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

Joanna Venator
Jason Fletcher

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. Between 2010 and 2017, Wisconsin passed three laws regulating abortion providers and two of five abortion clinics closed in Wisconsin, increasing the distance to the nearest clinic to 55 miles on average and to over 100 miles in the most affected counties. We use a difference-in-differences design to estimate the effect of changes in travel distance on births and abortions, using within-county variation across time in distance to identify the effect. We find that a 100-mile increase in distance to the nearest clinic is associated with 30.7 percent fewer abortions and 3.2 percent more births. We see no significant effect of increased congestion at remaining clinics on abortion rates. Interacting the legislative changes with distance, we find that the effects of distance on abortion are approximately 1.33 time stronger in the presence of laws requiring multiple physician visits to obtain an abortion. 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 where a greater number of clinics ceased operations. © *2020 by the Association for Public Policy Analysis and Management*

INTRODUCTION

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), resulting in many states being left with only a small number of abortion providers. 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.

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

[Correction: The article posted online on 3 November 2020 and has since been updated. Changes in regard to Wisconsin Act 217 have been made throughout. The updates have not affected the conclusions of the paper.]

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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 in which there were only a small number of clinics prior to the implementation of laws meant to restrict access?

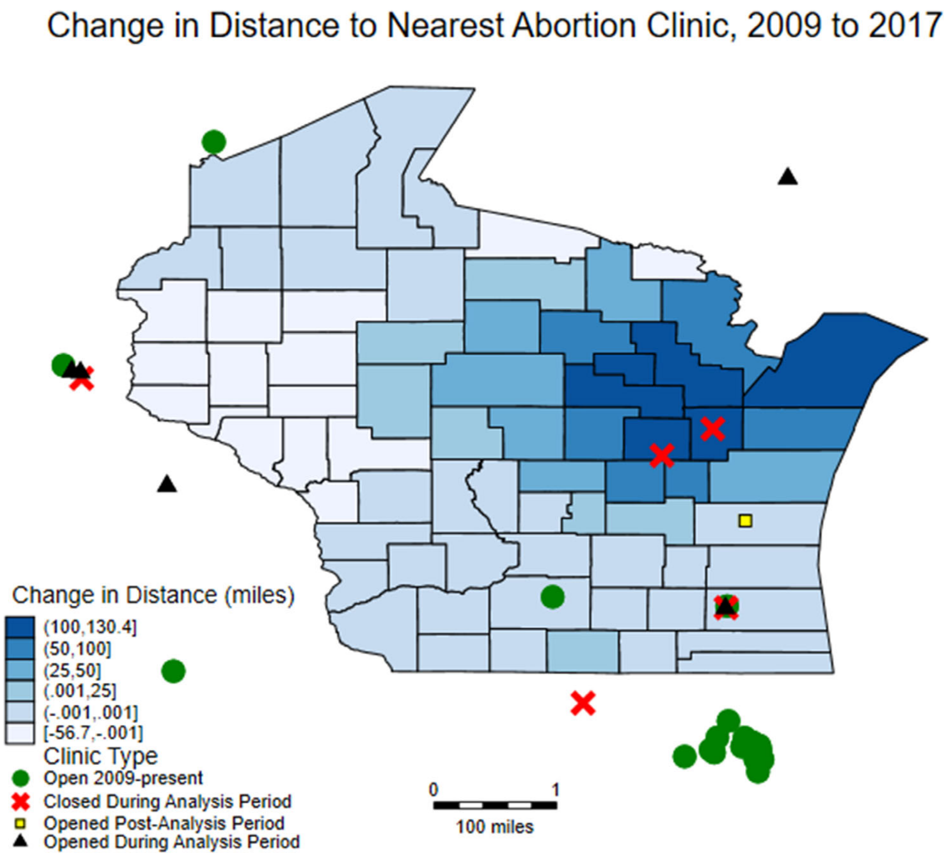
In this paper, we use one such state, Wisconsin, as a case study to explore the impacts of abortion clinic closures on abortion and birth rates. Much of the research on the impacts of abortion clinic closures has been focused on case studies in large states with many abortion providers that experienced large numbers of closures, such as the clinic closures in the wake of the 2013 Texas law HB2. By studying the effects of closures in Wisconsin, we are able to test the effects of clinic closures in medium-sized states that have few clinics even prior to implementation of targeted restrictions of abortion provider (TRAP) laws and have different geographic distributions of clinics.

For the purposes of this study, we focus on changes in abortion services access from 2009 through 2017. During this time period, two out of the five abortion clinics in Wisconsin closed. Figure 1 shows the locations of the abortion clinics in Wisconsin and surrounding states, as well as their closure status. 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 sites 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 used in Lindo et al. (2019) and Fischer, Royer, and White (2018), we then use these data to calculate the driving distance from each county's population weighted centroid to the nearest clinic during 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) that measures the population expected to be served by each clinic.

We find that the clinic closures in Wisconsin produced comparative if not more pronounced effects to those seen in past research on clinic closures. Our analysis shows that a 100-mile increase in linear distance to the nearest abortion clinic is associated with a 30.7 percent decline in abortion rates. We also see evidence of nonlinearity in the relationship between distance and abortion rates with the largest effects occurring for counties that start close to a clinic and experience a closure. Our estimates of abortion rate changes are large enough as a proportion of the state birth rates that we are able to identify a concurrent increase in births; moving from being within 25 miles of a clinic to being more than 100 miles of a clinic is associated with approximately 2.4 percent higher birth rates. These findings are of similar or larger magnitudes to the effects seen in Texas following HB2 (Fischer, Royer, & White, 2018; Lindo et al., 2019; Quast, Gonzalez, & Ziemba, 2017). Our results suggest that a small number of abortion clinic closures can have a similar impact on the use of abortion services as legislation that shuts down a large number of clinics in a short period of time (as in Texas), particularly if those clinic closures drastically change women's geographic access to clinics.

In addition to exploring the effects of travel distance on abortion access, we also look at alternative measures of reproductive healthcare access that may be affected by clinic closures. Past research in the context of Texas (Lindo et al., 2019) and



Notes: This figure shows the locations of all abortion clinics in Wisconsin and bordering states from 2009 through the present, along with 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. Green circles indicate clinics that were open in 2009 and remain open at present. Yellow squares represent clinics currently open that opened after our sample ended. Black triangle signs indicate clinics that opened during our sample period (2009 to 2017). Red X Marks represent clinics that closed during our sample period.

Figure 1. Locations of Abortion Clinics and Travel Distance, Wisconsin 2009 to the present.
[Color figure can be viewed at wileyonlinelibrary.com]

Pittsburgh (Kelly, 2019) suggest that one way in which clinic closures may decrease access to abortions is by increasing the caseloads faced by remaining clinics. We do not find evidence of this in Wisconsin. Our analysis shows that while clinic closures increased the service population per clinic in Wisconsin, the increases in congestion are not associated with lower abortion rates.

Lastly, we leverage policy changes in Wisconsin during the timeframe of our study to explore how the costs associated with travel distance change in the presence of mandatory waiting period laws. In April 2012, Wisconsin passed Act 217, which put in place regulations that require multiple, in-person appointments prior to an abortion. One would expect that this law would increase the costs associated with travel to an abortion clinic, requiring multiple trips or overnight stays. To test this, we regress county abortion rates on the interaction of distance to the nearest clinic and an indicator for whether the law was in place. We find significantly stronger negative effects of travel distance on abortion in the presence of the law; while a 100-mile increase in travel distance decreases the abortion rate by 24.5 percent in the absence of the policy, the same increase results in a 32.7 percent lower abortion rate in the presence of the policy.

These results contribute to the existing literature on the effects of abortion and family planning policy environments on abortion and birth rates. 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, Hershbein, & Miller, 2012), and the legalization of abortion, first at the state level and then nationally through *Roe v. Wade* (e.g., Myers, 2017).

As more states have passed laws intended to restrict access to abortion, a smaller body of research has documented the effects of such restrictions. Our paper complements the existing literature exploring the effects of TRAP laws, which have primarily been focused on legislation in Texas (e.g., Colman & Joyce, 2011; Fischer, Royer, & White, 2018; Girma & Paton, 2013; Grossman et al., 2017; Lindo et al., 2019; Lu & Slusky, 2016, 2019; Myers, Jones, & Upadhyay, 2019; Packham, 2017; Quast, Gonzalez, & Ziemba, 2017; Slusky, 2017)—an important case study due to its large size, heterogeneous population, and sharp law and policy changes. However, it is also a state that differs in size and demographic context from many other states. A small number of studies have tested the effects of TRAP laws outside of Texas, including Kelly (2019), which looks at the impacts of abortion clinic closures on clinic congestion and abortion rates in a single city (Pittsburgh), and Lu and Slusky (2016), which explores the effects of non-abortion clinic closures on preventative care take-up in both Texas and Wisconsin. As discussed in more depth in the next section, Wisconsin provides an opportunity to test whether the results in Texas generalize to smaller states with fewer clinics and different geographic distributions of populations and clinics.

Our analysis also complements past work exploring the effects of laws that increase the number of physician visits required to obtain an abortion. The literature on the impacts of mandatory waiting periods on abortion rates and abortion timing has primarily concentrated on laws passed pre-2000.[1] A series of papers analyzing Mississippi's 1992 mandatory waiting period law found that it was associated with lower abortion rates, delays in abortion, and more out-of-state abortions (Joyce & Kaestner, 2001; Joyce et al., 1997). However, studies of less stringent laws in other states (Joyce & Kaestner, 2001) and analyses of multiple state laws (Bitler & Zavodny, 2001) found that waiting periods had no significant effect on abortion rates but did delay timing of abortions. More recently, Lindo and Pineda-Torres (2019) find that Tennessee's implementation of a mandatory waiting period is associated with a significant increase in the share of abortions obtained in the second trimester and a decline in the abortion rate, though the latter finding is not statistically significant. Our paper, like Lindo and Pineda-Torres (2019), adds to our understanding of how more recent mandatory waiting period legislation has changed the abortion landscape. Recent analyses of the effects of parental consent laws on abortions (Joyce, Kaestner, & Ward., 2019; Myers & Ladd, 2020) suggest that the effects of abortion legislation vary across both geographic contexts and time periods, highlighting the importance of revisiting the effects of laws such as these.

[1] For a more comprehensive review of the effects of waiting periods on a wider array of outcomes, see Joyce et al. (2009).

