The London School of Economics and Political Science

Essays on Microeconomic Incentives in Public Policies

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Declaration

I certify that the thesis I have presented for examination for the MRes/PhD degree of the London School of Economics and Political Science is solely my own work other than where I have clearly indicated that it is the work of others (in which case the extent of any work carried out jointly by me and any other person is clearly identified in it).

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Statement of Conjoint Work

I confirm that Chapter 2 is jointly co-authored with S Roy, all authors contributed equally to the project. Chapter 3 is jointly co-authored with S Anukriti and Sonia Bhalotra, all authors contributed equally to the project.

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Abstract

This thesis consists of three chapters on public and development economics that study incentives in public policies from the perspective of microeconomics. It is devoted to understanding behavioural responses in public policies that are relevant to its design in each domain, including trading in market with frictions, enrolment in education and child marriage practice for young girls, and childcare resource allocation in family.

Chapter 1 studies how transaction tax policy affect market with frictions. Transaction tax in property market, with tax rate decreasing in holding period of property, received attention from governments in Asia for moderating speculation since 2000s. Using administrative transaction record of property, this chapter studies the behavioural response to the transaction tax in the timing of transaction, tax incidence and selection of buyers in Hong Kong and Singapore. I find that the inherent tax incentives, in the form of tax notches, induced tax avoidance behaviour in the timing of transaction, and average buyer and seller are willing to wait 3-4 weeks to avoid 1% of transaction tax. Exploiting discontinuity in tax liability at daily level, I find that the tax policy has impact on transaction activity that links closely to its rate and lower the overall chance of a property sold. Buyers bear significant tax burden on seller specific tax even when tax-free sellers are abundant in the market both evidence suggest strong search friction in property market. I also find that the differential tax rate in holding duration produce selection effect among buyers with different ex ante probability of trade in the taxable holding period. This chapter contributes to understanding the nature of transaction tax in markets with search friction.

Chapter 2, a joint work with S Roy, studies the impact of matrimonial laws introduced by the British in colonial India during 1800s and early 1900s. Legal reforms on marriage practices, including laws on minimum marriage age and female infanticide, were introduced in British Provinces - district that were under British direct rule. Exploiting quasi-random variations of districts that were former British Provinces within each post-independent Indian states, this chapter studies their

impact on female education and under age marriages in post-Independent India. From independent sources of large-scale micro data, including administrative records from schools and representative household surveys, we find that in former British Provinces females have 5% lower chances of marrying under the current legal age, and 1.6% higher chance of attending school at 10-16 years old. Child Marriage abolition Act was introduced in 1931, which raised the minimum age of marriage for female to 14. With newly digitized data on district level marriage pattern from Census of India 1901-1931, we find that the act distorts the marriage market in the short-run by increasing the likelihood of girls marrying at young age as it was pre-announced before its implementation, while district more aware of the law exhibit lower child marriage in the long run. It suggest that expansion of education for girls in India has demand side constraints from child marriage practice that has historical root.

The introduction of prenatal sex-detection technologies in India has led to a phenomenal increase in abortion of female fetuses. Chapter 3, a joint work with S Anukriti and Sonia Bhalotra investigates their impact on the relative chances of girls surviving after birth, fertility and parental investments. We find that it lead to reduction in excess female mortality, erosion of gender gaps in parental inputs such as breastfeeding and immunization, and moderation of son-stopping fertility. For every five aborted girls, we estimate that roughly one additional girl survives to age five. Our findings have implications that sex-selective abortion not only account for counts of missing girls but also for the later life outcomes of girls.

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Chapter 1

Behavioural response to time notches in transaction tax: Evidence from stamp duty in Hong Kong and Singapore

To moderate speculation in housing market, multiple Asian cities implemented transaction tax notches on holding period of property. Using administrative transaction record of property trading, this paper studies the behavioural response in the timing of transaction, tax incidence and selection of buyers using the policy changes in Hong Kong and Singapore. Tax notches on holding period generate significant tax avoidance bunching in the timing of transaction, and properties were less likely to be resold even one year after the tax last applies, suggesting plausible crowd out of transactions. I construct and use a new dataset on government estimated rental rate to estimate the tax incidence and find evidence that buyers bear significant tax burden even when tax-free alternatives are available in the market. Evidence shows that time notches on holding duration produce selection effect among buyers with different ex ante probability of trade in the taxable holding period. Estimates suggest that traders on average are willing to wait for 3-4 weeks to avoid 1% of transaction tax, and each week of delay in transaction would generate loss in the trading surplus at 0.3 % of property value.

1.1 Introduction

Transaction tax on short holding duration of real estate properties has become a popular policy tool for governments in east and south east Asia since 2010. Hong

Kong, Singapore, Macau, Taiwan and China introduced transaction tax with rates that vary with holding duration to deter speculation in the housing market.[1] Typically, residential property transactions are subject to an additional tax from 4-16% of the transaction price if the property is re-sold within 2 to 4 years. Moreover, the tax rate declines with the time seller holds the property. One intriguing feature of these tax instruments is that they have a flat tax rate for an extended period of time (e.g. 6 months), which then sharply decline once the time threshold is passed and thus constitute salient time notches. While transaction tax in the property market is present in many countries, there is very little evidence exist in the public finance literature for how would the market reacts to transaction tax and particularly those that are designed with objectives beyond raising tax revenue. This paper provide empirical evidence on understanding how the market react to transaction tax that have explicit inter-temporal design.

Is transaction tax on short holding duration an appropriate tool to stabilize the housing market? Speculators and middlemen re-sell faster than home-owners (Bayer et al., 2011), thus transaction tax on property with short holding duration may impose additional cost for speculators and create selection of buyers in the market. On the other hand, it distorts the trading behaviour for all traders including home-owners, by either the distinct tax avoidance incentives and the potential tax burden for new buyers. Recent studies in the UK shows that expected transaction tax increase of 1% in the property market could generate significant re-timing of transactions up to three weeks (Best and Kleven, 2015). The high transaction tax rate and sharp tax notches implemented in Asia may thus creates incentives for buyers and sellers to re-time their transaction, and discouraged welfare improving transactions if the cost of re-timing is too high.[2] In a way similar to the analysis of capital gain tax, properties could be "locked-in" by the tax, which could lead to equilibrium effect on transaction price where little is known empirically (Dai et al., 2008). Using comprehensive administrative record in property transactions of Hong Kong and Singapore, this paper analyses both the re-timing of transaction and extensive margin crowd out that conceptually constitute the two main sources of welfare cost of transaction tax with differential tax rate on holding duration.

[1] Hong Kong and Singapore introduced transaction tax that varies with holding duration in 2010; Macau and Taiwan had followed through in 2011.

[2] The design is similar to the long-term versus short-term capital gain tax in the U.S. (Shackelford and Verrecchia, 2002; Hurtt and Seida, 2004); Vermont had implemented capital gain tax on land with a similar declining tax rate with holding duration to deter speculation since 1970s. Daniels et al. (1986) argues that the capital gain tax on land had restricted supply and raise land price.; Ontorio had implemented land speculation tax in April 1974 on gains on property transfer within a 10 year horizon, and as the a fixed percentage of exemption is applied every full year, it created tax notches each year same as the tax notches analysed in this paper. (Smith, 1976)

The government of Hong Kong introduced the "Special Stamp Duty" on 20th November 2010, in which for any residential properties purchased thereafter and resold within 2 years, transaction tax rates of 5-15% were applicable. Using the administrative record of property transaction in Hong Kong from 2009-2015, I construct the full transaction history of residential properties from 2009-2015. Comparing properties that are subject to the "Special Stamp Duty" to those that are not, I provide graphical evidence on significant bunching in property transactions when properties were held for exactly 2 years as the tax rate drop from 5% to 0%. With an regression discontinuity design, I find that properties subject to the tax are less likely to be traded in the first three years since purchase by 13 percentage points, taking into account of re-timing of transaction. This gives evidence to the significant reduction in transaction at the extensive margin generated by the tax even when tax avoidance is feasible, which implies non-trivial cost of transaction re-timing in the housing market.

Contrasting cash buyers versus buyers who took a mortgage, I first provide evidence that cash buyers on average are more likely to sell their property in the first year by 4.6% before the reform. Using an event study setup, I find the introduction of Special Stamp Duty increase the percentage of buyer with mortgage by 5.5%, suggesting the Special Stamp Duty creates selection among the new buyers. The tax distorts the trading behaviour on both type of buyers at the intensive margin regarding tax-avoidance behaviour, while the extensive margin crowd-out is much larger among the cash buyers.

To estimate the tax incidence, I use difference-in-differences strategy comparing transactions carried out under the taxable holding period (before 2 years) to those sold after the taxable period, as well as its counterpart for properties that are not subject to the Special Stamp Duty at all. This controls for time-varying factors in the housing market that may be correlated with the execution time of the tax. To control for property heterogeneity, I construct a new data set on government estimated rental rate for properties being traded in 2009-2015. This makes sure the tax incidence is identified by comparing properties observationally equivalent but subject to different tax liability, a methodology that is close to Besley et al. (2014). I find that with transaction tax rate of 5% the buyer bear 4% of the burden, but from that on every 5 % additional tax buyer only bear 1%. The significant burden on buyer in a market where tax-free alternative is available suggests that buyer's property specific valuation is quantitatively important in housing transactions.

Under a simple search and match framework with Nash bargaining between buyer and seller with the option to re-time their transaction similar to that in Slemrod et al.

(2016), I apply the bunching estimation method in recent public finance literature (Best and Kleven, 2015; Chetty et al., 2011; Saez, 2010) to identify the time cost of deferring transaction. Exploiting both the time notches in the "Special Stamp Duty" and the "Seller's Stamp Duty" introduced in Singapore at January 2011, where its design is highly similar to the "Special Stamp Duty" in Hong Kong, I provide estimates on the time traders are willing to wait to close the contract for tax saving purpose.

I find that traders on average are willing to wait for 3-4 weeks to avoid 1% of tax. The empirical estimates provide a way to quantify the welfare loss arising from delay in transactions, and it suggests that each week of delay in transaction would generate 0.3% loss of the average value of the property traded. It also provides quantitative benchmark for calibration in macroeconomic models that explicitly model dynamics in housing market with forward looking buyers and sellers, for example when the market thickness changes with seasonality or when moving decision is endogenous, that buyers and sellers have to decide to enter the market now or later (Ngai and Tenreyro, 2014; Ngai and Sheedy, 2016). Taking into account of both the extensive margin crowd out and the cost of re-timing transactions, I find that the "Special Stamp Duty" imposed 0.31% of welfare loss in terms of property values per holding.

The paper is divided as follows. The context of holding duration tax in Hong Kong and Singapore is described in Section 1.2. The data and empirical strategy are described in Sections 1.3 and 1.4 respectively, followed by results in Section 1.5. Section 1.6 presents the conceptual framework to understand the welfare effect of the tax. Section 1.7 present the estimate of delay cost using the framework. Section 1.8 conducts welfare analysis. Section 1.9 discuss some further implications of the tax instrument.

1.2 Transaction tax on holding duration: Context

Transaction tax on real estate properties are termed stamp duty tax in Hong Kong and Singapore. It is applicable to both residential and commercial properties, where the tax rates varies according to the bracket of the transaction price, with progressive marginal tax rate on price ranges from 1.5 % to 4.25%.[3] While the marginal tax rate in each price bracket differs, there are no notches in which the average tax rate changes at sharp price cutoffs (e.g. in contrast to that in the UK stamp tax system).

[3]Figure 1.A.3 plots the tax rate in Hong Kong as a function of price. Leung et al. (2015) studies the extent of tax avoidance to kinks in the static non-linear tax schedule.

1.2.1 Special Stamp duty tax - Hong Kong

With the aim to prohibit speculative purchases and make residential properties more affordable, the Hong Kong government introduced the Special Stamp Duty (SSD) for residential units on 20th November 2010.

The Special Stamp Duty regulates that for any residential properties obtained after 20th November 2010, an additional transaction tax would be levied if its next transaction occur within 24 months from its last transaction.[4] For units that are resold within 6 months a tax rate of 15% is charged, 10% for 6-12 months, and 5% for 12-24 months. At each of these cutoff date, average tax rate drop one day before and after the time notch.[5] On 27th October 2012 the government raised the tax rate and increased the holding duration where the tax is chargeable to 3 years. For properties resold within 6 months, tax rate of 20% is levied, while for those held for 6-12 months tax rate of 15% is charged. For properties in which the holding period were from 1 to 3 years, the tax rate become 10%. The tax schedule as a function of holding duration is plotted in Figure 1.1.

Statutorily, both buyer and seller are jointly liable for paying tax. In any efficient bargaining between buyer and seller, buyer paying the tax would maximize the total surplus of the transaction.[6] It is also a common practice in the housing market in Hong Kong for buyer to pay the stamp duty tax, therefore in the following analysis I implicitly assume that buyer is responsible for paying the tax.

The Special Stamp Duty was completely unanticipated by the market in both phases, and apply immediately to those properties traded thereafter.[7] Property obtained before the government announcement of the policy is not affected. Non-residential units are also not subject to the Special Stamp Duty.

1.2.2 Seller's Stamp duty tax - Singapore

Singapore first introduced the Seller's Stamp Duty on 20th February 2010. Under the Seller's Stamp Duty, residential properties sold within one year is taxed at 1-3% of the transaction prices, depending on the price bracket of the property. The exact

[4] This tax is levied on top of the underlying tax schedule.

[5] The duration of holding is defined as the time lapse between two legal documents that are subject to stamp duty tax, which includes most provisional agreements signed

[6] Since this minimize the transaction price on paper and thus lower the amount of tax paid out of the joint surplus.

[7] There are conditions in which the tax could be exempted: 1. the property is transferred to close relatives; 2. re-selling by financial institutions under the condition of mortgage default 3. transfer of property to government/charitable organizations 4. mandated selling by the court 5. selling of inherited properties 6. selling under bankruptcy 6. transfer of residential properties between related corporate body http://www.ird.gov.hk/chi/faq/ssd.htm

tax rate is a function of the transaction price - which creates a tax structure similar to the one studied by Slemrod et al. (2016) in Washington D.C., where there is a simultaneous decision problem in tax avoidance by adjusting transaction price or time. Properties sold with holding period beyond one year face 0% rate from the Seller's Stamp Duty.

For properties obtained between 30th August 2010 to 13th January 2011, the taxable holding duration were extended to 3 years. Moreover, notches were introduced in 1 and 2 year, where the average tax rate drop by 1/3 at the 1 and 2 years cutoff. Therefore the size of the tax notch varies with the price bracket, for example, for properties with price smaller than 180,000SGD, selling within the first year implies 1% of tax rate, and the time notches have reduction in average tax rate of 0.33% in both the 1 and 2 years cutoffs.

In 14 Jan 2011, the tax rate were raised to 16% for selling within the first year, 12% for property sold in 1-2 year, 8% for property traded in 2-3 year, and 4% for property traded between 3-4 year. This tax structure follow exactly the same as that in Hong Kong, with a smaller drop in average tax rate at size of 4%. Also, the period that subject to the tax is longer than that in Hong Kong.[8] The tax schedule of the Seller's Stamp Duty after 14 Jan 2011 is plotted in Figure 1.11.[9].

1.3 Data

This paper uses government administrative transaction record of properties in Hong Kong, covering 2009-2015. The data contains information of more than 3 million contracts that is filed at the Land Registry for registration of tax purposes and is titled Memorial Day Book. It contains sets of contracts signed and lodged to the Land Registry for formal registration, including provisional/formal agreements of sales and purchases, assignment of the properties and mortgage agreement. The agreement of sales and purchases for residential property provide information on the date, address, value of the transaction as well as stamp duty paid. For each property I observe the history of transaction from 2009, I could therefore construct the duration of holding before its next sold, a key variable for the analysis that follows. From 2009-2015 there are 480,351 residential transaction records.[10]

The holding duration where the Special Stamp Duty applies counts from the first agreement that is registered to the Land Registry, which could either be a for-

[8]Source: Inland Revenue Authority of Singapore (2011)

[9]The transition of the seller's stamp duty for properties worth less than 180,000 SGD is summarized in Figure 1.11

[10]See data appendix for more detailed description.

mal or provisional agreement of sales and purchases. While provisional agreements were signed before the formal agreement as common practice, only those that had a significant lag between the two would require registration of the provisional agreements.[11] In the empirical analysis I correct for the date of purchases to the date of which the provisional agreement is signed, if such provisional agreement exist in the Land Registry record.

For each residential transaction, I match it with the mortgage record filed to the Land Registry, a common procedure for transaction where the property were purchased under mortgage. This allows for construction of an indicator for whether the buyer purchase the properties with cash or with mortgage.

I collect data on the rateable values, the government estimates of open market rents adjusted according to property specific characteristics, for each properties being bought and resold within the period 2009-2015, dated at October 2015. Rateable values are used for tax purpose, as property owner are liable to pay the rate for each property which is around 3-5 % of the rateable values, similar to the council tax in the UK. The Rating and Valuation department states that the value adjust for property characteristics, including age, size, location, floor, direction, transport facilities, amenities, quality of finishes, building maintenance/repair and property management.[12] This provides a single measure of observables of property to control for property heterogeneity in the empirical analysis. It is important to note that the assessment of the rateable value is not affected by any sales restrictions or ownership status of the property.

I also obtained administrative record for Singapore residential properties transaction record from the Real Estate Information System of the Urban Redevelopment Authority of Singapore, from 1995 to 2015, which consists of 431,359 records during the period.

1.4 Empirical Setup and Measure

1.4.1 Extensive margin response

I first use regression discontinuity design to estimate the effect of the Special Stamp Duty on the probability that a property is being resold before 2 years where the transaction is taxable and up to 3 years where the transaction is not taxed. This quantifies the extent in which the Special Stamp Duty change the trading pattern.

[11] Usually the provisional agreement had to be registered if its lapse with the formal agreement is mroe than 14 days

[12] It is published for the public every year from March to May upon search.

Specifically, I exploit the fact that for properties purchased one day before the 20th November 2010, there is no tax liability on all holding duration while those purchased right on the 20th November it is liable for the tax. By comparing the behaviour of purchaser before and after the application of the tax locally, who face almost identical demand condition in the market following the purchases, one could identify the behavioural response that comes from the side of seller. Moreover, with the assumption that purchaser who signed the agreement one day before and after the application of the tax are similar in their selling behaviour in the absence of tax conditional on observables, one could identify the treatment effect of the tax schedule on the trading pattern.

In particular, the Special Stamp Duty introduced on 20th November 2010 had taxable period of 2 years, and with distinct tax rate at transaction of holding period before and after the cutoff of 6 months, 1 year and 2 years. I thus construct indicator T_y of whether a property is traded within y years since last purchased, where y belongs to 6 months, 1 and 2 years.

To measure the extent of extensive margin response, one must take into account of the re-timing in transaction due to tax avoidance incentives. If transaction could be perfectly re-timed by incurring cost that is low relative to the surplus of the transaction, one would observe that the probability of a property being resold not to be affected at horizon after 2 years. However, if re-timing is costly and the magnitude of tax is large relative to the size of surplus, then the tax on holding duration could cause significant crowd out, which could in turn have first order impact in welfare. I thus further construct the indicator T_y for the 3 year horizon, allowing for one year of potential re-timing of transaction.

I estimate the following linear RDD regression for each property i and spell j around 20th November 2010,

$$T_{y,ij} = \alpha + \beta\tilde{t} + \rho D_t + \beta_2 D_t * \tilde{t} + \eta_{ij} \tag{1.1}$$

where t is measured in days. $\tilde{t} = t - \hat{t}$ where $\hat{t}$ represent 20th November 2010, thus $\tilde{t}$ is the time difference in days from 20th November 2010. D_t is an indicator for $t \geq \hat{t}$. $T_{y,ij}$ is an indicator if a property holding is resold within y years since its purchase.

The coefficient of interest is ρ, which capture the difference in probability of re-trade in y years locally between purchaser who are subject to the tax and those who are not.

1.4.2 Tax incidence

The reduced form effect of tax burden on buyers and seller could be estimated with the following equation

$$\ln\ P_{ijt} = \beta_1 * H_{ij,6} * post_{ij} + \beta_2 * H_{ij,6-12} * post_{ij} + \beta_3 * H_{ij,12-24} * post_{ij} \\ + \sigma_s + \gamma_t \ln\ RV_i + \gamma_{t2} \ln\ RV_i^2 + \delta_t + \epsilon_{it} \quad (1.2)$$

where P_{ijt} is pre-tax price for property i of spell j at time t.[13] H_{ijs} is an indicator for property i held for s months, $post_{ij}$ is an indicator that equals 1 if Special Stamp Duty is applicable to the property spell ij, i.e. bought between 20th Nov 2010-27th Oct 2012. σ_s is month of holding FE at broad group category of 0-6 months, 6-12 months, 12-24 months and beyond 24 months. δ_t is fixed effect for the time in which the transactions occur, this allows us to control for any changes in average buyer's valuation affected by the tax.

RV_i is rateable value of property i measured at Oct 2015. As described in the data section, this is the government estimated open market annual rental value in October 2015, which summarizes many property specific characteristics. This control for most relevant heterogeneity of the property. As rent-price ratio can varies in the long run (Gallin, 2008), I allow for arbitrary time variant fundamental relation between rental rate and transaction price by allowing γ_t and γ_{t2} to varies over time at month level. In general, I could include an polynomial of $\ln\ RV_i$ at higher order level, but as shown as the robustness section, the inclusion of 3rd or higher order makes very little difference to the empirical results.

1.4.3 Transaction Re-timing

To capture the overall effect and discontinuity in probability of trade generated by the Special Stamp Duty, I estimate the following hazard function flexibly at week intervals:

$$T_{ijt} = \sum_{j=0}^{J} w_j \mathrm{I}_j + \epsilon_{it} \quad (1.3)$$

where T_{ijt} is an indicator of whether a trade happen for property i at time t at the jth week since it was last purchased. I_j is a set of fully flexible dummy variables that capture the baseline hazard.[14] Equation 1.3 is estimated separately

[13] A spell is defined as a holding that start since a transaction occurred

[14] Estimating Equation 1.3 by OLS is equivalent to the Kaplan-Meier estimator

for residential properties in Hong Kong purchased in 3 periods: 1) 1st January 2009 - 20th November 2010 where there is no tax liability as function of holding duration; 2) between 20th November 2010-26th October 2012 where time notches are present in the 6, 12, 24 months and 3) on or after 27th October 2012 where time notches are present in 6, 12, 24, 36 months. As far as the data allows, I estimate equation 1.3 taking J up to 158 weeks. That allows one to capture any re-timing effect up to the third year.

