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Sex-selective abortion bans in the United States: evidence from laws passed 2010-2019

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Abstract

This paper attempts to estimate the effect of sex-selective abortion bans passed in the United States between 2010 and 2019 on the sex-ratio at birth among Asians and other Pacific Islanders. We use both difference-in-differences (DiD) across state and year and difference-in-difference-in-differences (DiDiD) across state, year, and child parity to identify this effect. While we find reason to expect that our DiD estimates will overstate the effect of a sex-selective abortion ban, our DiDiD estimates are robust to this. None of our estimates find statistically significant effects of sex-selective abortion bans, and the point estimates from our DiDiD estimates imply these bans leave the sex-ratio at birth among 3rd and higher parity children slightly elevated, changing from 108.1 boys to every 100 girls to 107.3 boys to every 100 girls.

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1 Introduction

The sex-ratio at birth, or the number of boys born for every one hundred girls, is usually between 103 and 107 (Anderson and Ray 2010).[1] However, elevated or son-biased sex-ratios at birth among 3rd and higher parity children of first-generation Chinese, Korean, and Indian immigrants after the birth of two girls have been found in the United States (Almond and Edlund 2008; Abrevaya 2009). These elevated sex ratios are commonly interpreted as providing evidence that first-generation Chinese, Korean, and Indian immigrants have continued the practice of selectively aborting female fetuses while living in the United States.

Largely in response to these studies (Kalantry 2015), states have begun implementing sex-selective abortion bans, which prohibit abortion providers from providing abortions when the reason for the abortion is the sex of the fetus. In 2010, Oklahoma became the first state to pass a ban on sex-selective abortions after Almond and Edlund (2008) and Abrevaya (2009).[2] By the end of 2020, 10 other states had followed Oklahoma in passing such bans.[3]

Despite the prevalence of sex-selective abortion bans, Nandi et al. (2015) is the only empirical study we are aware of to have considered their effect on the sex-ratio at birth. While Nandi et al. (2015) do not find any evidence that the sex-ratio at birth among Chinese, and other Asian and Pacific Islanders changes after sex-selective abortion bans, they consider only two, very early examples of these laws, which went into effect well before Almond and Edlund (2008) and Abrevaya (2009). Kalantry (2015) further notes the two specific laws considered in Nandi et al. (2015) were the only two laws in which legislators did not heavily cite the practice of sex-selective abortions among certain immigrant groups as a motivating factor for the ban.

Theoretically, there are reasons to believe that sex-selective abortion laws may be ineffective. First, most sex-selective abortion bans only prohibit abortion providers from knowingly providing a sex-selective abortion but impose no duty on the abortion provider to ask a patient the reason for the abortion.[4] Secondly, even if a patient is asked the reason they are getting an abortion, it is not clear why someone seeking a sex-selection would tell the truth, knowing that they would then be denied the abortion. Third, even if sex-selective abortion bans are effective within states that pass them, out-of-state abortions may be readily available to patients. Fourth, sex-selective abortions are not the only way to choose the sex of a child and sex-selective abortions generally do not apply to other methods of sex selection. Lastly, while the evidence in Almond and Edlund (2008) and Abrevaya (2009) is highly suggestive, as Anderson and Ray (2010) note, it is not conclusive evidence that sex-selective abortions are occurring. To the extent that biological or other factors are dominant factors in determining the elevated sex-ratios across higher parity children in these immigrant groups, we would not expect sex-selective abortion bans to influence these high sex ratios.

This paper considers the effect of sex-selective abortion bans on the probability an Asian or Pacific Islander has a boy by child parity using data from the National Vital Statistics System

[1]The range of 103 to 107 is fairly typical in the literature and comes from the sex-ratio at birth among African Americans and Asian Americans respectively (Anderson and Ray 2010).

[2]Illinois and Pennsylvania passed sex-selective abortion bans well before either study.

[3]These are Arizona, Arkansas, Indiana, Kansas, Kentucky, Missouri, North Carolina, North Dakota, South Dakota, and Tennessee.

[4]See the appendix table on pages 28-29 in Citro et al. (2014).

