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Sex-selective Abortion Bans are Not Associated with Changes in Sex Ratios at Birth among Asian Populations in Illinois and Pennsylvania

Abstract: Legal prohibitions on sex-selective abortions are proliferating in the United States. Eight state legislatures have banned abortions sought on the basis of the sex of the fetus, 21 states have considered such laws since 2009, and a similar bill is pending in U.S. Congress. These laws have been introduced and enacted without any empirical data about their impact or effectiveness. Prior studies of U.S. Census data found sex ratios among foreign-born Chinese, Korean and Indian immigrants were skewed in favor of boys, but only in families where there were already one or two girls. Using the variation in the timing of bans in Illinois and Pennsylvania as natural experiments, we compare the pre-ban and post-ban sex ratios of certain Asian newborn children in these states over 12-year periods. We then compare these ratios with the sex ratios of Asian newborn children in neighboring states during the same period. We find that the bans in Illinois and Pennsylvania are not associated with any changes in sex ratios at birth among Asians. In Illinois and its neighboring states, the sex ratio at birth of Asian children was not male-biased during our study period. On the other hand, the sex ratio at birth among Asians in Pennsylvania and its neighboring states was skewed slightly in favor of boys, but the enactment of the ban did not normalize the sex ratio. This strongly suggests that sex-selective abortion bans have had no impact on the practice of sex selection, to the extent that it occurs, in these states. This finding is highly relevant to legislative and policy debates in the U.S. Congress and state legislatures where sex-selective abortion laws are being considered.

Keywords: abortion; sex selection.

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1 Introduction

A preference for sons and the practice of sex selection have been documented in many countries in Asia over the past century, including China, India, Korea and Taiwan (Park and Cho 1995; Sen 2003; Lin et al. 2008; Zhu et al. 2009; Ebenstein 2010; Jha et al. 2011). Sex selection has resulted in very high male-to-female sex ratios in these countries (Das Gupta and Shuzhuo 1999; Das Gupta et al. 2003). The abortion of female fetuses is one method by which parents ensure they have a boy-child.

Recent studies have shown evidence of slightly male-biased sex ratios among the newborn children of foreign-born Chinese, Indian and Korean parents in the United States (Almond and Edlund 2008; Abrevaya 2009; Egan et al. 2011). These findings suggest that when these parents have two female children, they may preselect to ensure their third child is a boy.[1] There is no way of determining, however, whether abortion is used as the method of sex selection in these cases. Indeed, studies relying on census data alone cannot isolate the methods used for sex selection, which may take the form of abortion or pre-implantation techniques, such as sperm-sorting and genetic diagnosis.

In recent years, pointing to the practices of a very small segment of the U.S. population (<0.01%), opponents of abortion have pushed for the introduction of bans on sex-selective abortion in numerous states. Since 2011, Arizona, Kansas, North Carolina, North Dakota, Oklahoma and South Dakota have enacted laws prohibiting sex-selective abortions.[2] Twenty-one other states have considered such laws since 2009.[3] In 2008, Republican Trent Franks introduced the Susan B. Anthony Prenatal Nondiscrimination Act in the U.S. House of Representatives (H.R. 7016). The act would have instituted a federal ban on sex-selective abortions

1 A very small qualitative study of 68 Indian immigrant women has also found some direct evidence on sex-selective abortion practices (Puri et al. 2011).

2 Ariz. Rev. Stat. Ann. § 13-3603.02; Act of April 19, 2013, 2013 Kan. Laws Ch. 119 (H.B. 2253); Act of July 26, 2013, ch. 90, 2013 N.C. Laws 366; N.D. Cent. Code Ann. § 14-02.1-04.1; Okla. Stat. Ann. tit. 63, § 1-731.2; H.B. 1162, 89th Leg., Reg. Sess. (S.D. 2014) (enacted).

3 California, A.B. 2336, 2013–2104 Leg., Reg. Sess. (Cal. 2014); Colorado, see, e.g., H.B. 1131 and S.B. 56, 69th Gen. Assemb., Reg. Sess. (Colo. 2013); Florida, see, e.g., H.B. 845 and S.B. 1072, 115th Leg., Reg. Sess. (Fla. 2013); Georgia, H.B. 1155 and S.B. 529, 2009 Leg., Reg. Sess. (Ga. 2010); Iowa, S.F. 13, 85th Gen. Assemb. (Iowa 2014); Idaho, H.B. 693, 60th Leg., Reg. Sess. (Idaho 2010); Indiana, H.B. 1430 and S.B. 183, 118th Gen. Assemb., Reg. Sess. (Ind. 2013); Massachusetts, H.B. 1567, 188th Leg., Reg. Sess. (Mass. 2014); Michigan, see, e.g., H.B. 5731, 96th Leg., Reg. Sess. (Mich. 2012); Minnesota, see e.g., H.F. 1196 and S.F. 1073, 86th Leg., Reg. Sess. (Minn. 2010); Missouri, see, e.g., H.B. 1585, 97th Gen. Assemb., 2d Sess. (Mo. 2014); Mississippi, see, e.g., S.B. 2790, 129th Leg., Reg. Sess. (Miss. 2014); New Jersey, see, e.g., A.B. 2157, 215th Leg., Reg. Sess. (N.J. 2013); New York, see, e.g., A.B. 2533

"to prohibit discrimination against the unborn on the basis of sex or race, and for other purposes." The act died in committee, but was reintroduced several times. In 2012, lawmakers in the House of Representatives voted 246 to 168 in favor of the bill. However, it failed because it was considered under rules that required a two-thirds majority. The act was reintroduced by Republicans in the Senate in January 2013 (S. 138) and in the House in February 2013 (H.R. 447). Votes have not yet been scheduled on the two bills. Opponents of the bans on sex-selective abortions in the United States argue that the laws will reduce access to abortion and stigmatize Asian-American women (McDonough 2013).

To date, we are not aware of any empirical literature on the impact of sex-selective abortion bans in the United States. Prior to the recent legislative campaigns by anti-abortion groups, Illinois and Pennsylvania prohibited sex-selective abortions in 1984 and 1989, respectively.[4] This study is the first to evaluate the effectiveness of the bans in these two states, using the variation in the timing of the bans as a natural experiment. We use historical annual birth data from the National Center for Health Statistics (NCHS) during 1978–1994, and utilize difference-in-difference (DID) and triple difference (TD) regression frameworks that compare newborn sex ratios across different ethnic groups before and after the bans in Illinois and Pennsylvania. Outcomes in Illinois are compared with its neighboring states Wisconsin, Iowa, Missouri, Kentucky and Indiana during 1978 to 1989. Outcomes in Pennsylvania are compared with its neighbors New York, New Jersey, Delaware, Maryland, West Virginia and Ohio during 1983 to 1994. We employ several alternative econometric techniques to check the robustness of our findings.