The degree of re-timing would be captured by the estimate w_j where j equals the time of the notch. One would expect the probability of trade to be much higher right on the notch, and if there is any market friction that prevent trading happening exactly at time $\bar{t}$, one would also see that for the immediate weeks after the notch the estimate w_j would still be higher than the counter-factual w_j as estimated using properties without the tax liability. Correspondingly, one would observe w_j would be lower than counter factual for weeks immediately before the time notch.

Similarly for Singapore I estimate the hazard function of trade up to 5 years as the data allows, to capture any re-timing effect following the time notch at the 4 years since purchase.

1.4.4 Type selection

To identify the selection effect of the Special Stamp Duty on new buyers, I estimate the proportion of transaction where mortgages are filed in each week before and after the introduction of the tax. I estimate the following event-study equation using transactions of residential properties that has price below HKD 6,800,000, a price range with minimal change in loan-to-value ratio limit in the mortgage market for period 3 months before and after 20 November 2010[15]

$$m_{it} = \beta + \beta_1 * \sum_{b=-24}^{24} w_b \mathrm{I}_b + \mu_i + \gamma_{dm} + \epsilon_{it} \tag{1.4}$$

$m_{it} \in \{0, 1\}$ is an indicator equal to 1 if the property i at time t is purchased by mortgage buyer. I_b is an indicator of whether the transaction it occurred bth weeks before or after the application of Special Stamp Duty phase 1. To increase the power and control for seasonality, I estimate the following regression:

[15]If any of the exogenous reduction in loan-to-value ratio limit in this period have effect on mortgage demand, the estimate in this section provide a lower bound of the selection effect. See Figure 1.A.4 for a graphical summary of change in loan-to-limit ratio since 2009 in various price range.

$$m_{it} = \beta + \beta_2 * \text{After}_{it} + \mu_i + \gamma_{dm} + \epsilon_{it}$$

further including district-month of the year fixed effect γ_{dm} by extending the sample from July 2009 to March 2011, and property fixed effect μ_i to control for unobserved characteristics of the property.

1.5 Results

1.5.1 Re-timing and extensive margin response

Figure 1.2 plots the RDD graphs on the probability of re-sale in different time horizon since purchase by the day of the property was being obtained. The outcome variables are accumulative in each of the panels in Figure 1.2, thus the probability of being resold in 3 years in panel (a) is by definition larger than the probability of being resold in 6 months in panel (d). It is clear from all the panels in Figure 1.2 that there is a sharp discontinuity on 20th November 2010. For purchasers who purchase one day after 20th November 2010, the probability of resold in 6 months, 1 year, 2 years and 3 years are all significantly lower than those before it.

Table 1.1 reports the RD estimate of equation 1.1. Column (1) shows the coefficient for the effect of the tax on probability of re-sold in 3 years, where the estimate for ρ is -0.133 and is highly statistically significant, suggesting that purchaser who are subject to the tax are less likely to trade in 3 years by 13 percentage point. The baseline probability of a property being resold in 3 years is 29 percentage point, in which the treatment effect suggest a very large drop in the transaction at extensive margin. On the other hand, the estimates for ρ in column (4) for probability of resold in 6 months is exceptionally large at -0.0743 in relation to its baseline mean which is 7.59 percentage point, suggesting that the Special Stamp Duty reduced trade at very short duration dramatically.

I further estimate equation 1.1 by the type of buyers with or without mortgage. The treatment effect for buyers on the chance of resold in 3 years, for those without any mortgage is 16.6 percentage point, and is bigger than those without mortgage, where the treatment effect is 11.6 percentage point. This provide evidence that the tax has a stronger extensive margin effect on purchaser who are more likely to re-trade in early period.

Table 1.3 reports estimate of equation 1.1 for the holding duration conditional on a property holding is resold before 31st December 2015. Column (1) reports the estimate for all property holdings, while the average holding duration conditional on sales is 733 days before the treatment, the Special Stamp Duty increase the holding duration conditional on sales by 334 days. Comparing column (2) and (3), the magnitude of the treatment effect is of the same order of magnitude for property holding that is associated with or without mortgages.

Table 1.4 reports estimate of equation 1.1 on the probability of a property being resold in 5 years for the effect of Seller's Stamp Duty Phase 3 on 14 Jan 2011, where the maximum taxable holding period is up to 4 years and therefore capture the extensive margin effect taking potential re-timing of transaction into account. The coefficient estimate is -0.058, which suggest that property were less likely to be re-traded in 5 years by 5.8 percentage point. The related RDD graph is plotted in Figure 1.A.7.

I also conduct the test on continuity of the density around the cutoff date as suggested by McCrary (2008). Panel (a) and (b) of Figure 1.A.8 plot the density of trading day around 20th Nov 2010 for Hong Kong, and the graphs suggest that there is discontinuity in the density of trade around this period of time. The density mass of the transaction on the day the policy announced on 19 Nov 2010, which is a Friday, is not significantly higher than previous Friday, implying there is very little manipulation in the conventional sense. It suggests that the discontinuity is mainly driven by reduction in transaction after the application of the policy, which is by itself an outcome of the tax policy. The graphical evidences from Figure 1.2 display no discontinuous change in the outcomes to the left of the cutoff, again confirming that the RD estimate is unlikely to be confounded by manipulation.

1.5.2 Tax incidence

Table 1.5 shows the coefficient estimates for equation 1.2 for transaction price of residential properties sold in 2009-2015 in Hong Kong. Column (1) leaves out (log) rateable values as control while column (2) includes it. The coefficients of interest are *post*sold 6m*, *post*sold 6-12m* and *post*sold 1-2year*. The estimate for *post*sold 6m* is -0.0103 %, implying that property traded with 5% tax are associated with 1.03 % lower price for seller, and the buyer pay 3.97% higher price for the property that are observationally equivalent to property that are not subject to the tax in the same market. The coefficient *post*sold 6-12m* is -0.0505 and statistically significant, suggesting that property that are subject to 10% special stamp duty were traded at price 5.05% lower, and thus buyers are paying 4.95% higher price for the property. For property that are sold within the first 6 months are sold on average 8.85 % lower seller price, where it is subject to tax rate of 15%. This suggest that buyer pays around 6.15% higher price.

The difference between the estimates for *post*sold 6m*,*post*sold 6-12m* is around 0.038 and the difference between *post*sold 6-12m* and *post*sold 1-2year* is around 0.04. For every additional increase in the tax rate from 5% to 10% and 10% to 15%, the buyer only bear around 1 % of the additional tax, while the seller receive price that is 4% lower.

Comparing column (1) and (2), there are significant selection in the value of properties being traded under high transaction tax. In column (1) where I did not include rateable value as control, the *post*sold 6m* coefficient is positive, in line with the pattern observed in Figure 15, in which the median transaction price is plotted against the holding duration. The coefficient become significantly negative once we control for rateable values, suggesting that selection on observables explains a large proportion of the positive coefficient in column (1). Similarly, the coefficient *post*sold 6-12m* is not statistically significant and is close to zero in column (1), while it turns into negative and significant once I control for rateable value. This suggest that properties at a higher price range have a larger mass for having surplus beyond 10%.

Figure 1.A.10 plot the relationship between the log price and log rateable values for the sample in 2009. It shows the log rateable value is highly correlated with transaction price in a stable manner. In Table 1.6 I examined whether second order polynomial is a good fit for Rateable value as control for observables in the transaction price. I estimate equation 1.2 using 1-4 order polynomials from column (2)-(5). Upon the 2nd order polynomial, the estimates are very robust to adding additional polynomial terms. This suggest that 2nd order polynomial is a very good fit for rateable value to predict transaction price.

1.5.3 Transaction re-timing

Figure 1.5 plots the coefficients of w_j in equation 1.3 for properties that are purchased between January 2009 to 19th November 2010 and from 20th November 2010 to 26th October 2012. There is significant re-timing for transaction, and sharp discontinuities are observed in each of the time notch. There are three salient patterns 1) the tax schedule unambiguously decrease the probability of trade within the first 2 years of purchases; 2) the probability of trade increase dramatically at exactly 2 year, the last tax notch where the tax rate drop to 0; 3) the excess probability of trade is decreasing gradually after the 2 year time notch. Pattern (2) and (3) together provide sound evidence that many trades in the market are intentionally shifting their trading time to avoid the 5% tax rate within the 1-2 year window of holding period.

The reduction in the probability of trade within the first 2 years is proportional to the magnitude of the tax. Within the first year, the decline in the probability of trade is around 0.2-0.4%, which is almost equal to the pre-treatment average. In the 1-2 year period, the treatment effect is around 0.15%, and the probability of trade after the application of Special Stamp Duty is decreasing as one approach the 2 year time notch. Figure 1.4 plots the densities of holding duration for properties purchased in the two different periods and it shows highly similar pattern.

Figure 1.8 plot both the hazard function of trade and the coefficients of $week_j$ in equation 1.3 by whether the transaction has affiliated mortgages. Panel (a) plotted the hazard function by type of buyers for properties obtained under period that is unaffected by the tax, which shows cash buyer are more likely to trade within the first year, while starting from the second year the difference diminished or even reversed. The treatment effect of tax plotted from panel (b) shows that the Special Stamp Duty has a much stronger impact on cash buyer than buyer who filed mortgages, mainly due to the fact that cash buyer has a larger baseline probability of trading below 1 year. The probability of trade in the 1-2 years period on both type of buyers are indistinguishable after the application of the tax. And both exhibit similar degree of bunching towards the 2nd year time notch.

1.5.4 Selection of buyer

Figure 1.9 plot the estimates of coefficients of time t in equation 1.4. The base group is on the third week of November 2010. The coefficients are not statistically significant from zero before the introduction of Special Stamp Duty, and also exhibit no trends. The estimates of the time coefficients after third week of November 2012 are slightly positive and become statistically significant at around 0.05 percentage point after the 5th weeks.

Table 1.7 shows the estimation results of equation 1.4. Column (1) report the coefficient that cover sample of 2 months before and after the introduction of Special Stamp Duty. The estimates suggest that after the introduction of the special stamp duty tax, 1.7% of the buyer are buyers who would filed mortgage, compare to period in which the Special Stamp Duty was not in place. Column (2) and (3) extends the sample to 3 and 4 months before and after the Special Stamp Duty and the coefficient is largely similar in magnitude.

Column (4) extends the sample to July 2009 to March 2011, and control for district-month of the year fixed effect to control for seasonality in the last quarter of the year. The estimate become 2.92%, which is very close to estimates in column (1)-(3). Column (5) control for property fixed effects, and the coefficient become even bigger at magnitude of 5.52%. The evidence suggest that after the introduction of the Special Stamp Duty among the new buyer there is 2-5% more of those filed mortgages.

1.6 Conceptual framework

In this section I provide an empirically relevant framework to understand the welfare implication of transaction tax on holding duration. The market response and welfare cost of transaction tax with time notches largely depends on the time preferences of representative buyers and sellers. (Slemrod et al., 2016) If buyer and seller has low adjustment cost in transaction time, a time notch may generate significant bunching toward the side with lower average tax rate to avoid tax. (e.g. the right hand side of the time notch in Figure 1.1) This means the tax would shift the timing of transaction, but few matched transactions would be forgone. On the other hand if adjustment cost is high, there would be limited tax avoidance behaviour, but transaction with low surplus would be crowded out from the market depending on the size of the tax rate.

Central to this, I present a stylized model on buyer and sellers simultaneous decision on transaction price and trading time, simplifed from Slemrod et al. (2016). In the empirical section I would show that its prediction match the basic pattern of the market reaction, and argue that it provides a framework for identifying the adjustment cost in time for transaction. In the model, buyers and sellers explicitly prefer early trading date.[16] The trade-off in the optimal trading time at the intensive margin naturally features intertemporal decision between tax saving and disutility associated with delaying transaction.

Consider a pair of matched buyer and seller who value the property by H and u respectively, at time t counting from the seller last bought the house. The buyer and seller bargain over the price and trading time through Nash bargaining. A tax notch is set by the government, where the transaction is taxable at rate $\bar{\tau}$ if the trade is executed at time T when $T < \bar{t}$, otherwise tax rate $\underline{\tau}$ applies where $\underline{\tau} < \bar{\tau}$.

The seller's surplus in a transaction is

$$S_v = p - u - k_v(T - t)$$

where seller receive price p over losing the reservation value u. $k_v(T - t)$ capture the time preference of the seller, and $T - t$ is the time between the transaction and

[16]Where in Slemrod et al. (2016) it is modelled as buyer and seller having an ideal trade-date, a setup that is more relevant when considering an increase in tax rate after a fixed date

the match was first formed. Assume that $k_v > 0$ so that the seller prefer trading earlier. k_v is thus the utility cost in delaying transaction *per unit of time.* $k_v(T-t)$ could be considered as first order local approximation for a general function of T and t.

Similarly, the buyer's surplus is

$$S_b = H - p(1+\tau) - k_b(T-t)$$

where the buyer get the valuation H, pay the price p, and are subject to tax rate τ, where $\tau \in \{\underline{\tau}, \bar{\tau}\}$. $k_b(T-t)$ capture the time preference of buyer, with $k_b > 0$ so that the buyer also prefer to trade sooner than later.

The trading time T and price p is determined by Nash Bargaining that maximize the joint surplus, i.e.

$$max_{T,p}\,(S_b)^{1-\theta}(S_v)^{\theta}$$

If trade happen at time T, transaction price would follow the first order condition $(1+\tau)(1-\theta)S_v = \theta S_b$, explicitly given by the following

$$p(T,t,H,u) = \begin{cases} \frac{\theta}{1+\bar{\tau}}H + (1-\theta)u & \text{if } T = t \\ \frac{\theta}{1+\underline{\tau}}[H - k_b(\bar{t}-t)] + (1-\theta)(u + k_v(\bar{t}-t))] & \text{if } T = \bar{t} \end{cases} \tag{1.5}$$

To maximize the joint-surplus, only $T = \bar{t}$ and $T = t$ would be chosen, therefore a match between buyer and seller, depending on the respective size of H and u, would decide upon: 1) trade immediately at time $T = t$, 2) bunch at the time notch $T = \bar{t}$, or 3) no trade at all.

To understand better the trade-off, a pair of buyer and seller would choose to trade exactly at the time $\bar{t}$ if

$$S_b(T=\bar{t}) + S_v(T=\bar{t}) > max\{S_b(T=t) + S_v(T=t), 0\}$$

When $\underline{\tau} = 0$, the condition in which the match would wait until $\bar{t}$ to trade could be further reduced into

$$p(t,t,H,u)\Delta\tau_t > (k_b + k_v)(\bar{t}-t) \tag{1.6}$$

where $\Delta\tau = \bar{\tau} - \underline{\tau}$. The left hand side of equation 1.6 is the tax saving of executing the transaction at time $\bar{t}$ instead of time t, while the right hand side of the equation is the total utility cost of delaying the transaction by $\bar{t} - t$ unit of time.

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Proposition 1 *Assume that H and u follows some joint distribution of $G(H,u)$, if time preference for trade is weak, i.e. $k_b + k_v$ is small, then more transactions would trade exactly at $\bar{t}$, less transactions are carried out at $t < \bar{t}$*

Proposition 1 says that when time preference is weak, the buyer and seller has little to lost by delaying the trade date. This also implies that we may observe very few transactions realizing before $\bar{t}$. One thing to note is that the shifting in timing of transaction need not depends only on the time preference of the seller even if the tax is seller specific. In an economy where the seller has distinct characteristics than the buyer, in terms of demographics or economic wealth[18], both the time preference for buyer and sellers affect how the market response to the tax.

Proposition 2 *At the notch of which $\underline{\tau} = 0$, and some joint distribution of H and u $G(H,u)$*

1. *the probability of trading at time t decrease as t get closer to $\bar{t}$, for $t < \bar{t}$*
2. *The observed reduction in probability of trade at $t < \bar{t}$, where the tax rate drop to 0 after $\bar{t}$, is larger than the case with a tax policy that applies tax rate τ for all trading time t*

When $\underline{\tau} = 0$, the decision between trading immediately or delaying transactions depends on the relative rate of decline between the saving of tax with the utility lost of delaying. In general, the disutility decline with the rate $k_b + k_v$ when one get closer to $\bar{t}$, while the savings by tax decrease by the rate $((1-\theta)k_v - \theta k_b)\Delta\tau_t < k_b + k_v$. Thus the probability of trade at time t for $t < \bar{t}$ is decreasing in t.

The condition in which the three decisions are determined are illustrated in Figure 1.A.12. It plots the combination of different H and u with the relevant constraints when $\underline{\tau} = 0$. In the absence of tax, trade would happen for H and u when surplus is strictly positive.

When a transaction tax τ is introduced, only matches with surplus higher than τ percent of the reservation value of seller are traded. However if a time notch in $\bar{t}$ is present, some matches that would yield non-positive surplus under tax τ would then have positive surplus by waiting until $\bar{t}$ to trade (i.e. area B in Figure 1.A.12).

The empirical prediction in Proposition 2 can seen by conducting comparative static with respect to t. For matches formed far away from $\bar{t}$ implies a higher

[17]The general case where $\underline{\tau} \neq 0$ is $p(t,t,H,u)\Delta\tau_t > (k_b + k_v)(\bar{t} - t) - \Delta p_t \underline{\tau}$

[18]e.g. Lawrance (1991) finds that poorer households has higher rate of time preference

$(k_b + k_v)(\bar{t} - t)$. The line $H - u \geq k(t)$ and the downward sloping dotted line would both move to the right. The area $B + W$ would unambiguously decrease, and the area $T + A$ would increase. When matches are formed far away from the time notch $\bar{t}$, the probability of trading at time t thus increase.

Consider an exogenous increase in τ. Less transactions has the surplus over reservation value ratio higher than the tax rate, but at the same time it lower the cutoff of p_d in which it worth trading exactly at $\bar{t}$.

1.6.1 Selection on buyers and effect on equilibrium price

The application of transaction tax with rates that decrease with holding duration could create selection on type of buyers and changes the outside option of buyer in the search process.

If buyers have predetermined characteristics θ that could predict average trading time of the property in the future, buyers with different θ would foresee to face a different tax rate in expectation. In particular, the valuation of a buyer at the time of buying could be represented as

$$H = \phi(\theta)\underline{u}(\tau) + (1 - \phi(\theta))\bar{u} - z(\tau) \tag{1.7}$$

where $\phi(\theta)$ is probability of trade happen at $t < \bar{t}$ in the absence of tax, $\underline{u}(\tau)$ is the present discounted value of selling in period $t < \bar{t}$, $\bar{u}$ is the present discounted value of selling in period $t > \bar{t}$, as long as $\underline{u} < \bar{u}$, there would be more buyer with characteristics θ_l in the market, with $\theta_h > \theta_l$ and $\phi'(\theta) > 0$

Furthermore, by reducing the probability of buyer matched with seller with positive surplus, the tax can affect the outside option of the buyer in the bargaining process. For example, the outside option $z(\tau)$ could be decreasing in τ if the tax reduce the effective supply. On the other hand if the selection effect reduce the number of buyers in the market, it may increase the chance the buyer meeting another seller with positive surplus, in that case $z(\tau)$ would be increasing in τ. Formally endogenizing the outside option for buyer is beyond the scope of this paper.

1.7 Estimating waiting time and welfare cost of delayed transactions

To quantify the effect of the tax notch on trading time, I estimate the average waiting time for traders in the market $\bar{t} - t(H, u)$, where $t(H, u)$ is defined by rewriting equation 1.8 in the following form for each pair of H and u:

$$p(H,u)\Delta\tau_t = (k_b + k_v)(\bar{t} - t(H,u))) \tag{1.8}$$

$t(H,u)$ could be interpreted as the cutoff trading time for a transaction with price $p(H,u)$. If the seller meet a buyer at time t where $t \in \{t(H,u), \bar{t}\}$, the optimal trading time would be at time $\bar{t}$. For a given tax notch τ, expectation of the cutoff trading time and $\bar{t}$ across all the potentially profitable trades, $E(\bar{t} - t(H,u)|H - u > \tau u)$, would be the average waiting time for which the market would be willing to delay the transactions to time $\bar{t}$.

Assume that the density of arrival time of matches is $l(t)$, and the density of trading time in the presence of tax τ is $g(t)$, which could be estimated from the data, the excess mass B at $\bar{t}$ could be represented as

$$B \approx \int_{H-u>\tau u} \int_{t(H,u)}^{\bar{t}} l(t)dtdG(H,u) = g(\bar{t})(E(\bar{t} - t(H,u)|H - u > \tau u) \tag{1.9}$$

where $G(H,u)$ is any joint distribution function of H and u. [19]

The approximation is more accurate if the extensive margin response near the time notch is negligible. Under existence of extensive margin response locally at the time notch, the bunching estimates could provide a upper/lower bound of the quantity of interest, $(E(\bar{t} - t(H,u)|H - u > \tau u)$, by scaling the excess bunching mass with the density of the counter-factual density with tax rate 0 or with tax rate τ, which I provide more discussion on it in the appendix. [20] Following the existing literature as in Best and Kleven (2015) and Chetty et al. (2011), I estimate the counter-factual density $g(t)$ using the following equation:

$$b_j = c(j) + \sum_{k=\bar{t}-R}^{k=\bar{t}+R} \delta_k \mathrm{I}[j = k] + \mu_j \tag{1.10}$$

where b_j is the bin count in holding duration for jth weeks, $c(j)$ is a 7th order polynomial in terms of number of weeks, and $\mathrm{I}[j = k]$ is a set of dummies for R week before and after the tax notch at $\bar{t}$. [21] The prediction of this equation, $c(\bar{t})$, is going to provide an estimate of the counterfactual density $g(\bar{t})$, conceptually by excluding

[19] The estimate would be robust to the case where buyer and seller have exponential discounting with first order linearization, in that case $k_b = H\rho_b$ and $k_v = u\rho_v$, where ρ_b and ρ_v are discount rates of the buyer and seller respectively

[20] In the case of there is significant extensive margin response in at time $\bar{t}$, the estimate provide a bounding on $(E(\bar{t} - t(H,u)|H - u > \tau u)$. A detailed discussion is presented in the section 1.A.3.

[21] In practice I include the dummy variables for holding duration from 330 to 400 days to estimate the excess mass at 1 year notch, and 694 to 864 days for the 2 years notch.

the weeks immediately before/after $\bar{t}$. Thus we could compute the estimate using equation 1.9.

$$\hat{E}(\bar{t} - t(H,u)|H - u > \tau u) \approx \frac{\hat{B}}{g(\hat{t})}$$

where $\hat{B}$ is computed using the observed excess bin count around $\bar{t}$.

