How Far Is Too Far?

New Evidence on Abortion Clinic Closures, Access, and Abortions

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Abstract

We document the effects of abortion-clinic closures on clinic access, abortions, and births using variation generated by a law that shuttered nearly half of Texas' clinics. We find substantial and non-linear effects of travel distance on abortion rates: an increase in travel distance from 0-50 miles to 50-100 miles reduces abortion rates by 16 percent, and that the effects of increasing distance are smaller when the nearest clinic is already more than 50 miles away. We also demonstrate the importance of congestion with a proxy capturing effects of closures which have little impact on distance but which reduce clinics per-capita.

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I. Introduction

In June of 2016, the United States Supreme Court issued its first major abortion ruling in a quarter century, striking down components of Texas HB2 that had shuttered many of the clinics in the state and threatened to close all but a handful of those that remained (*Whole Woman's Health v. Hellerstedt* 2016a). This landmark case set a new precedent for evaluating abortion regulations against the "undue burden standard" established in *Planned Parenthood v. Casey* (1992). In particular, this 2016 decision stated that courts must "consider the burdens a law imposes on abortion access together with the benefits those laws confer" and highlighted a critical role for empirical evidence. Indeed, while little empirical evidence on the causal effects of changes in access existed at the time of this decision, a working paper version of this study was the basis for US District Judge Baker's calculation of the cumulative burden of the Arkansas law challenged in *Planned Parenthood of Arkansas and Eastern Oklahoma v. Jegley* (2018).[1]

In this study we aim to answer several related questions. What happens if/when laws are enacted that make it more difficult for abortion clinics to operate? What happens when they cause clinics to close? And to what degree are any effects of closures caused by increases in the distance women are required to travel as opposed to increased congestion at remaining clinics?

As Supreme Court Justice Elena Kagan observed during oral arguments in *Whole Woman's Health* (2016b), Texas' recent history is "almost like the perfect controlled experiment" to learn the answers to these questions. When Texas HB2 required physicians at abortion clinics to have admitting privileges at a hospital within 30 miles of the facility on November 1, 2013, nearly half of the abortion clinics in Texas immediately closed (Figure 1). On average this doubled a Texas resident's distance to her nearest clinic (Figure 1), but those in some counties were affected more than others (Figure 2).[2]

In this paper, we treat Texas' experience as a case study to document the causal effects of abortion clinic access. Specifically, we leverage plausibly exogenous geographic variation in abortion clinic access to estimate the effects using difference-in-differences models. To implement this research design, we construct panel data on abortion clinic operations from 2009 through 2015 in Texas and neighboring states. We use these data to measure driving distances from each county to its nearest clinic over time. While driving distance is a common measure that has been used in prior studies as well as referenced by the Supreme Court,[3] it notably *it fails* to capture potential changes in abortion access in areas where at least one clinic remains open. For example, closures in Dallas, Fort Worth, San Antonio, and Houston had trivial impacts on distances to clinics, because other nearby facilities remained open in each of these areas, but they dramatically increased the number of women each remaining clinic was expected to serve. An approach that focuses on distance alone ignores the possibility that clinic closures could influence abortion rates through increased congestion at remaining clinics. We explore this mechanism by proposing and constructing a new proxy for congestion.

Our econometric analysis indicates that travel distance has a substantial and non-linear effect on abortion rates. For example, we estimate that an increase in travel distance from 0 – 50 miles to 50 – 100 miles reduces abortion rates by 16 percent, and that the effects of increasing distance are smaller when the nearest clinic is already farther away. In addition to finding that even modest initial increases in distance have substantial effects on abortion rates, we also explore the effects of congestion as measured by the number of women served per abortion facility in a region. We find that congestion both reduces abortion rates and leads to delayed abortions as measured by gestational ages at the time of abortion.

Our results naturally raise the question: what are these women doing who would have obtained abortions if clinics had not closed? Though we cannot answer this question in a definitive manner, we do take some steps in this direction. One possibility is that women may be self-inducing abortions as a substitute for obtaining abortions at clinics. There is substantial anecdotal evidence suggesting that some women sought to self-induce abortions using an abortifacient sold over-the-counter at Mexican pharmacies under the brand name Cytotec (Eckholm 2013; Hellerstein 2014). We find especially large effects of clinic access for Hispanic women living near the Mexican border. This may mean that access to Cytotec plays an important role, but it may also reflect high poverty rates near the border. We also conclude that the estimated effects on abortions are too small relative to births to determine the degree to which the reductions in abortions from U.S. medical professionals translate to additional births, or whether they are offset by other behavioral changes.

II. Background

A. Prior evidence on the effects of "supply-side" abortion policies

Because "supply-side policies" targeting abortion facilities are a recent phenomenon relative to "demand-side policies" governing who can obtain an abortion, they have only recently received much attention from researchers. The most closely related paper that predates our own is Quast et al. (2017), which uses a research design similar to ours to evaluate the effects of "crow flies" distance to abortion clinics on abortion rates. While timely, this earlier study used operating licenses to measure clinic operations, which does not capture circumstances in which clinics had ceased providing abortion services though they had active licenses. We account for this scenario, and find *substantially* larger effects of distance, which is consistent with their acknowledgement that measurement error is likely to bias their estimates towards zero.[4]

Given the extant literature, we believe our study is the first to provide credible estimates of the causal effects of reduced access to abortion clinics using data on actual operations, the first to estimate the effects of a measure of congestion, and the first to consider heterogeneity using proxies for access to drugs to self-induce. We are also the first to argue that the magnitude of the effects on abortion are too small to be plausibly detected by an analysis of birth rates.[5]

B. Texas HB2 and its Aftermath

Texas HB2, which was enacted in July 2013, had two key provisions: (1) It required all abortion providers to obtain admitting privileges at a hospital located within 30 miles of the location at which an abortion was performed and (2) It required all abortion facilities to meet the standards of an ambulatory surgical center, regardless of whether they were providing surgical abortions or providing medication to induce abortions (Texas HB2 2013). In addition to these provisions, HB2 also prohibited abortions after 20 weeks gestation and required physicians to follow FDA protocols for medication-induced abortions, which restricted the use of abortion pills to pregnancies within 49 days gestational age and required that the medication be administered by a physician.[6]

Obtaining admitting privileges can be lengthy process, as it takes time for hospitals review a doctor's education, licensure, training, board certification, and history of malpractice. Moreover, many hospitals require admitting doctors to meet a quota of admissions.
After a lawsuit, decision, and a subsequent appeal, the admitting privileges requirement took effect on November 1, 2013 (*Planned Parenthood of Greater Teas Surgical Health Services v. Abbott* 2013) causing nearly half of Texas' abortion clinics to close.

The second major restriction of HB2, the ambulatory surgical center requirement, required clinics to meet additional size, zoning, and equipment requirements to meet the licensure standards for ambulatory surgical centers. This requirement was scheduled to take effect on September 1, 2014, 10 months after the admitting privileges requirement, and threatened most of Texas' remaining clinics.[7]

After another lawsuit, decision, and a subsequent appeal the ambulatory surgical center requirement went into effect on October 2, 2014 but its enforcement was blocked two weeks later by the US Supreme Court.[8]

In June of 2016, the United States Supreme Court struck down these two provisions of Texas HB2, issuing a majority opinion that Texas had failed to demonstrate that they served a legitimate interest in regulating women's health and that they imposed an undue burden on access to abortion (*Whole Woman's Health v. Hellerstedt* 2016a). As of June 2018, only three clinics that closed as a result of HB-2 have re-opened (Yaffe-Bellany 2018).

In the wake of the *Whole Woman's Health v. Hellerstedt* ruling, abortion opponents have continued to focus on supply-side abortion restrictions. Many states have continued to enforce these types of laws and to pass new ones (Guttmacher Institute 2016).[9] As such, policy considerations in the future are likely to depend on knowing what happens when abortion clinics close. The remainder of this paper focuses on answering this question, using the Texas experience as a case study.[10]

III. Data

Table 1 summarizes the variables used in our analysis: measures of abortion access, abortion rates, birth rates, and variables measuring county demographics: age and racial composition (SEER 2016) and unemployment (BLS 2016).