2 Data

We use publicly available NCHS data from 1978 to 1994.[5] The NCHS collects annual cross-sectional data related to childbirth in the United States. Although the scope

and S.B. 2286, 237th Leg., Reg. Sess. (N.Y. 2014); Ohio, H.B. 570, 129th Gen. Assemb., Reg. Sess. (Ohio 2012); Oregon, see, e.g., H.B. 4034, 78th Leg. Assemb., Reg. Sess. (Or. 2014); Rhode Island, see, e.g., H.B. 7383 and S.B. 2376, 2014 Leg., Reg. Sess. (R.I. 2014); Texas, see, e.g., H.B. 309, 83d Leg., Reg. Sess. (Tex. 2013); Virginia, see, e.g., H.B. 98, 2014 Leg., Reg. Sess. (Va. 2014); West Virginia, see, e.g., H.B. 2371, 81st Leg., Reg. Sess. (W. Va. 2014); Wisconsin, see, e.g., A.B. 217 and S.B. 201, 101st Leg., Reg. Sess. (Wis. 2014). A Utah state senator initially proposed a sex-selective abortion ban but changed the proposed legislation to a bill to collect statistics on the reasons women report for seeking abortions. See S.B. 60, 60th Leg., Gen. Sess. (Utah 2013).

4 720 Ill. Comp. Stat. Ann. § 510/6-8; 18 Pa. Cons. Stat. Ann. § 3204.

5 These data were downloaded from the National Bureau of Economic research website. http://www.nber.org/data/vital-statistics-natality-data.html.

of the surveys has expanded over time, information collected generally includes demographic characteristics of the parents, birth history, antenatal and postnatal care, and health outcomes of the newborn. Prior to 1972, NCHS only captured a 50% sample of all reported births. Since 1985, a full 100% of reported births in all states are captured by NCHS surveys, effectively making it a census.

The NCHS survey started reporting the races of parents in 1978. From 1978 to 1991, only four categories of Asian races were reported: (1) Chinese, (2) Filipino, (3) Japanese, and (4) other Asian or Pacific Islanders.[6] The last category included Asian Indians and Koreans. Beginning in 1992, some states started reporting Asian Indians and Koreans as separate race categories.[7] To avoid a mismatch between the reported Asian race categories in NCHS from 1978 to 1991 on the one hand and after 1992 on the other hand, we combined Asian Indians, Koreans and Pacific Islanders into one category in the post-1992 data. This new post-1992 category is equivalent to the category "other Asian or Pacific Islanders" in the 1978 to 1991 data.

Our analysis thus considers two groups of newborn children: (1) Chinese and (2) other Asian or Pacific Islanders. We do not include Japanese or Filipinos in the latter category. A newborn baby is considered of a particular race only if both parents are of that race.

The NCHS data does not provide information on whether Asian parents were born in the United States or abroad. However, a rapid increase in Asian migration to the United States during the 1980s (Lee 1998; Kandel 2011) implies that a majority of the newborn Chinese and other Asian or Pacific Islander children in our data were likely to have foreign-born parents.[8] This is important for our analysis, as prior studies that have shown male-biased sex ratios at birth focused on foreign-born Asian immigrants (Almond and Edlund 2008; Abrevaya 2009; Egan et al. 2011).

We compile the 1979 to 1994 NCHS data into two different datasets for our analysis. The first dataset contains a total of 6,142,600 childbirths during 1978 to 1989 in Illinois and its neighboring states (i.e., 6 years before and 6 years after Illinois banned sex-selective abortion in 1984). Of these newborn children, 6296 were Chinese and 42,472 were other Asian or Pacific Islanders. The second dataset contains data on 9,010,467 newborn children during 1983 to 1994 in Pennsylvania

6 The other race categories were White, Black, American Indian (includes Aleuts and Eskimos), Hawaiian (includes Part-Hawaiian), and other nonWhite. NCHS uses the term "race" to classify all these population groups. For simplicity, we also use their definition.

7 Also reported separately were Samoan, Vietnamese, and Guamanian.

8 The Asian population in the US grew by 70% in the 1980s between 1980 and 1988 (Hollmann 1990). More than 65% of the US Asian population (including children) was foreign-born in 1990, and the proportion increased to almost 70% in 2000 (Lopez et al. 2013).

and its neighboring states (i.e., 6 years before and 6 years after Pennsylvania banned sex-selective abortion in 1989). Of these newborn children, 50,230 were Chinese and 116,068 were other Asian or Pacific Islanders. The raw estimates of the male-to-female sex ratios in these two datasets are presented in Tables 3 and 4 in the Appendix.

3 Methods

3.1 Pooled Analysis

Using the above data, we first conduct a pooled regression analysis separately for the two state groups. The pooled model provides an overview of newborn sex ratios among the Asian races over the study periods in the two state groups. Our regression model is a state and year fixed linear probability model of the likelihood of a newborn child being male. It includes the following covariates: indicators of child's ethnicity, live birth order, and father's and mother's education levels. We also separately run pooled regressions among children of 2nd or higher and 3rd or higher birth orders in both the Illinois and Pennsylvania state group samples.

3.2 Difference-in-difference and Triple Difference Analysis

Next, in order to analyze the effect of the abortion ban on sex ratios, we construct a linear difference-in-difference regression model of the following form:

$$Y_{igt} = \beta_0 + \beta_1 * treat_g + \beta_2 * post_t + \beta_3 * treat_g * post_t + X_{igt}\gamma + u_{igt} \quad (1)$$

for the i-th child in the g-th state and time t. Sex of the child (whether male) is denoted by Y. Binary variables $treat_g$ and $post_t$ denote the treatment status (i.e., whether the state is Illinois or Pennsylvania) and post-ban period (i.e., $t \geq 1984$ for the IL analysis, and $t \geq 1989$ for the PA analysis), respectively. The set of variables X includes father's and mother's education, and the live birth order of the child. u_{igt} is the iid error term of the model. The estimated coefficient β_2 will show the differential effect of the ban on the gender of the children born in a treatment state, as compared with children from other states, and also in comparison with the pre-ban period.

Evidence from more recent data points towards slightly male-biased sex ratios at higher birth parities among certain Asian groups (Almond and Edlund

2008; Abrevaya 2009; Egan et al. 2011). Therefore, it is important to further evaluate the DID effect of the Illinois and Pennsylvania bans at higher birth parities in comparison with lower birth parities. To this end, we estimate a triple difference (difference-in-difference-in-difference) model by modifying equation (1) in the following way. First, we drop the live birth order variable from X, and instead include a binary indicator of high birth parity for the i-th child in the g-th state and time t as below:

$$High_{igt}\begin{cases}=1 & \text{if the live birth order is} \geq 3\\ =0 & \text{if the live birth order is } 1 \text{ or } 2\end{cases}$$

Then, we fully interact this high parity indicator with all other explanatory variables of the regression. The coefficient of the interaction term ($treat_g * post_t * high_{igt}$) in essence will estimate the triple difference (TD) effect, i.e., the differential impact of the ban on higher birth parities (as compared with lower parities) in the treatment state as compared with the control state group.