The ratio of waiting time to tax bear an intuitive interpretation for estimating the utility cost of $k_b + k_s$ as percentage of trading price, by taking appropriate expectation over equation 1.8 and rearranging terms:

$$\frac{\Delta\tau}{E(\bar{t} - t(H,u)|H - u > \tau u)} = \frac{k_b + k_s}{E(p|H - u > \tau u)} \tag{1.11}$$

1.7.1 Results

Figure 1.13 panel (a) plots the actual and estimated counter-factual density of the 2 years notch of the Special Stamp Duty, for properties purchased between 20th November 2010 to 26th October 2012. Using the gap between the actual density and the counter-factual to calculate the excess mass, I find that on average the cutoff waiting time, $E(\bar{t} - t(H,u)|H - u > \tau u)$, is 17.39 week when the size of the notch is 5%.

Panel (b) plot the actual and estimated counter-factual density of the 4 years notch of the Seller's Stamp Duty, for properties purchased between 14 January 2011, in which the 3rd phase of the Seller's Stamp Duty begins, to those that are purchased before 14 January 2012. I restrict plotting the data to the one year windows because for those properties that are purchased in mid-2012 the transaction after 4 years since purchased is still not being observed at the time of writing. I find that the bunching estimates of the average cutoff waiting time is 13.32 week, slightly smaller than the estimate from the Special Stamp Duty but is associated with a smaller tax notch at 4%.

Using equation 1.11, one can apply the estimates obtained for $E(\bar{t} - t(H,u)|H - u > \tau u)$ and the related size of tax notch to identify $\frac{k_b+k_v}{E(p)}$ which is the proportional welfare cost of delaying a transaction by a week. From the 2 years tax notch of the Special Stamp Duty phase 1, the implied $\frac{k_b+k_v}{E(p|H-u>\tau u)}$ is 0.29%, while from the last tax notch of the 3rd phase of Seller's Stamp Duty, the implied estimate for $\frac{k_b+k_v}{E(p|H-u>\tau u)}$ equals 0.3%, suggesting delay in a single transaction would results in lost of trading surplus by 0.3% of the value of the transacted property.

1.8 Welfare analysis

Following the conceptual framework, the welfare gain from transaction in the economy per unit of property held less than 3 years when the tax rate is 0 is given by

$$W(0) = \int_t \int_{H-u>0} (H-u) dG(H,u) d\tilde{L}(t) \tag{1.12}$$

With tax rate $\tau > 0$, the welfare is given by

$$W(\tau) = \int_t \int_{A,T} (H-u) dG(H,u) d\tilde{L}(t) + \int_t \int_{W,B} (H-u) \\ - (k_b + k_v)(\bar{t} - t) dG(H,u) d\tilde{L}(t) + \lambda R(\tau) \tag{1.13}$$

where A, T denote the value of H, u such that trading immediately is optimal, W, B denote the value of H, u where wait is the optimal decision, as shown in Figure 1.A.12. Denoting the tax revenue by $R(\tau)$ and the shadow value of tax revenue as λ, the change in welfare is given by

$$W(\tau) - W(0) + \lambda R(\tau) = \\ -\int_t \int_{NT} H - u dG(H,u) d\tilde{L}(t) - \int_t \int_{W,B} (k_b + k_v)(\bar{t} - t) dG(H,u) d\tilde{L}(t) + (\lambda - 1) R(\tau) \tag{1.14}$$

To bound the welfare loss, I use the estimates of $k_b + k_v$ from the previous section. The welfare cost associated with the second term, for which a property were forced to wait until 2 years, could be approximated by

$$\int_t \int_{W,B} (k_b + k_v)(\bar{t} - t) dG(H,u) d\tilde{L}(t) \\ \approx (k_b + k_v) * \frac{E(\bar{t} - t(H,u) | H - u > \tau u)^2}{2} * P(H - u > \tau u) * l(\bar{t}) \tag{1.15}$$

We could obtain an estimate of $P(H - u > \tau u) * l(\bar{t})$ by the probability of a property trade per week after for $t < \bar{t}$, which is roughly 0.0005 under the 5% tax. Assuming the the tax rate was 5% everywhere with $\bar{t}$ equals to 2 years, the welfare cost associated with the second term is $\frac{17.36^2}{2} * 0.29\% * 0.0005 * E(p) = 0.02\% * E(p)$

of the average value of property in the market.

To quantify the welfare loss at the extensive margin, it would be important to understand how trading surplus is distributed among transactions that were crowded out. For example, at a tax rate of 5%, all the transaction that were crowded out must have surplus lower than 5% of the seller's reservation value. Whether the trading surplus concentrated mostly towards 0% versus 5% determine the size of the lost in trading surplus.

To understand better the distribution of trading surplus, I take advantage of the fact that under the same context both the Special Stamp Duty and the Seller's Stamp Duty generate exogenous variation in tax rate that results in different trading hazard at various tax level. The hazard rate of trade at a particular tax level depends on likelihood that trading surplus in proportional sense, $z \equiv \frac{H-u}{u}$, is greater than the transaction tax level. In particular, at different tax rate τ and τ' abstracting from the re-timing close to the notch, with $h(\tau)$ being the hazard rate of trade at time t with tax rate τ, we can calculate the density of the trading surplus

$$\lim_{\tau' \to \tau} \frac{h(\tau) - h(\tau')}{h(0) - h(\bar{\tau})} \frac{1}{(\tau - \tau')} == -j(\tau | 0 < z < \bar{\tau}) \tag{1.16}$$

for any τ and τ' that is close, $\tau < \tau'$, and $\bar{\tau} > \{\tau, \tau'\}$. [22]

That is, by comparing the rate in which the hazard rate decline with respect to tax rate, one could infer the density of the conditional distribution of the surplus. Moreover, with enough data point at each tax rate, one could estimate non-parametrically the density of the trading surplus.

For the Special stamp Duty, I get estimates from the earlier section on $h(\tau)$ where $\tau \in \{0, 5, 10, 15\}$

In practice, I use the hazard rate $h(\tau)$ where $\tau = \{0, 5\}$ to estimate the conditional density at $\tau = 0$, and use $\tau = \{5, 10\}$ to estimate the conditional density at $\tau = 5$ so and so forth.

The results for both trading surplus in Hong Kong and Singapore is presented in graph 1.A.11, and the shape is very similar using the two procedures. It suggest that the trading surplus decline very rapidly at around 4-10%, and there is little mass above the threshold of 10%.

Using a linear approximation of the density between 0% and 5%, I find that the

[22]

$$\lim_{\tau' \to \tau} \frac{h(\tau) - h(\tau')}{h(0) - h(\bar{\tau})} \frac{1}{(\tau - \tau')} = \lim_{\tau' \to \tau} \frac{(P(z > \tau) - P(z > \tau'))l(t)}{(P(z > 0) - P(z > \bar{\tau})(\tau - \tau')l(t)}$$
$$= -j(\tau | 0 < z < \bar{\tau})$$

conditional expectation of the trading surplus is $E(\frac{H-u}{u}|0 < \frac{H-u}{u} < 0.05) = 2.23\%$. Therefore the extensive margin lost would be $-0.13 * 2.23\% = -0.29\%$ of the value of property.

In total, the lower bound of welfare loss would be the sum of the intensive and extensive margin, which amounts to 0.31% of the value of a representative property. This is the welfare loss per each new holding, and is increasing in the number of properties purchased in which the policy is in place. Given that very little tax revenue is generated from the tax schedule, this suggest non-trivial welfare loss to the property market as a whole.

1.9 Discussion

1.9.1 Comparison with capital gain tax

The extensive margin response or lock-in effect shown previously, in comparison to previous literature, is reasonably large. For example, Shan (2011) finds that 10,000 USD increase in capital gains tax liability reduces semiannual sales by 0.1 percentage points. Take the upper bound on the special stamp duty tax and assume the tax rate of 15% apply to all the transaction below 2 years, for a residential property at median price approximately 4,000,000 HKD, this only translates into 0.77 percentage point reduction in sales rate every half year, which in turns translate into 4.5 percentage point reduction in probability of sales in three years. This is only one-third of the extensive margin response from the RDD estimate. [23] It suggests that transaction tax have larger impact of on the extensive margin of transaction than capital gain tax in terms because the tax applies regardless of the magnitude of appreciation of the asset.

1.9.2 Simulation of short-holding trades and equilibrium price effect

To evaluate the effect on equilibrium price from reduction in effective supply of short-holdings properties, I proxy for the short-run supply by

$$SS_{dt} = \sum_{j=1}^{150} \hat{w}_{jt} N_{d,t-j}$$

[23] $1 - (1 - 0.0077)^6 = 0.045$, where $(1 - 0.0077)^6$ is the probability of a property remain unsold in 3 years.

where $N_{d,t-j}$ is the number of transaction in district d in week $t-j$.[24] The effective supply is proxied by probability of the properties being sold in the market, $w_{j,t-j}$, multiplied by the number of properties purchased in period $t-j$ in district d.

$w_{j,t-j}$ is time-varying. For the three period with different tax liability, from Jan 2009 to 19 November 2010, 20 November 2010 to 27 October 2012, 27 October 2012 to 31 December 2015, it equals to the estimated coefficient w_j in equation 1.3 estimated for properties purchased in the relevant period. The evolution of the variable SS_{dt} is plotted in Figure 1.10. During the period, predicted short-run re-sales drop continuously from December 2011 to December 2015. The continuous decline in predicted re-sales in 2012 coincides with rise in transaction price even within each districts, and the increase in transaction price come to a slow down in the last quarter of 2012 together with the increase in the predicted re-sales.

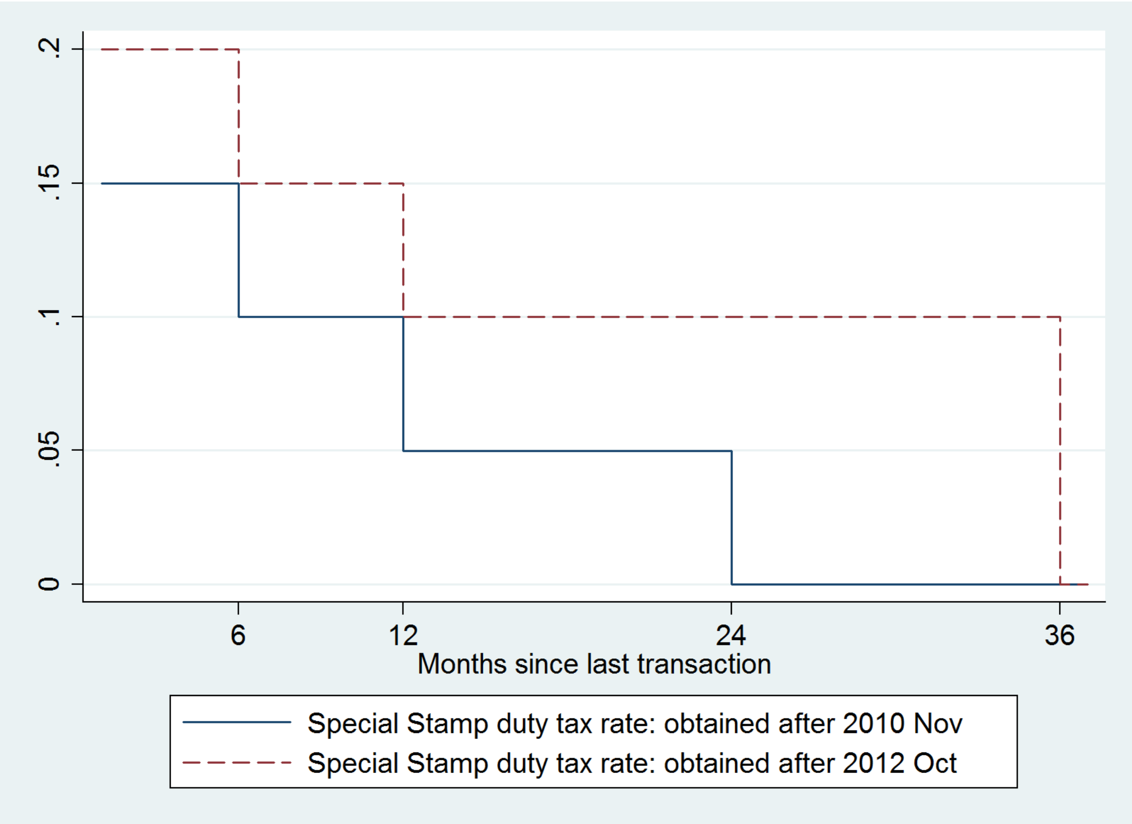
1.10 Conclusion

In this paper I show that tax notches at holding duration in Hong Kong and Singapore generate distinct incentives for buyer and seller in re-timing the transaction, on average for 3-4 weeks for each 1% of tax savings. Combined with the actual significant size of the tax, it reduced the probability of a property being traded by 13 percentage point even one year after the tax dropped to zero, which could be due to the search frictions in the property market. Buyer share up to 41-80% of the burden of the transaction tax, which may have high social cost in a policy that gives larger welfare weight to buyer than seller. This paper also provides evidence on the selection effect of this type of tax instrument. Tax on holding duration were used as public policy tool to deter speculation in the 1970s in North America, but very little were known about the tool before the new wave of introduction of similar tax instrument in east Asia in 2010s. This paper suggest that the aggregate welfare loss could be large for an understudied tax instrument that survived through history.

[24]Number of district is 3, the largest natural division for Hong Kong

1.11 Figures

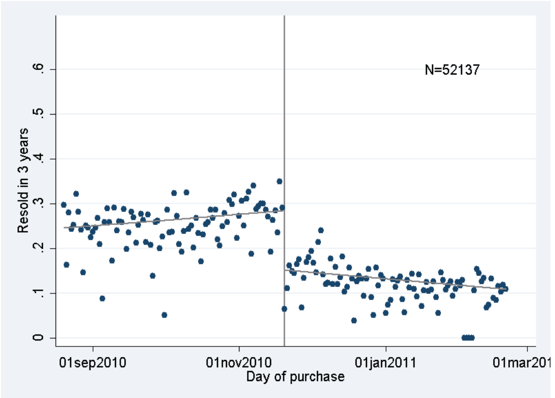
Figure 1.1: Tax on holding duration: Hong Kong



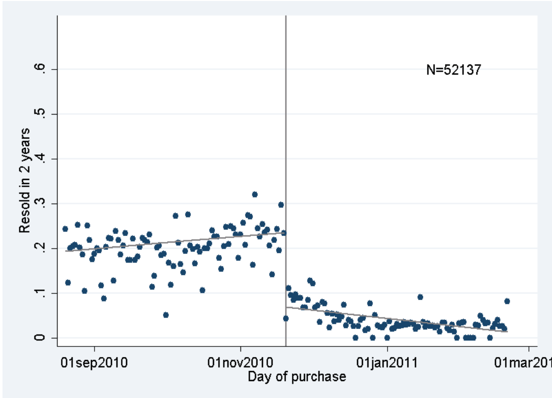
Notes: The graph plots the transaction tax rate as function of holding duration of properties since last purchased (the Special Stamp Duty), which applies to residential properties obtained after 20 November 2010 in Hong Kong (the solid line), and for those obtained after October 2012 (the dotted line). It is levied on top of the basic stamp duty rate that does not depend on the holding duration of a property.

Figure 1.2: RDD plot: all purchases

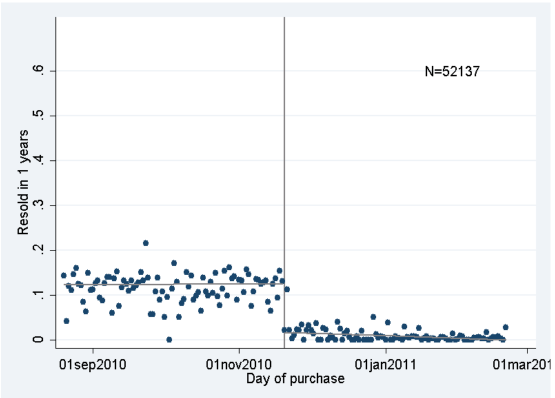
(a) Probability of resold in 3 years



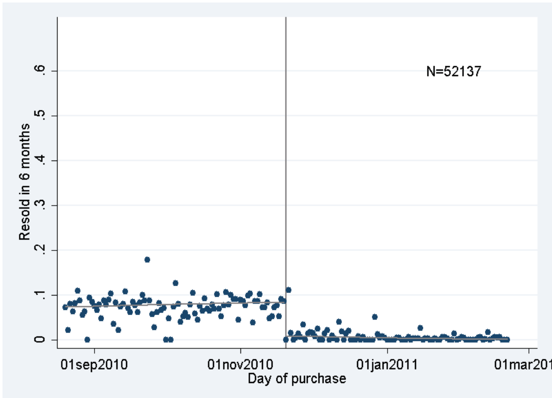
(b) Probability of resold in 2 years



(c) Probability of resold in 1 year

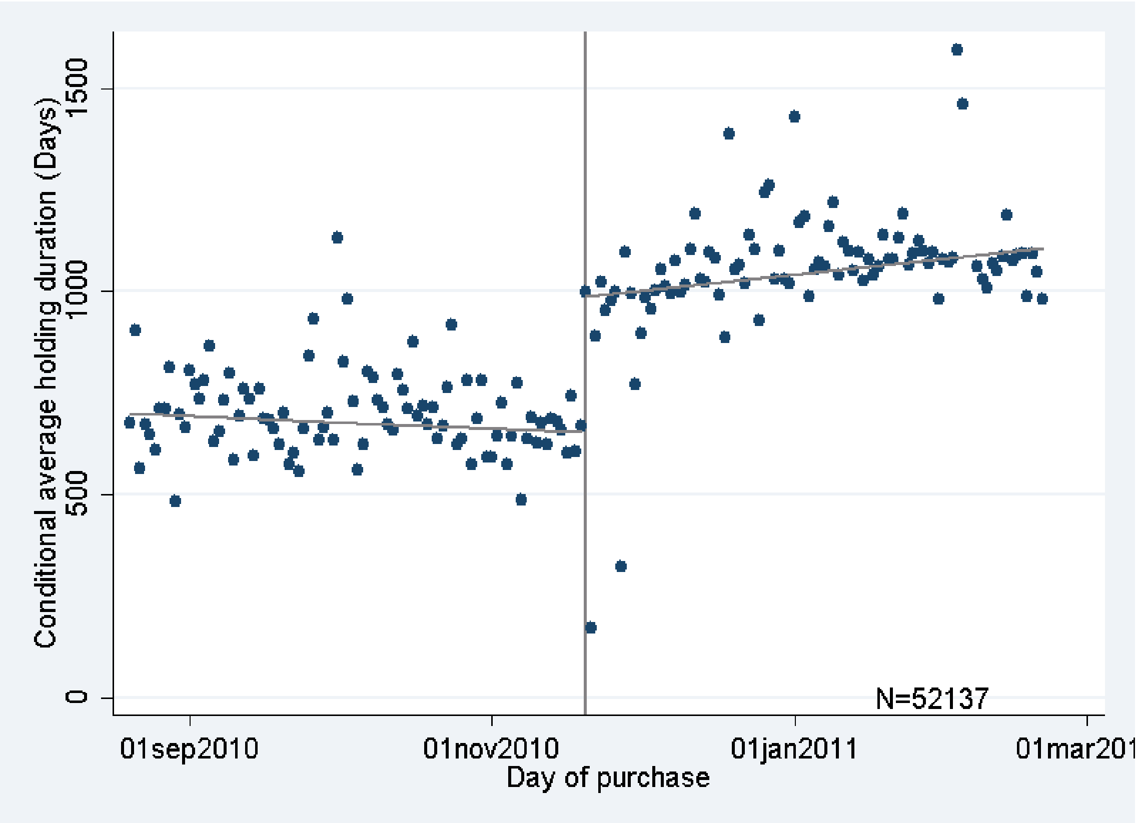


(d) Probability of resold in 6 months



Note: The graphs plot the probability of a property being resold in 3,2,1 and 0.5 years since last purchased by the day in which they are obtained. The vertical line shows the day of 20 November 2010, in which the Special Stamp Duty first applies. The dotted lines are fitted linear trend in the probability of resold in the respective horizon, estimated separately for period before and after 20 November 2010.

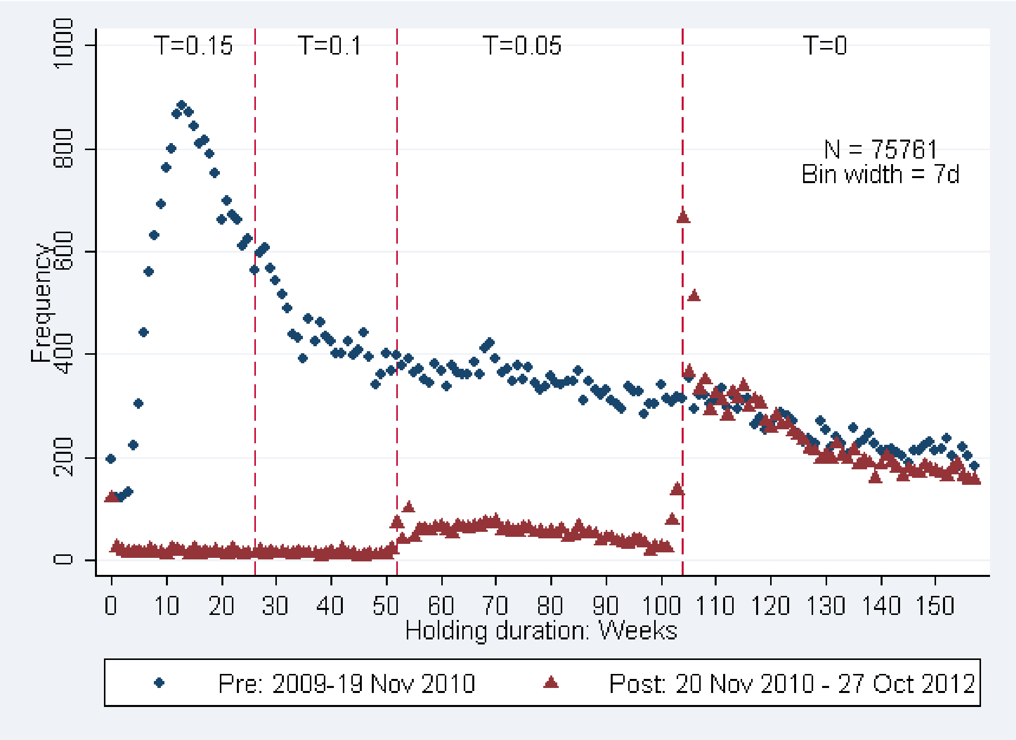
Figure 1.3: RD: Conditional average holding duration



Note: The graph plots the holding duration for residential properties in Hong Kong purchased in the window of 20 Aug 2010 to 20 Feb 2011, conditional on the properties being resold before Dec 2015. The dotted lines are fitted linear trend in the conditional holding duration in the respective horizon, estimated separately for period before and after 20 November 2010.

