The Effects of Reduced Women's Health Clinic Capacity on Abortion, Fertility, and Sexual Health

Andrea M. Kelly*

November 1, 2019

Click here for latest version

Abstract

The impacts of access to healthcare on health outcomes are important to understand, yet difficult to measure. Researchers have documented the importance of health insurance, quality of local hospitals, and distance to the nearest healthcare facility, but have struggled to identify the impacts of constrained local capacity. I estimate the effect of reduced clinic capacity for abortion services using a natural experiment in Pennsylvania. In 2011, Pennsylvania passed new regulations requiring any facility that provided abortion services to meet ambulatory surgical facility standards, which ultimately caused the closure of almost half of the state's abortion facilities—but all closing facilities were near facilities that remained open. This meant that distance to the nearest clinic was largely unchanged, but local clinic capacity was reduced. I use both a difference-in-differences and a synthetic control design and find similar results. Reduced clinic capacity caused at least 13 percent fewer abortions to occur in the first 8 weeks of gestation and more abortions to occur at later gestational ages. There is also suggestive evidence that reduced clinic capacity reduced the overall abortion rate, increased rates of sexually transmitted infections, and increased the birth rate for black women, highlighting the potential for racial disparities in access to reproductive control technology.

JEL Classification: I18, J13, J18

Keywords: abortion, childbearing, sexual health, sexually transmitted infections

*Department of Economics, Texas A&M University, amkelly@tamu.edu. I thank Thomas Garvish at the Pennsylvania Office of Administration, Division of Health Informatics for data and useful conversations about the setting. I also thank Ann Chronister at the Pennsylvania Department of Health for data and assistance regarding abortion facilities in Pennsylvania. I also thank Kim Evert of Planned Parenthood Western Pennsylvania for useful conversations regarding the operations of an abortion-providing facility under the new regulations. This paper has benefitted from helpful comments from Jason Lindo, Joanna Lahey, Jennifer Doleac, Annalisa Packham, CarlyWill Sloan, Brittany Street, Meradee Tangvatcharapong, Chelsea Temple, Joshua Witter, and conference participants at the Western Economic Association International annual conference, and the 2019 Stata Texas Empirical Microeconomics Conference.

1 Introduction

Healthcare spending makes up approximately 18% of American GDP (Martin et al. (2018)), yet it is difficult to measure the degree to which healthcare services affect health outcomes. Access to healthcare is largely endogenous: healthy (or unhealthy) people may differentially select into areas with better local healthcare systems; wealthier people may spend more on healthcare but may also have better underlying health than the less wealthy; communities with better healthcare facilities may have invested in these facilities because their population was already sicker than average populations, etc. In any of these cases, comparisons of those with better access to healthcare services to those with worse access will not reflect the causal effects of access to healthcare. To establish a causal relationship between healthcare and health outcomes, researchers need a source of exogenous variation in access to healthcare. Toward this end, researchers have used changes in insurance (e.g. Anderson et al. (2012); Courtemanche et al. (2017); Finkelstein (2007); Finkelstein et al. (2012); Kolstad and Kowalski (2012)), shocks to local hospital quality (Doyle (2011)), and exogenous closures to local healthcare facilities (e.g. Countouris et al. (2014); Fischer et al. (2017); Lindo et al. (2017); Venator and Fletcher (2019); Quast et al. (2017)). This paper follows in the footsteps of the facility closures literature: I use exogenous closures to study the impact of reduced access to healthcare services on health outcomes.

Healthcare facility closures have been used to study the impacts of healthcare on various outcomes, including mortality (Countouris et al. (2014)) and reproductive health (Colman and Joyce (2011); Fischer et al. (2017); Lindo et al. (2017); Venator and Fletcher (2019); Quast et al. (2017)). In each of their settings, authors argue that the closure of a facility provides an exogenous shock to healthcare access.[1] In either case, closure of any facility results in two potential effects: increases in distance for some clients to reach a provider, and a higher number of potential clients at the facilities that remain open—yet it is difficult to separately identify the effects of these two mechanisms. Both mechanisms

[1]Sometimes these closures occur due to lack of profitability (Countouris et al. (2014)), while new governmental regulations force closures in other settings (Colman and Joyce (2011); Fischer et al. (2017); Lindo et al. (2017); Venator and Fletcher (2019); Quast et al. (2017)).

could be important: traveling further for healthcare can be costly and could prohibit use of services, but a congested facility may not be able to service demand and may be forced to turn away patients. Research has documented the importance of distance in access to healthcare services (Colman and Joyce (2011); Countouris et al. (2014); Fischer et al. (2017); Lindo et al. (2017) Venator and Fletcher (2019); Quast et al. (2017)), but does a change in the number of potential clients per clinic alone impact healthcare use?

In this paper, I exploit a natural experiment in Pennsylvania in which all new and existing abortion clinics were required to meet the same standards as ambulatory surgical facilities. These regulations were primarily related to construction of the building and staffing requirements. The new standards were costly to implement and caused the closure of 9 of the state's existing 22 abortion providers.[2] Notably, these closures all occurred in urban areas, where other clinics remained open. As such, they provide a setting in which a region's total clinic capacity changed while distance to the nearest clinic did not. I use a difference-in-differences approach to estimate the causal effect of reduced clinic capacity on abortion and fertility outcomes by comparing counties in Pennsylvania with few or no clinic closures (and therefore little to no change in clinic capacity) to those with major clinic closures.

Results suggest that this reduction in clinic capacity reduced abortion access to women in Pennsylvania—in the first year the laws took effect, the overall abortion rate was approximately 3.5% lower than would have been expected in the absence of the closures, though this effect is not statistically significant. Effects estimated by the synthetic control method indicate that reduced clinic capacity reduced overall abortion rates by an average of 8.5% per year in each of the years after the closures, which is statistically significant. Both the difference-in-difference and the synthetic control method results also indicate reduced clinic capacity delayed the timing of abortions. Difference-in-difference estimates suggest reduced clinic capacity reduced the rate of abortions occurring within the first 8 weeks of gestation by 13.8%, and increased the rate of abortions occurring in weeks 9–10

[2]While these types of regulations are increasingly popular among the United States, Pennsylvania's legislation was pushed forward after the discovery of an illegally-operating clinic in Philadelphia. The clinic was not meeting the standards in place at the time, yet the stories that came from this particular rogue clinic gave legislators the public support they needed to pass these laws.

by 22.6%, weeks 11–12 by 55.1%, weeks 13–14 by 15.8%, and in weeks 15 and beyond by approximately 32.3%. These effects are statistically significant and consistent across various robustness checks, such as adding in control variables, adjusting functional form, redefining the comparison group, among others. Synthetic control estimates are similar in direction and statistical significance, though magnitudes are somewhat different.[3] Using the difference-in-difference estimates, this amounts to a reduction of approximately 2700 abortions taking place in the first 8 weeks of gestation and an increase of 1200, 1500, 240, and 770 abortions in weeks 9–10, 11–12, 13–14, and 15+, respectively. In addition, I test for effects of reduced clinic capacity on rates of sexually transmitted infections and birth rates by mother's race. Using the difference-in-differences approach, I find suggestive evidence of an increase in rates of gonorrhea and syphilis and suggestive evidence of an 11% increase in the birth rate for black women and no change in birth rates overall or for white women, though synthetic control estimates for these outcomes are too noisy to provide conclusive evidence.

Obtaining an abortion later in the pregnancy could have serious implications for a few different reasons: first, the choice set for abortion procedures falls as gestational age increases—medical abortions are only available in Pennsylvania through the 10th week of gestation. Second, abortion services get more expensive for pregnant women as time goes on.[4] Third, as gestation continues, the risk of serious complications from abortion procedures increases (Sajadi-Ernazarova and Martinez (2019)).

Finally, I employ a synthetic control method to create a more fitting comparison group. The results from the synthetic control method largely support the results from the difference-in-differences method and suggest even larger effects on abortion rates: synthetic control results suggest a reduction in the overall abortion rate of approximately 8.5%, which is statistically significant at the 5% confidence level. Results from the syn-

[3]Synthetic control results suggest a larger reduction in early-term abortions (22.3%), a smaller increase in abortions in weeks 9–10, 11–12, and 13–14 (13.5%, 24.6%, and 9.3%, respectively). The average estimated effect for 15+ weeks is actually a reduction of 3%, though the rarity of this outcome makes estimated effects quite noisy.

[4]Planned Parenthood of Western Pennsylvania currently reports fees increasing from $435 in the first 11 weeks of gestation to $540 in weeks 12-13, over $815 in weeks 14-16, and over $915 after the start of week 17.

thetic control estimation for the timing of abortions and birth rates are consistent with difference-in-difference estimates, though the increase in birth rates for black women loses statistical significance.

This paper also contributes to the literature on access to reproductive control technology in the modern landscape. Access to reproductive control technology improves women's ability to avoid unintended births, which has been documented to improve economic conditions for women (Bailey (2006); Bailey et al. (2012); Bailey et al. (2018); Goldin and Katz (2002); Myers (2017)). Abortion is also one of the safest medical procedures to obtain,[5] yet abortion providers have become an increasingly regulated medical body. According to the Guttmacher Institute (2019), 24 states currently have "laws or policies that regulate abortion providers and go beyond what is necessary to ensure patients' safety." These regulations have caused clinics to close their doors (Colman and Joyce (2011); Fischer et al. (2017), Lindo et al. (2017), Quast et al. (2017)), and the closures have reduced access to both abortion and family planning services—all with no evidence that these regulations actually improve the safety of abortions (Roberts et al. (2018)).[6] These new barriers to abortion access—in addition to better knowledge and use of effective contraception—imply that changes in access to aboriton services may generate different effects than changes in earlier decades.

Supply-side abortion regulations, or regulations that target the abortion-providing facilities rather than individuals seeking abortions, have become increasingly popular over time. Work has shown that distance to the nearest abortion clinic is a crucial factor in access to abortion services (Colman and Joyce (2011); Fischer et al. (2017); Lindo et al. (2017); Venator and Fletcher (2019); Quast et al. (2017)). However, to date there

[5]According to National Academies of Sciences and Medicine (2018), abortions are safe and effective, and complications are rare. All four of the main abortion methods (medication, aspiration, dilation and evacuation, and induction) were studied. Additionally, according to the Pennsylvania Department of Health's Annual Abortion Reports, the total complication rate in any given year of this study ranged from 0.001 to 0.005. This means that in Pennsylvania, in a given year, only 1/10th of a percent to 1/2 of a percent of all abortions had some kind of complication.

[6]To my knowledge, there is no causal work that shows the impact of ambulatory surgical facility standards on complication rates or other adverse outcomes related to abortion. Roberts et al. is a correlational study that finds that differences in abortion-related adverse events for women who obtained abortions in ambulatory surgical centers relative to women who obtained abortions in office-based settings are not statistically significant.

is little evidence of the importance of clinic congestion, and the evidence that exists is somewhat conflicting. Lindo et al. (2017) find that clinic congestion reduced abortion rates, but Venator and Fletcher (2019) find no effect of increased clinic congestion on abortion rates. In a setting in which distance remains unchanged, does an increase in the number of potential clients per open clinic impact abortion rates, the timing of abortions, or birth rates?

The remainder of this paper is organized as follows. In the next section I provide background information on abortion provider regulation guidelines and the natural experiment setting in Pennsylvania. Then I discuss the data and methods I use to analyze the causal effects of reduced clinic capacity on abortion and fertility outcomes and present the results of the analysis. I next show heterogeneous treatment effects and robustness of the estimates, and discuss synthetic control estimation and results. Lastly, I conclude and discuss the implications of this and similar policies.

2 Background

2.1 Abortion provider regulations and their effects

Abortion facility regulations have been growing in popularity in the United States. While different states have passed slightly different packages of regulations, they often include staffing requirements, hospital admitting privileges, and building requirements. The implementation of such regulations has been studied in health economics literature: for example, Lindo et al. (2017), Fischer et al. (2017), and Quast et al. (2017) use a major legislative change in Texas as a natural experiment to estimate the effects of supply-side restrictions on abortion access and find that these restrictions reduce abortion rates.[7] These studies all examine the effect of Texas' House Bill 2 (HB2), a bill that required

[7]Colman and Joyce (2011) also studies a law change in Texas called the Women's Right to Know Act, which had a new requirement of information to be provided to women considering an abortion in addition to a new requirement that abortions after 16 weeks gestation be obtained in ambulatory surgical facilities. This change reduced the number of abortions occurring after 16 weeks gestational age, but increased out-of-state travel for abortions and the number of abortions obtained at 15 weeks of gestation. Effects were persistent: even after new facilities opened that were qualified to perform abortions after the 16th week of gestation, the number of abortions obtained in Texas at this gestational age remained well below pre-Right-to-Know levels.

all doctors who provided abortions to have admitting privileges at nearby hospitals (no further than 30 miles from the abortion clinic), required abortion facilities to meet surgical facility standards, and banned abortions after 20 weeks of gestation. The bill caused the closure of 22 of the state's existing 41 clinics to close their doors, and eventually was overturned by the United States Supreme Court in 2016, stating that the 'provisions constituted an undue burden and are therefore unconstitutional' (Domonoske (2016)). Fischer et al. (2017), Lindo et al. (2017), and Quast et al. (2017) all study the impact of clinic closures on abortion rates, focusing on the impact of increases in distance to the nearest clinic on abortion behavior. Each study finds substantial reductions in abortion rates: the estimated reductions in abortion rates range from 10-20 percent, with variation in estimate size coming from different estimation strategies.[8] Given a pre-regulation abortion rate of approximately 12 abortions per 1,000 women of childbearing age in Texas, these estimated effects imply a reduction in abortions of 1-2 abortions per 1,000 women of childbearing age.[9] Additionally, Venator and Fletcher (2019) studies a similar setting in Wisconsin, in which new regulations forced the closure of two of the state's five existing clinics. Venator and Fletcher (2019) document an average increase in distance to the nearest clinic of 55 miles, with some women experiencing significantly larger increases. The increase in distance in Wisconson cased a reduction in the abortion rate of 25% and an increase in the birth rate of 4%. Both Venator and Fletcher (2019) and Lindo et al. (2017) attempt to separate effects of changes in distance from changes in congestion: Venator and Fletcher (2019) find no effect of clinic congestion on abortion or birth rates, and Lindo et al. (2017) find that both distance and congestion reduce abortions, but that increased congestion may account for 59 percent of the effects of clinic closures on abortion and find that an increase in the average number of women per open clinic in an area of 100,000 reduces abortion rates by 5 percent. The setting I study is unique: distance to the nearest clinic is largely unchanged, so I am able to separately identify the effects of reduced local clinic capacity. Additionally, the new regulations in Texas

[8]Fischer et al. (2017) assume a linear relationship between distance and effects; Lindo et al. (2017) allow for non-linearities; and Quast et al. (2017) use a linear regression but have fewer post-HB2 data-points.

[9]Lindo et al. (2017) estimate a reduction of 11,900 abortions in the two years of the law's enactment.

massively cut funding for non-abortion providing family planning clinics, which may be interacting with abortion clinic closures and could be contributing to the effects of clinic closures alone. In Wisconsin, both distance and congestion change simultaneously, and the state has lower pre-regulation access to abortion than Pennsylvania, which could contribute to the detected effects of clinic closures.

I also contribute to this literature by studying a new population: Pennsylvania is different from Texas both geographically and demographically. Texas shares a border with Mexico; Pennsylvania is almost entirely bordered by other states in the US (with the small exception of the Erie area, which borders Lake Erie). Texas is also much larger than Pennsylvania, geographically. Driving from El Paso to Dallas or El Paso to Houston (representing West to East travel across the state) takes approximately 9 or 10.5 hours, respectively, and the public transportation options increase potential travel time significantly. Driving from Pittsburgh to Philadelphia (again representing West to East travel across the state) can take less than 4 hours, and public transportation options can take as little as one hour longer than driving. Texas also has a large Hispanic population—falling second in the nation with a Hispanic population of over 36 percent—while only approximately 5 percent of Pennsylvanians are Hispanic.[10] Fertility behavior for Hispanic women has been declining much more dramatically than fertility behavior for other ethnic groups over the past decade (Tavernise (2019)), so abortion responses in Texas may not be representative of abortion responses in less-Hispanic states. For each of these reasons, evidence from Pennsylvania helps to inform how similar policy changes may impact other states. Pennsylvania and Wisconsin are somewhat more similar, with more similar demographic and geographic profiles. However, Wisconsin had fewer abortion clinics per population of childbearing-aged women prior to the closures than Pennsylvania, which could impact the effects of the closures in either direction: perhaps women in Wisconsin were already adjusted to limited access to abortion services, which would predict smaller effects of clinic closures; perhaps closures in Wisconsin are more binding

[10] Wee (2016) shows that this places Texas in second for largest share of the population that is Hispanic; in the tenth-place state, Illinois, only 17 percent of the state's population is Hispanic. This means that Texas has more than double the 90th percentile of the US's share of population that is Hispanic.

than in Pennsylvania due to the scarcity of services, which would predict larger effects of clinic closures.

2.2 Pennsylvania SB732

In December of 2011, Pennsylvania SB732 (also known as Act 122 of 2011) was enacted into law, though clinics were given a “grace period” to meet the new standards. This law had several components, all of which had the stated goal of improving the safety of abortion services. First, the law redefined “abortion facilities” to include any public or private hospital, clinic, center, medical school, medical training institution, physicians office, infirmary, or other institution which provides surgical services meant to terminate the clinically disposable pregnancy of a woman.[11] Second, abortion facilities were required to meet the same fire and safety standards, personnel and equipment requirements, and quality assurance checks as ambulatory surgical facilities. These standards included increased hallway width, increased operating room size, increased staffing requirements, each facility had to have transfer priviliges to a hospital, and elevators had to meet certain size guidelines. Third, this legislation also enacted annual and random inspections of abortion facilities in order to ensure facilities were meeting the requirements. Prior to the passage of this law, annual inspections were not standard.

Before the law was passed, Pennsylvania had 22 open abortion clinics. Between April and December of 2012, 9 abortion facilities permanently closed their doors, and still others closed temporarily to make the necessarily construction changes. Most of these closures occurred in urban areas, resulting in changes in the number of women of childbearing age per open clinic, while the distance from each county’s population centroid to the nearest open clinic remained constant. In fact, 5 of the 9 clinic closures that occurred in 2012 were within the city of Pittsburgh, leaving only 2 open clinics in the entire Pittsburgh service region.[12] This setting therefore provides a unique opportunity to understand the effects

[11]Facilities that only provided medical abortion services were exempt, although prior to the law’s passage there were no facilities that only provided medical abortions.

[12]Three of the other clinic closures were in Philadelphia, leaving 9 open clinics in the Philadelphia region; 1 clinic closure was in Allentown, leaving 2 clinics open in the Allentown region. Because these closures did not reduce their respective region’s clinic supply as dramatically as the Pittsburgh closures, Allentown and Philadelphia will be a part of the comparison group. However, since Philadelphia and

of reduced clinic capacity, rather than distance from the county's population centroid to the nearest open clinic, on abortion rates, abortion timing, and birth rates.

While I cannot directly measure the congestion in a given clinic, anecdotal evidence suggests that clinic congestion did increase after the closures for the clinics that remained open. I spoke with Planned Parenthood of Western Pennsylvania's CEO and President, Kimberlee Evert, who said that the Planned Parenthood in Pittsburgh, which was one of the two clinics in the Pittsburgh area that remained open, had to close for some time in 2012 to meet the new standards. Despite their temporary closure, the percentage of abortions in the Pittsburgh area that were performed by Planned Parenthood grew from 31% in 2011 to 42% in 2013, although the number of abortions fell over time. Additionally, wait times for abortion services increased dramatically after the closures began. Some of the increase in wait times is persistent to this day: women calling to request abortion services typically have to wait one week for a medical abortion, or two weeks for a surgical one. Ms. Evert says this is an improvement—at one point after the law's passage, the wait times were at least double that for each type of abortion procedure. Given this anecdotal evidence, I argue that the mechanism for any observed effects is clinic congestion (Evert (2019)).