A. Abortion access in Texas

To evaluate the effects of Texas HB2 on abortion-clinic access, we compile a database of abortion clinic operations in Texas and adjacent states (Colorado, Louisiana, New Mexico, and Oklahoma) based on a variety of sources including licensure data maintained by the Texas DSHS, clinic websites, judicial rulings, newspaper articles, and websites tracking clinic operations maintained by both advocacy and oppositional groups.[11] We use the clinic operations database to construct two county-level measures of abortion access for each quarter: distance to the nearest abortion provider and a measure of congestion we term the "average service population."[12]

Distance to the nearest provider is calculated using the Stata *georoute* module (Weber and Péclat 2016) to estimate the travel distance from the population centroid of each county (United States Census Bureau 2016) to the nearest operating abortion clinic, including those in the neighboring states of Colorado, Louisiana, New Mexico, and Oklahoma.

Figure 1, Panel A illustrates that the distance the average Texas woman had to travel to reach an abortion clinic increased from 21 miles in the quarter prior to HB2 to 44 miles in the quarter immediately after. The percentage of women who had to travel more than 100 miles (one-way) to reach a clinic increased from 5 to 15 percent.[13] Figure 2 describes the spatial patterns of clinic closures occurring between Quarter 2 2013 and Quarter 4 2013 when HB2's first major requirement went into effect. The central-western region of Texas exhibits the largest increases in travel distances, in many cases in excess of 100 miles. Travel distance to the nearest clinic was unchanged for women whose nearest abortion clinic was already located in a major city – Houston, Dallas, Fort Worth, San Antonio, Austin or El Paso –because at least one clinic remained open in these cities.

Ideally, we would like to measure wait times or facility capacities as additional proxies for abortion access, but this is impossible because, to our knowledge, no such data were collected prior to the implementation of HB2. We therefore propose an alternative measure of abortion access that captures the increasing patient loads faced by a reduced number of facilities. We call this variable the "average service population." To construct it, we first assign each county *c* in time period *t* to an "abortion service region" *r* according to the location of the closest city with an abortion facility.[14] The average service population is the ratio of the population of women aged 15-44 in the service region to the number of facilities in the service region:

$$(1)\ Average\ service\ population_{c,r,t} = \frac{\sum_{c\in r} population_{c,t}}{number\ of\ clinics_{c,t}}$$

This measure represents a novel attempt to capture a potentially important dimension of abortion access with available data. It also has necessary limitations, most notably in that the denominator sums the number of abortion facilities in a region without accounting for potential differences in capacities of these facilities.[15] Figure 1, Panel B illustrates time trends in this measure of congestion.[16] In the immediate aftermath of HB2, the average service population rose from 150,000 to 290,000 in Texas. This occurred for two reasons: (1) As clinics closed in small cities, women had to travel to clinics that remained in larger cities, shrinking the number and expanding the sizes of service regions; and (2) As clinics closed in large cities, there were fewer providers of abortion services. Figure 3 summarizes spatial variation the *change* in the average service population between the second and fourth quarters of 2013. As was the case for the distance measure of access, there is substantial variation in how this congestion measure of access changed across Texas following HB2. The average service population did not change in eastern Texas, where only one clinic closed in the fourth quarter of 2013 (though several did close the following year). But in the Dallas-Fort Worth region, where distances had not changed, the average service population increased by 250,000. Clinic closures continued through 2014, and Figure 1 shows that the average service population continued to rise. In Dallas-Fort Worth an additional clinic closure in June 2015 increased the average service population from 380,000 to 480,000. Over the same period, TPEP (2015a) reports that wait times at Dallas and Fort Worth clinics increased from 2 to 20 days.[17]

B. Abortion Rates in Texas

We use publicly available data on Texas abortions by county of residence (TDSHS, 2017a; 2017b). To produce these data, the Texas DSHS combines in-state abortions, which providers are mandated to report, with information on out-of-state abortions it obtains via the State and Territorial Exchange of Vital Events (STEVE) system. To construct abortion rates, we use population denominators based on annual estimates of county populations by race, gender and age from SEER (2016). We use these same population data to construct demographic control variables.

These abortion rates account for interstate travel so far as the Texas DSHS is able to observe abortions to Texas residents reported via the STEVE system. Based on information we obtained from the state health departments in nearby and neighboring states, we estimate that the abortion data provided by the Texas DSHS may be missing up to 1,164 abortions obtained in these states in 2014 and 1,418 in 2015, roughly 2 percent of total abortions to Texas residents.[18]

In subsequent sections we estimate that these abortions obtained in nearby states can only account for a small fraction of the observed effects. We also demonstrate our main results are robust to focusing only on counties where it is unlikely for many women to seek abortions out of state in any year.[19]

Figure 4 illustrates the change in abortion rates between 2013 and 2014, aggregating rates to the public health region to reduce visual "noise" in rates for counties with small populations. Figure 4 illustrates that abortion rates declined across the state in the year following HB2, but the reductions were most dramatic in the regions that experienced the largest reductions in abortion access. In the Rio Grande Valley and Texas Panhandle and west Texas, abortion rates declined by more than 30 percent, while in the Houston and Austin areas, they declined by less than 10 percent. This visualization foreshadows the results of our difference-in-difference analysis.

Our analysis also makes use of county-level abortion counts provided to us by the Texas DSHS for three gestational age groupings: less than 7 weeks gestation, 7 to 12 weeks gestation, and greater than 12 weeks gestation (TDSHS 2018). An important limitation of these data, which we address through sensitivity analysis, is that the Texas DSHS suppresses abortion counts by gestational age in cells where the count is between 1 and 9 abortions.[20] We also note that the Texas DSHS switched from reporting abortions by gestational age to reporting abortions by post-fertilization age in 2014. To make the data series comparable, we adjust the categories using a 2-week difference.

C. Births Rates in Texas

We use restricted-use natality files provided by the National Center for Health Statistics from 2009 – 2015 (NCHS 2017). These data consist of a record of every birth taking place in the

United States over this time period. To construct county birth rates, we use population denominators based on annual estimates of county populations from SEER (2016).

IV. Empirical Strategy

We estimate the effects of access to abortion clinics using a generalized difference-in-differences design, which exploits within-county variation over time while controlling for aggregate time-varying shocks. The identifying assumption underlying this approach is that changes in abortion rates for counties with small changes in access provide a good counterfactual for the changes in abortion rates that would have been observed for counties with larger changes in access if their access had changed similarly.

Given the discrete nature of abortions, and because we encounter cells with zero abortions when looking at some subgroups, we operationalize this strategy with a Poisson model.[21] In particular, our approach to estimating the effect of changes in abortion access on the abortion rate corresponds to the following equation:

$$\textbf{(2)}\ E\left[AR_{c,t}\middle|\boldsymbol{access_{c,t}}, \alpha_c, \theta_t, \boldsymbol{X_{c,t}}\right] = exp(\boldsymbol{\beta}\boldsymbol{access_{c,t}} + \alpha_c + \theta_t + \boldsymbol{\gamma}\boldsymbol{X_{c,t}})$$

where $AR_{c,t}$ is the abortion rate for residents of county *c* in year *t*; $access_{c,t}$ is a set of measures of access to abortion clinics for residents of county *c* in year *t*; α_c are county fixed effects, which control both observed and unobserved county characteristics with time-invariant effects on abortion rates; θ_t are year fixed effects, which control for time-varying factors affecting abortion rates in all Texas counties in the same manner; and $X_{c,t}$ can include time-varying measures of county demographics, unemployment, and family-planning access. Specifically, the demographic control variables include the fraction of the 15-44 female population in each five-year grouping and the fraction of each of these age groups that is non-Hispanic white, non-Hispanic black, or Hispanic (versus other race/ethnicity). Our approach to controlling for family planning follows Packham (2016) who evaluates the effects of Texas' decision to cut funding to family planning clinics by two-thirds in 2012. In particular, we control for whether a county had a publicly funded family planning clinic prior to the funding cut interacted with the time period after the funding cut occurred (post-2012).

Because Poisson models are more typically thought of as considering counts, not rates, we note that this model can be expressed alternatively as estimating the natural log of the expected count of abortions while controlling for the natural log of the relevant population and constraining its coefficient to be equal to one. All of the standard-error estimates we report allow errors to be correlated within counties over time.[22]

As described in Section III, our measures of abortion clinic access are based on travel distance to the nearest clinic (from the county population-weighted centroid) and the "average service population," which measures the number of people each clinic is expected to serve in each "service region".[23] To separately identify the effects of these two measures of access, there must be independent variation. As we noted in Section III, such variation is expected because closures in areas where some clinics remained open increase congestion without affecting distance-to-nearest-clinic whereas closures in areas where no clinics remained open increase both congestion and distance-to-nearest-clinic. This is evident from a comparison of Figures 2 and 3 which depicted changes in the two measures across different Texas counties.[24]

For ease of interpretation for readers, in the main text we present estimates with the average service population included as a linear term and using a set of indicator variables to measure distance, including an indicator for being 50-100 miles from the nearest clinic, an indicator for being 100-150 miles to the nearest clinic, and an indicator for being 150-200 miles to the nearest clinic (such that being 0-50 miles from the nearest clinic is the excluded category). In Appendix C we report estimated effects from a model in which abortion rates are a quadratic and the average service population.[25]

V. Results

A. Establishing the Validity of the Research Design

The primary goal of our paper is to estimate the *causal* effects of abortion-clinic access on abortions provided by US medical professionals. The identifying assumption underlying our differences-in-difference strategy is the changes in abortion rates for counties with small changes in access provide a good counterfactual for the changes in abortion rates that would have been observed for counties with larger changes in access if their access had changed similarly.