The remainder of the paper is organized as follows. The following section describes Wisconsin's policy environment between 2009 and 2017, situating this paper within the broader national policy context surrounding reproductive healthcare access. The third section describes the data used and provides descriptive evidence on the pre- and post-clinic closure setting in Wisconsin. The fourth section describes our empirical strategy for estimating the effects of clinic closures on abortion and birth rates. The fifth section describes our results, including a discussion of tests of our identification strategy, our main results, and a set of heterogeneity analyses. Finally, we conclude and offer policy implications of our findings in the sixth section.

POLICY CONTEXT

For the purposes of this study, we focus on changes in abortion services access in the state of Wisconsin from 2009 through 2017. At the beginning of this time period, Wisconsin had five abortion clinics: Planned Parenthood Madison East, Planned Parenthood Milwaukee-Jackson Street, Affiliated Medical Services-Milwaukee, Planned Parenthood Appleton, and OB/GYN Associates of Green Bay. In August of 2013, OB/GYN Associates of Green Bay stopped providing abortions as part of a buyout by Bellin Health Systems.[2] In October 2015, Planned Parenthood Appleton stopped providing abortions.

Figure 1 shows the locations of the clinics that closed and describes the spatial variation in the effects of these closures. The gradient of colors represents the change in distance to nearest clinic between January 2009 and June 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.[3] Western Wisconsin actually experienced a slight decrease in distance due to Planned Parenthood–Rochester, Minnesota opening in late 2015.

During this same time period, there were multiple legislative changes related to reproductive healthcare access. 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 sole federal Title X grantee in Wisconsin 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. In April 2012, the state passed Act 217, which included measures that require multiple, in-person appointments prior to an abortion. Next, in July 2013, Act 37 implemented a series of TRAP laws, including a requirement that women receive fetal ultrasounds. The law also originally included a requirement that physicians have admitting privileges at a hospital within 30 miles, but this was never implemented in full due to a court injunction shortly after it was passed. 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 paper focuses primarily on the effects of the two abortion clinic closures, rather than the legislative changes. Unlike Texas, the closures of the abortion clinics

[2] Prior to the buyout, the practice was privately owned and had one doctor providing abortions. As a condition of continued employment in the system, the doctor was required to agree not to provide abortion services to any patients.
[3] 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.

did not occur simultaneously with law changes and were not directly caused by the changes in legislation.[4] Though our analysis controls for policy changes that affect all counties equally using year fixed effects, some legislation, such as the 24-hour waiting period, may have larger effects in counties further from an abortion provider. We therefore include controls for these policies in a series of additional analyses in the fifth section of the paper when we discuss our results.

There are three reasons why studying Wisconsin's clinic closures adds to our understanding of the changing abortion policy landscape. First, many of the states whose legislation targeting abortion providers are currently making their way through the circuit courts have less than five clinics. In this paper, we focus on a case study of a state more similar to these settings: Wisconsin, which had five abortion clinics in 2009, and by 2017, had only three. 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. Table A1 reports the number of abortion clinics and the percent of women living in a county without an abortion clinic for Wisconsin, Texas, the United States as a whole, and a set of states that are currently facing court challenges to TRAP legislation (i.e., Arkansas, Indiana, Kentucky, Louisiana, Mississippi, Missouri, and Ohio).[5] We see that the average number of clinics in states facing court challenges is 3.85 and about 75 percent of women in these states live in counties without clinics compared to 70 percent of Wisconsin women. Texas, in comparison, had 21 abortion clinics as of 2017 and only 43 percent of women live in a county without a clinic (Jones et al., 2019).

Second, we are able to make causal estimates in a different demographic context than Texas can provide, which may be more similar to that of states facing upcoming judicial challenges to TRAP legislation. Wisconsin over-represents the White population relative to the country as a whole but is more representative than Texas of the states where abortion access is most under legislative threat. Table A1 shows that there is a larger proportion non-Hispanic White population in Wisconsin (0.81) and states with court cases (0.76) relative to Texas (0.41), driven primarily by the much larger share of Hispanics in Texas. Importantly, 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. In the context of research on reproductive policies, Latinx heritage is a particularly important demographic characteristic to consider given notably higher fertility rates among Hispanic Americans relative to Non-Hispanics. In part due to these demographic differences, annual birth rates per 1000 women are lower in Wisconsin (61.3) and states with court cases (64.9) than in Texas (70.6), as are abortion rates (WI: 5.7; States with court cases: 7.5; TX: 11.3).

Lastly, Wisconsin is a medium-sized state and individuals can more easily cross state borders to access abortion providers than in a large state like Texas. This context allows us to test if declines in in-state abortions are driven by substitution to out-of-state providers, an important consideration for states with more lenient

[4] Conversations with the provider from OB/GYN Associates of Green Bay indicate that the buyout was not linked to legislation or changes in demand for abortions from patients, but rather due to retirements of senior practitioners and increasing costs of maintaining a private practice. Planned Parenthood in Appleton cited staffing shortages and security concerns as their reason for closing this clinic, based on stricter security protocols put in place at a national level following a shooting at Colorado Planned Parenthood.

[5] See Table A1 in the Appendix at the end of this article.

abortion policies bordering states with TRAP laws. In a series of robustness checks, we 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. Additionally, we use births to in-state residents as a second outcome of interest. Unlike abortions, these births encompass all births to Wisconsin residents, meaning that the birth estimates include the effect of women substituting to out-of-state abortion care.

DATA

Measures of Abortion Clinic Access

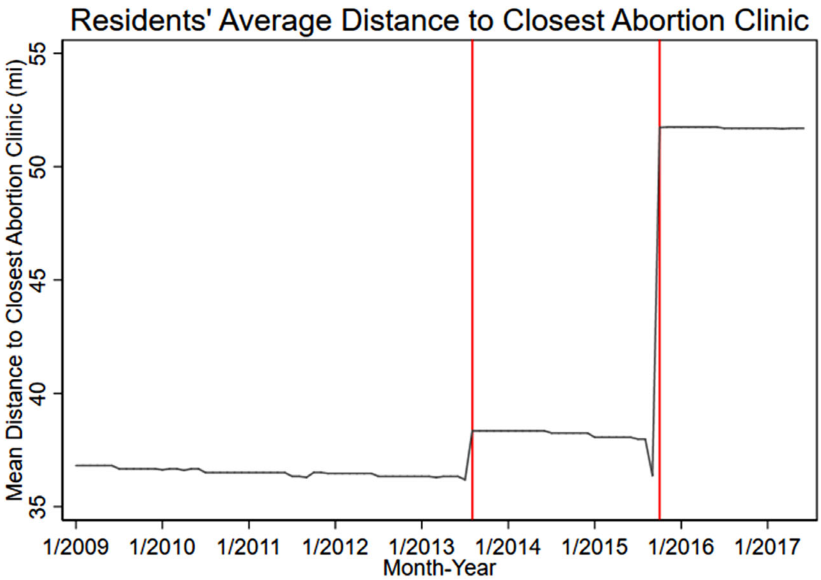
We create a database of all abortion clinics operating in Wisconsin and the bordering states (Minnesota, Iowa, Illinois, and Michigan) between 2009 and 2017 using a variety of sources to verify the dates of openings and closings. Because of the small number of clinics in Wisconsin, the closings of both 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 & Peclat, 2017) to calculate the driving distance between the population-weighted centroid of the county (as calculated by the Census based on the 2010 population count) and each of the 32 clinics in our data set and rank them from closest to farthest.[6] 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.

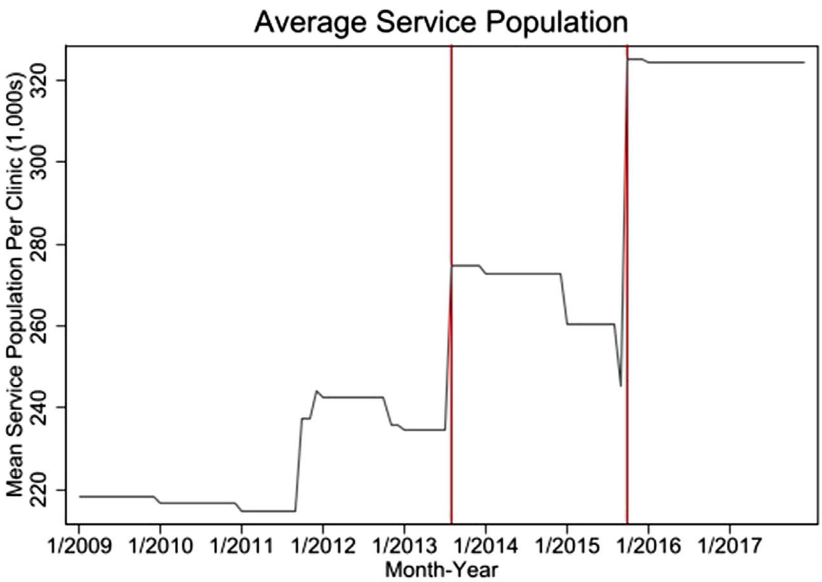
Figure 2, panel A shows the average distance to the nearest abortion clinic for each county, weighted by the population of women age 15 to 44 in the county, from January 2009 to June 2017. The two vertical lines correspond to the closing of OB/GYN Associates of Green Bay in August 2013 and Planned Parenthood Appleton in October 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

[6] County is the most granular level of geographic information we have for both abortions and births, limiting our ability to precisely measure a woman's travel distance. Lu and Slusky's (2019) research on family planning clinic closures suggests that effects are more precisely estimated for zip code-level data; we therefore expect our less granular data to be a conservative estimate of whether there are significant effects of closures.

Panel A



Panel B



Notes: This figure shows abortion access in terms of distance traveled 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 (in 1000s), 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 2. Abortion Clinic Access, Wisconsin 2009 to 2017.
[Color figure can be viewed at wileyonlinelibrary.com]

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 county's population of women to the nearest clinic in each month between January 2009 and June 2017. In addition to the in-state populations, we map all counties in bordering states to the nearest abortion clinic and assign those counties' populations if they are nearest to a clinic within the sample.[7] We define a service area as having multiple clinics if the clinics are within the same commuting zone.[8] 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:

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

Figure 2, panel B shows the average change in service population for the nearest clinic to each county with the two vertical lines again representing clinic closings. By the end of the period, the average service population for each county's nearest clinic increased by approximately 110,000 women, going from around 220,000 women per clinic in 2009 to around 330,000 in 2017.[9] 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 Wisconsin closures resulted in larger changes in average congestion per clinic than the Texas closures.

