State Abortion Context and U.S. Women's Contraceptive Choices, 1995–2010

CONTEXT: The number of women in the United States exposed to restrictive abortion policies has increased substantially over the past decade. It is not well understood whether and how women adjust their contraceptive behavior when faced with restrictive abortion contexts.

METHODS: Data from 14,523 women aged 15–44 were drawn from the 1995 and 2010 cycles of the National Survey of Family Growth. A difference-in-differences approach was employed to examine the relationship between state-level changes in women's access to abortion and their contraceptive choices. Multinomial logistic regression analysis was used to determine the relative risk of using highly effective or less effective methods rather than no method for women exposed to varying levels of restrictive abortion contexts.

RESULTS: Women who lived in a state where abortion access was low were more likely than women living in a state with greater access to use highly effective contraceptives rather than no method (relative risk ratio, 1.4). Similarly, women in states characterized by high abortion hostility (i.e., states with four or more types of restrictive policies in place) were more likely to use highly effective methods than were women in states with less hostility (1.3). The transition to a more restrictive abortion context was not associated with women's contraceptive behavior, perhaps because states that introduced restrictive abortion legislation between 1995 and 2010 already had significant limitations in place.

CONCLUSION: To prevent unwanted pregnancies, it is important to ensure access to highly effective contraceptive methods when access to abortions is limited.

By Josephine Jacobs and Maria Stanfors

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The United States has one of the highest rates of unintended pregnancy in the developed world—an estimated 49% of all pregnancies are unintended (i.e., unwanted or mistimed).[1] Of these, 52% are due to contraceptive nonuse, and 48% to imperfect use. Thus, while more than half of all unintended pregnancies are the result of women's opting not to use a contraceptive, a significant proportion occur because women choose methods that require user compliance and, therefore, have relatively low effectiveness rates with typical use.[2] Given that unintended pregnancies can generate costs to society and have negative consequences for women's employment, children's health and families' welfare,[3] the federal government has made improved access to effective contraceptive methods a public health objective.[4]

A number of studies have examined variables that may be associated with a woman's choice of contraceptive, with a focus on how method cost may be correlated with contraceptive decisions.[5–7] Researchers have placed less emphasis on whether costs that may be faced in the future to avoid a birth are associated with current contraceptive choices. The relationship between state-regulated abortion laws or context and a woman's contraceptive choice, for instance, has rarely been examined, even though abortion and its associated costs become very relevant in the case of a contraceptive failure.[8]

In the past decade, in particular, there has been a substantial increase in the proportion of women of reproductive age who were exposed to restrictive abortion policies. Gold and Nash estimated that the proportion of women aged 15–44 and living in "hostile" abortion policy environments (i.e., states with at least four types of abortion restrictions) increased from 31% to 55% between 2000 and 2011.[9] The transition to a more restrictive abortion context can limit women's choices regarding abortion by increasing the financial, time and emotional costs of obtaining the procedure.[10] In light of the country's increasingly restrictive policy environments, exploring the association between state abortion context—and changes thereto—and contraceptive behavior should be of particular interest to policymakers. Yet evidence on this issue is scarce.

A handful of studies have used individual-level data to indirectly explore the role of abortion legislation on teenagers' decision to use any contraceptive; these studies showed mixed results. Levine found a positive association between state restrictions and teenage contraceptive use.[11] In contrast, neither Averett et al.[12] nor Sen[13] found statistically significant associations between parental consent laws or Medicaid funding restrictions and contraceptive use among teenagers. Another Levine study showed that Medicaid funding restrictions and mandatory waiting period laws were positively associated with contraceptive use, and that parental consent

laws were negatively associated with use, though the author did not consider these findings to be robust.[14]

Overall, abortion context has not been widely explored in the contraceptive choice literature, and when it has, the results have not been consistent or robust. In part, these inconsistencies have been attributed to the cross-sectional nature of the data used in some studies, which can impede conclusions about causality if a model does not control for unobservable beliefs and attitudes in a state's population.[13] Furthermore, the literature does not take into account that almost one-half of all unintended pregnancies occur among women who are using a contraceptive and that the majority of unintended pregnancies are among women 18 or older.[15] Only one study has explored the association between abortion context and the use of highly effective contraceptive methods among all women of reproductive age, but the author used aggregate-level data for a single year.[10] No study has attempted to incorporate state abortion context into an individual-level model of the contraceptive choices of all women of reproductive age.

In this article, we use individual-level data to explore the role that abortion context plays in the contraceptive choices of all women aged 15–44. We also examine whether the transition to a more restrictive abortion context is associated with contraceptive choices. We focus on the period 1995–2010, which witnessed a significant increase in the proportion of women exposed to restrictive abortion policies and contexts.[9]

CONCEPTUAL FRAMEWORK AND HYPOTHESES

Levine outlined a demand model for contraceptive use,[8] which we use as the basis for this study. This model assumes that women take all relevant economic considerations into account when making contraceptive choices, that contraceptives can be both costly and imperfect, and that they are used solely to prevent pregnancy. Knowing the costs and benefits associated with a birth, a woman can take measures to reduce the likelihood of pregnancy. Each additional measure that she takes (e.g., using more effective methods) is assumed to be increasingly costly. To avoid a birth, a woman can use a contraceptive and, in case of method failure, have an abortion. Thus, abortion acts like a form of insurance for method failure. If a woman chooses a less effective method, she will have an increased likelihood of experiencing an unwanted pregnancy and potentially facing an abortion and its associated costs. The model predicts that if abortions are available at a high cost, women may increase their efforts at avoiding pregnancy. Our first hypothesis is that a woman who lives in a state where abortion is not widely accessible—for instance, because of limited provider access or restrictive state abortion policies—will be more inclined to use a more effective contraceptive method than will a woman who lives in a state with greater accessibility. Moreover, if a state transitions to a more restrictive abortion

context, we hypothesize that women will adapt their behavior and choose more effective methods.