To assess the identifying assumption, we plot data over time for each of four groups categorized according to their changes in distance-to-nearest-clinic between the second quarter of 2013 (before HB2) and the fourth quarters of 2013 (after HB2). One group consists of counties with no increase in distance-to-nearest-clinic over this time period. The other three groups of counties are in terciles based on the amount that their distance-to-nearest-clinic increased over the same period. Panel A of Figure 5, shows that the average distance-to-nearest-clinic was flat for all four groups of counties prior to 2013. This implies we can use pre-2013 years to evaluate the credibility of the common trends assumption. Panels B and C of Figure 5 show similar plots for log of the abortion and birth rates. From 2009 to 2012, log abortion rates were changing very similarly for counties that would subsequently experience a major increase in distance-to-nearest-clinic and counties that would subsequently experience smaller (or no) increases. Panel C similar evidence of common pre-HB2 trends in birth rates. Overall, Figure 5 provides empirical support for our identifying assumption—that these common trends would have continued into subsequent years in the absence of differential changes in abortion clinic access.

In addition to providing support for the validity of our identification strategy, Figure 5 also provides some visual evidence of the effects of distance on abortion and birth rates. In particular, counties experiencing the greatest increase in distance exhibit correspondingly greater decreases in abortion rates. Some readers may also note that distances decreased somewhat for the top two terciles between 2014 and 2015 and also that that there is a corresponding "rebound" in the abortion rate. This could be taken as further evidence that abortion rates respond to changes in distance to clinics. That said, the magnitude of the rebound in abortion rates is such that it could reflect that the effects of the earlier, larger, increases in distance are short lived. We explore this possibility in sensitivity checks for our main results.[26] Figure 5 shows no evidence of an increase in births corresponding to the decrease in abortions, which we discuss in greater detail below.[27]

B. The Causal Effects of Distance and Congestion

Having provided evidence to support the key identifying assumption underlying our difference-in-differences research design, we now present estimates of the causal effects of access to abortion clinics that are based on this research design. In order to give a general sense of the effects, Table 2 reports estimates in which abortion rates are measured as a step-function in distance and as a linear function of a our measure of congestion, the average service population. The estimates in Column 1 of Table 2 indicates that relative to having the nearest abortion provider within 50 miles, having the nearest abortion provider 50-100 miles away reduces abortions by 16 percent, having the nearest abortion provider 100-150 miles away reduces abortions by 28 percent, having the nearest abortion provider 150-200 miles away reduces abortions by 38 percent, and having the nearest abortion more than 200 miles away reduces abortions by 44 percent.[28] They also indicate that a 100,000 woman/clinic increase in average service population reduces abortions by 7 percent.[29,30]

Based on our estimates, if access to abortion clinics had remained at pre-HB2 levels, Texas women would have had 119,730 legal abortions in 2014-2015 rather than the 107,830 observed in the abortion surveillance data. This represents an estimated reduction of 11,900 abortions due to HB2 in these two years after which it was enacted. We estimate that 41 percent of this two-year reduction was due to increased driving distances, and 59 percent was due to increased congestion.[31]

C. Addressing Interstate Travel

As discussed in Section III.B, abortion surveillance practices vary in neighboring states. Summing up abortions to Texas residents in states not participating in STEVE, we estimated that the 53,882 abortions to Texas residents reported in 2014 by the Texas DSHS (2017) may be missing up to 1,164 abortions in Arkansas, Colorado, Oklahoma, and New Mexico. Similarly, the 54,310 abortions reported in 2015 may be missing up to 1,418 abortions obtained in neighboring states.

Might these abortions obtained in other states explain our results? In the previous section we estimated that Texas HB2 reduced the number of abortions by 11,900 over 2014 and 2015 (and by a smaller number in 2013 which was only partially affected by closures). This estimated effect is far in excess of the 2,582 abortions we are potentially missing in nearby states during this two-year period, but we do note that they could account for as much as 20 percent of the estimated reduction. That said, in Column 2 of Table 2 and Appendix Table C1 we confirm that our estimates are robust to the exclusion of counties in the Texas Panhandle, the counties we might be most concerned about due to missing data on abortions obtained in New Mexico.[32]

D. Effects on Abortions by Gestational Age

Thus far we have found evidence that increasing distance and congestion both cause reductions in observed abortions. It is also possible that these factors may delay abortions because of increased wait times, or because it takes additional time for women to make plans and assemble the resources required for longer trips. To empirically assess whether reduced access increases delay, we obtained county-level abortion counts from the Texas DSHS for three gestational age groupings: less than 7 weeks gestation, 7 to 12 weeks gestation, and greater than 12 weeks gestation.

Columns 3–5 of in Table 2 and Appendix Table C1 show the estimated effects on the abortion rate for abortions at these different gestational ages. These results suggest that, holding congestion constant, increased travel distances reduce abortions in all of the gestational age categories. This pattern of estimates prevents us from determining whether increases in travel distance causes delay. In contrast, the pattern of the estimated effects of congestion indicate that increased congestion at abortion facilities does causes delays. Specifically, the estimates in Table 2 indicate that a 100,000 increase in the average service population causes a 5.6 percent increase in second trimester abortions, no impact on abortions at 7-12 weeks, and a 13.4 reduction in abortions prior to 7 weeks. This suggests that increasing congestion shifted the distribution of gestational age to the right. The estimates in Appendix Table C1 lead to the same general conclusion.[33,34]

An important caveat to these analyses is that the Texas DSHS suppresses abortion counts by gestational age in cells where the count is between 1 and 9 abortions.[35] That said, we come to the same conclusions if we evaluate a balanced panel of counties for which abortions by gestational age are never suppressed.[36]

E. Heterogeneity by Ethnicity and Distance to Mexico

Anecdotal evidence suggests that as access to abortion clinics decreased in Texas, some women sought to self-induce abortions by using Cytotec, a drug that is sold over-the-counter at Mexican pharmacies for the treatment of ulcers (Eckholm 2013; Hellerstein 2014). Cytotec is a brand name for Misoprostol, a drug that induces uterine contractions. This drug is the second in a two-part drug combination that is the FDA protocol for medical abortions. Taken alone in the first trimester, Misoprostol is successful at inducing an abortion about 90 percent of the time, with decreasing efficacy as the pregnancy progresses (von Hertzen et al. 2007). The prescribing information for Cytotec reports that it can cause incomplete abortion and that it increases the risk of congenital anomalies (skull defects, cranial nerve palsies, facial malformations, and limb defects) for pregnancies that continue after the drug is taken.[37]

In 2008-2009, 1.2 percent of patients at abortion clinics reported that they had used Misoprostol on their own to self-induce abortion at some point in the past (Jones 2011). Rates may be higher in Texas because women can more easily travel to Mexico to obtain the drug. In 2012, prior to the enactment of HB2, 7 percent of Texas abortion patients reported that they had tried to "do something" on their own to end the pregnancy (Grossman et al. 2014). In 2014, the Texas Policy Evaluation Project surveyed 779 Texas women; 2 percent reported attempting to self-induce an abortion and 22 percent reported knowing someone else who had done so (TPEP 2015b). Based on this finding, the authors estimate that 2 to 4 percent of Texas women aged 18-

49 may have attempted to self-induce an abortion, and that rates are higher for Hispanic women living in counties bordering Mexico.