Figure 3 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 seven different regions of the state served by different clinics, the post-2016 map has six clinic clusters, only two of which are in Wisconsin: Madison (one clinic); Milwaukee (two clinics); Upper Peninsula Michigan (one clinic); Rochester, Minnesota (one clinic); Duluth, Minnesota (one clinic); and Minneapolis/St. Paul (three clinics).

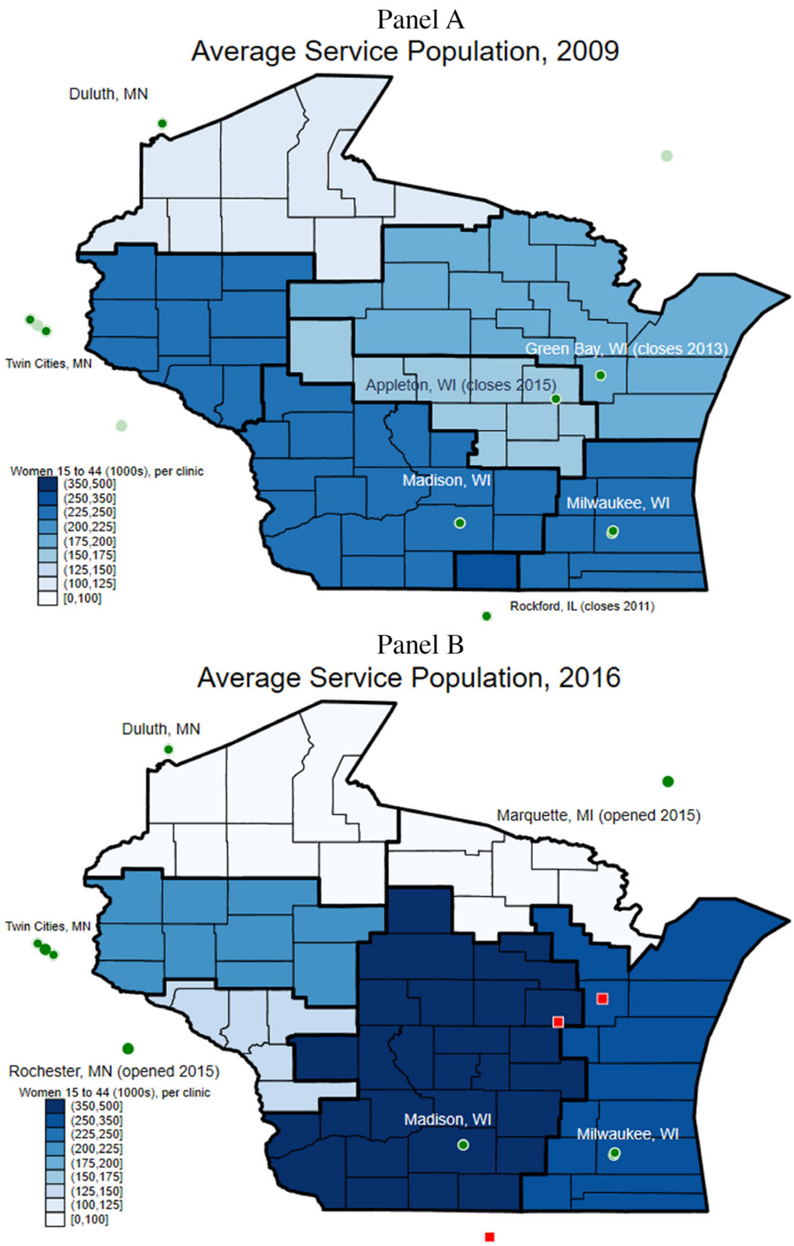
Outcome Measure #1: Abortion Rates in Wisconsin

Abortion data come from the Wisconsin Department of Health Services (DHS), 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, 2017; Ninneman, 2012). All counts are bottom-coded, with any

[7] To assign border states' counties to the nearest clinic, we extend our list of clinics to all clinics in bordering states and states bordering border states (e.g., North and South Dakota clinics are included to calculate which Minnesota counties are nearest those clinics versus the Minneapolis/St. Paul clinics). Note that we do not consider any Canadian clinics as possible substitutes in our analysis since we consider Lake Superior a large enough barrier to travel to forestall travel across national borders.

[8] Despite being relatively close to each other, the Green Bay and the Appleton clinics are in separate commuting zones. We have run our analysis combining these clinics into one service region and it does not substantively change the results.

[9] The spike in 2011 comes from the closure of a Rockford, IL clinic that was previously Rock County's closest clinic; the Madison clinic then became the closest clinic for Rock County.



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 (in 1000s) 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. The clinics that correspond to each service region are labeled; clinics are excluded that are in the sample but were never part of a service region. Dark green dots indicate clinics open in the year in question; light green dots in the 2009 graph are clinics that open during the study period; red squares in the 2016 graph are clinics that close prior to 2016.

Figure 3. Size of Average Service Population of Nearest Abortion Clinic, 2009 and 2016.
[Color figure can be viewed at wileyonlinelibrary.com]

counties reporting zero to five abortions listed as "<5" and counties reporting zero abortions listed as "0." We impute the bottom-coded values to be three for all counties that do not report a zero 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 abortions per 1000 women ages 15 to 44 in a county. Population levels for each county (the denominator of the abortion rate) are taken from the yearly county estimate from the National Institute of Health's Surveillance, Epidemiology and End Results (SEER) U.S. population data (SEER, 2018).

For cross-state abortion analyses, we use annual abortion counts to residents of a state within the state. Data are taken from individual state department of health and human service surveillance reports. Because not all states publicly report abortion counts, data are available for only 38 out of the 50 states.[10]

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 (January 2009 to December 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 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 are defined as before. This leaves us with a data set of month-county observations from January 2009 through June 2017.

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.[11] 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 study 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 calculate the distance

[10] We use data from individual state surveillance reports rather than the CDC national surveillance report because the CDC did not have data available post-2014 at the time of analysis. The full data set, including a list of states included and documentation for the state sources, is available upon request from authors.

[11] 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 percent of the state's Title X patients.

between the clinic and the county population centroids for each county-clinic pair. We then construct our measure of access as the distance to nearest Planned Parenthood open 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 was open in July of that year.

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 the 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.

EMPIRICAL STRATEGY

For this analysis, we follow the empirical strategies used in Lindo et al. (2019) and Fischer, Royer, and White (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, Royer, and White (2018), we compare trends in abortion and birth rates at the state-level to trends in the same time period for other states. Fischer, Royer, and White (2018) conduct a synthetic control method analysis of Texas' HB2 implementation and find a significant increase in monthly fertility rates in their synthetic analysis. We therefore employ a synthetic control method in the context of Wisconsin, using the method described in Abadie, Diamond, and Hainmueller (2015). Figures 4 and 5 show the time-series patterns for 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 January 2016 to June 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, Diamond, & Hainmueller, 2015).

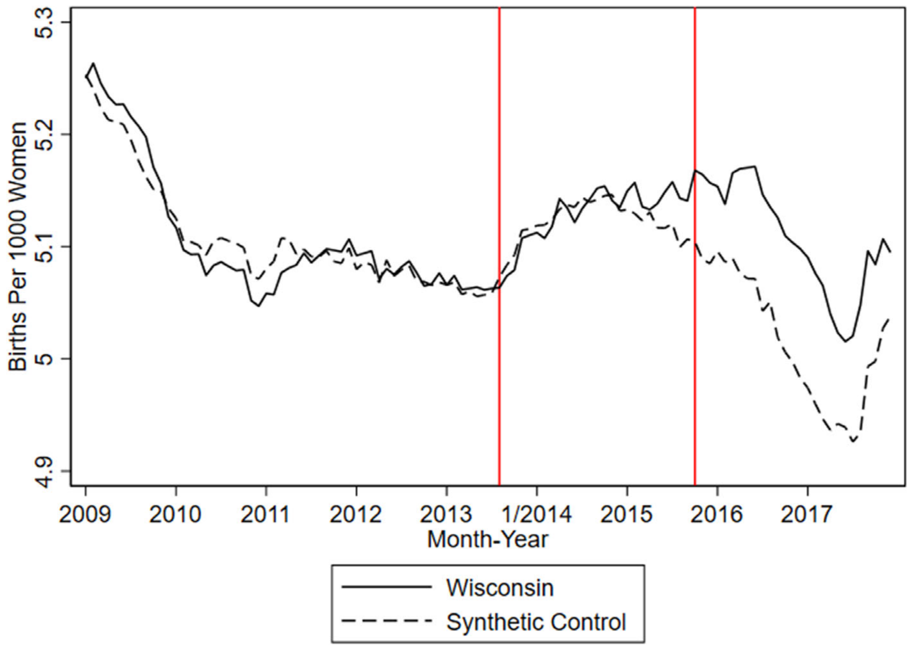
One possible reason for why we do not see significant effects in the synthetic analysis is that the permutation-based inference relies on the assumption that other states do not themselves have policies going into place during this time period affecting their birth and abortion rates. Implementing placebo treatments would only be valid 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

Table 1. Summary statistics.