METHODS
Data
We used data from Cycles 5 and 7 of the National Survey of Family Growth (NSFG). The NSFG is a nationally representative survey of U.S. women aged 15–44 in the civilian, noninstitutionalized population. For the 1995 and 2006–2010 cycles,* 10,847 and 12,279 women, respectively, were interviewed. Sample weights may be used to correct for oversampling, nonresponse and noncoverage.[16]

We also used data from NARAL Pro-Choice America,[17,18] Henshaw,[19] and Jones and Kooistra[20] on macro-level abortion context indicators. The NARAL publications from 1995 to 2010 provided data on what abortion policies were in place each year. The other two publications provided state-level abortion access data for 1995–1996 and 2007–2008, respectively. The 2007–2008 data were used to capture provider access for individuals surveyed in the 2006–2010 NSFG cycle, as these were the most relevant data available. Use of access data for 2010–2011 did not change our results.

Our base sample consisted of women who were identified as being at risk for an unintended pregnancy and met the following eligibility criteria: They were currently not pregnant or seeking to get pregnant, they and their partners were not sterile (because of natural causes), and they reported having had heterosexual intercourse in the three months preceding the interview. This screening reduced our sample to 7,006 women from the 1995 cycle and 7,770 women from 2010. We also conducted alternate analyses that excluded women using nonreversible methods—that is, women who were, or whose partners were, sterile as a result of surgical procedures. These women were at very minimal risk of getting pregnant.[21,22]

Measures
•*Dependent variable.* Our dependent variable is contraceptive choice, and women are subdivided into three groups: those who used highly effective methods (surgical sterilization, pill, injectable, implant, patch, ring or IUD); those who used less effective methods (diaphragm, male or female condom, foam, cervical cap, sponge, suppository, jelly, cream, natural family planning, calendar rhythm, withdrawal, emergency contraception or other method); and those who used no method. Women were shown a list of options and asked which method they currently used (i.e., their "current contraceptive status"). In cases of multiple method use, we used the method that women identified as their main one. The rationale for our contraceptive categories follows from the theoretical model and research question at hand. With typical use, highly effective methods have effectiveness rates ranging from 92% to 99.9%.[1] These are also the most costly methods—at least initially— and cannot be obtained without consulting a physician.[22] The remaining methods range in effectiveness from 73% to

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*Technically, the latest data are from 2006–2010, but for simplicity, we refer to the year 2010.

85% with typical use;[1] these tend to be much less expensive and more easily available than the highly effective methods.* Finally, the annual rate of unintended pregnancy among sexually active women using no method of birth control is 85%.[1]

•*Independent variables.* Our key independent variables are state-level indicators that act as proxies for the state abortion context. We captured the context in three ways. First, we considered state-level provider access. Using data from Henshaw[19] and from Jones and Kooistra,[20] we determined the proportion of women in a state who had no abortion provider in their county in 1995–1996 and 2007–2008. In descriptive analyses, we subdivided women into three categories: women living in states with a low level of abortion access (i.e., at least 50% of women had no county provider), women living in states with medium access (i.e., 25–49% had no county provider) and women living in states with high access (i.e., 0–24% had no county provider). In multivariate analyses, we combined the medium and high access categories, because we did not observe significant variation between them over time, whereas we did between the medium and low categories. We would expect that having less access to abortion providers would result in the use of more effective contraceptives.

Second, we applied criteria from Gold and Nash[9] to construct an abortion hostility index for states. We identified eight categories of potential abortion restrictions: parental involvement laws, mandatory delay periods, Medicaid funding restrictions, laws requiring an extra premium for or prohibiting insurance coverage of abortions, unconstitutional bans on abortions throughout pregnancy, unconstitutional bans on abortions prior to fetal viability, mandated non–medically necessary ultrasounds prior to abortions and requirements that medication abortions be performed by physicians. Using data from NARAL Pro-Choice America,[17,18] we determined what types of restrictions each state had in 1995 and in 2006–2010. We categorized each state as less hostile, if it had two or fewer restrictions; moderately hostile, if it had three restrictions; or hostile, if it had four or more. In multivariate analyses, we combined the less hostile and moderately hostile categories, as limited variation was observed between them. There was a high correlation (rho=0.86) between our index and existing measures of abortion hostility (i.e., NARAL Pro-Choice America's annual letter grade[18]). We expected that living in a more hostile state would result in the use of more effective contraceptives.

Third, in line with previous work,[10] we captured the state abortion policy context by indicating whether a state had an enforceable mandatory delay law, parental consent law or Medicaid funding restriction in place. The last indicator was based on whether, in 1995 and 2010, there were Medicaid funding restrictions for abortion outside the federally mandated rape, incest or life endangerment requirements. We note, however, that while some states did not have these restrictions in place, very few abortions were funded through Medicaid over this time period.[23]

Therefore, in alternate specifications, these states were classified as having Medicaid restrictions. We expected that women in states with restricted abortion access would choose more effective contraceptive methods than would women in states with wider abortion access.

In all of our models, we accounted for a number of other variables that might be related to a woman's attitude toward births, abortions and contraceptive methods. We controlled for basic demographic characteristics, including age, race and ethnicity (Hispanic, white, black, other), having a husband or cohabiting partner† and number of births. Some of these characteristics could be correlated with a woman's degree of pregnancy ambivalence, which has been found to be associated with less contraceptive use.[24]

We also included a number of socioeconomic indicators, such as having health insurance (private, public, none), educational level (less than high school, completed high school, some college, completed college) and employment status (full-time, part-time, not working). Having health insurance and levels of education and employment have been found to have positive associations with contraceptive use.[5,13,25–27] Household income was inflation-adjusted to 2010 dollars and divided into four categories: less than $20,000, $20,000–39,999, $40,000–70,000 and more than $70,000. The expected association between income and contraceptive choice was not clear from the literature.[26,28,29] We controlled for degree of urbanization by using metropolitan statistical area categories of central city, other metropolitan and nonmetropolitan to capture women's potential access to contraceptive and abortion services.