(NVSS). Our main contribution to the literature is considering the effects of sex-selective abortion bans using the roll-out of sex-selective abortion bans passed subsequent to Almond and Edlund (2008) and Abrevaya (2009) from 2010 to 2019. We use both difference-in-differences (DiD) and (DiDiD) methods to estimate the effects of sex-selective abortion bans. While we present evidence that the DiD model may be biased to overstate the effect of a sex-selective abortion ban, the DiDiD is robust this.

We do not find any statistically significant effect of sex-selective abortion bans on the probability that an Asian or other Pacific Islander has a boy, regardless of child parity. The point estimate from our preferred model suggests that such bans leave the sex-ratio at birth slightly elevated, decreasing this number from 108.1 boys for every 100 girls to 107.3. While our estimates are not precise enough to rule out that sex-selective abortion bans may cause the sex-ratio at birth to fall under 107, a power analysis shows that our estimates did possess substantial power to reject the null hypothesis of no effect given changes in the probability of having a boy large enough to lower the sex-ratio at birth to 107. We therefore conclude that sex-selective abortion bans have little effect on the sex-ratio at birth among Asians and other Pacific Islanders.

2 Data

Our main data source on births is NVSS natality data accessed via the Center for Disease Control's Wide-ranging On-Line Data for Epidemiologic Research (CDC WONDER). This system tracks all births occurring within the United States to United States' residents.[5] The data was pulled from CDC WONDER organized into state, year, child sex, and live birth order cells, which is the highest cell level which would allow calculating the probability of having a boy by child parity at the state-year level. The main sample is further limited to births occurring to mothers who were Asian or other Pacific Islanders. The resultant sample includes 3,504,912 births from 2007-2019.

The high cell level was selected to preserve as many births as possible as CDC WONDER suppresses any cells with fewer than 10 births. To check how severe of a problem cell suppression may pose to our estimates, we pulled all births to Asian and other Pacific Islander mothers between 2007 and 2019. This showed 3,525,936 births over the sample period. Thus, even with the cell suppression, our final sample includes nearly 99.5% of births to Asians and Other Pacific Islanders over the relevant time frame.

The finest racial category available in CDC WONDER during our time frame is Asian and other Pacific Islander and nativity is not reported. This is not ideal, as elevated sex-ratios have only been observed in the United States in births to first-generation Chinese, Korean,[6] and Indian immigrants (Almond and Edlund 2008; Abrevaya 2009).[7] To get an idea of how many Asian and other Pacific Islanders in our sample may be first-generation Chinese, Korean, or Indian Immigrants, we used the Integrated Public-Use Microdata Series extract of the American Community Survey (IPUMS

[5] See https://wonder.cdc.gov/wonder/help/Natality.html#. Last accessed 2/26/21.

[6] Almond and Sun (2017) have also found using the 2010 Census that Korean-Americans no longer have elevated sex ratios at birth, which may indicate that excluding first-generation Koreans from the analysis would also be desirable

[7] Almond et al. (2013) find similarly elevated sex-ratios at birth among second-generation Chinese, Indian, and Korean immigrants in Canada. However, the United States Census does not record nativity to the second generation, so the degree to which this occurs in the United States may not be directly tested.

ACS). From this, we calculated that approximately 31% of Asian and other Pacific Islanders over our sample period were first-generation Chinese, Korean, or Indian immigrants, and this proportion rises slightly to 33% in our treatment group. Lastly, we note that while the use of such a broad racial category is not ideal, it is in line with Nandi et al. (2015) who look at births among Chinese and Asians and other Pacific Islanders.

It would also be ideal to incorporate information on the sex composition of children already in the household, as prior literature has generally documented that sex-ratios are birth are the most elevated following the birth of two girls. The data, unfortunately, do not allow for this as it is not matched across siblings. Thus, we can only observe the sex of the child that is born in the current year.

Table 1 presents a summary of the sex-ratio at birth births by treatment. As expected, the estimated probability of having a boy across first and second parity children place the sex-ratio at birth slightly elevated from that of non-Hispanic Whites (around 105.9 (Anderson and Ray 2010)), but well within the biologically normal range of 103-107 boys for every 100 girls. The sex ratios across 3rd-plus parity children, however, are higher for both states which never pass a sex-selective abortion ban and in states before a ban is effective. The sex-ratio at birth for these groups respectively are 107.9 and 108.1, respectively. Similar results are found over our sample period as shown in Figure A.1 in the appendix. In states with an effective sex-selective abortion ban, the probability of having a boy appears to fall and implies a sex-ratio at birth of 105.8. However, we caution our reader from attributing this drop to the sex-selective abortion ban itself, as changes in demographics or other variables could be the cause of such a change.