	2009–2016		2012		2016	
	Mean	SD	Mean	SD	Mean	SD
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)	249.294	(75.037)	244.656	(74.271)	327.512	(118.302)
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	42908.9	(7388.6)	43014.7	(6578.3)	47506.7	(7958.3)
Observations	576		72		72	

Notes: Summary statistics calculated for Wisconsin counties (N = 72) for the pooled sample (2009 to 2016) and individually 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).

permutation-inference exercise 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. Additionally, this synthetic analysis does not distinguish between the effects of the statewide legislative changes in Wisconsin in 2011 to 2013 and the effects of the clinic closures in 2013 and 2015,

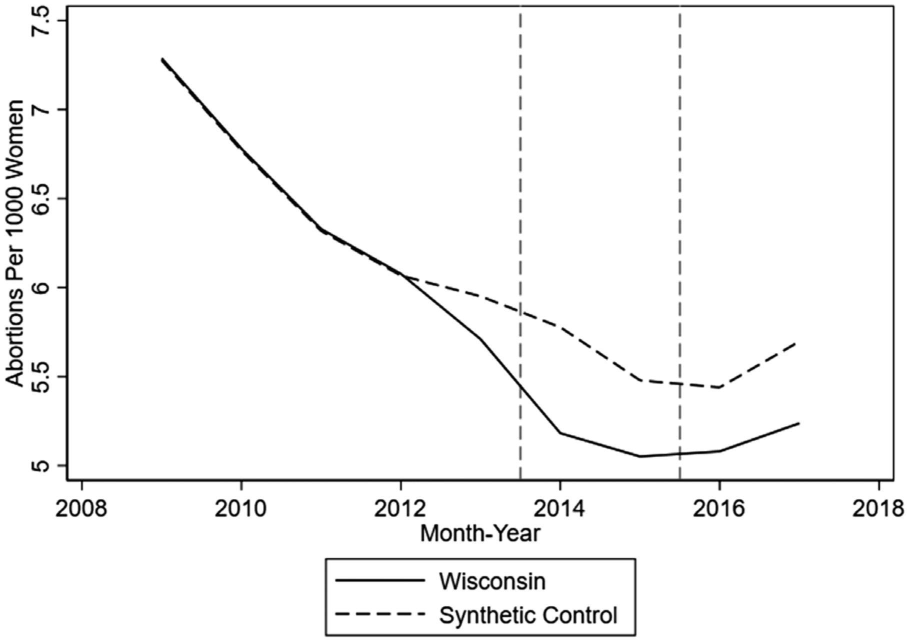


Notes: Synthetic Wisconsin is constructed by matching on the 12-month moving average in the 16 months prior to 10/2013, unemployment rate average over that time period, per capita income averaged over that time period, percent black mothers averaged over that time period, and percent college-educated 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 the District of Columbia. 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 had 0.10 more births per 1000 women than the synthetic control state. Following Abadie, Diamond, and Hainmueller (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 4. Statewide 12-month Moving Average of Birth Rate: Synthetic Control. [Color figure can be viewed at wileyonlinelibrary.com]

nor does it allow us to the explore heterogeneity of the effects of closure based on distance to clinics.

We therefore turn to the methods employed in both Lindo et al. (2019) and Fischer, Royer, and White (2018). In this method, we estimate the effects of access to abortion clinics on abortion and birth rates using a generalized difference-in-differences (D-in-D) design, which uses within-county variation over time in distance to a clinic, controlling for cross-county time-varying shocks. This method identifies the effect of access to clinics based on the assumption that changes in abortion and birth rates for counties with no change in distance over time are a reasonable control group for counties that experience changes in distance. That is, the trajectory of abortions and births over time for counties with no change 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.



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; collected from state DHS Abortion Surveillance reports). 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 to 2017) is -0.48, indicating Wisconsin had 0.48 fewer abortions per 1000 women than the synthetic control state. Following Abadie, Diamond, and Hainmueller (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 5. Statewide Annual Abortion Rate: Synthetic Control.

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.[12] For regressions using counts of births and abortions as the outcome, we use the following estimating equations:

$$\mathbb{E}[abortion\ count_{ct}|dist_{ct}, \alpha_{ct}, \theta_{ct}, X_{ct}] = \exp\left(\beta_1 dist_{ct} + \alpha_c + \theta_t + X'_{ct}\beta_2\right) \quad (1)$$

$$\mathbb{E}[birth\ count_{c,t+6}|dist_{ct}, \alpha_{ct}, \theta_{ct}, X_{ct}] = \exp\left(\beta_1 dist_{ct} + \alpha_c + \theta_t + X'_{ct}\beta_2\right) \quad (2)$$

[12] 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 & Trivedi, 2005).

where abortioncount$_{c,t}$ is the number of abortions in a county c in year t and 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 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 (1(50 > dist > 25 miles), 1(100 > dist > 50 miles), 1(dist > 100 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 distance to the nearest Planned Parenthood.[13]

We also include log of county population of women 15 to 44 and constrain the coefficient to be one, following Fischer, Royer, and White (2018). This is equivalent to Lindo et al.'s (2019) decision to use abortion rate as the outcome in a Poisson model: the log of the rate is the same as the log of the county minus the log of the population. For comparison, we also report a linear fixed effect 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 ages 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 more than 100 miles from the nearest clinic. One might not think that the first mile of distance has the same effect on 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.

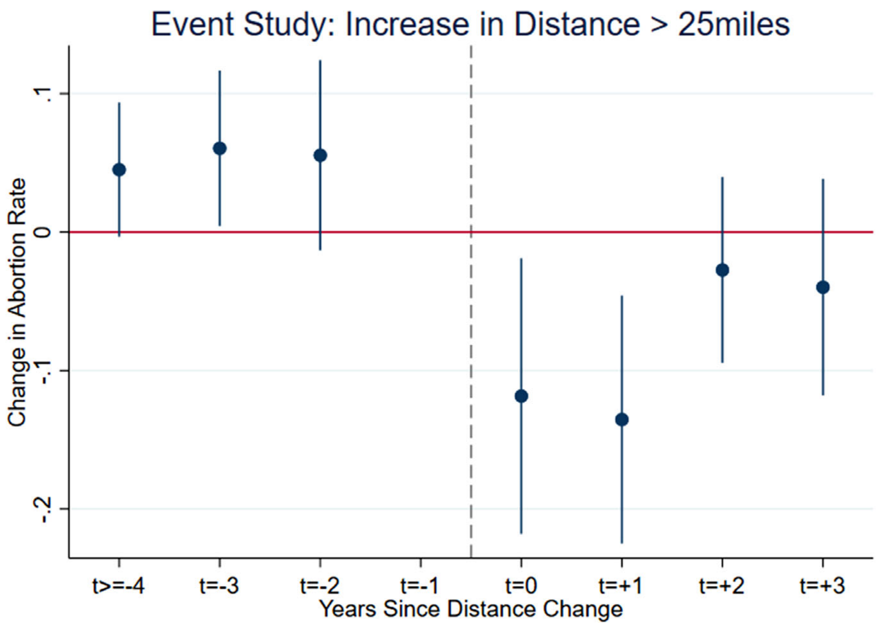
RESULTS

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 sole factor 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 formally test the assumption that abortion and fertility trends pre-clinic closure did not differ across counties systematically by distance to clinic, we run two tests. First, we estimate an event study, where we define the event in question as a

[13] 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 fund 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.



Notes: This figure plots the coefficients of a Poisson regression of abortions on indicators for whether the travel distance to the nearest clinic increased by 25 miles or more, as well as leads and lags around said change. Period $t = -1$ is omitted. The regression includes 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 45; county populations of White, Black, and Other Races groups, and the distance to nearest Planned Parenthood clinic.

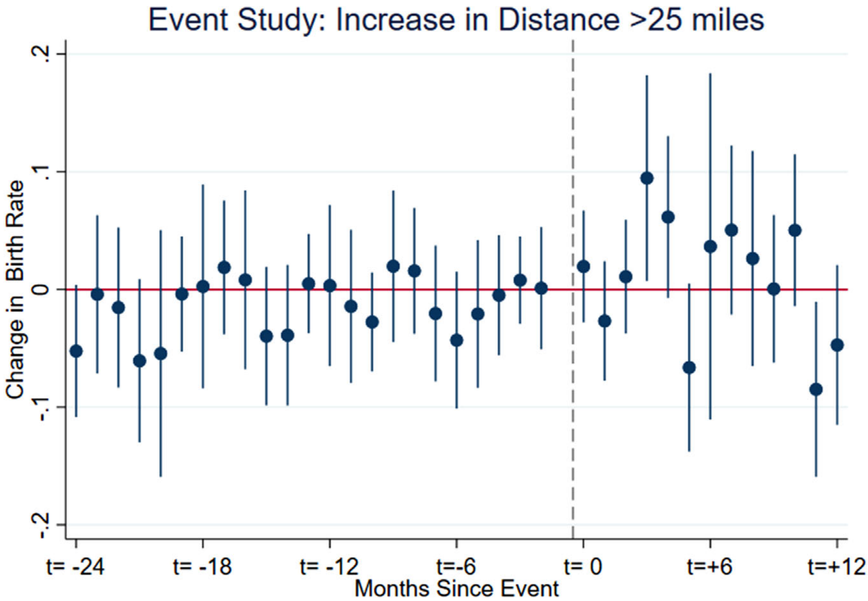
Figure 6. Pre-Trend Analysis: Effect of a Closure that Increases Travel Distance by More Than 25 miles on Abortions.
[Color figure can be viewed at wileyonlinelibrary.com]

closure that causes a 25-mile increase in travel distance to the nearest clinic. We estimate equation (1) with the measure of travel distance replaced by indicator variables equal to one if the change in travel distance since the last period exceeds 25 miles,[14] as well as leads and lags for the years surrounding the reference period, t. The indicator for period $t-1$ is omitted, meaning that the coefficients can be interpreted as the effect of a clinic closure that increases travel distance by more than 25 miles on abortion rates relative to the abortion rates in the year prior to the clinic closure.

Figure 6 plots the coefficients for the three years prior to the closure and for the three years following the closure for the regression with abortions as the outcome. We see no significant difference in pre-closure abortion rates for counties that experience a closure relative to those that do not. We also see a significant decline in abortions in the year of and the year following the closure.

Figure 7 plots the coefficients for the 24 months prior to the closure and for the 12 months following the closure for the regression with births as the outcome, with the month prior to the closure as the omitted category. Unlike the main analyses, we do not offset births by six months to make clearer the timing of the "event"—a closure that increases distance—relative to the timing of the birth. We do not see any

[14] We have also run specifications with 10-mile increases and 50-mile increases; though the magnitude of the effects changes, the direction and significance do not.



Notes: This figure plots the coefficients of a Poisson regression of monthly births on indicators for whether the travel distance to the nearest clinic increased by 25 miles or more, as well as leads and lags around said change. Period $t = -1$ is omitted. The regression includes 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 45; county populations of White, Black, and Other Races groups, and the distance to nearest Planned Parenthood clinics located within the county.

Figure 7. Pre-Trend Analysis: Effect of a Closure that Increases Travel Distance by More Than 25 miles on Births.
[Color figure can be viewed at wileyonlinelibrary.com]

significant differences in pre-closure birth rates in the months before the closures of clinics or in the months immediately following the clinic closure, consistent with the assumption that women who are already well into their pregnancy at the time of the closure are unaffected by the change in travel distance. Post-closure, the estimates are noisily estimated but primarily in the expected direction.