Both the DID and TD regression models are estimated separately for Chinese and other Asian or Pacific Islander children. Standard errors are clustered at the state level and corrected using the wild bootstrap (1000 repetitions) method (Cameron et al. 2008).

3.3 Robustness Checks, Matching, and Synthetic Controls

We test the robustness of our results in two different ways. First, we use two alternative econometric techniques to estimate the impact of the ban – the matching methods and synthetic control approach – and compare them with the outcomes of the TD linear model. Second, we test the linear model with a series of placebo TD analyses. The results from our original TD models will be considered robust if they do not differ significantly from the outcomes of these additional analyses. These methods are described below.

There remains a possibility that newborn children in a state with the ban are systematically different in characteristics from those in the neighboring states. Such differences may bias the outcomes of the DID or the triple difference models. Propensity score matching methods (Rosenbaum and Rubin 1983; Heckman et al. 1997; Dehejia and Wahba 1998) can reduce such disparities between treatment and control observations by matching on the estimated propensity of the treatment. We check the robustness of our results by using an alternative propensity score matching (PSM) methodology to evaluate the effect of the bans in Illinois and Pennsylvania. First, for all newborn children in the 1978 to 1983 pre-ban data,

we estimate the propensity score of the birth happening in Illinois as compared with the neighboring states (with X as covariates). Then, Illinois newborns are matched with similar newborns in neighboring states on the basis of the estimated propensity score. The average difference in the probability of being male between the matched Illinois and neighboring state observations will therefore be the average treatment effect on the treated (ATT_{pre}) in the pre-ban period. We estimate a similar ATT_{post} between propensity-score-matched Illinois and neighbor state observations in the post-ban 1984 to 1989 data. The difference ($ATT_{post} - ATT_{pre}$) is the DID PSM effect of the ban in Illinois. Similarly, we also estimate the DID PSM effect of the ban in PA, ($ATT^{PA}_{post} - ATT^{PA}_{pre}$), using data from 1983 to 1994.[9]

Furthermore, we check the robustness of our results by using the generalized DID methodology of synthetic controls (Abadie and Gardeazabal 2003). In this analysis, the outcomes of a treatment state are not compared with the outcomes in neighboring states, but with a single "synthetic" control state. The control state is a convex combination of other states, such that its characteristics are similar to the treatment state. In order to construct the synthetic control, we include 48 other states and Washington, D.C. in our analysis. The analysis for Illinois excludes Pennsylvania from the comparison states, and the analysis of Pennsylvania excludes Illinois. We create balanced state-year panel datasets with the mean values of the outcome and X variables. Two separate datasets are created: one over 1978 to 1989 for Illinois and all other states (except Pennsylvania), and another over 1983 to 1994 for Pennsylvania and all other states (except Illinois). We then report the DID effect of the ban in a treatment state in comparison with its synthetic control created in the data. For simplicity, we do not separately conduct the TD exercise (for higher parities) with this method.[10]

Finally, we run a series of placebo TD linear regression models. This is required to evaluate possible confounding factors that may bias our results, e.g., other public policies that may have coincided with the implementation of the ban. Our placebo models consider hypothetical scenarios where instead of year 1984 in Illinois and 1989 in Pennsylvania, the bans are assumed to be implemented in a different year. We consider four placebo years of ban implementation for each state group – years 1982, 1983, 1985 and 1986 for Illinois, and years 1987, 1988, 1990 and 1991 for Pennsylvania. Triple difference regressions separately for

9 We use a 5-nearest-neighbor matching (with replacement) method. Standard errors are bias-corrected (Abadie and Imbens 2006). For simplicity, we do not conduct TD analysis (for higher parities) using PSM.

10 We use the Stata programs *synth.ado* for the synthetic control analysis, *psmatch2.ado* for the PSM analysis, and *cgmwildboot.ado* for the linear models.

Chinese and other Asian or Pacific Islander children are then run under each of these scenarios. For example, we consider a scenario in which the sex-selective abortion ban is implemented in Illinois in 1982. Then, the period 1978–1981 in our data is considered to be the pre-ban period, and 1982–1989 is considered post-ban. The outcomes in Illinois are evaluated between the post- and pre-ban periods, and also in comparison with the neighboring states.

To account for the fact that NCHS data includes both foreign-born and US-born Asian parents, we conduct further tests to ensure our analysis has enough statistical power to detect a small change in sex ratios at birth. Because the U.S. census indicates that as many as 70% of the Asians are foreign-born, our test assumes that only 70% of the Chinese and other Asian or Pacific Islander newborn children in our analysis had foreign-born parents (see footnote 8). Therefore, we consider samples sizes in Illinois, Pennsylvania and their neighboring state groups that are 30% smaller than those mentioned in Appendix Table 5. Then, we test whether this smaller sample will produce enough statistical power for a predetermined effect size for the law.

Since there is no current benchmark effect size for sex-selective abortion bans in the United States or a similar socioeconomic context,[11] we simulate absolute effect sizes (reduction in the probability of a newborn child being male) in the range of 0 to 0.1, with 0.001 increments. In order to avoid complexities related to a small number of clusters in our data, we consider a simple calculation of statistical power without clustering the standard errors within states. The statistical power estimates from the simulations (with a 5% level of significance) are presented in Appendix Figures 1 and 2.

We find that with a power of 0.8 the 70% samples in Illinois and its neighboring states are adequate for detecting changes in the probability of a boy birth among Chinese and other Asian or Pacific Islander children that are as small as 0.028 and 0.011, respectively. Similarly, the smallest effect sizes that can be detected by the 70% samples with a 0.8 power in Pennsylvania and its neighboring states are 0.028 and 0.012 among Chinese and other Asian or Pacific Islander newborn children, respectively. These effect sizes are similar to, or lower than, our estimates shown in Table 2. Also, the impact of the bans on sex-selective abortion in Illinois and Pennsylvania among children of foreign-born parents, if any, are likely to be higher than the effects seen in the overall sample, which includes

11 A recent study by Nandi and Deolalikar 2013 present some possible effect sizes of a ban on pre-natal sex determination in India. However, the socioeconomic and cultural context of this study, along with the nature of the law and its enforcement are completely different from our present study. Therefore, it does not provide a comparable benchmark.

both foreign-born and U.S.-born parents. Furthermore, any estimated effect size that is below the threshold minimum effect size produced by our sample will be of little importance, even if it is statistically significant (Ziliak and McCloskey 2007).

4 Results

First, we present some trend analysis using annual NCHS data in Figures 3 and 4 in the Appendix. We analyze sex ratios among newborn Chinese and other Asian or Pacific Islander children to evaluate the so-called "parallel trends assumption," i.e., similar trends in the treatment and control states leading to the implementation of the sex-selective abortion ban. In addition, if the ban has any discernable impact, we should also observe a break in the trend for the treatment state from the pre-ban to the post-ban period. Since annual data on sex ratios at birth suffer from a large variance, we present a smoother 3-year moving average of sex ratios over time. We use data from an extended period during 1979 to 1999. It should be noted that during this period (in 1993) the law prohibiting sex-selective abortion in Illinois was declared partially unenforceable by a federal court.[12] Criminal sanctions were no longer allowed under the law for pre-viability abortions. However, civil penalties (such as fines) continue to be available for pre-viability sex-selective abortions and both criminal and civil penalties are available for post-viability abortions under the law.