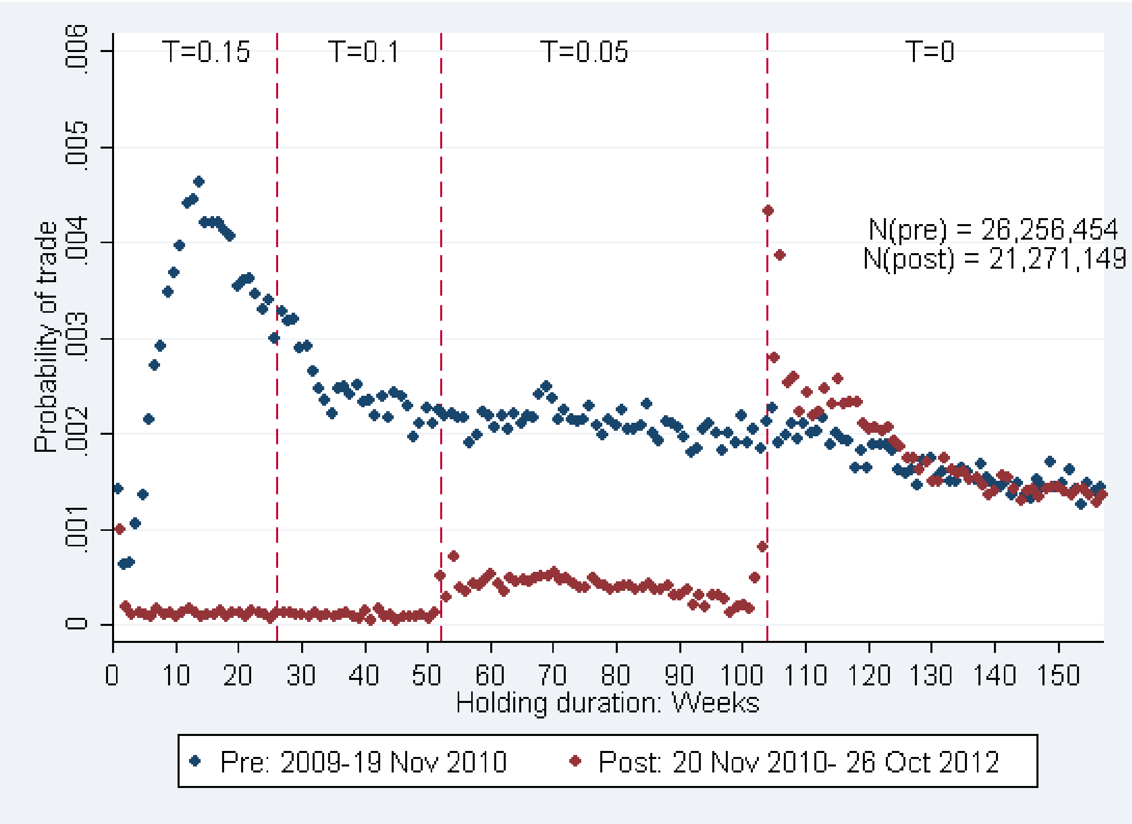
Figure 1.4: Duration of residential property holdings: 2009-2015

(a) Purchased before and under SSD phase 1



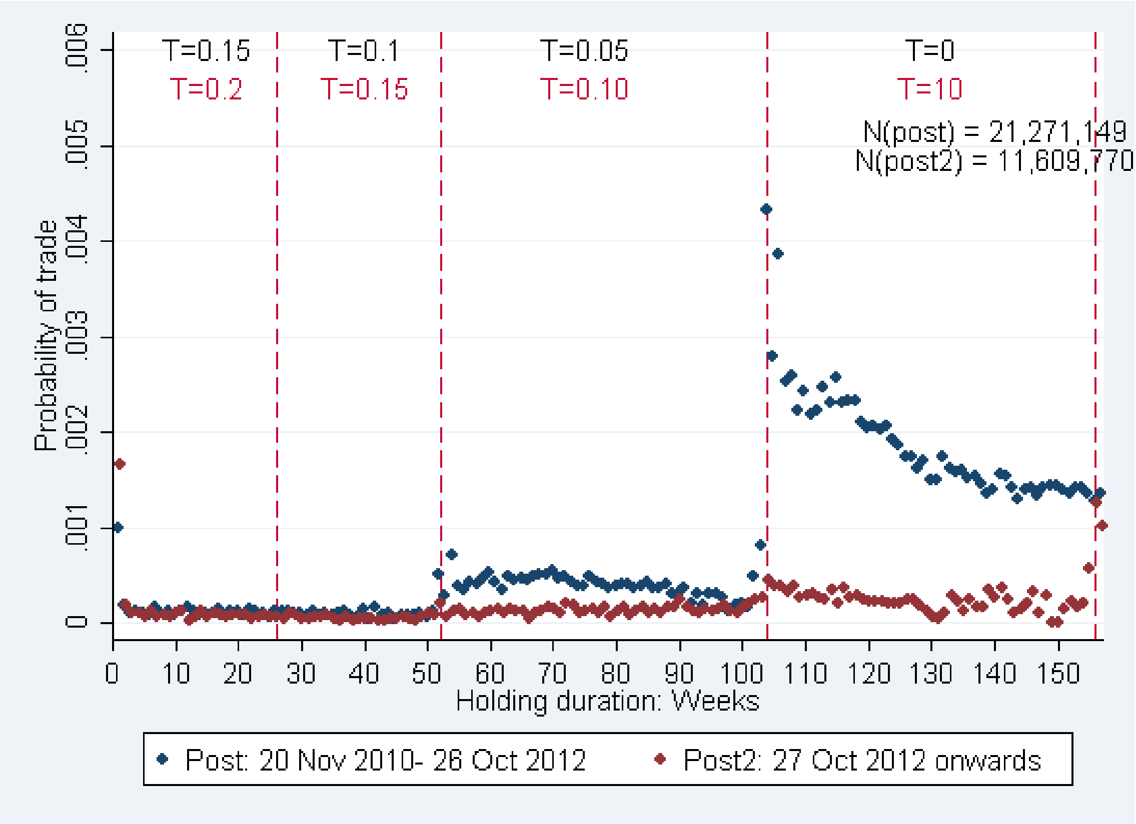
Notes: The graphs plot the distribution of holding duration for properties obtained in 2009 in Hong Kong, between 1 to 158 weeks. Each dot represent the frequency count of a bin width of 7 days period. The blue/circle dots plot the density of the holding duration for those obtained between January 2009 and 20 November 2010, where no tax rate is applied as function of holding duration. The red/triangle dots plot in addition the density of the holding duration for those obtained between 20 November 2010 to November 2012, where the 1st phase of special stamp duty applies. Sample includes all residential properties transactions as described in the data appendix.

Figure 1.5: Hazard function of trade: holding duration by week



Notes: The graph plots the hazard function estimated using OLS of weekly dummies for the holding period, where the outcome variable is an indicator of a trade happening in week t, with a panel of trading record at each week since a property is obtained using the transaction history for properties. The blue dots plot the hazard rate at each week for properties purchased from January 2009 to 19 November 2010, and the red dots plot the hazard rate for properties purchased from 20 November 2010 to 26 October 2012. Trading date is adjusted to the date of the provisional contract first signed when applicable. The hazard functions are estimated separately for properties obtained in each period.

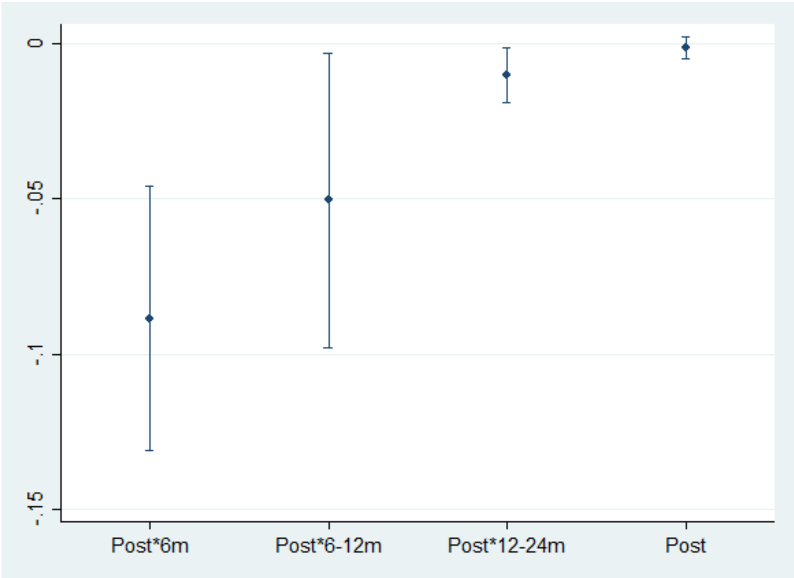
Figure 1.6: Hazard of trade: Special stamp duty phase 1 and phase 2, Hong Kong



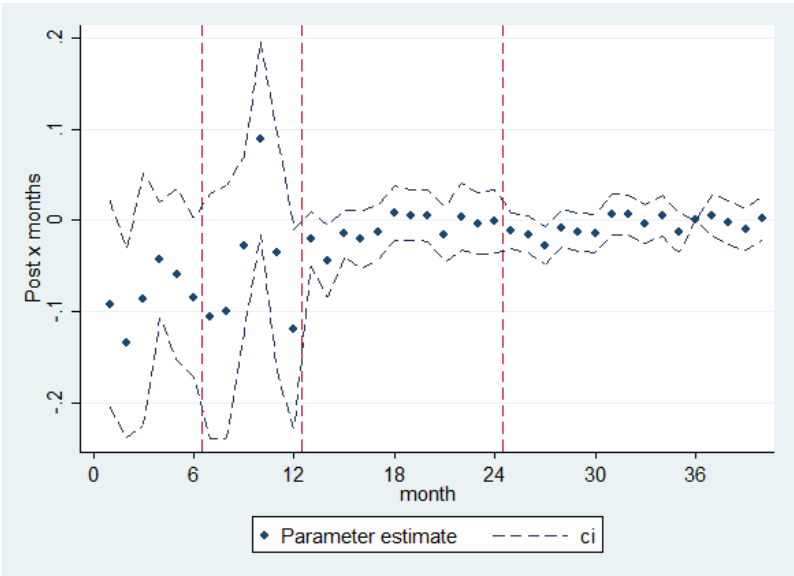
Notes: See note of Figure 1.5. The graph plots the hazard rate for properties purchased between 20 Nov 2010 to 27 Oct 2012 and from 28 Oct 2012 to 31 Dec 2015.

Figure 1.7: Pre-tax price regression: coefficient plots

(a) Broad category



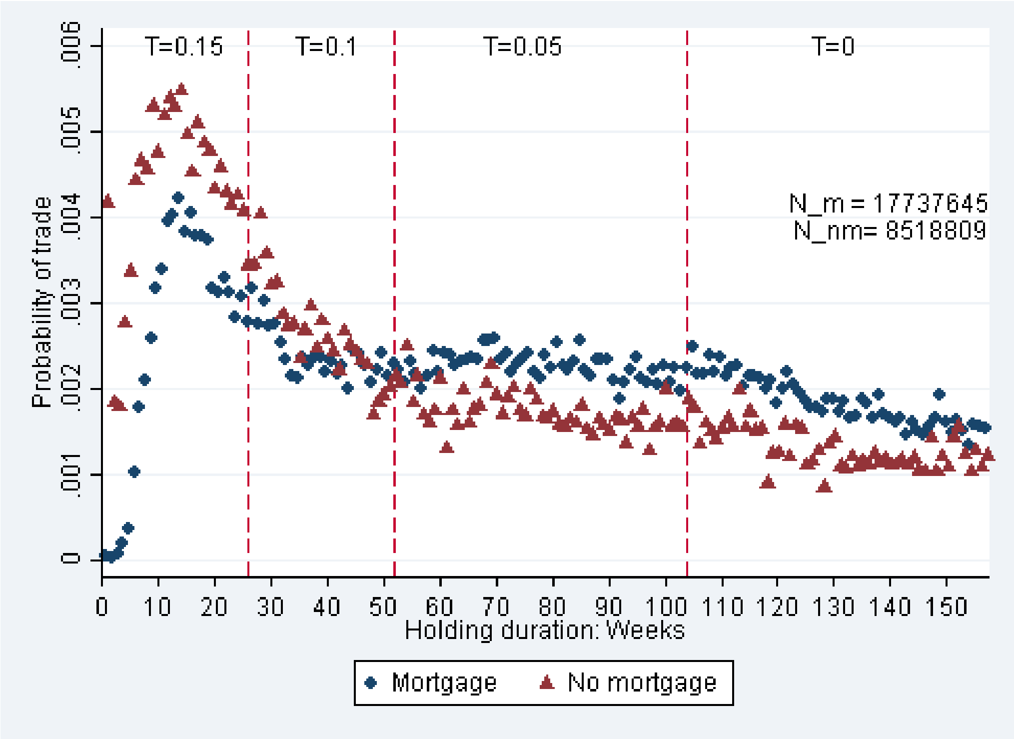
(b) Monthly non-parametric



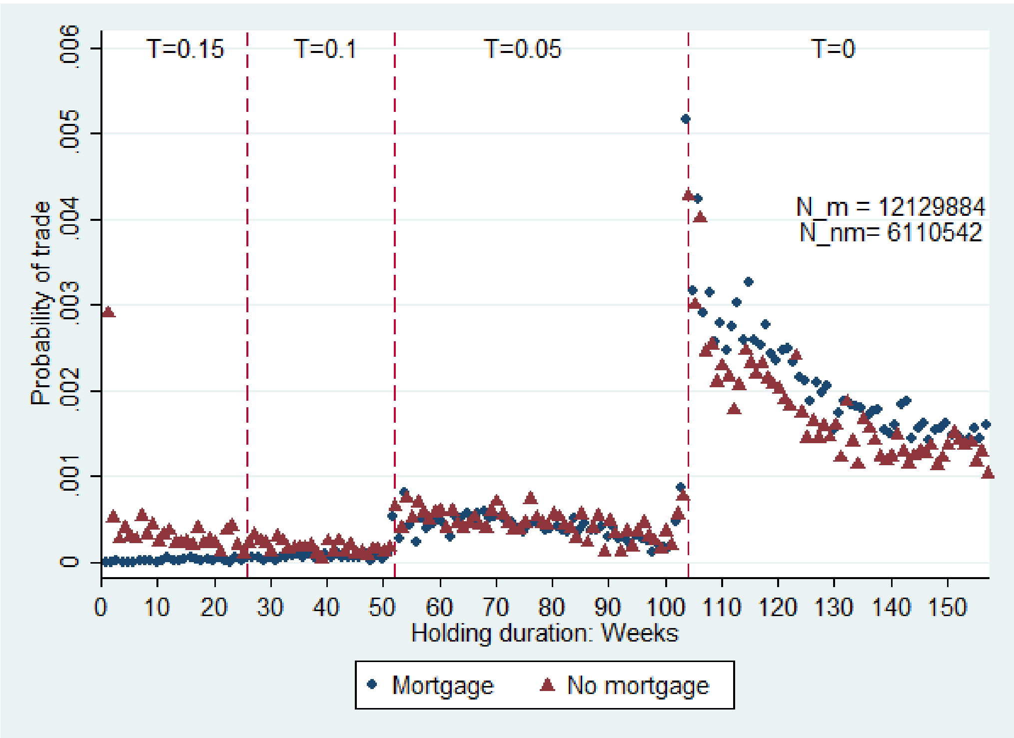
Notes: Panel (a) - The graph plot the $post * H$ coefficients of the price regression in column 2 of Table 1.5; the coefficient of $post * H_{i,j}$ indicate that the property is sold within the i and j month of holding the property; Outcome variable is log of pre-tax price; Controls includes holding duration FE in broad group, date of sales FE (Year-month), 2nd order polynomial of log of rateable values and its interaction with year-month dummies; Panel (b) - The graph plots the $post * i^{th} months$ of holding duration coefficients of the price regression in column 2 of Table 1.5 in the flexible specification; the coefficient of $post * i^{th} month$ indicate that the property is sold on the i^{th} month of holding the property; Outcome variable is log of pre-tax price; Controls includes holding duration FE in broad group, date of sales FE (Year-month), 2nd order polynomial of log of rateable values and its interaction with year-month dummies.

Figure 1.8: Hazard function of trade: by mortgage buyer and cash buyer

(a) Hazard of trade: pre

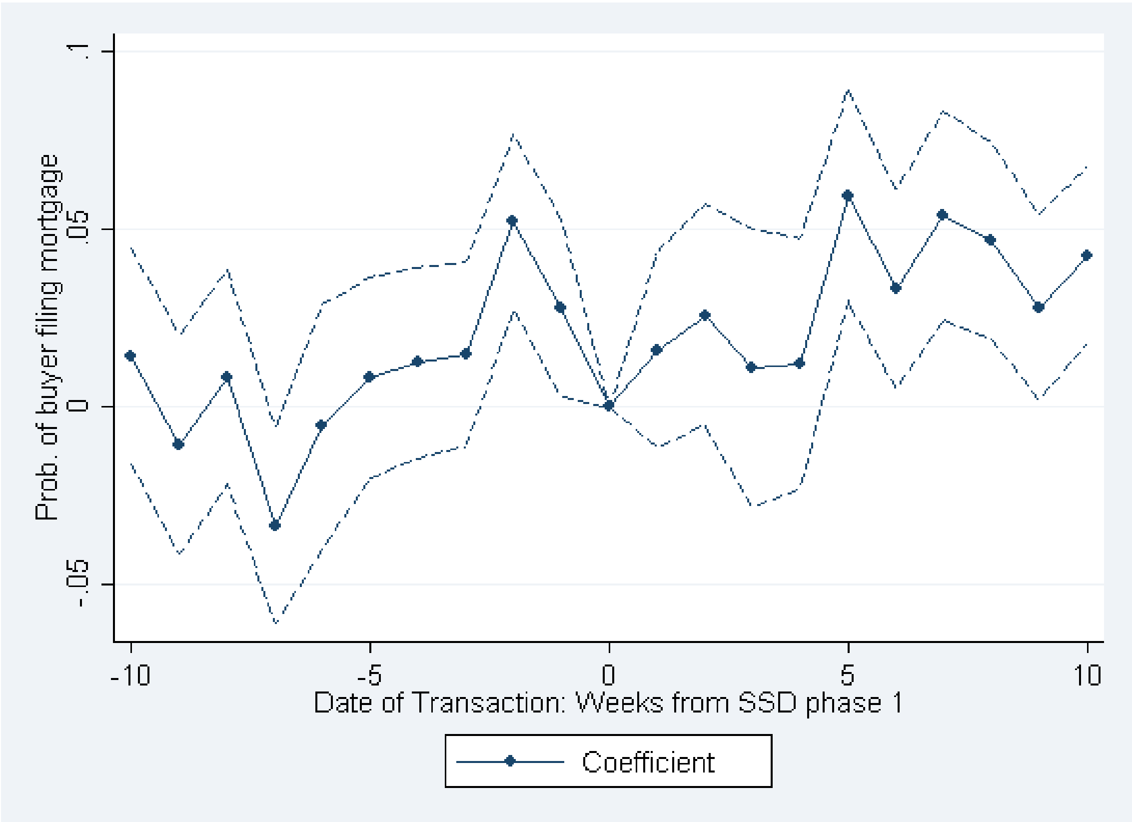


(b) Hazard of trade: post



Notes: See note of Figure 1.5. Panel (a) present the estimate for the hazard function for transaction spell that has an associated mortgage record and those that do not have an associated mortgage record. Panel (b) present the estimate for the hazard function for properties purchased under Special Stamp Duty phase 1, between 20 November 2010 to 26 October 2012.

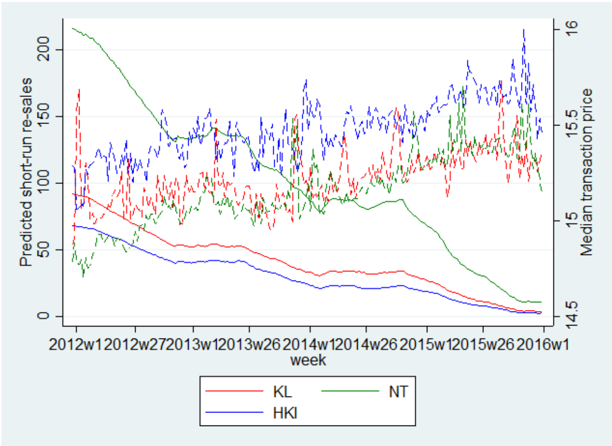
Figure 1.9: Effect of SSD on selection of type of buyers: event study



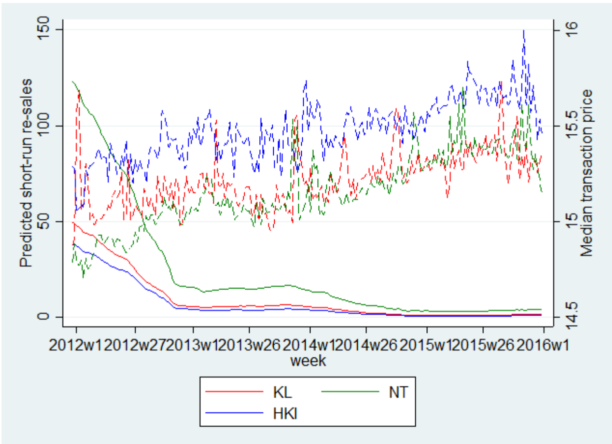
Notes: The graph presents estimates for equation 1.4. Sample include transaction in 20 Aug 2010 to Jan 20 March 2011 for properties purchased with price below HKD 6800000. Control includes districts fixed effects.