Ideally, we would be able to evaluate the effects of abortion-clinic access on self-induced abortions as well those that are provided at clinics in order to measure the degree to which women substitute the former for the latter. However, these self-induced abortions take place out of sight of public health authorities tracking legal abortions in licensed facilities, which makes a rigorous analysis along these lines impossible. Instead, we examine whether the effects on abortions provided by US medical professionals are relatively large among Hispanic women and women who live close to the Mexican border, as we anticipate that such women would have better access to Cytotec than the average woman.[38]

To implement this analysis, we separately estimate the effects for Hispanic and non-Hispanic women, and we estimate a model with clinic distance entered linearly and by interacting this variable and "average service population" with in indicator that a given county is less than 100 miles from the nearest border crossing.[39]

The results of this analysis are shown in Table 3. The results of this analysis indicate that there are heterogeneous effects of across ethnicity and across counties near versus far from Mexico. For Hispanics, both the estimated effects of both access measures are significantly larger for those in counties near Mexico. For non-Hispanics, the estimates are far too imprecise us to discern whether the effects are significantly different for those in counties near Mexico relative to those in counties farther from Mexico. These estimates also indicate that the effect of distance is significantly larger for Hispanic women than for non-Hispanic women, among those who live more than 100 miles from Mexico. Interestingly, we do not find evidence of a statistically significant effect of the average service population among Hispanic women who live more than 100 miles from Mexico. Estimates from our quadratic model (Figure C3) indicate that the effects are similar for Hispanics and non-Hispanics more than 100 miles from the border and suggest that decreasing access to abortion providers had larger effects closer to the border.

These results provide empirical support to the anecdotal and survey evidence that substitution to self-induced abortion may have been common, especially in areas close to Mexico. The results could also be explained by the higher poverty rates among Hispanic women and residents of counties near the Mexican border. We also acknowledge that the estimates are imprecise given the small number of border counties.

F. Do the Effects on Abortions Show Up in Birth Rates?

This question naturally arises from the preceding set of results. In a mechanical sense, one might expect fewer abortions to lead to more births. However, reductions in abortions provided by medical professionals could be offset by increases in self-induced abortions. Moreover, reduced access to abortion clinics could lead to changes in sexual behavior and contraceptive use, which could also offset impacts on abortion. These are reasons to believe that the full reduction in observed abortions may not be reflected in an increase in birth rates.

To get a sense of the largest magnitudes that we might expect from an analysis of birth rates, we have estimated effects that we would expect solely from the estimated effects on abortions for each county-year based on the estimates reported in Table 2 (Column 1) and using a similar model applied to the quarterly birth rate data for each county.[40] Specifically, we estimate the effect on birth rates that we would obtain if all of the "missing abortions" caused by reductions in access since 2012 were to show up as births two quarters later. These estimates indicate that at a maximum, we might expect to increases in distance from 0-50 miles to increase birth rates by 2.1 percent for an increase to 50-100 miles, by 2.5 percent for an increase to 100-

150 miles, by 2.7 percent for an increase to 150-200 miles, and by 3.6 percent for an increase to greater than 200 miles. They also indicate that a 100,000 increase in the average service population could increase the birth rate through impacts on abortion by as much as 0.5 percent.

The estimated effects we obtain when we evaluate the birth rates that are actually observed are shown in Table 4. The estimated effects of changes in distance are typically smaller than we would expect based solely on our estimated effects on abortions, suggesting that the full reduction in observed abortions is not reflected in an increase in birth rates. This is consistent with the mechanisms described above. Our analysis of various subgroups defined by age, race, ethnicity, birth parity, and marital status also do not provide any robust evidence of effects on birth rates—the estimates are sometimes positive, sometimes negative, and they are typically not statistically significant at conventional levels. These estimates also offer little evidence of statistically significant effects of our measure of congestion. Estimated effects based on our model that considers a quadratic in distance and the average service population generally lead to similar conclusions.[41]

VI. Discussion and Conclusion

The results of our empirical analysis demonstrate that regulation-induced reductions in access to abortion clinics can have sizable effects. For women living within 200 miles of an abortion clinic, we document substantial and statistically significant effects of increasing distance to abortion providers. The finding that even small initial increases in distance have significant effects is notable in light of previous Supreme Court opinions suggesting that travel up to 150 miles not be considered an undue burden (*Planned Parenthood v. Casey* 1992). Moreover, our estimates also indicate that increased travel distances are not the only burden imposed by clinic closures. Indeed, our results indicate that the effects of congestion, as measured by clinics-per-capita in a service region, can play an even larger role. The effects operating through this channel account for 59 percent of the overall effect of reduced clinic access caused by closures following Texas HB2. We also find that impacts through this channel shift the gestational age distribution (at the time of abortion) to the right, which is consistent with the impacts on congestion causing delays.

Based on our estimated models, if access to abortion clinics had remained at pre-HB2 levels, Texas women would have had nearly 12,000 more abortions from U.S. medical professionals in 2014-2015 than were actually observed. We hope that future research can address what explains these “missing” abortions. It is possible that some women responded to the reduction in access to abortion facilities by decreasing risky sexual behaviors. It also is possible that they can be explained by more women giving birth, though our analysis of birth rates does not reveal statistically significant effects for the most part. A third possibility, based on anecdotal and survey evidence, is that some Texas women resorted to “do-it-yourself abortions” using an over-the-counter drug available in Mexican pharmacies (Hellerstein 2014; TPEP 2015b). Consistent with this story, we find the largest effects on abortions obtained from U.S. medical professionals among Hispanic women living near the Mexican border, but this could also be explained by higher poverty in these areas. Data limitations make it difficult to directly investigate the possibility of self-induced abortions.