Second, we test whether abortion or fertility trends pre-August 2013 are predictive of the change in travel distance faced by counties following the closures. We regress the change in distance between 2013 and 2017 on the change in abortions between 2009 and 2013, in the following regression:

$$Dist_{c,2017} - Dist_{c,2013} = \beta_0 + \beta_{1(Abortions_{c,2013} - Abortions_{c,2009})} + \varepsilon_c.$$

We then repeat this exercise using change in monthly births as the explanatory variable, where the difference is now calculated as the difference in distance between December 2017 and July 2013 and the difference in births is calculated as the difference in a 12-month moving average of birth rates (offset by six months) between July 2009 and July 2013.

Table A2 reports the results of this regression. There is no significant effect of abortion or birth trends in the pre-period on the subsequent change in travel distance following the clinic closures, consistent with our assumption that closures were not due to changes in demand for abortions.

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 (100 mi.)	−0.307***	−0.466***	
	(0.0410)	(0.137)	
Distance Squared (1002 mi.)		0.119	
		(0.115)	
1(50 >Closest Clinic > 25 miles)			−0.161***
			(0.0352)
1(100 >Closest Clinic > 50 miles)			−0.252***
			(0.0359)
1(Closest Clinic > 100 miles)			−0.360***
			(0.0375)
N	470	470	470
County Fixed Effect	Y	Y	Y
Time Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y

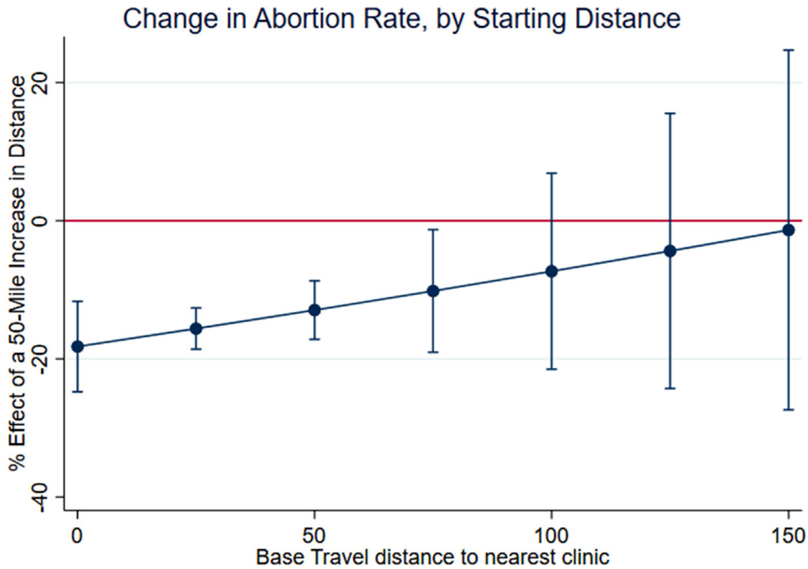
Notes: This table reports coefficients of a Poisson model of abortion counts as a function of measures of travel distance from nearest abortion clinic to county population centroid, where each observation is a county-year and the coefficient on county population is constrained to be one. Counties with fewer than five abortions in a year are excluded from regressions. All regressions include county fixed effects, period 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 45; county populations of White, Black, and Other Races groups, and the distance to nearest Planned Parenthood in state. 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 to 50 miles away, 50 to 100 miles away, and more than 100 miles away. Standard errors are in parentheses. $^{\dagger}p < 0.10$; $^{*}p < 0.05$; $^{**}p < 0.01$; $^{***}p < 0.001$.

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; Table A3 shows the linear model with abortion rates as the outcome. In both tables, column 1 reports the linear effect of distance on abortion rates; column 2 reports the quadratic effect of distance. Column 3 reports the distance bin dummy regressions. As demonstrated in the third column, there are strong nonlinearities in the effect of distance, with larger impacts for counties that switched from being less than 25 miles from a clinic to being more than 100 miles from a clinic.

In our Poisson specification, a 100-mile increase is associated with 30.7 percent fewer abortions per county.[15] In the quadratic version, however, we see that there is a nonlinear relationship with each additional mile increasing the cost at a diminishing rate. Figure 8 plots the effect of a 50-mile increase in distance to the

[15] 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.



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 column 2 of Table 2.

Figure 8. Effect of a 50-mile Increase in Distance on Abortion.
[Color figure can be viewed at wileyonlinelibrary.com]

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 nonlinear 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 16.1 percent, whereas a county that is now 50 to 100 miles from a clinic experiences an average decline of 25.2 percent. Moving to not having a clinic within 100 miles decreases abortions more than a change to being 50 to 100 miles from a clinic: a decline of 36.0 percent. In Figure A1, we plot the coefficients for each distance bin for annual abortions; Figure A2 does the same for monthly births.

Similar patterns are seen in the linear fixed effect model, with a 100-mile increase corresponding to a 0.928 lower abortion rate per 1000 women, which corresponds to a 15.2 percent decline given the average abortion rate in the pre-period of 6.117. 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 a county that began within 25 miles of a clinic and became more than 100 miles from a clinic due to the closures would have 1.5 fewer abortions per 1000 women, which corresponds to a 24.4 percent decline on the base rate.

The nonlinear pattern with lower marginal effects further from a clinic is consistent with the pattern seen in Texas analyses and is of comparable, if not larger, magnitudes. In both the Lindo et al. (2019) and the Fischer, Royer, and White (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, Royer, and White (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 16.6 percent lower abortion rate, 50 to 100 miles is associated with a 16.7 percent lower abortion rate, and more than 100 miles is associated with a 22.1 percent lower abortion rate. The size of our effects is larger than those seen in Fischer, Royer, and White (2018), but smaller than the effects seen in Lindo et al. (2019) for very long travel distances (more than 200 miles).

Births Analysis

Though we see a significant decline in abortions in response to clinic closings, this may not necessarily translate to higher birth rates. Women may go out of state to obtain an abortion, which we are unable to measure in our data. Though Minnesota abortion rates are fairly constant during the time period of our sample, Illinois has experienced large fluctuations in the number of Wisconsin residents obtaining abortions with increases between 2012 and 2014 and declines between 2014 and 2016. Additionally, women may find methods to self-induce abortions if they do not have access to official channels to obtain an abortion. By looking at the effects of clinic closures on birth rates, we are able to address concerns that the "declines" in abortions are not true declines, but instead are substitution to out-of-state or non-official methods of abortion.

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 a similar method to the one described in Fischer, Royer, and White (2018).[16] Our abortion estimates imply that moving from being less than 25 miles from a clinic to the 25- to 50-mile distance bin is associated with a potential increase in the birth rate of 1.17 percent if all missed abortions convert to births. For the 50- to 100-mile distance bin, this method predicts a maximum birth rate increase of 1.51 percent. Lastly, in counties that were more than 100 miles from a clinic, the expected increase in birth rates is 2.8 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 births. Table 3 and Table A4 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 nonlinearities 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 3.18 percent increase in the number of births per month. The bin estimates of distance change also show non-significant and small effects

[16] Specifically, 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 example, for counties that were in the 25- to 50-mile bin in 2017, there were 3,096 abortions from 2009 to 2012 and 42,404 births in the same time period. This gives us an abortion to births ratio of 0.073, which is then multiplied by 0.161.

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

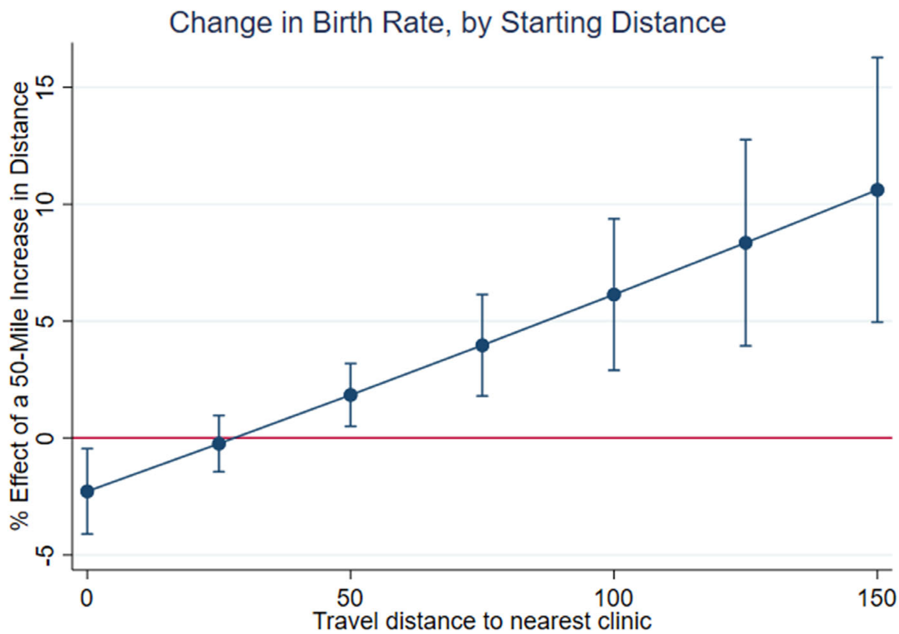
	(1) Birth rate	(2) Birth rate	(3) Birth rate
Distance from Nearest Clinic (100 mi.)	0.0318*	−0.0697*	
	(0.0155)	(0.0299)	
Distance Squared (1002 mi.)		0.0742**	
		(0.0237)	
1(50 >Closest Clinic > 25 miles)			0.00930
			(0.00866)
1(100 >Closest Clinic > 50 miles)			−0.0161
			(0.0157)
1(Closest Clinic > 100 miles)			0.0239†
			(0.0139)
County Fixed Effect	Y	Y	Y
Time Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y
N	6,541	6,541	6,541

Notes: 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, period 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 45; county populations of White, Black, and Other Races groups, and the distance to nearest Planned Parenthood. 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. Standard errors are in parentheses. †p < 0.10; *p < 0.05; **p < 0.01; ***p < 0.001.

for shorter distances, consistent with the smaller effects predicted in the back-of-the-envelope estimates. The bin for counties further than 100 miles from a clinic show a similar effect to what we would expect based on the back-of-the-envelope calculation; having the closest clinic become further than 100 miles away is associated with 2.4 percent higher birth rates compared to our predicted estimate of 2.8 percent higher birth rates. The linear model, reported in Table A4, 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.

Figure 9 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. This suggests increasing marginal effects of distance, which contrasts with the diminishing marginal effects of distance on abortions. To understand this difference across the models, it's important to be clear about what this graph is showing: the percent change in 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 different second derivatives of the quadratic function of distance.[17]

[17] 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-mile bin to the 50- to 100-mile bin and then increasing as we move from the 50- to 100-mile bin to the more than 100-mile bin. This results in a smaller predicted increase in the



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 column 2 of Table 3.