Figure 1.10: Predicted number of short-run resales

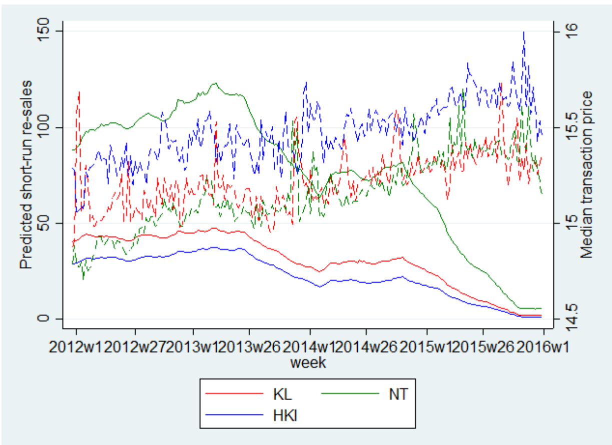
(a) All



(b) Midterm

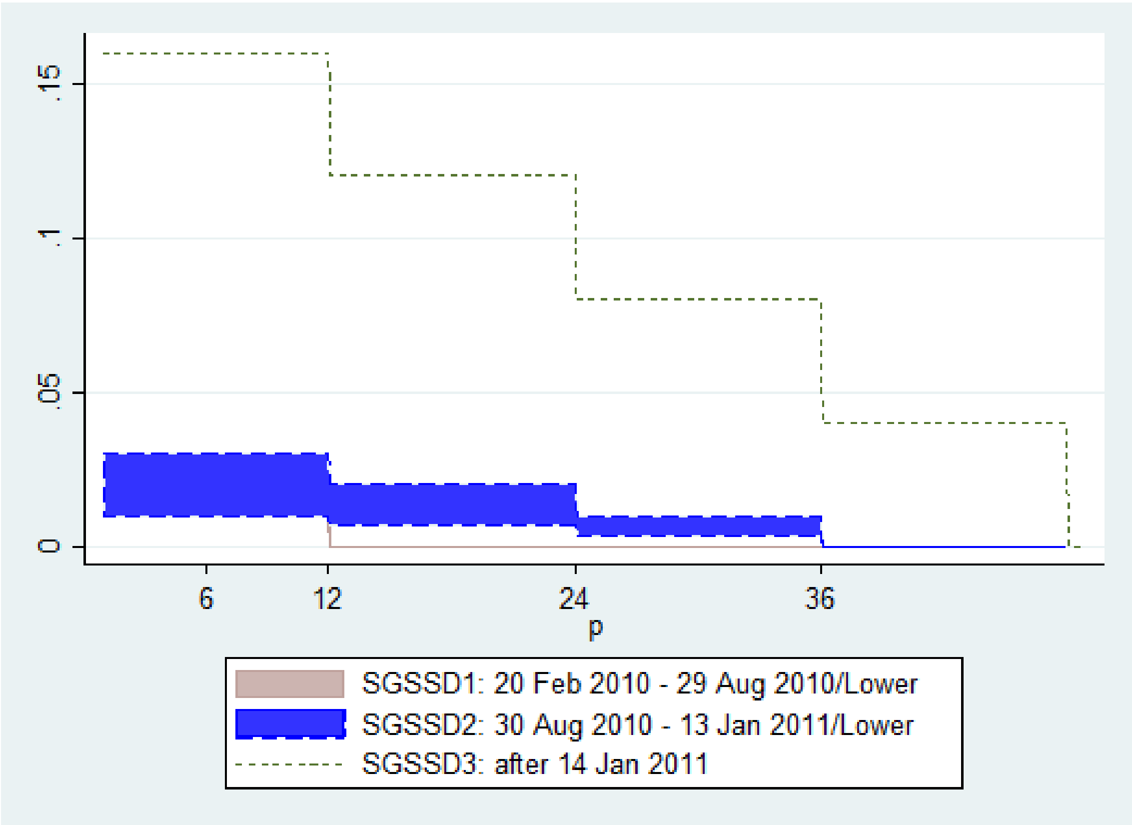


(c) Longterm



Notes: The graph plots the predicted number of sales from properties purchased in the past 3 years, using the number of transaction in jth week before time t, weighted by the estimates of hazard function w_j, according to the tax liability of $t - j$. It cover data from the December 2011 to Dec 2015. The dash-line plot the median (log) transaction price in the district in week t.

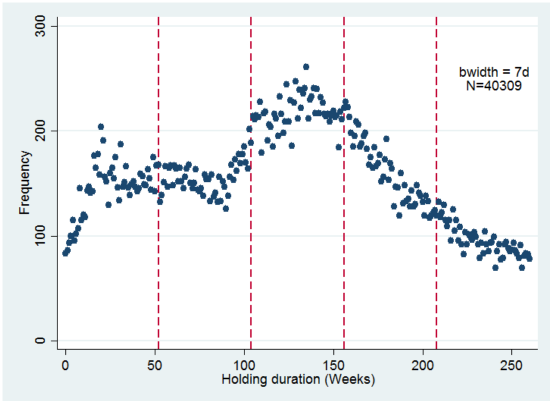
Figure 1.11: Tax on holding duration: Singapore



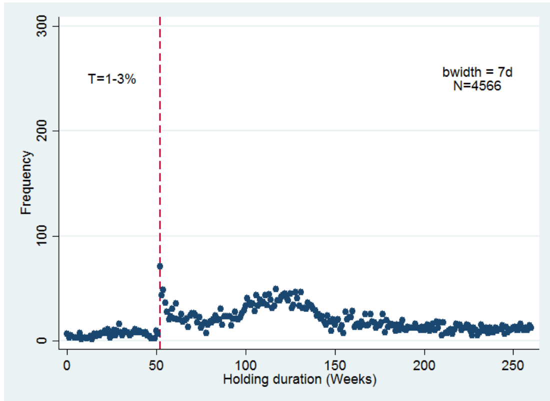
Notes: The graph plots the tax rate for residential properties under the Seller's Stamp Duty as function of holding period, for properties purchased in different period in Singapore from Phase 1 to Phase 3. In Phase 1 and 2 the Seller's Stamp Duty to be paid follows the rate at which the marginal rate is $\{1, 2, 3\}\%$ in the three brackets of $\{\leq 180000, 180000 - 360000, \geq 360000\}$ SGD in the 1st year; for Phase 2 it follows the same tax structure of rate $\{\frac{2}{3}, \frac{4}{3}, 2\}\%$ in the 2nd year and $\{\frac{1}{3}, \frac{2}{3}, 1\}\%$ in the 3rd year.

Figure 1.12: Holding duration: Singapore

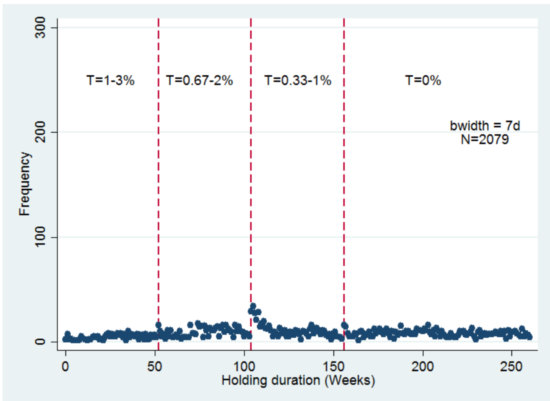
(a) Purchased from 2006 to 20 Feb 2010



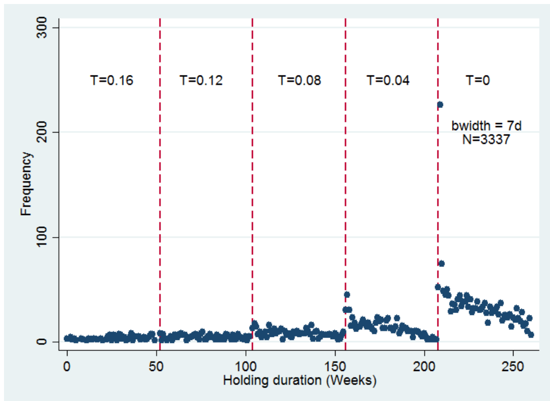
(b) Purchased between 20 Feb 2010, 29 Aug 2010



(c) Purchased between 29 Aug 2010, 13 Jan 2011



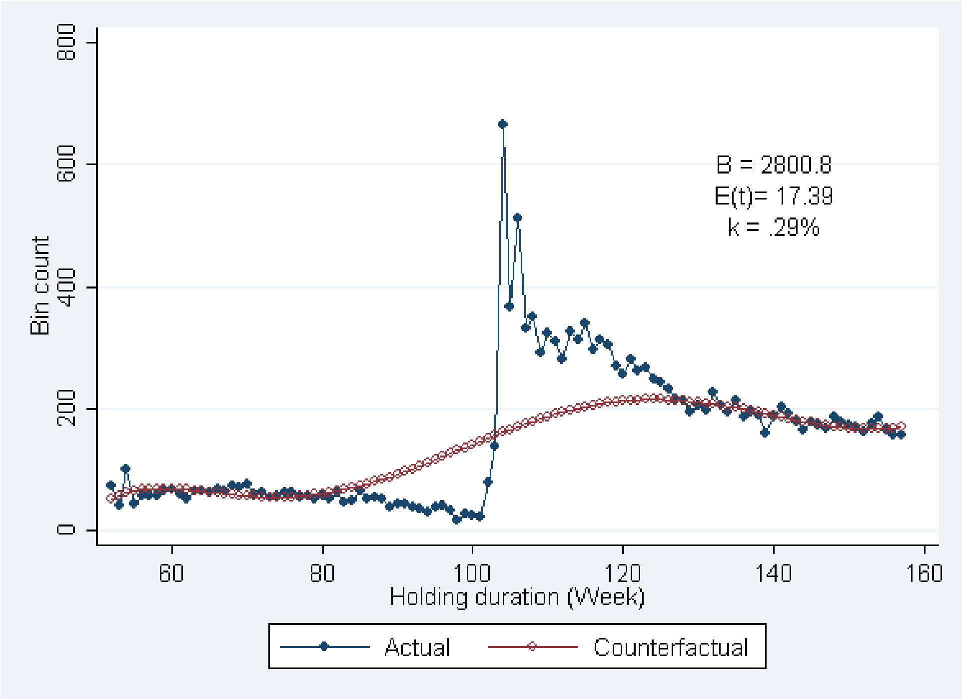
(d) Purchased on or after 14 Jan 2011-31 Sep 2016



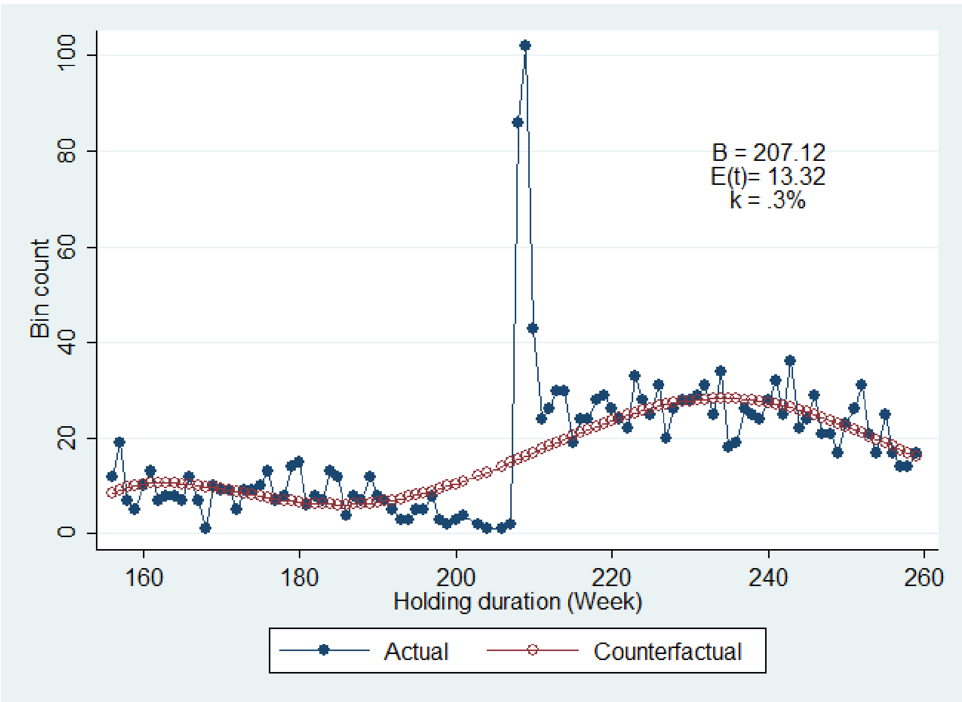
Notes: The graph shows the holding duration distribution of residential units in Singapore from same day to 5 years. Each point shows the frequency of holding at 7 days bin.

Figure 1.13: Estimation of average cutoff of waiting time

(a) Hong Kong SSD phase 1: 2 years notch



(b) Singapore SGSSD phsae 3: 4 years notch



Notes: Panel (a) - The graph plot the frequency of holding period at the bin of week same as Figure 1.5, for residential properties purchased between 20 Nov 2010 to 27 Oct 2012 in Hong Kong, and from 52th to 158th week. The data excluded are from 92th week to 124th week. The excess mass are defined as mass above the counterfactual density from 104 to 124th week. Panel (b) - The graph plot the frequency of holding period at the bin of week same for residential properties purchased between 14 Jan 2011 to 14 Jan 2012 in Singapore. Excess mass defined from 208th week to 213th week.

1.12 Tables

Table 1.1: RDD: HK SSD Phase 1

	Outcome: Probability of resold in			
	3 years	2 years	1 year	6 months
	(1)	(2)	(3)	(4)
RD estimate=1	-0.133***	-0.168***	-0.110***	-0.0743***
	(0.00934)	(0.00947)	(0.00477)	(0.00358)
Observations	52137	52137	52137	52137
Baseline mean	0.290	0.225	0.135	0.0759

Notes: Sample includes residential properties purchased from 20 Aug 2010-20 Feb 2011 in Hong Kong. Cut off is defined for purchases before or after 20 November 2010, the specification allow differential linear trend function for properties purchased before and after 20 November 2010. Column (1) includes all transactions; Column (2) includes transactions that has no mortgage record associated; Column (3) includes transactions that has mortgage record associated. Standard errors are clustered at the level of day of purchase.

Table 1.2: RDD: HK SSD Phase 1, by type of purchases

	Outcome: Resold in 3 years		
	All	No mortgage	Mortgage
	(1)	(2)	(3)
RD estimate=1	-0.133***	-0.166***	-0.116***
	(0.00934)	(0.0160)	(0.00968)
Observations	52137	16687	35450

Notes: Sample includes residential properties purchased from 20 Aug 2010-20 Feb 2011 in Hong Kong. Cut off is defined for purchases before or after 20 November 2010, the specification allow differential linear trend function for properties purchased before and after 20 November 2010. Column (1) includes all transactions; Column (2) includes transactions that has no mortgage record associated; Column (3) includes transactions that has mortgage record associated. Standard errors are clustered at the level of day of purchase.

Table 1.3: RDD: HK SSD Phase 1, conditional holding duration

	Outcome: Cond. holding duration		
	All	No mortgage	Mortgage
	(1)	(2)	(3)
RD estimate=1	334.1***	371.3***	314.2***
	(18.98)	(31.35)	(21.77)
Observations	15650	4731	10919
Baseline mean	733.9	649.0	772.9

Notes: Outcome variable is the holding duration in days conditional on the property is resold before 31st Dec 2015. Sample includes residential properties purchased from 20 Aug 2010-20 Feb 2011 in Hong Kong. Cut off is defined for purchases before or after 20 November 2010, the specification allow differential linear trend function for properties purchased before and after 20 November 2010. Column (1) includes all transactions; Column (2) includes transactions that has no mortgage record associated; Column (3) includes transactions that has mortgage record associated. Standard errors are clustered at the level of day of purchase.

Table 1.4: RDD: Singapore SSD Phase 3

	Outcome: Resold in 5 years		
	All	HDB	Private
	(1)	(2)	(3)
RD estimate	-0.0577***	-0.0866***	-0.0382**
	(0.0144)	(0.0199)	(0.0181)
Observations	6271	2487	3784

Notes: Sample includes residential properties purchased from 15 Dec 2010-15 Feb 2011 in Singapore. Cut off is defined at purchased after 14 Jan 2011, the specification allow differential linear trend function for properties purchased before and after 14 Jan 2011. Column (1) includes all transactions; Column (2) includes transactions where the buyer address is from HDB housing; Column (3) includes transactions where the buyer address is from private address. Standard errors are clustered at the level of day of purchase.

Table 1.5: Estimation of tax incidence

	Ln of pre-tax price	
	(1)	(2)
post=1	-0.105***	-0.00137
	(0.00469)	(0.00175)
post=1 × sold < 6m=1	0.273***	-0.0885***
	(0.0477)	(0.0218)
post=1 × sold 6-12m=1	0.147***	-0.0505**
	(0.0426)	(0.0242)
post=1 × sold 1-2 years=1	-0.00948	-0.0103**
	(0.0127)	(0.00452)
Observations	97017	94503
Year-month of sales FE	Y	Y
ln Rateable value*month FE	N	Y
ln Rateable value^2*month FE	N	Y

Notes: Sample include properties bought between Jan 2009 and 26 Oct 2012, and are sold before 31 December 2015. *post* is an indicator for property bought between 20 Nov 2010 to 26 Oct 2012. Column (2) includes 2nd order polynomial of log rateable values and its interaction with date of sales (Year-month) dummies. Standard errors are clustered at property level;

Table 1.6: Robustness of price regression: polynomial

	(1)	(2)	(3)	(4)	(5)
post=1	-0.105***	-0.00182	-0.00137	-0.000824	-0.00102
	(0.00469)	(0.00181)	(0.00175)	(0.00175)	(0.00175)
post=1 × sold < 6m=1	0.273***	-0.0507**	-0.0885***	-0.0988***	-0.0942***
	(0.0477)	(0.0225)	(0.0218)	(0.0220)	(0.0220)
post=1 × sold 6-12m=1	0.147***	-0.0306	-0.0505**	-0.0549**	-0.0566**
	(0.0426)	(0.0247)	(0.0242)	(0.0243)	(0.0243)
post=1 × sold 1-2 years=1	-0.00948	-0.00324	-0.0103**	-0.0115**	-0.0115**
	(0.0127)	(0.00478)	(0.00452)	(0.00449)	(0.00447)
Observations	97017	94503	94503	94503	94503
Degree of polynomial	0	1	2	3	4

Notes: See notes of Table 1.5. Column (2)-(5) controls for time varying polynomial of log rateable value of properties with 1-4 degree respectively. Standard errors clustered at property level.

Table 1.7: SSD phase 1: selection in types of buyers

	Outcome: Filing mortgage				
	Sample: before/after 20 Nov 2010				
	Sep10-Jan11	Aug10-Feb11	Jul10-Mar11	Jul09-Mar11	Jul09-Mar11
	(1)	(2)	(3)	(4)	(5)
After SSD 1=1	0.0170***	0.0266***	0.0142***	0.0292***	0.0552**
	(0.00535)	(0.00426)	(0.00361)	(0.00354)	(0.0266)
Observations	30645	46335	64637	164551	164551
Baseline mean	0.687	0.696	0.699	0.688	0.688
District FE	x	x	x	x	x
District*month of the year FE				x	x
Property FE					x

Notes: Outcome variable is an indicator of whether the transaction has an associate mortgage record. Column (1) contains sample of properties purchased from 20 Sep 2010 to 20 Jan 2010; Column (2) contains sample of properties purchased from 20 Aug 2010 to 20 Feb 2010; Column (3) contains sample of properties purchased from 20 Jul 2010 to 20 20 March 2010; Column (4) contains sample of properties purchased from 20 Jul 2009 to 20 20 March 2011; Sample includes residential properties purchased in the relevant period in Hong Kong. Standard errors clustered at property level.

Appendix

1.A Appendix of chapter 1

Figure 1.A.1: Timeline of the stamp duty tax policy change in Hong Kong since 2010

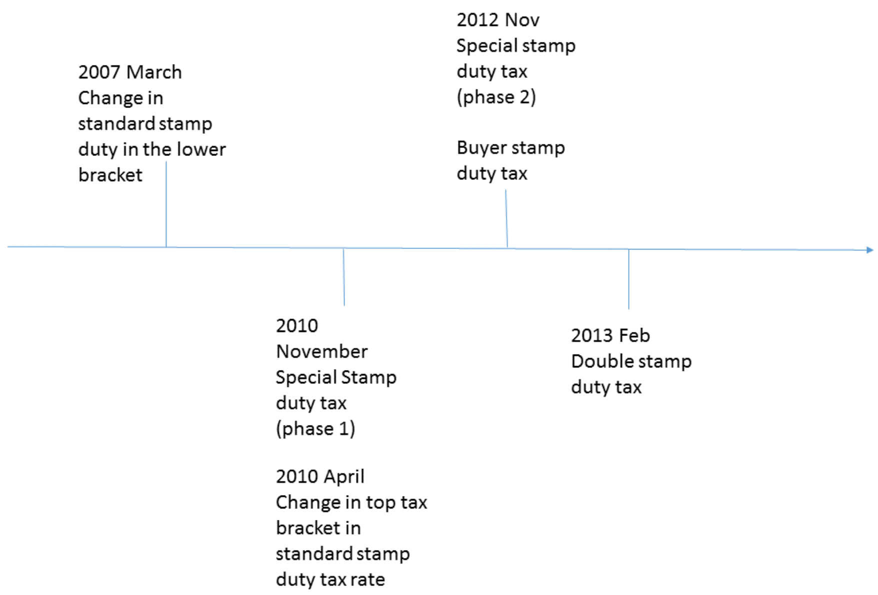


Figure 1.A.2: Timeline of the Seller's Stamp Duty policy change in Singapore since 2010