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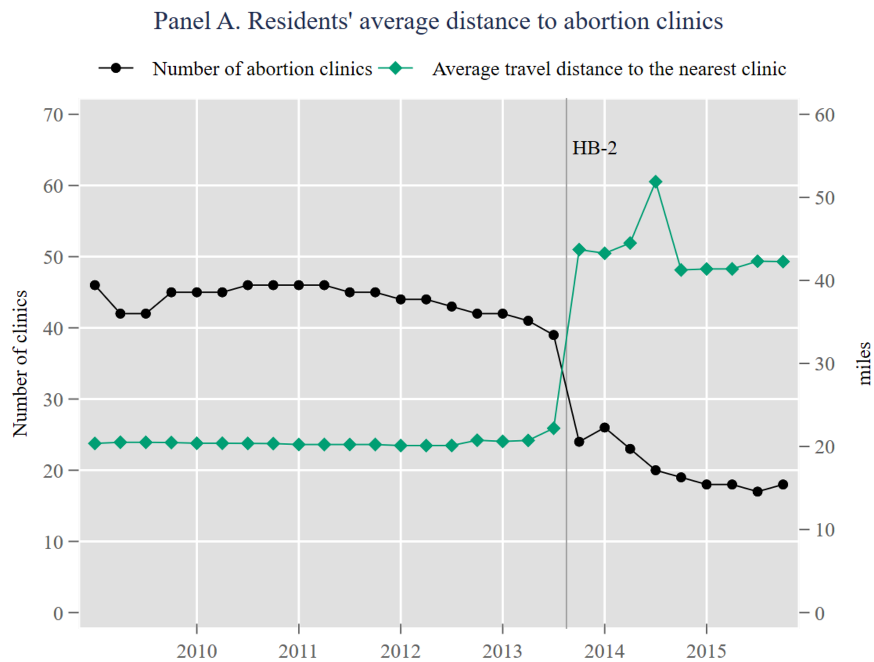
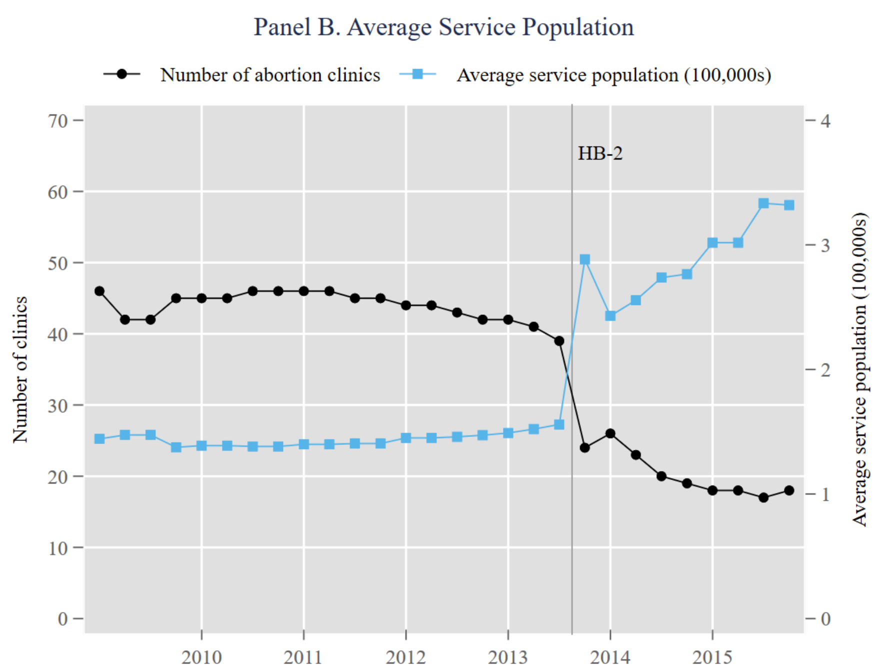
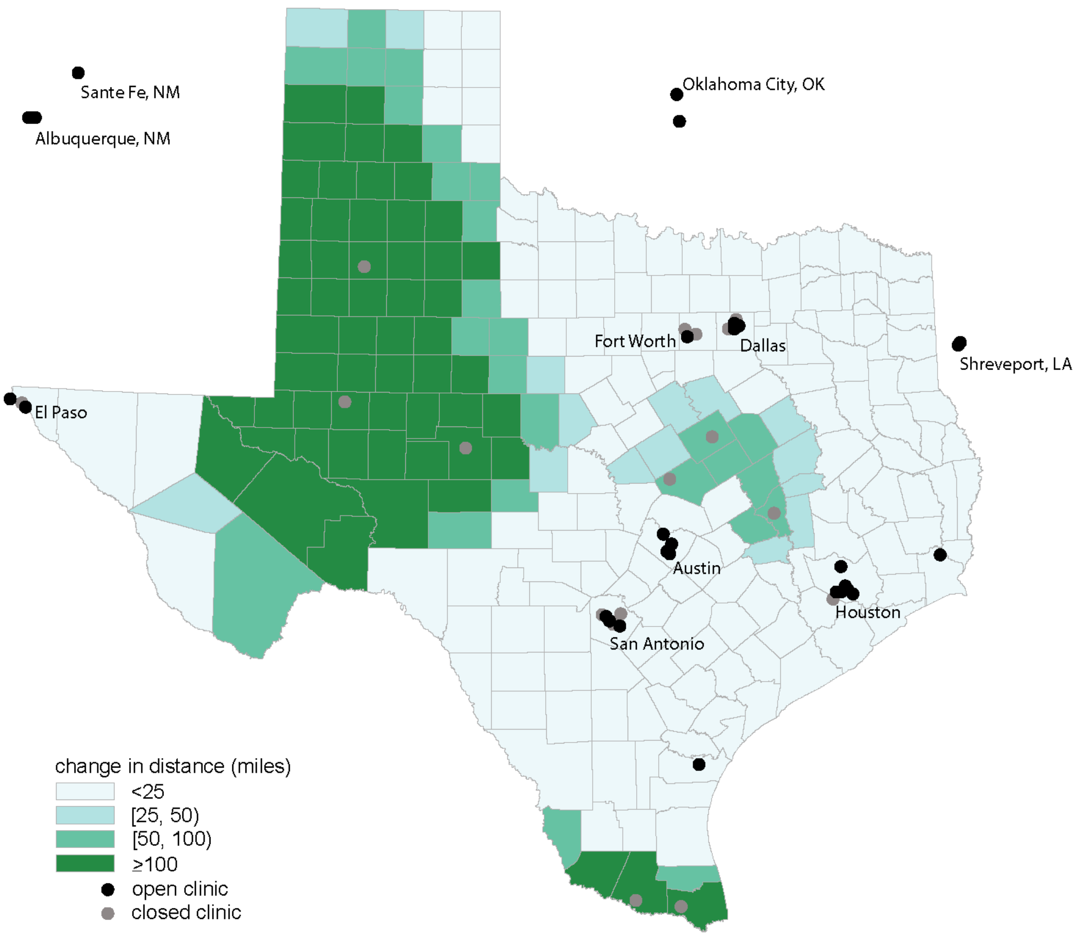
Figure 1: Abortion clinics and abortion access, Texas 2009-2015





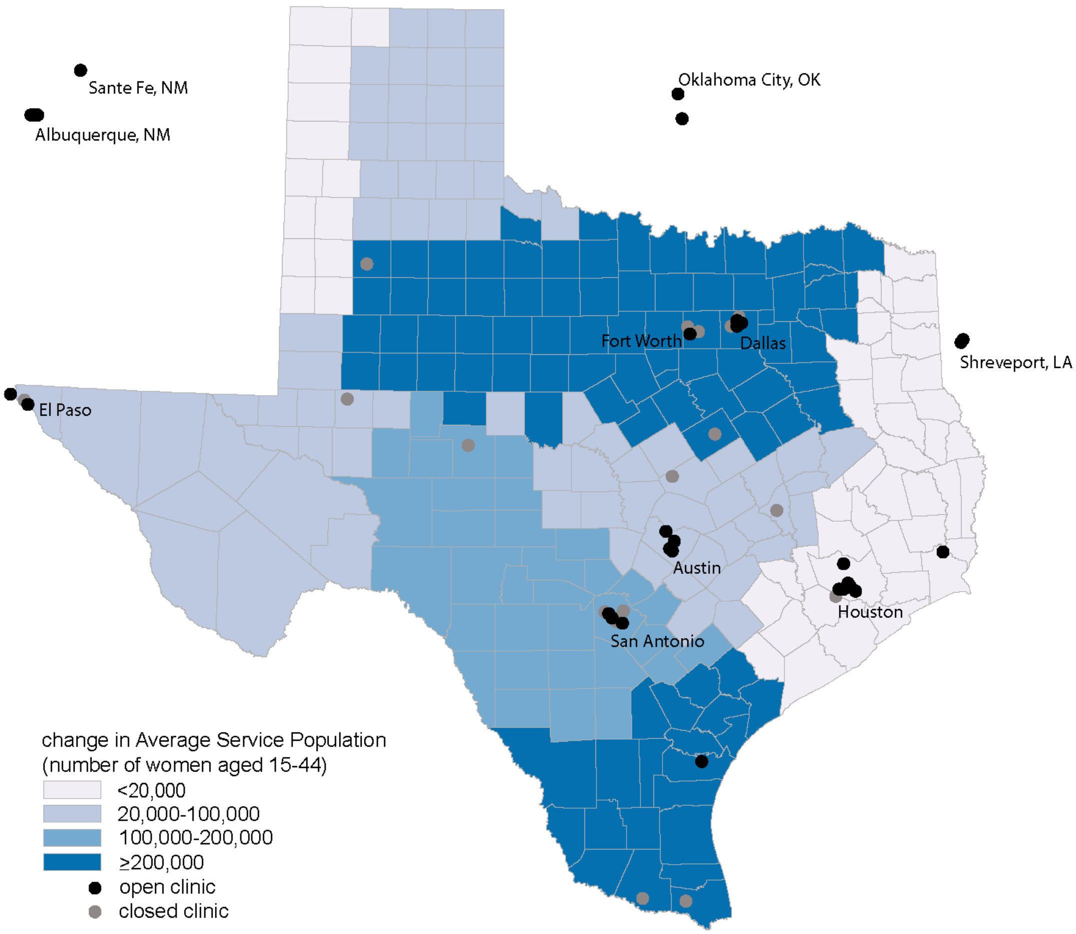
Notes: Notes: Distances are population-weighted average travel distances from county population centroids to the geographic coordinates of the nearest open abortion facility. Facility operations are measured quarterly, and a facility is considered "open" if it provided surgical or medical abortions for at least half of a given quarter. Sources: Facility operations data were compiled by the authors, annual county-level population estimates were obtained from SEER (2016), and geographic coordinates of county population centroids were obtained from the United States Census Bureau (2016).

