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UNDUE BURDEN BEYOND TEXAS:
AN ANALYSIS OF ABORTION CLINIC CLOSURES, BIRTHS, AND ABORTIONS IN WISCONSIN

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Undue Burden Beyond Texas: An Analysis of Abortion Clinic Closures, Births, And Abortions in Wisconsin
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ABSTRACT

In this paper, we estimate the impacts of abortion clinic closures on access to clinics in terms of distance and congestion, abortion rates, and birth rates. Legislation regulating abortion providers enacted in Wisconsin in 2011-2013 ultimately led to the closure of two of five abortion clinics in Wisconsin, increasing the average distance to the nearest clinic to 55 miles and distance to some counties to over 100 miles. We use a difference-in-differences design to estimate the effect of change in distance to the nearest clinic on birth and abortion rates, using within-county variation across time in distance to identify the effect. We find that a hundred-mile increase in distance to the nearest clinic is associated with 25 percent fewer abortions and 4 percent more births. We see no significant effect of increased congestion at remaining clinics on abortion rates. We find significant racial disparities in who is most affected by abortion clinic closures, with increases in distance increasing birth rates significantly more for Black, Asian, and Hispanic women. Our results suggest that even small numbers of clinic closures can result in significant restrictions to abortion access of similar magnitude to those seen in Texas when a greater number of clinics closed their doors.

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1 INTRODUCTION

In the last decade, a growing number of state-level policies targeting abortion providers have drastically changed access to abortion and reproductive care in the United States. Since 2010, states have enacted 436 separate restrictions targeting abortion and reproductive health providers (Nash et al., 2016; Nash et al., 2017; Nash et al., 2018b; Nash et al., 2018a). In some states, the effects of these restrictions have resulted in sudden and wide-spread closures of clinics. Texas is a good example of this – when the state passed HB2, a law which mandated that doctors providing abortions must have admitting privileges at a nearby hospital, 14 clinics closed in the span of four months. However, a more common result of targeted restrictions of abortion provider (TRAP) laws is the slow drain of access from a state, in which the increasing burden of many pieces of legislation slowly amps up the costs and regulations faced by abortion providers until the number of abortion clinics in a geographic region dwindles down to one or two clinics.

In this paper, we use policy changes in the state of Wisconsin to estimate the causal effects of clinic closures on abortion and birth rates. Between 2011 and 2013, Wisconsin passed three major pieces of legislation which reduced state funding to family planning clinics affiliated with abortion services, increased regulations faced by abortion providers, and introduced restrictions on how and when women could obtain abortions. Wisconsin is a critical case study for understand how the combination of increasing legislative restrictions and closures of abortion clinics can slowly erode women's access to and use of abortion services, because it allows us to understand the effects of restrictive legislation in medium-sized states which have few clinics even in the absence of TRAP laws. Much of the research on the impacts of abortion clinic closures has been focused on case studies such as Texas in which a single policy closed many clinics. Wisconsin instead began with only five clinics as of 2010, meaning that the policy changes of 2011 to 2013 resulted in a drastic change in abortion access even though only two abortion clinics closed in subsequent years.

Many of the states which have put in place policies targeting abortion providers are more similar to Wisconsin than Texas in terms of size, number of clinics, and demographics. As we see an increasing number of these regulations being challenged in the courts, it is important to evaluate whether legislation that chips away at a small number of clinics has a similar impact on access as legislation that shuts down a large number of clinics in a short period of time. Does a 'death by a thousand cuts,' in which repeated measures to restrict access slowly whittle away the number of abortion providers accessible to women, constitute undue burden?

The undue burden standard was originally established in *Planned Parenthood v. Casey* (1992) and stated that legislation enacted for "the purpose or effect of placing a substantial obstacle in the path of a woman seeking an abortion of a nonviable fetus" is unconstitutional. Recently, the Supreme Court case *Whole Woman's Health v. Hellerstedt* (2016) evaluated the constitutionality of Texas' HB2 law and established a new precedent for the definition of the undue burden standard; the court emphasized that determination of undue burden must weigh the costs of reduced access to abortion against any purported benefits of the law. A growing number of similar cases have been working their way through the appellate courts, including laws that target abortion providers through regulations on admitting privileges of providers (e.g., Louisiana's *June Medical Services v. Gee*), waiting periods between fetal ultrasounds and an abortion (e.g., Indiana's *Planned Parenthood of Indiana & Kentucky v. Commissioner of the Indiana State Department of Health*), and laws banning abortion after 6 weeks (e.g., Mississippi's *Jackson Women's Health Organization v. Thomas Dobbs*). For many of these states, the abortion access landscape mirrors that of Wisconsin's during the period of our study: a small number of clinics that are in danger of closing and leaving the state with only one or two clinics.

The decision in *Whole Woman's Health v. Hellerstedt* (2016) makes clear the importance of empirical evidence on the causal impacts of these laws on people's ability to obtain abortions. When states put in place restrictions on what types of organizations can provide abortions or add regulations meant to make it more costly to obtain an abortion, how many clinics will remain open in a state? If clinics close, do fewer people obtain abortions and does this result in an observable change in the number of births? Do these effects differ in a setting such as Wisconsin in which there were only a small number of clinics prior to the implementation of laws meant to restrict access?

For the purposes of this study, we focus on changes in abortion services access in the state of Wisconsin from 2009 through 2017. In June 2011, then Wisconsin governor Scott Walker signed into law Act 32, which implemented a budget that barred entities that provide abortion services from receiving any state and federal family planning funds. Planned Parenthood, which was the federal Title X grantee in Wisconsin for almost 40 years until 2018 and which administrates the majority of family planning care clinics in Wisconsin, lost approximately $1 million dollars in state funding due to this act. Planned Parenthood clinics also constituted three out of the five state abortion clinics in the state at the time. Act 32 was the beginning of a series of legislative pushes to place financial and bureaucratic restrictions on abortion and reproductive health care providers in the state of Wisconsin, including a large number of restrictions on women seeking abortions and physicians providing abortions in April 2012 through Act 217.[1] Next, in July 2013, Act 37 implemented a series of TRAP laws, including a requirement that physicians have admitting privileges at a hospital within 30 miles and a requirement that women receive fetal ultrasounds.[2] Taken cumulatively, these policies put increased regulatory and financial burdens on providers of both abortion services and other family planning services in the state, particularly Planned Parenthood.

This legislative backdrop provides some context for the resulting changes in access to abortion clinics and other reproductive health clinics. At the beginning of this time period, Wisconsin had five abortion clinics in Wisconsin: Planned Parenthood Madison East, Planned Parenthood Milwaukee-Jackson Street, Affiliated Medical Services- Milwaukee, Planned Parenthood Appleton, and OB/GYN Associates of Green Bay. While Act 32, Act 217, and Act 37 did not shut down any clinics directly, these provisions challenged clinics' ability to keep their doors open. For example, Act 217 mandated that the physician prescribing medication abortion must perform a physical exam, obtain voluntary consent for the abortion, allow a mandatory 24-hour waiting period to pass, and then directly observe the patient take mifepristone. Affiliated Medical Services and Planned Parenthood Madison temporarily halted provision of medication abortions from April 2012 until a court blocked this provision two months later. Both sites did continue providing surgical abortions during this period.[3] In August of 2013, OB/GYN Associates of Green Bay stopped providing abortions

[1]These 2012 restrictions were focused on measures that require multiple, in-person appointments prior to an abortion: physicians must confirm in person that the woman wishes to get an abortion, physicians must provide oral and in-person information about right of refusal and abuse resources 24 hours prior to the procedure, and for drug-induced abortions, physicians must meet with the woman 24 hours prior to the receipt of drugs and then be physically present when the drug is given to the woman the next day.

[2]The admitting privileges component of the law was never implemented in full due to a court injunction shortly after it was passed and a final court decision that this provision was unconstitutional in December 2013. The other provisions, such as the fetal ultrasound provision, were implemented as planned in July 2013.

[3]Prior to Act 217, records from Wisconsin's Department of Health Services indicate that approximately 25 percent of abortions in Wisconsin were medication abortions. This dropped to 14 percent in 2012 after medication abortions were halted and then reinstated.

as part of a buyout by Bellin Health Systems.[4] In October 2015, Planned Parenthood Appleton stopped providing abortions and then closed its doors for good in 2016. Figure 1 shows the locations of the abortion clinics in Wisconsin and surrounding states, as well as their closure status. Additionally, multiple Planned Parenthood sites that do not provide abortions closed between 2011 and 2013 in direct response to the loss of state funds. As a result of these closures, the distance to the closest abortion clinic increased drastically for many women, particularly those in Northern Wisconsin for whom the Green Bay and Appleton sites were previously the only site within 100 miles.

We leverage this geographic variation in where and when clinics closed to estimate the effects of distance from abortion clinics on abortion and birth rates using a difference-in-differences design at the county level. We create a data set of the location and operating dates of abortion clinics in Wisconsin and neighboring states' border counties between January 2009 and December 2017. Following methods pioneered in Lindo et al. (2019) and Fischer et al. (2018), we then use this data to calculate the driving distance from each county's population weighted centroid to the nearest clinic in each month. Driving distance captures one facet of the costs associated with clinic closures; as clinics close, increases in travel time make it more costly in terms of both time and money to obtain an abortion. However, even for those who live in counties that do not experience a change in distance traveled, these clinic closures may affect access to reproductive care by increasing the case load for remaining sites in the state. We therefore also create a proxy for congestion based on the measure from Lindo et al. (2019) which measures the population expected to be served by each clinic.

As previously noted, our paper is not the first to use these methods to test the impacts of restrictive abortion legislation on abortion and birth rates. Economists and public health experts have previously explored this question using the changes in the number of abortion and women's health clinics in Texas in response to HB2 as a case study (Lu and Slusky, 2016; Grossman et al., 2017; Packham, 2017; Quast et al., 2017; Fischer et al., 2018; Lu and Slusky, 2018; Lindo et al., 2019; Myers et al., 2019). Our paper is most similar in methods to Fischer et al. (2018) and Lindo et al. (2019) which both use a difference-in-differences design,

[4]Prior to the buy out, the practice was owned by an abortion provider who sold the practice to Bellin and as a condition of continued employment in the system, the doctor was required to agree not to provide abortion services to any patients. The buy out was not directly linked to a single piece of legislation, but did occur in the same summer as Act 37.

exploiting county-level variation in distance to the nearest clinic post-HB2 to estimate the causal impacts of abortion clinic closures on abortion and birth rates in Texas.

However, limited research exists on recent abortion restrictions outside of Texas. Expansions of access to abortion and family planning have been studied primarily in the context of the changes in access in the mid twentieth century through the roll out of Title X clinics (see Bailey, 2013 for an overview of the literature), the legalization and modernization of contraception (e.g., Bailey et al., 2012), and the legalization of abortion first at the state level and then nationally through *Roe v. Wade* (e.g., Myers, 2017). As more and more states have passed laws intended to restrict access to abortion, a smaller body of research has documented the effects of state variations in laws targeted at individuals seeking abortions, such as waiting periods and parental consent laws for minors, with earlier literature in this area findings null impacts of the waiting periods on abortions and more recent papers showing declines in abortions associated with waiting periods for teenagers seeking abortions.[5] However, the causal literature on TRAP laws – legislation targeting clinics with the intention of increasing regulations and costs associated with provision of abortions – is less developed. This literature has almost entirely focused on legislation in Texas – an undeniably large state and an important case study, but also a state that differs in size and context from many other states. One exception is Lu and Slusky (2016), which finds that women's health clinic closures are associated with lower rates of preventative care take up in both Texas and Wisconsin, but does not estimate the effects of abortion clinic closures on health care or reproductive outcomes in Wisconsin.

The focus on Texas as a case study is partially because Texas' legislation was one of the earliest and more high profile examples of a TRAP law. Additionally, the state of Texas provides high quality annual data at the county level on abortions; the policies' implementation dates and the clinic closure dates provide a clean natural experiment; and Texas's funding provisions for family planning are separate from abortion, allowing researchers to isolate impacts of the two services separately. Nonetheless, there are features of this policy landscape that are drawbacks when considering extrapolation to other states' policy settings. Here we highlight three reasons to conduct a similar analysis in another setting.

[5] For a review of the literature on waiting periods, see Joyce et al. (2009) and Bitler and Zavodny (2001). For a review of the literature on parental consent laws, see Joyce et al. (2019).

First, Texas' implementation of HB2 garnered a large amount of both popular media and research attention partially due to the large number of clinic closures associated with the law. Between the month the law passed and when the law went into effect (i.e., July to November 2013), 14 out of 42 abortion clinics in Texas closed (Lindo et al., 2019). If a state experiences a closure of only one or two clinics from a relatively smaller pool, it is not as clear whether we should expect the same additional burden on the affected population. Many of the states whose legislation targeting abortion providers are currently making its way through the circuit courts have less than five clinics (e.g., Mississippi, Alabama, Arkansas). In this paper, we focus on a case study of a state more similar to these settings: Wisconsin, which had five abortion clinics in 2011 and by 2017, had only three in response to legislative restrictions on funding and facilities.We test the effects of clinic closures in a setting where, prior to legislation, many residents already faced limited geographic access. In doing so, we are able to demonstrate that even laws that close only two clinics are also potentially in violation of the undue burden standard, particularly if these clinics are in the more geographically remote areas of the state.

Second, by using Wisconsin as a case study, we are able to estimate causal estimates in a different demographic context than Texas can provide. Texas is second only to New Mexico in the proportion of its residents who are Hispanic. As of 2016, 39.4 percent of Texas residents are of Hispanic or Latinx ethnicity. The United States as a whole is only 18.1 percent Hispanic and over half of all states have Hispanic populations that make up less than 10 percent of all residents. In the context of research on reproductive policies, Latinx heritage is a particularly important demographic characteristic to consider given significantly different fertility rates among Hispanic, White, and Black Americans. As of 2017, the total fertility rate (i.e., the expected number of lifetime births per 1,000 women) for Hispanic women was 2006.5 births compared to 1666.5 for White women and 1824.5 for Black women (NVSS, 2018). Additionally, analyses in Texas find that the negative effect of distance from the nearest clinic on abortion rates is stronger for Hispanics than Non-Hispanic Whites (Lindo et al., 2019), suggesting that researchers should take caution when applying average treatment effects from Texas to demographically dissimilar environments. Wisconsin's population, in contrast, is 81 percent White, 6 percent Black, 7 percent Hispanic, 3 percent Asian, and 3 percent other races (Census, 2018). Wisconsin over represents the White population relative to the country as a whole, but is more representative of the states where abortion access is most under legislative threat (e.g., Midwestern states such as Missouri or Indiana or Southern states such as Arkansas).