Figure 9. Effect of a 50-mile Increase in Distance on Births.
[Color figure can be viewed at wileyonlinelibrary.com]

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.

We therefore look at the effects of our measure of clinic congestion, average service population, on annual abortion rates. In Table 4, columns 1 and 2 show the results of a Poisson model of annual abortions in a county regressed on the average service population of the nearest clinic, scaled to be in 1000s. The effect is non-significant and has the wrong sign. When we put both the service population and travel distance in the model, the effect size of distance is of smaller magnitude than the primary specification but remains significant. This suggests that travel distance,

percent increase in birth rates from the 25- to 50-mile distance bin to the 50- to 100-mile bin (1.51–1.17 = 0.34) than from the 50- to 100-mile bin to the more than 100-mile bin (2.8–1.51 = 1.29). 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 –0.16 to –0.25 ($\Delta = 0.9$) and then from –0.36 to –0.25 ($\Delta = 1.1$). 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 has different signs.

Table 4. Effect of clinic congestion from clinic closures on annual abortion rates and monthly birth rates.

	(1) Abortion rate	(2) Abortion rate	(3) Birth rate	(4) Birth rate
Service Pop. (1000s)	0.000211	0.000381	0.00007	0.00005
	(0.000205)	(0.000198)	(0.00004)	(0.00004)
Distance from Nearest Clinic (100 mi.)		-0.189^{*}		0.0193
		(0.0791)		(0.0162)
N	647	647	6,541	6,541
County Fixed Effect	Y	Y	Y	Y
Time Fixed Effect	Y	Y	Y	Y
County Level Controls	Y	Y	Y	Y

Notes: This table reports coefficients from Poisson models of annual abortion counts (columns 1 and 2) and Poisson models of monthly birth counts (columns 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, period 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 45; county populations of White, Black, and Other Races groups, and distance to the nearest Planned Parenthood. Columns 2 and 4 add distance to the nearest clinic. Standard errors are in parentheses. $^{*}p < 0.05$; $^{**}p < 0.01$; $^{***}p < 0.001$.

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. Columns 3 and 4 of Table 4 show the results of regressing monthly county births on the average service population in a Poisson model with and without controls for travel distance, respectively. We do see a small increase in births in response to increases in service population, with a 100,000 women per clinic increase resulting in a 0.7 percent increase in births in the Poisson model, though this is not significant at traditional levels ($p = 0.104$). Recall from Figure 3 that the average service population increased from around 220,000 to around 330,000, making this magnitude of an increase comparable to the change seen in the data between 2012 and 2016. These effects are not, however, precisely estimated and the addition of distance increases the imprecision.

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 7.3 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 245,000 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 327,500 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

Table 5. Heterogeneity analysis: Birth rates by maternal age, race, and marital status.

	(1) Non-White births	(2) White births	(3) Non-Teen births	(4) Teen births	(5) Unmarried births	(6) Married births
Panel A:						
Distance (100 miles)	0.00685	−0.109*	−0.118*	0.0472	−0.128*	−0.0664
	(0.149)	(0.0358)	(0.0307)	(0.164)	(0.0401)	(0.0455)
Distance^2 (100 miles)	0.0112	0.0987*	0.0998*	−0.0232	0.111*	0.0681**
	(0.11)	(0.0272)	(0.0218)	(0.0983)	(0.0295)	(0.033)
Panel B:						
Marginal Effect: 50 mi.	0.0119	0.0199*	0.0161**	0.0062	0.0190**	0.0195**
	(0.0491)	(0.0067)	(0.0075)	(0.0267)	(0.0083)	(0.0093)
T-stat	0.161		0.357		0.04	
N	4,840	6,505	6,539	4,967	6,458	6,495

Notes: This table reports coefficients from Poisson models of county birth counts for six different demographic groups, defined by the characteristic of the mother, regressed on a quadratic of distance squared, with the coefficients reported in panel A and the marginal effect of a 50-mile increase in distance conditional on starting 50 miles from a clinic. T-statistic for difference in mean marginal effect across demographic sub-groups reported in panel B as well. Each observation is a county-year and the coefficient on county population is constrained to be one. All regressions include county fixed effects, period 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 45; county populations of White, Black, and Other Races groups, and the distance to nearest Planned Parenthood clinic. County-months with zero subgroup births are omitted from the regression of births. Standard errors are in parentheses. *$p < 0.10$; **$p < 0.05$; ***$p < 0.01$.

few clinics, more congestion on top of the existing high service populations may not result in as large a decline in abortions.

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 (Teen or Non-Teen), race (White, Non-White), and marital status (Married, Unmarried) and repeat the Poisson model of birth counts with subsample counts as the outcome. Table 5 shows the results of these regressions of subsample births using a quadratic of distance (panel A). To make the estimates from the quadratic more interpretable, we also report the quadratic model prediction of the marginal effect of a 50-mile increase in distance for a county that starts 50 miles from the clinic in panel B. All models include the same controls as the primary specifications. We drop county-month observations where there were zero births to that subgroup.

We see significant impacts of the closures on older mothers, White mothers, and both married and unmarried mothers. For example, a 50-mile increase is associated with 1.99 percent increase in the White birth rate, compared to a non-significant 1.19 percent for the non-White birth rate. The marginal effect for non-teen mothers is a 1.6 percent increase in births relative to a 0.6 percent increase for teen mothers. However, smaller subgroups (i.e., Non-Teens; Non-White) have less precisely estimated effects and we cannot reject the null hypothesis that the marginal effects are equal across demographic groups.[18]

Threats to Identification

The previous analyses focus on variation in clinic closures to identify the effects of travel distance to abortion providers on abortion and birth rates. The identification of the causal effect of clinic closures relies on the assumption that the only factor changing within the county at the exact time of the clinic closures was the distance to the nearest abortion clinic. By controlling for time period fixed effects, we control for any changes that affect all counties equally—for example, during the period this study focuses on, the Affordable Care Act was implemented and Wisconsin's Medicaid program, BadgerCare, changed the income thresholds Wisconsin used for eligibility. These time period fixed effects also ostensibly control for the effects of the TRAP laws implemented in either Wisconsin or neighboring states.

However, there are a particular set of TRAP laws that we might expect to differentially impact counties based on their distance to an abortion clinic: laws that increase the number of physician visits required to obtain an abortion. We might expect that this kind of law is more of an obstacle for women who live further from a clinic. In April 2012, Wisconsin implemented Act 217, a law that increased the number of appointments required prior to receiving an abortion including a requirement that 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 women twice: one to administer the first round of the drug and then be physically present when the second round of the drug is given to the woman the next day. We therefore might be concerned that travel distance has a differential effect on abortion rates in counties where the closest clinic requires multiple visits to receive an abortion.

To test this, we reestimate the Poisson model of abortions with the measures of distance interacted with an indicator for if the nearest clinic is in a state/time period in which Act 217 is in place. This indicator is equal to one in 2012 and later. Table 6 shows the results of this regression; column 1 uses the linear measure of distance and column 2 uses the binned measures of distance.

We see that a 100-mile increase in travel distance is associated with a 24.5 percent decrease in the abortion rate in the absence of the law, but that in the presence of the law, distance matters more. Act 217 decreases the effects, with an 100 mile increase now being associated with a 32.7 percent decrease. The binned estimators show that these effects are strongest for very large changes in travel distance. While the effects of distance are not significantly different for the 25- to 50-mile bin or the 50- to 100-mile bin, the law significantly increases the effects of a change in travel distance that moves a county from being close to a clinic to being more than 100 miles from a clinic. In the absence of the law, having the closest clinic be further than 100 miles away is associated with 19.8 percent lower abortion rates. This increases to 36.7 percent lower abortion rates in the presence of the law.

[18] To test differences across groups, we use the test statistic for equality of coefficients described in Clogg, Petkova, and Haritou (1995): $T = \frac{\beta_1 - \beta_2}{\sqrt{SE^2_{\beta_1} + SE^2_{\beta_2}}}$.

Table 6. Interaction between Act 217 legislation and clinic closures.

	(1) Abortion rate	(2) Abortion rate
Distance (100 mi.)	-0.245^{*}	
	(0.0583)	
Act 217 × Distance (100 mi.)	$-0.0824^{\dagger}$	
	(0.0467)	
1(50 >Closest Clinic > 25 miles)		-0.201^{*}
		(0.0470)
1(100 >Closest Clinic > 50 miles)		-0.249^{*}
		(0.0512)
1(Closest Clinic > 100 miles)		-0.198^{*}
		(0.0541)
Act 217 × 1(50 >Closest Clinic > 25 miles)		0.0605
		(0.0535)
Act 217 × 1(100 >Closest Clinic > 50 miles)		−0.0124
		(0.0425)
Act 217 × 1(Closest Clinic > 100 miles)		-0.169^{*}
		(0.0502)
N	470	470

Notes: 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 interacted with an indicator for if Act 217 is in place, where each observation is a county-year and the coefficient on county population is constrained to be one. Counties with fewer than five abortions in a year are excluded from regressions. All regressions include county fixed effects, period 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 45; county populations of White, Black, and Other Races groups, and the distance to nearest Planned Parenthood in state. Column 1 reports a regression of abortion rate on distance from the nearest clinic. Column 2 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. Standard errors are in parentheses. $^{\dagger}p < 0.1$; $^{*}p < 0.05$; $^{**}p < 0.01$.

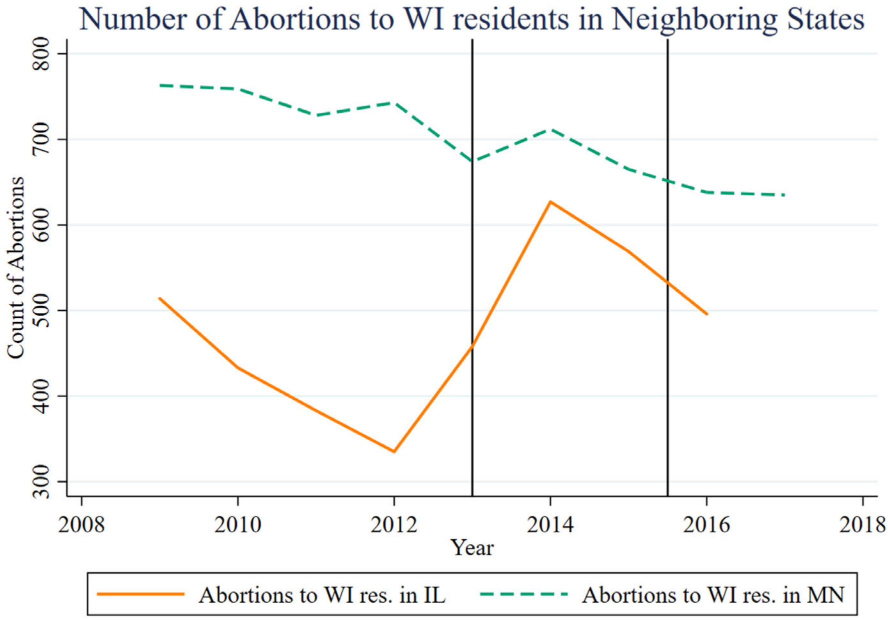
Robustness Checks

We conduct three additional exercises to test that our findings are robust to alternative specifications. Full descriptions of these specifications are described in the Appendix.