3 Empirical approach

This section describes the data and empirical approach I use to estimate the causal effects of reduced clinic capacity caused by the passage of SB732 and the resulting clinic closures.

3.1 Data

To define treatment and comparison groups, I first define a county's treatment-defining abortion service region as the nearest city in 2006 that had an abortion clinic open.[13]

Allentown experienced some closures, I consider my results to be a lower bound for the true effects of increased clinic congestion.

[13]All distance calculations are based on geolocations for abortion-providing facilities and county population centroids. I calculate distance to the nearest provider using the Stata *georoute* program (Weber

Any counties for which the original, treatment-defining service region experienced an endogenous closure prior to the passage of SB732 are dropped from the analysis.[14] Because Pittsburgh is the abortion service region most affected by SB732, I use counties that were first observed in the Pittsburgh abortion service region as my 'treated' counties. I use all other counties as my 'comparison' counties.[15] Figure 1 shows the counties defined as treated in blue, those defined as comparison counties in orange, and omitted counties in gray.[16] Panel A shows the open clinics in 2011, prior to the law's passage, and Panel B shows the open clinics in 2013, after the 2012 closures which resulted from the law's passage. Figure 2 shows the clinic locations in Pittsburgh, in 2011 and 2013.[17]

In order to estimate the effects of clinic congestion caused by reduced clinic capacity, I create a measure of the "abortion service population," following Lindo et al. (2017). While this measure does not perfectly capture actual clinic congestion, it does capture the expected increase in patient loads faced by the reduced number of clinics in operation.[18] To construct the average service population, I first assign each county c in time period t to an "abortion service region" r according to the location of the closest city with an abortion clinic. The average service population is the ratio of the population of women aged 15–44 in the service region to the number of clinics in the service region:

$$ASP_{c,r,t} = \frac{\sum_{c \in r} population_{c,t}}{number\ of\ clinics_{r,t}} \tag{1}$$

To create abortion rates, I use data from the Pennsylvania Department of Health, which

and Peclat (2017)). This program estimates the travel distance from the population centroid of each county (United States Census Bureau, 2018) to the geocoded address of the nearest in-state operating abortion clinic. Figure 4 shows that the population-weighted average distance to the nearest clinic changes very minimally over the time under observation, so I argue that observed effects are caused by the dramatic increases in the average number of women of childbearing age per open clinic.

[14]The abortion service regions being dropped are East Stroudsburg, Erie, Huntingdon, and State College.

[15]Omitting counties with endogenous closures still keeps 36 counties in the analysis, and most of the excluded counties are rural. Additionally, Figures ?? and ?? and Tables A2 and A3 test the robustness of the results to the inclusion of omitted counties in the comparison group. Results from this robustness test are further discussed in Section 5.

[16]Counties are omitted if their nearest clinic in 2006 closed prior to the passage of the law. These closures cannot be seen as exogenous.

[17]Four of the 7 clinics in Pittsburgh were within the same suite of offices: the top right dot in 2011 actually represents 4 unique abortion facilities; in 2013 only one facility remained open in that location.

[18]However, if clinics react in such a way that they reduce their wait times, estimated effects will be an underestimate of the effects of clinic congestion.

tracks various abortion statistics over time. Importantly, these data contain the number of abortions obtained per county per year by age group, as well as the number of abortions obtained per county per year by gestational age at the time of the abortion. I will use both measures, as well as population denominators from to construct my outcomes of interest, namely county-level abortion rates by age group as well as by gestational age at the time of the abortion. In future work, I plan to also estimate the effects on birth rates by age group and race, both for women living in Pennsylvania and neighboring states. To do so, I will use natality data from the National Vital Statistics System from the Center for Disease Control.

Table 1 summarizes the variables used in my analysis: abortion rates by age and gestational age by mother's county of residence, average service population (the number of women of childbearing age per open clinic in the region), abortion rates, and variables measuring county demographics: age and racial composition (SEER, 2018), poverty rate (Census Bureau, 2018) and unemployment (BLS, 2018). Data in this table are broken down into the period before the law was enacted (2006–2011) and the period after the law was enacted (2012–2016). Notably, both groups have similar pre-period poverty and unemployment rates, and both are predominantly white.

3.2 Identification strategy

I use a generalized difference-in-differences approach with county-specific linear time trends to estimate the causal effects of reduced clinic capacity. This approach exploits within-county variation over time and controls for aggregate time shocks, as well as fixed differences across counties over time and differences in pre-regulation trends. In order for this approach to be valid, it must be true that changes in abortion rates for comparison counties provide a good counterfactual for the changes in abortion rates that would have been observed for reated counties (relative to pre-regulation trends), if clinic capacity had remained unchanged. I also report estimates that do not control for county-specific linear time trends, but focus on estimates controlling for trends, which are more conservative.

Because abortions are discrete occurrences rather than continuous, and because there are age-county-year observations with zero observations, I use a Poisson model for this analysis.[19] Thus, my approach to estimating the effects of changes in average service population on the abortion rates corresponds to the following equation:

$$E[y_{ct}|capacity_{c,t-k}, \alpha_c, \alpha_t, X_{ct}] = exp(\sum_{k=1}^{5} \theta_k capacity_{c,t-k} + \alpha_c + \alpha_t + \gamma_c * t) \quad (2)$$

where y_{ct} is the outcome of interest for residents of county c in year t; $capacity_{c,t-k}$ is an indicator for whether county c experienced reduced capacity in year $t-k$; α_c are county fixed effects; α_t are year fixed effects; and $\gamma_c * t$ are county-specific linear time trends.[20] All reported standard-error estimates are clustered on the county to account for correlation within counties over time. I use this model to estimate effects on abortion rates for women of various age groups, abortion rates for various gestational ages, and birth rates by mother's race.

I also estimate effects on the share of abortions occurring at a given gestational age. Since this outcome is a share rather than a continuous variable, I use population-weighted least squares. The model uses the same fixed effects as the Poisson regression, but uses a county's total population as its weighting measure. The weighted least squares equation is as follows:

$$E[ShareGestAge_{ct}|capacity_{c,t-k}, \alpha_c, \alpha_t, X_{ct}] = \sum_{k=1}^{5} \theta_k capacity_{c,t-k} + \alpha_c + \alpha_t + \gamma_c * t \quad (3)$$

and the variable definitions are the same as the prior regression, with the exception that $ShareGestAge_{ct}$ is the share of the county's existing abortions that takes place within a given gestational age. For example, the $Share \leq 8$ would be the number of abortions that occurred for women residing in a county at $\leq$8 weeks of gestational age divided by the total number of abortions in that county in that year. However, since this outcome

[19]Because this is a Poisson model, not negative binomial or other non-linear model, the model is not subject to inconsistency caused by the incidental parameters problem associated with fixed effects. The main argument against Poisson is the possibility of overdispersion, which is corrected for by calculating sandwiched standard errors Cameron and Trivedi (2005). Additionally, Allison and Waterman (2002) shows that the negative binomial model is not a true fixed effects model.

[20]Because treatment may affect the trend in outcomes in the post-period, the county-specific linear trends are based only on pre-period outcome data. This allows the treatment to affect post-period trends in outcome, but controls for any differential trends between counties in the pre-period.

is conditional on an endogenous variable (the occurrence of abortion), I consider these results to be supplementary and the results on rate of abortion at a given gestational age to be the primary results of interest.

To further test the robustness of my results, and to improve the match of the comparison group to the treated group in the pre-period, I use the synthetic control method. I use this method to estimate the effect of reduced clinic capacity on logged abortion rates and birth rates, comparing the outcomes for the Pittsburgh area to the outcomes of a "Synthetic Pittsburgh Area" (Abadie et al. (2010)). First, I create a "Pittsburgh Area" observation: I collapse outcomes for treated counties by a population-weight. I then use data on abortion and fertility behavior from comparison counties. I identify the weighted average of comparison counties that best matches the outcome of interest observed in the Pittsburgh area prior to the closures. Here the identification assumption is that the synthetic Pittsburgh area provides a good counterfactual for abortion and fertility outcomes that would have been observed in the Pittsburgh area, absent the new regulations. If the assumption holds, the difference between outcomes for the Pittsburgh area and the synthetic control will provide an unbiased estimate of the causal effect of reduced clinic capacity. In order to execute this strategy, I select non-negative weights for each potential "donor county" to minimize the function:

$$(X_{Pitt} - X_{SC}W)'V(X_{Pitt} - X_{SC}W) \tag{4}$$

where X_{Pitt} is a (K×1) vector of variables measuring abortion or fertility outcomes from 2006-2011, X_{SC} is a (K×J) matrix containing the outcome variables for other counties in Pennsylvania, W is a (J×1) vector of weights summing to one, and the diagonal matrix V contains the "importance weights" assigned to each variable in X. I include the outcome of interest (abortion rate, rate of abortion at a given gestational age, share of abortions in a given gestational age, STI rate, or birth rate) observed in each pre-regulation year in X, putting 10% of the weight on 2011 and splitting the remainder of the weight evenly among outcomes in 2006-2010.[21]

[21] Splitting the weights evenly among all pre-period years created convexity issues that made the code

To conduct inference, I estimate the distribution of estimated treatment effects under the null hypothesis of a zero treatment effect and reassign treatment separately to each county in the donor pool to estimate a placebo effect for each county. I then construct p-values for the estimated effect for the Pittsburgh area, given the ratio of the post-period root mean squared error to the pre-period root mean squared error. I use this approach for each outcome of interest: abortion rates by age group, abortion rates by gestational age, share of abortions occurring at each gestational age, rates of sexually transmitted infections, and birth rates by race.

4 Results

4.1 Graphical Evidence of the Proposed Mechanism

First, to demonstrate the increase in average service population, I plot the number of childbearing-aged women per open clinic in an abortion service region. Figure 3 shows the average service population by treatment status. This figure demonstrates that average service population was relatively constant prior to the passage of the new regulations, and that the treated and comparison counties' average service populations tracked prior to the regulations. However, after the regulations were passed and clinics closed, we see both treated and comparion counties' average service population increase, but the treated counties' average service population increases much more dramatically.[22] Meanwhile, Figure 4 shows that distance to the nearest clinic did not change for the treated or the comparison counties over time—meaning that any observed effects should be a result of reduced clinic capacity rather than changes in distance to the nearest clinic.

---

unable to run. My results are robust to different weighting of pre-period outcomes, but I have chosen to put the least weight on the latest pre-regulation year because this provides a slight test of the fit of the synthetic control in the pre-period.

[22] There is a jump in average service population in both the treated and comparison counties in 2010. Figures A3 and A4 show that the main results are robust to excluding this 'odd' pre-period year.

4.2 Graphical Evidence for Outcomes from Difference-in-Difference Models

Before discussing my preferred estimates of the effects of clinic congestion on abortion rates, abortion timing, rates of sexually transmitted infections, and births, I first present graphical evidence to support my main results and the validity of my research design. In Figure 5 I present results graphically for the overall abortion rate, as well as the abortion rate for teenaged women and non-teenaged women. The results shown in these figures are from my baseline specification, which includes county and year fixed effects and a county-specific linear time trend. Results without the inclusion of county-specific linear time trends demonstrate similar patterns (see Figures A1 and A2), though results from my preferred specification (with linear time trends) are more conservative. Estimates prior to the new regulations provide a placebo test for my model, and the model passes these tests since the estimates are not statistically different from zero. While estimated effects on abortion rates overall and for non-teen women are negative, the effects are not statistically significant. In all cases, any reduction in abortion rates seems to disappear by 2015. This could be evidence clinics were able to adjust their capacity over time, perhaps by constructing new space or by hiring more physicians.

Next, I look at abortions occurring at various gestational ages. Figure 6 shows the average share of abortions occurring in each gestational age I consider separately for treated and comparison counties. These figures provide additional evidence that the comparison counties provide a good counterfactual for the treated counties, as the share of abortions occurring at any given gestational age is very similar in treated and comparison counties prior to the law's passage. These figures also show that abortions seem to be shifting toward later gestational ages—from $\leq$8 weeks to 9–10, 11–12, and possibly 13–14 and 15+ weeks of gestational age for women living in counties that experienced a reduction in clinic capacity. Figure 7 shows estimated effects using the same baseline regression as the abortion rate-by-age figures (Figure 5). These figures demonstrate that the model passes the pre-regulation placebo tests in most cases, though it does fail in some of the later gestational age outcomes. The estimated effect is negative for the rate of abortions occurring in the first 8 weeks of gestation, and is statistically significant and economically meaningful. The estimated effects are positive for the rate of abortions occurring in weeks 9–10 and 11-12. There also appear to be positive effects in the rate of abortions occurring in weeks 13–14 and 15+, though evidence for these effects is more suggestive. Figure 8 shows the results when using a population-weighted least squares, with the share of abortions taking place within a given gestational age as the outcome. The results follow the same pattern as those for the rate of abortion occurring at a given gestational age. The estimated effect for share of abortions taking place within the first 8 weeks of gestational age is large and negative from 2012–2014, at the same time that the effect for the share of abortions occurring in weeks 9–12 is positive in the treated counties relative to what would have been expected in the absence of clinic closures.

Abortion-providing facilities also provide preventative healthcare and testing and care for sexually transmitted infections (STIs). Reduced abortion clinic capacity could, therefore, reduce access to preventative healthcare or care for STIs. The state of Pennsylvania publishes annual, county-level data on STIs, with data supression at county-year observations of 1-5. Using these data and Equation 2, Figure 9 shows the estimated effects of reduced clinic capacity on STI rates. STI rates are noisy, and the model does not pass every pre-period placebo test. However, there does appear to be some suggestive evidence of increases in some STI rates: the estimated effects for gonorrhea and syphilis are positive and statistically significant.

Given the documented disparities in access to reproductive healthcare based on race and ethnicity, I would ideally run a similar analysis for abortion rates by race and ethnicity. Unfortunately, these data do not exist at the county-by-year level for the state of Pennsylvania. Insteaad, I estimate effects for birth rates occurring by mother's race and ethnicity. Figure 10 shows the estimated results using the same specification as was used in the abortion figures, and passes the pre-period placebo tests for the rate of births occurring to the total population, white mothers, or black mothers.[23] While there is

[23] I also have birth rates for Hispanic women, but the population of Hispanic women in Pennsylvania is so small that estimated effects are too noisy to provide any information. Results are available upon request.

very little movement for the total population or the population of white mothers, there is a statistically significant and positive estimated effect in births to black mothers. This suggests that clinic closures impact women of different races differentially.

4.3 Graphical Evidence for Outcomes from the Synthetic Control Approach

In order to test whether estimated effects are persistent across other models, I next present graphical evidence from a synthetic control approach. First, I look at abortion rates by age group. Figure 11 presents synthetic control estimates for abortion rates by age group on the left-hand side, with the corresponding randomization inference figures on the right-hand side. Estimated effects generally follow the same pattern as the difference-in-differences estimates, but using the synthetic comparison group generates larger and are statistically significant. Effects broken down by age group (teen and non-teen) are also large but are not statistically significant. Reductions in abortion rates are persistent across the entire post-period, suggesting that the difference-in-differences estimation may be understating later period effects.

Next, I look at timing of abortions. Figures 12-13 present the synthetic control estimates for abortion rates by gestational age group on the left-hand side, with the corresponding randomization inference figures on the right-hand side. Again, the synthetic control largely supports the difference-in-differences findings: early-term abortion rates drop dramatically as a result of reduced clinic congestion. Results for increases in later abortions, though following the same pattern as DiD results, are not statistically significant. Figures 14-15 show the synthetic control estimates for the share of abortion rates occurring at a given gestational age on the left-hand side, with the corresponding randomization inference figures on the right-hand side. Results indicate that the abortions that did occur after the new regulations occurred later in the pregnancy than would have been expected in the absence of the closures. Taken together, these results suggest that any abortions taking place after the clinic closures did so at a later gestational age than they would have if the clinics had remained open.

Next, Figure 16 presents the synthetic control estimates for STI rates on the left-hand side, with the corresponding randomization inference figures on the right-hand side. These results differ from the difference-in-difference estimates: the positive estimated effects for gonorrhea and syphilis rates are either not obvious or do not exist, and the only effect that seems to exist is an increase in chlamydia rates. Taken with the difference-in-difference results, this suggests that there may have been non-negative effects on rates of sexually transmitted infections, but the evidence is suggestive at best.

Finally, Figure 17 presents synthetic control estimates for birth rates on the left-hand side, with the corresponding randomization inference figures on the right-hand side. Estimates generated by the synthetic control method are noisy, but follow the same general pattern as estimates from the difference-in-differences approach. These results suggest that reduced abortion clinic capacity had little to no effect on birth rates, with suggestive evidence of an increase in the birth rate for black women.

4.4 Difference-in-Difference Estimates

The figures discussed in the previous section show graphical evidence of effects. In this section, I discuss point estimates from the main specification using results from the difference-in-differences approach. Results from the baseline specification for each age group are shown in Table 2.[24] All columns include year fixed effects and county fixed effects, as well as county-specific linear time trends. The results from Table 2 indicate that the total abortion rate fell in the treated counties in 2012–2014 relative to what would have been expected in the absence of closures (had the abortion rate continued on pre-period trends) though the estimates are never statistically significant.

Because reduced clinic capacity seems to have impacted the timing of abortions, estimated effects for abortion rates at various gestational ages are shown in Table 3. All columns include year and county fixed effects, as well as county-specific linear time trends. These results indicate that the rate of abortions occurring within the first 8 weeks of pregnancy fell by an average of 13.8% in each year after from 2012–2016, which is significant

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

at the 99%-confidence-level. The effects are largest and most statistically significant in the first 3 years, with the largest estimated effects occurring in 2013 with a reduction of 28.2%. This corresponds to an average decrease of 546 abortions in the first 8 weeks of gestation per year (or 2733 over the entire post-period). Additionally, the estimated effects for the rate of abortions occurring in weeks 9–10 and 11–12 is positive: the average estimated effecst over the entire post periods are roughly 22.6% and 55.1% respectively. This corresponds to an increase of 255 abortions in weeks 9–10 per year (or 1276 over the entire post-period) and an increase of 317 abortions in weeks 11–12 per year (or 1588 over the entire post-period). The estimated effects are also positive for the rate of abortions taking place in weeks 13-14 and 15+, but these estimates are noisier due to the rarity of these outcomes. These results suggest that reduced clinic capacity reduced the rate of abortions occurring in the first 8 weeks of gestational age and increased the rate of abortions occurring at later gestational ages.

Next, I present results for the estimated effects of reduced clinic capacity on the share of abortions occurring in given gestational age groups. Table 4 shows results for the weighted least squares estimates on the share of abortions occurring within given gestational ages. These results show that the share of abortions occurring within the first eight weeks of gestational age fell by approximately 10.8 percentage points on average between 2012 and 2016. Similarly, the estimated effect of the share of abortions occurring in later gestational ages was positive: 4.3 percentage points on average for weeks 9–10, 4.6 percentage points for weeks 11–12, and 1.3 percentage points for weeks 15 and beyond. The share of abortions occurring in weeks 13–14 also increased, though the estimated effect is only statistically significant in the second year of the new regulations. These results suggest that for abortions that did occur after the clinic closures, more abortions occurred later in the pregnancy than would have been expected in the absence of clinic closures.