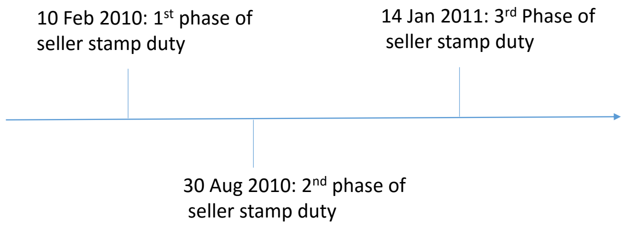
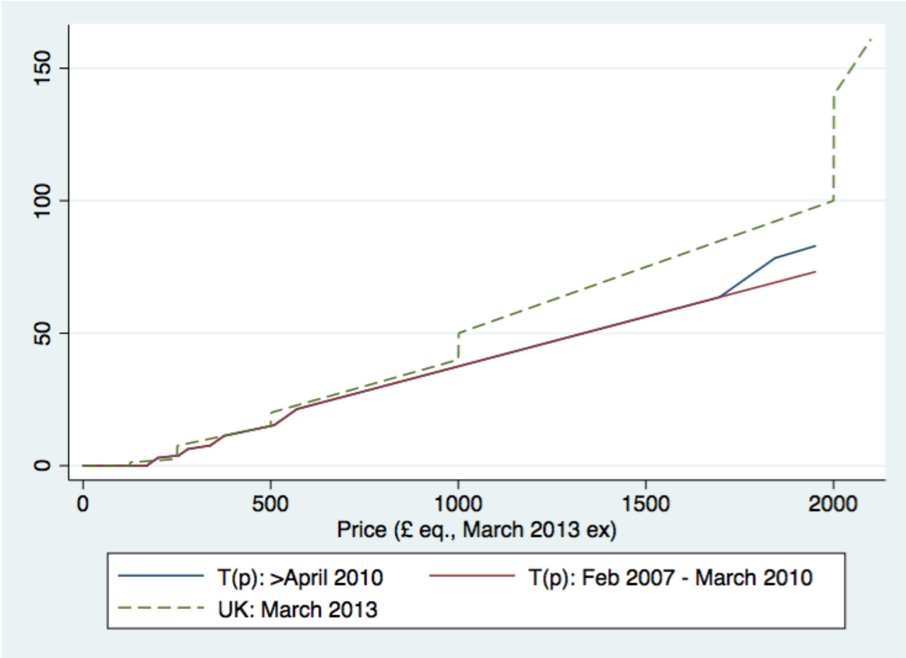
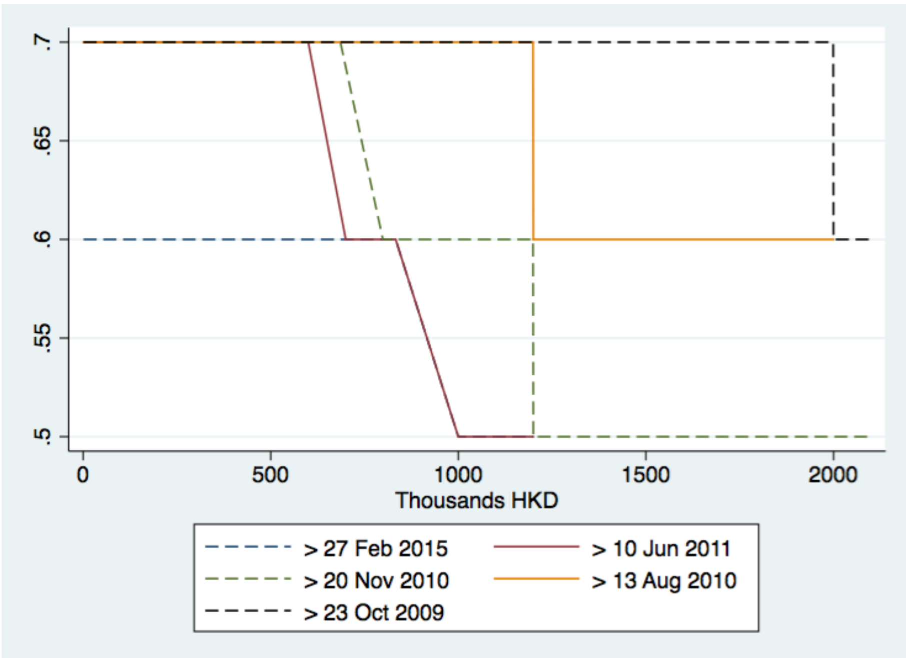


Figure 1.A.3: Property stamp duty tax schedule of Hong Kong in various time point



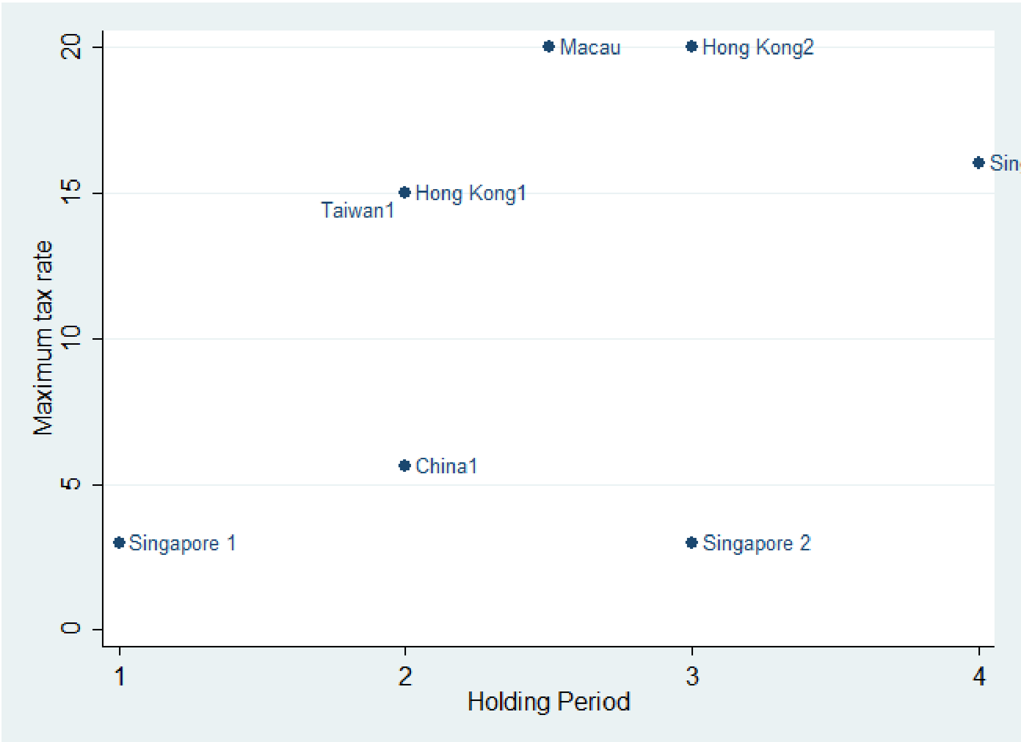
Note: Unit in thousands

Figure 1.A.4: Changes in mortgage loan-to-value limit ratio



Note: For residential properties

Figure 1.A.5: Use of transaction tax on holding period in Asia



Note: The graph plots the region/country that had implemented transaction tax on residential properties on holding period since 2010 in Asia, by the maximum transaction tax rate and when the last notch of the tax applies, of each phase of it being implemented.

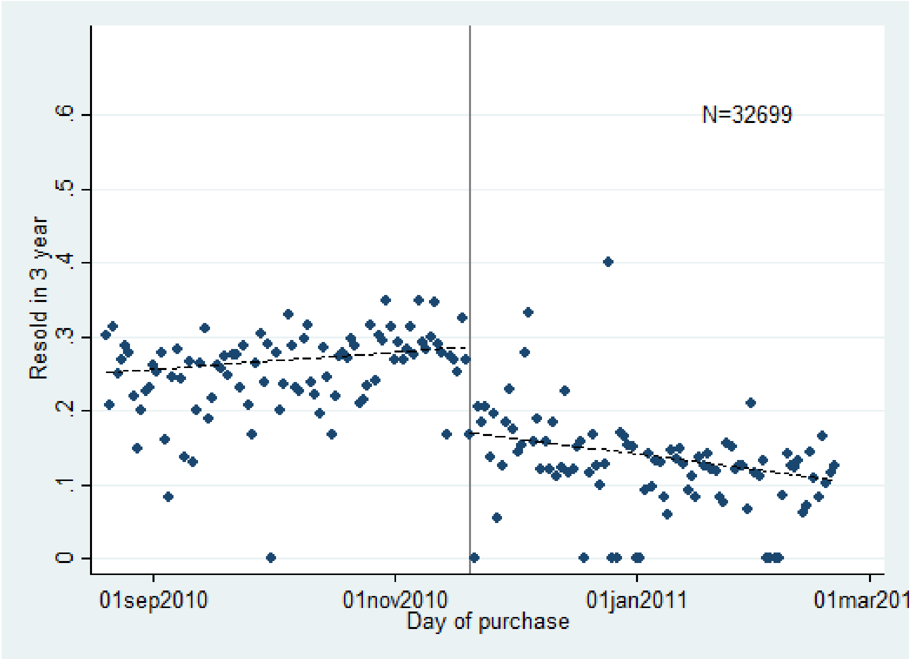
Table 1.A.1: Probability of selling within 2 years since purchases

	Probability of selling in:		
	6 months	1 year	2 years
	(1)	(2)	(3)
Mortgage	-0.0418***	-0.0461***	-0.0206***
	(0.00136)	(0.00170)	(0.00201)
Observations	201709	201709	201709
District FE	x	x	x
Week of purchases FE	x	x	x

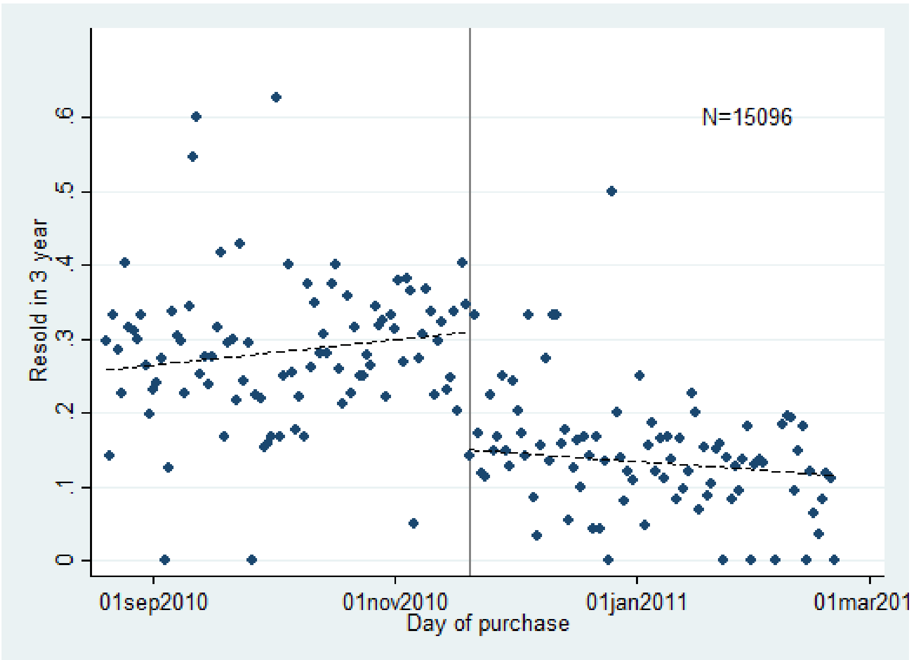
Notes: Sample includes residential properties purchased between January 2009 to 19 November 2010. Outcome variable is an indicator of the property being resold within jth month since purchase. $Mortgage$ is an indicator for a transaction having an associated mortgage record.

Figure 1.A.6: RDD plot: Probability of resold in 3 years - by mortgage record

(a) With mortgage

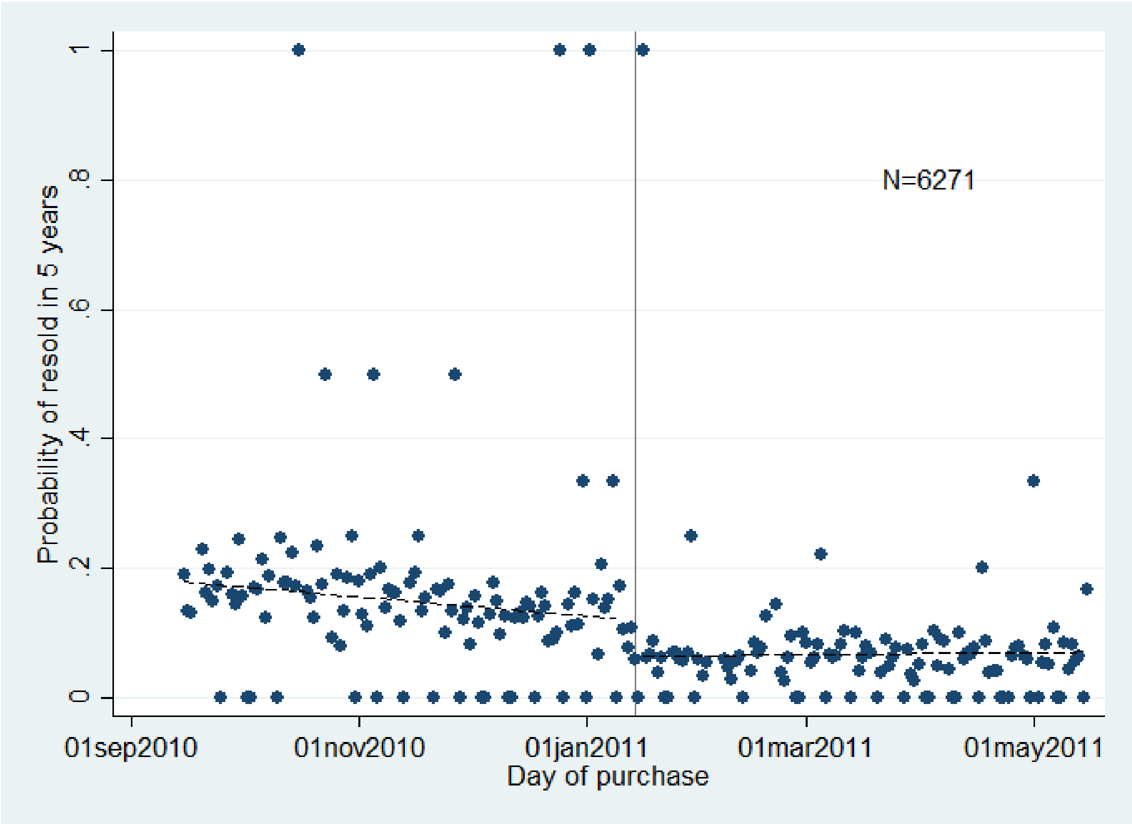


(b) Without mortgage



Note: See note of Figure 1.2. Panel (a) plot the holdings in which there is an associated mortgage record. Panel (b) plot the holdings in which there is no associated mortgage record.

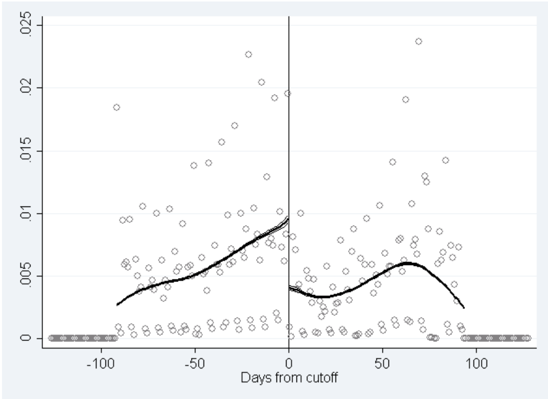
Figure 1.A.7: RDD: Singapore SSD Phase 3



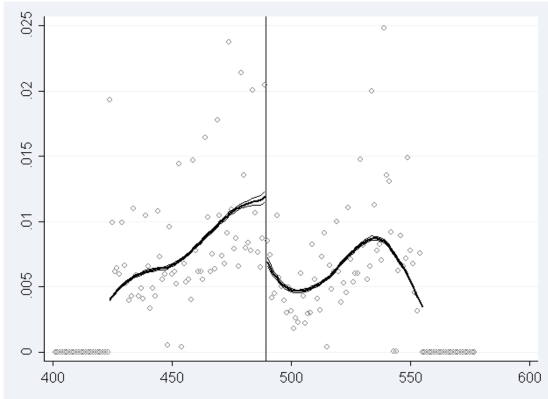
Note: The graph plots the probability of a property being resold in 5 years since its purchased, by the day they are last obtained. The vertical line represent 14 Jan 2011. The dotted line are estimated linear trend in probability of resold in 5 years, separately for properties obtained before and after 14 Jan 2011.

Figure 1.A.8: Density of trading dates

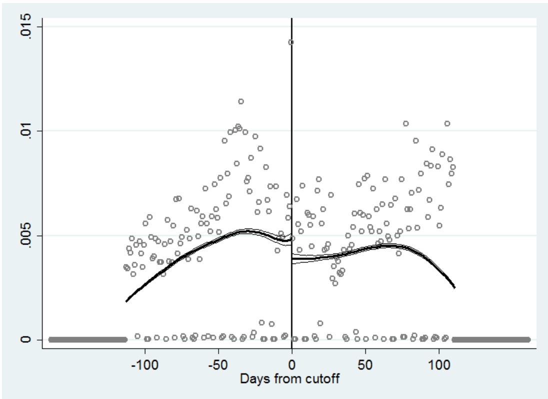
(a) HK: All days



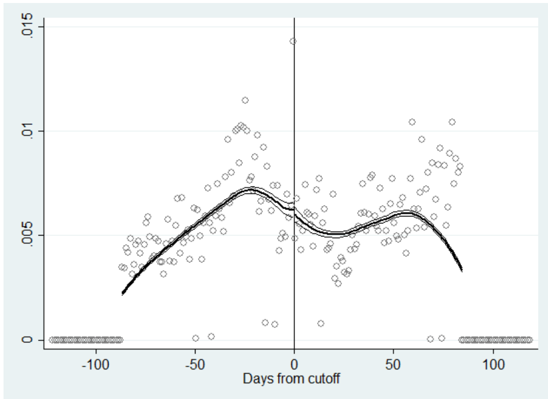
(b) HK: Excluding weekend



(c) SG: All days



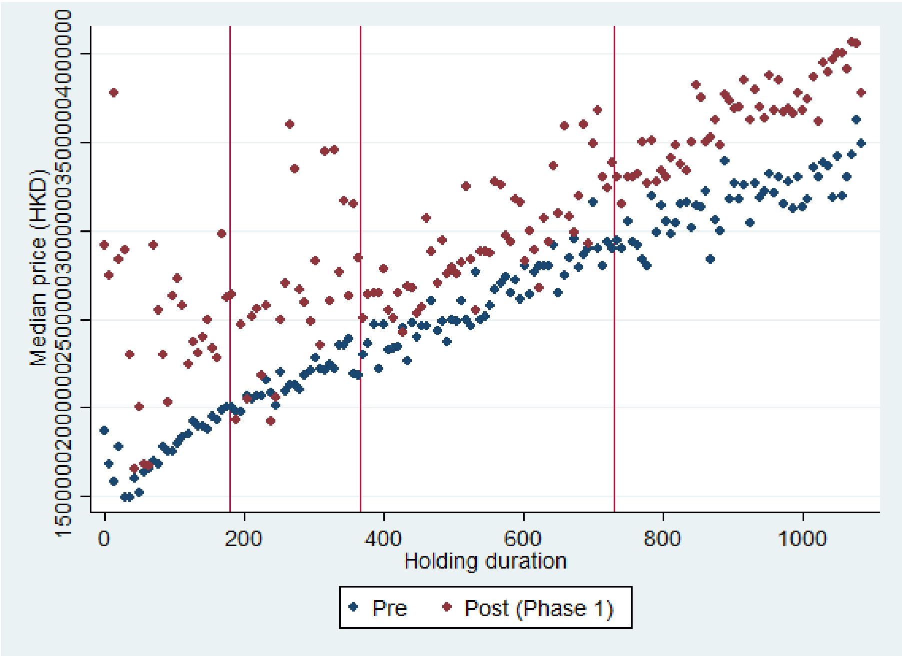
(d) SG: Excluding weekend



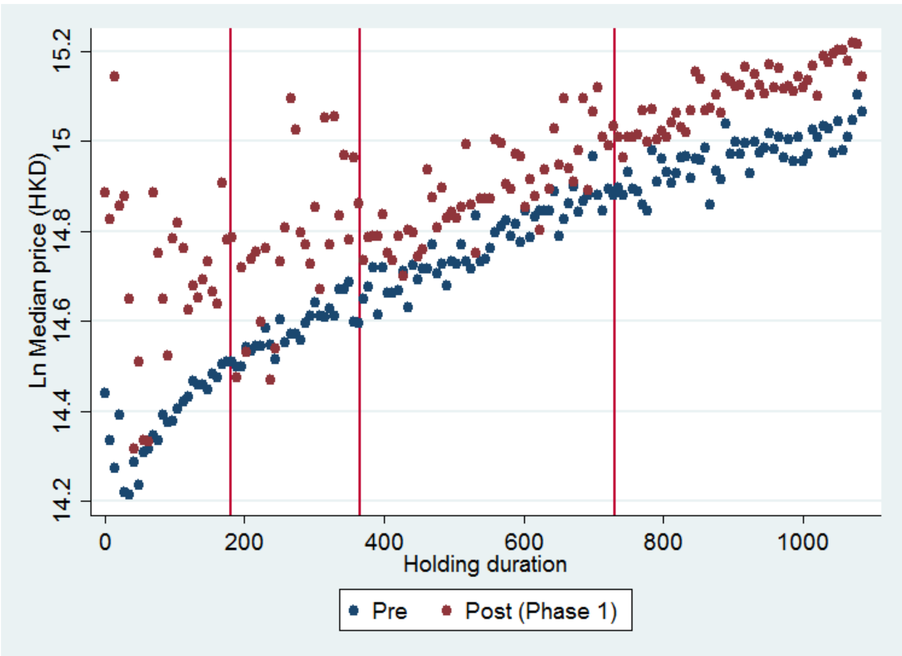
Panel (a) and (b) plot the density of transaction in Hong Kong by days from 20 Nov 2010, with all days and excluding weekends respectively; Panel (c) and (d) plot the density of transaction in Singapore by days from 20 Nov 2010, with all days and excluding weekends respectively.