Table 1: Sex-Ratio at Birth among Asians and Other Pacific Islanders by Treatment

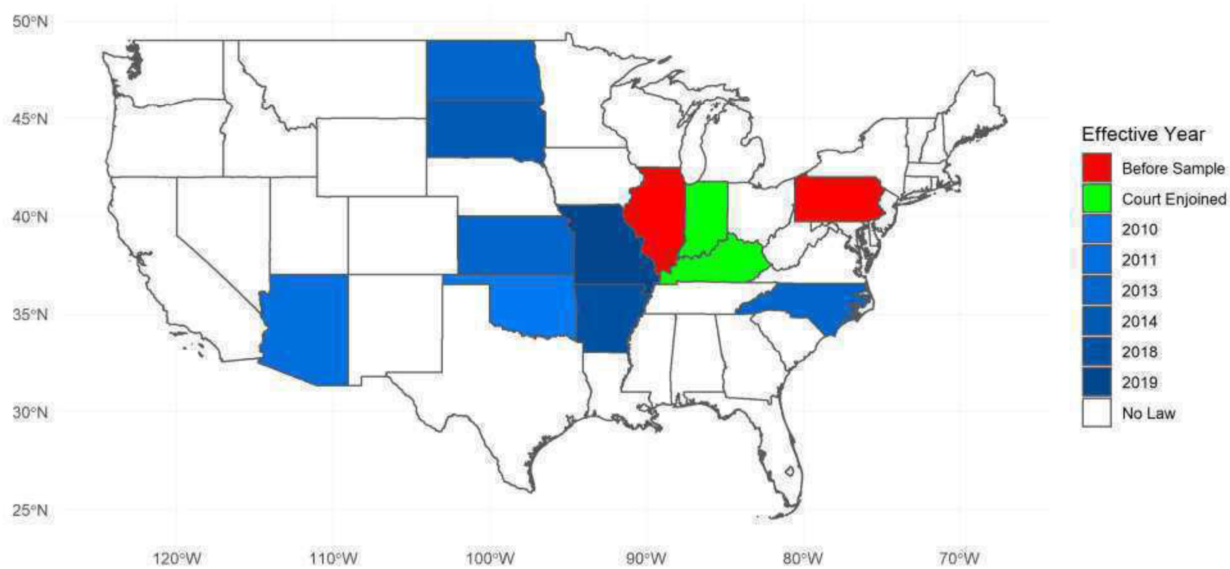
	No Ban	Pre-Ban	Post-Ban
1st Parity	106.1	106.7	105.4
	[105.8, 106.4]	[104.6, 108.8]	[103.5, 107.3]
N	1,489,678	39,716	46,044
2nd Parity	106.6	106.1	104.2
	[106.2, 107.0]	[103.8,108.5]	[102.1, 106.3]
N	1,193,285	31,140	37,889
3rd+ Parity	107.9	108.1	105.8
	[107.4, 108.5]	[105.1, 111.1]	[103.2, 108.5]
N	622,207	19,929	25,024

95% confidence intervals, based on the standard deviation of the probability of having a boy, are given in brackets. No Ban indicates the state in which the birth occurred did not pass a sex-selective abortion ban on our sample, Pre-Ban indicates the state in which the birth occurred implemented a sex-selective abortion ban after the birth occurred, and Post-Ban indicates that the birth occurred in a state with an effective sex-selective abortion ban. The estimate is built from birth certificate data from CDC WONDER encompassing births to Asian and other Pacific Islanders in the United States from 2007 to 2019.

To control for state-level demographics and other factors which may be changing over our sample period, we also use data from IPUMS ACS and the Current Population Survey's Merged Out-Going Rotation Group (IPUMS CPS-MORG). From the IPUMS ACS, we calculated the percent of the population that was non-Hispanic White, non-Hispanic Black, Asian, a first-generation Chinese, Indian, or Korean immigrant, and Hispanic. We also calculated the percent of the population which had a high-school diploma, some college, or a Bachelor's degree or higher, the number of Females aged 15-44, and the unemployment rate. From the IPUMS CPS-MORG, we calculated the median wage for females aged 15-44. A summary of these controls by treatment is given in the appendix.