Finally, since data for abortions by race and ethnicity do not exist at the county-level in Pennsylvania, I test for differential access to abortion services by race and ethnicity, using birth rates as the outcome of interest. Because there is not a significant reduction in abortions for women overall, estimated effects for the birth rate for all women should be zeroes. Table 5 shows the results for the Poisson estimation of effects on birth rates by race. Results suggest little to no effect on birth rates overall with positive estimated effects for birthrates for black women. The average estimated effect over the post-period is a 12 percent increase in the black birth rate. However, results are sensitive to the inclusion of linear time trends, so I interpret observed effects as suggestive.[25]

5 Validity and Robustness

In this section, I present a set of robustness checks to provide additional support for my identifying assumption. First, one might be concerned that the counties that remain in the sample are somehow different from the counties that are omitted in fertility and abortion behavior. This could create a problem for external validity. To test this, I include the previously excluded counties in the comparison. That is, I keep the treated group the same, but add the counties that experienced endogenous closures of their nearest abortion facility (closures prior to 2011) into the comparison group. Point estimates for this analysis can be found in Table A2. The point estimates in this table are quite similar to those shown in Table 2. Similarly, estimated effects for abortion rates at various gestational ages remain robust: Table A3 presents the point estimates. All results are similar to those shown in the previous section.

Next, I provide further support that the mechanism for the effects is, in fact, clinic congestion. One may be concerned that small changes in distance in urban areas have a meaningful impact on abortion access for women living in those urban areas. To address this concern, I re-run the main analysis, dropping Allegheny County (home of Pittsburgh). If all of my effects were due to women in the Pittsburgh losing access to these facilities (perhaps via increased difficulty in using public transport), this analysis would show no effects from the closures. Tables A4 and A5 present the point estimates from this analysis. Both the direction and the magnitude of the estimates are quite similar to those presented

[25] The data include birth counts by county for Hispanic women, but the share of Pennsylvania that is Hispanic is so small that estimates are noisy and cannot be conclusive.

in the main analysis. The takeaways from these tables are largely the same as those from the full sample, which supports the idea that effects are driven by clinic congestion rather than changes in distance.

Finally, I consider the possibility of inter-state travel for women wishing to obtain abortions. Due to the gravity of the potential outcomes of not obtaining an abortion when desired, women may travel to nearby states to obtain an abortion. If this is the case, then women traveling into Pennsylvania for abortions could also be impacted by reduced clinic capacity. This would be especially true for women traveling to the Pittsburgh area rather than other parts of the state. To test for effects on out-of-state women obtaining abortion in Pennsylvania, I plot the natural log of the rate of abortion for women traveling to Pennsylvania for abortions, for each of Pennsylvania's six neighboring states.[26] Figures A5 and A6 show the natural log of the abortion rates for each of these six states, with a vertical line drawn in at the passage of the new regulations. There do appear to be some declines in abortion rates for some states and some age groups: teenagers in all neighboring states seem to experience a reduction after 2011; West Virginia also appears to demonstrate a reduction for almost all age groups. The figures for abortion rates by gestational age overall do not exhibit evidence of delays in abortion timing. This suggests that closures within Pennsylvania may impact abortion or fertility behaviors for women in neighboring states, particularly in states with limited access to abortion services.

The other possibility for inter-state travel is Pennsyvlania women traveling to other states for abortion services. If closures in Pennsylvania force women in treated counties to obtain abortions out-of-state rather than in Pennsylvania, their abortions would not be collected in the Pennsylvania Department of Health data. This means that any negative estimated effects in abortion rates *could* be a result of women traveling out-of-state for abortions, rather than abortions actually falling. In order to test for effects on this

[26] The reason I use these six states is because the Pennsylvania Department of Health reports the number of abortions occurring within Pennsylvania per age group and per gestational age for each of its six neighboring states (Delaware, Maryland, New Jersey, New York, Ohio, and West Virginia)—but not for any other states. CDC Abortion Surveillance data also show the number of women from other states obtaining abortions in Pennsylvania, and include more than just the six neighboring states. However, I am choosing to look that the PA Department of Health data in order to have the age of the women obtaining abortions, as well as the gestational age at the time of the abortion.

behavior, I rely on the CDC Abortion Surveillance Data, which is available from 2009-2015. These data do not provide information on the age of the woman obtaining the abortion, or on the gestational age at the time of the abortion, so figures can only show the total abortion rate for women living in Pennsylvania obtaining abortions out of state. Figure A7 shows the natural log of the abortion rate of women in Pennsylvania obtaining abortion in other states, grouping the neighboring states based on which of Pennsylvania's borders they share. Since Pittsburgh is the treated city and is near the West border of the state, I expect any changes resulting from the new regulations to appear in the West Border States group.[27] Abortion rates for Pennsylvania residents traveling to North, South, and East border states remain relatively constant. Abortion rates for Pennsylvania residents traveling to West border states was falling sharply before the clinic closures, then rose in the first two years after Pennsylvania clinics closed. This suggests that some Pittsburgh-area women are responding to the closures by traveling out-of-state when they otherwise may have obtained an abortion in Pittsburgh.

6 Conclusion and Discussion

While the popularity of supply-side restrictions on abortion access grow, it is important to understand the impacts of such policies. Previous work has shown that distance to the nearest abortion clinic matters. Pennsylvania's unique natural experiment, however, allows me to estimate the effects of *congestion* alone on abortion access, as Pennsylvania's clinic closures only minimally affected distance to the nearest clinic for most women in the state. My findings suggest that clinic congestion matters. The results offer suggestive evidence that abortion rates fall in counties that experience major increases in congestion, relative to counties that experience only a minor change in congestion. Additionally, the number of early-term ($\leq$8 weeks gestational age) abortions falls by approximately 13.8% per year in the counties experiencing major congestion changes, relative to the compari-

[27] Since I do not know the county of residence for women obtaining out-of-state abortions, the thought process here is that women are likely to travel to the nearest out-of-state clinic if they choose to leave the state. This means that West border states are treated, East are not, and the predicted effects for North and South border states is ambiguous.

son counties. Many of the abortions that would have occurred during this gestational age are likely still occurring, but at a later time in the pregnancy. These changes in abortion timing are important for three reasons: first, the choice set for abortion procedures falls as gestational age increases. In the state of Pennsylvania, women can obtain a medical abortion through their tenth week of gestation. After that point, they are only able to access surgical abortion, which is a much more invasive procedure. Second, abortion services get more expensive as gestational age increases. In July of 2019, Planned Parenthood of Western Pennsylvania listed prices for abortion services by gestational age. For a surgical abortion with local anesthetic (the cheapest surgical abortion option), the cost of an abortion was $435 up through week 11 of gestation, then jumped to $540 in weeks 12–13, $815 in weeks 14–16, and $915 in weeks 17–18. This particular clinic also does not offer abortion services after week 18 of gestation, though the state legally allows abortions through week 24. This increase in the cost of abortion is particularly concerning, since nearly half of all women obtaining abortions in the United States have an income below the federal poverty level, and this cost increase is not considering any other potential costs a woman may incur due to delaying her abortion (Jones and Jerman (2017)). Third, while abortion is overall a very safe procedure, the risk of dangerous complications grows as gestation goes on. Typically, abortions are safest early in the pregnancy, and grow less safe as the pregnancy goes on. Figure 18 shows Pennsylvania's average state-wide complication rate by gestational age at the time of abortion and type of complication, for the years of 2006-2011. Retained products of conception refers to a complication in which the abortion was unsuccessful and a second 'abortion' must take place. This complication is most common with medical abortions, so seeing this type of complication rate fall as gestational age increases (as the medical abortion is no longer an option) is unsurprising. However, the risk of complications like infection or bleeding increases with gestational age. While women obtaining abortions at any gestational age are quite unlikely to experience any complications, the increase in the risk of these dangerous complications is a concern.

This line of research is especially relevant given the growing popularity of regulations of this nature. While the stated aim of these regulations—and similar ones in other states—is to improve the safety of abortions obtained, results show that they actually force the closure of many existing clinics. Clinic closures may increase distance to the nearest clinic, which has been documented to be important to health outcomes, and may increase clinic congestion, which I show also has significant effects on access to services obtained at the clinics that remain open. Evidence on the effects of clinic congestion is relevant to discussions about healthcare access: being geographically near an open clinic is only part of the issue.

References

Abadie, A., A. Diamond, and J. Hainmueller (2010). Synthetic control methods for comparative case studies: Estimating the effect of californias tobacco control program. *Journal of the American statistical Association 105*(490), 493–505.

Allison, P. D. and R. P. Waterman (2002). Fixed–effects negative binomial regression models. *Sociological methodology 32*(1), 247–265.

Anderson, M., C. Dobkin, and T. Gross (2012). The effect of health insurance coverage on the use of medical services. *American Economic Journal: Economic Policy 4*(1), 1–27.

Bailey, M. J. (2006). More power to the pill: The impact of contraceptive freedom on women's life cycle labor supply. *The Quarterly Journal of Economics 121*(1), 289–320.

Bailey, M. J., B. Hershbein, and A. R. Miller (2012, July). The opt-in revolution? contraception and the gender gap in wages. *American Economic Journal: Applied Economics 4*(3), 225–54.

Bailey, M. J., O. Malkova, and Z. M. McLaren (2018). Does access to family planning increase childrens opportunities? evidence from the war on poverty and the early years of title x. *Journal of Human Resources*, 1216–8401R1.

Cameron, A. C. and P. K. Trivedi (2005). *Microeconometrics: methods and applications*. Cambridge university press.

Colman, S. and T. Joyce (2011). Regulating abortion: impact on patients and providers in texas. *Journal of Policy Analysis and Management 30*(4), 775–797.

Countouris, M., S. Gilmore, and M. Yonas (2014). Exploring the impact of a community hospital closure on older adults: A focus group study. *Health & place 26*, 143–148.

Courtemanche, C., J. Marton, B. Ukert, A. Yelowitz, and D. Zapata (2017). Early impacts of the affordable care act on health insurance coverage in medicaid expansion and non-expansion states. *Journal of Policy Analysis and Management 36*(1), 178–210.

Domonoske, C. (2016, Jun). Supreme court strikes down abortion restrictions in texas.

Doyle, J. J. (2011, July). Returns to local-area health care spending: Evidence from health shocks to patients far from home. *American Economic Journal: Applied Economics 3*(3), 221–43.

Evert, K. (2019, 07). Personal communication.

Finkelstein, A. (2007). The aggregate effects of health insurance: Evidence from the introduction of medicare. *The quarterly journal of economics 122*(1), 1–37.

Finkelstein, A., S. Taubman, B. Wright, M. Bernstein, J. Gruber, J. P. Newhouse, H. Allen, K. Baicker, and O. H. S. Group (2012). The oregon health insurance experiment: evidence from the first year. *The Quarterly journal of economics 127*(3), 1057–1106.

Fischer, S., H. Royer, and C. White (2017). The impacts of reduced access to abortion and family planning services: Evidence from texas.

Goldin, C. and L. F. Katz (2002). The power of the pill: Oral contraceptives and women's career and marriage decisions. *Journal of Political Economy 110*(4), 730–770.

Guttmacher Institute, . (2019). Targeted regulation of abortion providers. `https://www.guttmacher.org/state-policy/explore/targeted-regulation-abortion-providers`. [Online, accessed 15-April-2019].

Jones, R. K. and J. Jerman (2017). Population group abortion rates and lifetime incidence of abortion: United states, 2008–2014. *American Journal of Public Health 107*(12), 1904–1909.

Kolstad, J. T. and A. E. Kowalski (2012). The impact of health care reform on hospital and preventive care: evidence from massachusetts. *Journal of Public Economics 96*(11-12), 909–929.

Lindo, J. M., C. Myers, A. Schlosser, and S. Cunningham (2017). How far is too far? new evidence on abortion clinic closures, access, and abortions. Technical report, National Bureau of Economic Research.

Martin, A. B., A. M. Sisko, S. A. Glied, J. M. Lambrew, Z. Cooper, G. Claxton, D. M. Cutler, and G. F. Anderson (2018, Dec). National health care spending in 2017: Growth slows to postgreat recession rates; share of gdp stabilizes.

Myers, C. K. (2017). The power of abortion policy: Re-examining the effects of young women's access to reproductive control. *Journal of Political Economy 125*(6), 2178–2224.

National Academies of Sciences, E. and Medicine (2018). *The Safety and Quality of Abortion Care in the United States*. Washington, DC: The National Academies Press.

Quast, T., F. Gonzalez, and R. Ziemba (2017). Abortion facility closings and abortion rates in texas. *Inquiry 54*, 1–7.

Roberts, S. C., U. D. Upadhyay, G. Liu, J. L. Kerns, D. Ba, N. Beam, and D. L. Leslie (2018). Association of facility type with procedural-related morbidities and adverse events among patients undergoing induced abortions. *Jama 319*(24), 2497–2506.

Sajadi-Ernazarova, K. R. and C. L. Martinez (2019, Jul). Abortion complications.

Tavernise, S. (2019, Mar). Why birthrates among hispanic americans have plummeted.

Venator, J. and J. Fletcher (2019). Undue burden beyond texas: An analysis of abortion clinic closures, births, and abortions in wisconsin. *NBER Working Paper No. 26362*.

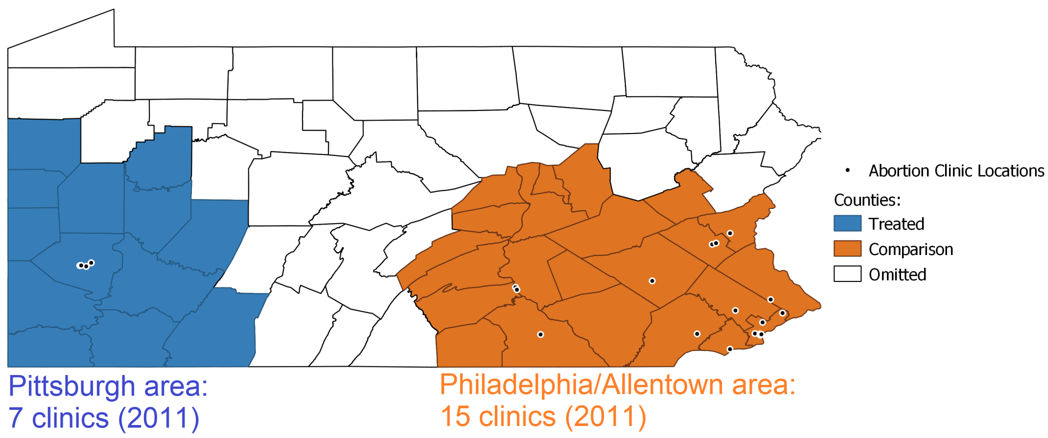
Weber, S. and M. Peclat (2017). A simple command to calculate travel distance and travel time. *Stata Journal 17*(4), 962–971.

Wee, R. Y. (2016, Mar). States with the largest latino and hispanic populations.

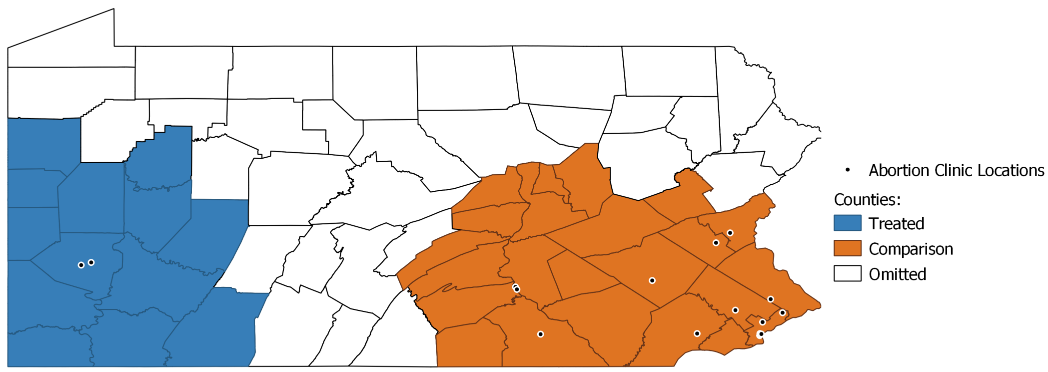
Figures

Figure 1
Abortion Clinic Locations

2011



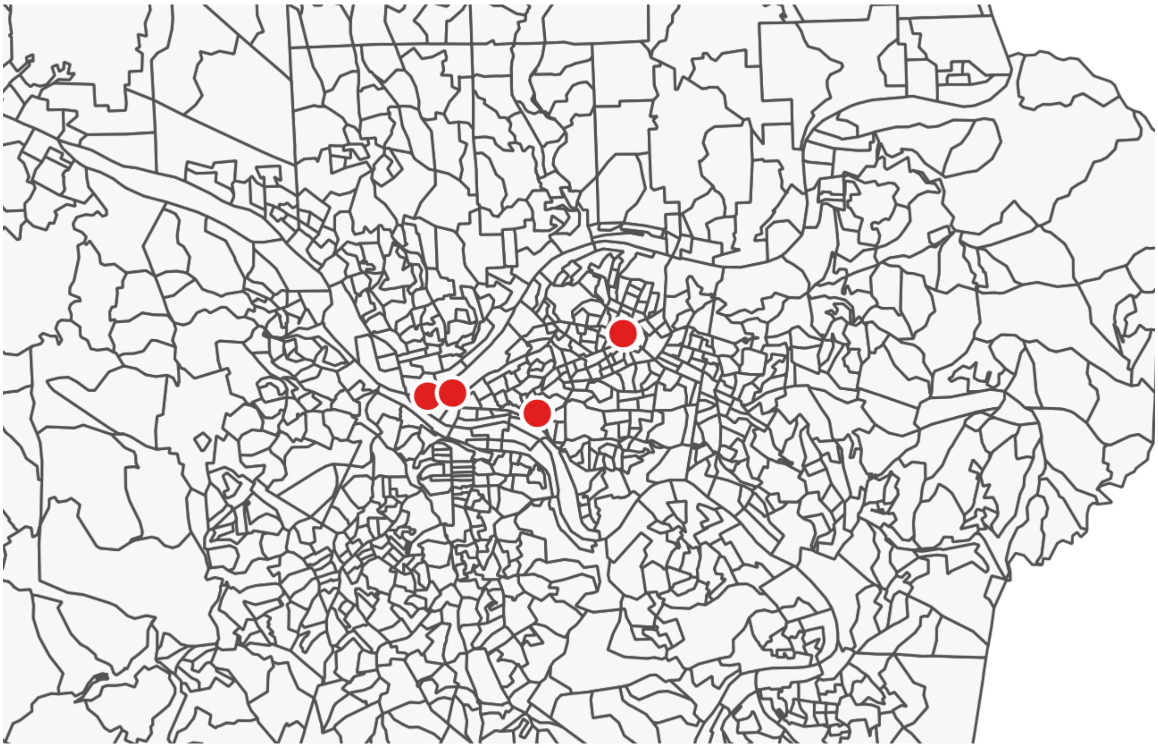
2013



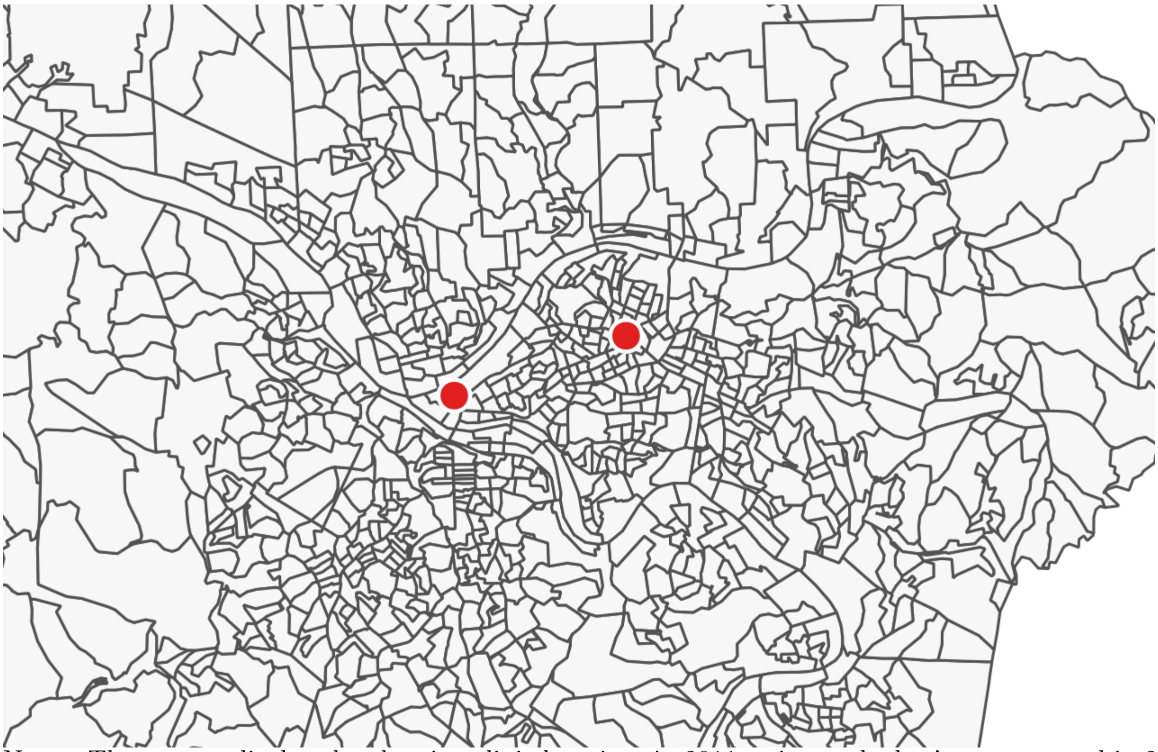
Notes: These maps display the abortion clinic locations in 2011, prior to the law's passage, and in 2013, after the law had taken effect and clinics had closed. Counties shaded in blue (on the west side of the state) are the treated counties, while counties shaded in burnt orange (on the east side of the state) are the comparison counties. Counties in white are omitted from the main analysis, as the closest clinic in 2006 (the first year of clinic location data) closed prior to the law change.