We find that the parallel trends assumption is generally satisfied. We also find no significant changes in the trend of sex ratios at birth in Illinois or Pennsylvania after the bans on sex-selective abortion were enacted. However, the trend lines of the Chinese newborn sex ratios are less conclusive because of fluctuations in the data due to small sample sizes.[13]

Next, results from our pooled linear probability regression models of the odds-ratio of a boy birth are presented in Table 1. From the regression of Illinois and its neighboring states during 1979 to 1989, we find that Chinese and other Asian or Pacific Islander children are no more statistically likely to be male as compared with children of other races, controlling for parental education, live birth order of the child, and state and year fixed-effects. The results are similar when we separately analyze children of 2nd or higher birth orders.

12 *Herbst v. O'Malley*, 1995 WL 55252 (not reported in F. Supp.).

13 Table 5 in the Appendix provides the annual sample sizes of newborn children across different population groups in our study.

Table 1 Pooled Linear Regressions of the Probability of a Boy Birth in Illinois, Pennsylvania and their Neighboring States.

	Illinois and neighboring states			Pennsylvania and neighboring states		
	All children	2nd or higher birth order children	3rd or higher birth order children	All children	2nd or higher birth order children	3rd or higher birth order children
Chinese parents	0.001	0.008	0.002	0.006**	0.008**	0.020**
	(−0.006)	(−0.009)	(−0.016)	(−0.002)	(−0.003)	(−0.006)
Other Asian-PI parents	0.001	0.000	0.000	0.005**	0.008**	0.016**
	(−0.002)	(−0.003)	(−0.005)	(−0.001)	(−0.002)	(−0.003)
Live birth order	−0.001**	−0.001**	−0.001**	−0.001**	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Mother: educated but less than high school	−0.007**	−0.009**	−0.012**	−0.008**	−0.010**	−0.011**
	(−0.003)	(−0.003)	(−0.005)	(−0.001)	(−0.001)	(−0.002)
Mother: high school but less than college	−0.006**	−0.008**	−0.010**	−0.007**	−0.009**	−0.009**
	(−0.003)	(−0.003)	(−0.005)	(−0.001)	(−0.001)	(−0.002)
Mother: completed college or higher	−0.004	−0.007*	−0.008	−0.007**	−0.007**	−0.006**
	(−0.003)	(−0.004)	(−0.005)	(−0.001)	(−0.001)	(−0.002)
Father: educated but less than high school	0.004**	0.005**	0.006**	0.007**	0.009**	0.009**
	(−0.001)	(−0.001)	(−0.002)	(−0.001)	(−0.001)	(−0.001)
Father: high school but less than college	0.004**	0.005**	0.005**	0.006**	0.008**	0.008**
	(−0.001)	(−0.001)	(−0.001)	(−0.001)	(−0.001)	(−0.001)
Father: completed college or higher	0.006**	0.006**	0.005**	0.006**	0.009**	0.009**
	(−0.001)	(−0.001)	(−0.002)	(−0.001)	(−0.001)	(−0.002)
Constant term	0.517**	0.518**	0.523**	0.516**	0.513**	0.514**
Sample size	6,092,268	3,605,336	1,581,269	9,674,889	5,569,421	2,409,697
p-value of F-statistic	0	0	0	0	0	0

Data are from NCHS 1978–1989 for Illinois and its neighboring states, and 1983–1994 for Pennsylvania and its neighboring states. Regression coefficients that are significant at 10% and 5% levels are respectively marked by * and **. Standard errors are heteroskedasticity-robust, and presented in the parentheses. Fathers and mothers with no formal education are considered as the excluded categories.

However, from the same analysis of Pennsylvania and its neighboring states during 1984 to 1994, we find that Chinese and other Asian or Pacific Islander children are slightly more likely to be male as compared to children of parents of other races, controlling for the covariates. The likelihood of male birth in these race groups increases with parity: among children of the 3rd or higher birth order, Chinese and other Asian or Pacific Islander children are 2% and 1.6% more likely to be male, respectively. However, the sample sizes of higher birth order Chinese children are generally low in our study and the results could be distorted as a result.

Results from the TD linear regression, PSM method and synthetic control analysis are presented in Table 2. For a concise presentation, we only show the main DID parameter of interest – change in the probability of a boy birth in the treatment state (i.e., Illinois or Pennsylvania) from the pre-ban to the post-ban period, as compared with the change over the same period of time among its neighboring control states without abortion bans. In addition, the TD parameter (impact of the law among children of 3rd or higher birth parities) is also shown in the case of linear models. The outcomes of the synthetic control analysis are

Table 2 Triple Difference Linear Model, Propensity Score Matching and Synthetic Control Analysis of the Probability of a Boy Birth among Chinese and other Asian-Pacific Islanders.

	Children born in IL and neighboring states		Children born in PA and neighboring states	
	Chinese	**Other Asian-PI**	**Chinese**	**Other Asian-PI**
TD linear model coefficients:				
DID effect of law ($treat_g * post_t$)	−0.040	−0.008	−0.010	0.004
TD effect ($treat_g * post_t * high_{igt}$)	0.057	−0.009	0.014	0.018
PSM DID effect of law (ATT_{post}–ATT_{pre})	−0.195	−0.005	0.035	0.153
DID effect of law (synthetic control)	0.000	0.018	−0.036	−0.003

Data are from NCHS 1978–1989 for Illinois and its neighboring states, and 1983–1994 for Pennsylvania and its neighboring states. Estimators that are significant at 10% and 5% levels are respectively marked by * and **. In the synthetic control analysis, all 50 states and District of Columbia are included (except Illinois is excluded from Pennsylvania analysis, and Pennsylvania is excluded from Illinois analysis). However, a few states are dropped (missing values of state-year) in order to create a balanced panel required for the synthetic control analysis. The propensity score matching method uses 5-nearest-neighbor matching with replacement, and the standard errors are bias corrected (Abadie and Imbens 2006). Standard errors of the linear models are corrected using the wild bootstrap method (Cameron et al. 2008), 1000 repetitions with null imposed. TD=triple difference; DID=difference-in-difference; PSM=propensity score matching.

shown as trend graphs in Figures 5 and 6 in the Appendix. More detailed results are available from the authors on request.

The DID estimates from all three econometric methods are statistically insignificant at the 10% level or lower. This indicates that bans on sex-selective abortion in Illinois and Pennsylvania were not associated with any changes in sex ratios at birth of Chinese and other Asian or Pacific Islander children from the pre-ban to the post-ban periods. The TD regression coefficient for higher birth parities is also insignificant in all cases, indicating that the ban did not have any further effect on children of higher birth parities.