The treatment dates used in the analysis are based on the year in which a sex-selective abortion ban went into effect. These dates are from the author's own research and were collected through a search of state legislative websites and state-wide newspapers. Figure 1 summarizes these dates, while the exact dates are given in Table A.1 in the appendix. The reader will note that there is some concern of clustering in the treatment. The states that passed sex-selective abortion bans are our sample do tend to be smaller states, with a larger percent of the population being white, and a lower percentage of college graduates. These states also seem to be diversifying over the sample period. We generally attempt to correct for this by 1) inclusion of the above controls, and 2) trying alternative control groups for the DiD model.

Figure 1: Treatment Dates



Alaska and Hawaii do not have sex-selective abortion bans. Tenessee's sex-selective abortion ban did not take effect until after the sample period in November of 2020.

3 Model and Identification

Our initial model is a generalized DiD estimator given by the following

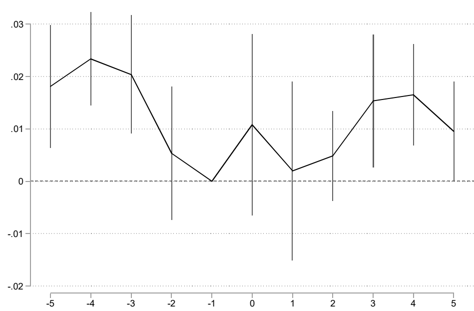
$$Boy_{it} = \beta_0 + \beta_1 Ban_{st} + X_{st}\gamma + \delta_s + \tau_t + \epsilon_{it} \quad (1)$$

In the above, Boy_{it} is an indicator equal to one if child i born in year t is a boy and zero otherwise. Ban_{st} is an indicator equal to one if state s had an effective sex-selective abortion ban in year t. X_{st} is a vector of state-level characteristics including controls for racial composition, educational composition, the log number of females aged 15-44, and the median female wage. δ_s and τ_t are state and year fixed effects respectively, and ϵ_{it} is an idiosyncratic error term.

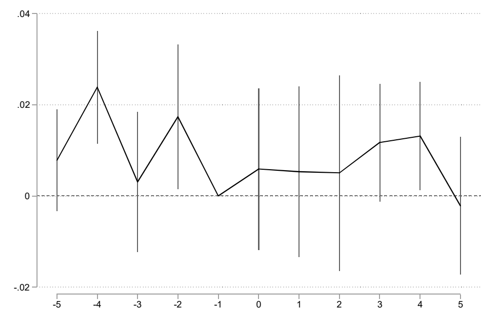
All models are fit with ordinary least squares (OLS) and are linear probability models. Standard errors are clustered at the state level. Models are fit separately by child parity for 1st, 2nd, and 3rd-plus parity children. β_1 is our coefficient of interest and gives the DiD estimate of the change in probability a child of the given parity is a boy due to a sex-selective abortion ban. As elevated sex-ratios at birth are only documented across higher parity children, we report the DiD estimate on 1st parity children as a falsification test.

The main identifying assumption for our DiD estimates is parallel trends. We check this assumption graphically by plotting trends in the probability of having a boy by treatment over our sample period and conducting an event study based on equation (1) in which Ban_{st} is replaced with a series of leads and lags for Ban_{st}. The results of the event study are presented in Figure 2, while the overall trends are given in the appendix. While the trend for our treatment group is extremely noisy, the probability of having a boy in the treatment group does generally fluctuate around a similar level to the control over the sample. However, the event study reveals a pretreatment "bump" in the probability of having a boy in the treatment group. Since this bump leaves the treatment group downward sloping relative to the control group prior to the treatment, we believe that the DiD will be negatively biased. As we expect sex-selective abortion bans to have a non-positive effect on the probability of having a boy, this bias should then overstate the effect of a sex-selective abortion ban.