Figure 2
Abortion Clinic Locations - Pittsburgh

2011

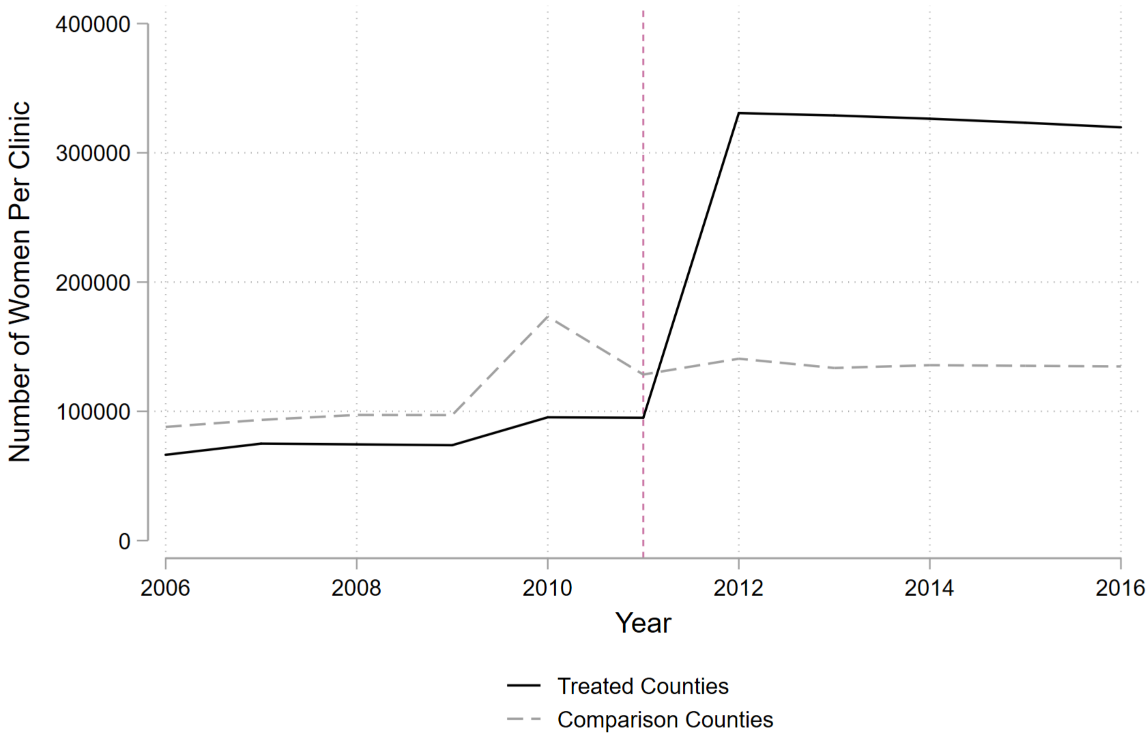


2013



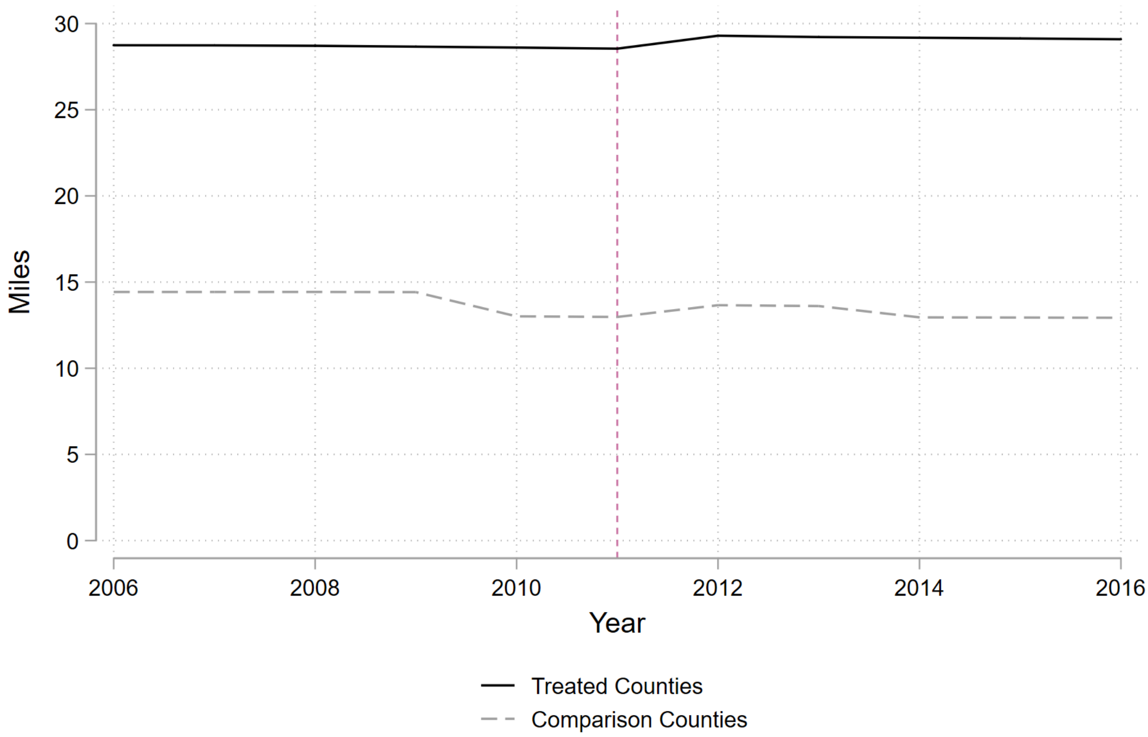
Notes: These maps display the abortion clinic locations in 2011, prior to the law's passage, and in 2013, after the law had taken effect and clinics had closed. Counties shaded in blue (on the west side of the state) are the treated counties, while counties shaded in burnt orange (on the east side of the state) are the comparison counties. Counties in white are omitted from the main analysis, as the closest clinic in 2006 (the first year of clinic location data) closed prior to the law change.

Figure 3
Service Populations Over Time



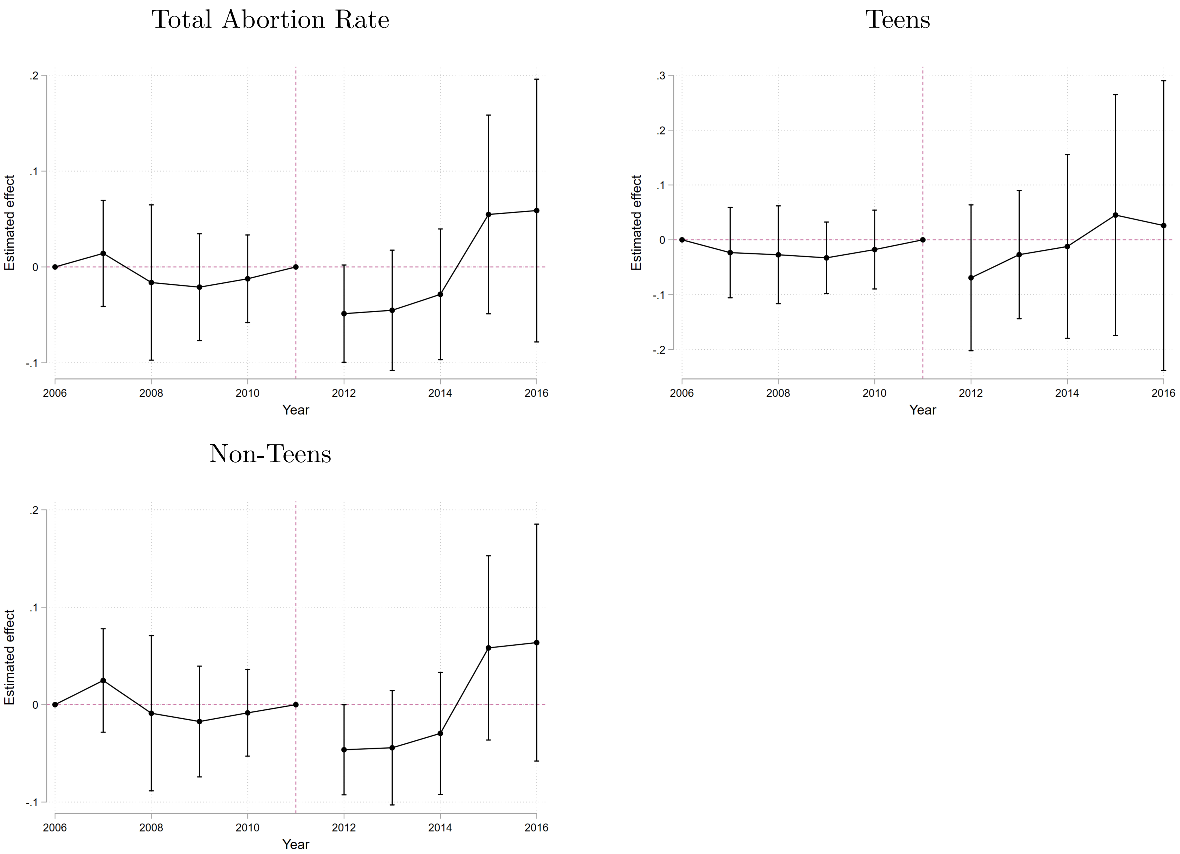
Notes: This figure shows the average service population (number of childbearing aged women divided by number of open abortion clinics) for treated and comparison counties over time. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006.

Figure 4
Distances over time



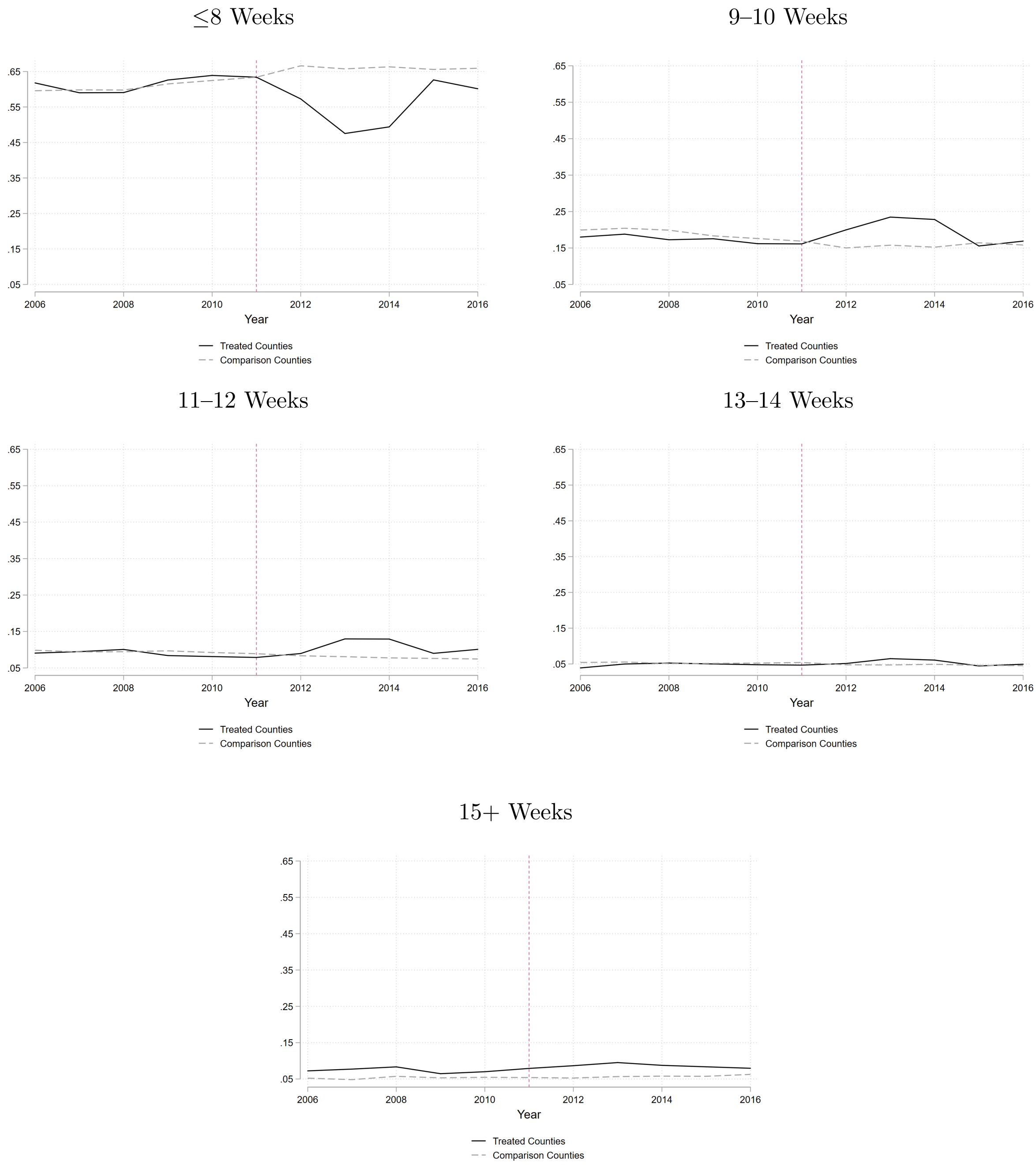
Notes: This figure plots the average distance from the county population centroid to the nearest abortion-providing facility over time, for treated and comparison counties. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006.

Figure 5
Effects for Various Age Groups



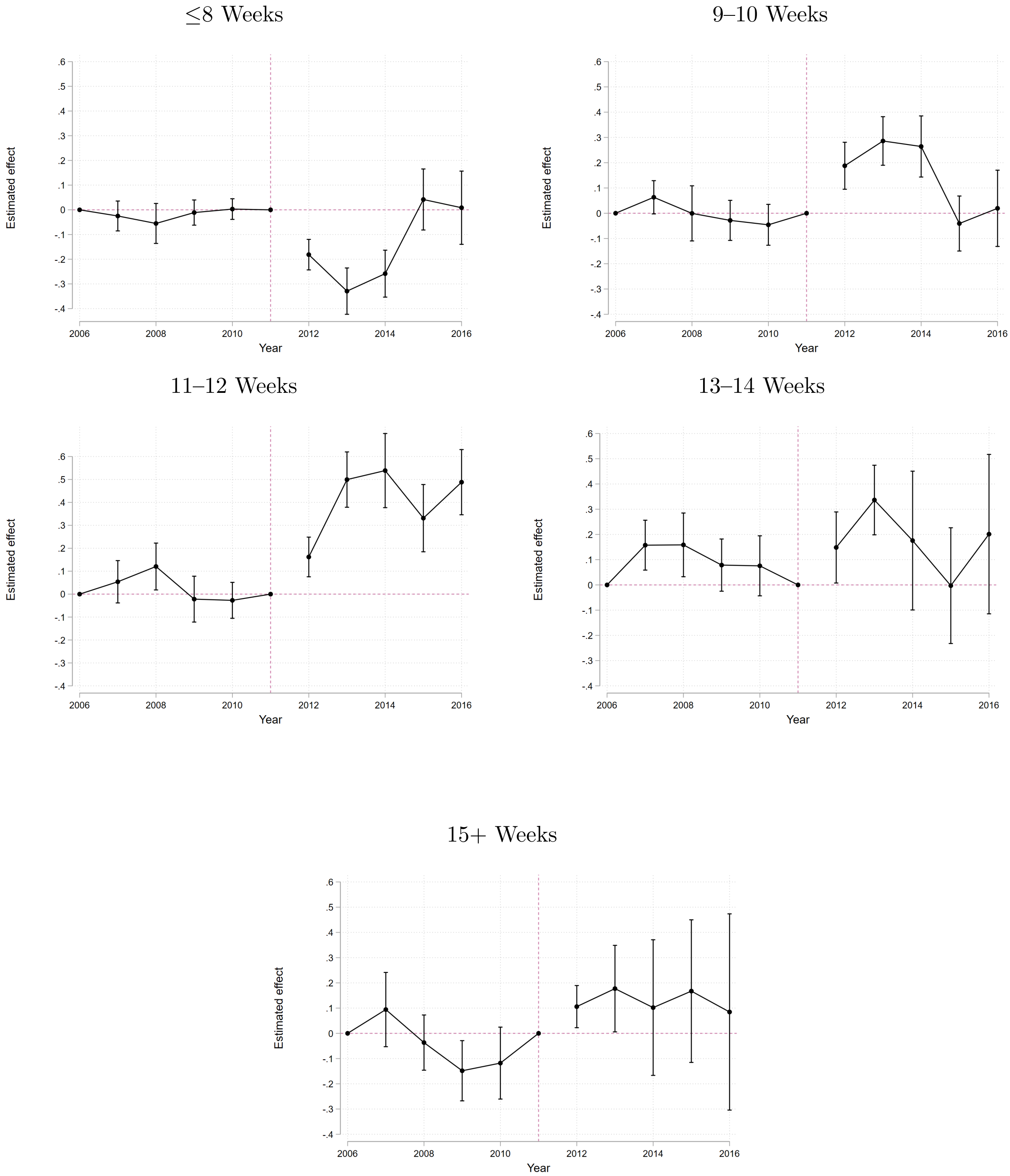
Notes: This figure plots the estimated effect of the passage of the law on abortion rates for various age groups. Estimates come from a Poisson model which controls for county and year fixed effects and county-specific linear time trends. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006.