To control for attitudinal variables, we considered women's degree of religiosity and religious affiliation. Religiosity was measured by the frequency of attendance at religious services (weekly, less than weekly, never); more frequent attendance might be expected to have a negative influence on contraceptive use, depending on religious affiliation, which was categorized as Protestant, Catholic, other or none. Finally, the number of sexual partners in the previous 12 months was included because research has indicated that women with multiple or casual partners may be more likely than others to use condoms or less likely to use the pill.[30–32]

Analysis

We studied differences in women's contraceptive choices to see which variables were associated with their choices and how this changed over time. The multiple-choice setting of highly effective contraceptives versus less effective methods or none called for use of a multinomial logit model that applies individual-specific data and state identifiers.

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*While most of these methods can be obtained over the counter in the United States, the cervical cap and diaphragm require a health care professional for fitting. Their cost and effectiveness, however, are still lower than those of the methods categorized as highly effective.

†We combined married and cohabiting women in our analyses, and while we acknowledge that there may be differences regarding union stability, this variable was used to control for living with a partner.

Our empirical model was one of optional outcomes of contraceptive use. The model contains a set of explanatory variables, comprising individual characteristics, indicators of state-level abortion access, a time trend and state group fixed effects. We included a time trend to control for factors that change over time. The state group fixed effects took into consideration that states may differ regarding important unobserved state-level determinants of contraceptive use. For example, if a general liberal attitude or progressive social policy climate is positively associated with wider abortion access and wider use of highly effective contraceptives, excluding these variables may bias the results upward. Because certain states had limited observations, they were combined into groups with nearby states that had similar abortion, political and attitudinal characteristics.* The association between individual characteristics and contraceptive choice was assessed with and without state group fixed effects.

Our models also include interaction terms between abortion context and time. We employed a model similar to a difference-in-differences approach that exploited change over time and accounted for unobserved heterogeneity between the treated and comparison groups.†[33] We multiplied the policy variable by a dummy variable indicating that the individual was from the second (i.e., 2006–2010) wave of data. For the abortion hostility index, abortion access and Medicaid funding restriction variables, the state categorizations did not change between 2006 and 2010. However, some changes in mandatory delay and parental consent laws occurred over this time period, and in these cases, only women who were exposed to new abortion laws were categorized as such. For instance, if a state law became effective in January 2008, women in that state who were interviewed in 2006 and 2007 would not be coded as being exposed to that law, while women interviewed from 2008 to 2010 would be coded as having been exposed. The association between a change in abortion policy context and contraceptive choice can be determined by the interaction term (i.e., abortion context multiplied by the 2010 dummy). We predict that reduced access to abortion meaningfully increases women's use of highly effective contraceptive methods in states affected by changes in provider access or policy. It is important to note that the estimated effects of the access and policy changes will be based only

TABLE 1. Number of states that transitioned to more restrictive abortion contexts over the period 1995–2010

Type of transition	No.
Reduced abortion provider access	10
Increased abortion hostility	9
Introduced mandatory delay law	17
Introduced parental consent law	14
Introduced Medicaid funding restriction	3†

†Includes the District of Columbia. *Notes:* Restrictive abortion contexts are defined in the Methods section of the text. See Appendix Table 1 (Supporting Information) for details. *Sources:* references 17–20.

on states that changed over the time period in question. Therefore, variation in a number of states is necessary for a robust analysis.

The models we present combine the 1995 and 2010 waves. We also ran a number of regressions with each individual wave; however, to address our main research question, it was necessary to include multiple waves to determine within-state effects over time by exploiting changes in abortion context.

Because we used complex sample survey data, the statistical estimation technique requires the use of weights to properly compute regression coefficients. Further adjustments were required for the estimation of standard errors, which used the Huber-White sandwich estimator to correct for heteroskedasticity. All results are presented as relative risk ratios. We compared each of the first two contraceptive choice alternatives—highly effective and less effective methods—with the base category of no method. The relative risk ratio indicates how the probability of choosing a given alternative, as opposed to no method, changes if the independent variable is increased by one unit. It represents the ratio of two relative risks—in this case, the risk of using a method, rather than no method, for those subject to abortion restrictions relative to the same risk for those without restrictions. For our study, it is likely that a woman would make a contraceptive choice by comparing the alternatives available to her. Furthermore, a base category of no method is likely to factor into her decision-making process, as it represents an extreme alternative in terms of cost and effectiveness.

We conducted a number of sensitivity analyses. We ran models to test whether Medicaid funding restrictions were associated with the contraceptive choices of women receiving Medicaid, and whether parental consent laws were associated with the choices of women younger than 18. Following Medoff,[10] we ran all regressions for two age-groups (18–25 and 26–44). We also altered the hostility index definitions of states to determine if altering these thresholds changed our results. Finally, we constructed regression models that altered our dependent variable. First, we ran all regressions including women using only reversible methods (i.e., excluding female sterilization). Next, we ran specifications in which the use of long-acting reversible contraceptives (i.e., the implant or IUD) and of condoms were separate categories, as these methods may differ regarding cost and utility for disease prevention.

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*These combined groups were Idaho, Nebraska, North Dakota and South Dakota; Connecticut, Delaware, Maine, Maryland, New Hampshire, Rhode Island and Vermont; Alaska and Montana; Indiana and Minnesota; Oregon and Washington; Hawaii, Nevada and New Mexico; Colorado, Utah and Wyoming; Arkansas, Kansas, Kentucky, Missouri and Oklahoma; Alabama, Georgia, Mississippi and South Carolina; North Carolina and West Virginia; and Florida and Washington, DC.

†Strictly speaking, a difference-in-differences approach identifies changes in behavior between states that are treated and those that are not. In some cases, it implies that all states are untreated in the first time period, but in our analysis, some states already had restrictive abortion policies and contexts in place at that time. This has, however, no bearing on the estimation of our models, since our analysis focuses on change versus no change.