Lastly, there is value to exploring whether the effects seen in Texas carry over to smaller states where individuals can more easily cross state borders to access abortion providers. On one hand, Texas' size makes the natural experiment of HB2 more convincing; it would be very difficult for many people to travel outside the state to obtain an abortion, meaning that the decline in abortions cannot be explained as individuals substituting to providers in other states. However, though Texas' size is beneficial from a research design stand point, it also represents another way in which Texas differs from the average state facing this legislation. Our study is able to show not only that abortions decline in Wisconsin as clinics close, but also that the number of abortions to Wisconsin residents in other states do not increase enough to fully explain the declines. Minnesota is the state that is closest to the Wisconsin counties for whom the clinic closures most decreased access; however, abortions taking place in Minnesota to Wisconsin residents are declining over this time period rather than increasing. The second closest state (Illinois) has seen increases in the number of out of state abortions, but the increases for abortion to Wisconsin residents are not large enough to explain the declines of in-state abortions we see in Wisconsin. Moreover, the Illinois clinics are further away from affected counties than remaining in-state clinics. We also see large declines in the number of out of state abortions obtained by residents of Michigan's Upper Peninsula which was previously served by clinics in north-eastern Wisconsin. These analyses demonstrate the substitution to out-of-state care we might expect in medium-sized states that implement TRAP legislation.

Our goal in this paper is to test whether the case study of Texas is generalizable to other states. This is particularly important when using the estimates from Texas to extrapolate the effects of future policies. For example, Myers et al. (2019) uses the parameters estimated in Lindo et al. (2019) to predict the impacts of clinic closures associated with possible future judicial decisions across the country. Their estimates rely on the assumption that all states would behave as Texas behaved in response to these laws and that the effects of clinic closures on women's use of abortion services would be the same in states other than Texas. Our paper tests that assumption by applying the methods of Lindo et al. (2019) and Fischer et al. (2018)

to Wisconsin.

We find that the small number of clinic closures in Wisconsin produced comparative if not more pronounced effects to those seen in Texas. Our analysis shows that a 100 mile increase in linear distance to the nearest abortion clinic is associated with a 24.9 percent decline in abortion rates. However, we also see evidence of non-linearity in the relationship between distance and abortion rates; in a non-parametric specification, we see that clinic closures that change a county from being less than 25 miles to the nearest clinic to being 25 to 50 miles are associated with a 20.9 percent decline in abortion rates in those counties. Counties which experience a distance increase from being less than 25 miles to being 50 to 100 miles from the nearest clinic see a 25.9 percent decline in abortion rates and those that increase to being further than 100 miles from a clinic see a 31.7 percent decline. This is consistent with past findings in similar studies of clinic closures in Texas which found that moving from being within 25 miles of a clinic to being 25 to 50 miles is associated with a 21.9 percent decline (Fischer et al., 2018) or that moving from being less than 50 miles from a clinic to being between 100 to 150 miles is associated with 35.9 percent lower abortion rates (Lindo et al., 2019).

Our estimates of abortion rate changes are also large enough as a proportion of the state birth rates that we are able to identify a concurrent increase in births. In Lindo et al. (2019), they are unable to reject the null hypothesis that abortion clinic closures had no effect on birth rates and show that the predicted decline in abortions would be too small to detect an effect even if all non-abortions concluded in a birth. Fischer et al. (2018), however, do estimate the effects of closure on births and find positive effects in a parsimonious regression model without time-varying controls such as unemployment rates or demographic trends. These effects are no longer significant in the presence of covariates. In the Wisconsin case, we see positive impacts even with controls; moving from being within 25 miles of a clinic to being more than 100 miles of a clinic is associated with approximately 3.3 percent higher birth rates.

This paper therefore provides additional support for the results seen in the existing literature on Texas' reproductive health access restrictions and suggest that distance to the nearest clinic has similar impacts in states with few clinics remaining to take on the service burden. Despite the fact that Wisconsin had fewer clinics close in response to the legislation, the small number of remaining clinics meant that even the closure of only two clinics reduced access by similar amounts to the closure of fourteen clinics in Texas. However, we do not see the same negative impacts of congestion in the context of Wisconsin. While Lindo et al. (2019) found that increased service population for remaining clinics resulted in lower abortion rates, we see no significant association between population serviced and abortion, despite Wisconsin's clinics having a larger service burden on average than Texas clinics both before and after closures. Lastly, we are able to conclusively document that this pattern holds in a very different demographic context, as well as demonstrating that the effects of such policies are more restricting for non-White residents.

The remainder of the paper is organized as follows. Section 2 describes the data used, as well as describing the policy setting of Wisconsin. Section 3 describes our empirical strategy for estimating the effects of clinic closures on abortion and birth rates. Section 4 describes our results, including a discussion of tests of our identifying strategy, our main results, and a set of heterogeneity analyses. Finally, we conclude and offer policy implications of our findings in section 5.

2 DATA

2.1 Measures of Abortion Clinic Access

To evaluate the effects of Wisconsin's restrictions on abortion provision on births, we create a database of all abortion clinics operating in Wisconsin and the bordering states (Minnesota, Iowa, Illinois, and Michigan) between 2009 and 2018 using a variety of sources to verify the dates of openings and closings between the years of 2009 and 2018. Because of the small number of clinics in Wisconsin, the closings of both the Green Bay and the Appleton clinics received a good amount of press coverage and we are able to use archival versions of the clinic websites, newspaper articles, and websites tracking clinic operations maintained by both pro- and anti- abortion advocacy groups to collect sources for dates of openings (for clinics that open post-2009) or confirmation that the clinic was in existence and providing abortions in the year 2009. For bordering states with larger numbers of clinics, we use the currently existing clinics in states within 100 miles of the Wisconsin border as the starting point and use the same methods to verify dates of operations for these clinics. We then use archived versions of the Planned Parenthood website, the National Abortion Federation website (pro- abortion advocacy group), and AbortionDocs.org (anti-abortion advocacy group) to find any abortion sites in bordering states that have closed in the relevant time period and verify the dates of openings/closings for those sites.

We then use this clinic database to construct a county-level measure of abortion access based on the distance to the closest abortion clinic for each month in the period 2009 through 2017. To calculate the closest clinic, we use the Stata program georoute (Weber and Péclat, 2017) to calculate the driving distance between the population-weighted centroid of the county (as calculated by the Census based on 2010 population count) and each of the 32 clinics in our data set and rank them from closest to farthest. Then, for each month, the distance to closest clinic is assigned based on the shortest distance to an open clinic, using the opening and closing dates collected in the data described above. For annual analyses, a clinic is considered open if it is open for at least six months of the year, coded based on whether the clinic is open in July of a year.

Figure 2, panel A shows the average distance to the nearest abortion clinic for each county, weighted by the population of women age 15 to 44 in the county, from January 2009 to June 2017. The two red lines correspond to the closing of OB/GYN Associates of Green Bay in 8/2013 and Planned Parenthood Appleton in 10/2015. Though the Green Bay site closing only slightly increased average distance traveled, this was due to the fact that the Appleton site was still open in Northeastern Wisconsin. Once that closed, there was a spike in average distance from around 35 miles to around 55 miles – an increase comparable to the change in Texas at the time of HB2 from 21 miles in the quarter prior to HB2 to 44 miles in the quarter afterwards (Lindo et al., 2019).

We next calculate the average service population per clinic by assigning each counties' population of women to the nearest clinic in each month between January 2009 and June 2017. For the two service areas with multiple clinics (Milwaukee and St.Paul/Minneapolis), we combine all counties nearest to one of the clinics in that service area and then divide by the number of clinics in the service area:

$$\text{Service Population}_{it} = \frac{\sum_k \text{ (Pop., Women 15-44 in county } k)\mathbb{1}(\text{service region } i \text{ closest to county } k \text{ in time } t)}{\text{Number of clinics in service region } i}$$

Figure 2, panel B shows the average change in service population for the nearest clinic to each county with the two red lines again representing clinic closings. By the end of the period, the average service population served for each county's nearest clinic increased by approximately 100,000 women, going from around 181,000 women per clinic to 283,000 in June 2017. For comparison, following the HB2 legislation in Texas, the average service population rose from around 146,000 to 262,000 (Lindo et al., 2019). Wisconsin's population per clinic was larger than Texas' pre-closures and the closures resulted in approximately the same end-level of burden per clinic in the two states.

Figure 3 and figure 4 describes the spatial variation in the effects of these closures. In figure 3, the gradient of colors represent the change in distance to nearest clinic between 1/2009 and 6/2017 for each county, with darker colors indicating larger changes in distance to the nearest clinic. Northeastern Wisconsin has the largest change in distance to the nearest clinic, with increases of over 100 miles of travel.[6] Western Wisconsin actually experienced a slight decrease in distance due to Planned Parenthood– Rochester, Minnesota opening in late 2015. Figure 4 shows the size of average service population for the nearest clinic to each county in 2009 (panel A) and in 2016 (panel B) after the closures. The map demonstrates not only the increase in service population, but also the dearth of clinics available after the closure of the Green Bay and Appleton clinics. While the 2009 map shows many distinct clinic population levels, indicating that different regions of the state were served by different clinic populations, the post-2016 map only has four gradations because there are only five clinic clusters near to Wisconsin: Madison (1 clinic), Milwaukee (2 clinics), Upper Peninsula Michigan (1 clinic), Duluth, Minnesota (1 clinic), and Minneapolis/St. Paul (3 clinics).

2.2 Outcome Measure #1: Abortion Rates in Wisconsin

Abortion data comes from the Wisconsin Department of Health Services which requires all medical facilities in Wisconsin to report select information on all patients who obtained induced abortions, including state and county of residence, age, marital status, race, and education. Wisconsin DHS provides annual reports by county of residence of counts of abortions obtained by Wisconsin residents in Wisconsin (Garcia-Lago,

[6] Due to a Planned Parenthood clinic starting to offer abortions in January 2015 in the Upper Peninsula of Michigan, counties closest to the state border had a small decline in distance traveled.

2017; Ninneman, 2012). All counts are bottom-coded, with any counties reporting 0-5 abortions listed as "<5" and counties reporting 0 abortions listed as "0". We impute the bottom-coded values to be 3 for all counties that do not report a 0 count of abortions. We use abortion counts from 2009 to 2017 to construct an annual abortion rate for each county, defined as the number of births per 1000 women ages 15 to 44 in a county. Population levels for each count are defined as before.

Figure 5, panel A shows the change in abortion rates for each of the full public health regions of Wisconsin, as defined by the Wisconsin Department of Health Services. Abortion rates decreased the most in the Southeastern region which contains Milwaukee (-2.1 abortions per 1000 women decrease) and the Eastern region which contains Green Bay and Appleton (-1.97 abortion decrease). They decreased by the least in the Western region (-0.6 abortions per 1000 women). Note that abortions decreased the least in the places birth rates increased the most (panel B) and vice versa, suggesting that abortion rate decreases are not necessarily translating to higher birth rates. That said, the regions where abortions decreased most are the two regions closest to the abortion clinic closures.

2.3 Outcome Measure #2: Birth Rates in Wisconsin

This analysis uses the full universe of births to women ages 15 to 44 in Wisconsin as reported in the restricted-use natality files provided by the National Center for Health Statistics in the National Vital Statistics System (NVSS) from June 2009 through December 2017 (NVSS, 2018). We restricted to this period to match the years we have collected data on openings/closings of abortion clinics (Jan. 2009- Dec. 2017), forward dated by six months. Since the relevant time period for when the clinic is open is during the pregnancy not at birth, we assign a birth in month-year t to period $t-6$ for matching to the month-year that we observe a clinic's status to reflect access to abortion at approximately 13 weeks into the pregnancy.

A birth is assigned to a county based on the county of residence reported by the mother. Monthly and annual birth rates are then calculated as the number of births per 1000 women ages 15 to 44 in a county. Population levels for each county (the denominator of the birth rate) are taken from yearly county estimate from the National Institute of Health's Surveillance, Epidemiology and End Results (SEER) U.S. population data (SEER, 2018). This leaves us with a data set of month-county observations from January 2009 through June 2017 and year-county observations from 2009 through 2016.[7]

Figure 5, panel B shows the change in birth rates for each of the full public health regions of Wisconsin, as defined by the Wisconsin Department of Health Services. Birth rates increased the most in the Northeast, by 5.1 births per 1000 women in the Northern region and 1.27 births per 1000 women in the Western region. They increased by the least in the Southeastern region which contains Milwaukee (0.71 birth increase) and the Eastern region which contains Green Bay and Appleton (0.91 birth increase).

2.4 Access to Family Planning Clinics

During the period of interest of this study, access to family planning clinics that did not provide abortions was also changing. The primary organization that provides subsidized family planning care in Wisconsin is Planned Parenthood, which was the Title X grantee for the entire time period of the study.[8] Due to the state funding cuts, Planned Parenthood WI closed down five clinics that did not provide abortions: Chippewa Falls Health Center, Beaver Dam Health Center, Johnson Creek Health Center, Shawano Health Center, and Fond du Lac Health Center.

To control for variation over time in access to family planning care, we collect a data set containing the location of every Planned Parenthood within Wisconsin borders during the period of the time period and the opening/closing date for each clinic if it occurred between 2009 and 2018 using newspaper coverage of openings/closings and archival versions of the Planned Parenthood website to confirm existence of the clinic and the street address of the clinic. We then use Google maps to confirm the latitude and longitude of each clinic based on this address and assign each clinic to a county. We then construct our measure of access as the number of Planned Parenthood clinics open within a county in each month, where a clinic is considered closed in a month if the closing date was during that month. For annual measures, we characterize a clinic as open during a year if it was open for at least six months during that year, measured based on whether it

[7] Due to backdating, there is only half a year of data for 2017, making the annual birth rate non-comparable to other full-years.