Figure 6
Share of Abortions Occurring within the First 8 Weeks of Gestation



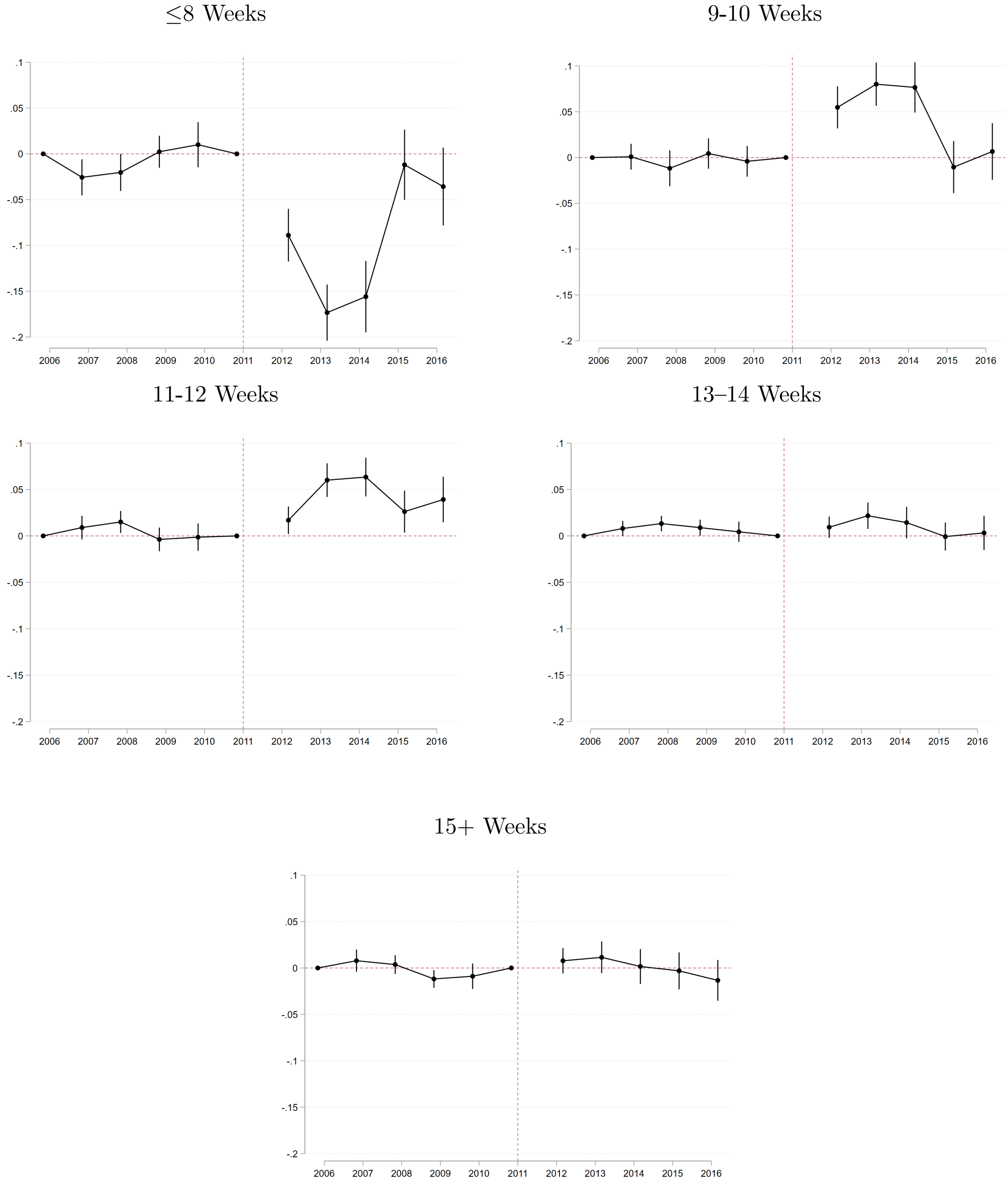
Notes: This figure plots the average percentage of abortions occurring within various gestational ages, for treated and comparison counties. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006.

Figure 7
Effects for Various Gestational Ages



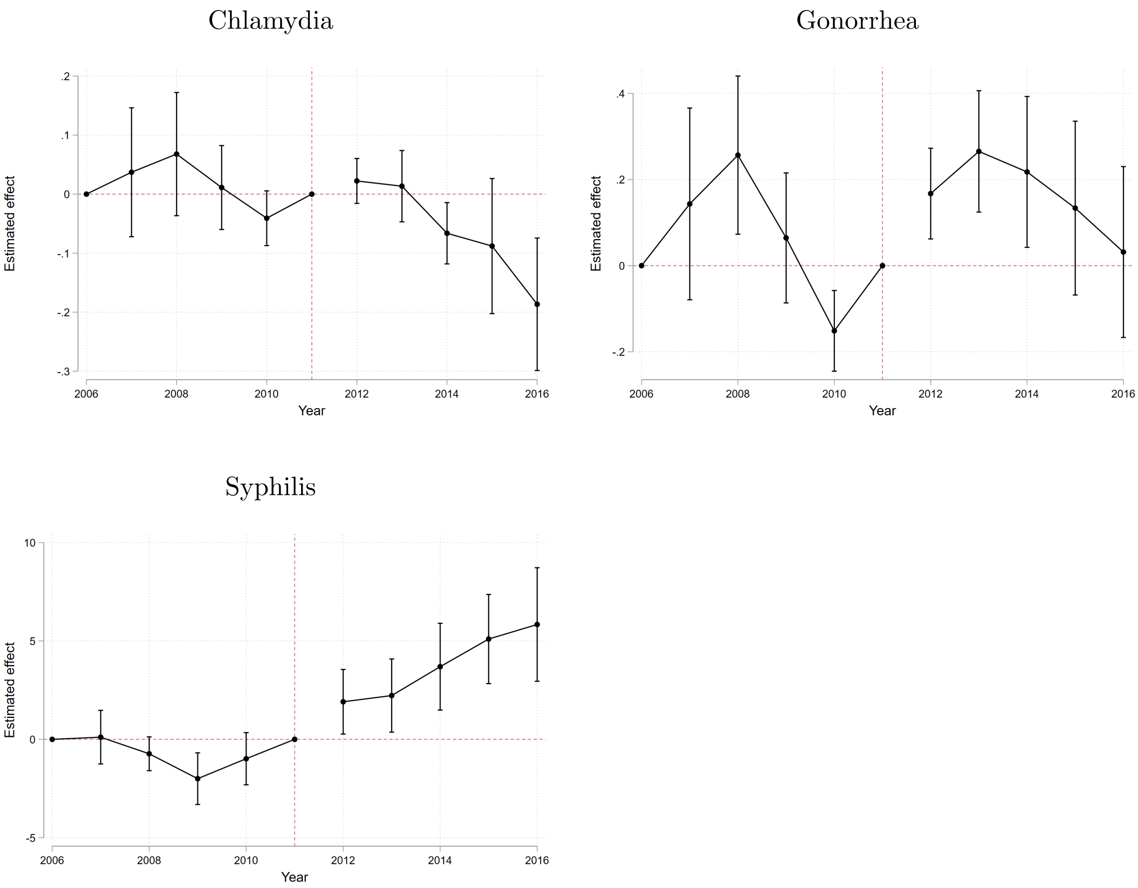
Notes: This figure plots the estimated effect of the passage of the law on abortion rates for various gestational ages. Estimates come from a Poisson model which controls for county and year fixed effects and county-specific linear time trends. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006.

Figure 8
Effects for Various Gestational Ages (Pct of Abortions Occurring at a Given Gestational Age)



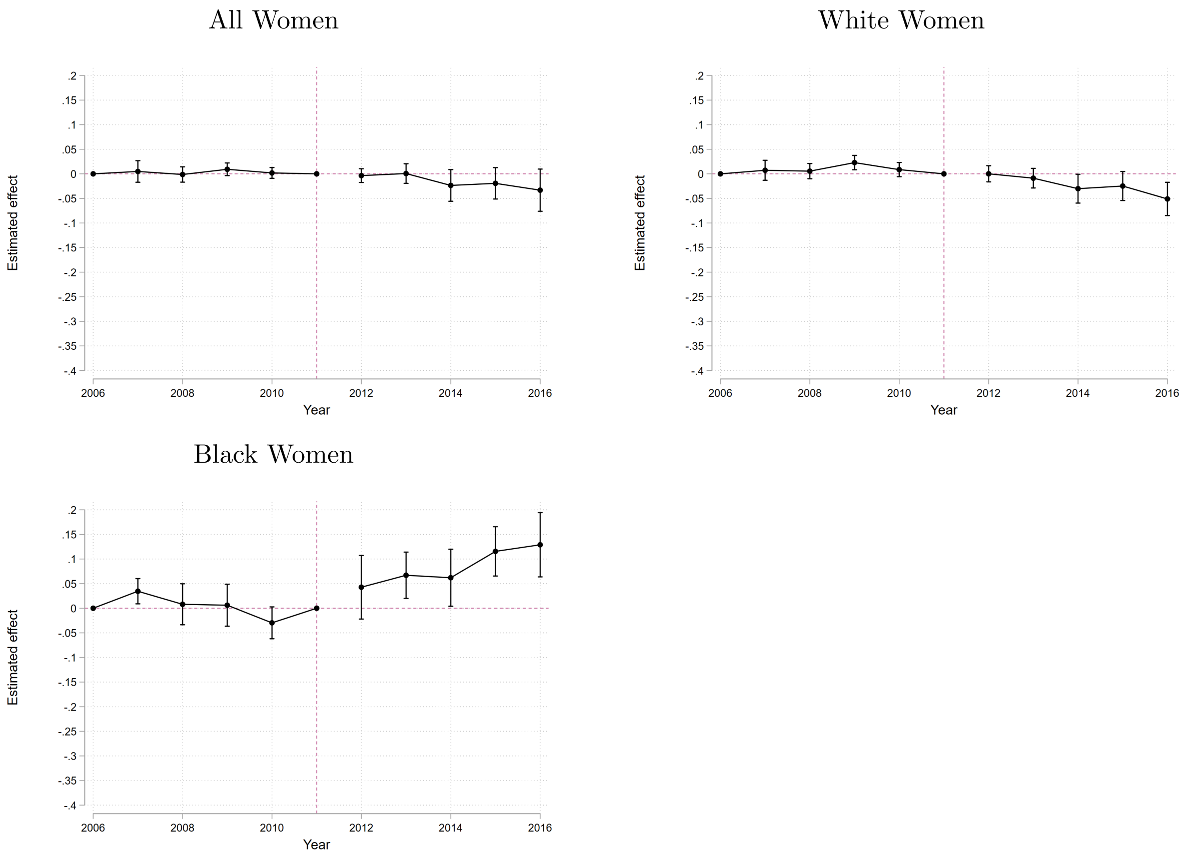
Notes: This figure plots the estimated effect of the passage of the law on the share of abortions occurring within a given gestational age. Estimates come from a weighted least squares model which controls for county and year fixed effects. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006.

Figure 9
Sexually Transmitted Infection Rates



Notes:

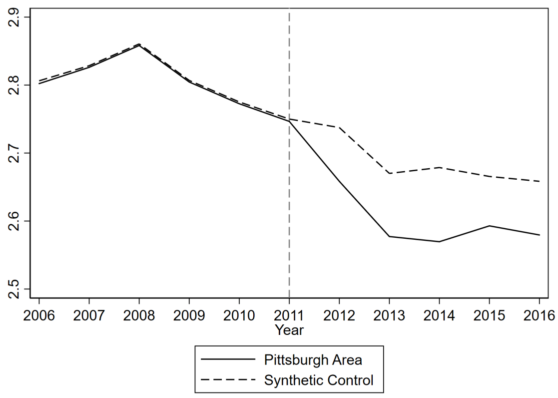
Figure 10
Effects on Births by Race of Mother



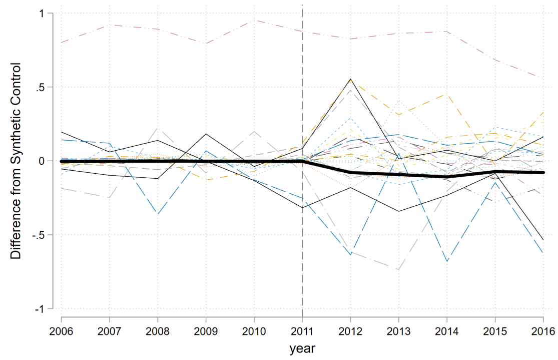
Notes: This figure plots the estimated effect of the passage of the law on births to mothers of various races. Estimates come from a Poisson model which controls for county and year fixed effects and county-specific linear time trends. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006.

Figure 11
Synthetic Control

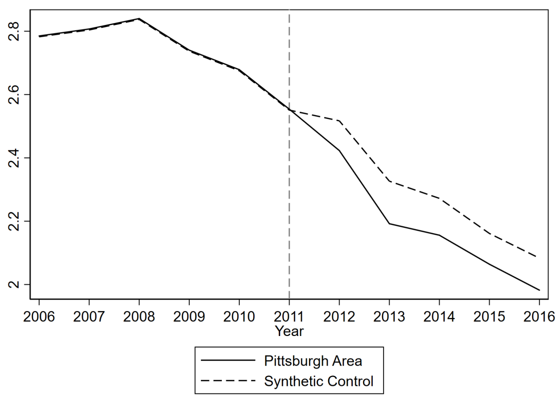
Log Total Abortion Rate



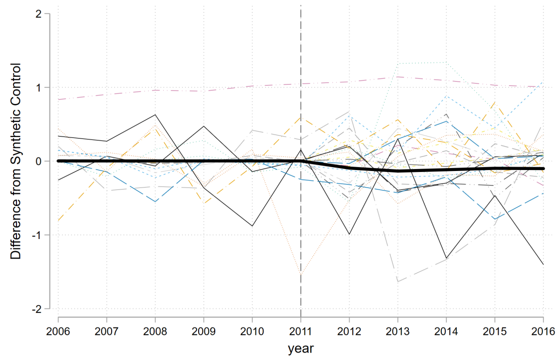
Inference: Log Total Abortion Rate



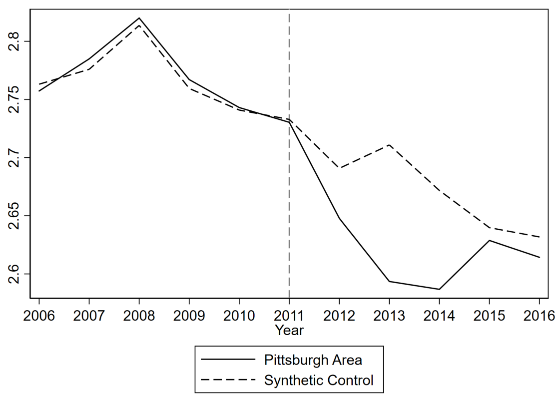
Log Abortion Rate, Teens



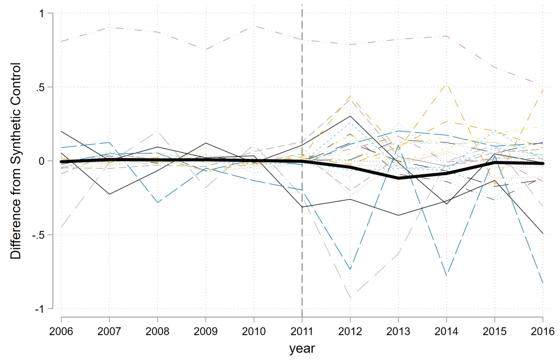
Inference: Log Abortion Rate, Teens



Log Abortion Rate, Non-Teens



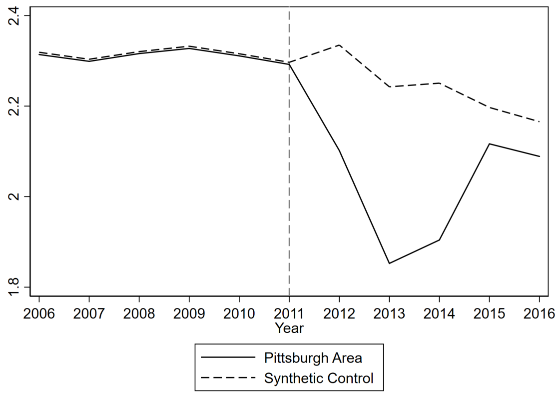
Inference: Log Abortion Rate, Non-Teens



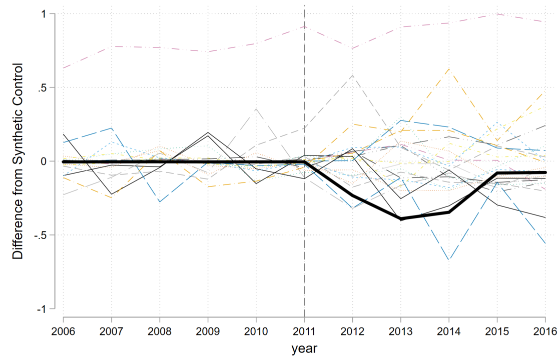
Notes:

Figure 12
Synthetic Control

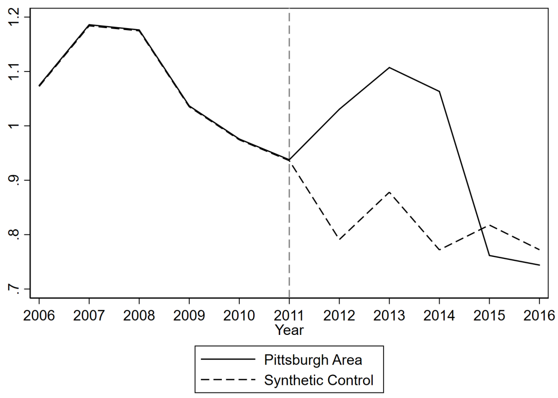
Log Abortion Rate, ≤8 Weeks



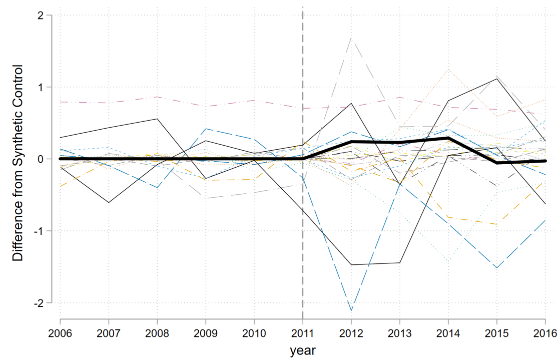
Inference: Log Abortion Rate ≤8 Weeks



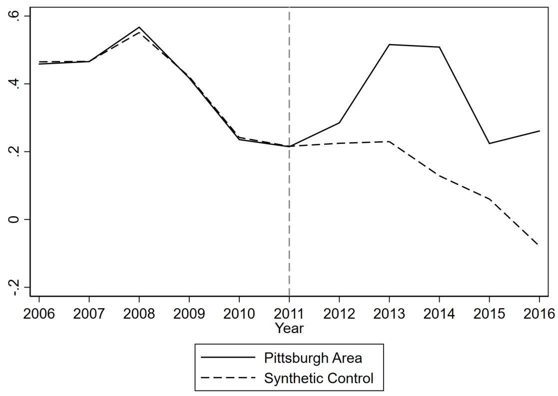
Log Abortion Rate, 9–10 Weeks



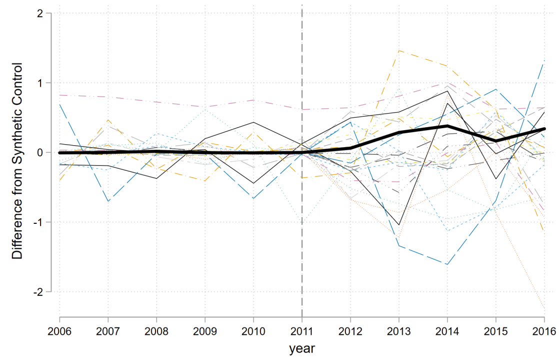
Inference: Log Abortion Rate 9–10 Weeks



Log Abortion Rate, 11–12 Weeks



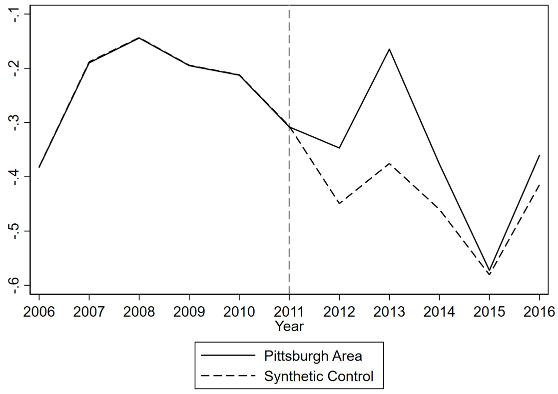
Inference: Log Abortion Rate 11–12 Weeks