RESULTS

Descriptive Findings

Nine states transitioned from having medium abortion provider access to having low access over the study period, while one state (Florida) transitioned from high to medium access (Table 1). Nineteen states had low provider access in 1995–1996, while 27 states had low access by 2007–2008. A single state (Ohio) transitioned from low to high access, and two states went from medium to high access. In the nine states that crossed the low access threshold, 39–49% of women lived in a county with no provider in 1995. A more detailed state-by-state summary of these changes can be found in Appendix Table 1 (Supporting Information).

In 1995, 12 states and the District of Columbia were less hostile to abortion, 13 states were moderately hostile and 25 states were hostile. By 2010, 13 states and the District of Columbia were less hostile, seven states were moderately hostile and 30 states were hostile. Overall, two states

transitioned from less hostile to moderately hostile over the study period, and seven states transitioned to having a hostile context (four or more restrictions). Six of the latter states, however, already had three restrictions in place in 1995.

Only seven states had mandatory delay laws in 1995. By 2010, however, an additional 17 states had introduced enforceable delay laws. Twenty-six states and the District of Columbia did not have parental consent or notice laws in 1995, and 14 states introduced such laws by 2010.

Finally, two states (Idaho and North Carolina) and the District of Columbia introduced Medicaid funding restrictions over the study period; however, two other states (Montana and Arizona) no longer had these restrictions in place by 2010. Hence, little variation was found in Medicaid restrictions. In total, 17 states and the District of Columbia did not have such restrictions in 1995, while 17 states did not have them in 2010.

TABLE 2. Selected characteristics of women aged 15–44 who were at risk for unintended pregnancy, and their state abortion contexts, by survey year, National Survey of Family Growth

Characteristic	1995 (N=6,945)	2010 (N=7,578)	Characteristic	1995 (N=6,945)	2010 (N=7,578)
PERCENTAGE DISTRIBUTIONS			PERCENTAGE DISTRIBUTIONS		
Contraceptive method†			**Educational level**		
Highly effective method	61 (0.01)	65 (0.01)	<high school	17 (0.01)	20 (0.01)
Less effective method	30 (0.01)	23 (0.01)	Completed high school	38 (0.01)	26 (0.01)
None	8 (0.01)	12 (0.01)	Some college	19 (0.01)	21 (0.01)
			Completed college	27 (0.01)	33 (0.01)
Abortion provider access					
High	40 (0.02)	36 (0.03)	**Employment status**		
Medium	35 (0.02)	23 (0.03)	Full-time	45 (0.01)	46 (0.01)
Low	24 (0.02)	42 (0.03)	Part-time	20 (0.01)	24 (0.01)
			Not working	36 (0.01)	30 (0.01)
Abortion hostility index					
Less hostile	19 (0.02)	28 (0.02)	**Household income (in 2010 $)**		
Moderately hostile	32 (0.02)	7 (0.02)	<$20,000	25 (0.01)	24 (0.01)
Hostile	49 (0.02)	65 (0.03)	$20,000–39,999	30 (0.01)	27 (0.01)
			$40,000–70,000	29 (0.01)	28 (0.01)
Mandatory delay law			>$70,000	16 (0.01)	21 (0.01)
Yes	12 (0.02)	55 (0.03)			
No	88 (0.02)	45 (0.03)	**Metropolitan residence**		
			Central city	49 (0.01)	47 (0.02)
Parental consent law			Other metropolitan	21 (0.01)	21 (0.02)
Yes	42 (0.02)	70 (0.02)	Nonmetropolitan	31 (0.01)	32 (0.02)
No	58 (0.02)	30 (0.02)			
			Attendance at religious services		
Medicaid funding restrictions			Weekly	31 (0.01)	28 (0.01)
Yes	58 (0.02)	65 (0.03)	<weekly	45 (0.01)	49 (0.01)
No	42 (0.02)	35 (0.03)	Never	24 (0.01)	24 (0.01)
Race/ethnicity			**Religious affiliation**		
Hispanic	11 (0.01)	17 (0.02)	Protestant	52 (0.01)	48 (0.01)
White	70 (0.01)	62 (0.02)	Catholic	29 (0.01)	25 (0.01)
Black	15 (0.01)	15 (0.01)	Other	5 (0.01)	8 (0.01)
Other	4 (0.01)	7 (0.01)	None	13 (0.01)	19 (0.01)
Has husband/cohabiting partner			Total	100	100
Yes	62 (0.01)	60 (0.01)			
No	38 (0.01)	40 (0.01)	MEANS		
			No. of births	1.3 (0.02)	1.3 (0.01)
Insurance			**Age**	30.8 (0.12)	30.7 (0.19)
Private	71 (0.01)	66 (0.01)	**No. of partners in past 12 months**	1.2 (0.01)	1.2 (0.01)
Public	17 (0.01)	19 (0.01)			
None	12 (0.01)	15 (0.01)			

†Highly effective methods are surgical sterilization, pill, injectable, implant, patch, ring and IUD; less effective methods are the diaphragm, male and female condom, foam, cervical cap, sponge, suppository, jelly, cream, natural family planning, calendar rhythm, withdrawal, emergency contraception and other methods. *Notes:* All percentages are weighted. Figures in parentheses are standard errors. Percentages may not total 100 because of rounding.

We had complete information for 6,945 women in 1995. The majority were contraceptive users; 61% used highly effective methods (Table 2). Only 12% were subject to mandatory delay laws, while 42% lived in states with parental consent laws. The majority of women (58%) lived in states where there were Medicaid funding restrictions. Forty-nine percent lived in the most hostile states, while 24% lived in states with low abortion access.