Figure 1.A.9: Median price conditional on holding duration

(a) Median price

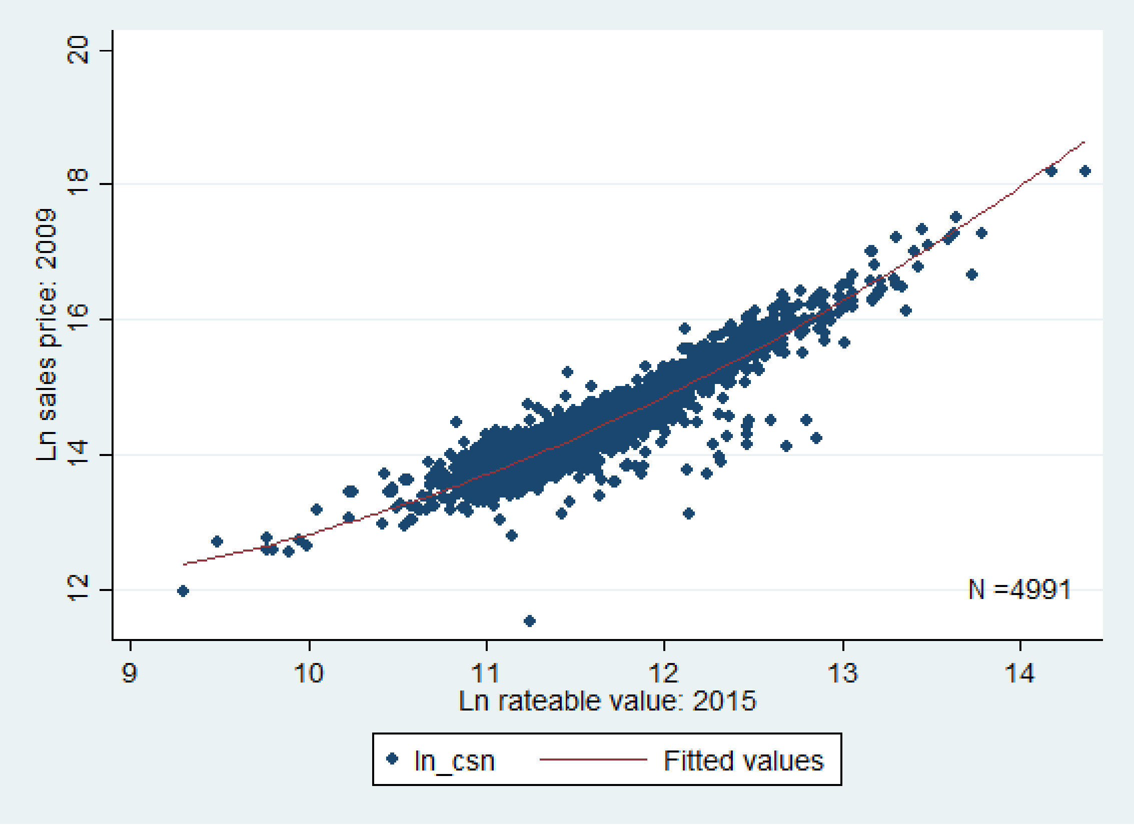


(b) Ln median price



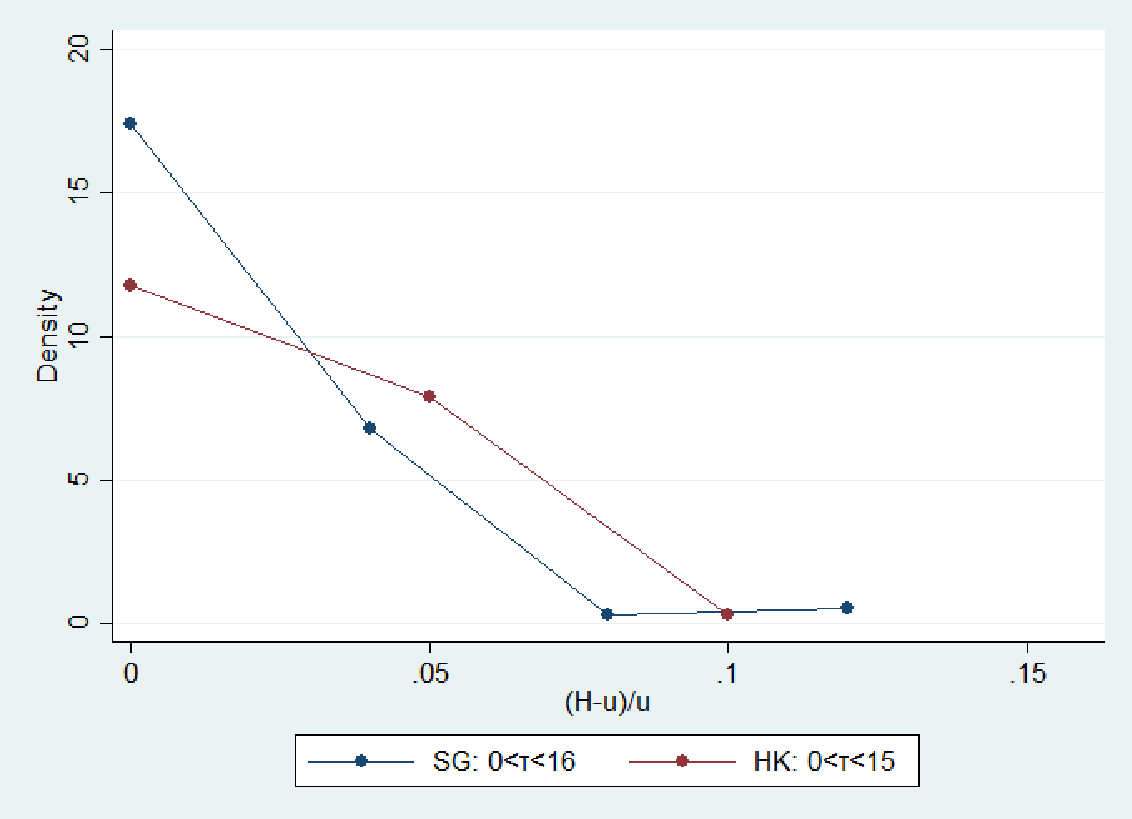
Notes: The graph plots the median price (current price) and median log price of residential properties traded at each holding period at week interval. The blue dots represent properties purchased between Jan 2009 to 19 Nov 2010. The red dots represent properties purchased between 20 Nov 2010 to 26 Oct 2012. The vertical lines represent holding period in which time notch is present in the Hong Kong Special Stamp Duty Phase 1, at 6 months, 1 year, 2 years intervals.

Figure 1.A.10: Ratio of transacted price and government estimated rental rate



Notes: The graph plots the relationship between transacted price and the government estimate rateable values measured in 2015, for properties purchased and traded in 2009. It also shows fitted line using 2nd order polynomial.

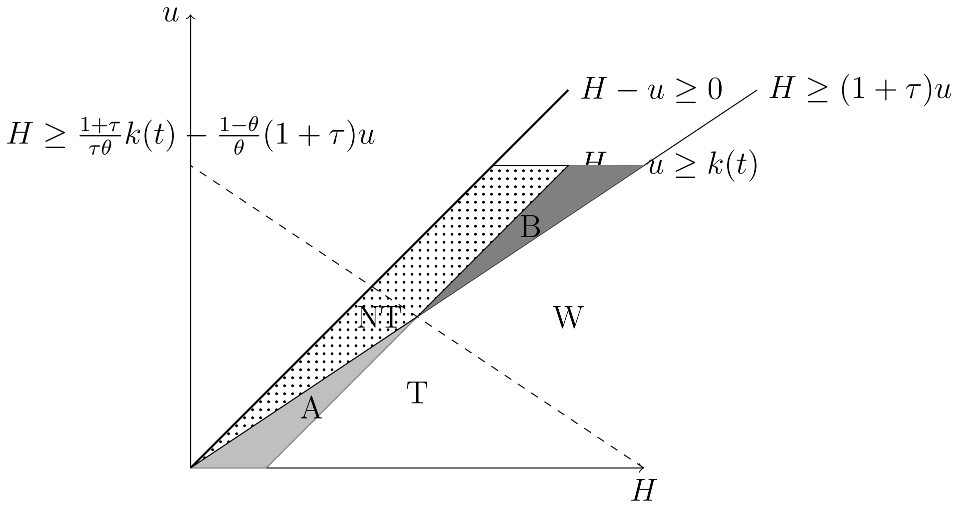
Figure 1.A.11: Estimated distribution of trading surplus



Note: The graph plots the estimated density of the trading surplus distribution using equation 1.16, using the 1st phase of Special Stamp Duty and the 3rd phase of the Seller's Stamp Duty, computed from the hazard rate of trade in the flat region in each of distinct period with different tax rate within the same tax scheme. The red line plot the point estimates of the trading surplus at $\{0, 0.050.1\}$ using the trading hazard of properties purchased under the Special Stamp Duty phase 1 in Hong Kong. The blue line plot the point estimates of the trading surplus at $\{0, 0.04, 0.08, 0.12\}$ using the trading hazard of properties purchased under the Seller's Stamp Duty phase 3 in Singapore.

1.A.1 Explanatory note for conceptual framework

Figure 1.A.12: Illustration of trade off between waiting and trade immediately



The area filled with dots is the strictly no trade area. The grey area represent the area in which trading would not happen in a universal transaction tax at rate τ, but with the time notch design bunching expand the possibility of trade; the light grey area represent the case in which trading immediately dominates the bunching response. Area W and T are trade would have happen, while all the cases in W would choose to wait until the time notch instead of trading immediately.

When a transaction tax τ is introduced, only matches with surplus higher than τ percent of the reservation value of seller are traded. This typically reduced the number of transactions (i.e. area NT and B on the graph). However, with a specified time notch, a newly created area B exist that would yield positive surplus if the buyer and seller wait until $t = \bar{t}$ to trade. The size of this area would depends on the utility cost of delaying transaction $k_b + k_v$.

The dotted line with negative slope indicate the trade off of bunching versus trade immediately even for those matches that would yield positive surplus if traded immediately. For matches that lies on the area W, both H and u are high, and thus the agreed price after bargaining would also be high, this implies that the proportional tax would be exceptionally costly for these transaction to happen. Therefore for matches in area W (and B) they would wait until time t to trade to evade the tax.

On the other hand, on the left of the downward sloping dotted line, H and u are smaller than those on the right, which implies a smaller price under Nash bargaining. The tax to evade is too small relative to the total unit cost of waiting $k_b + k_v$, and thus for matches in area T (and A), trade would happen immediately.

When there is an exogenous increase in τ, area A and T would decrease, but the cutoff $H - u \geq k(t)$ remains unchanged, and thus area of B (and also W) would indeed increase. Increase the tax rate would increase the number of people who wait until $\bar{t}$ to trade.

1.A.2 Proofs for the conceptual framework

Proposition 1

To see proposition 1, assume an arbitrary joint distribution of H and u - $g(h,u)$, and the density of match formation at time t - $f(t)$. Note that for every pair of H and u, we can map it into a price p. Then we can define the density of the conditional price distribution of p, the price distribution undistored by the tax conditional on $H - u > (k_b + k_v)(T - t) \cap H - u > 0$, as $P(p,k,t)$. Then there is a critical price where the amount of buncher is given by the mass of price where for each time t being greater than $p_d(\tau,t,k)$

$$p_d(\tau,t,k)\Delta\tau_t = (k_b + k_v)(\bar{t} - t) \tag{1.17}$$

Thus the amount of buncher is

$$B = \int_0^{\bar{t}} \int_{p_d(\tau,t,k)}^{\infty} P(p,k,t) Prob(H - u > (k_b + k_v)(T - t)) dp\, f(t) dt \tag{1.18}$$

$$\begin{aligned} \frac{dB}{d(k_v + k_b)} = & \int_0^{\bar{t}} -P(p_d(H,u,\tau,t,k)) \frac{dp_d}{d(k_v + k_b)} f(t) dt + \\ & \int_0^{\bar{t}} \int_{p_d(\tau,t,k)}^{\infty} P(p,k,t) \frac{dProb(H - u > (k_b + k_v)(T - t))}{d(k_v + k_b)} dp\, f(t) dt \end{aligned} \tag{1.19}$$

where $\int_0^{\bar{t}} -P(p_d(H,u,\tau,t,k)) \frac{dp_d}{d(k_v+k_b)} f(t)dt < 0$ as $\frac{dp_d}{d(k_v+k_b)} > 0$, and $\frac{dProb(H-u>(k_b+k_v)(T-t))}{d(k_v+k_b)} dp < 0$. Thus if $k_v + k_b$ is smaller, then the amount of bunchers is larger.

Propositions 2

We prove the case where $\tau_{\bar{t}} = 0$

Define $(1+\tau)u = L(\tau, u)$ and $\frac{1+\tau}{\tau\theta}k(t) - \frac{1-\theta}{\theta}(1+\tau)u) = U(\tau, t, u)$ and some arbitrary joint distribution of H and u - $g(h, u)$
Probability of trade at time t =

$$Prob(L(\tau, U) < H < U(\tau, t, \theta, u))$$

$$\int_0^{\frac{k(t)}{\tau}} \int_{L(\tau,U)}^{U(\tau,t,\theta,u)} g(h, u)dhdu \tag{1.20}$$

Since the upper bound of the two integrals are decreasing in t, where the joint density is strictly positive and independent of t, we have $\frac{dProb(L(\tau,U)<H<U(\tau,t,\theta,u))}{dt} < 0$

Probability of forgone trade at time t =

$$\int_{\frac{k(t)}{\tau}}^{\infty} \int_u^{k(t)+u} g(h, u)dhdu + \int_0^{\frac{k(t)}{\tau}} \int_u^{(1+\tau)u} g(h, u)dhdu \tag{1.21}$$

where $\frac{dProb(forgone)}{dt} = \int_{\frac{k(t)}{\tau}}^{\infty} g(k(t) + u, u)k'(t)dhdu < 0$
which is decreasing with t

1.A.3 Explanatory note on the bunching estimates

Define the distribution of timing of arrival of a match (q) for a seller is $L(t) = Prob(q \leq t)$, then the density of the time of sales is

$$
\begin{aligned}
P(T \leq t) =& \text{Prob}(q \leq t)\text{Prob}(H - u > 0) \\
=& L(t) \int_{H-u>0} dG(H, u) \\
=& F(t)
\end{aligned}
$$

Define f(t) as the pdf of F(t)

$$f(t) = l(t) \int_{H-u>0} dG(H, u)$$

and g(t) as

$$g(t) = l(t) \int_{H-u>\tau u} dG(H, u)$$

Then the expected waiting duration could be estimated using $\hat{B}$,$\hat{f}(\bar{t})$ and $\hat{g}(\bar{t})$

$$
\begin{aligned}
B =& P(T = t) \in \{\underline{B}, \bar{B}\} \\
\bar{B} =& \int_0^{\bar{t}} \text{Prob}(q \geq t(H,u) \cap H - u > 0) dL(q) \\
=& \int_0^{\bar{t}} \int_{q \geq t(H,u) \cap H-u>0} dG(H,u) dL(q) \\
=& \int_0^{\bar{t}} \int_{H-u>0} I[q \geq t(H,u)] dG(H,u) dL(q) \\
=& \int_{H-u>0} \int_0^{\bar{t}} I[q \geq t(H,u)] dL(q) dG(H,u) \\
=& \int_{H-u>0} \int_{t(H,u)}^{\bar{t}} dL(q) dG(H,u) \\
\approx& \int_{H-u>0} \int_{t(H,u)}^{\bar{t}} l(\bar{t}) dq dG(H,u) \\
=& \int_{H-u>0} \int_{t(H,u)}^{\bar{t}} dq dG(H,u) l(\bar{t}) \\
=& \int_{H-u>0} [\bar{t} - t(H,u)] dG(H,u) l(\bar{t}) \\
=& \int_{H-u>0} [\bar{t} - t(H,u)] \frac{dG(H,u)}{\int_{H-u>0} dG(H,u)} l(\bar{t}) \int_{H-u>0} dG(H,u) \\
=& E(\bar{t} - t(H,u) | H - u > 0) l(\bar{t}) \int_{H-u>0} dG(H,u) \\
=& E(\bar{t} - t(H,u) | H - u > 0) f(\bar{t})
\end{aligned}
$$

and

$$
\begin{aligned}
\underline{B} =& \int_0^{\bar{t}} \text{Prob}(q \geq t(H,u) \cap H - u > \tau u) dL(q) \\
=& E(\bar{t} - t(H,u) | H - u > \tau u) g(\bar{t})
\end{aligned}
$$

This implies

$$
E(\bar{t} - t(H,u) | H - u > 0) \geq \frac{\hat{B}}{f(\bar{t})}
$$

and

$$E(\bar{t} - t(H,u)|H - u > \tau u) \leq \frac{\hat{B}}{g(\bar{t})}$$

And as $p(\tau, H, u)\tau = \bar{t} - t(H,u)$

$$\begin{aligned}
E(\bar{t} - t(H,u)|H - u > \tau u) &= E(p(\tau, H, u)\tau|H - u > \tau u) \\
&= \tau E(p(\tau, H, u)|H - u > \tau u) \\
&= \tau \frac{E(p(\tau, H, u)|H - u > \tau u)}{E(p|H - u > 0)} * E(p|H - u > 0) \\
&\geq \tau \frac{E(p(\tau, H, u)|H - u > \tau u)}{E(p|H - u > 0)} * E(p(\tau, H, u)|H - u > 0) \\
&= \frac{E(p(\tau, H, u)|H - u > \tau u)}{E(p|H - u > 0)} * E(\tau p(\tau, H, u)|H - u > 0) \\
&= \frac{E(p(\tau, H, u)|H - u > \tau u)}{E(p|H - u > 0)} * E(t - t(H, u)|H - u > 0) \\
&\geq \frac{E(p(\tau, H, u)|H - u > \tau u)}{E(p|H - u > 0)} * \frac{B}{f(\bar{t})}
\end{aligned}$$

Thus

$$\frac{\hat{B}}{g(\bar{t})} \geq E(\bar{t} - t(H,u)|H - u > \tau u) \geq \frac{E(p(\tau, H, u)|H - u > \tau u)}{E(p|H - u > 0)} * \frac{B}{f(\bar{t})}$$

1.A.4 Explanatory notes on welfare analysis

The welfare cost of having traders wait for the time notch is represented by the term

$$\int_0^{\bar{t}} \int_{W,B} (k_b + k_v)(\bar{t} - t) dG(H,u) d\tilde{L}(t) \tag{1.22}$$

This could be reduced to

$$\begin{aligned}
&\int_0^{\bar{t}} \int_{W,B} (k_b + k_v)(\bar{t} - t) dG(H,u) d\tilde{L}(t) \qquad = \\
&\geq \int_0^{\bar{t}} \int_{W} (k_b + k_v)(\bar{t} - t) dG(H,u) d\tilde{L}(t) \\
&= \int_{H-u \geq \tau u} \int_{t(H,u)}^{\bar{t}} (k_b + k_v)(\bar{t} - t) dG(H,u) d\tilde{L}(t) \\
&= (k_b + k_v) \int_{H-u \geq \tau u} \int_{t(H,u)}^{\bar{t}} (\bar{t} - t) d\tilde{L}(t) dG(H,u) \\
&\approx (k_b + k_v) \int_{H-u \geq \tau u} \int_{t(H,u)}^{\bar{t}} (\bar{t} - t) l(\bar{t}) dG(H,u) \\
&= (k_b + k_v) l(\bar{t}) \int_{H-u \geq \tau u} \int_{t(H,u)}^{\bar{t}} (\bar{t} - t) dt dG(H,u) \\
&= (k_b + k_v) l(\bar{t}) \int_{H-u \geq \tau u} \frac{(\bar{t} - t)^2}{2} dG(H,u) \\
&= (k_b + k_v) l(\bar{t}) E(\frac{(\bar{t} - t)^2}{2} | H - u \geq \tau u) * \int_{H-u > \tau u} dG(H,u) \\
&\geq (k_b + k_v) l(\bar{t}) \frac{E((\bar{t} - t) | H - u \geq \tau u))^2}{2} * \int_{H-u > \tau u} dG(H,u) \\
&= (k_b + k_v) \frac{E((\bar{t} - t) | H - u \geq \tau u))^2}{2} * \int_{H-u > \tau u} dG(H,u) l(\bar{t})
\end{aligned}$$

which is the expression in the main text, and conceptually excluding those traders with surplus less than τu but who chose to trade at $\bar{t}$.

1.B Data appendix

Hong Kong

- **Source** The data used in this paper originated from the Memorial Day Book provided directly by The Land Registry of Hong Kong, and standardized by the author.
- **Transaction Record** The data used in this paper contains the "Agreement of Sales and Purchases" and its variations in the Memorial Day Book.
- **Sample construction** The data used in the current version of the paper does not include transactions that has multiple unit of real estate properties transacted in one single contract that constitute very small amount of all the contracts. The original data contains transactions of *all* real estate properties that are chargeable for stamp duty, that include properties for all uses including for example car parks and lands; the residential sample used in the current version of the paper is determined by the full address with official list of residential estates and other publicly available information, supplemented by information from the Rating and Valuation Department.
- **Date of purchase** The date of purchase is defined as the day in which the contract is signed, instead of the day it arrives in the Land Registry, under the column of "Date of Instrument" as opposed to "Date of Delivery". If there is a provisional agreement appears in the Memorial Day Book for the same property under the exact same price of formal agreement of sales and purchases and there is no other formal agreement of sales and purchased between the provisional agreement and the formal agreement in which it has the exact same price, it is considered as under the same transaction. The day of purchases is thus corrected for the day in which the provisional agreement is signed.
- **Mortgage record** For mortgages approved by financial institution, the Land Registry contains an entry and it appears in the Memorial Day Book. The description of mortgage entry in the Memorial Day Book allows identification between first time mortgage of the property, and for those transaction in which there is a mortgage entry record appearing in the 3 months that follows, before any other formal transaction record appears, it is defined as associated with a mortgage record.

Singapore

- The transaction in which a property is transacted on the same day with exact same price is omitted from the data which is likely due to an entry mistakes in the original data from the Urban Redevelopment Board, a procedure also applied in Giglio et al. (2015)

Chapter 2

British colonial gender laws and gender differential human capital investment in India[1]

We study the long run impact of historical legal reforms on matrimonial law introduced by the British within British provinces in 1800s and early 1900s on female education and under age marriages in post-Independent India, exploiting quasi-random variations of districts that were former British Provinces within each post-independent Indian states. From three independent sources of large scale micro data, that includes administrative records from schools and representative household surveys, we find that in former British Provinces females have 5% lower chances of marrying under the current legal age of 18 years, and 1.6% higher chance of attending school between the ages of 10-16 years, than those in the Princely States, where legal reforms in family matters were scant before 1947. We further digitize data on marriage status of population between 0-15 years at district level, from historical Census of India 1901-1931, to estimate the impact of Child Marriage abolition Act (1931) which raised the minimum age of marriage for female to 14. We find that the announcement of the law in 1929 before its implementation increased the likelihood of girls getting married at 5-15 years old by 2.8 percentage point, while districts that were more aware of the law exhibit lower child marriage in the long run. This provides evidence that the colonial matrimonial laws reduce child marriages in the long run, and that the regional differences in child marriage practices in India has a strong historical root.

[1] This chapter is a joint work with S Roy.

2.1 Introduction

Gender inequality in education is part of traditional cultures in many developing countries, where women get married earlier, are less educated and have poorer health outcomes than men (World Bank, 2012). Besides being a serious concern in terms of equality of opportunity, it may also has adverse implication to economic development (Klasen, 2002a).[2] This paper investigates whether the regions of India that have historically had legal institutions that fostered women's rights have better gender outcomes in the modern day. The analysis provides the basis for direct policy interventions on gender-biased social norms and practices in society. This is the first paper, to our knowledge, that investigates historical legal reforms to understand their long-run impact on gender outcomes.

The paper contributes to the literature on the impact of colonial institutions on modern day outcomes. While the literature on colonial institutions mainly focuses on changes in modern institutions (Acemoglu et al. (2001); Acemoglu and Johnson (2005)), our paper is closer to within country analysis, as in Michalopoulos and Papaioannou (2013) and Dell et al. (2015), showing that modern economic outcomes can be explained by the persistence of informal institutions. We contribute to the literature by studying the long-run impact of colonial institutions on household decisions on education and marriage, holding modern institutions constant. First, we map gender inequality in terms of education and marriage in modern India to historical political institutions. Next, we closely examine the short run and the long run impact of legal reforms introduced in two different polities.