Figure 2: Change in distance to the nearest abortion clinic, Q2 2013 to Q4 2013



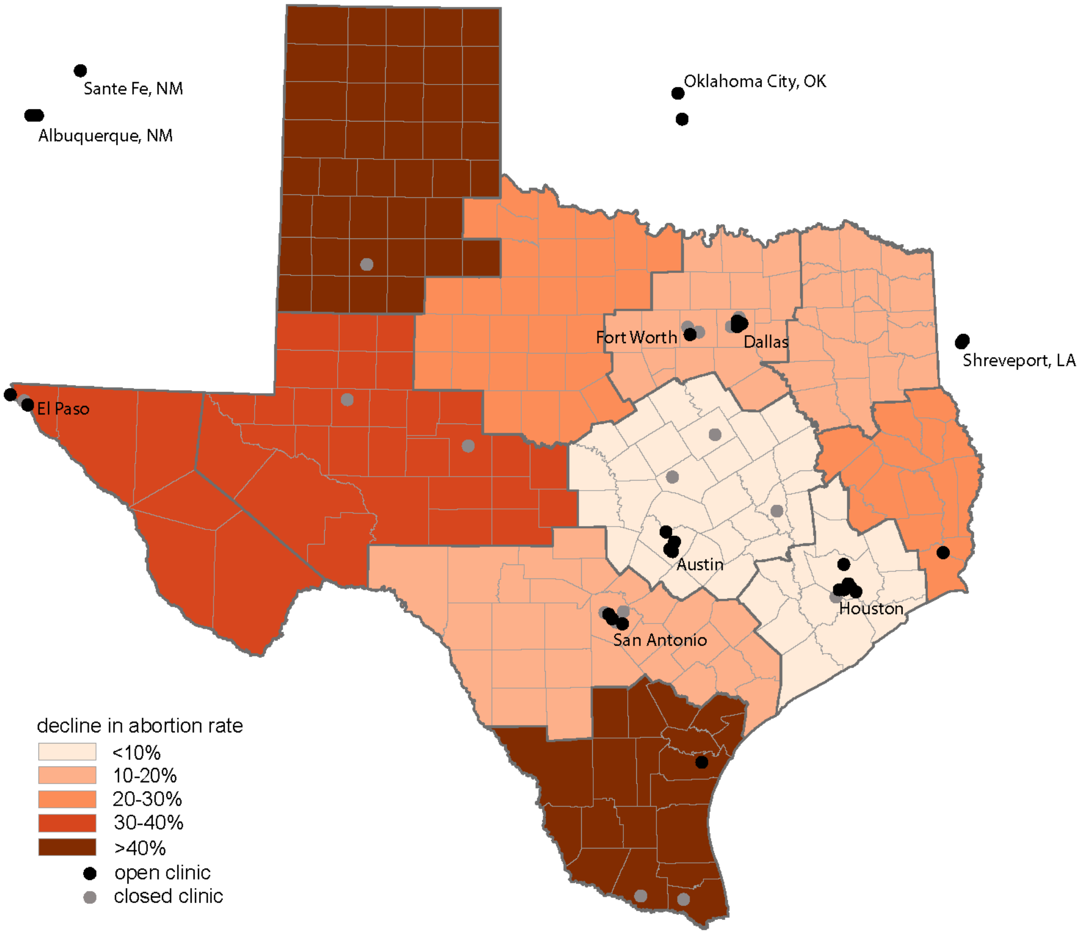
Notes: County-level change in the average distances to the nearest open abortion facility measured in Quarter 2 2013 and Quarter 4 2013. Distances are the estimated travel distances from county population centroids to the geographic coordinates of the nearest open abortion facility. A facility is considered "open" if it provided surgical or medical abortions for at least 2 months in a given quarter. Sources: The clinic operations data were compiled by the authors, annual county-level population estimates were obtained from SEER (2016), and geographic coordinates of county population centroids were obtained from the United States Census Bureau (2016).

Figure 3: Change in Average Service Population, Q2 2013 to Q4 2013



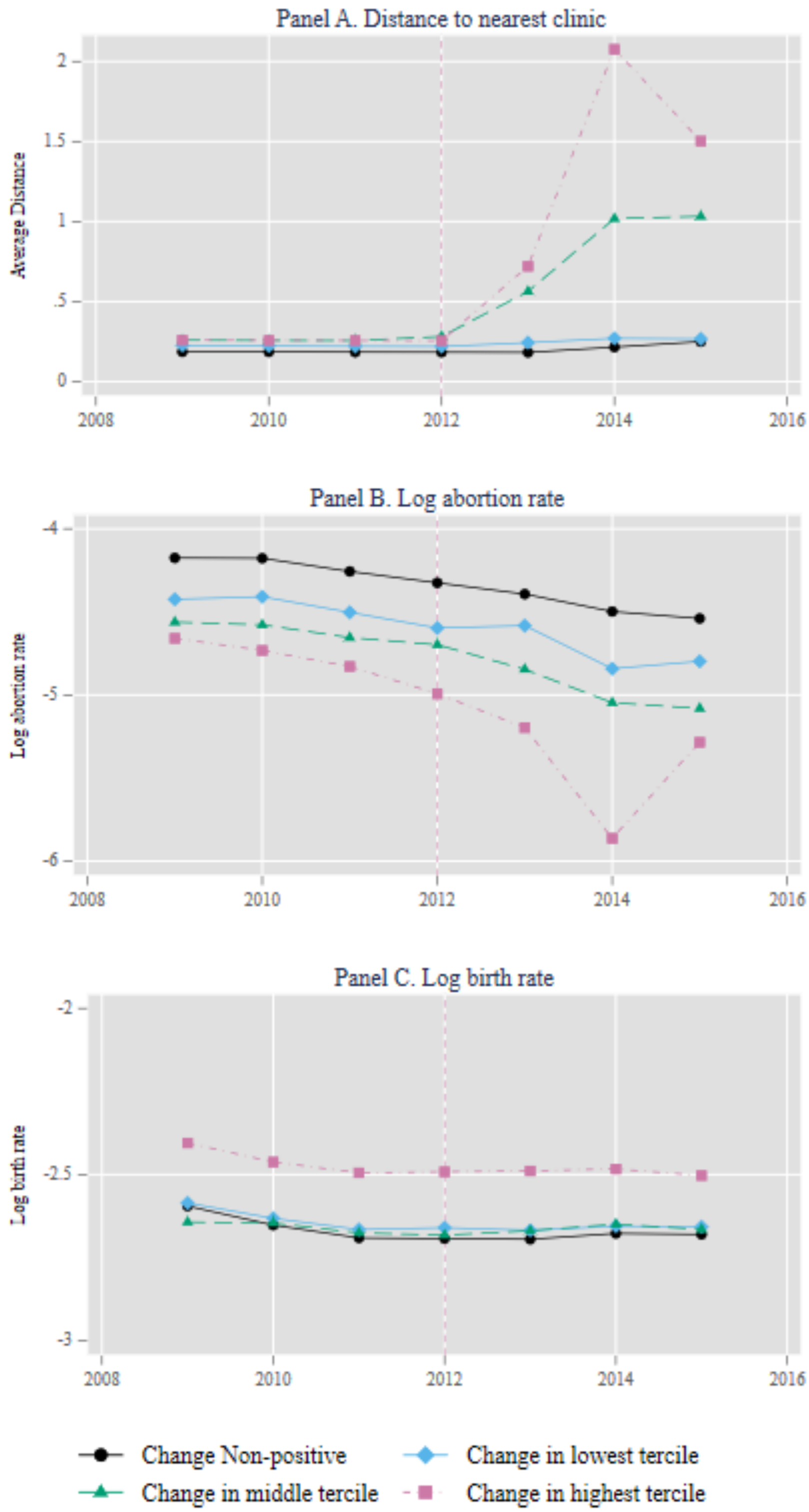
Notes: County-level change in the average service population in Quarter 2 2013 and Quarter 4 2013. The average service population associated with a county in a given year is based on the population (women aged 15-44) and the number of clinics in its abortion service region in that year. Service regions are defined annually by spatial proximity to the nearest city with an abortion clinic. Sources: The clinic operations data were compiled by the authors, annual county-level population estimates were obtained from SEER (2016), and geographic coordinates of county population centroids were obtained from the United States Census Bureau (2016).

Figure 4: Percent change in abortion rates by public health region, 2013 to 2014



Notes: Percent change in abortions per 1,000 women aged 15-44, calculated for each of Texas' 11 Public Health Regions. Clinics are coded as open or closed based on their status in the second quarter of 2014.

Figure 5: Trends in distance, abortions, and births across treatment intensity groups, where treatment intensity is the change in distance between Q2 2013 and Q4 2013



Notes: The vertical line highlights the final year of data before HB2 was enacted.