We had full information for 7,578 women from the 2010 sample. Most women were using highly effective contraceptive methods (65%). By 2010, the majority were exposed to mandatory delay laws (55%), parental consent laws (70%) and Medicaid funding restrictions (65%). Sixty-five

TABLE 3. Relative risk ratios (and standard errors) from multinomial logit regression analyses assessing associations between state-level abortion provider access (and selected characteristics of women) and women's use of a highly effective or less effective contraceptive method, rather than no method, in models with and without state fixed effects

Characteristic	Excluding state fixed effects		Including state fixed effects	
	Highly effective	Less effective	Highly effective	Less effective
Abortion provider access				
High/medium access (ref)	1.00	1.00	1.00	1.00
Low access	1.35 (0.17)*	0.94 (0.12)	1.02 (0.19)	0.69 (0.15)†
Low access x 2010	1.07 (0.19)	1.25 (0.24)	1.13 (0.21)	1.37 (0.29)
Survey year				
1995 (ref)	1.00	1.00	1.00	1.00
2010	0.64 (0.06)***	0.44 (0.04)***	0.66 (0.06)***	0.45 (0.05)***
Age				
14–17 (ref)	1.00	1.00	1.00	1.00
18–24	1.67 (0.30)**	1.29 (0.27)	1.67 (0.30)**	1.27 (0.27)
25–34	1.50 (0.29)*	1.16 (0.26)	1.48 (0.29)*	1.13 (0.26)
35–44	1.10 (0.21)	0.82 (0.19)	1.09 (0.21)	0.80 (0.19)
Race/ethnicity				
White (ref)	1.00	1.00	1.00	1.00
Hispanic	1.06 (0.12)	1.14 (0.14)	1.03 (0.12)	1.10 (0.15)
Black	0.72 (0.07)***	0.80 (0.08)*	0.71 (0.07)***	0.77 (0.08)*
Other	0.72 (0.11)*	1.59 (0.31)*	0.76 (0.12)†	1.64 (0.33)*
Has husband/cohabiting partner				
Yes	1.40 (0.13)***	1.94 (0.20)***	1.38 (0.13)***	1.95 (0.20)**
No (ref)	1.00	1.00	1.00	1.00
No. of births				
0 (ref)	1.00	1.00	1.00	1.00
1	1.00 (0.11)	1.02 (0.12)	1.01 (0.11)	1.05 (0.13)
2	2.58 (0.35)***	1.47 (0.20)**	2.54 (0.35)***	1.48 (0.21)**
≥3	3.91 (0.60)***	1.30 (0.22)	3.97 (0.61)***	1.34 (0.23)†
Insurance				
Private	1.41 (0.16)**	0.99 (0.13)	1.42 (0.16)**	1.00 (0.13)
Public	1.30 (0.16)*	0.86 (0.12)	1.35 (0.17)*	0.87 (0.13)
None (ref)	1.00	1.00	1.00	1.00
Educational level				
<high school (ref)	1.00	1.00	1.00	1.00
Completed high school	1.02 (0.12)	0.93 (0.13)	1.03 (0.12)	0.94 (0.13)
Some college	1.35 (0.18)*	1.46 (0.21)**	1.37 (0.18)*	1.49 (0.22)**
Completed college	1.39 (0.20)*	1.84 (0.28)***	1.41 (0.21)*	1.83 (0.28)***
Employment status				
Full-time	1.41 (0.14)***	1.06 (0.11)	1.41 (0.14)***	1.06 (0.11)
Part-time	1.25 (0.14)*	1.09 (0.13)	1.25 (0.14)*	1.08 (0.13)
Not working (ref)	1.00	1.00	1.00	1.00
Attendance at religious services				
Weekly (ref)	1.00	1.00	1.00	1.00
<weekly	1.18 (0.11)†	1.03 (0.10)	1.21 (0.12)*	1.03 (0.10)
Never	1.17 (0.14)	1.13 (0.15)	1.20 (0.14)	1.13 (0.15)
No. of sexual partners	0.80 (0.04)***	1.14 (0.06)*	0.80 (0.04)***	1.14 (0.06)**
Constant	*2.13 (0.51)**	*1.14 (0.33)*	*1.71 (0.61)*	*1.04 (0.53)*

*p<.05. **p<.01. ***p<.001. †p<.10. *Notes:* Regression models also controlled for household income, metropolitan residence and religious affiliation. ref= reference group.

percent lived in the most hostile states, and 42% lived in states with low abortion access.

In both samples, the mean age was 31. The proportion of women who were Hispanic was 11% in 1995 and 17% in 2010, while the proportions who were white were 70% and 62% in these years, respectively. At both survey dates, most women had a husband or cohabiting partner, and their average number of births was 1.3. The majority of women in both waves were privately insured, but 12% in 1995 and 15% in 2010 had no insurance coverage.

TABLE 4. Relative risk ratios (and standard errors) from multinomial logit regression analyses assessing associations between state-level abortion hostility index (and selected characteristics of women) and women's use of a highly effective or less effective contraceptive method, rather than no method, in models with and without state fixed effects