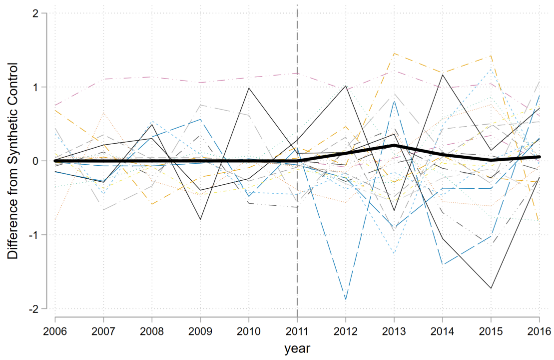
Notes:

Figure 13
Synthetic Control

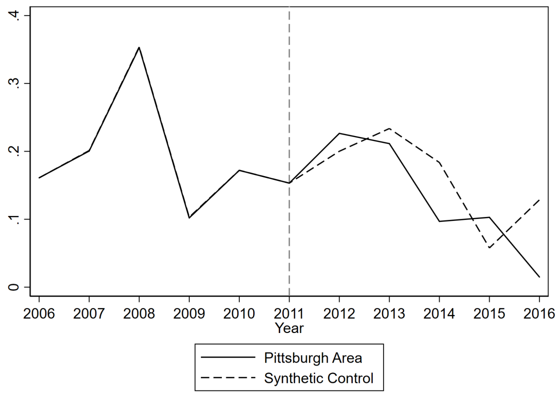
Log Abortion Rate, 13–14 Weeks



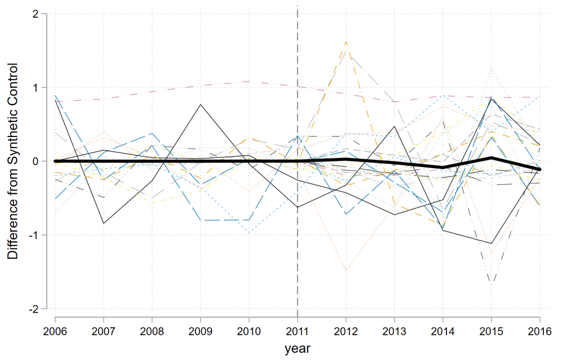
Inference: Log Abortion Rate 13–14 Weeks



Log Abortion Rate, 15+ Weeks

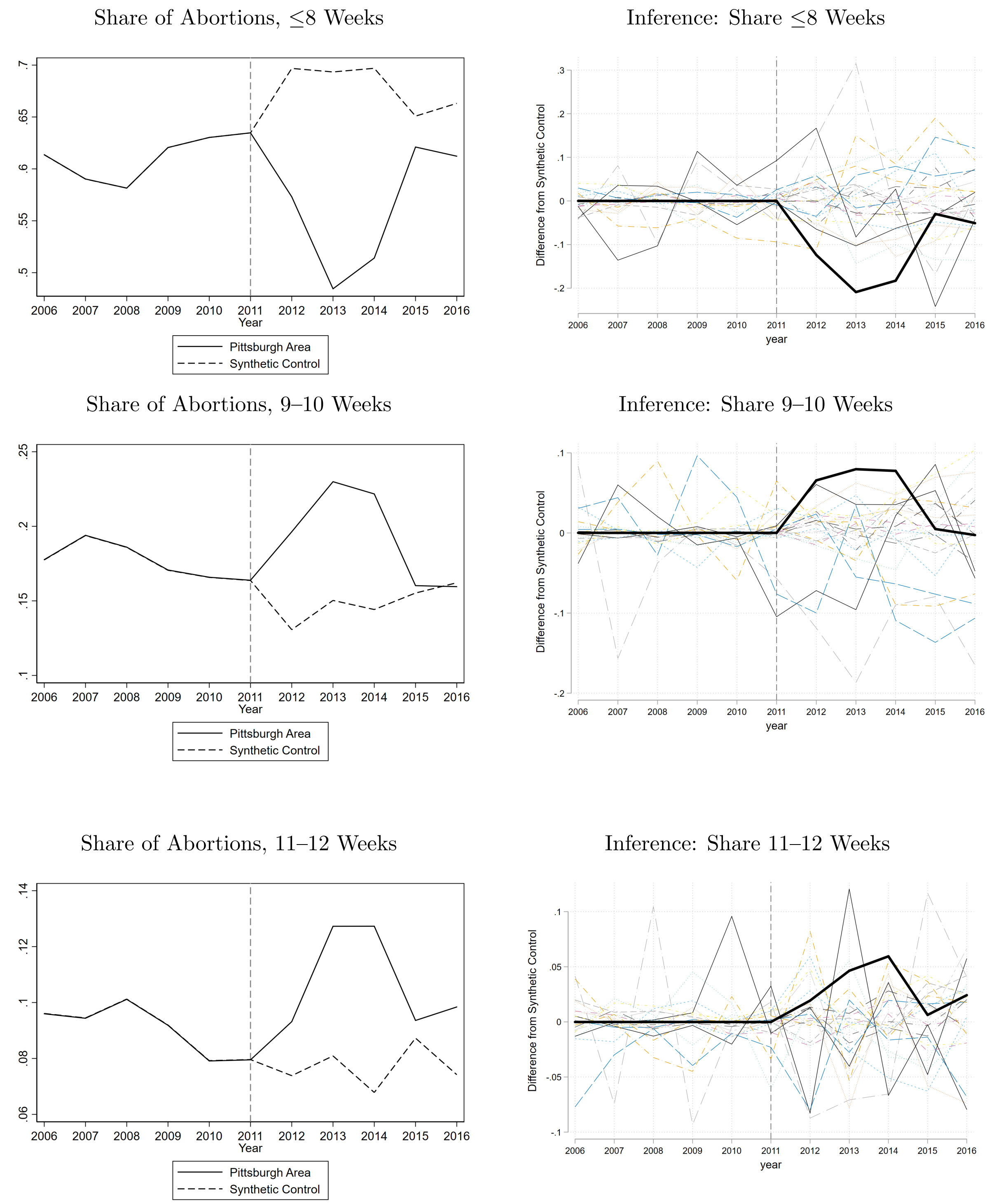


Inference: Log Abortion Rate 15+ Weeks



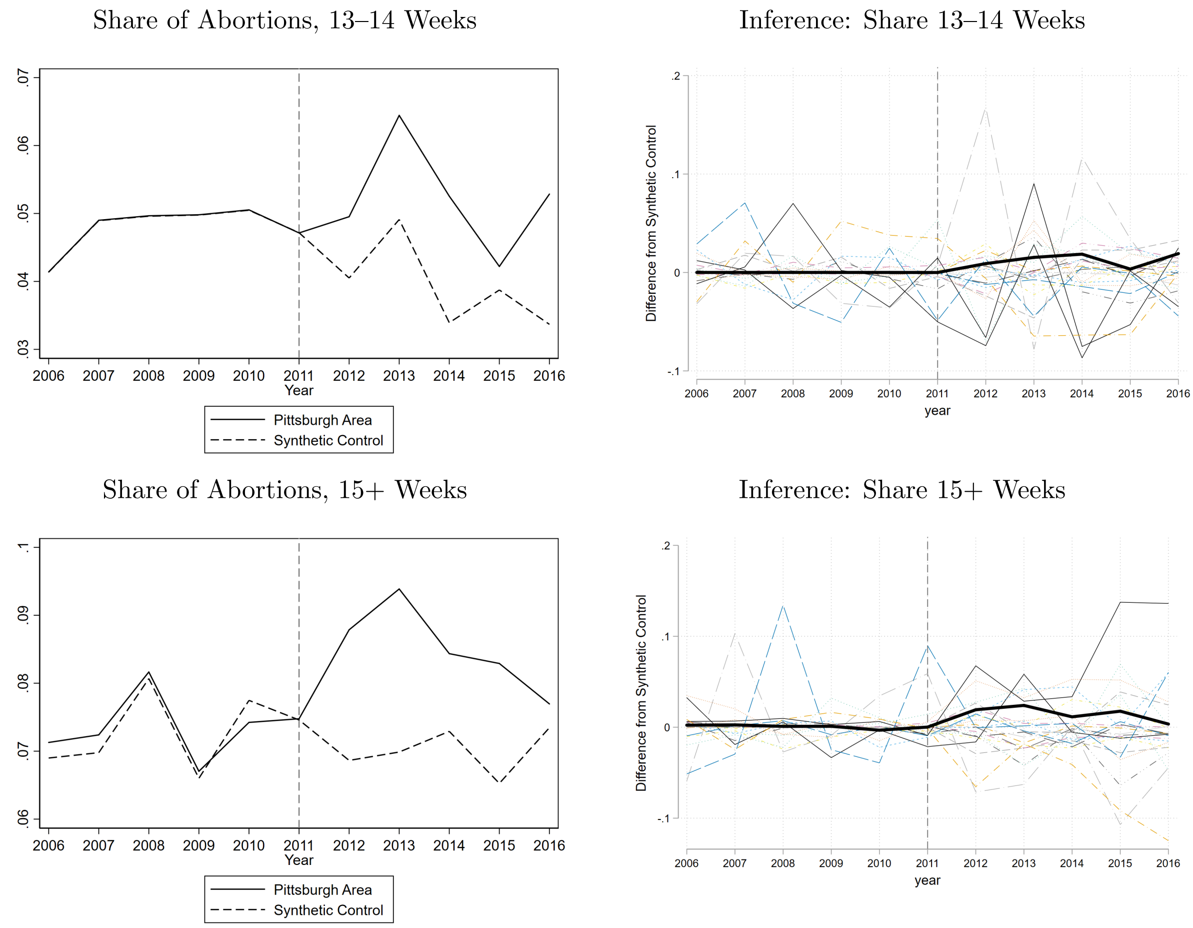
Notes:

Figure 14
Synthetic Control



Notes:

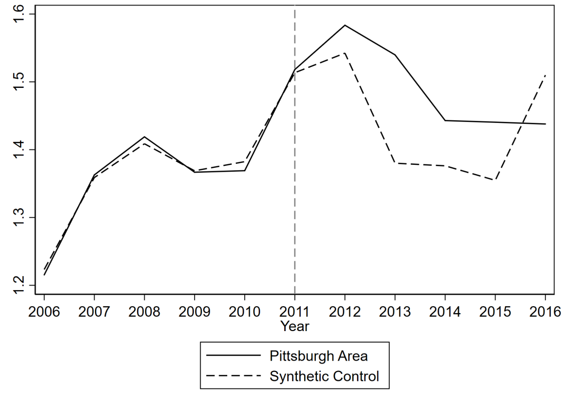
Figure 15
Synthetic Control



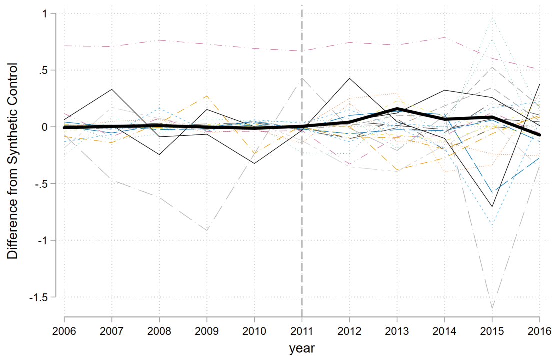
Notes:

Figure 16
Synthetic Control

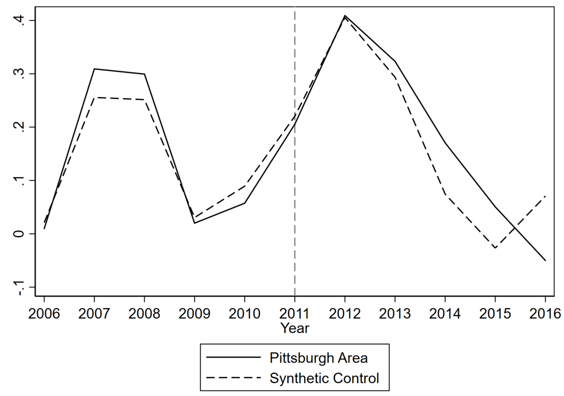
Log Chlamydia Rate



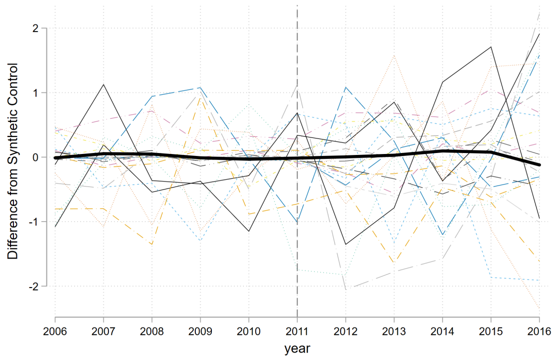
Inference: Log Chlamydia Rate



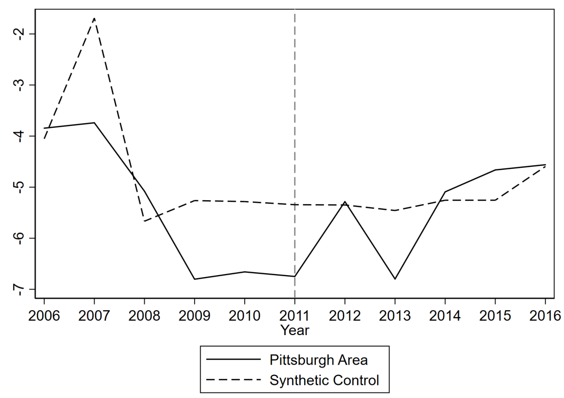
Log Gonorrhea Rate



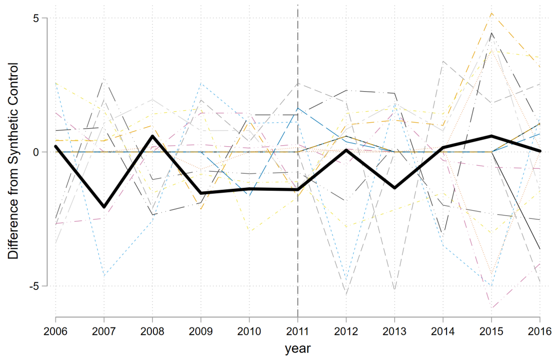
Inference: Log Gonorrhea Rate



Log Syphilis Rate



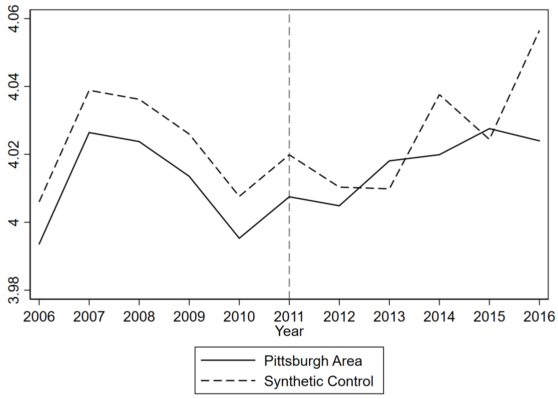
Inference: Log Syphilis Rate



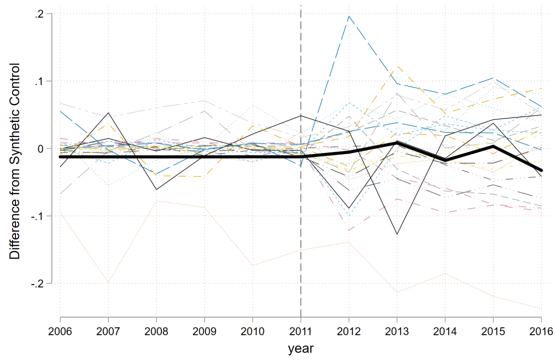
Notes:

Figure 17
Synthetic Control

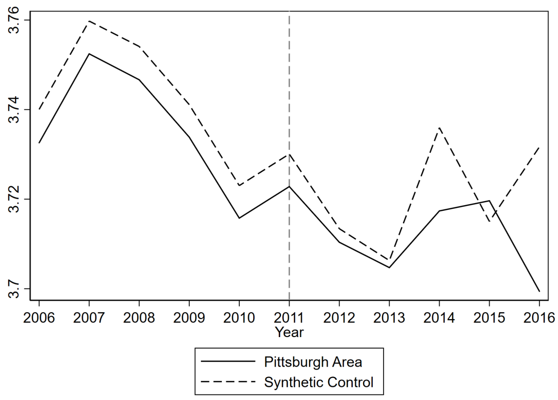
Log Birth Rate, Total Births



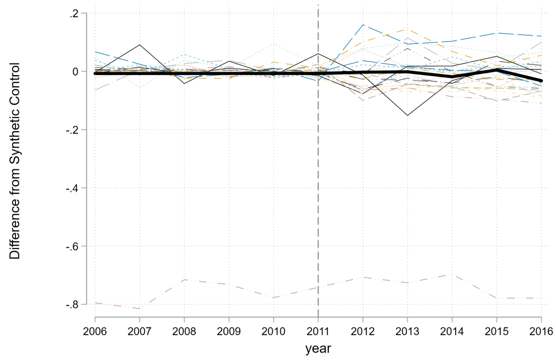
Inference: Log Birth Rate, Total Births



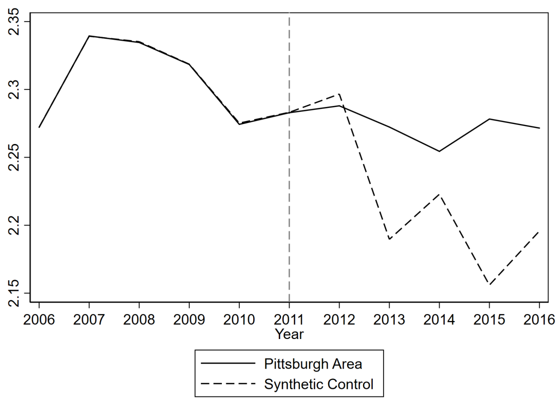
Log Birth Rate, White Births



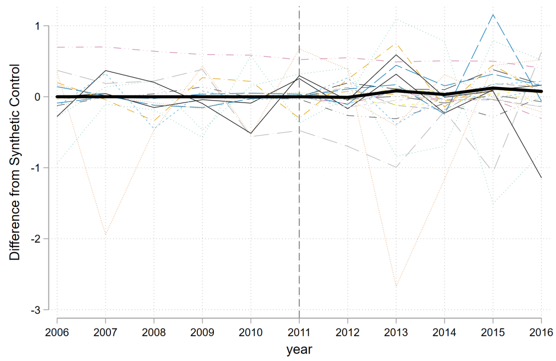
Inference: Log Birth Rate, White Births



Log Birth Rate, Black Births

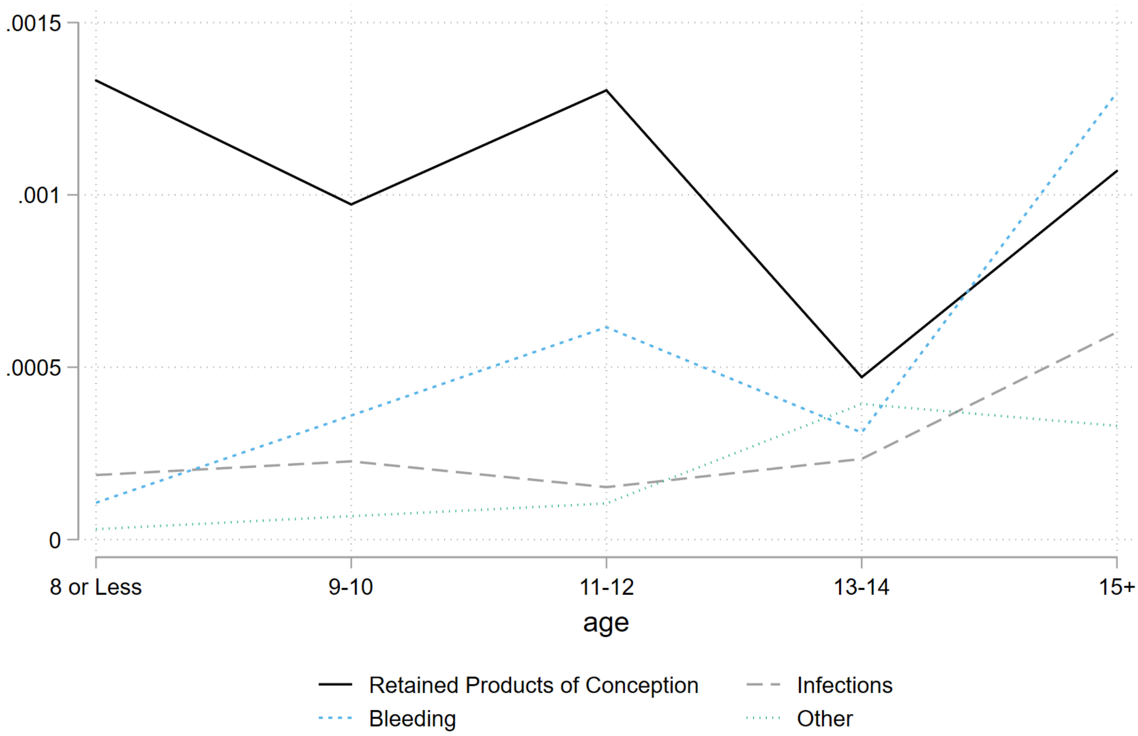


Inference: Log Birth Rate, Black Births



Notes:

Figure 18
Complication Rates by Gestational Age at Abortion



Notes: This figure plots the complication rate by gestational age and type of complication. Data come from the Pennsylvania Department of Health Annual Abortion files from 2006-2011.