[8] Until the federal legislation barring Title X patients from receiving care at Planned Parenthood in 2018, Planned Parenthood had been the Title X grantee for decades and served 87% of the state's Title X patients.

was open in July of that year.

2.5 Additional Data

We also use supplementary data on county-level characteristics that vary over time. Data on seasonally unadjusted unemployment rates by county and month are from the publicly-available Bureau of Labor Statistics Local Area Unemployment Statistics data from January 2009 through December 2017. Population counts by demographic group (age, race, gender) are taken from yearly county estimate from the National Institute of Health's Surveillance, Epidemiology and End Results (SEER) U.S. population data SEER, 2018. Per capita income (PCI) comes from the publicly-available U.S. Bureau of Economic Analysis Local Area Personal Income accounts, 'Annual Personal Income by County.' Because PCI is measured at yearly level, we interpolate between years for any analyses at the monthly level, assigning PCI in year t to year t month 1 and then interpolating across months linearly

Summary statistics for these covariates, as well as for the measures of access, births, and abortions, are presented in Table 1.

3 EMPIRICAL STRATEGY

For this analysis, we rely heavily on the empirical strategies used in Lindo et al. (2019) and Fischer et al. (2018) in their analyses of the effects of clinic closings in Texas to allow for better comparison of the impacts of Wisconsin's legislation on abortion access to the impacts in Texas.

First, following the analyses in Fischer et al. (2018), we compare trends in abortions or birth rates to at the state-level to trends in the same time period for other states. Fischer et al. (2018) conduct a synthetic control method analysis of Texas' HB2 implementation and find a significant increase in monthly fertility rates (effect size =0.170, one-sided p-value = 0.04) in their synthetic analysis. We therefore employ a synthetic control method in the context of Wisconsin, using the method described in Abadie et al. (2015). Figure 6 and figure

7 show the time-series patterns births and abortions, respectively, for Wisconsin from 2009 to 2017 alongside a synthetic control for Wisconsin (see figure notes for construction of synthetic Wisconsin). We see a lower abortion rate in Wisconsin from 2014 to 2017 than the synthetic Wisconsin (Average treatment effect = -0.48 abortions per 1000 women) and a larger birth rate, particularly following the closure of the second clinic (ATE = 0.10 births per 1000 women for 1/2016 to 6/2017). Though the overall patterns are supportive of the hypothesis that the closure of abortion clinics in Wisconsin led to higher birth rates and lower abortion rates, the effect sizes are non-significant based on the p-values calculated using permutation-based inference Abadie et al. (2015).

One possible reason for why we do not see significant effects in the synthetic analysis is that the permutation-based inference relies on the assumption that other states do not themselves have policies going into place during this time period affecting their birth and abortion rates. Implementing placebo "treatments" are only valid "placebos" if there truly are no changes happening in the other states. In reality, 17 states passed a total of 57 new abortion regulations in 2015 alone. Five out of the seven "placebo" treatments that had larger effect sizes in the birth rate analysis were states that had increasing restrictions on abortion access during the 2013 to 2015 time period: North Dakota, Alabama, Arkansas, Kansas, and Kentucky. Because of the nation-wide fluctuations in abortion access over this time period, state-level synthetic analyses are limited in their ability to causally estimate the impacts of Wisconsin's legislation.

We therefore turn to the methods employed in both Lindo et al. (2019) and Fischer et al. (2018). In this method, we estimate the effects of access to abortion clinics on abortion and birth rates using a generalized difference-in-differences design, which uses within-county variation over time in distance to a clinic, controlling for cross-county time-varying shocks. This method identifies the true effect of access to clinics based on the assumption that changes in abortion and birth rates for counties with small changes in distance over time are a reasonable control group for counties with larger changes in distance. That is, the trajectory of abortions and births over time for counties with small changes in distance to the nearest clinic is the path that they would have taken in counties with large changes in the absence of the clinic closings.

Following Lindo et al. (2019), we operationalize this strategy with a Poisson model of abortions. The authors note that they use this method because abortions are discrete counts and the frequent small/zero counts of abortions in some counties in some time periods make a Poisson model a more natural model than a linear regression model of abortion rates.[9] For regressions using counts of births and abortions as the outcome, we use the following estimating equations :

$$\begin{aligned}
\mathbb{E}[\text{abortion count}_{c,t}|\text{dist}_{ct}, \alpha_c, \theta_t, X_{ct}] &= exp(\beta_1 \text{dist}_{ct} + \alpha_c + \theta_t + X_{ct}'\beta_2) \\
\mathbb{E}[\text{birth count}_{c,t+6}|\text{dist}_{ct}, \alpha_c, \theta_t, X_{ct}] &= exp(\beta_1 \text{dist}_{ct} + \alpha_c + \theta_t + X_{ct}'\beta_2)
\end{aligned}$$

where $\text{abortioncount}_{c,t}$ is the number of abortions in a county c in year t and $\text{birth count}_{c,t+6}$ is the number of births in a county c in month $t+6$, where t indexes the month-year of the abortion clinic's open-close status. For all abortion analyses, the time period is yearly; for birth analyses, we run regressions at the month-year level back-dated by six months. dist_{ct} is a dummy variable equal to a distance measure of the closest clinic to county c's population-weighted centroid at time period t, where the measures include a linear measure of distance, a quadratic of distance, and distance bin dummies equal to one if the closest clinic is a given distance away. ($\mathbb{1}(50 > \text{dist} > 25 \text{ miles})$, $\mathbb{1}(100 > \text{dist} > 50 \text{ miles})$, $\mathbb{1}(> 100 \text{ miles})$). α_c are county fixed effects; θ_t are time period fixed effects. X_{ct} contains time-varying county characteristics including unemployment rate, per capita income, number of women in five year age bins from 15 to 44 (e.g., 15 to 19, 20 to 24, etc.), and population counts by race (white, Black, Asian, and Native Americans). We also control for the number of Planned Parenthood clinics open per county in the time period.[10]

We also include log of county population of women 15 to 44 and constrain the coefficient to be one, following (Fischer et al., 2018). This is equivalent to Lindo et al. (2019)'s decision to use abortion rate as the outcome in a Poisson model: the log of rate is that same as the log of the county minus the log of the population. For comparison and a possibly more easily interpretable outcome variable, we also report a linear fixed effect

[9] They also note that this modeling decision results in consistent estimates in a D-in-D design, saying, "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). "

[10] This control differs slightly from the control for family planning access used by Lindo et al. (2019), who controlled for whether a county had a publicly funded family planning clinic, not a Planned Parenthood. In Texas, Title X funds recipients were not primarily Planned Parenthoods. In Wisconsin during the period of the study, all family planning clinics that received Title X funds were either Planned Parenthoods or a small number of contracted clinics determined by Planned Parenthood. While we do not control for the contracted clinics, our measure covers the substantive bulk of clinics receiving Title X funds.

model of abortion rates and birth rates in the appendix, regressing these measures on the same covariates as in the Poisson model. In both cases, rates are defined as the number of births (abortions) per 1000 women 15 to 44 in a county. Coefficients in the Poisson model can be interpreted as the percent change in the birth (abortion) rate; coefficients in the linear model can be interpreted as the level change in the birth (abortion) rate.

In each of the regression models, we model the effects of distance first assuming a linear effect of distance, then a quadratic, and then finally in a non-parametric manner with bins of distance: 25 to 50 miles from the nearest clinic, 50 to 100 miles from the nearest clinic, and >100 miles from the nearest clinic. One might not think that the first mile of distance has the same effect of ease of access as the 50th or the 100th additional mile traveled. By using the binned measures of distance, we are better able to illustrate how far is too far when considering the additional travel burden that a clinic closing induces. The percentage of counties in each of these bins is reported in Table 1.

4 RESULTS

In this section, we discuss the causal impacts of abortion-clinic access on reproductive outcomes in Wisconsin.

4.1 Identification

Before discussing the regression analyses, we first evaluate whether the assumptions underpinning our identification arguments are valid. Our causal claim rests on the idea that the only thing changing at the exact time of the clinic closures that impacted births and abortions was the distance to the nearest abortion clinic and the subsequent congestion at the remaining clinics due to increased service populations.

To assess this claim, we first look at birth rates and abortion rates over time for four types of counties: counties which experienced no positive change in distance to the nearest clinic between 2009 and 2017; counties in the bottom tercile of distance changes (i.e., <33 mile increase), counties in the middle tercile (i.e., between 33 and 95 mile increase), and counties in the top tercile (i.e., > 95 mile increase). Panel A of figure 8 shows the average travel distance over time for each of the groups. As demonstrated in this figure, distance to the nearest clinic was unchanged for all groups prior to 2013 when the Green Bay abortion clinic closed, suggesting that pre-2013 trends in birth rates and abortions should be similar across groups if the common trends assumptions underpinning the differences-in-difference design holds.

Panel B and C of figure 8 shows the trends in monthly log birth rate and annual log abortion rate for each tercile of distance change and the no-distance change group, respectively.[11] In both cases, the trend for the four groups of counties are similar prior to 2013, with the exception of the middle tercile for birth rates. This is suggestive evidence of the common trends assumption that birth rates and abortion rates would have followed similar patterns post-2013 in the absence of the clinics closing.

Post- clinic closings in 2013, there is divergence of trends for the groups in the expected direction for both measures, particularly when comparing the group with no change in access to the group with the biggest change in distance. For abortions, there is a slight downward trend for all county groups prior to clinic closings, which then reverses after 2013 for the counties with no change in distance and continues to decline for counties with large changes in distance. For births, trends are fairly flat prior to 2013, but begin to increase slightly post-2013 for counties experiencing changes in distance with the largest incline for the counties in the top tercile of distance change. Though these descriptive analyses are only crude measures of the impact of closures, this is suggestive evidence that the clinic closures in Wisconsin had an impact on both births and abortions.

4.2 Abortion Analysis

Having established evidence in support of our identifying assumptions, we next move to our primary results from the difference-in-differences analysis. The first outcome we look at is annual county-level abortions. Table 2 shows the Poisson model with abortion counts as the outcome; Appendix Table A-1 shows the linear

[11] We use log of the rates rather than the rates themselves to make the graphs comparable to what Lindo et al. (2019) present. For monthly birth rates, we smooth the seasonal trends by taking a 12 month moving average for each month before taking the log.

model with abortion rates as the outcome. In both tables, column 1 reports the linear effect of distance on birth rates; column 2 reports the quadratic effect of distance. Column 3 reports the distance bin dummy regressions. As demonstrated in these final 3 columns, there are strong non-linearities in the effect of distance, with larger impacts for counties who switched from being less than 25 miles from a clinic to being more than 100 miles form a clinic.

In our Poisson specification, a one-mile increase is associated with approximately 0.25 percent fewer abortions per county.[12] In the quadratic version, however, we see that there is a non-linear relationship with each additional mile increasing the cost at a diminishing rate. Figure 9 plots the effect of a 50 mile increase in distance to the nearest clinic conditional on starting distance; we see stronger declines in abortion rates at lower base distances with the increase no longer significantly affecting abortion rates if a county starts 75 miles away from the nearest clinic. This is consistent with diminishing marginal costs to an additional mile.

This non-linear pattern is then verified in the distance bin analyses which show that even a county that moves from being less than 25 miles from a clinic to being 25 to 50 miles from a clinic experiences an average decline of 20.4 percent, whereas a county that is now 50 to 100 miles from a clinic experiences an average decline of 25.9 percent. Moving to not having a clinic within 100 miles decreases abortions only marginally more than 50 to 100: a decline of 31.7 percent. In appendix figure A-1, we plots the coefficients for each distance bin for annual abortions; appendix figure A-2 does the same for monthly births.

Similar patterns are seen in the linear fixed model, with a 100 mile increase corresponding to a 0.613 lower abortion rate per 1000 women. We also see similar effects in the non-parametric specification, with further distance changes being associated with larger declines in abortion rates. We estimate that county that began within 25 miles of a clinic and became more than 100 miles from a clinic due to the closures would have 1.2 fewer abortions per 1000 women.

The non-linear pattern with lower marginal effects further from a clinic is consistent the pattern seen in Texas analyses and of comparable, if not larger, magnitudes. In both the Lindo et al. (2019) and the Fischer et al.

[12] Note that coefficients in a Poisson model can be thought of as the effect of a one-unit increase in the independent variable in terms of log-units of the outcome variable; an increase in log-abortions can be approximated as a percentage change.

(2018) paper, further distance increases were associated with larger declines in abortion rates. Moreover, the effect sizes we see are remarkably similar to those seen in Texas. Fischer et al. (2018) use the same distance bins and find that moving from less than 25 miles to being 25 to 50 miles from a clinic is associated with a 16 percent lower abortion rate, 50 to 100 miles is associated with a 17 percent lower abortion rate, and more than 100 miles is associated with a 22 percent lower abortion rate. Our effect sizes are larger than those seen in Fischer et al. (2018), but smaller than the effects seen in Lindo et al. (2019) for very long travel distances (>200 miles).

4.3 Births Analysis

Though we see a significant decline in abortions in response to clinic closings, this may not necessarily translate to lower birth rates. Women may go out of state to obtain an abortion, which we are unable to see in our data. Though Minnesota abortion rates are fairly constant during the time period of our sample, Illinois has experienced a large increase in the number of out of state patients obtaining abortions. Due to data limitations, we cannot know how much of that increase comes from Wisconsin women seeking treatment. Additionally, women may find methods to self-induce abortions if they do not have access to official channels to obtain an abortion. For comparison, in Lindo et al. (2019)'s analysis of Texas clinic closures, they concluded that even if all the 'missing' abortions resulted in births six months later, the magnitude of the effect would not be statistically different from zero. In contrast, Fischer et al (2018) do find a significant effect, with births increasing by 3 percent in counties without a clinic within 50 miles.

Before turning to regression analysis to answer this question, we can do a back of the envelope calculation based on the abortion estimates to approximate what we might expect the maximum effect on birth rates to be given our estimates of abortion declines. To do this, we use the same method described in Fischer et al. (2018). We take each of the bin estimates from column 3 of Table 2 and use the ratio of abortions to births in Wisconsin pre-2013 in those distance categories to calculate how many more births there would be if all 'lost' abortions post-2013 became births. We then multiply the magnitude of the bin estimate coefficient for each group by the ratio of abortions to births in that group. This gives us the percent increase in the number of births for being a clinic in that distance bin, assuming that in the absence of the clinic closures the ratio of abortions to births would have remained the same and the only increase in births comes from missed abortions.