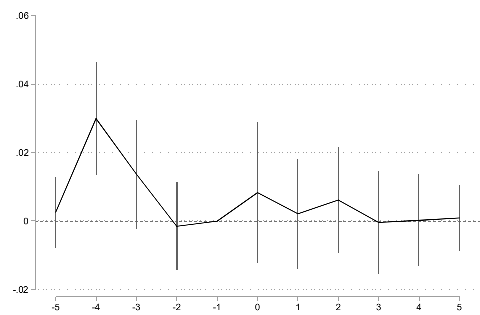
Figure 2: Event Study on the Probability an Asian or Pacific Islander Newborn is a Boy without State-Level Controls



(a) First



(b) Second



(c) Third Plus

The horizontal axis gives years relative to treatment. The horizontal line is the point estimate of the coefficient while the vertical line gives a 95% confidence interval. The regression also included state and year fixed effects. Estimates including state-level controls are given in the appendix.

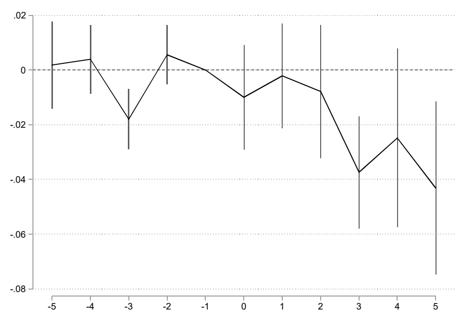
As the DiD estimate is biased, we further extend our analysis to a DiDiD estimator based on Yelowitz (1995) which is more robust. This model is given by,

$$
\begin{aligned}
Boy_{it} &= \beta_0 + \beta_1 Ban_{st} * HighParity_{it} + \delta_s + \tau_t \\
&\quad + \delta_s * HighParity_{st} + \tau_t * HighParity_{it} + \rho_{st} + \epsilon_{it}
\end{aligned} \tag{2}
$$

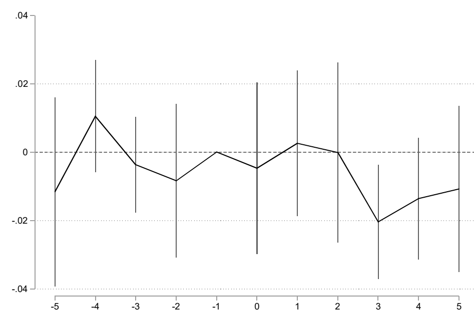
The major differences from equation (1) in (2) are that now Ban_{st} is interacted with $HighParity_{it}$, which is an indicator equal to one if child i born in year t is high parity and zero otherwise. We define high parity as 2nd-plus and 3rd-plus parity alternatively. We further interact each of the state and year fixed effects with the same indicator for high parity, and include ρ_{st} which is a series of state-year effects. We do not control for any state-specific covariates in (2), because the effects of these variables are removed by the state-year effects. The coefficient of interest in equation (2) is β_1 and is the DiDiD estimate of the effect of a sex-selective abortion ban on the probability a high-parity child is a boy relative to lower parity children.

The main identifying assumption in (2) is that the difference in the probability of having a boy between high and low parities is trending similarly between treated and control states (Olden and Møen 2020). The aforementioned plot of the overall trends in the appendix provides a provides a visual check of this assumption. We further attempt to demonstrate the validity of this assumption with an event study based on (2) with $Ban_{st} * HighParity_{it}$ replaced with a series of lead and lags and fit on our sample of treated states. This then checks the main identifying assumption for the second DiD in our DiDiD. The results of this event study are presented in Figure 3 and generally confirm that the probability of having a boy at high and low parities are trending similarly within the treated states prior to a sex-selective abortion ban. Therefore, our preferred estimates are from equation (2).

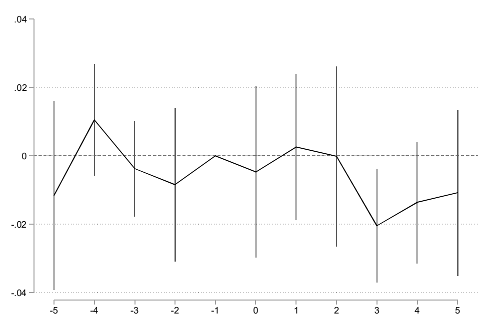
Figure 3: Event Study on the Probability a High-Parity Asian or Pacific Islaner Newborn is a Boy within Treated States



(a) Second Plus Relative to First



(b) Third Plus Relative to First and Second



(c) Third Plus Relative to First

The horizontal axis gives years relative to treatment. The horizontal line is the point estimate of the coefficient while the vertical line gives a 95% confidence interval. The regression also included state by parity, year by parity, and state by year fixed effects.