Results from placebo TD linear regressions are presented in Table 6 in the Appendix. We find that the hypothetical bans on sex-selective abortion in these additional years are also not associated with any changes in the sex ratios at birth among Chinese and other Asian or Pacific Islander children from the pre-ban to the post-ban periods. The TD coefficients are also generally insignificant, showing no additional effect of the ban among children at higher birth parities.[14]

In sum, our findings indicate that sex-selective abortion bans in Illinois and Pennsylvania are not associated with any significant changes in sex ratios at birth of Chinese and other Asian or Pacific Islander children in the two states.

5 Discussion and Conclusion

Since Almond and Edlund (2008) published a study finding male-biased sex ratios at higher birth parities among the children of foreign-born Chinese, Indian and Korean parents, six states have banned sex-selective abortion and 21 states and U.S. Congress have considered similar bans. Prior to our study, no other empirical analysis has been conducted to determine whether laws banning sex-selective abortion impact sex ratios at birth. We used annual birth data from the National Center for Health Statistics from 1978 to 1994 to determine if legal prohibitions on sex-selective abortion enacted in Illinois in 1984 and Pennsylvania in 1989 were associated with any changes in sex ratios at birth among Asians in those states.

Even though NCHS data does not distinguish between births of foreign-born and U.S.-born Asians, our results remain highly relevant to policy debates in the United States given that many believe both foreign-born and U.S.-born Asians

14 In a few cases, the DID or TD coefficients are marginally significant at 10% level. However, the magnitude of these effects are very small. Also, since there is no definite patterns, these idiosyncratic results are very likely due to small sample sizes.

sex select against girls. For example, the federal Prenatal Nondiscrimination Act (PRENDA) (H.R. 3541) that would ban sex-selective abortion throughout the United States claims: "The evidence strongly suggests that some Americans are exercising sex-selection abortion practices within the US consistent with discriminatory practices common to… the country to which they trace their ancestry."[15]

We find that legal prohibitions on sex-selective abortion enacted in Illinois and Pennsylvania are not associated with any changes in sex ratios at birth among Asians in the 6-year period after the bans were enacted in each state. In Illinois and its neighboring states, the sex ratio at birth of Asian children was not male-biased during our study period. On the other hand, while the sex ratio at birth among Asians in Pennsylvania and its neighboring states was skewed slightly in favor of boys, the enactment of the ban did not normalize the ratio. This strongly suggests that sex-selective abortion bans have had no impact on the practice of sex selection, to the extent that it occurred, during the study period in the two states.

There may be several reasons why sex-selective abortion bans are not associated with changes in sex ratios at birth. The bans only prohibit sex selection through abortion; there are no prohibitions on the use of pre-implantation reproductive technologies, such as sperm sorting and genetic diagnosis, for the purposes of sex selection. As a result, people may be using these techniques to sex select. The bans may also be hard to enforce due to the difficulty in determining a woman's reason for seeking an abortion.

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15 Source: The US Library of Congress, http://thomas.loc.gov/cgi-bin/query/z?c112:H.R.3541.

Appendix

Table 3 Raw Estimates of the Male-to-Female Sex Ratios of Chinese and other Asian Pacific Islander Newborn Children in Illinois and Neighboring States.

	Illinois		Neighboring states of Illinois	
	Pre-law period (1978–1983)	Post-law period (1984–1989)	Pre-law period (1978–1983)	Post-law period (1984–1989)
Chinese newborn:				
Male-to-female sex ratio	1.062	1.071	0.986	1.126
Number of children	1994	1630	1301	1348
Other Asian-PI newborn:				
Male-to-female sex ratio	1.055	1.082	1.012	1.071
Number of children	10,152	13,539	7103	11,351

Data are from NCHS 1978–1989 for Illinois and its neighboring states.

Table 4 Raw Estimates of the Male-to-Female Sex Ratios of Chinese and other Asian Pacific Islander Newborn Children in Pennsylvania and Neighboring States.

	Pennsylvania		Neighboring states of Pennsylvania	
	Pre-law period (1983–1988)	Post-law period (1989–1994)	Pre-law period (1983–1988)	Post-law period (1989–1994)
Chinese newborn:				
Male-to-female sex ratio	1.115	1.063	1.080	1.072
Number of children	1379	2288	20,326	29,137
Other Asian-PI newborn:				
Male-to-female sex ratio	1.068	1.090	1.086	1.066
Number of children	7480	10,147	40,100	64,781

Data are from NCHS 1983–1994 for Pennsylvania and its neighboring states.

Table 5 Sample Sizes of Chinese and other Asian Pacific Islander Newborn Children in Illinois, Pennsylvania and their Neighboring States.

	Chinese		Other Asian-PI	
Illinois and neighboring states	**IL**	**Neighboring states of IL**	**IL**	**Neighboring states of IL**
1978	310	220	1238	674
1979	358	195	1410	829
1980	334	232	1745	1138
1981	338	239	1996	1432
1982	309	207	2011	1507
1983	345	208	1752	1523
1984	399	195	1497	1637
1985	247	191	1574	1737
1986	233	189	1487	1913
1987	231	219	2776	1849
1988	286	239	3121	2037
1989	234	315	3084	2178
Pennsylvania and neighboring states	**PA**	**Neighboring states of PA**	**PA**	**Neighboring states of PA**
1983	219	2681	1141	5299
1984	201	2865	1112	5434
1985	193	3198	1163	5975
1986	208	3336	1217	6631
1987	255	3758	1397	7564
1988	303	4488	1450	9197
1989	295	4633	1460	9403
1990	372	4811	1613	10,909
1991	382	4301	1644	11,393
1992	374	4810	1762	10,866
1993	447	5055	1820	11,138
1994	418	5527	1848	11,072

Data are from NCHS 1978–1994.

Table 6 Placebo Triple Difference Linear Models of the Probability of a Boy Birth.

	Chinese	Other Asian-PI
Illinois and neighboring states		
Hypothetical 1982 law:		
DID effect of law ($treat_g * post_t$)	−0.043	−0.018
TD effect ($treat_g * post_t * high_{igt}$)	0.083	0.018
Hypothetical 1983 law:		
DID effect of law ($treat_g * post_t$)	−0.043	−0.007
TD effect ($treat_g * post_t * high_{igt}$)	0.053	0.010
Hypothetical 1985 law:		
DID effect of law ($treat_g * post_t$)	−0.005	−0.015
TD effect ($treat_g * post_t * high_{igt}$)	0.048	−0.009
Hypothetical 1986 law:		
DID effect of law ($treat_g * post_t$)	0.009	−0.022*
TD effect ($treat_g * post_t * high_{igt}$)	0.063	0.009
Pennsylvania and neighboring states		
Hypothetical 1987 law:		
DID effect of law ($treat_g * post_t$)	0.008*	0.009
TD effect ($treat_g * post_t * high_{igt}$)	−0.035	0.021
Hypothetical 1988 law:		
DID effect of law ($treat_g * post_t$)	−0.008*	0.008
TD effect ($treat_g * post_t * high_{igt}$)	−0.024	0.014
Hypothetical 1990 law:		
DID effect of law ($treat_g * post_t$)	−0.016	0.001
TD effect ($treat_g * post_t * high_{igt}$)	0.025	0.026
Hypothetical 1991 law:		
DID effect of law ($treat_g * post_t$)	−0.023	0.000
TD effect ($treat_g * post_t * high_{igt}$)	0.033*	0.021

Data are from NCHS 1978–1989 for Illinois and its neighboring states, and 1983–1994 for Pennsylvania and its neighboring states. Estimators that are significant at 10% and 5% levels are respectively marked by * and **. All standard errors are corrected using the wild bootstrap method (Cameron et al. 2008), 1000 repetitions with null imposed. TD=triple difference; DID=difference-in-difference.