For counties that were in the 25 to 50 mile bin in 2017, there were 3,096 abortions from 2009-2012 and 42,404 births in the same time period. This gives us an abortion to births ratio of 0.073, which when multiplied by 0.204 gives us a potential increase in the birth rate of 1.49 percent for having a clinic 25 to 50 miles away if all missed abortions convert to births. For the 50 to 100 mile bin, there were 3387 abortions from 2009-2012 and 56,587 births in the same time period. The abortion to births ratio is then 0.060, which combined with the abortion coefficient estimate gives us a maximum birthrate increase of 1.55 percent. Lastly, from 2009-2012, there were 4799 abortions in counties which were more than 100 miles from a clinic in 2017 and 61331 births. The abortion to birth ratio is then 0.078 and the expected increase in birth rates is 2.5 percent.

When we repeat the difference-in-differences analysis with births as the outcome, we do see a significant effect of distance to the nearest clinic on Table 3 and Appendix Table A-2 report the regression results respectively from the Poisson model of monthly births and the linear model of monthly birth rates as a function of three different measures of travel distance to the nearest clinic. As in the abortion analysis, column 1 reports the linear effect of distance on birth rates; column 2 reports the quadratic effect of distance. Column 3 reports the distance bin dummy regressions. We again see strong non-linearities in the effect of distance, with the primary effects coming through large changes (i.e., an increase of more than 100 miles).

Column 1 of Table 3 reports that a 100 mile increase in distance from the nearest clinic is associated with a 3.71 percent increase in the number of births per month. The bin estimates of distance change also show non-significant and small effects for shorter distances, consistent with the smaller effect sizes predicted in the back-of-the-envelope estimates. The bin for counties further than 100 miles from a clinic show a larger effect size than we would expect based on the back-of-the-envelope calculation; having the closest clinic be further than 100 miles away is associated with 3.3 percent higher birth rates compared to our estimate of 2.7 percent higher birth rates. However, the 95 percent confidence interval for the estimates ranges from [0.007, 0.063], meaning we cannot reject the hypothesis that this estimate is significantly different from the

2.7 percent predicted increase predicted by the back-of-the-envelope calculation.

The linear model results in noisy predictions for which we cannot reject the null hypothesis for either the linear effect of distance on birth rates or the non-parametric bin estimates.

Why might we see these larger effect sizes? For one, we are using a more exact measure of closures in the monthly analysis of births than we are able to use in the analysis of annual abortion counts. In the annual analysis, we classify a clinic as open in any year where it is open for at least the first half of the year, even though it may be closed for the latter half of the year, biasing our estimates of the effect of a closure downward.

A surprising aspect of the results is that the quadratic estimates suggest that there are increasing marginal effects of distance, which at first glance seem inconsistent with the diminishing marginal effects we saw on abortions. Figure 10 plots the effect of a 50 mile increase in distance to the nearest clinic conditional on starting distance; we see higher increases in births for counties that started further from a clinic. How do we square these results with the results we see for abortions? To understand this seemed conundrum, it's important to be clear about what this graph is showing: the percent change in the the number of births per 1000 women. It is possible for there to be symmetric increases in the number of births as there are decreases in the number of abortions and have differentsecond derivatives of the quadratic function of distance. [13]

4.4 Service Population Analysis

In addition to clinic closures resulting in greater travel distance to the nearest clinic, we also would expect that the number of women served by the remaining clinics would increase, resulting in greater clinic congestion and less access to abortion services even for women who do not experience a change in travel distance.

[13] This is demonstrated in the back of the envelope calculation based on the non-parametric analysis, in which we assume that all "lost abortions" become births. The abortion to birth ratio is decreasing as we move from the 25 to 50 bin to the 50 to 100 bin and then increasing as we move from the 50 to 100 mile bin to the $>$ 100 mile bin. This results in a smaller predicted increase in the percent increase in birth rates from the 25 to 50 distance bin to the 50 to 100 bin (1.55-1.49 = 0.06) than from the 50 to 100 bin to the $>$ 100 bin (2.5-1.55 = 0.95). This holds even though the corresponding increases for percent change in abortion rates in these bins has a similar marginal effect in terms of percent change as distance increases going from -0.20 to -0.26 ($\Delta = 0.6$) and then from -0.32 to -0.26 ($\Delta = 0.6$). Even though the count of abortion declines and birth increases are the same in the back of the envelope, the second derivative of the percent change in the rates have different signs.

We therefore look at the effects of our measure of clinic congestion, average service population, on annual abortion rates. Table 4 shows the results of both linear (col. 1 and 2) and Poisson models (col. 3 and 4) of monthly births in a county regressed on the average service population of the nearest clinic, scaled to be in 1000s. In both the linear and the Poisson, the effect is non-significant and has the wrong sign in the Poisson model. Recall from figure 4 that the average service population increased from around 190,000 to around 280,000, making this magnitude of an increase comparable to the change seen in the data between 2012 and 2016. When we put both the service population and the bin of distance in the model, the effect size of distance is larger and now significant in the linear model (column 3) and of similar size to the primary specification in the Poisson model (column 4). This suggests that travel distance, not congestion, is the more relevant cost to consider for women when abortion clinics closed in Wisconsin.

We also explore whether average service population is associated with higher birth rates. Table 5 shows the results of regressing monthly county births on the average service population in a linear (col. 1) and a Poisson model (col. 3), as well as adding the bin estimator for not having a clinic within 100 miles, which was the distance bin that had a significant effect on births in the monthly analysis. We do see a significant increase in births in response to increases in service population, with a 100,000 women per clinic increase resulting in a 16.4 percent increase in births in the Poisson model. These effects are robust to including a measure of distance.

Unlike our analysis of distance from the nearest clinic, these findings differ substantively from analyses of the impact of increased service population in Texas. In comparable regressions in Lindo et al. (2019), the authors find that an increase in service population of 100,000 decreases abortions significantly by 7.5 percent. They find no significant effect of population served on birth rates. This demonstrates that the relative impacts of increased service populations may vary depending on the number of abortion clinics in place prior to the implementation of TRAP legislation. As previously noted, the population-weighted average service population of clinics serving Wisconsin was 191,000 in 2012 whereas the average service population of clinics serving Texas was 145,000 prior to HB2 in 2012 (Lindo et al., 2019). In Wisconsin, the average service population in 2017 was 283,000 compared to 253,000 in Texas post HB2 in 2015. While both states had similar increases in congestion in response to clinic closures, Wisconsin started and ended at higher service population levels on average, and the increase had less of an impact on abortion rates. This suggests that there may be diminishing marginal effects of congestion; in states with few clinics, more congestion on top of the existing high service populations may not result in as large a decline in abortions.

4.5 Heterogeneity Analysis

In addition to looking at the overall effects of the clinic closures on births, we also can test whether certain demographic groups were more or less affected by the closures. NVSS reports the age, race, and marital status of the mother, allowing us to calculate the number of births to mothers by age (15 to 19, 20 to 29, 30 or older), race (Non-Hispanic White, Black, Hispanic, or Asian), and marital status (Married, Unmarried) and repeat the Poisson model of birth counts with sub-sample counts as the outcome. Table 6 shows the results of these regressions of sub-sample births on a linear measure of distance (panel A) and separate regressions on the distance bins (panel B). Both models include the same controls as the primary specifications.

We see consistently larger impacts of the closures on older mothers and non-white mothers compared to White mothers. For example, a 100-mile increase is associated with 0.223 more births to Black mothers, compared to 0.039 more births to White mothers. Effects on births to Hispanic and Asian mothers are not significantly different than zero in the model regressing births on distance, but the distance bin analysis shows significant increases in births to Asian mothers. We also can test if the effect on Black births is significantly larger than that on White births, using the test statistic for equality of coefficients described in Clogg et al. (1995) which calculates the following Z-test statistic:

$$Z = \frac{\beta_1 - \beta_2}{\sqrt{\left(SE^2_{\beta_1} + SE^2_{\beta_2}\right)}}$$

Testing equality of coefficients in the model of distance from nearest clinic for Black and White births, gives us a z-statistic of 2.97 which says that we can reject the null that the coefficient for the Black births is equal to the White births coefficient at a p = 0.0015 level. The effect on Asian births is also marginally significantly larger than the effect on White mothers for the linear measure of distance (Z-stat = -1.35, p =0.085). Though effect sizes are larger for mothers over thirty, we cannot reject the null that the coefficients are the same as in the teen birth model. When we compare the coefficients on the distance bins, the effect of increasing distance bins on births are significantly higher for both Asian and Black birth rates compared to the coefficients for White birth rates. The coefficient on moving from being within 25 miles to being 50 to 100 miles from a clinic is also significantly higher for Hispanic births than for White births.

There is a slightly stronger impact of closures on births to unmarried mothers, but this effect is not statistically distinguishable from that for married births. The coefficients on married and unmarried births regressed on distance are not significantly different from each other (Z-stat= 0.044, p = 0.488).

The racial differences in the effects of access to clinics on births are consistent with the findings in Lindo et al. (2019) that the effect of increased distance from a clinic had larger effects for abortions among Hispanic women than non-Hispanic women. They posited that part of the higher effect sizes may be due to Hispanic women in Texas near the Mexican border having greater access to self-induced abortifacients, but this explanation cannot explain similar findings in Wisconsin. Wisconsin does not report abortions by race at the county level other than for the four largest counties, so we cannot do a direct comparison using abortions by race as an outcome variable. Nonetheless, our results suggest that there may be racial barriers to access to abortion care that are exacerbated by the closures of clinics.

4.6 Robustness Checks

4.6.1 *Sub-Sample Analyses*

In our analyses of pre-trends in birth and abortion rates, counties in the middle tercile of distance change (i.e., counties that experience an increase in distance between 33 and 95 miles) experience different trends in birth rates that then other county groups, with spikes in births in the years prior to the clinic closures. To test whether our results to excluding these counties, we re-run the Poisson model regressions of Tables 7 and 8 excluding counties which experience a distance change of between 33 and 95 miles.

This robustness check shows that excluding these counties does not substantively change the results found in the full sample. The effect sizes are somewhat attentuated in the births analysis, but are of the same sign and similar magnitudes as our main findings. Notably, the exclusion of these counties does not change the signs of coefficients in the quadratic model of distance for the birth analysis. The differing pre-trends therefore cannot explain the pattern of increasing marginal effects of distance on births.

4.6.2 *Out of State Abortions*

In all analyses, the abortion counts do not include any abortions obtained by Wisconsin residents in other states. Even prior to the closure of clinics in North Western Wisconsin, the closest clinics to Wisconsin residents in North Eastern Wisconsin were out of state in either Duluth or St.Paul/Minneapolis, meaning that the abortion counts are likely to be undercounts of the true number of abortions obtained by Wisconsinites in both the pre-and post- period. We might be concerned that the effect sizes we see in our analysis of clinic closures are not a true reduction in women's ability to obtain an abortion, but rather that women are merely substituting to obtaining abortions out of state. While we cannot fully rule out this concern, we can look at trends in abortion receipt in neighboring states and test how robust our specification is to excluding counties closest to clinics in neighboring states.

Figure 11 shows the trends over time in the abortions obtained in the two primary neighboring states, Illinois (Jatlaoui et al., 2018) and Minnesota (Jatlaoui et al., 2018 for year 2009-2015; Garcia-Lago, 2017 for 2016, 2017). Trends for Minnesota remain fairly flat across this time period. Since Illinois does not provide annual out-of-state abortion patient numbers broken out by state for years after 2015, we cannot see if there were changes in the number of Wisconsin residents who went to Illinois after the Appleton clinic closed. Overall out of state residents obtaining abortions in Illinois increased by a large number in 2016 and 2017, going from a steady rate of around 3100 from 2011-2015 up to 4543 in 2016 and 5528 in 2017 (Shah, 2017). It is unclear how much of Illinois' increased out-of-state patient load came from Wisconsin residents or from other states, such as Missouri or Iowa, which faced increased restrictions on abortion access during this period as well.

However, the counties for which the closures had the largest impact in terms of changing travel distance are not counties closest to Illinois and Minnesota clinics, either pre- or post- closures. As made clear in figure 3, the counties in north-western Wisconsin that were served by Green Bay Ob/Gyn Services and Appleton Planned Parenthood in 2009 primarily ended up in the Milwaukee service region or the Madison service region. To the extent to which our results are primarily driven by effects on reproductive outcomes of women in these highly-affected regions, these populations are less likely to go out-of-state when the closer clinics are in-state.

We test whether our results are robust to excluding counties for which out-of-state providers are likely to be the primary source of care. We re-run the Poisson model regressions of col. 1 and 3 of Tables 2 and 3 excluding any counties for which the nearest clinic was ever in a state other than Wisconsin. Table 9 and 10 report results for annual abortions and monthly births, respectively.

Our findings are robust to the exclusion of counties served by non-Wisconsin counties. We find that a 100 mile increase in distance is associated with a 22.5 percent decline in abortions, compared to 24.9 percent in the full sample of counties. In the analysis using non-parametric measures of distance, we see somewhat smaller, but still significant coefficients on each distance bin. For example, the coefficient on the indicator for not having a clinic within 100 miles is now -0.238 compared to -0.317 in the full sample.[14] The effect sizes in the model of births are similar in both the full sample and the sample excluding out-of-state access; a 100 mile increase is associated with a 2.9 percent increase in births in both samples.

Additionally, while the missing data on Wisconsin residents traveling outside of the state to obtain abortions in response to the closures may overstate the size of the decline in abortions, we also are understating the effects by ignoring the declines in abortion access to Michigan residents in the upper peninsula for whom the Appleton clinic was one of the closest clinics in terms of travel time. Figure 12 shows the annual trends in abortions obtained by Michigan residents in Michigan for counties in the Upper Peninsula versus counties not in the Upper Peninsula. There is a clear substitution towards in-state abortions for residents of the Upper Peninsula.[15]

[14]Using the Clogg et al. (1995) method of testing difference in coefficients across model, we find that we can reject the null that the coefficients are equivalent at a p = 0.08 level.

