurance mandates among white women.[5]

Our key outcome variable for this analysis is an indicator for whether the woman has ever obtained infertility treatment.[6] Table 2 contains summary statistics for our outcome variable, for all women, and by age group (under 30 versus 30 and older) and completed education (no college versus at least some college). While about 14.5% of all women have ever obtained infertility treatment, treatment is more common among older women and among more highly educated women. Only about 8.2% of women under 30 have ever obtained treatment, compared with nearly 20% of women 30 or older.

There is also some evidence that the type of infertility visit varies by age (not shown in table). For women 30 and older who had an infertility visit in the 2002 NSFG, 73% sought help to get pregnant as opposed to help preventing a miscarriage (compared to 57% of women under 30), 34% ever were given ovulation inducing drugs (compared to 21% of women under 30), and fully 12% ever had artificial insemination or in vitro fertilization (as compared to 3.8% of women under 30). This suggests that the older women are disproportionately obtaining more expensive services.

Our empirical specification is as follows:

$$treatment_{ist} = \beta_1 * mandate_{st} + \beta_2 * age\ 30\ or\ older_{ist} + \beta_3 * some\ college_{ist} + \beta_4 * mandate_{st} * age\ 30\ or\ older_{ist} + \beta_5 * mandate_{st} * some\ college_{ist} + \beta_6 * some\ college_{ist} * age\ 30\ or\ older_{ist} + \beta_7 * mandate_{st} * age\ 30\ or\ older_{ist} * some\ college_{ist} + urban_{ist}\delta + X_{st}\alpha + \gamma_s + \nu_t + \varepsilon_{ist}.$$

For the reasons outlined above, we expect that the mandates will have the most impact on older, college educated women, since they are most likely to have private health insurance and most likely to have the greatest demand for fertility treatments.[7] Thus, our key estimated effect, β_7, is on the three-way interaction between the woman's state having an infertility insurance mandate; the woman's age being at least 30, and the woman having attained at least some college. We also control separately for mandate, age, and education effects; all two-way interactions between mandate, age, and education; as well as state and year fixed effects. Thus, in effect, we have estimated a differences-in-differences-in-differences specification.

We control for whether the woman lives in an urban area and for a host of time-varying state characteristics including the share black and Hispanic, the Medicaid eligibility threshold for a pregnant woman, the real maximum AFDC/TANF benefit for a family of 4, real median income for a family of 4, the unemployment rate and the employment growth rate, the share of the population under the Federal Poverty Level, and the share of births to unmarried women. These state-level controls have been found to be associated with fertility behavior in other work.

We estimate these regressions on the sample of women who have had sex and are past menarche. In addition to reporting results for all women, we also report results for the subsample of white women — our own previous work has shown that white women are much more likely to report having had such visits (Bitler and Schmidt, 2006) and the CPS tabulations provided earlier indicate that white women are more likely to be privately insured (see also Rhoades and Chu, 2000).

We estimate logistic regressions. We report both the underlying odds ratios (Table 3) as well as the relevant marginal effects (Table 4). We weight the data to be population representative, and report heteroskedasticity robust standard errors, with clustering at the state-by-year level. Since our model is nonlinear and our variables of interest are interactions, we focus on the marginal effects for our key variables, because coefficients on interaction effects in nonlinear models are not equal to the marginal effects of the interaction terms. As Ai and Norton (2003) point out, the magnitude of the marginal effects of the interaction depends on the value of the covariates in the model across the full sample and could even be of a different sign than the coefficient on the interaction term. The marginal effects are averaged over the full samples with each observation's *X*s (except those for the key coefficients) set to their actual values, and standard errors calculated via the delta method.

III. Results

Our key results are reported in Tables 3 and 4. The first column presents odds ratios (Table 3) or marginal effects (Table 4) from logistic regressions of the determinants of any infertility treatment for the sample of all women, while the second column presents the same values for the subsample of white women. The marginal effects for the main mandate variable in Table 4 replicates our earlier finding that the mandate itself has no statistically significant effect on reports of infertility treatment, either for the sample of all women or for the subsample of white women. In addition, the two-way interactions show no effect of the mandates on the subgroup of women aged 30 and older or on the subgroup of women who have attended at least some college.

However, the marginal effect for the three-way interaction of mandate, age at least 30, and education at least some college is a positive 0.046 (standard error 0.023) and is statistically significant at the 5% level. This suggests that for highly educated older women, living in a mandate state is associated with a 4.6 percentage point increase in the probability of ever having had an infertility visit. Given the higher rates of private insurance coverage for white women, we expect to find a larger marginal effect for the sample of white women, and do. For white women, the marginal effect is a positive and significant 0.061 (0.027) and is also significant at the 5% level. These magnitudes are large, given the pre-reform mean of around 15% percent of women who ever had such a visit. This suggests two things: first, that the mandates have an economically significant effect on utilization of infertility services; and second, that even mandates that have a large effect on a particular subgroup may have no discernable impact on the entire population.