Characteristic	Excluding state fixed effects		Including state fixed effects	
	Highly effective	Less effective	Highly effective	Less effective
Abortion hostility index				
Less/moderately hostile (ref)	1.00	1.00	1.00	1.00
Hostile	1.32 (0.14)**	0.98 (0.10)	1.00 (0.18)	0.82 (0.18)
Hostile x 2010	0.92 (0.14)	1.06 (0.17)	0.95 (0.15)	1.12 (0.20)
Survey year				
1995 (ref)	1.00	1.00	1.00	1.00
2010	0.70 (0.07)***	0.46 (0.05)***	0.71 (0.08)***	0.46 (0.06)***
Age				
14–17 (ref)	1.00	1.00	1.00	1.00
18–24	1.67 (0.30)**	1.28 (0.27)	1.67 (0.30)**	1.27 (0.27)
25–34	1.50 (0.29)*	1.15 (0.26)	1.49 (0.29)*	1.13 (0.26)
35–44	1.09 (0.21)	0.81 (0.19)	1.09 (0.21)	0.80 (0.19)
Race/ethnicity				
White (ref)	1.00	1.00	1.00	1.00
Hispanic	1.04 (0.11)	1.12 (0.14)	1.03 (0.12)	1.11 (0.15)
Black	0.73 (0.07)**	0.80 (0.08)*	0.71 (0.07)***	0.77 (0.08)*
Other	0.74 (0.11)*	1.60 (0.31)*	0.76 (0.12)†	1.65 (0.33)*
Has husband/cohabiting partner				
Yes	1.41 (0.13)***	1.94 (0.20)***	1.39 (0.13)***	1.94 (0.20)***
No (ref)	1.00	1.00	1.00	1.00
No. of births				
0 (ref)	1.00	1.00	1.00	1.00
1	1.00 (0.11)	1.03 (0.13)	1.01 (0.11)	1.05 (0.13)
2	2.55 (0.35)***	1.47 (0.20)**	2.55 (0.35)***	1.48 (0.21)**
≥3	3.87 (0.59)***	1.31 (0.22)	3.97 (0.61)***	1.34 (0.23)†
Insurance				
Private	1.40 (0.16)**	0.99 (0.13)	1.42 (0.16)**	1.00 (0.13)
Public	1.28 (0.16)*	0.86 (0.12)	1.35 (0.17)*	0.87 (0.13)
None (ref)	1.00	1.00	1.00	1.00
Educational level				
<high school (ref)	1.00	1.00	1.00	1.00
Completed high school	1.02 (0.12)	0.93 (0.13)	1.03 (0.12)	0.94 (0.13)
Some college	1.35 (0.18)*	1.47 (0.22)**	1.37 (0.18)*	1.49 (0.22)**
Completed college	1.39 (0.20)*	1.84 (0.28)***	1.41 (0.20)*	1.83 (0.28)***
Employment status				
Full-time	1.41 (0.14)***	1.05 (0.11)	1.40 (0.14)***	1.06 (0.12)
Part-time	1.26 (0.14)*	1.09 (0.13)	1.25 (0.14)*	1.08 (0.13)
Not working (ref)	1.00	1.00	1.00	1.00
Metropolitan residence				
Central city	0.69 (0.08)***	1.04 (0.14)	0.73 (0.09)**	1.03 (0.15)
Other metropolitan	0.77 (0.08)*	1.20 (0.15)	0.83 (0.09)†	1.17 (0.15)
Nonmetropolitan (ref)	1.00	1.00	1.00	1.00
Attendance at religious services				
Weekly (ref)	1.00	1.00	1.00	1.00
<weekly	1.18 (0.11)†	1.03 (0.10)	1.20 (0.12)†	1.03 (0.10)
Never	1.17 (0.14)	1.13 (0.15)	1.20 (0.14)	1.14 (0.15)
No. of sexual partners	0.80 (0.04)***	1.14 (0.06)*	0.80 (0.04)***	1.14 (0.06)**
Constant	*2.13 (0.53)** *	*1.17 (0.34)*	*1.70 (0.60)*	*1.06 (0.52)*

*p<.05. **p<.01. ***p<.001. †p<.10. *Notes:* Regression models also controlled for household income and religious affiliation. ref=reference group.

TABLE 5. Relative risk ratios (and standard errors) from multinomial logit regression analyses assessing associations between state-level restrictive abortion policies (and selected characteristics of women) and women's use of a highly effective or less effective contraceptive method, rather than no method, in models with and without state fixed effects

Characteristic	Excluding state fixed effects		Including state fixed effects	
	Highly effective	Less effective	Highly effective	Less effective
Restrictive abortion policies				
No policy (ref)	1.00	1.00	1.00	1.00
Mandatory delay law	0.78 (0.15)	1.06 (0.20)	1.06 (0.29)	1.81 (0.54)*
Parental consent law	1.07 (0.14)	1.08 (0.15)	0.84 (0.14)	0.92 (0.17)
Medicaid funding restriction	1.38 (0.16)**	0.87 (0.11)	1.13 (0.19)	0.76 (0.16)
Mandatory delay law x 2010	1.90 (0.47)**	1.25 (0.34)	1.24 (0.38)	0.71 (0.25)
Parental consent law x 2010	0.96 (0.21)	0.83 (0.21)	0.93 (0.22)	0.77 (0.19)
Medicaid funding restriction x 2010	0.63 (0.13)*	1.05 (0.24)	0.76 (0.17)	1.31 (0.30)
Survey year				
1995 (ref)	1.00	1.00	1.00	1.00
2010	0.73 (0.08)**	0.46 (0.06)***	0.78 (0.09)*	0.48 (0.06)***
Age				
14–17 (ref)	1.00	1.00	1.00	1.00
18–24	1.68 (0.30)**	1.27 (0.27)	1.68 (0.31)**	1.27 (0.27)
25–34	1.52 (0.30)*	1.14 (0.26)	1.49 (0.29)*	1.13 (0.26)
35–44	1.10 (0.22)	0.80 (0.19)	1.09 (0.21)	0.80 (0.19)
Race/ethnicity				
White (ref)	1.00	1.00	1.00	1.00
Hispanic	1.03 (0.11)	1.13 (0.14)	1.03 (0.12)	1.09 (0.15)
Black	0.70 (0.07)***	0.79 (0.08)*	0.71 (0.07)***	0.77 (0.08)
Other	0.74 (0.11)*	1.60 (0.31)*	0.77 (0.13)†	1.64 (0.33)*
Has husband/cohabiting partner				
Yes	1.41 (0.13)***	1.95 (0.20)***	1.39 (0.13)***	1.95 (0.20)***
No (ref)	1.00	1.00	1.00	1.00
No. of births				
0 (ref)	1.00	1.00	1.00	1.00
1	0.99 (0.11)	1.03 (0.13)	1.01 (0.12)	1.05 (0.13)
2	2.53 (0.35)***	1.47 (0.21)**	2.55 (0.35)***	1.47 (0.21)**
≥3	3.84 (0.59)***	1.30 (0.22)	3.97 (0.61)***	1.33 (0.23)
Insurance				
Private	1.40 (0.16)**	0.99 (0.13)	1.42 (0.16)**	1.00 (0.13)
Public	1.30 (0.16)*	0.85 (0.12)	1.34 (0.17)**	0.87 (0.12)
None (ref)	1.00	1.00	1.00	1.00
Educational level				
<high school (ref)	1.00	1.00	1.00	1.00
Completed high school	1.02 (0.12)	0.93 (0.13)**	1.03 (0.12)	0.93 (0.12)
Some college	1.35 (0.18)*	1.46 (0.21)**	1.36 (0.18)*	1.47 (0.21)**
Completed college	1.39 (0.20)*	1.83 (0.28)***	1.41 (0.20)*	1.81 (0.27)***
Employment status				
Full-time	1.41 (0.14)***	1.06 (0.12)	1.40 (0.14)***	1.07 (0.12)
Part-time	1.26 (0.14)*	1.09 (0.13)	1.26 (0.14)*	1.08 (0.13)
Not working (ref)	1.00	1.00	1.00	1.00
Metropolitan residence				
Central city	0.72 (0.08)**	1.07 (0.14)	0.74 (0.09)*	1.05 (0.15)
Other metropolitan	0.80 (0.09)*	1.22 (0.16)	0.84 (0.09)	1.20 (0.16)
Nonmetropolitan (ref)	1.00	1.00	1.00	1.00
Attendance at religious services				
Weekly (ref)	1.00	1.00	1.00	1.00
<weekly	1.18 (0.11)†	1.02 (0.10)	1.20 (0.11)†	1.03 (0.10)
Never	1.17 (0.14)	1.12 (0.15)	1.19 (0.14)	1.13 (0.15)
No. of sexual partners	0.80 (0.04)***	1.14 (0.06)*	0.80 (0.04)***	1.15 (0.06)**
Constant	*1.95 (0.49)**	*1.19 (0.35)*	*1.63 (0.60)*	*1.39 (0.75)*