[15]In 2015, the Marquette Health Center in the Upper Peninsula began providing abortions, meaning this substitution is likely

5 CONCLUSIONS

Our analysis demonstrates that the closures of two abortion clinics in Wisconsin, OB/GYN Associates of Green Bay and Planned Parenthood of Appleton, induced significant reductions in the number of abortions obtained within Wisconsin. We show that there are significant impacts of distance to the nearest clinic, but that there are diminishing marginal effects of an additional mile with counties originally nearest to the closed clinics experiencing the largest declines. Our estimates for the effects of distance on abortions are comparable to those found in analyses of Texas' clinic closures following HB2. This demonstrates that even relatively small numbers of clinic closures can have large magnitude effects if the clinics that close are geographically remote from the next nearest service provider. Though only two clinics closed in Wisconsin compared to fourteen in Texas, Wisconsinites experienced comparable increases in average distance traveled because the two clinics that closed were the only providers in Northern Wisconsin.

We do not, however, find that congestion – measured using the average number of women served by the nearest clinic – is associated with the number of abortions obtained. In their analysis of the impacts in Texas, Lindo et al. (2019) find that congestion was not only associated with lower abortion rates, but that the majority of the declines in abortions in Texas following HB2 occurred due to this increased congestion. The fact that we do not replicate this result in Wisconsin suggests that the impact of TRAP legislation and the subsequent clinic closures on the clinics that remain open may vary depending on a state's existing clinic infrastructure. Wisconsin had only five clinics even before the implementation of the TRAP legislation and the average population served by those clinics was higher in the pre-legislation period in Wisconsin than in Texas. When considering the impacts of similar legislation in other states with few clinics, we should be careful about generalizing the results on congestion from Texas to other states.

Lastly, our findings suggest that the clinic not only reduced the number of abortions in the state of Wisconsin, but resulted in significantly higher births. We show that an increase in distance from the nearest abortion of 100 miles is associated with a 3.3 percent increase in births. Moreover, we see large racial differences in

only partially due to the loss of the Appleton Clinic as the closest clinic and is instead partially due to a new provider opening in-state.

which populations experienced the greatest increases in birth rates. Births to non-White mothers increased the most, with Black births increasing the most, followed by Asian births, and then Hispanic births. This suggests that legislation that targets abortion providers may have racially disparate effects on access to abortion services.

Our research is a case study of just one state, Wisconsin, and is subject to the same questions of generalizabilty as we raise about the analyses in Texas. Though our findings provide additional support for the finding that increased distance to abortion clinics reduces uses of abortions, more research is needed in a variety of legislative contexts to fully understand how the increased regulation of abortion care will affect fertility across the United States. Additionally, while our results suggest that the reduction in abortions may at least partially result in more births, the changing birth patterns in response to clinic closures are not fully explained by the decline in abortions. We see increasing effects of distance in counties that started further from a clinic, which contrasts with the diminishing marginal effects of distance in the abortion analysis. This suggests that women may be responding to decreased access to abortion clinics in more ways than just their decisions to obtain an abortion or not. Future analyses on the reproductive choices people make to compensate for less access to abortion providers would be an important next step in understanding the long run impacts of restrictions targeting abortion providers.

Nonetheless, this paper demonstrates an important takeaway for policy makers and courts evaluating legislation that impact abortion access. Policies that slowly amp up the regulations and costs faced by abortion service providers can have as much of an impact on access as singular large-scale policy changes. When courts are weighing the costs of these laws, it is therefore important to consider each law in the context of the state's overall legislative environment, as well as considering other characteristics of the state such as the proportion of religiously affiliated health systems or the geographic make up of a state. With an ever growing number of states implementing laws targeting abortion providers – and the subsequent legal challenges making their way to the Supreme Court – it is more important than ever that researchers and policy makers have a comprehensive understanding of how these laws impact women's access to abortion services.

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6 Figures

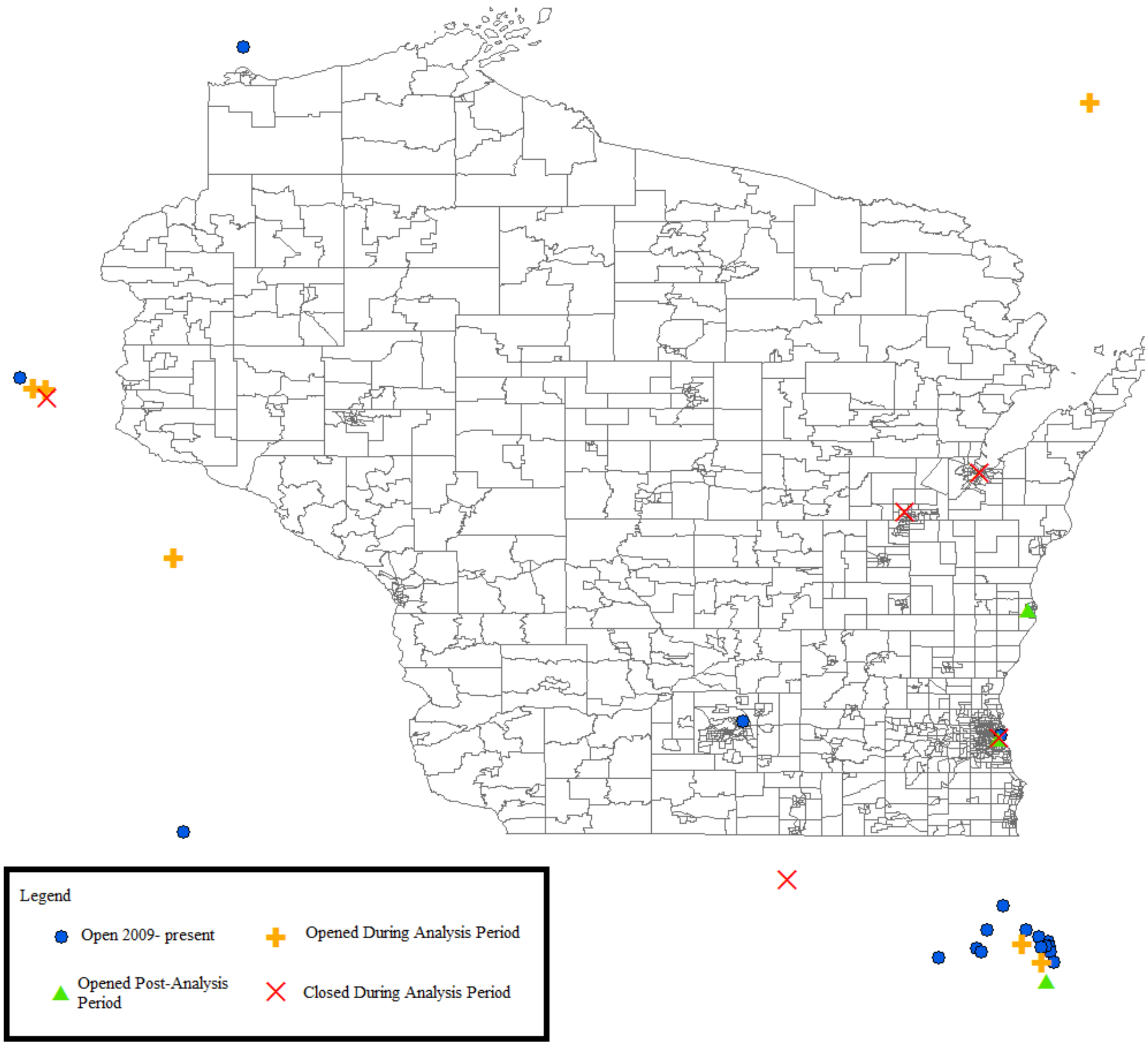
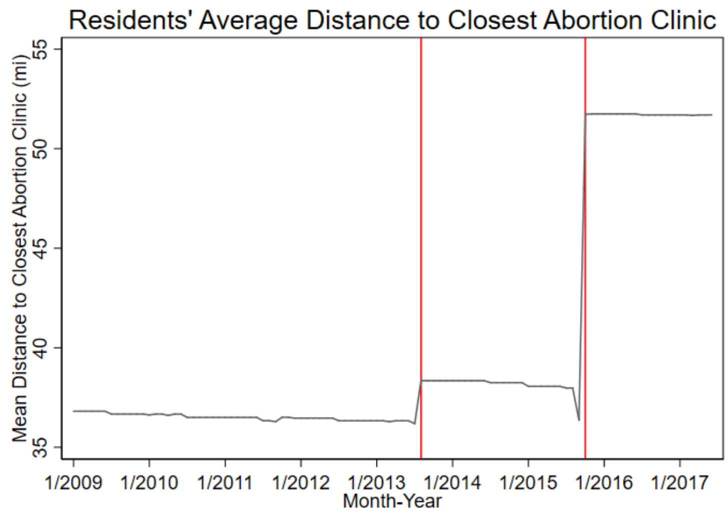


Figure 1: *Locations of abortion clinics, Wisconsin 2009-present*
Notes: This figure shows the locations of all abortion clinics in Wisconsin and bordering states from 2009 through present. Blue circles indicate clinics that were open in 2009 and remain open in present. Green triangles represent clinics currently open that opened after our sample ended. Orange plus signs indicate clinics that opened during our sample period (2009-2017). Red X Marks represent clinics that closed during our sample period.

Panel A



Panel B

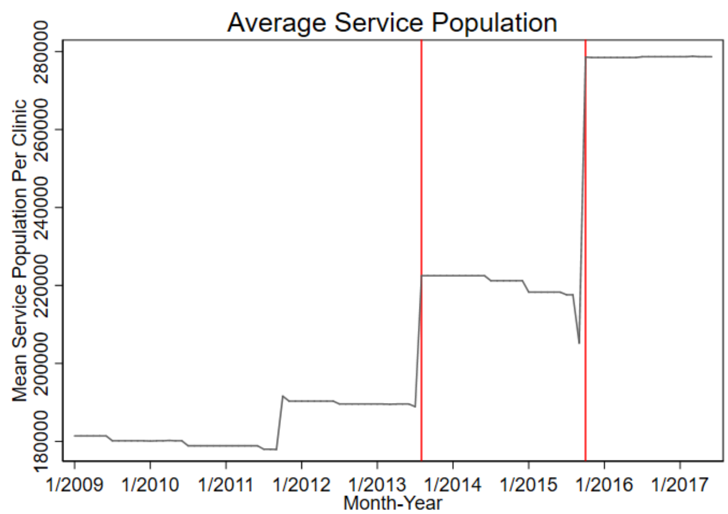


Figure 2: *Abortion clinic access, Wisconsin 2009-2017*

Notes: This figure shows abortion access in terms of distance travelled to nearest clinic (panel A) and size of clinic service population (panel B). Distances are the average travel distance to the nearest clinic from county population centroids, weighted by county population. Average service population is defined as the number of women age 15 to 44 served per abortion clinic, calculated as the population of counties for which a clinic or cluster of clinics was the nearest clinic. Facility operations are measured monthly and are characterized as open if it provided either surgical or medication abortions in that month. The two red lines correspond to the closing of OB/GYN Associates of Green Bay in 8/2013 and Planned Parenthood Appleton in 10/2015.

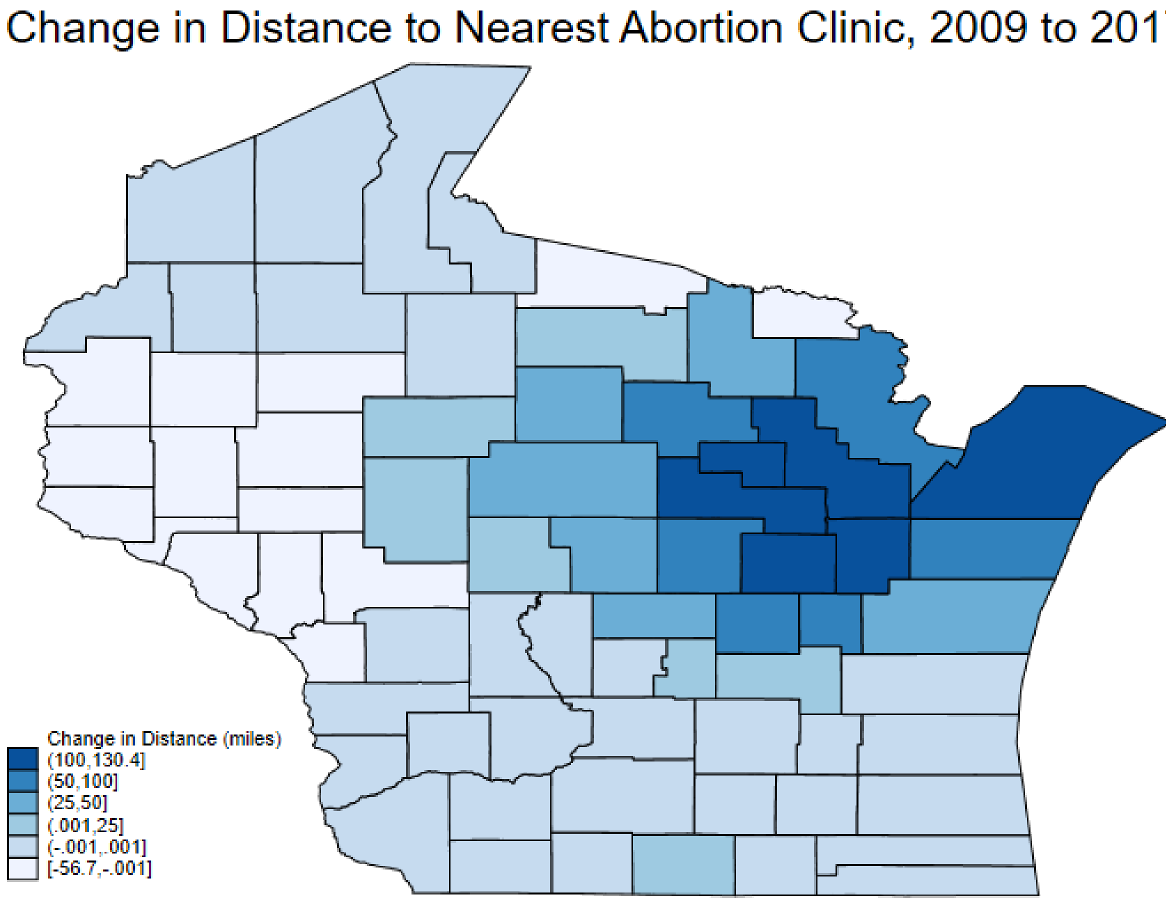
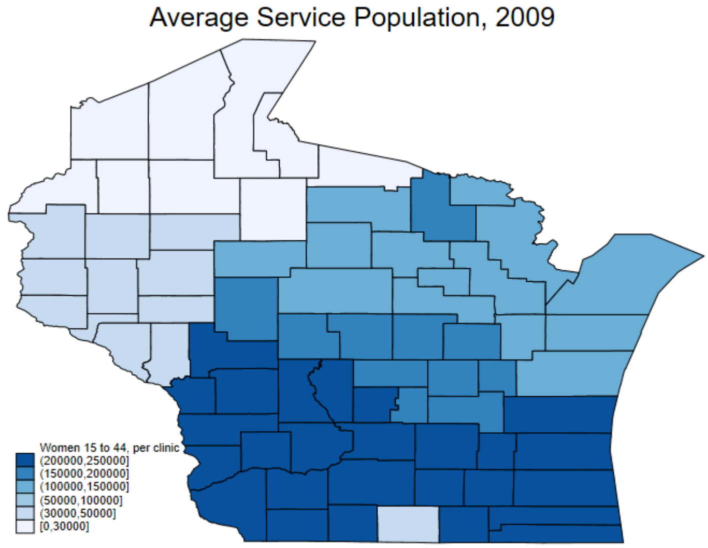


Figure 3: *Change in distance to the nearest abortion clinic, 2009 to 2017*
Notes: This figure shows the change in the distance to the nearest abortion clinic for each county between January 2009 and June 2017. Distances are the average travel distance to the nearest clinic from county population centroids, weighted by county population. Facility operations are measured monthly and are characterized as open if it provided either surgical or medication abortions in that month.