Tables

Table 1
Summary Statitstics, Treated vs. Control Counties

	Treated	Comparison
Pre Period (Prior to 2012)		
Abortion Rate per 1000 Women aged 15-44	12.4	18.0
Abortion Rate per 1000 Women aged 15-19	11.3	16.3
Pct of Population that is aged 15-44	37.3	40.0
Pct White	89.2	73.8
Pct Hispanic	1.2	7.6
Pct Black	7.9	14.6
Poverty Rate	12.1	11.7
Unemployment Rate	6.9	6.3
Post Period (2012-2016)		
Abortion Rate per 1000 Women aged 15-44	10.3	16.2
Abortion Rate per 1000 Women aged 15-19	6.6	10.1
Pct of Population that is aged 15-44	36.7	38.9
Pct White	88.1	71.2
Pct Hispanic	1.5	9.2
Pct Black	8.1	14.9
Poverty Rate	12.2	12.8
Unemployment Rate	6.8	6.0

Notes: This table shows summary statistics for treated and comparison counties, before and after the law's passage. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006.

Table 2
Estimated Effects of Abortion Regulations on Abortion Rates by Age

	Total	Teens	Non-Teens
First Year of the Law	-0.034	-0.051	-0.034
	(0.036)	(0.076)	(0.033)
Second Year of the Law	-0.028	-0.008	-0.029
	(0.043)	(0.073)	(0.039)
Third Year of the Law	-0.009	0.007	-0.011
	(0.046)	(0.092)	(0.042)
Fourth Year of the Law	0.077	0.064	0.080
	(0.064)	(0.125)	(0.058)
Fifth Year of the Law	0.084	0.045	0.089
	(0.080)	(0.145)	(0.072)
Average effect	0.018	0.011	0.019
P-value (test average effect = 0)	0.722	0.905	0.669
Observations	396	396	396

Notes: Results come from a Poisson model with county and year fixed effects and county-specific linear time trends.
***, **, and * represent p-values less than 0.01, 0.05, and 0.10, respectively.

Table 3
Estimated Effects of Abortion Regulations on Abortion Rates by Gestational Age

	≤8 Weeks	9–10 Weeks	11–12 Weeks	13–14 Weeks	15+ Weeks
First Year of the Law	-0.179***	0.227***	0.176***	0.103	0.215***
	(0.039)	(0.040)	(0.042)	(0.081)	(0.052)
Second Year of the Law	-0.331***	0.336***	0.524***	0.302***	0.308***
	(0.055)	(0.040)	(0.049)	(0.079)	(0.101)
Third Year of the Law	-0.264***	0.324***	0.574***	0.151	0.254*
	(0.055)	(0.050)	(0.072)	(0.152)	(0.146)
Fourth Year of the Law	0.033	0.030	0.377***	-0.017	0.341**
	(0.067)	(0.051)	(0.061)	(0.131)	(0.155)
Fifth Year of the Law	-0.004	0.101	0.545***	0.198	0.280
	(0.080)	(0.078)	(0.069)	(0.178)	(0.204)
Average effect	-0.149	0.204	0.439	0.147	0.280
P-value (test average effect = 0)	0.006	0.000	0.000	0.192	0.029
Observations	396	396	396	396	396

Notes: Results come from a Poisson model with county and year fixed effects and county-specific linear time trends.
***, **, and * represent p-values less than 0.01, 0.05, and 0.10, respectively.

Table 4
Estimated Effects of Abortion Regulations on the Share of Abortions Occurring at a Given Gestational Age

	≤8 Weeks	9–10 Weeks	11–12 Weeks	13–14 Weeks	15-+ Weeks
First Year of the Law	-0.096***	0.056***	0.019***	0.005	0.016***
	(0.013)	(0.011)	(0.006)	(0.005)	(0.006)
Second Year of the Law	-0.184***	0.082***	0.063***	0.018***	0.021***
	(0.014)	(0.010)	(0.008)	(0.007)	(0.007)
Third Year of the Law	-0.171***	0.078***	0.068***	0.011	0.013*
	(0.018)	(0.012)	(0.009)	(0.008)	(0.008)
Fourth Year of the Law	-0.030*	-0.009	0.032***	-0.004	0.011
	(0.017)	(0.012)	(0.010)	(0.007)	(0.008)
Fifth Year of the Law	-0.058***	0.008	0.047***	0.001	0.002
	(0.019)	(0.013)	(0.011)	(0.009)	(0.009)
Average effect	-0.108	0.043	0.046	0.006	0.013
P-value (test average effect = 0)	0.000	0.000	0.000	0.284	0.055
Observations	396	396	396	396	396

Notes: Results come from a weighted least squares model with county and year fixed effects. Controls include race by age demographics, unemployment rate, and the percent of the county living in poverty. ***, **, and * represent p-values less than 0.01, 0.05, and 0.10, respectively.

Table 5
Estimated Effects of Abortion Regulations on Birth Rates by Race of Mother

	Total	White	Black
First Year of the Law	-0.006	-0.009	0.059**
	(0.007)	(0.008)	(0.025)
Second Year of the Law	-0.002	-0.019	0.091***
	(0.012)	(0.012)	(0.028)
Third Year of the Law	-0.026	-0.041**	0.092***
	(0.018)	(0.017)	(0.031)
Fourth Year of the Law	-0.022	-0.036**	0.152***
	(0.019)	(0.016)	(0.023)
Fifth Year of the Law	-0.036	-0.063***	0.171***
	(0.026)	(0.020)	(0.034)
Average effect	-0.018	-0.033	0.113
P-value (test average effect = 0)	0.229	0.011	0.000
Observations	396	396	396

Notes: Results come from a Poisson model with county and year fixed effects and county-specific linear time trends.
***, **, and * represent p-values less than 0.01, 0.05, and 0.10, respectively.

Table 6
Synthetic Control: Estimated Effects of Abortion Regulations on Abortion Rates by Age

Total Abortion Rate

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	-.078	-.091	-.108	-.071	-.078	-.085
P-Value	.042	.042	.042	.042	.042	.042

Teen Abortion Rate

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	-.098	-.137	-.119	-.100	-.103	-.111
P-Value	.167	.167	.250	.167	.208	.250

Non-Teen Abortion Rate

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	-.054	-.117	-.101	-.035	-.024	-.065
P-Value	.250	.208	.250	.417	.500	.292

Notes:

Table 7
Synthetic Control: Estimated Effects of Abortion Regulations on STI Rates

Chlamydia Rate

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.040	.159	.066	.086	-.073	.056
P-Value	.292	.167	.250	.292	.250	.208

Gonorrhea Rate

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.006	.030	.095	.072	-.126	016
P-Value	.958	.833	.333	.547	.292	.458

Syphilis Rate

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	-.392	-1.74	-.347	.082	-.522	-.584
P-Value	.917	.625	.875	.958	.917	.917

Notes:

Table 8
Synthetic Control: Estimated Effects of Abortion Regulations on Abortion Rates by Gestational Age

≤8 Weeks

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	-.231	-.387	-.344	-.078	-.075	-.223
P-Value	.042	.042	.042	.083	.083	.042

9–10 Weeks

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.240	.229	.291	-.057	-.029	.135
P-Value	.083	.125	.083	.250	.208	.125

11–12 Weeks

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.060	.285	.378	.164	.342	.246
P-Value	.333	.167	.167	.167	.167	.167

13–14 Weeks

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.104	.212	.085	.010	.055	.093
P-Value	.083	.042	.083	.250	.167	.083

15+ Weeks

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.027	-.023	-.088	.046	-.114	-.030
P-Value	.250	.292	.083	.208	.125	.167

Notes:

Table 9
Synthetic Control: Estimated Effects of Abortion Regulations on Share of Abortions in a Given Gestational Age

≤8 Weeks

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	-.122	-.207	-.181	-.028	-.050	-.118
P-Value	.042	.042	.042	.042	.083	.042

9–10 Weeks

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.066	.080	.077	.057	-.003	.045
P-Value	.083	.083	.083	.333	.417	.083

11–12 Weeks

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.019	.046	.059	.006	.024	.031
P-Value	.042	.042	.042	.083	.042	.042

13–14 Weeks

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.009	.015	.019	.003	.019	.013
P-Value	.042	.042	.042	.125	.042	.042

15+ Weeks

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.019	.025	.012	.018	.005	.016
P-Value	.250	.292	.333	.417	.625	.375

Notes:

Table 10
Synthetic Control: Estimated Effects of Abortion Regulations on Birth Rates by Race

Total Birth Rate

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.011	.025	-.001	.020	-.016	.008
P-Value	.583	.208	.958	.375	.500	.542

White Birth Rate

	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	.000	.002	-.016	.008	-.029	-.007
P-Value	.958	.917	.500	.667	.333	.667

Black Birth Rate

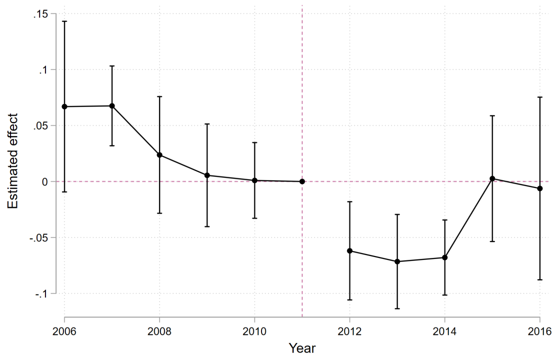
	Year 1	Year 2	Year 3	Year 4	Year 5	Avg Y1-5
Est. Effect	-.011	.078	.028	.119	.073	.057
P-Value	.458	.333	.375	.125	.333	.375

Notes:

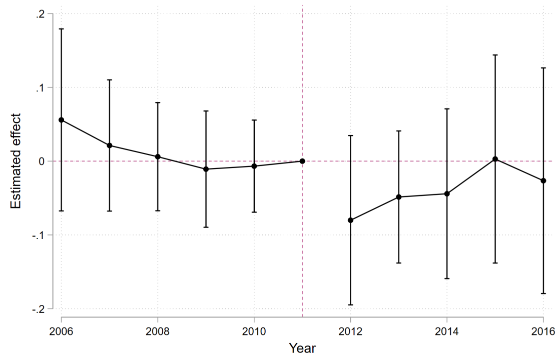
Appendix

Figure A1
Effects for Various Age Groups, Without Linear Time Trends

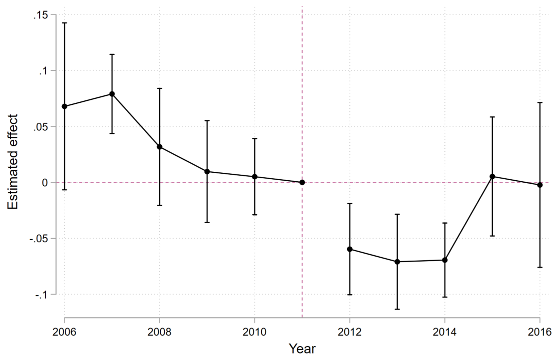
Total



Teens



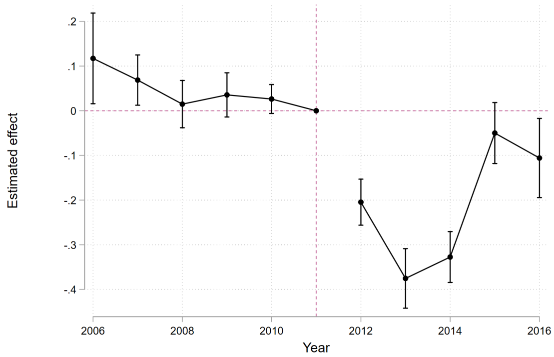
Non-Teens



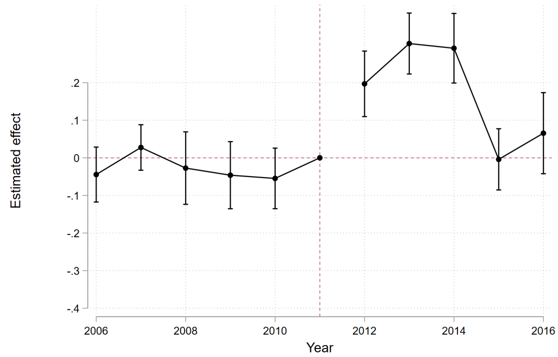
Notes: This figure plots the estimated effect of the passage of the law on abortion rates for various age groups. Estimates come from a Poisson model which controls for county and year fixed effects. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006. All other counties in the state are included in the comparison group.

Figure A2
Effects for Various Gestational Ages, Without Linear Time Trends

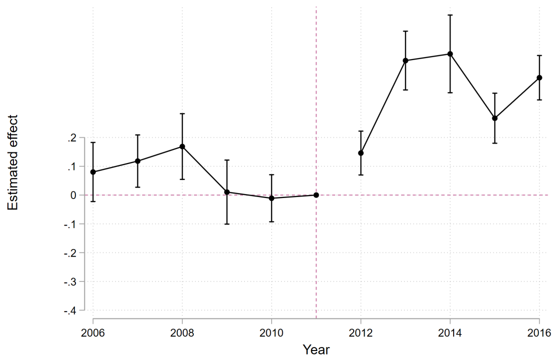
≤8 Weeks



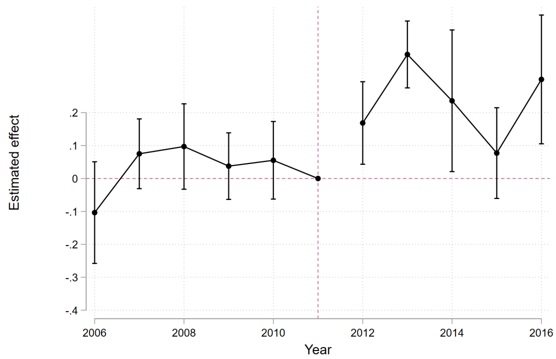
9–10 Weeks



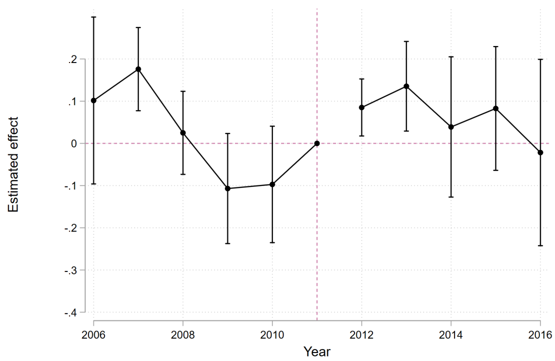
11–12 Weeks



13–14 Weeks



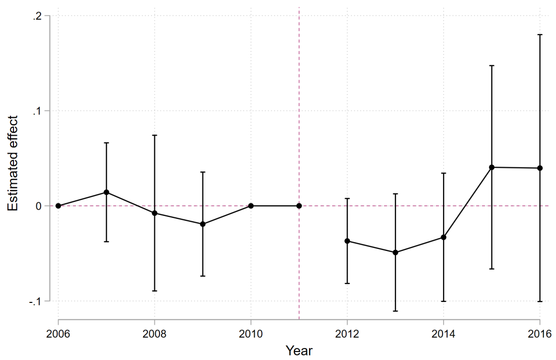
15+ Weeks



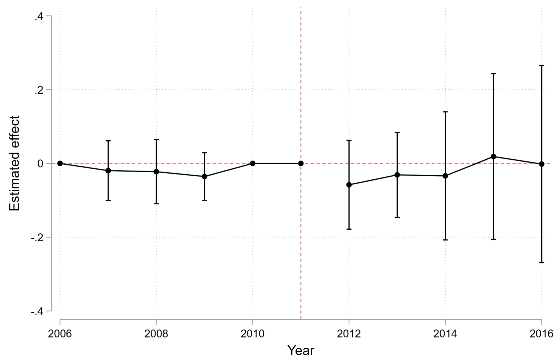
Notes: This figure plots the estimated effect of the passage of the law on abortion rates for various age groups. Estimates come from a Poisson model which controls for county and year fixed effects. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006. All other counties in the state are included in the comparison group.

Figure A3
Effects for Various Age Groups, Omitting 2010

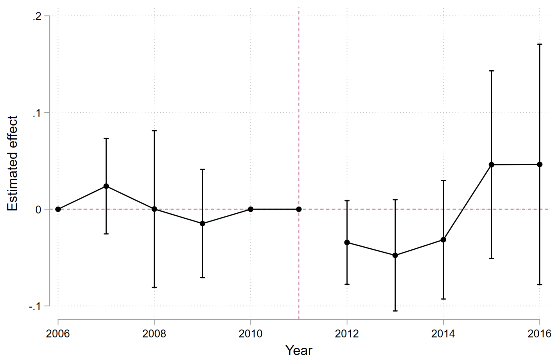
Total



Teens



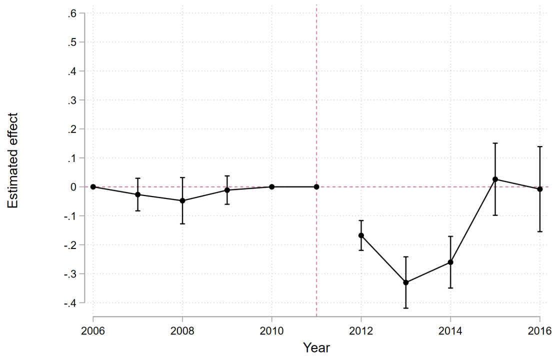
Non-Teens



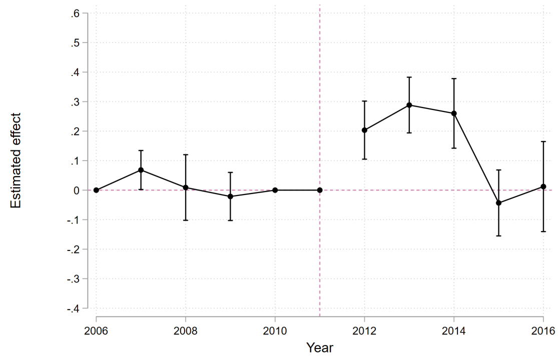
Notes: This figure plots the estimated effect of the passage of the law on abortion rates for various age groups. Estimates come from a Poisson model which controls for county and year fixed effects. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006. All other counties in the state are included in the comparison group.