Table 1: Summary Statistics

Variable	2009 to 2015 mean	2009 to 2015 s.d.	2012 mean	2012 s.d.	2014 mean	2014 s.d.
Abortion rate (per 1,000 women)						
Total	11.68	5.05	11.78	4.98	9.46	4.32
Non-Hispanic	12.37	5.83	12.48	5.96	10.31	4.59
Hispanic	10.71	4.65	10.78	4.29	8.28	4.29
< 7 weeks gestation†	5.45	2.53	5.84	2.32	3.31	1.56
7-12 weeks gestation†	5.84	2.57	5.66	2.6	5.46	2.27
>12 weeks gestation†	1.03	0.46	0.98	0.46	1.12	0.47
Birth rate (per 1,000 women)						
Total	71.11	9.74	69.65	9.38	70.6	9.33
Age 15 to 19	44.86	16.98	44.34	15.52	37.74	13.94
Age 20 to 29	111.74	24.12	108.95	23.38	109.9	23.6
Age 30 to 39	71.74	10.58	70.73	9.84	75.3	9.97
Age 40 to 44	10.08	2.85	9.98	2.74	10.29	2.93
White	62.41	8.89	61.74	9.12	63.64	9.12
Black	63.99	10.06	62.3	9.61	64.33	10.32
Hispanic	83.01	11.12	80.44	9.87	80.17	10
Other	62.06	11.41	63.51	12.35	63.14	10.36
1st birth†	23.42	3.23	22.99	2.97	23.03	3.19
2nd birth†	47.65	7.46	46.62	7.04	47.52	7.37
Married†	41.21	5.53	40.33	5.26	41.15	5.16
Unmarried†	29.91	8.6	29.32	8.36	29.45	8.46
Measures of abortion access						
Distance (100s of miles)	0.28	0.48	0.2	0.33	0.45	0.7
Average Service Population (100,000s)	1.92	0.92	1.46	0.49	2.62	0.66
Race						
White	40.04	19.02	40.05	19.04	39.2	18.65
Black	13.05	8.39	13.02	8.38	13.19	8.37
Hispanic	41.35	21.44	41.4	21.46	41.73	21.17
Other	5.56	3.93	5.53	3.89	5.88	4.14
Age distribution						
15 to 19	16.72	2.01	16.6	1.96	16.41	1.94
20 to 29	34.03	3.9	34.08	4.03	34.2	3.72
30 to 39	33.1	2.73	32.97	2.71	33.15	2.61
40 to 44	16.15	1.9	16.34	1.94	16.25	1.88
Economic conditions						
Unemployment rate	6.59	1.86	6.77	1.41	5.17	1.22

Notes: Population-weighted summary statistics calculated for Texas counties (n=254) for the pooled sample period (2009-2015) and individually for 2012 (the year prior to HB-2) and 2014 (the year after HB-2). † indicates that rate is calculated using population of women aged 15-44 as denominator. All other rates are calculated using population of women in the indicator racial, ethnic, or age group as denominator. Sources: Authors' compilation of clinic operations, annual county-level population estimates from SEER (2016), abortions by county of residence from the Texas DSHS (2017a; 2017a; 2018), geographic coordinates of county population centroids from the United States Census Bureau (2016), and unemployment rates from the BLS (2016).

Table 2: Estimated effects of distance and congestion on abortions

	(1) Total	(2) Panhandle excluded	(3) < 7 weeks	(4) 7-12 weeks	(5) >12 weeks
I(50<Distance < 100)	-0.179***	-0.180***	-0.295***	-0.143*	-0.226***
	(0.032)	(0.031)	(0.109)	(0.076)	(0.072)
I(100 < Distance < 150)	-0.333***	-0.393***	-0.615**	-0.157	-0.410***
	(0.09)	(0.088)	(0.305)	(0.102)	(0.095)
I(150 < Distance < 200)	-0.477***	-0.482***	-0.949***	-0.284***	-0.271
	(0.068)	(0.069)	(0.254)	(0.103)	(0.247)
I(200 < Distance)	-0.588***	-0.444***	-0.267	-0.652***	-0.469***
	(0.097)	(0.091)	(0.282)	(0.149)	(0.155)
Average Service Population (100,000s)	-0.073**	-0.077**	-0.134*	-0.009	0.056*
	(0.03)	(0.032)	(0.073)	(0.028)	(0.029)
No. of counties	254	213	110	127	59
N	1775	1488	552	611	321

Notes: Estimates are based on a Poisson model evaluating expected abortion rates among women aged 15 to 44 using county-level data for all 254 Texas counties from 2009-2015. All models include county and year fixed effects as well as the following time-varying county control variables: the fraction of the 15-44 female population in five year groupings; the fraction of each of these age groups that is non-Hispanic white, non-Hispanic black, or Hispanic versus other race/ethnicity); family planning control variables as described in the text; and the county unemployment rate. Standard errors (in parentheses) allow errors to be correlated within counties over time. *, **, and *** indicate statistical significance at the ten, five, and one percent levels, respectively.

Table 3: Estimated effects on abortion by ethnicity and distance from Mexican border

	Non Hispanics	Hispanics
	(1)	(2)
Distance (100s miles)	-0.133***	-0.258***
	(0.020)	(0.036)
Average Service Population (100,000s)	-0.092***	<|0.001|
	(0.031)	(0.041)
Distance X I(<100 miles to Mexican border)	0.339	-0.222**
	(0.287)	(0.089)
Avg Service Pop X I(<100 miles to Mexican border)	-0.255	-0.120***
	(0.158)	(0.045)

Notes: Estimates are based on a Poisson model evaluating expected abortion rates among women aged 15 to 44 using county-level data for all 254 Texas counties from 2009-2015. Distance to the Mexican border is calculated as travel distance between county population centroids and the nearest Mexican border crossings that can be used by pedestrians and/or private vehicles. All columns control for county fixed effects, year fixed effects, demographics, the unemployment rate, an indicator for the presence of a family planning clinic in the county, and this indicator's interaction with post-2012. Standard errors (in parentheses) allow errors to be correlated within counties over time.
*, **, and *** indicate statistical significance at the ten, five and one percent levels, respectively.

Table 4: Estimated effects of distance and congestion on births

Group	All	15-19	20-29	30-39	40-44	White	Hispanic	Black	Other	1st Birth†	2nd+ Birth†	Married†	Unmarried†
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
I(50<Dist≤100)	0.016*	0.043**	0.012	0.015	0.078**	0.008	0.014	0.006	0.012	-0.025*	0.036***	-0.002	0.039***
	(0.009)	(0.019)	(0.011)	(0.016)	(0.037)	(0.010)	(0.012)	(0.018)	(0.031)	(0.015)	(0.011)	(0.015)	(0.014)
I(100<Dist≤150)	0.004	-0.010	0.009	0.015	0.064	0.033**	-0.002	-0.034	0.018	-0.033*	0.022**	0.019*	-0.01
	(0.007)	(0.018)	(0.008)	(0.011)	(0.043)	(0.013)	(0.008)	(0.044)	(0.057)	(0.018)	(0.010)	(0.011)	(0.013)
I(150<Dist≤200)	-0.013	0.028	-0.008	-0.021	0.154*	-0.006	-0.007	-0.008	0.037	-0.02	-0.009	-0.026*	0.003
	(0.011)	(0.031)	(0.013)	(0.023)	(0.086)	(0.015)	(0.013)	(0.044)	(0.094)	(0.023)	(0.014)	(0.015)	(0.015)
I(200<Dist)	-0.001	0.036**	0.009	0.022	0.089	0.046***	-0.001	-0.019	0.002	-0.038*	0.015	0.013	-0.017*
	(0.010)	(0.016)	(0.013)	(0.014)	(0.056)	(0.012)	(0.015)	(0.035)	(0.046)	(0.020)	(0.011)	(0.016)	(0.009)
Avg. Service	0.003	0.008	0.002	-0.002	0.004	0.005	0.006**	-0.002	-0.005	0.015***	-0.003	0.002	0.003
Pop. (100,000s)	(0.002)	(0.006)	(0.003)	(0.003)	(0.015)	(0.004)	(0.003)	(0.006)	(0.008)	(0.005)	(0.003)	(0.003)	(0.004)

Notes: Estimates are based on a Poisson model evaluating the effects of abortion access in quarter t on expected births in quarter t+2. Births are measured for all women aged 15 to 44 (Column 1) and for various sub-groups of women (Columns 2-13) using county-level data for all 254 Texas counties over all quarters between 2009 and 2015. Dagger signifies using overall population of women aged 15-44 as denominator because population estimates for the relevant sub-group are not available. All models include county fixed effects, year fixed effects, demographics, the unemployment rate, an indicator for the presence of a family planning clinic in the county, and this indicator's interaction with post-2012. Standard errors (in parentheses) allow errors to be correlated within counties over time.
*, **, and *** indicate statistical significance at the ten, five, and one percent levels, respectively.

[1] These calculations are described in detail beginning in the section "Burdens Imposed: Women Who Will Forgo an Abortion" of the preliminary injunction order issued on July 2, 2018. This injunction order is available online at https://cases.justia.com/federal/district-courts/arkansas/aredce/4:2015cv00784/102375/142/0.pdf.

[2] All measures of abortion access account for potential travel to clinics in neighboring states.

[3] See *Planned Parenthood* v. *Casey* (1992) and *Whole Woman's Health* v. *Hellerstedt* (2016a), as well as transcripts of oral arguments in *Whole Woman's Health* v. *Hellerstedt* (2016b) in which travel distances are repeatedly discussed.