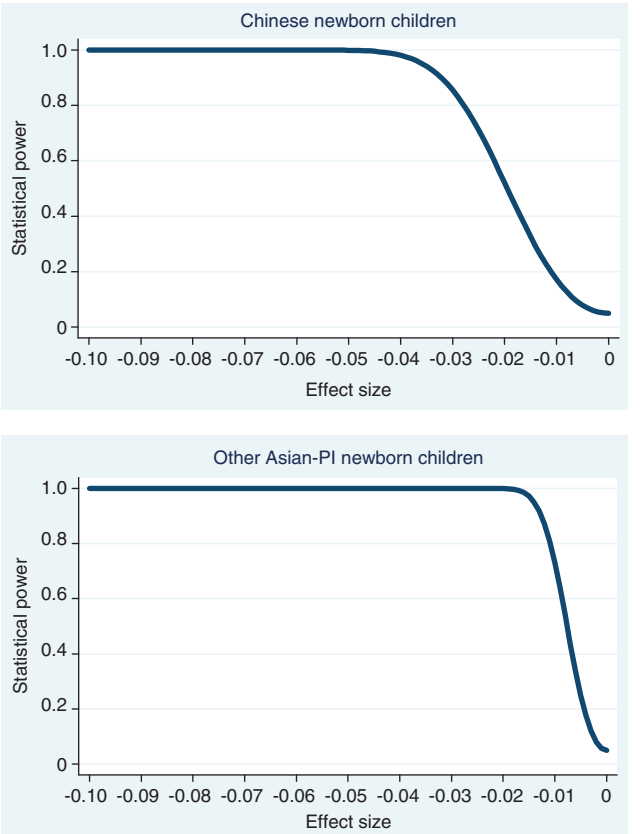

Figure 1 Statistical Power of Simulated Effect Sizes of the Ban on Sex-selective Abortion among Illinois and its Neighboring States (1978 to 1989), 70% Sample.
Source: Effect size is the absolute reduction in the probability of a newborn child being male, with a 5% level of significance. We use the *samplesize* command in Stata, with the *sampsi_reg* option, and actual standard deviations of the proportion of male child in Illinois and its neighboring states.

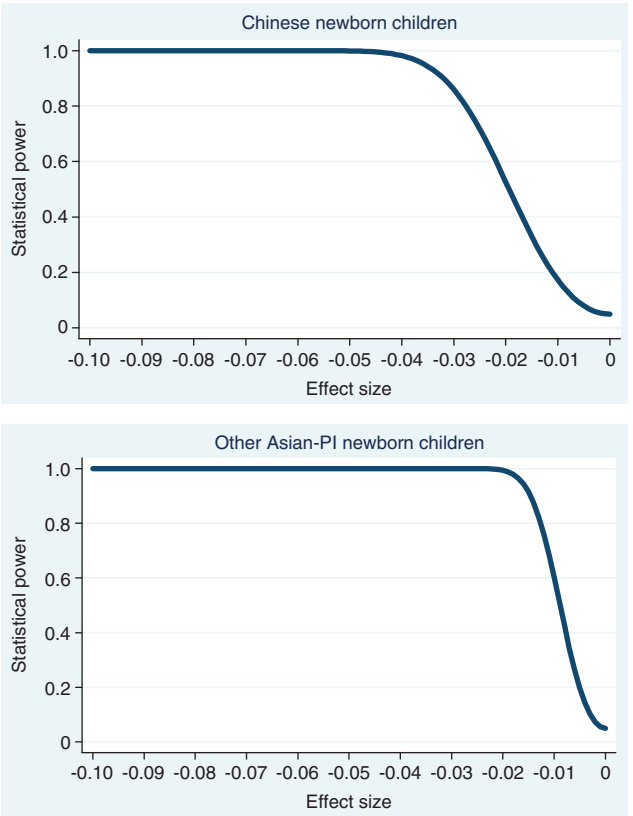

Figure 2 Statistical Power of Simulated Effect Sizes of the Ban on Sex-selective Abortion among Pennsylvania and its Neighboring States (1983 to 1994), 70% Sample.
Source: Effect size is the absolute reduction in the probability of a newborn child being male, with a 5% level of significance. We use the *samplesize* command in Stata, with the *sampsi_reg* option, and actual standard deviations of the proportion of male child in Pennsylvania and its neighboring states.

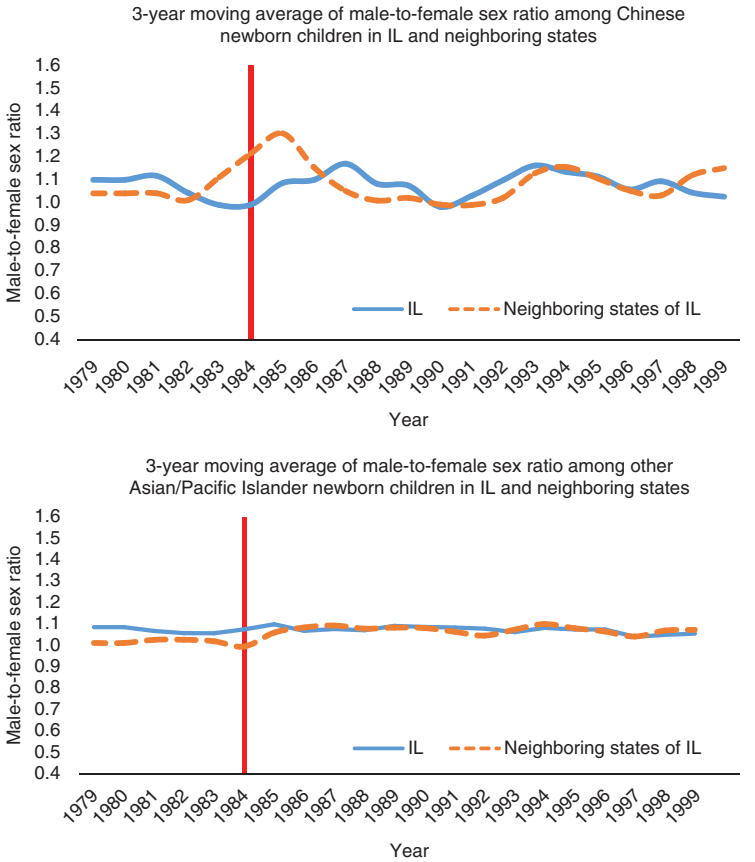

Figure 3 Raw Sex Ratio at Birth among Certain Ethnicities in Illinois and Neighboring States (1979 to 1999).
Source: Estimated from NCHS Data 1978–2000. The vertical line in each graph marks the year of the ban on sex-selective abortions in Illinois (1984). Each data point is a 3-year average of the year and its preceding and succeeding years.