Panel A



Panel B

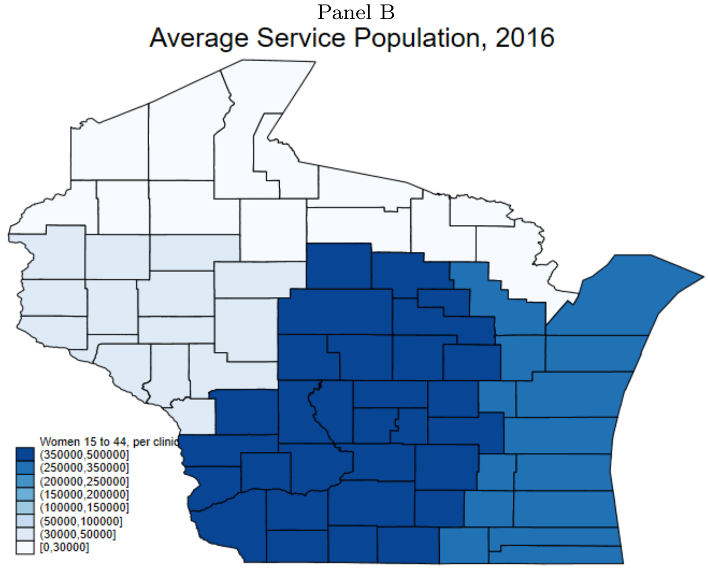
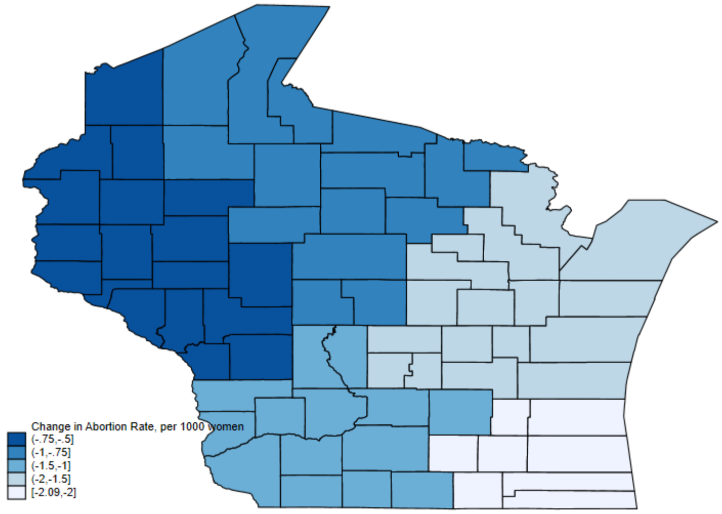


Figure 4: *Size of average service population of nearest abortion clinic, 2009 and 2016*

Notes: This figure shows the average number of Wisconsin residents served by the clinic nearest to each county in terms of travel distance for 2009 (panel A) and 2016 (panel B). Average service population is defined as the number of women age 15 to 44 served per abortion clinic, calculated as the population of counties for which a clinic or cluster of clinics was the nearest clinic.

Panel A

Change in Abortion Rate by Public Health Region, 2009 to 2017



Panel B

Change in Birth Rate by Public Health Region, 2009 to 2016

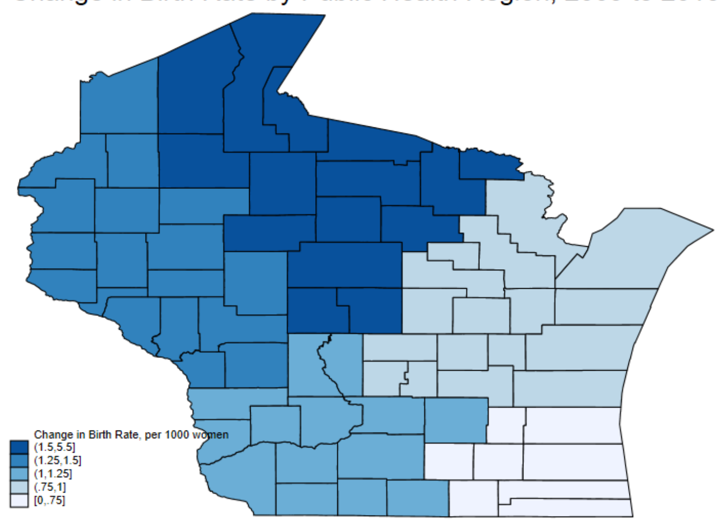


Figure 5: *Change in birth rate and abortion rate by public health regions, 2009 to 2016*

Notes: This figure shows the change in the annual birth rate from 2009 to 2016 (panel A) and the annual abortion rate (panel B) by county from 2009 to 2017. Birth rate is measured as the number of births per 1000 women age 15 to 44. Abortion rate is measured as the number of abortions per 1000 women age 15 to 44, with imputed values of 3 abortions for bottom-coded counties with less than 5 and greater than 0 abortions.

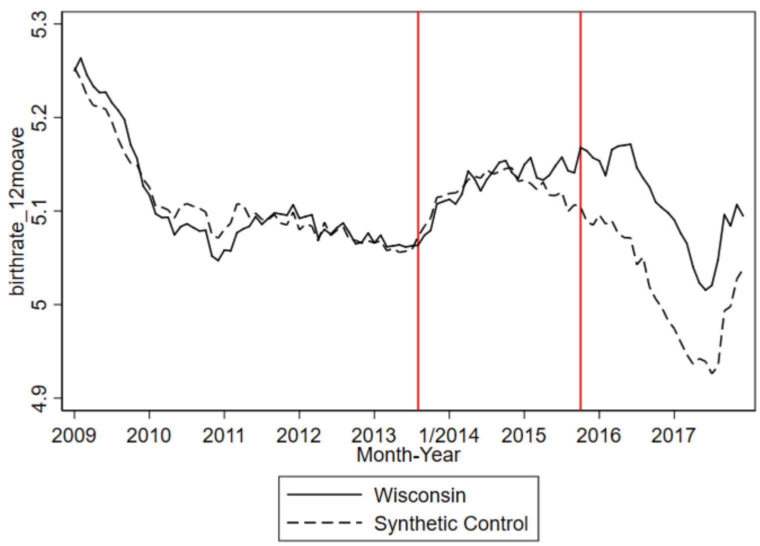


Figure 6: *Statewide 12-month moving average of birth rate: Synthetic control*

Notes: Synthetic Wisconsin is constructed by matching on the 12 month moving average in the 16 months prior to 10/2013, unemployment rate average over that time period, per capita income averaged over that time period, % black mothers averaged over that time period, and % college-educated mothers averaged over that time period. Births are dated back six months so that the dates in the graph correspond to the treatment's effects on mothers who are mid-pregnancy, when abortion clinic access is most likely to matter. We use as possible donors all 49 other states plus DC. The states contributing to synthetic Wisconsin are DE (0.16), IN (0.33), MA (0.06), MT (0.01), NM (0.10), ND (0.06), OR (0.09), and VT (0.19). The red lines correspond to the closure dates of the Green Bay and the Appleton abortion clinics. The average treatment effect in the months following the second clinic closure is 0.10, indicating Wisconsin's had 0.10 more births per 1000 women than the synthetic control state. Following Abadie et al. (2010), we simulate placebo implementation of the treatment in all states in 10/2013 and then calculate the p-value of the average treatment effect by calculating the likelihood that Wisconsin's ratio of average treatment to mean square error of the pre-period is larger in magnitude than the placebo treatment effects; we cannot reject the null, p = 0.333. When we do a one-sided comparison and only look at how many treatment effects are positive and larger than Wisconsin, we still cannot reject the null, p = 0. 176.

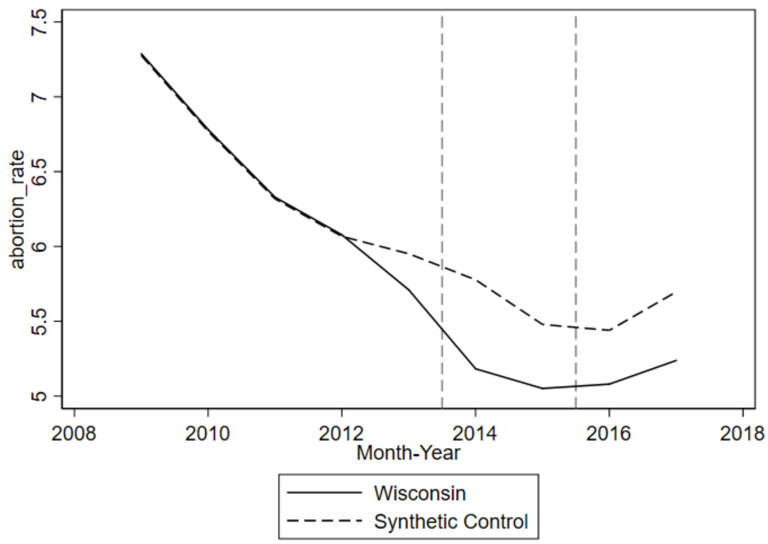


Figure 7: *Statewide annual abortion rate: Synthetic control*

Notes: Synthetic Wisconsin is constructed by matching on annual abortion rates in 2009, 2010, 2011, and 2012 using as possible donors the 36 states we have data on in-state abortion rates (AL, AK, AZ, AR, CO, DE GA, HI, ID, IL, IN, IA, KS, ME, MI, MN, MS, MO, NE, NM, NY, NC, ND, OH, OK, OR, PA, SC, SD, TN, TX, UT, VT, VA, WA, and WV). All states are donors, but the only states donating more than 5 percent of the synthetic sample are Idaho (0.502), Kansas (0.185), and New Mexico (0.077). The dotted lines are the date of the closure of the Green Bay and the Appleton clinics. The average treatment effect (averaged over 2014-2017) is -0.48, indicating Wisconsin's had 0.48 fewer abortions per 1000 women than the synthetic control state. Following Abadie et al. (2010), we simulate placebo implementation of the treatment in all states in 2010, 2011, 2012, 2013, 2014, and 2015 and then calculate the p-value of the average treatment effect by calculating the likelihood that Wisconsin's ratio of average treatment to mean square error of the pre-period is larger in magnitude than the placebo treatment effects; we cannot reject the null, p = 0.54. When we do a one-sided comparison and only look at how many treatment effects are negative and larger than Wisconsin, we still cannot reject the null, p = 0.43

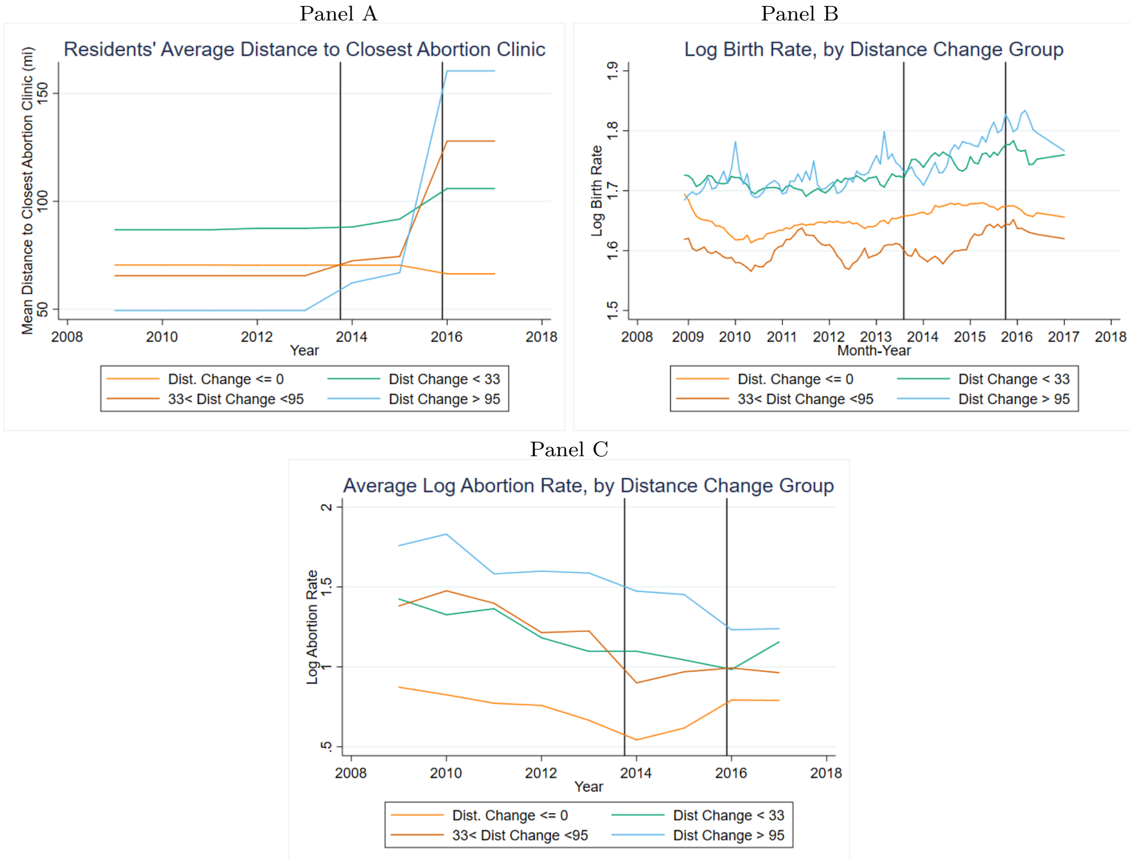


Figure 8: *Pre-trend analysis, by tercile of change in distance to nearest clinic*

Notes: This figure shows the trends in the distance to nearest clinic (Panel A), log of a 12 month moving average of monthly birth rates (panel B) and the log of annual abortion rate (panel C) by tercile of distance change to nearest clinic from 2009 to 2016. Distances are the average travel distance to the nearest clinic from county population centroids, weighted by population of the group. Birth rate is measured as the number of births per 1000 women age 15 to 44. Abortion rate is measured as the number of abortions per 1000 women age 15 to 44, with imputed values of 3 abortions for bottom-coded counties with less than 5 and greater than 0 abortions.The two vertical lines correspond to the closing of OB/GYN Associates of Green Bay in 2013 and Planned Parenthood Appleton in 2015.