[4] In a related descriptive study over a similar period of time, Grossman et al. (2017) shows 2012-2014 changes in distance-to-nearest-abortion-facility are negatively correlated with abortion rates across Texas counties. Causal studies examining the effects of different Texas' regulation of abortion providers include Colman and Joyce (2011), Lu and Slusky (2016a), Lu and Slusky (2016b), and Packham (2017).

[5] Since we initially released our study, Fischer et al. (2018) have released an analysis that also leverages variation induced by Texas HB2. Specifically, they estimate the effects of distance using a similar research design and similarly controlling for access to family planning clinics. They find similar effects of distance on abortion rates but do not consider the effects of congestion. Though they find some statistically significant effects on birth rates, we show that there is little evidence of effects overall, which is consistent with our evidence that impacts on abortion rates are too small relative to birth rates to be plausibly detected.

[6] The fraction of medication abortions fell from 27 percent in 2012 to 9 percent in 2014.

[7] At the time HB2 was passed, only 6 facilities in 4 cities – Austin, Fort Worth, Houston and San Antonio—met the standards of an ambulatory surgical center. In response to the law, Planned Parenthood opened an additional facility in Dallas in the summer of 2014 and another in San Antonio in the following year (Martin 2014; Stoeltje 2014) while Alamo Women's Reproductive Services in San Antonio relocated to a surgical facility in 2015 (Garcia-Ditta 2015). These openings and relocations cost $6M, $6.5M, and $3M, respectively.

[8] At the same time the US Supreme Court blocked enforcement of the admitting privileges requirement for the clinics in McAllen and El Paso.

[9] Two days after the Supreme Court struck down HB2, Texas legislators proposed new rules requiring that abortion providers bury or cremate fetal remains. Similar laws have been proposed in Indiana and Louisiana, and could add substantially to the cost of an abortion (Zavis 2017).

[10] One important part of this context is that Texas has a law requiring a 24-hour waiting period after a counseling session before an abortion can be performed for women who live less than 100 miles from the clinic. Texas is not atypical in having such laws: 35 states have counseling requirements, 27 have waiting periods, and 24 hours is the most common waiting period (Guttmacher Institute 2017).

[11] Appendix A contains detailed information on abortion clinic operations in Texas.

[12] See Appendix A for additional details on how these measures are constructed based on clinic operations dates.

[13] These are population-weighted county averages using estimates of the populations of women aged 15-44 (SEER 2016).

[14] See Appendix A for details on how we combine groups of counties in the same commuting zone.

[15] Under the assumptions that this generates classical errors-in-variables in measuring congestion, in which the mean measurement error is zero and this measurement error is uncorrelated with abortion rates, then we can expect this to lead to attenuation bias in the estimated effects of congestion on abortion.

[16] Appendix Figure B1 shows the service regions in each quarter.

[17] Consistent with the evidence provided by our measure, the number of physicians providing abortions in the state dropped from 48 to 28 in the aftermath of HB2 (TPEP 2016). As the number of clinics and providers shrank, wait times to obtain an abortion likely increased. The Texas Policy Evaluation Project documented wait times of three weeks in some Austin, Dallas and Fort Worth clinics (TPEP 2015a) based on telephone surveys.

[18] See Appendix A for specific details for nearby and neighboring states.

[19] It is important to note that it is not clear whether the ideal data would or would not include abortions obtained out of state, since it was indicated in *Whole Woman's Health v Hellerstedt* that a woman's ability to obtain an abortion in Texas was the relevant consideration for whether the Texas laws placed an undue burden on women. For this reason, our estimates focusing only on counties where it is unlikely for many women to seek abortions out of state may have the most legal relevance.

[20] Whereas the sample size is 1,775 for our analysis of abortions overall, we are limited to sample sizes of 552, 611, and 321 for our analyses of abortions at less than 7 weeks gestation, at 7 to 12 weeks gestation, and at more than 12 weeks gestation, respectively.

[21] Like linear models, the Poisson model is not subject to inconsistency caused by the incidental parameters problem associated with fixed effects. While the possibility of overdispersion is the main theoretical argument that might favor alternative models, overdispersion is corrected by calculating sandwiched standard errors (Cameron and Trivedi 2005). Moreover, the conditional fixed effects negative binomial model has been demonstrated to not be a true fixed effects model (Allison and Waterman 2002).

[22] We have also examined standard-error estimates that instead cluster on initial abortion service regions; they are typically very similar or smaller than those that we report.

[23] These data are constructed quarterly; however we use the annual average in our analysis of abortion data, which is not available quarterly. When we examine quarterly birth data, we use the quarterly measures of access to correspond to quarterly birth rate data.

[24] We also illustrate this point in Appendix Figure B2, which plots county-level changes in the average service population against county-level changes in distance-to-nearest-clinic. There is a positive relationship between changes to these measures of abortion-clinic access but the relationship is not strong and there is substantial independent variation.

[25] In this appendix we also explain why the indicator specification is useful for representing the effects in broad strokes but the quadratic specification is better suited to predicting the effects of changes in access.

[26] We have also investigated the counties underlying this variation in greater detail, which we discuss in Appendix B.

[27] Figures B3 and B4 in the appendix show common trends in abortion and birth rates for different subgroups of women. Figure 5 shows similar plots for county demographics (race, ethnicity, age), the unemployment rate, and the number of family planning clinics. Figure B6

plots more-disaggregated trends in outcomes prior to the enactment of HB2. These figures also support the common trends assumption.

[28] Note that percent effects from the Poisson model are calculated as $(e^{\beta} - 1) \times 100$ %.

[29] Appendix Table C1 reports estimates from the model in which abortion rates are evaluated as a quadratic function of both travel distance and the average service population. They are qualitatively similar for the effects of distance, and also indicate non-linear effects of our congestion measure such that increases in the average service population have a larger effect when it is already high.

[30] Appendix tables D1 through D4 show the results of several sensitivity tests. We present results using geodesic ("as the crow flies") distance, using travel time, using an Inverse Hyperbolic Sine Transformation, using alternative approaches to controlling for access to family planning, and excluding various regions or years from the analysis. All of these specifications support the conclusions we reach based on the main analysis.

[31] These estimates are based on our measures of abortion-clinic access in 2012 and the results of the estimated model in Table C1, Column 1. Estimates based on Table C1, Column 1 yield smaller estimated effects of distance and larger estimated effects of the average service population (and a larger effect overall). This occurs because that model is unable to capture the effects of increases in distance that don't cause counties to change distance categories and because it does not account for the non-linear effects of our congestion measure.

[32] We have also confirmed that our estimates are robust to the exclusion of any county that ever has an out-of-state clinic as its nearest clinic during the time frame for our analysis. These results are shown in Appendix Table D4.

[33] Specifically, they indicate that from an initial level of 300,000, a 100,000 woman increase in average service population is predicted to reduce abortions prior to 7 weeks by 25 percent, has no statistically significant effect on abortions at 7 to 12 weeks, and increases second trimester abortions by 14 percent.

[34] A graphical representation of the results of these analyses are shown in Figure C2.

[35] Whereas the sample size for all specifications in Table 2 showing estimated effects for total abortions is 1,775, the sample size is 552 when evaluating abortions at less than 7 weeks gestation, 611 when evaluating abortions at 7 to 12 weeks gestation, and 321 when evaluating abortions at more than 12 weeks gestation. We also note that the Texas DSHS switched from reporting abortions by gestational age to reporting abortions by post-fertilization age in 2014. To make the data series comparable, we adjust the categories using a 2 week difference.

[36] See Appendix Table D5. Also note that we also come to the same conclusions if we evaluate the share of births in each gestational age category using weighted least squares. These results are reported in Lindo, Myers, Schlosser, and Cunningham (2018).

[37] The prescribing information can be found on the U.S. FDA website: http://www.accessdata.fda.gov/drugsatfda_docs/label/2002/19268slr037.pdf.

[38] Survey evidence suggests self-induction rates may be greater for this population (TPEP, 2015b).

[39] We obtained the geographic coordinates of U.S./Mexico border crossings from the Texas Department of Transportation (TXDOT 2017), limiting the analysis to crossings that can be accessed by pedestrians or private vehicles. We then calculated the travel distance from the population centroid of each county to the geographic coordinates of the nearest border crossing.

[40] We assume that the effects on abortion are the same for each quarter in any given year.

[41] These results are shown in Table C2 and depicted graphically in Figure C4. These estimates are consistent with the results shown in Table 4 in that the estimated effects of increases congestion on birth rates are not statistically significant on average.