Figure A4
Effects for Various Gestational Ages, Omitting 2010

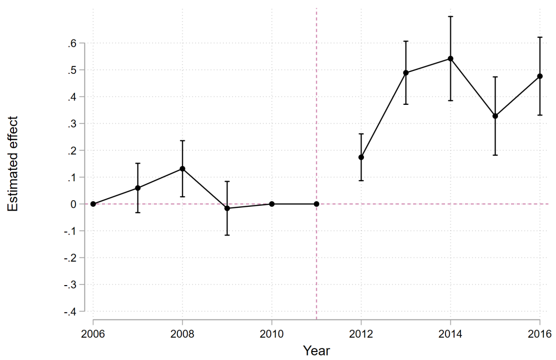
≤8 Weeks



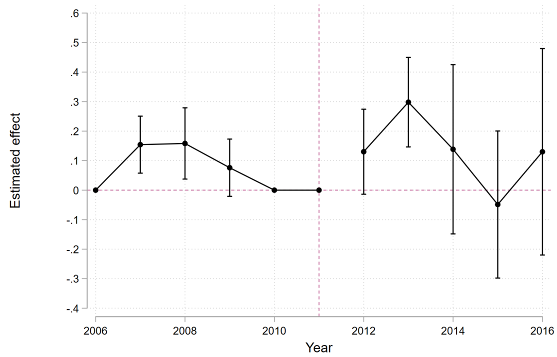
9–10 Weeks



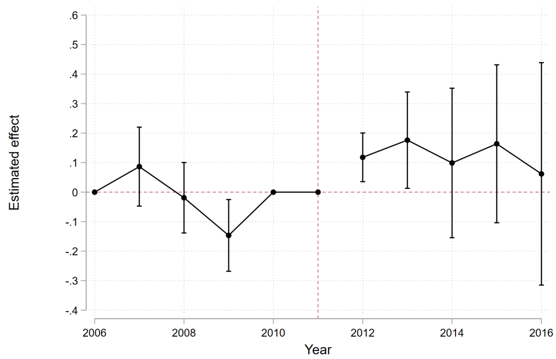
11–12 Weeks



13–14 Weeks



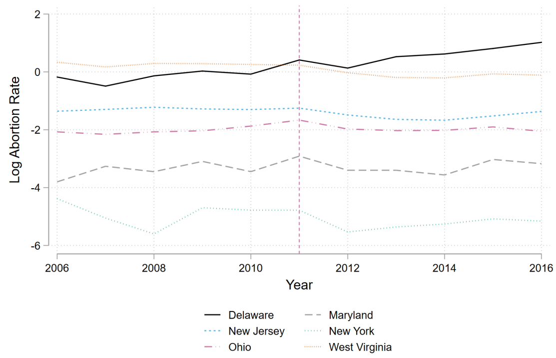
15+ Weeks



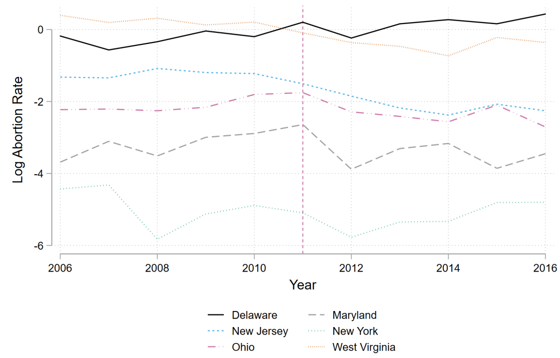
Notes: This figure plots the estimated effect of the passage of the law on abortion rates for various age groups. Estimates come from a Poisson model which controls for county and year fixed effects. Treated counties are those for which Pittsburgh was the nearest abortion-providing city in 2006, comparison counties are those for which Allentown, Harrisburg, Philadelphia, Reading, Upland, Warminster, or West Chester was the nearest abortion-providing city in 2006. All other counties in the state are included in the comparison group.

Figure A5
Log Abortion Rate, Out-of-State Women, By Age Group

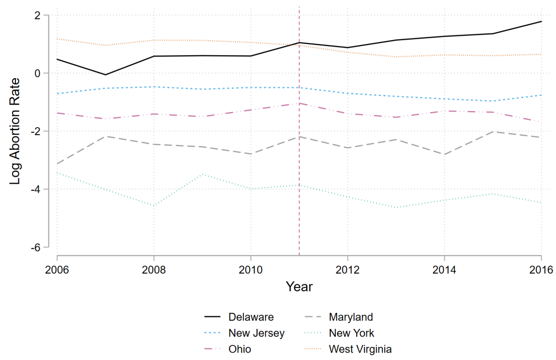
Total



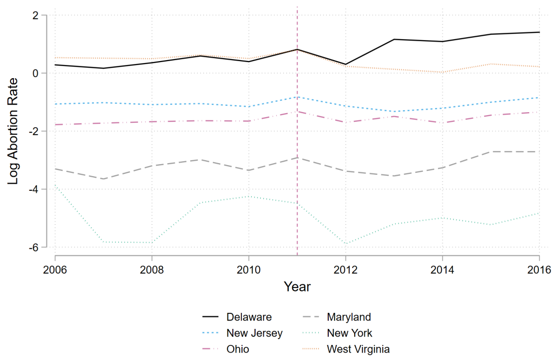
15–19 Year Olds



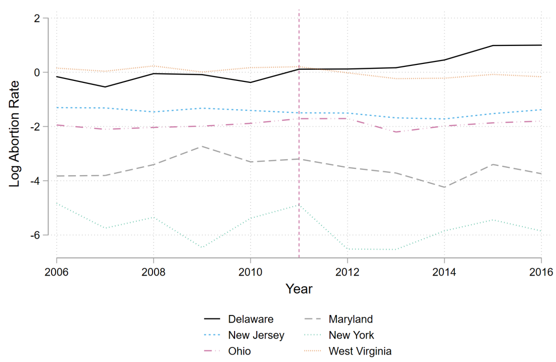
20–24 Year Olds



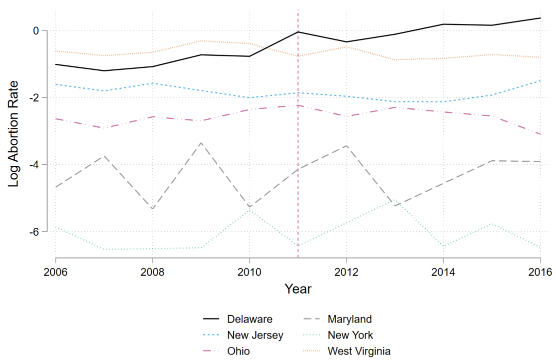
25–29 Year Olds



30–34 Year Olds

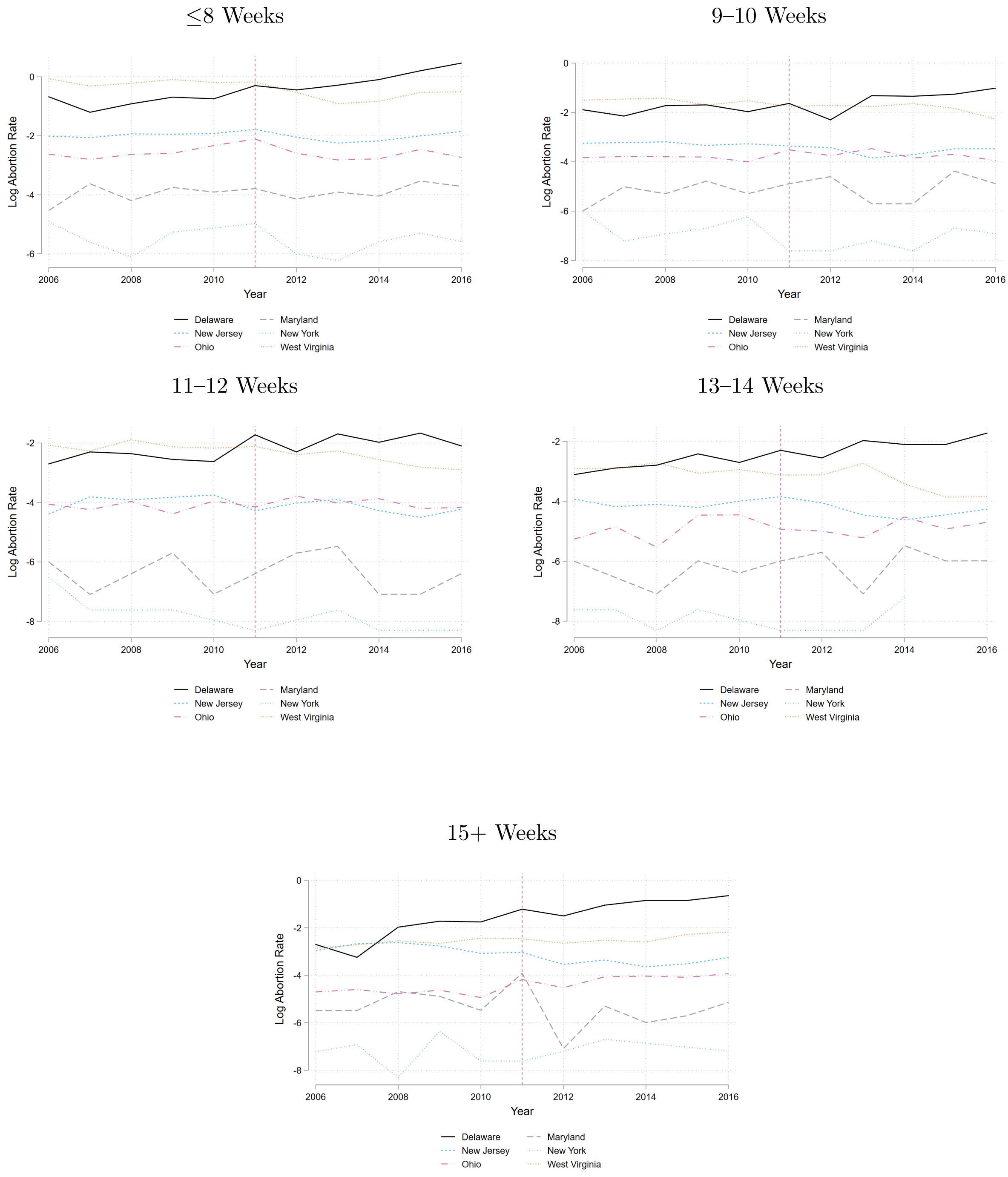


35–39 Year Olds



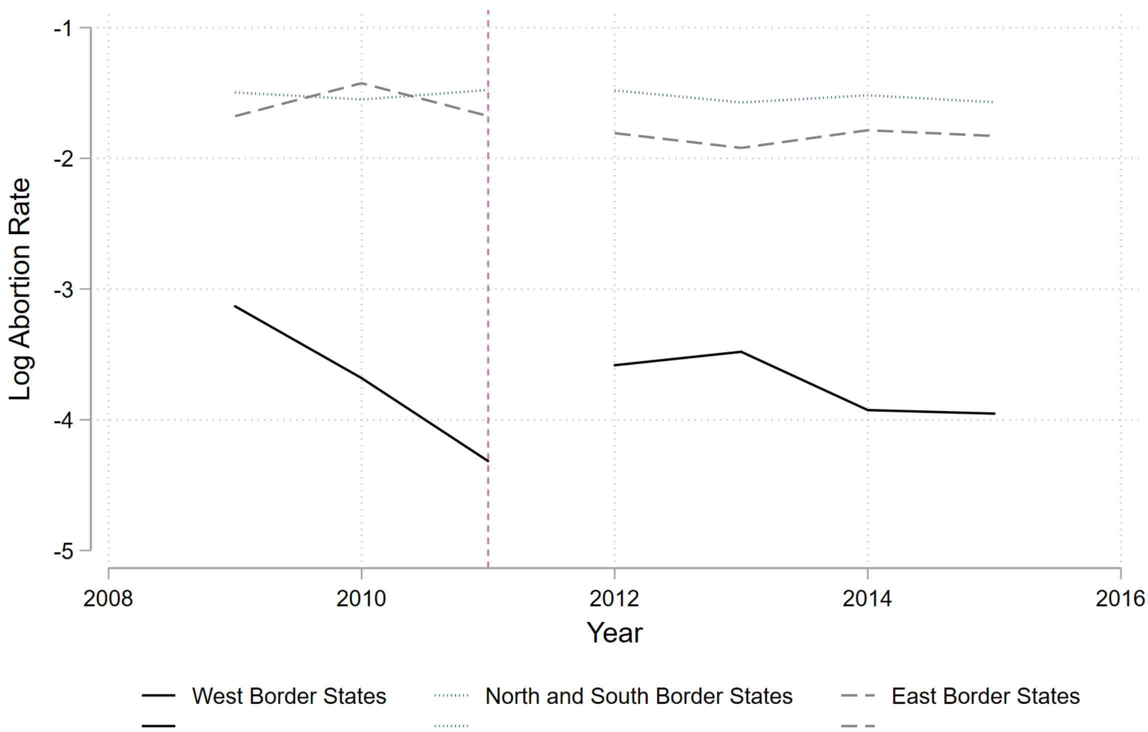
Notes: This figure plots the natural log of the abortion rate for women coming to Pennsylvania from other states to obtain an abortion. The states shown are the only states for which state-specific data is provided in the Pennsylvania Department of Health's Annual Abortion Report.

Figure A6
Log Abortion Rate, Out-of-State Women, By Gestational Age



Notes: This figure plots the natural log of the abortion rate for women coming to Pennsylvania from other states to obtain an abortion. The states shown are the only states for which state-specific data is provided in the Pennsylvania Department of Health's Annual Abortion Report.

Figure A7
Log Abortion Rate for PA Residents Traveling Out of State for Abortions



Notes: This figure plots the natural log of the abortion rate for Pennsylvania women obtaining abortions outside of Pennsylvania. Data are from the CDC Abortion Surveillance dataset.

Table A1

Summary Statistics for Pennsylvania vs. Texas vs. Rest of US

Percent of Population that is aged 15-44	39.34	43.80	41.51
Percent Poverty Rate	11.68	16.31	13.40
Population	12581173.94	24301626.35	2.66e+08
Percent White	81.13	47.68	66.81
Percent Hispanic	5.15	36.27	14.09
Percent Black	10.81	11.79	12.92

Table A2
Estimated Effects of Abortion Regulations on Abortion Rates by Age, Including All Counties in PA

	Total	Teen	Non-Teen
First Year of the Law	-0.026	-0.038	-0.026
	(0.031)	(0.068)	(0.030)
Second Year of the Law	-0.036	-0.010	-0.037
	(0.043)	(0.070)	(0.040)
Third Year of the Law	-0.018	-0.012	-0.018
	(0.047)	(0.093)	(0.043)
Fourth Year of the Law	0.058	0.041	0.062
	(0.067)	(0.125)	(0.061)
Fifth Year of the Law	0.060	0.022	0.065
	(0.083)	(0.144)	(0.075)
Average effect	0.008	0.001	0.009
P-value (test average effect = 0)	0.879	0.993	0.841
Observations	737	737	737

Notes: Results come from a Poisson model with county and year fixed effects and county-specific linear time trends.
***, **, and * represent p-values less than 0.01, 0.05, and 0.10, respectively.

Table A3
Estimated Effects of Abortion Regulations on Abortion Rates by Gestational Age, Including All Counties in PA

	≤8 Weeks	9–10 Weeks	11–12 Weeks	13–14 Weeks	15+ Weeks
First Year of the Law	-0.167***	0.231***	0.187***	0.087	0.214***
	(0.032)	(0.041)	(0.040)	(0.084)	(0.051)
Second Year of the Law	-0.333***	0.325***	0.514***	0.265***	0.292***
	(0.052)	(0.039)	(0.048)	(0.091)	(0.103)
Third Year of the Law	-0.267***	0.306***	0.578***	0.116	0.234
	(0.052)	(0.050)	(0.070)	(0.159)	(0.146)
Fourth Year of the Law	0.016	0.013	0.375***	-0.061	0.319**
	(0.068)	(0.056)	(0.060)	(0.143)	(0.156)
Fifth Year of the Law	-0.022	0.077	0.535***	0.129	0.236
	(0.079)	(0.082)	(0.069)	(0.196)	(0.209)
Average effect	-0.155	0.191	0.438	0.107	0.259
P-value (test average effect = 0)	0.003	0.000	0.000	0.389	0.045
Observations	737	737	737	737	737

Notes: Results come from a Poisson model with county and year fixed effects and county-specific linear time trends.
***, **, and * represent p-values less than 0.01, 0.05, and 0.10, respectively.

Table A4

Estimated Effects of Abortion Regulations on Abortion Rates by Age, Excluding Pittsburgh

	Total	Teen	Non-Teen
First Year of the Law	0.003	-0.070	0.015
	(0.041)	(0.126)	(0.038)
Second Year of the Law	-0.021	0.069	-0.030
	(0.058)	(0.086)	(0.056)
Third Year of the Law	0.031	0.005	0.040
	(0.061)	(0.147)	(0.055)
Fourth Year of the Law	0.146**	0.099	0.158**
	(0.070)	(0.166)	(0.064)
Fifth Year of the Law	0.091	-0.017	0.110
	(0.096)	(0.186)	(0.089)
Average effect	0.050	0.017	0.059
P-value (test average effect = 0)	0.406	0.890	0.286
Observations	385	385	385

Notes: Results come from a Poisson model with county and year fixed effects and county-specific linear time trends.
***, **, and * represent p-values less than 0.01, 0.05, and 0.10, respectively.

Table A5
Estimated Effects of Abortion Regulations on Abortion Rates by Gestational Age, Excluding Pittsburgh

	≤8 Weeks	9–10 Weeks	11–12 Weeks	13–14 Weeks	15+ Weeks
First Year of the Law	-0.156***	0.304***	0.156**	0.213	0.285***
	(0.047)	(0.065)	(0.075)	(0.143)	(0.083)
Second Year of the Law	-0.365***	0.380***	0.559***	0.374**	0.383***
	(0.070)	(0.071)	(0.102)	(0.165)	(0.139)
Third Year of the Law	-0.285***	0.410***	0.624***	0.392*	0.394**
	(0.065)	(0.099)	(0.127)	(0.229)	(0.170)
Fourth Year of the Law	0.096	0.081	0.357***	0.143	0.514***
	(0.074)	(0.097)	(0.123)	(0.220)	(0.173)
Fifth Year of the Law	-0.041	0.203	0.545***	0.185	0.406*
	(0.087)	(0.151)	(0.139)	(0.254)	(0.231)
Average effect	-0.150	0.276	0.448	0.262	0.397
P-value (test average effect = 0)	0.013	0.001	0.000	0.152	0.006
Observations	385	385	385	385	385

Notes: Results come from a Poisson model with county and year fixed effects and county-specific linear time trends.
***, **, and * represent p-values less than 0.01, 0.05, and 0.10, respectively.