We have verified that these results are robust to functional form by specifying the models as ordinary least squares, where the analogous coefficient for all women is 0.041 (significant at the 10% level), and that for white women is 0.059 (significant at the 5% level). We have checked that our results are not driven by endogeneity of the passage of mandates by including leads of the mandate variables in our specifications; these leads are not statistically significant.

IV. Discussion

Our results suggest that, unlike many other types of health insurance mandates, mandates related to coverage of infertility treatment are associated with an increase in utilization of services. However, these effects are only present among a subgroup of older, more-educated women.

There are several possible explanations for why these particular mandates seem to have an effect on utilization. First, infertility treatment may be less stigmatizing than other types of health services (e.g., mental health services), leading to a greater responsiveness by individuals to the state legislation. Second, despite the fact that there is little empirical evidence that mandates lead firms to self-insure, it is possible that the relatively low average costs of insurance coverage for infertility treatment lead to less of a response by companies in terms of their decisions to stop offering insurance coverage. Griffin and Panak (1998) estimate the insurance cost of covering infertility treatment at 0.4% of total medical costs per month (about $3.08 per month), compared to an average U.S. family premium per month of $771 (AHRQ, 2005). At the same time, for the women with high demand for treatment, insurance coverage may be quite valuable — a recent study suggested a median cost per live delivery resulting from IVF of $56,419 (Collins, 2001).

In addition, a larger share of firms and employees may be affected by this legislation than by some other kinds of mandates about insurance coverage. We use state-specific data from the 2003 Medical Expenditure Panel Survey-Insurance Component to estimate the share of private-sector employees with employer-provided insurance who could be affected by the mandates. We incorporate the type of insurance in which employees are enrolled (i.e., whether the plan is self-insured and ERISA exempt), as well as firm-size mandate exemptions in Illinois (firms under 25 employees) and Maryland (firms under 50 employees) in calculating the number of employees affected by the mandate. (The denominator for this calculation includes all employees enrolled in employer-sponsored insurance.)

Among states with mandates passed by 2000, we estimate that between 33 and 45% of private-sector employees with employer-provided insurance were enrolled in non-ERISA exempt insurance plans and not exempt from the mandates because of firm-size restrictions, and were therefore likely to be covered by the mandates. This translates into about 14–19% of the total number of private-sector employees that are enrolled in employer-provided insurance for the entire U.S. For comparison, Buchmueller et al. (2006) find that in 2003, only about 3% of private-sector employees enrolled in employer-sponsored health plans were covered by mental parity laws that applied to all mental illnesses.

Finally, 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 the mandate 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.

V. Conclusion

Evidence concerning the effect of various health insurance mandates suggests many such mandates have little impact on health care utilization. In this paper, we pool data from waves of the National Survey of Family Growth to see whether mandates for infertility treatment affect use of infertility treatment among women 15–44. Our own previous work has found such mandates have no effects for all women 15–44 (Bitler and Schmidt, 2006). Here, we find evidence that infertility treatment mandates do increase use of treatment, but only among highly educated older women, who have high demand for services and high rates of private insurance coverage. Since mandates are enacted to affect utilization of services and, ultimately, health outcomes, understanding why certain mandates affect these variables is important for policy. Additional research is necessary to disentangle these potential explanations.

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Table 1 State mandated infertility insurance

State	Year law enacted
Arkansas	1987[a]
California	1989[a]
Connecticut	1989
Hawaii	1987
Illinois	1991
Louisiana	2001
Maryland	1985
Massachusetts	1987
Montana	1987
New Jersey	2001
New York	1990[a]
Ohio	1991
Rhode Island	1989
Texas	1987
West Virginia	1977[a]

Note: Data come from Resolve (www.resolve.org) and state laws (see Appendix A of Schmidt, 2005).

[a] Arkansas, California, New York, and West Virginia first passed mandates in the years shown. These mandates were subsequently revised but remained in place.

Table 2 Summary statistics for ever having had infertility treatment, all women and by group, pooled NSFG data

Sample	Mean
All women	0.145
By age	
Under 30	0.082
At least 30	0.196
By completed education	
No college	0.127
Some college	0.167

Notes: Shown are weighted averages among women who have ever had sex after menarche for the variable ever having had an infertility visit. Value in first row is mean for all women, that in second row for women under 30, that in third row for women 30 and older, that in fourth row for women with no college attendance, and that in fifth row for women with some college. Data are from pooled 1982, 1988, 1995, and 2002 waves of the NSFG.