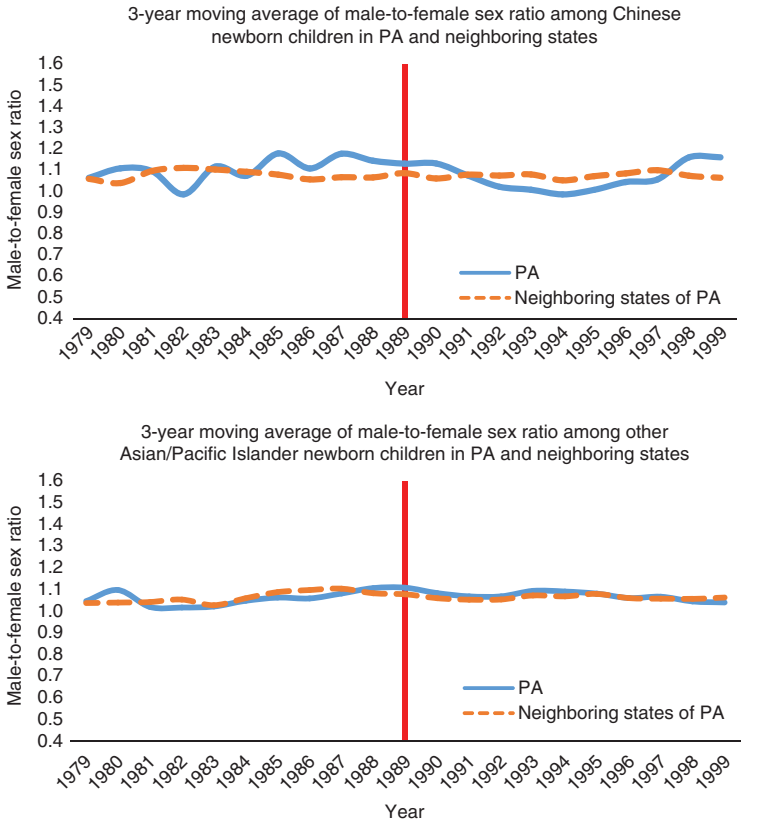

Figure 4 Raw Sex Ratio at Birth among Certain Ethnicities in Pennsylvania and Neighboring States (1979 to 1999).

Source: Estimated from NCHS Data 1978–2000. The vertical line in each graph marks the year of the ban on sex-selective abortions in Pennsylvania (1989). Each data point is a 3-year average of the year and its preceding and succeeding years.

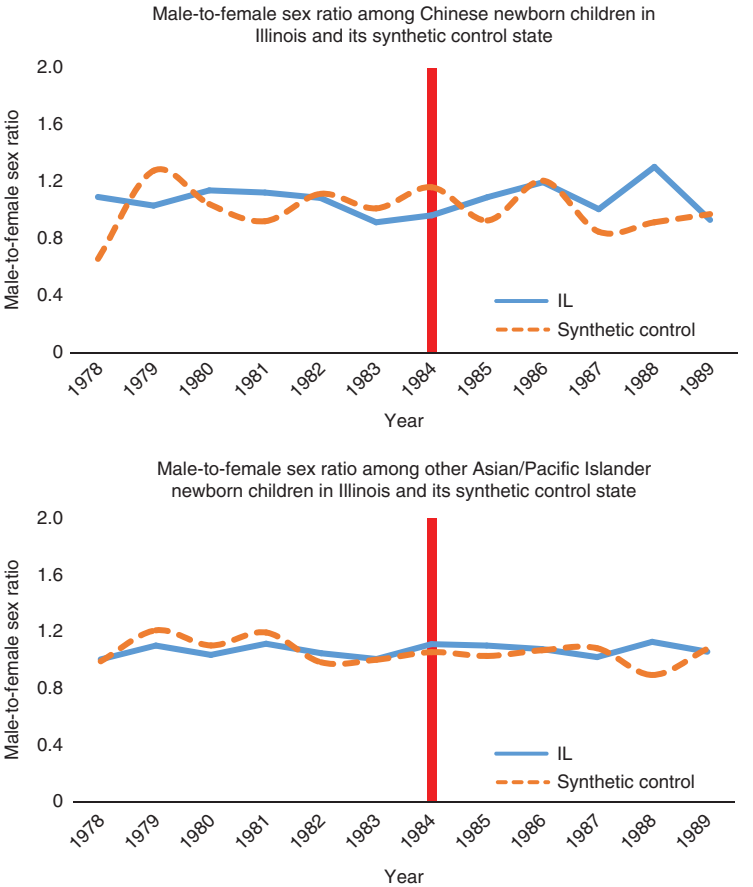

Figure 5 Sex Ratio at Birth among Certain Ethnicities in Illinois and its Synthetic Control State (1978 to 1989).
Source: Estimated from NCHS Data 1978–1989. The vertical line in each graph marks the year of the ban on sex-selective abortions in Illinois (1983). The synthetic control state is created from all 50 states and Washington DC (except Pennsylvania).