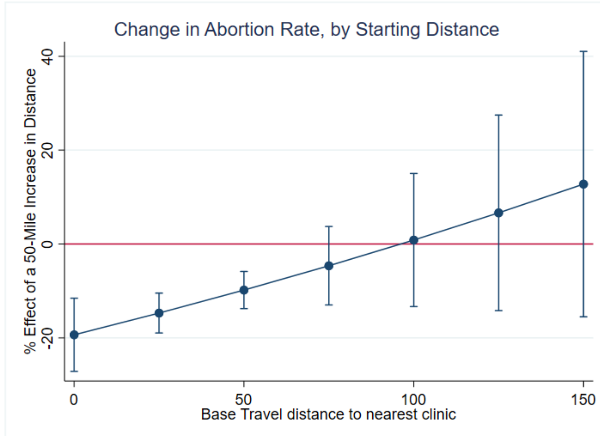


Figure 9: *Effect of a 50 mile increase in distance on abortion*
Notes: This figure shows the percent increase in annual abortion rates associated with a 50 mile increase in distance from the nearest clinic, conditional on starting distance from a clinic. The predictions are based on the estimates in the Poisson model regressing abortions on a quadratic of distance, reported in col. 2 of Table 2.

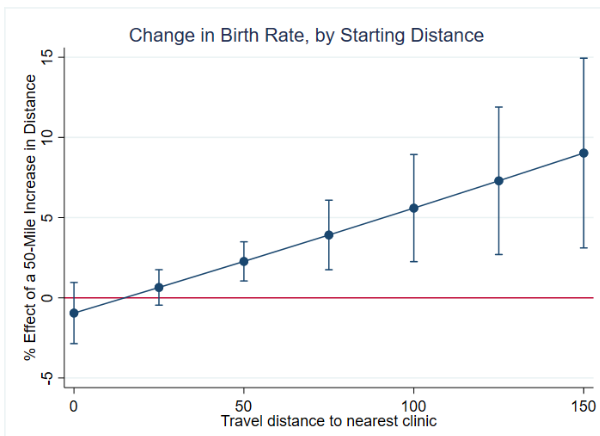


Figure 10: *Effect of a 50 mile increase in distance on births*
Notes: This figure shows the percent increase in monthly birth rates associated with a 50 mile increase in distance from the nearest clinic, conditional on starting distance from a clinic. The predictions are based on the estimates in the Poisson model regressing abortions on a quadratic of distance, reported in col. 2 of Table 2.

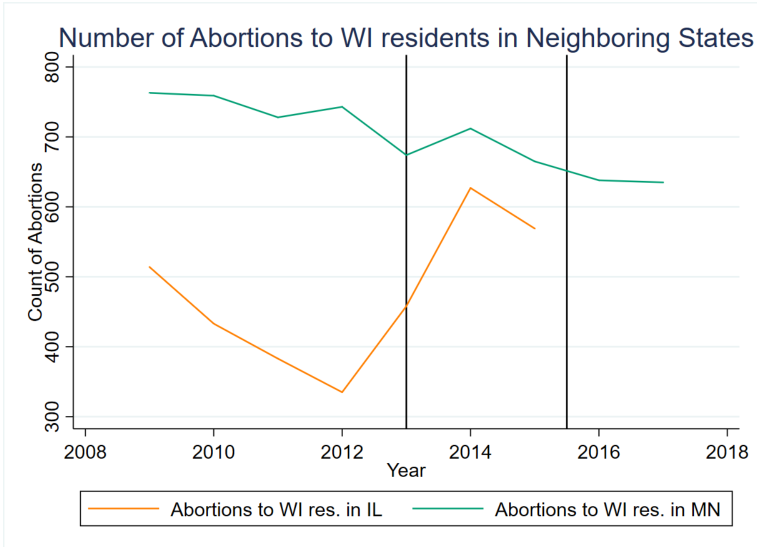


Figure 11: *Number of abortions to WI residents obtained in MN and IL*
Notes: This figure shows the number of abortions obtained in Minnesota and Illinois by Wisconsin residents annually. Data for 2009 to 2015 come from CDC Abortion Surveillance Data. Minnesota abortion data from 2016 and 2017 come from Wisconsin Department of Health Services Abortion Reports 2016 and 2017.

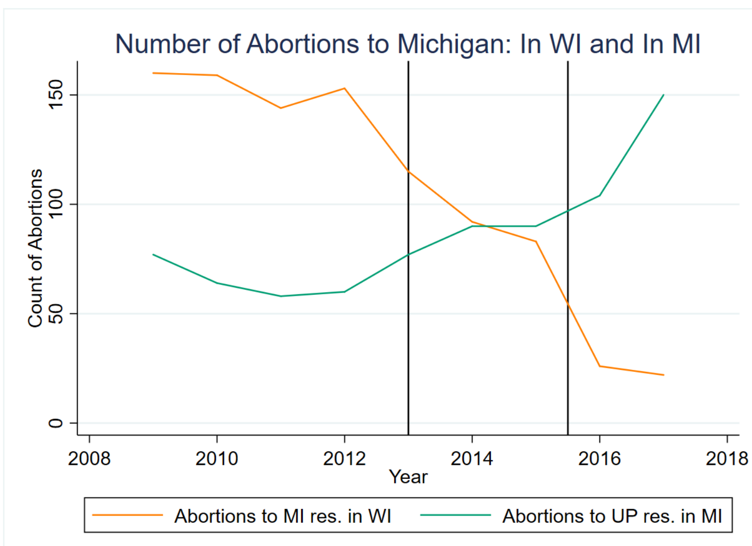


Figure 12: *Number of abortions to MI residents obtained in WI and MI*
Notes: This figure shows the number of abortions obtained by all Michigan residents in Wisconsin from 2009 through 2017, as well as the number of abortions obtained by residents of the Upper Peninsula of Michigan within Michigan.

7 Tables

Table 1: Summary statistics

	2009-2016		2012		2016	
	Mean	Standard dev.	Mean	Standard dev.	Mean	Standard dev.
Birth Rate, per 1000 women	61.135	(7.820)	60.775	(7.695)	61.168	(8.070)
Abortion Rate, per 1000 women	5.992	(4.013)	6.117	(3.846)	5.094	(3.483)
Distance to Closest Clinic	38.842	(39.393)	36.376	(37.900)	52.388	(47.622)
Average Service Pop. (1000s)	205.061	(85.945)	191.346	(71.142)	282.778	(126.839)
Percent Within 25 miles of Clinic	0.128	(0.335)	0.139	(0.348)	0.083	(0.278)
Percent 25 to 50 miles from Clinic	0.208	(0.406)	0.222	(0.419)	0.139	(0.348)
Percent 50 to 100 miles from Clinic	0.429	(0.495)	0.431	(0.499)	0.403	(0.494)
Percent More than 100 miles from Clinic	0.234	(0.424)	0.208	(0.409)	0.375	(0.488)
Women 15 to 44, Percent White	0.873	(0.121)	0.874	(0.121)	0.862	(0.123)
Women 15 to 44, Percent Black	0.078	(0.109)	0.078	(0.110)	0.082	(0.111)
Women 15 to 44, Percent Other Race	0.049	(0.035)	0.048	(0.035)	0.056	(0.037)
Percent Age 15 to 19	0.177	(0.047)	0.176	(0.048)	0.185	(0.046)
Percent Age 20 to 24	0.162	(0.023)	0.163	(0.021)	0.161	(0.025)
Percent Age 25 to 29	0.163	(0.014)	0.163	(0.012)	0.170	(0.011)
Percent Age 30 to 34	0.154	(0.012)	0.148	(0.010)	0.164	(0.011)
Percent Age 35 to 39	0.162	(0.020)	0.165	(0.020)	0.151	(0.016)
Percent Age 40 to 44	0.183	(0.032)	0.185	(0.033)	0.170	(0.026)
Birth Rate, White	51.083	(11.044)	50.867	(10.117)	50.897	(15.037)
Birth Rate, Black	73.154	(31.945)	72.804	(25.985)	79.935	(30.927)
Birth Rate, 15 to 19	21.273	(12.273)	21.937	(10.727)	14.654	(8.747)
Birth Rate, 20 to 29	98.497	(20.899)	98.001	(20.031)	92.004	(22.215)
Birth Rate, 30 and over	52.469	(10.489)	51.305	(9.552)	58.928	(9.351)
County Unemployment Rate	7.395	(2.232)	7.848	(1.600)	4.723	(0.995)
County Per Capita Income	42908.948	(7388.684)	43014.708	(6578.352)	47506.781	(7958.321)
Observations	576		72		72	

Note. Summary statistics calculated for Wisconsin counties (N= 72) for the pooled sample (2009-2016) and individuall for 2012 (the year prior to Green Bay Ob/Gyn closing) and 2016 (the year following Appleton Planned Parenthood closing), weighted by population of the county. Population-level rates are calculated using the population of women aged 15 to 44 in that demographic group as the denominator (with the exception of age-specific birth rates which use the age range specified).

Table 2: Effect of distance increases from clinic closures on annual abortion counts, Poisson fixed effects model

	(1) Abortion rate	(2) Abortion rate	(3) Abortion rate
Distance from Nearest Clinic	-0.00249***	-0.00535***	
	(0.000427)	(0.00160)	
Distance Squared		0.0000214	
		(0.0000123)	
$\mathbb{1}(50 >$Closest Clinic > 25 miles)			-0.204***
			(0.0421)
$\mathbb{1}(100 >$Closest Clinic > 50 miles)			-0.259***
			(0.0436)
$\mathbb{1}$(Closest Clinic > 100 miles)			-0.317***
			(0.0513)
N	420	420	420
County Fixed Effect	Y	Y	Y
Month-Year Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y

Standard errors in parentheses; † $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Note. This table reports coefficients of a Poisson model of abortion counts as a function of measures of travel distance from nearest abortion clinic to county population centroid, where each observation is a county-year and the coefficient on county population is constrained to be 1. Counties with fewer than 5 abortions in a year excluded from regressions. All regressions include county fixed effects, year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. Column 1 reports a regression of abortion rate on distance from the nearest clinic; column 2 adds a quadratic of distance. Column 3 uses binned measures of distance to the nearest clinic: 25-50 miles away, 50-100 miles away, and more than 100 miles away.

Table 3: Effect of distance increases from clinic closures on monthly birth counts, Poisson fixed effects model

	(1) Birth rate	(2) Birth rate	(3) Birth rate
Distance from Nearest Clinic	0.000371**	-0.000511	
	(0.000142)	(0.000305)	
Distance Squared		0.00000640**	
		(0.00000238)	
$\mathbb{1}$(50 >Closest Clinic > 25 miles)			0.00892
			(0.0108)
$\mathbb{1}$(100 >Closest Clinic > 50 miles)			-0.00410
			(0.0164)
$\mathbb{1}$(Closest Clinic > 100 miles)			0.0330*
			(0.0139)
County Fixed Effect	Y	Y	Y
Month-Year Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y
N	6541	6541	6541

Standard errors in parentheses; * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Note. This table reports coefficients of a Poisson model of births as a function of measures of travel distance from nearest abortion clinic to county population centroid, where each observation is a county-month-year and the coefficient on county population is constrained to be 1. All regressions include county fixed effects, month-year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. Column 1 reports a regression of birth rate on distance from the nearest clinic; column 2 adds a quadratic of distance. Column 3 uses binned measures of distance to the nearest clinic: 25 to 50 miles away, 50 to 100 miles away, and more than 100 miles away.

Table 4: Effect of clinic congestion from clinic closures on annual abortion rates

	(1) Abortion rate	(2) Abortion rate	(3) Abortion rate	(4) Abortion rate
Service Pop. (1000s)	-0.000119	0.000661	0.000250	0.000466
	(0.000901)	(0.000913)	(0.000241)	(0.000250)
Distance from Nearest Clinic		-0.00708*		-0.00181*
		(0.00301)		(0.000759)
N	575	575	575	575
County Fixed Effect	Y	Y	Y	Y
Year Fixed Effect	Y	Y	Y	Y
County Level Controls	Y	Y	Y	Y

Standard errors in parentheses; * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Note. This table reports coefficients from linear models of abortion rates (col. 1 and col. 2) and Poisson models of abortion counts (col. 3 and 4) regressed on the average population served by the nearest abortion clinic, measured as the number of women 15 to 44 living in a region with the same nearest clinic or clinic cluster. All regressions include county fixed effects, year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. Column 2 and 4 add a control for distance to the nearest clinic.