Table 3 Odds ratios, logistic regressions of determinants of having ever obtained infertility treatment as a function of state mandated infertility insurance

	All women		White women	
Any infertility mandate	1.049		1.108	
	(0.135)		(0.187)	
Age 30 or older	2.064	***	2.172	***
	(0.138)		(0.164)	
At least some college	0.821	**	0.795	**
	(0.073)		(0.081)	
Any mandate x age 30 or older	0.872		0.799	
	(0.120)		(0.135)	
Any mandate x at least some college	0.808		0.521	***
	(0.124)		(0.107)	
Age at least 30 * at least some college	1.717	***	1.698	***
	(0.180)		(0.208)	
Mandate x age 30 or older x at least some college	1.518	**	2.195	***
	(0.291)		(0.530)	
MSA	1.067		1.052	
	(0.063)		(0.066)	
White	1.356	***	--	
	(0.067)			
% Hispanic	0.976		0.985	
	(0.018)		(0.023)	
% Black	0.950		0.955	
	(0.041)		(0.052)	
Medicaid eligibility threshold	1.000		1.001	
	(0.001)		(0.001)	
Real maximum AFDC/TANF benefits for family of 4	0.972		0.949	
	(0.043)		(0.055)	
Real median income for family of 4	1.026	*	1.021	
	(0.015)		(0.019)	
Unemployment rate (/100)	8.054		0.246	
	(22.610)		(0.820)	
Employment growth rate	0.065		0.003	*
	(0.197)		(0.010)	
Share of population under FPL	0.228		0.204	
	(0.407)		(0.419)	
Share of births to unmarried women	2.592		0.910	
	(3.380)		(1.516)	
Observations	31,047		17,323	
Log-likelihood	-12,246.3		-7,201.67	
Psuedo R-squared	0.05		0.05	

Notes: Shown are odds ratios (standard errors) from logistic regressions of determinants of ever having had an infertility visit. Regressions are weighted, with standard errors clustered at the state-by-year level, and also include state and year of interview fixed effects. Data are from pooled 1982, 1988, 1995, and 2002 waves of the NSFG. Sample in first column is all women who have had sex post-menarche and sample in second column is all non-Hispanic white women who have ever had sex post-menarche. ***, **, and * denote that the underlying logistic regression coefficient is significantly different from zero at the 1, 5, and 10 percent levels.

Table 4 Marginal effects from logistic regressions of determinants of having ever obtained infertility treatment as a function of state mandated infertility insurance

	All women		White women	
Any infertility mandate	0.001		-0.008	
	(0.013)		(0.018)	
Age 30 or older	0.112	***	0.126	***
	(0.014)		(0.017)	
At least some college	0.027	**	0.020	**
	(0.011)		(0.013)	
Any mandate x age 30 or older	0.006		0.004	
	(0.015)		(0.020)	
Any mandate x at least some college	0.012		-0.006	
	(0.017)		(0.021)	
Age at least 30 * at least some college	0.079	***	0.082	***
	(0.015)		(0.018)	
Any mandate x age 30 or older x at least some college	0.046	**	0.061	**
	(0.023)		(0.027)	

Notes: Shown are the marginal effects (differences in predicted probabilities), with standard errors in parentheses, for various dummy variables from logistic regressions of determinants of ever having had an infertility visit (regressions in Table 3). Regressions are weighted, with standard errors clustered at the state-by-year level, and also include state and year of interview fixed effects. Data are from pooled 1982, 1988, 1995, and 2002 waves of the NSFG. Sample in first column is all women who have had sex post-menarche and sample in second column is all non-Hispanic white women who have ever had sex post-menarche. ***, **, and * denote that the delta method-calculated p-value is statistically significant at the 1, 5, and 10 percent levels. Marginal effects are averaged over all observations, with all Xs evaluated at their actual values except that relevant dummy variables were set to 0 or 1.

[1] One notable exception is the case of early postpartum discharge laws. Liu et al. (2004) find a positive significant effect of these laws on length of hospital stay. Evans and Wei (2006) find that these postpartum discharge laws may also have decreased the probability that infants were readmitted to the hospital.

[2] For additional detail on the mandates, see Schmidt (2005).

[3] A number of papers have considered fertility or health impacts of the insurance mandates (e.g., Schmidt, 2005; Bitler, 2006; Buckles, 2005; Bundorf, Henne, and Baker, 2005). Others have looked at one measure of utilization — cycles of in vitro fertilization or other advanced fertility treatments (e.g., Hamilton and McManus, 2004; Jain, Harlow, and Hornstein, 2002), but do so without the benefits of pre-mandate data on utilization that we use here. Additionally, our data measure all visits for infertility treatment rather than use of in vitro fertilization or other advanced techniques. In vitro fertilization is much less common than overall infertility treatment, thus effects of mandates on IVF likely could not be detected in survey data with typical sample sizes.

[4] Authors' tabulations from 2003 March Current Population Survey.

[5] In our previous work, we have found that black and Hispanic women are more likely to be infertile than white non-Hispanic women (Bitler and Schmidt, 2006). However, white non-Hispanic women are more likely to seek treatment even before the mandates were imposed.

[6] This variable includes women who sought help from a doctor to become pregnant or as well as those who sought help to prevent miscarriage. We cannot analyze these two states separately in the pooled regressions due to differences in the way the survey questions were asked in different waves of the NSFG.

[7] We cannot observe private insurance coverage in our data and likely would not want to use it as a control in any case, as it could conceivably respond to the mandates.