*p<.05. **p<.01. ***p<.001. †p<.10. *Notes:* Regression models also controlled for household income and religious affiliation. ref=reference group.

Sixty-five percent of women were working full-time or part-time at the beginning of the study period, while 70% were doing so by the end.

Multivariate Findings

In multivariate analysis, women living in states with a low level of provider access were more likely than women in states with greater provider access to use highly effective contraceptives rather than no method (relative risk ratio, 1.4—Table 3). However, when state fixed effects were controlled for, the coefficient declined in magnitude and lost significance. In models including the interaction between low access and the 2010 dummy variable, both with and without state fixed effects, no associations were found, indicating that women in states where providers became less accessible over time did not change their contraceptive behavior.

In regression analysis that assessed abortion hostility while excluding state fixed effects, a positive association was found between the use of highly effective contraceptives and hostility level (Table 4). Compared with women living in a less or moderately hostile state, women in a hostile state had an elevated risk of using highly effective contraceptives as opposed to no method (relative risk ratio, 1.3). (When we separately tested the less and moderately hostile categories, the results were the same.) In the model that included state fixed effects, this coefficient lost significance. For states that had transitioned to a hostile abortion policy context, no associations were found for the use of either highly effective or less effective methods.

In analyses that assessed other restrictive abortion policies, we found a positive association between Medicaid funding restrictions and women's use of highly effective contraceptives rather than no method (relative risk ratio, 1.4); in the state fixed effects model, however, the coefficient was not significant (Table 5). No associations were found for parental consent laws, while the presence of a mandatory delay law was correlated with the use of less effective methods, as opposed to none, in the state fixed effects model (1.8). When we interacted each policy variable with the year, we found that while the interactions with the delay law and Medicaid restrictions were significant for women's use of highly effective methods (1.9 and 0.6, respectively), these coefficients were not significant in models that included state fixed effects.

We found several noteworthy results in the multivariate analyses that included state group fixed effects, which controlled for state-specific factors that do not vary over time. Women aged 18–24 or 25–34 were more likely than younger teenagers to use highly effective methods rather than no method (relative risk ratios, 1.5–1.7), and blacks were less likely than whites to use such a method instead of none (0.7). Compared with women who did not have a husband or cohabiting partner, those who had one were more likely to use highly effective methods and less effective methods as opposed to none (1.4 and 2.0, respectively).

Across all of our state fixed effects regression models, women who had had two or more births were more likely than those with no births to use highly effective contraceptives instead of none (relative risk ratios, 2.5–4.0), and women with private or public health insurance were more likely than those with no insurance to do so (1.3–1.4). Women who had at least some college education were more likely than those with less than a high school degree to use a highly effective method rather than no method (1.4), and women who worked full-time or part-time were more likely than nonworking women to use these methods (1.3–1.4). Household income was not associated with contraceptive use. (In alternative specifications, we ran stepwise regressions that excluded collinear socioeconomic variables and did not find that household income became significant.) In the abortion access regression model, women who attended religious services less than weekly had a greater likelihood of using a highly effective method than did women who attended weekly (1.2). Finally, the greater the number of women's sexual partners in the last 12 months, the less likely they were to use highly effective methods (0.8), and the more likely they were to use less effective ones (mostly male condoms; 1.1–1.2), instead of none. (We note the potential for reverse causality and therefore a biased coefficient for the number of sexual partners.)

The results of our sensitivity analyses yielded some notable findings (Appendix Tables 2–4, Supporting Information). In the regressions exploring our base case models, we found that 18–25-year-olds in states with low abortion access and in states that had transitioned to a mandatory delay law were more likely than their counterparts to use highly effective methods as opposed to none (relative risk ratios, 2.0 and 3.3, respectively). Meanwhile, the relative risk ratio for women exposed to Medicaid funding restrictions was elevated only in the older age-group (1.8). However, with the addition of state fixed effects, these associations lost significance (not shown).

When we altered the definition of abortion hostility categories and the number of categories in the index (i.e., ranging from two to four), we consistently found that having four or more restrictions in place was positively associated with hormonal contraceptive use. We saw a smaller but still significant association when we used a threshold of at least three restrictions (relative risk ratios, 1.27 vs. 1.33). In models that excluded female sterilization, we did not find significant differences compared with our base case results. Finally, in models where the dependent variable had separate categories for long-acting reversible contraceptives and condoms, we found no differences between condoms and other less effective methods, or between long-acting reversible contraceptives and other hormonal methods.