First, we test whether our birth analysis is sensitive to our decision to offset births by six months prior to the timing of the birth by testing alternative offsets, including nine and 12 months (results shown in Table A5). We find no significant association between birth rates and travel distance to a clinic nine months prior to birth or 12 months prior to a birth.

Next, we test whether our results are robust to excluding counties in the middle tercile of distance change, which experienced different trends in birth rates pre-closure from other counties (columns 1 and 2 of Table A6). We find that this exclusion does not substantively change the results found in the full sample.

Lastly, 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. In all analyses, the abortion counts do not include any abortions obtained by Wisconsin residents in other states, which might lead to the concern that our abortion results are driven by substitution to out-of-state abortion providers rather than women no longer being able to obtain abortions. We rerun the Poisson model regressions excluding any counties for which the nearest clinic was ever in a state other than



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

Figure 10. Number of Abortions to WI Residents Obtained in MN and IL. [Color figure can be viewed at wileyonlinelibrary.com]

Wisconsin (columns 3 and 4 of Table A6). The effects of distance are somewhat attenuated when these counties are excluded but remain statistically significant in all specifications.

Additionally, we do not observe increases in abortions to Wisconsin residents in neighboring states large enough to explain the declines in in-state abortions. Figure 10 shows the trends over time in the abortions obtained in the two primary neighboring states, Illinois (Jatlaoui et al., 2018) and Minnesota (Garcia-Lago, 2017; Jatlaoui et al., 2018). Trends for Minnesota remain flat across this time period. There are large fluctuations in abortions to Wisconsin residents in Illinois, but these fluctuations do not coincide with the years of clinic closures and are at a similar level post-closures (in 2016) as they are pre-closure (in 2013). The Appendix discusses these out-of-state substitution patterns in more detail.

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 14 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.

Third, our findings suggest that the clinic closures 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 provider of 100 miles is associated with a 3.2 percent increase in births. Since this measure includes births to all Wisconsin residents, this estimate is robust to concerns about closures causing women to substitute to out-of-state providers.

Lastly, this research demonstrates that legislation restricting access to abortions exacerbate the effects of increased travel distance on women's ability to use abortion services. Specifically, we find that the declines in abortions associated with an increase in distance are approximately 33 percent larger in the presence of a law that increases the number of physician visits required for abortions. This finding suggests that the spatial context of abortion access in a state is important to consider when analyzing the impacts of TRAP legislation. The impacts of these policies in states with less geographic coverage of abortion providers may be stronger than in states with more coverage.

Our research is a case study of just one state, Wisconsin, and is subject to the same questions of generalizability as we raise about the analyses in Texas. Though our findings support past research showing increased distance to abortion clinics reduces use 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. 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 policymakers have a comprehensive understanding of how these laws and subsequent clinic closures would impact women's access to abortion services.

JOANNA VENATOR is a Ph.D. Candidate in Economics at the University of Wisconsin–Madison, Department of Economics, 1180 Observatory Drive, Madison, WI 53706 (e-mail: venator@wisc.edu).

JASON FLETCHER is a Professor of Public Affairs and Sociology at the La Follette School of Public Affairs at the University of Wisconsin–Madison, 1225 Observatory Drive, Madison, WI 53706-1211 (e-mail: jason.fletcher@wisc.edu).

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APPENDIX

Robustness Checks

Timing of Births and Closures

For the birth analyses, we look at the effect of driving distance six months prior to the timing of the birth to adjust for the fact that access to abortion services are most relevant to realized births in the first trimester of the pregnancy, not at the time of birth. However, we might be concerned that part of what we capture in our analysis of birth rates is not only how clinic closures affected access to abortion services, but also how they changed access to contraceptive services. We therefore test how much of the decline in births remain if we offset births by nine months or 12 months. The driving distance to an abortion clinic 12 months prior to a birth would be known at the time of conception, meaning that if we see increases in 12-month offset births, women may be responding to abortion clinic closures through changes in pre-conception behavior such as contraception use in addition to changes in abortion decisions post-conception.

Table A5 reports the coefficients for a Poisson model of births for our primary specification of a six-month offset (column 1), a nine-month offset (column 2), and a 12-month offset (column 3). There is no significant association between births and distance to the nearest clinic nine months prior to birth or 12 months prior. We can marginally reject that the coefficient is larger in our primary specification than the 12-month offset in a one-sided test of equality of coefficients, $p = 0.08$. This suggests that the increases in births are not primarily working through decreases in access to contraceptive services.[19]

Subsample Analyses

When looking at pre-trends in counts of births and abortions, 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 than other county groups, with spikes in births in the years prior to the clinic closures. To test whether our results are robust to excluding these counties, we re-run the regressions of abortions and births on linear travel distance and bins of distance excluding counties that experience a distance change of between 33 and 95 miles. Columns 1 and 2 of Table A6 report the results.

This robustness check shows that excluding these counties does not substantively change the results found in the full sample. The effect sizes are somewhat attenuated in the births analysis but are of the same sign and with 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.

[19] These findings contrast with the findings of Lu and Slusky (2019) and Packham (2017), who use cuts to family planning funding in Texas to show that family planning clinic closures result in lower birth rates. The difference in these results may stem from the fact that these abortion clinics are only two of many family planning clinics in the state, making the context very different from Texas. While there are few abortion providers in Wisconsin, family planning clinics are more plentiful and geographically well-distributed throughout the state.

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 an indication that women are merely 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.

To test this, we re-run the primary specification excluding all counties that ever have an out-of-state abortion provider as the closest provider. Columns 3 and 4 of Table A6 report the results. Results are attenuated for both the linear and the nonparametric specifications, but the exclusion of these counties does not substantively change our findings. For example, we find that a 100-mile increase in distance is associated with a 22.5 percent decline in abortions, compared to 30.7 percent in the full sample of counties. The coefficient on the indicator for not having a clinic within 100 miles is now -0.238 compared to -0.360 in the full sample.[20] The effect sizes in the model of births are larger in the sample excluding out-of-state access; a 100-mile increase is associated with a 3.9 percent increase in births in both samples.

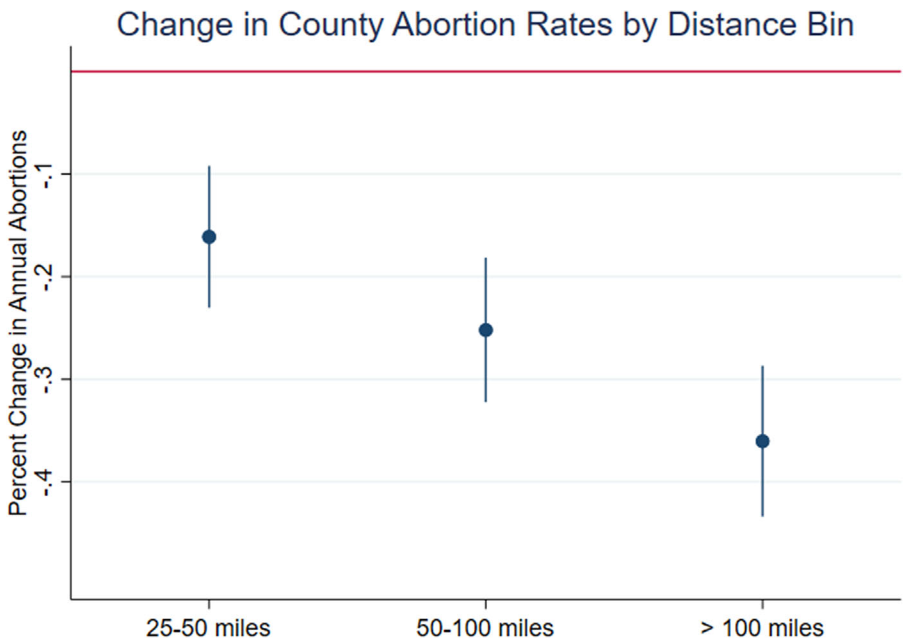
Figure 10 shows the trends over time in the abortions obtained in the two primary neighboring states, Illinois (Jatlaoui et al., 2018) and Minnesota (Jatlaoui et al., 2018, for years 2009 to 2016; Garcia-Lago, 2017, for 2017). Trends for Minnesota remain fairly flat across this time period. Since Illinois does not provide annual out-of-state abortion patient numbers broken out by state for years after 2017, we cannot see one year of changes in the number of Wisconsin residents who went to Illinois after the Appleton clinic closed. There are a similar number of Wisconsin residents obtaining abortions in Illinois in 2016 as in 2009 and the number is declining between 2014 and 2016. This suggests that the declines in abortions in-state in Wisconsin cannot be explained by substitution to out-of-state providers.

Moreover, 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 shown in Figure 3, the counties in Northwestern 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.

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 A3 shows the annual trends in abortions obtained by Michigan residents in Michigan for counties

[20] Using the Clogg, Petkova, and Haritou (1995) method of testing difference in coefficients across a model, we find that we cannot reject the null that the coefficients are equivalent at a $p = 0.14$ level.