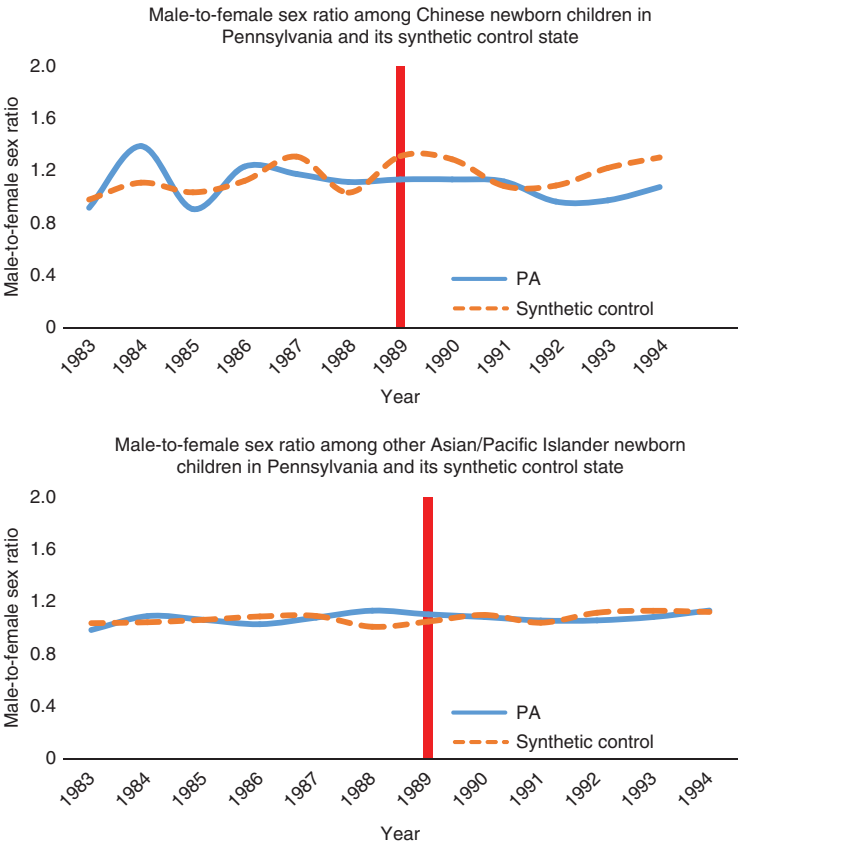

Figure 6 Sex Ratio at Birth among Certain Ethnicities in Pennsylvania and its Synthetic Control State (1983 to 1994).
Source: Estimated from NCHS Data 1983–1994. The vertical line in each graph marks the year of the ban on sex-selective abortions in Pennsylvania (1989). The synthetic control state is created from all 50 states and Washington DC (except Illinois).

References

Abadie, A., and J. Gardeazabal (2003) "The Economic Costs of Conflict: A Case Study of the Basque Country," The American Economic Review, 93:113–132.

Abadie, A., and G. W. Imbens (2006) "Large Sample Properties of Matching Estimators for Average Treatment Effects," Econometrica, 74:235–267.

Abrevaya, J. (2009) "Are There Missing Girls in the United States? Evidence from Birth Data," American Economic Journal: Applied Economics, 1:1–34.

Almond, D., and L. Edlund (2008) "Son-Biased Sex Ratios in the 2000 United States Census," Proceedings of the National Academy of Science USA, 105:5681–5682.

Cameron, A., J. Gelbach, and D. Miller (2008) "Bootstrap-based Improvements for Inference with Clustered Errors," The Review of Economics and Statistics, 90:414–427.

Das Gupta, M., and L. Shuzhuo (1999) "Gender Bias in China, South Korea and India 1920–1990: Effects of War, Famine and Fertility Decline," Development and Change, 30:619–652.

Das Gupta, M., J. Zhenghua, X. Z. Li Bohua, B. H. Woojin Chung, L. Bohua, X. Zhenming, W. Chung, and B. Hwa-Ok (2003) "Why is Son Preference so Persistent in East and South Asia? A Cross-Country Study of China, India, and the Republic of Korea," The Journal of Development Studies, 40:153–187. doi:10.1080/00220380412331293807.

Dehejia, R. H., and S. Wahba (1998) "Causal Effects in Non-Experimental Studies: Re-Evaluating the Evaluation of Training Programs," Journal of the American Statistical Association, 94:1–053.

Ebenstein, A. (2010) "The 'Missing Girls' of China and the Unintended Consequences of the One Child Policy," Journal of Human Resources, 45:87–115.

Egan, J. F. X., W. A. Campbell, A. Chapman, A. A. Shamshirsaz, P. Gurram, and P. A. Benn (2011) "Distortions of Sex Ratios at Birth in the United States; Evidence for Prenatal Gender Selection," Prenatal Diagnosis, 31:560–565.

Heckman, J. J., H. Ichimura, and P. E. Todd (1997) "Matching as an Econometric Evaluation Estimator: Evidence from Evaluating a Job Training Programme," Review of Economic Studies, 64:605–654.

Hollmann, F. W. (1990) United States Population Estimates, by Age, Sex, Race, and Hispanic Origin: 1980 to 1988. Current Population Reports, Series P-25, No. 1045, US Census Bureau.

Jha, P., M. A. Kesler, R. Kumar, F. Ram, U. Ram, L. Aleksandrowicz, D. G. Bassani, S. Chandra, and J. K. Banthia (2011) "Trends in Selective Abortions of Girls in India: Analysis of Nationally Representative Birth Histories from 1990 to 2005 and Census Data from 1991 to 2011," Lancet, 377:1921–1928.

Kandel, W. A. (2011) *The U.S. Foreign-Born Population: Trends and Selected Characteristics*. Washington DC: Congressional Research Service.

Lee, S. M. (1998) "Asian Americans: Diverse and Growing," Population Bulletin, 53:1–44.

Lin, M.-J., N. Qian, and J.-T. Liu (2008) More Women Missing, Fewer Girls Dying: The Impact of Abortion on Sex Ratios at Birth and Excess Female Mortality in Taiwan (No. w14541). National Bureau of Economic Research.

Lopez, M. H., P. Taylor, C. Funk, and A. Gonzalez-Barrera (2013) *On Immigration Policy, Deportation Relief Seen As More Important Than Citizenship: A Survey of Hispanics and Asian Americans*. Washington DC: Pew Research Center.

McDonough, K. (2013) *Sex-Selective Abortion Bans: Just Another Right-Wing Ruse*. Salon. http://www.salon.com/2013/08/19/sex_selective_abortion_bans_just_another_right_wing_ruse/. Accessed November 11, 2014.

Nandi, A., and A. B. Deolalikar (2013) "Does a Legal Ban on Sex-Selective Abortions Improve Child Sex Ratios? Evidence from a Policy Change in India," Journal of Development Economics, 103:216–228. doi:10.1016/j.jdeveco.2013.02.007.

Park, C. Bin, and N.-H. Cho (1995) "Consequences of Son Preference in a Low-Fertility Society: Imbalance of the Sex Ratio at Birth in Korea," Population and Development Review, 21:59–84.

Puri, S., V. Adams, S. Ivey, and R. D. Nachtigall (2011) " 'There is Such a Thing as Too Many Daughters, but not Too Many Sons': A Qualitative Study of Son Preference and Fetal Sex Selection Among Indian Immigrants in the United States," Social Science and Medicine, 72:1169–1176.

Rosenbaum, P. R., and D. B. Rubin (1983) "The Central Role of the Propensity Score in Observational Studies for Causal Effects," Biometrika, 70:41–55.

Sen, A. (2003) "Missing Women – Revisited," British Medical Journal, 327:1297–1298. doi:10.1136/bmj.327.7427.1297.

Zhu, W. X., L. L. Lu, and T. Hesketh (2009) "China's Excess Males, Sex Selective Abortion, and One Child Policy: Analysis of Data From 2005 National Intercensus Survey," British Medical Journal 338:b1211–b1211. doi:10.1136/bmj.b1211.

Ziliak, S. T., and D. N. McCloskey (2007) *The Cult of Statistical Significance: How the Standard Error Costs Us Jobs, Justice, and Lives*. Ann Arbor, MI: University of Michigan Press.