DISCUSSION

In the foregoing analysis, we attempted to determine whether state abortion context, as measured through provider access, overall legislative hostility and restrictive abortion policies, was associated with the contraceptive

Contraceptive choice [seems to be] most strongly influenced by individual-level variables.

choices of women of reproductive age in the United States. While provider access and legislative hostility appeared to be associated with the use of highly effective contraceptives, the inclusion of state group fixed effects rendered these associations insignificant. This is an indication that the associations we identified may be attributable to other time-invariant state-level variables—for example, other reproductive health policies or women's attitudes.

Furthermore, the transition to less accessible and more hostile abortion policy contexts was not associated with the use of highly effective methods. This finding would, at first glance, appear to disprove our hypothesis that women respond to an increase in the cost (monetary, time or other) of abortions by using more effective contraceptives. However, a more nuanced analysis of which states decreased accessibility and introduced restrictive legislation highlights another possible explanation for these findings. As noted earlier, the majority of the states that transitioned to less accessible and more hostile legislative abortion contexts already had relatively low access and high hostility. The nine states that crossed the low access threshold by 2008 were relatively close to this threshold in 1995. Likewise, six of the seven states that transitioned to the hostile category by 2010 already were close to it in 1995. Thus, women in these states may have already adjusted their contraceptive behavior to fairly restrictive abortion environments.

Even a cursory overview of the states that most frequently made antiabortion headlines in 2013 supports this idea. In that year, North Dakota passed a previability ban that criminalized abortions after a fetal heartbeat can be detected, Oklahoma repeatedly attempted to restrict medication abortions and State Senator Wendy Davis held an 11-hour filibuster to block laws that would close a number of abortion clinics in Texas. These and many other antiabortion stories took place in states that already had highly hostile legislative abortion contexts. The long-term debates over restrictive abortion laws that precede a policy change likely lead to a general public awareness of the issue and the potential policy changes. Concerned women may come to expect such changes and may take these future changes into consideration when choosing a contraceptive. Of course, short-term responses to changes in abortion context may have showed up only in the year or two following these changes, and we may not have captured these responses over the more extensive time period between survey waves.

It is also important to consider that the number of states that transitioned to more restrictive abortion contexts is relatively limited. Our approach was reliant on a high degree of variation in the abortion context variables across the two waves. Some of our context measures varied little, which provides an alternative explanation for the nonsignificance of our difference-in-differences results.

Earlier studies that examined similar questions found no association between policies (except parental consent laws) and teenage contraceptive use.[11–14] Overall, these studies found that the coefficients were often inconsistent and not robust after the addition of state fixed effects. We identified similar inconsistencies when the population being considered was extended beyond teenagers, and when more comprehensive indicators for state abortion context were considered.

Our results seem to indicate that contraceptive choice is most strongly influenced by individual-level variables. The fact that the coefficients for individual-level characteristics were highly robust, irrespective of the model, highlights the importance of the relationship between individual attributes and contraceptive choice. States differ regarding general attitudes toward family planning and sex, as well as gender roles, family matters and social progressiveness; however, individual characteristics may be more important in informing a woman's choice of contraceptive method, irrespective of the larger abortion context.

Limitations

A number of limitations should be kept in mind when interpreting our results. First, there is potential bias from endogeneity, which makes it difficult to identify the impact of certain state-level attributes and policy factors on contraceptive choice and on each other. Specifically, if there were time-varying state-level characteristics associated with women's contraceptive choices that we did not capture (e.g., level of religiosity or other reproductive health regulation), this is a potential source of bias. In the absence of a valid instrument, it is difficult to remedy this. However, we employed a residual inclusion technique. In the first equations, we predicted the likelihood of a state's having a restrictive abortion context by using variables indicating whether a state's legislative bodies had a Republican majority or at least 20% representation by female legislators.[34] Because these are still theoretically weak instruments and the results did not differ significantly, we did not use these in our base case analysis and did not attempt to infer causality in a strict sense.

A second limitation is the concern that the relatively little variation in state abortion policies makes their impact very difficult to differentiate from the wider state-level context.[11,13] Although some states did transition to higher hostility and lower access categories, they were already quite close to the thresholds we set for these categories. Our difference-in-differences results are based on these states, and this lack of variation likely explains the absence of significant findings for the state abortion context variables when state group fixed effects were added and when we looked at transitions over time.

Another limitation is that when a subsample of sexually active women is used for analysis, there is an implicit assumption that contraceptive decision making is undertaken only after a woman has made the decision to be sexually active. This is a potential source of bias. To correct for this bias using a two-stage sample selection correction method would require the identification of variables that are associated with a woman's decision to become sexually active, but not with her choice of method. Because we did

not have variables that met this criterion, we opted not to try to correct for sample selection bias, and acknowledge the applicability of our results only to women whose choice to become sexually active was not influenced by the policy variables we examined.

In addition, the NSFG is not a panel, in which individual respondents are followed over time. This analysis, therefore, was not longitudinal. However, the use of repeated cross-sections allowed us to conduct a trend analysis that compared contraceptive use patterns between two time periods. Finally, we did not have access to county-level data or data on proximity to a state's border. In instances where women live in a county bordering another state, there is the potential for these women to cross the border into less restrictive states. We could not control for this directly; however, this might be addressed partly by employing state-level fixed effects, and would apply to women who were living in a more restrictive state that bordered a less restrictive one.

Conclusions

Our results indicate that women living in states with more restrictive abortion contexts tend to use highly effective contraceptives. However, increases in restrictiveness do not appear to be associated with the use of these methods. The likely explanation is that the states introducing this restrictive legislation already have significant limitations on abortion in place, and that women living in these states already have adjusted their behavior to these restrictions. However, not all women are equally able to adapt to restrictive abortion contexts. From a reproductive health policy standpoint, this implies that to avoid unwanted pregnancies, it is important to ensure access to highly effective contraceptive methods for all women when access to abortions is limited.

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