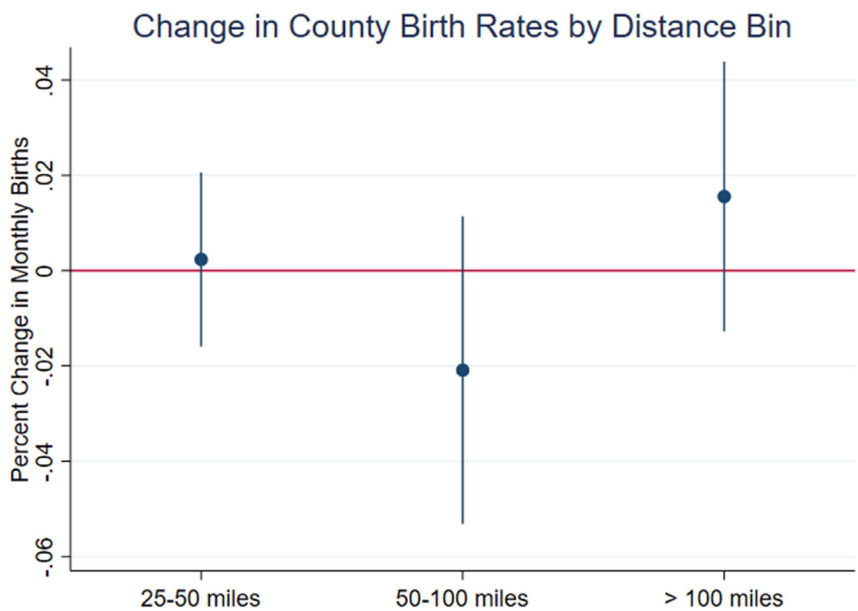
in the Upper Peninsula versus counties not in the Upper Peninsula. There is a clear substitution towards in-state abortions for residents of the Upper Peninsula.[21]



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, period 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 45; county populations of White, Black, and Other Races groups, and the distance to nearest Planned Parenthood clinic.

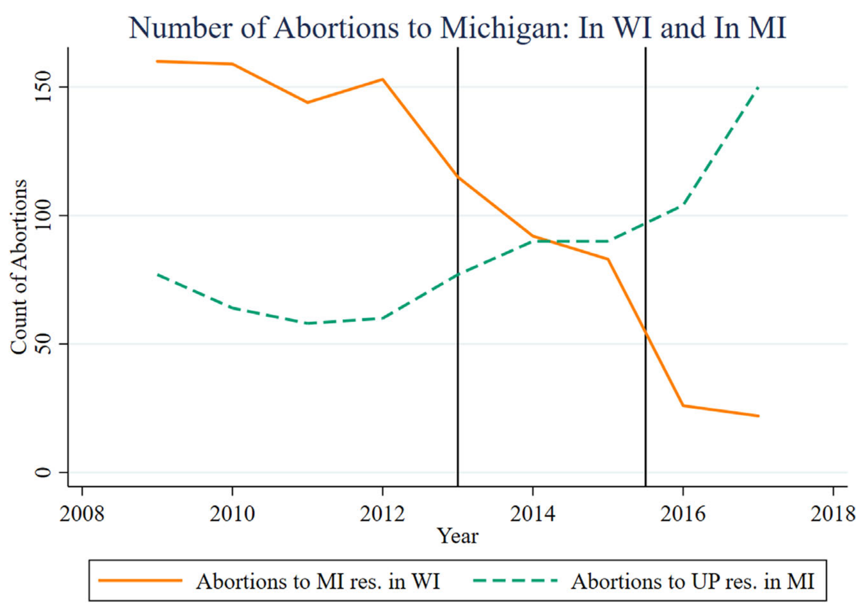
Figure A1. Effect of Distance from Nearest Clinic on Abortion Counts.

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



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, period 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 45; county populations of White, Black, and Other Races groups, and the distance to nearest Planned Parenthood clinic.

Figure A2. Effect of Distance from Nearest Clinic on Birth Counts.



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.

Figure A3. Number of Abortions to MI Residents Obtained in WI and MI.

Table A1. Demographic comparison: Wisconsin vs. other states.

	Wisconsin	States with Court Cases	USA	Texas
Birth Rate, per 1000 women	61.29	64.90	63.36	70.58
	(0.446)	(1.851)	(7.619)	(1.122)
Abortion Rate, per 1000 women	5.745	7.448	9.756	11.29
	(0.679)	(2.293)	(5.888)	(1.886)
Women 15 to 44, Percent Non-Hispanic White	0.810	0.756	0.680	0.413
	(0.0107)	(0.0908)	(0.169)	(0.0126)
Women 15 to 44, Percent Black	0.0751	0.165	0.122	0.131
	(0.0026)	(0.0868)	(0.111)	(0.0020)
Women 15 to 44, Percent Hispanic	0.069	0.0496	0.120	0.397
	(0.0043)	(0.0159)	(0.108)	(0.0067)
Number of Abortion Clinics Per State	3	3.85	16.16	21
Percent Women Living in a County w/o a Clinic	0.70	0.75	0.38	0.43

Notes: This table compares state characteristics of Wisconsin to states with upcoming judicial challenges to TRAP law (Arkansas, Indiana, Kentucky, Louisiana, Missouri, Mississippi, Ohio), the United States, and Texas. Row 1 reports birth rate per 1000 women, averaged over 2010 to 2016. Row 2 reports in-state abortion rate per 1000 women averaged over 2010 to 2016; the sample only includes states with comparable data (see Figure 5 for full list) meaning that column 2 only includes Arkansas, Indiana, and Ohio. Rows 3 to 5 report percent women 15 to 44 who are non-Hispanic White, percent Black, and percent Hispanic. Row 6 reports the average number of abortion clinics per state in 2017 (Jones et al., 2019). Row 7 reports the percent of women living in a county without an abortion clinic in 2017 (Jones et al., 2019).

Table A2. Pre-trend analysis: Effect of fertility trends prior to closures on distance change post-closure.

	(1) Δ Distance, 2013–2017	(2) Δ Distance, 2013–2017
Δ Abortions, 2009–2013	−1.868	
	(3.674)	
Δ Births, 2009–2013		−0.157
		(2.759)
Constant	14.88**	16.33**
	(5.513)	(4.915)
N	72	72

Notes: This table regresses the change in average travel distance between 2013 and 2017 on annual abortion rates between 2009 and 2013 (column 1) and monthly birth rates between 7/2009 and 7/2013 (column 2). Standard errors are in parentheses. *$p < 0.05$; **$p < 0.01$; ***$p < 0.001$.

Table A3. 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 (100 mi.)	−0.928**	−2.10*	
	(0.327)	(0.845)	
Distance Squared (1002 mi.)		0.690	
		(0.562)	
1(50 >Closest Clinic > 25 miles)			−0.628*
			(0.290)
1(100 >Closest Clinic > 50 miles)			−0.924***
			(0.259)
1(Closest Clinic > 100 miles)			−1.483***
			(0.281)
N	485	485	485
County Fixed Effect	Y	Y	Y
Time Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y

Notes: 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 five abortions in a year are excluded from regressions. All regressions include county fixed effects, period 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 45; county populations of non-Hispanic White, Hispanic, Black, and Other Races groups, and distance to nearest Planned Parenthood in-state. 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 to 50 miles away, 50 to 100 miles away, and more than 100 miles away. Standard errors are in parentheses. †p < 0.10; *p < 0.05; **p < 0.01; ***p < 0.001.

Table A4. 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 (100 mi.)	0.0765	−0.668*	
	(0.129)	(0.325)	
Distance Squared (100^2 mi.)		0.430*	
		(0.199)	
1(50 >Closest Clinic > 25 miles)			−0.0390
			(0.146)
1(100 >Closest Clinic > 50 miles)			−0.0956
			(0.155)
1(Closest Clinic > 100 miles)			−0.0721
			(0.188)
County Fixed Effect	Y	Y	Y
Time Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y
N	6,541	6,541	6,541

Notes: 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, period 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 45; county populations of non-Hispanic White, Hispanic, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the county. 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. Standard errors are in parentheses. $^{*}p < 0.05$; $^{**}p < 0.01$; $^{***}p < 0.001$.

Table A5. Impact of driving distance on birth rate, offset by 6, 9, and 12 months.

	(1) Births, 6 mo. Offset	(2) Births, 9 mo. Offset	(3) Births, 12 mo. Offset
Distance from Nearest Clinic (100 mi.)	0.0318*	0.0238	−0.00052
	(0.0155)	(0.0168)	(0.0220)
County Fixed Effect	Y	Y	Y
Time Fixed Effect	Y	Y	Y
County Level Controls	Y	Y	Y
N	6,541	6,331	6,115

Notes: This table reports coefficients of Poisson models of birth 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 one. Column 1 is our primary specification where we look at births relative to driving distance six months prior; column 2 uses births relative to driving distance nine months prior; column 3 uses births relative to driving distance 12 months prior. All regressions include county fixed effects, period 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 45; county populations of non-Hispanic White, Hispanic, Black, and Other Races groups, and the number of Planned Parenthood clinics located within the county. Standard errors are in parentheses. † $p < 0.10$; * $p < 0.05$; ** $p < 0.01$.

Table A6. Effect of distance increases from clinic closures on annual abortions and monthly births, robustness checks.

	Excluding Middle Tercile		Excluding Out of State	
	Abortion rate	Birth rate	Abortion rate	Birth rate
Panel A				
Distance from Nearest Clinic (100 mi.)	−0.208***	0.0356*	−0.225***	0.039*
	(0.07)	(0.0144)	(0.022)	(0.017)
Panel B				
1(50 >Closest Clinic > 25 miles)	−0.255***	0.006	−0.149	0.045
	(0.0473)	(0.011)	(0.0473)	(0.034)
1(100 >Closest Clinic > 50 miles)	−0.369***	−0.022	−0.191***	0.001
	(0.075)	(0.021)	(0.075)	(0.024)
1(Closest Clinic > 100 miles)	−0.366***	0.025*	−0.238***	0.047*
	(0.0563)	(0.012)	(0.0563)	(0.022)
N	349	3,361	290	5,722
County Fixed Effect	Y	Y	Y	Y
Time Fixed Effect	Y	Y	Y	Y
County Level Controls	Y	Y	Y	Y

Notes: This table reports coefficients of Poisson models of abortion and birth 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 one. In columns 1 and 2, we omit all counties that experience a change in distance between 33 and 95 miles. In columns 3 and 4, we omit counties where the closest clinic is ever an out-of-state clinic. Counties with fewer than five abortions in a year are excluded from regressions. All regressions include county fixed effects, period 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 45; county populations of non-Hispanic White, Hispanic, Black, and Other Races groups, and the distance to nearest Planned Parenthood clinic. Panel A reports a regression of abortion rate on distance from the nearest clinic; panel B 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. Standard errors are in parentheses. †$p < 0.10$; *$p < 0.05$; **$p < 0.01$; ***$p < 0.001$.