Table 5: Effect of clinic congestion from clinic closures on monthly birth rates

	(1) Birth rate	(2) Birth rate	(3) Birth rate	(4) Birth rate
Service Pop. (1000s)	0.000325	0.000266	0.000164***	0.000150***
	(0.000404)	(0.000395)	(0.0000359)	(0.0000389)
Distance from Nearest Clinic		0.000550		0.000115
		(0.00120)		(0.000119)
N	6541	6541	6541	6541
County Fixed Effect	Y	Y	Y	Y
Month-Year Fixed Effect	Y	Y	Y	Y
County Level Controls	Y	Y	Y	Y

Standard errors in parentheses; * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Note. This table reports coefficients from linear models of birth rates (col. 1 and col. 2) and Poisson models of birth counts (col. 3 and 4) regressed on the average population served by the nearest abortion clinic, measured as the number of women 15 to 44 living in a region with the same nearest clinic or clinic cluster. All regressions include county fixed effects, month-year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. Column 2 and 4 add distance to the nearest clinic.

Table 6: Heterogeneity analysis: Birth rates by maternal age, race, and marital status

	(1) Teen births	(2) Twenties births	(3) Thirty Plus births	(4) Black births	(5) White births
Panel A:	0.000119	0.000269	0.000530*	0.00223*** [a]	0.000390*
Distance	(0.000416)	(0.000219)	(0.000212)	(0.000589)	(0.000162)
N	6541	6541	6541	6273	6541
Panel B:					
$\mathbb{1}$(50 >Closest Clinic > 25 miles)	0.0377	-0.00795 [b]	0.0322*	0.0126	0.00631
	(0.0289)	(0.0165)	(0.0132)	(0.0319)	(0.0139)
$\mathbb{1}$(100 >Closest Clinic > 50 miles)	-0.0169	-0.0362*	0.0365	0.312***[a]	-0.0215
	(0.0488)	(0.0166)	(0.0267)	(0.0800)	(0.0179)
$\mathbb{1}$(Closest Clinic > 100 miles)	0.0135	0.0126	0.0597**	0.246***[a]	0.0334*
	(0.0455)	(0.0190)	(0.0216)	(0.0591)	(0.0155)
N	6541	6541	6541	6541	6541

	(6) Hispanic births	(7) Asian births	(8) Unmarried births	(9) Married births
Panel A:	0.000328	0.000724 [b]	0.000356	0.000373*
Distance	(0.000326)	(0.000473)	(0.000215)	(0.000164)
N	6541	6541	6541	6541
Panel B:				
$\mathbb{1}$(50 >Closest Clinic > 25 miles)	0.00132	0.0667* [a]	-0.0328	0.0277* [b]
	(0.0228)	(0.0269)	(0.0171)	(0.0111)
$\mathbb{1}$(100 >Closest Clinic > 50 miles)	0.0805 [a]	0.107*[a]	-0.0139	0.000507
	(0.0491)	(0.0458)	(0.0277)	(0.0186)
$\mathbb{1}$(Closest Clinic > 100 miles)	0.0171	0.0784[b]	0.0306	0.0328*
	(0.0392)	(0.0467)	(0.0270)	(0.0152)
N	6541	6541	6541	6541

Standard errors in parentheses

Null$_1$: $\beta = 0$ * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$; Null$_2$:$\beta = \beta_{\text{base}}$, where base = teen, white, unmarried [a] $p < 0.05$ [b] $p < 0.10$

Note. This table reports coefficients from Poisson models of county birth counts for nine different demographic groups, defined by the characteristic of the mother, regressed on distance from the nearest clinic (panel A) and an indicator for not having a clinic within 100 miles (panel B). Each observation is a county-year and the coefficient on county population is constrained to be 1. All regressions include county fixed effects, year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. County-months with zero black births are omitted from the regression of black births, resulting in 268 county-month observations being dropped.

Table 7: Effect of distance increases from clinic closures on annual abortion rate, Poisson fixed effects model, excluding middle tercile

	(1) Abortion rate	(2) Abortion rate	(3) Abortion rate
Distance from Nearest Clinic	-0.00208***	-0.00550***	
	(0.0007)	(0.00166)	
Distance Squared		0.0000236**	
		(0.000009)	
$\mathbb{1}$(50 >Closest Clinic > 25 miles)			-0.255***
			(0.0473)
$\mathbb{1}$(100 >Closest Clinic > 50 miles)			-0.369***
			(0.075)
$\mathbb{1}$(Closest Clinic > 100 miles)			-0.366***
			(0.0563)
N	349	349	349
County Fixed Effect	Y	Y	Y
Month-Year Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y

Standard errors in parentheses; † $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Note. This table reports coefficients of a Poisson model of abortion counts as a function of measures of travel distance from nearest abortion clinic to county population centroid, where each observation is a county-year and the coefficient on county population is constrained to be 1. We omit all counties that experience a change in distance between 33 and 95 miles. Counties with fewer than 5 abortions in a year excluded from regressions. All regressions include county fixed effects, year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. Column 1 reports a regression of abortion rate on distance from the nearest clinic; column 2 adds a quadratic of distance. Column 3 uses binned measures of distance to the nearest clinic: 25-50 miles away, 50-100 miles away, and more than 100 miles away.

Table 8: Effect of distance increases from clinic closures on on monthly birth rate, Poisson fixed effects model, excluding middle tercile

	(1) Birth rate	(2) Birth rate	(3) Birth rate
Distance from Nearest Clinic	0.000356*	-0.00025	
	(0.000144)	(0.00022)	
Distance Squared		0.00000423**	
		(0.00000177)	
$\mathbb{1}$(50 >Closest Clinic > 25 miles)			0.006
			(0.011)
$\mathbb{1}$(100 >Closest Clinic > 50 miles)			-0.022
			(0.021)
$\mathbb{1}$(Closest Clinic > 100 miles)			0.025*
			(0.012)
County Fixed Effect	Y	Y	Y
Month-Year Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y
N	3361	3361	3361

Standard errors in parentheses;† $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Note. This table reports coefficients of a Poisson model of births as a function of measures of travel distance from nearest abortion clinic to county population centroid, where each observation is a county-month-year and the coefficient on county population is constrained to be 1. We omit all counties that experience a change in distance between 33 and 95 miles. All regressions include county fixed effects, month-year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. Column 1 reports a regression of birth rate on distance from the nearest clinic; column 2 adds a quadratic of distance. Column 3 uses binned measures of distance to the nearest clinic: 25 to 50 miles away, 50 to 100 miles away, and more than 100 miles away.

Table 9: Effect of distance increases from clinic closures on annual abortion rate, Poisson fixed effects model, excluding out of state clinic use

	(1) Abortion rate	(2) Abortion rate	(3) Abortion rate
Distance from Nearest Clinic	-0.00225***	-0.0033***	
	(0.0007)	(0.0015)	
Distance Squared		0.0000076	
		(0.00007)	
$\mathbb{1}(50 >$Closest Clinic > 25 miles)			-0.149
			(0.0473)
$\mathbb{1}(100 >$Closest Clinic > 50 miles)			-0.191***
			(0.075)
$\mathbb{1}($Closest Clinic > 100 miles)			-0.238***
			(0.0563)
N	290	290	290
County Fixed Effect	Y	Y	Y
Month-Year Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y

Standard errors in parentheses; † $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Note. This table reports coefficients of a Poisson model of abortion counts as a function of measures of travel distance from nearest abortion clinic to county population centroid, where each observation is a county-year and the coefficient on county population is constrained to be 1. We omit all counties that ever have an out of state clinic as the nearest clinic. Counties with fewer than 5 abortions in a year excluded from regressions. All regressions include county fixed effects, year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. Column 1 reports a regression of abortion rate on distance from the nearest clinic; column 2 adds a quadratic of distance. Column 3 uses binned measures of distance to the nearest clinic: 25-50miles away, 50-100 miles away, and more than 100 miles away.

Table 10: Effect of distance increases from clinic closures on monthly birth counts, Poisson fixed effects model, excluding out of state clinic use

	(1) Birth rate	(2) Birth rate	(3) Birth rate
Distance from Nearest Clinic	0.00039*	-0.00049	
	(0.00017)	(0.00022)	
Distance Squared		0.0000064†	
		(0.00000177)	
$\mathbb{1}$(50 >Closest Clinic > 25 miles)			0.045
			(0.034)
$\mathbb{1}$(100 >Closest Clinic > 50 miles)			0.001
			(0.024)
$\mathbb{1}$(Closest Clinic > 100 miles)			0.047*
			(0.022)
County Fixed Effect	Y	Y	Y
Month-Year Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y
N	5722	5722	5722

Standard errors in parentheses; * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Note. This table reports coefficients of a Poisson model of births as a function of measures of travel distance from nearest abortion clinic to county population centroid, where each observation is a county-month-year and the coefficient on county population is constrained to be 1.We omit all counties that ever have an out of state clinic as the nearest clinic.All regressions include county fixed effects, month-year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. Column 1 reports a regression of birth rate on distance from the nearest clinic; column 2 adds a quadratic of distance. Column 3 uses binned measures of distance to the nearest clinic: 25 to 50 miles away , 50 to 100 miles away, and more than 100 miles away.

8 Appendix

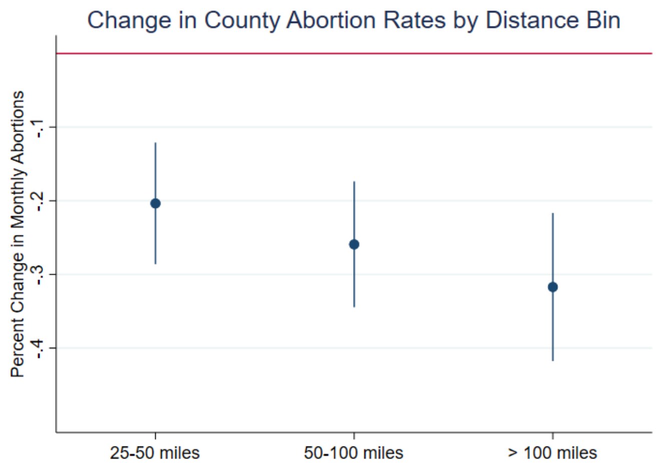


Figure A-1: *Effect of distance from nearest clinic on abortion counts*
Notes: This figure plots the coefficients on indicators for distance from the nearest clinic in a Poisson model regressing annual county abortion counts on these indicators, county fixed effects, year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country.

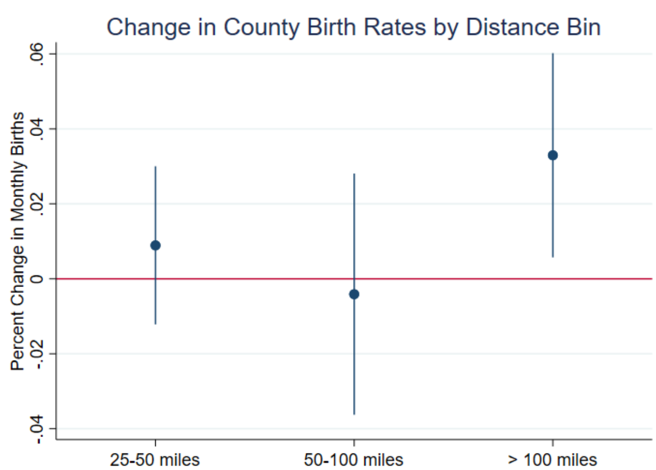


Figure A-2: *Effect of distance from nearest clinic on birth counts*
Notes: This figure plots the coefficients on indicators for distance from the nearest clinic in a Poisson model regressing monthly county birth counts on these indicators, county fixed effects, month-year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country.

Table A-1: Effect of distance increases from clinic closures on annual abortion rates, linear fixed effects model

	(1) Abortion rate	(2) Abortion rate	(3) Abortion rate
Distance from Nearest Clinic	-0.00613*	-0.0202**	
	(0.00296)	(0.00702)	
Distance Squared		0.0000813	
		(0.0000438)	
$\mathbb{1}(50 >$Closest Clinic > 25 miles)			-0.631
			(0.357)
$\mathbb{1}(100 >$Closest Clinic > 50 miles)			-0.909**
			(0.330)
$\mathbb{1}$(Closest Clinic > 100 miles)			-1.192***
			(0.325)
N	428	428	428
County Fixed Effect	Y	Y	Y
Month-Year Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y

Standard errors in parentheses; † $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Note. This table reports coefficients of a regression of abortion rate on measures of travel distance from nearest abortion clinic to county population centroid, where each observation is a county-year. Counties with fewer than 5 abortions in a year excluded from regressions. All regressions include county fixed effects, year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. Column 1 reports a regression of abortion rate on distance from the nearest clinic; column 2 adds a quadratic of distance. Column 3 uses binned measures of distance to the nearest clinic: 25-50 miles away, 50-100 miles away, and more than 100 miles away.

Table A-2: Effect of distance increases from clinic closures on monthly birth rates, linear fixed effects model

	(1) Birth rate	(2) Birth rate	(3) Birth rate
Distance from Nearest Clinic	0.00101	-0.00534	
	(0.00132)	(0.00313)	
Distance Squared		0.0000365	
		(0.0000203)	
$\mathbb{1}$(50 >Closest Clinic > 25 miles)			-0.126
			(0.121)
$\mathbb{1}$(100 >Closest Clinic > 50 miles)			-0.0540
			(0.131)
$\mathbb{1}$(Closest Clinic > 100 miles)			-0.0544
			(0.161)
County Fixed Effect	Y	Y	Y
Month-Year Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y
N	6541	6541	6541

Standard errors in parentheses; * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Note. This table reports coefficients of a regression of birth rate on measures of travel distance from nearest abortion clinic to county population centroid, where each observation is a county-month-year. All regressions include county fixed effects, month-year fixed effects, and controls for county per capita income; county unemployment rate; county population; number of women in age groups 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40, and 40 to 33; county populations of White, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the country. Column 1 reports a regression of birth rate on distance from the nearest clinic; column 2 adds a quadratic of distance.Column 3 uses binned measures of distance to the nearest clinic: 25 to 50 miles away , 50 to 100 miles away, and more than 100 miles away.