There is a concern that the changes in the probability of having a boy necessary to decrease the sex-ratio at birth across third-plus parity children on our sample are quite small. To decrease the sex-ratios from Table 1 to 107 would require a decrease in the probability of having a boy of .0022 and .0025 in the control group and the treatment group before the treatment respectively. As such, the reader may be skeptical of our ability to detect such small changes. To address this, we conducted a power analysis largely following Nandi et al. (2015). For the power analysis, we used 30% of our full sample size, as our estimates with IPUMS ACS suggest that approximately 30% of Asians and other Pacific Islanders would have been first-generation Chinese, Indian, or Korean immigrants over our sample period. The actual standard deviations of the indicator for having a boy and treatment were used. The null hypothesis was that the slope in the simple linear regression is zero. The results of the power analysis are presented in Figure A.4 in the appendix. We find that we have a power of .8 at approximately a decrease in the probability of having a boy of .0014, and have power of well over .8 at changes large enough to lower the sex-ratio at birth to 107. We therefore believe our estimates have substantial power to detect any economically significant changes in the probability of having a boy due to a sex-selective abortion ban.

As Kalantry (2015) argues, there is some concern that sex-selective abortion bans may be correlated with other abortion restrictions. If this is the case, then other abortion restrictions in states with a sex-selective abortion ban may be influencing pretreatment trends or be responsible for any observed results. To test this hypothesis, we gathered data on abortion incidence and availability from the CDC and Guttmacher Institute. This was used to conduct a preliminary analysis on the effect of sex-selective abortion bans on the number of abortions. The results of the analysis are presented in Figures A5-A7 and Table A4 of the appendix and do not appear to be consistent with the hypothesis that abortions were being further restricted in states with a sex-selective abortion ban.

Lastly, the reader may have some concern that family size is endogenous. For instance, Dahl and Moretti (2008) documents that households with a first-born daughter tend to have higher fertility, and Blau et al. (2020) finds this preference among immigrants. We, however, do not think this preference will harm our estimates in this case, as we are directly analyzing the probability of having a boy. Absent sex-selection, the probability of having a boy should not depend on household size, as each the sex of the child at birth should still be an independent event.

4 Results

Table 2 presents our DiD estimates of the effect of sex-selective abortion bans by child parity. All models are linear probability models, and the coefficient is interpreted as the change in the probability that an Asian or other Pacific Islander child of the noted parity is a boy due to a sex-selective abortion ban. Standard errors are clustered at the state level.

Across child parity, our DiD estimates are not statistically different from zero at any traditional confidence level. Further, the point estimate for the effect of sex-selective abortion bans on first-parity children is smaller than the estimate across 3rd-plus parity children; however, the difference is small and unlikely to be statistically significant. This is somewhat suggestive that these policies have little effect on the sex-ratio at birth as the estimate across first-parity children are included as a falsification test. Despite the statistical insignificance of the coefficients, the estimated size of the coefficients across 3rd-plus parity children are not exactly trivial. The point estimates imply a change in the sex-ratio at birth across 3rd-plus parity children from 108.1 before a ban to 105.9

and 106.3 with and without controls respectively. However, we again caution our reader that we expect our DiD estimates to overstate the true effect of a sex-selective abortion ban.[8]

Table 2: DiD Estimates of the Effect of Sex-Selective Abortion Bans on the Probability of Having a Boy by Parity

	(1) 1st Child	(2) 2nd Child	(3) 3rd Plus
Ban	-0.0033	-0.0050	-0.0043
	(0.003)	(0.004)	(0.003)
Controls	No	No	No
N	1,575,438	1,262,314	667,160
Ban	-0.0017	-0.0033	-0.0052
	(0.004)	(0.004)	(0.004)
Controls	Yes	Yes	Yes
N	1,575,438	1,262,314	667,160

Standard errors are clustered at the state level and reported beneath the estimated coefficient in parentheses. The dependent variable is an indicator equal to one if the child was a boy and zero otherwise. Thus, all models are linear probability models and the estimated coefficient is interpreted as the DiD estimate of the change in the probability that a newborn Asian or other Pacific Islander was a boy due to a sex-selective abortion ban. Controls include the log number of females, the unemployment rate, the percent of the population that was black, white, Asian, a first-generation Chinese, Korean, or Indian immigrant, or Hispanic, the percent of the population who have a high-school diploma, some college, or a Bachelor's or higher. The time frame of the sample is 2007-2019.
* $p < .1$ ** $p < .05$ *** $p < .01$

Table 3 presents our DiDiD estimates, which corrects for the potential of our DiD estimates to overstate the effect of a sex-selective abortion ban. Models are linear probability and standard errors are clustered at the state, child-parity level, which is the level of identification. The reported coefficient is the DiDiD estimate and is interpreted as the change in the probability a high-parity

[8] We further considered several alternative control groups with our DiD estimator. These included boarding states, states which also considered a sex-selective abortion ban from 2010-2014, and states which passed a sex-selective abortion ban that was court enjoined. While this did not seem to further smooth the pretreatment trends, estimates appeared to be similar in all cases but the last, in which case our estimates became very imprecisely estimated. These results are presented in Table A3 of the appendix.

Asian or other Pacific Islander child is a boy relative to a lower parity due to a sex-selective abortion ban.

The point estimates for the DiDiD estimate are somewhat smaller in magnitude than point estimates for the DiD estimates for 3rd-plus parity children, which is in line with our reasoning that the DiD estimates may be biased to overstate the effect of a sex-selective abortion ban. While the sex-ratio at birth across second-plus children is not above the biological norm of 103-107 in our control group or before a ban is effective, the DiDiD estimate only implies a small change in the sex-ratio at birth from 106.9 to 106.1. The point estimate for the DiDiD across 3rd-plus children relative to first and second-parity children further leaves the sex-ratio at birth slightly elevated and implies a change from 108.1 before a ban to 107.3. The point estimate for the DiDiD across 3rd parity children relative to first implies a similar change from 108.1 to 107.1.

Table 3: DiDiD Estimates of the Effect of Sex-Selective Abortion Bans on the Probability of Having a Boy by Parity

	(1) All Parity	(2) All Parity	(3) 1st and 3rd
Ban*2nd Plus	-0.0018		
	(0.003)		
Ban*3rd Plus		-0.0019	-0.0022
		(0.005)	(0.005)
N	3,504,912	3,504,912	2,242,598

Standard errors are clustered at the state, child-parity level, which is the level of identification, and reported beneath the estimated coefficient in parentheses. The dependent variable is an indicator equal to one if the child was a boy and zero otherwise. Thus, all models are linear probability models and the estimated coefficient is interpreted as the DiDiD estimate of the change in the probability that a high-parity, newborn Asian or other Pacific Islander was a boy due to a sex-selective abortion ban. The time frame of the sample is 2007-2019.
* $p < .1$ ** $p < .05$ *** $p < .01$

While the point estimates suggest that the sex-ratio at birth across 3rd-plus parity children remains slightly elevated after a sex-selective abortion ban, it should be noted the imprecision of our estimates mean that we cannot directly rule out a change in the sex-ratio at birth to below 107 with any meaningful degree of confidence. We therefore must rest on the previously mentioned power analysis to conclude that we had substantial ability to detect changes large enough to be economically meaningful.

5 Conclusions

We consider the effect of sex-selective abortion bans in the United States on the probability an Asian and other Pacific Islander child is a boy by child parity, using both a DiD and a DiDiD methodology. While our DiD estimates are biased to overstate the effect of a sex-selective abortion ban, our DiDiD model corrects for this.

We do not find any statistically significant effects of a sex-selective abortion ban on the probability that an Asian or other Pacific Islander child is a boy. The point estimates from our DiDiD model also imply that a sex-selective abortion ban leaves the sex-ratio at birth among third-plus parity Asian and other Pacific Islander children slightly elevated. While our estimates are not precise enough to directly rule out a change in the sex-ratio at birth below 107, we conduct a power analysis which suggests that our estimates had substantial power to detect changes in the probability an Asian or other Pacific Islander child is a boy large enough to decrease the sex-ratio to 107. We therefore view our estimates as suggesting that sex-selective abortion bans likely have little, if any, effect on the sex-ratio at birth among Asians and other Pacific Islanders.

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