After England took over India in 1858, India was divided into two different administrative institutions: the Princely States and British Provinces. This division ceased to exist post-independence; the State Re-Organisation Act 1956 re-divided India on the basis of linguistic identity. This led to a quasi-random distribution of Princely States and British Provinces within each modern state that made up independent India. Herein, we compare, within each modern state, the gender differential human capital investment between the regions that were under direct British rule and those that were Princely States in pre-independent India. Most of the variations in formal institutions are at the state level in India after independence. This implies that a comparison between regions within a state would allow us to control for almost all differences in formal institutions. Furthermore, the re-division of India along the lines of linguistic ethnicity allows us to further compare the

[2]Klasen (2002) finds that gender inequality in education is correlated with slower economic growth, both directly by lowering average human capital, and indirectly through its impact on investment and population growth.

impact of colonial social reforms on groups that share very similar ethnic identities. We compare gender-related outcomes between regions that were once under the direct rule of the British, termed as British Provinces, and regions that were ruled by hereditary Indian rulers, known as native states or Princely States. We find that in former British Provinces, 5% fewer females marry under the current legal age of 18 years, and females have 1.6% higher chance of attending school between the ages of 10-16 years than those in the Princely States. This shows that regions that have different historical experiences behave differently, even after coming under the same common law.

Our hypothesis is that the legal reforms introduced by the British rulers forcibly changed the behaviour of the natives in the British provinces resulting in a positive long term effect on gender equality in India today. Before examining the long term effect of British laws, we first test the short run impact of the law using historical data to determine whether the introduction of legal reforms in British provinces in the past changed the behaviour of the natives in that region in the past. To examine this we use historical census data on marriage and literacy from 1911-1931 to estimate the impact of Child Marriage Restraint Act (1929) using the difference-in-differences strategy. The Child Marriage Restraint Act 1929, passed on 28 September 1929 in the British India Legislature of India, fixed the age of marriage for girls at 14 years and boys at 18 years. It is popularly known as the Sarda Act, after its sponsor Harbilas Sarda. It came into effect six months later on April 1, 1930 and it applied to all of British India. This created significant incentives for families to marry their children before April 1930. We use the Census data of 1911 to 1931 to capture the effect of the announcement of the law, with the Princely States as our control group. We find that announcement of the law increased the likelihood of girls married at age 5-10 by 2.8 percentage point more among the natives in British provinces, compared to the natives in the Princely States.

Next, we examine whether the regions that exhibited excess marriage in 1929-30 have different marriage and education outcomes today as compared to the regions that did not. One may argue that the British provinces where the natives married their girls off early in response to the announcement of the child marriage abolition bill being passed were different in an unobservable way and this also affected marriage outcomes in the long run. Therefore, we instrument for the awareness of the law by using distance from the birthplace of the reformers who advocated for the law. The relevance of the instrument comes from the assumption that a reformer would have family links to the place of birth and would be more likely to go and spread their propaganda against child marriage in their birth place. So, the people living in places (even urban places) that are far away from the places of birth of the reformer are less likely to be aware of the law and hence are less likely to respond to the legislation. We use the instrument to estimate the impact of long run effect of Sarda Act (1931) on marriage under legal age and schooling decision in 2002, controlling for regional variation in the practice of child marriage up to 1921.

The OLS estimates of the long run impact of the Sarda Act show that regions that were more aware of the law in 1929-1930 were less likely to marry their girls at young age for the cohorts born in 1958-1984. If a district had one percentage point larger proportion of girls married between 5-10 years old in 1921, in later cohorts female were 0.45% more likely to marry between 14-17; however, one percentage point increase in the proportion of married girls from 1921 to 1931 predict that female are 0.2% less likely to marry between 14-17 for subsequent cohorts. The IV estimates show the same pattern and are larger in magnitude. Our estimates provide evidence that the awareness of the law reduce the degree of child marriages in the long run, it also suggest that a significant part of child marriage in India has a strong historical root going beyond 1921.

Our findings highlight the importance of understanding social background when we think about how society responds to the development of the labour market. Social norms in society can persist for many years and can affect the decision to participate in education and the labour market for certain demographic groups. Even with the same formal institutions and economic environment, a society riddled with prejudices may not take full advantage of its economic transformation and development. Our paper also explains a significant part of the large regional variations in the degree of gender bias in India. The regions that were formerly British Provinces have better female education outcomes and fewer females marrying under the legal age compared to former Princely State regions. This allows for both academic and policy discussions about gender in India to go beyond geographical differences by states or by social class. We provide some explanations regarding why such differences continue to persist, but are unable to clearly determine the impact of each historical law. This would require a more elaborate analysis of historical data, which we hope to accomplish in future work.

The paper is divided as follows. Historical background is provided in Section 2, followed by a conceptual framework in Section 3. The data and empirical strategy are described in Sections 4 and 5 respectively. The Results and Discussions are provided in Sections 6 and 7, followed by a conclusion.

2.2 Historical Overview

The British first arrived in India through a trading company called the East India Company. They signed their first commercial treaty in the year 1612, granted by the Mughal Emperor Jahangir. It was not until 1757 that the British had their first military conquest. The East India Company had experimented with a number of political arrangements to maximise their commercial profits and minimise their administrative liabilities. Some states were brought directly under their control and some states entered into political and commercial treaties with the British. This experiment came to an end with the Great Revolution of 1857, when the British Government took control. The British divided areas under British rule into two territories: British India and Native (or Princely) States. British India represented all territories under the Majesty's dominion that were ruled by the Queen through the Governor-Generals. The Native States represented independent kingdoms of all the Indian kings who accepted British suzerainty. They came under the governance of the Viceroy or the Governor-General, who was the head of the administration in India and a representative of the Monarch in India. A clear distinction between "dominion" and "suzerainty" was supplied by the jurisdiction of the courts of law: the laws of British India rested upon the laws passed by the British Parliament and the legislative powers of those laws vested in the various governments of British India, both central and local; in contrast, the courts of the Princely States existed under the authority of the respective rulers of those states (The Interpretation Act 1889, British Parliament). Although the East India Company enforced indirect control over the Princely States, the rulers of those regions were not passive figures. The indigenous rulers had their own customs and laws which they insisted on pursuing. (Ramusack, 2003).

India became independent in 1947, at which time it was still administratively divided into regions of British India, regions ruled by other European colonisers like the French or Danish and the Princely States. This division rendered it difficult for the administrators to rule the country. The State Re-Organisation Act was passed in 1956 that re-divided India on the basis of linguistic ethnicity. This is discussed further in the identification section.

2.2.1 Social reforms

Before the British came to administer the Indian territories, matters of marriage, maintenance, succession and legitimacy were solved using different religious laws for Hindus (such as Dayabhaga and Mitakshara law), literary traditions of Ithna Ashari and Hanafi for Muslims, and several customary laws for tribal communities. When the British took control of India, they promised not to interfere with personal laws such as marriage, succession etc. (Carroll, 1983). However, they reserved the right to intervene using statutory laws, which would override all religious laws in personal matters. Social reforms that were introduced by the British depended upon the discretion of the Governor-Generals in charge and the native social reformers (see Chitnis and Wright (2007)). All the British reforms that were introduced by the Governor-Generals were in direct conflict with the existing laws of Indian society (Lord William Bentinck (1829), Carroll (1983)). Most of the social reforms were not in the interest of the British, as they created tension between the natives and their British rulers. However, the laws were passed after much deliberation by the reformist Governor-Generals. The first of the most important social reforms introduced in colonial India was the abolition of Sati in 1829. Sati was only practiced by upper caste Hindus in Bengal, Rajputana and Central India. It was a practice that involved a widow immolating herself on her husband's funeral pyre. The reform was pushed forward by a native social reformer, Raja Ram Mohan Roy. Lord William Bentinck introduced this law, arguing that the general masses of India were uncivilised and would continue this custom if the British did not bring forward a legal reform making it a punishable offence. In a speech in 1829, he pointed out that Britain could afford to abolish Sati without fearing rebellion from the natives because the majority of Indian soldiers in the British army belonged to the tribes that did not practice Sati (Fisch, 2000). Since Sati was only practiced by few ethnic groups in India, it was possible to extend the law outside British jurisdictions. The British negotiated with the Princely States to abolish Sati - Rajputana was the last native state to abolish it in 1861 (Ramusack, 2003).

Since then, most of the social reforms were implemented within British Provinces but were not enforced in the Princely States. With the initiative of the educationalist Pandit Iswar Chandra Vidyasagar, the British passed the Hindu Widow Remarriage Act of 1856. Until then, widow remarriage among upper caste Hindus had been prohibited, and Hindu widows were expected to live a life of austerity (Peers, 2013). It was introduced with the rationale of reducing female infanticide (Law Commission, 1837) and was very unpopular among the natives. The law, however, deprived childless widows of inheritance (Law Commission Report, 1856).

Although Sati was abolished in all of India, as a practice, it was not as widespread as female infanticide and child marriage (Grey, 2013), which existed across all of India and in all religions. Unlike Sati, the practice of female infanticide was not restricted to upper caste Hindus. The abolition of female infanticide (1870) and child marriage were harder to implement as they went directly against the widespread age-old customs of the natives across castes and tribes (Grey, 2011). The laws related to these practices were again confined to the British Provinces. In 1891, the Age of Consent Law was passed that raised the age of consent to 12 years. This bill created a lot of tension among the native population (Chitnis and Wright (2007); Ramusack (2003)). The reforms were slow. It took the British almost forty years to pass the Child Marriage Abolition Act (also called the Sarda Act) in 1929, which raised the age of consent to 14 years.

In our paper, we will closely examine the impact of the Sarda Act on both historical and modern marriage outcomes. The Child Marriage Restraint Act 1929, passed on 28 September 1929 in the British India Legislature of India, fixed the age of marriage for girls at 14 years and boys at 18 years. It is popularly known as the Sarda Act, after its sponsor Harbilas Sarda.[3] It came into effect six months later on April 1, 1930 and it applied to all of British India.[4] With protests from the Muslim organisation in undivided India, a personal law called as Shariat Act was passed in 1937 that allowed child marriages among Muslims with the consent of the childs guardian. Family matters were in general governed by personal religious laws such as the Shastric law for Hindus, and Shariat law for Muslims etc. The British social reforms mostly interfered with Hindu Shastric law, using statutory laws to override customary religious laws (Carroll, 1983). Hence our analysis focuses on the Hindu population of both the British provinces and the Princely States.

First we compare the impact of the Sarda Act on marriages in the British Provinces and the Princely states. However, due to paucity of census data in 1941 and 1951, we can only analyse the immediate impact up to 1931, which capture mainly the announcement effect of the law. To further analyse the causal impact of the awareness of the law on marriage outcomes in the long run, we use an IV strategy using distance from the birth places of the reformers who pushed for the bill to be passed in the legislative assembly.

The Sarda Act was the first social reform that was brought about by the efforts of organised women committees (Raman, 2009; Mukherjee, 2006). Pro-reform politicians, such as Motilal Nehru, were caught off guard when the organised women's association met with leaders to ask for their support in the bill. The members of All-India Women's conference (AIWC), Women's Indian Association and National

[3]Before the Sarda Act (1931), a cult group called Brahmo Samaj pioneered by Raja Ram Mohan Roy abolished the marriage of girls below 14 years of age in 1872 under an act called as the Native Marriage Act. But it only applied to the members of that group

[4]Hatekar et al. (2007) found that after the Sarda Act the probability of girls marrying below the age of 14 years dropped dramatically among the upper caste using micro data from family genealogies.

Council of Women in India pressured politicians for their support to the bill, standing outside their delegations holding placards and shouting slogans such as 'if you oppose Sarda's bill, the world will laugh at you'. It was also this group who pushed for, and eventually succeeded in having Gandhi address the evils of child marriage in his speeches. Apart from the members of organised associations, there were civil servants, academics and members of other reformist group such as the Brahmo Samaj that propagated the ideas of child marriage as a social evil. The reformers came from both the British provinces and the princely states (Mukherjee, 2006) . However, the reformers could pass the child marriage reforms only in the British provinces (Sinha, 2006; Mahmood, 2002).

In contrast to the reforms in the British Raj, there were very few gender reforms in the Princely States. The only Princely States that implemented gender-related reforms were the Mysore and Kathiawar Agency of Baroda. Dewan Sheadari Iyer of Mysore in 1894 abolished the marriage of girls below the age of 8, and marriage between girls under 16 years old to men over 50. This law was less stringent than the British Sarda Act. In the face of widespread discontent among the masses, the Mysore Princely State mostly implemented this reform by occasionally prosecuting the powerless lower castes (Ramusack, 2003). The political agent, Alexander Walker, of Kathiawar agency tried to abolish female infanticide among the Jhareja and Jetwa tribes, with little success. (Walker, 1856)

2.3 Conceptual framework

The social reforms implemented under the British rule in India may explain the differences in educational outcomes between the former Princely State regions and those that were under direct British rule, through more than one channel. We first discuss how specific reforms in British India directly affected the decision-making of the household, before discussing three potential mechanisms that could generate the long-term persistence of gender inequality, years after Indian independence from the British rule: the persistence of social norms, an information friction channel and the impact on the re-allocation of household resources.

Early marriages bring monetary savings and reduction of effort cost to the families of daughters, as after marriage they will no longer need to be taken care of at home. If parents are happy to see their daughters married at an early age, they may only choose to educate their daughters when the net return on education is very high. Raising the legal age of marriage increases the total amount of time the daughters stay at home, and thus the cost of raising them. If there exists economic opportunities for skilled labour, the households will have an incentive to educate their daughters to participate in the labour market to reduce the net cost of looking after them for a prolonged period.

With the passing of the State Reorganisation Act, the Princely States and the British Provinces came under the same jurisdiction and laws. Observed differences in gender bias in education after this reunification could either be explained by differences in the perceived return of education *specific for female* or the historical persistence of cultural bias/dis-utility generated by female participation in activities outside the household.

Our empirical exercise attempts to highlight that social reforms have a slow but persistent effect. One possible explanation for the persistent gap may be due to information friction. Correctly inferring the returns of education could be costly, and households may only make inferences based on limited experiences of other members of the same village. A larger initial stock of human capital among females could help the community identify market opportunities that are suitable for females. Thus, differences in initial human capital stock generated by historical reform in British India could translate into differences in the perceived returns of education, particularly when the return of human capital rose rapidly after trade liberalisation in India.

An alternative explanation could be linked to current debates on the subject of women empowerment (Duflo (2012) provided an in-depth discussion). Higher female education may have a direct impact on resource allocation and decision-making within the household (e.g. Quisumbing (1994) found that better educated mothers invested more in girls; Breierova and Duflo (2004) found evidence of female education on reductions in fertility and child mortality). If the mechanism holds, an exogenous shock that increases female education would have intergenerational persistence simply because better educated mothers allocate more resources to educating their daughters. This intergenerational transmission mechanism will have a larger effect if the other two mechanisms are also in operation.

With the mechanisms discussed, we test the hypothesis that regions that were historically under British rule have better gender outcomes in the short run and in the long run, compared to regions that were Princely States. In the following sections, we discuss the data and empirical strategy used.

2.4 Data

Our main source of information on the administrative division between the Princely States and the British Raj is Baden-Powell (1892), which included a detailed map on the division between the Princely States and areas under direct British rule together with the year of acquisition for each district. As the landscape of the Princely States and direct British rule was mostly settled by 1857, we define a district to be under British direct rule according to Baden-Powell (1892), otherwise it is defined as a Princely State. The geographical distribution is presented in Figure 2.2.

Our measure of human capital investment comes from two independent sources: District Information System for Education and the National Sample Survey. The District Information System for Education (DISE) provides administrative records for enrolment at the school level in India. The data is designed to cover all regions of India in terms of the administrative information for each school in each academic year, including the number of students of each gender enrolled and the number of classrooms in each school. As the distinction between the Princely States and regions of direct British rule is mostly at the district level, we aggregate the information at the district level.[5] For the analysis, we aggregate all schools in each district in terms of the number of students enrolled in each class by gender for each year between 2005-2013; this gives us estimates of the ratio of male to female students enrolled in each class in each academic year for 433 districts. Summary statistics are reported in Table 2.1. On average, the schools in India have 9 % more boys enrolled in Class 6 compared to girls. [6]

The National Sample Survey (NSS) 64-66th round (2007-2008; 2009-2010) is another important data source that allows us to measure school attendance at the individual level. We focus on school attendance for children aged between 10-16 years old to study human capital investment decisions beyond basic literacy. It gives us approximately 155,989 individual records (of 10-16 years old) regarding their principal activities in the past 365 days, including school attendance, participation in domestic work, and casual waged work. We report the summary statistics in Table 2.2 for the sample we used. The average school attendance rate is 0.85, with, however, very high variance.

We further look into the percentage of marriages under the legal age in the year

[5] We excluded Karnataka in the analysis in this sample due to the lack of data availability at the time of writing.

[6] This is the ratio of raw enrolment, i.e. it does not take into account the gender ratio of the population; taking the NSS estimates of the proportion of females from 10-16 reported in 2.2 as 0.464 (which is by itself a number that shows very high gender bias), there are 15 % more boys than girls in this age range.

2006-2007 at the district level from the District Level Household and Facility Survey (DLHS Round 3) by the Ministry of Health and Family Welfare of India. The data reported marriages under the legal age of 18 for women and recorded all marriage ceremonies held during the three years preceding the survey, covering 570 districts. [7] We use the micro data of the survey for the year 2002-2004 (DLHS Round 2) to conduct analysis with respect to Hindu female who are beyond the age of 18 at the time of interview, there are 86,214 individuals that we could merge where they are now with the historical census data we have.

In addition, we obtained the district level GDP per capita from the Planning Commission of the Government of India. The geographical controls, such as latitude and distance to the coast[8] for each district, are defined at the centroid of the districts.

To study the persistence of the marriage pattern and the impact of the Sarda Act in 1929, we digitized the data from the Census of India regarding the population and marriage status of male and females at the district level for 1901-1941, covering the major British provinces and Princely States. [9] The census data are available at ten years interval for 1901, 1911, 1921, 1931 and 1941. We mainly limit our analysis to the data from 1911-1931, as the data were available for all major provinces and have consistent definition of variables across years, while in 1941 data do not exist for some regions and the definitions are inconsistent with those reported in previous years. There are changes in district names since the independence. To map the historical data to modern data, we geocode the historical district name, and compute which modern district it falls into. If more than one historical district falls into the same district in administrative division post-independence, we associate the average of records from the historical districts to the modern district. This maps into 126 modern districts.

We document the reformers for the Sarda Act. The reformers were coded by recording the names of the members from All India's Women Conferences (AIWC), Womens International Conference and National Council of Women in India who were directly associated with the Sarda Act. The information came from various journal articles on the Child Marriage Restraint Act, supplemented by information from Wikipedia. Apart from the women members of organized women groups, other reformists groups such as the Brahmo Samaj, members of civil service, journalists and politicians were involved in supporting the Act and attempted to create a strong public opinion against child marriage. These names were similarly extracted from

[7] The data is released through DevInfo 6.0 by UNICEF.

[8] Physical distance instead of travel distance.

[9] This includes Madras, Bombay, Bengal, Rajputana, Central Provinces, Central India Agencies, Mysore, Travancore, Hyderabad, Ajmer and Punjab

articles on the Child Marriage Restraint Act. For each reformer, we coded the birth place, birth year, places where they were active and the reforms they were involved in.[10]

2.5 Empirical strategy

To study the long run impact of social reforms on human capital investment, a common challenge is to control for modern institutions and ethnicity. Different ethnic groups may be starting with different social norms. Moreover, each ethnic group may have different laws and social institutions that endogenously emerge according to the customs and culture of the group. We will describe how the State Re-Organisation Act of 1956 could help us control for both ethnicity and modern institutions.

After independence in 1947, the State of India was divided into three main regions: regions that were formerly British Provinces, regions that were under the rule of hereditary Indian rulers, and regions that were formerly under other European rulers. This division proved difficult for administrative purposes. Thus, the government of India decided to divide India on the basis of linguistic ethnicity. This proposal was very popular among the masses. The Telegu-speaking people formed the state of Andhra Pradesh, Marathi-speaking people formed the state of Maharashtra, and Kannada-speaking people formed the state of Karnataka, etc.

Linguistic ethnicity is an important determinant of identity in India. Modern India is adversely affected by conflict and riots triggered on the basis of differences in language. Since the 1920s, there has been conflict between Assamese and Bengali-speaking people. In recent times, Bihari-speaking people have been targeted in Assam. In Maharashtra, Marathi-speaking people target migrants from Bihar and South India. In recent times, there has been a movement towards the compulsory use of the Marathi language in Mumbai, including in the Municipal Corporation. (see Baruah (2003); Weiner (2015); Murthy (2006); Menon (1989); Mitra (1995))

Therefore, each modern state in India has people speaking the same language but with different historical experiences in terms of direct and indirect British rule. Residents of each state were subjected to same state law after 1956; this is our key source of identification. We argue that the distribution of Princely States and British Provinces are quasi-random within each state. We assume that the British did not select groups of people with a particular type of gender preferences within ethnicities to subject them to direct British rule. People with the same ethnicity

[10] There are 36 reformers in our sample.

tend to share norms. It is hard to imagine that people that were subjected to direct British rule had systematically different gender preferences to those of the same ethnicity living in native states at the beginning of the British India era.

In this section, we investigate the effect of Princely States (as opposed to being directly ruled by the British) on modern gender differential human capital investment. The key differences between the two forms of control were the gender-related social reforms that were highlighted in the historical section of this paper. However, there are potential confounding factors, such as differences in income and geographical characteristics, which we try to control for.

We use the following specification to test the impact of the rule of Princely States on the male/female enrolment ratio in the DISE data.

$$MFR_{sdct} = \alpha * I[princelystates]_{sd} + X'_{sd}\xi + \delta_s + \gamma_t + \mu_{sdt} \tag{2.1}$$

MFR_{sdct} measures the ratio of male/female students enrolled in class c in district d within state s in year t. α, the coefficient of interest, captures whether in Princely States there are systematically more male children enrolled in school. δ_s is the state fixed effect that captures the systematic differences between states, such as the gender ratio, unobserved gender bias in social norms, and the provision of schools. X'_{sd} is the district level controls that include the proportion of rural schools in district d, the average number of classrooms in schools in district d, log GDP per capita (in 2000), and latitude and distance to the coast. γ_t is the year fixed effect that controls for yearly variations in gender differences in school enrollment. [11]

Moreover, we use the following specification to test the impact of the rule of Princely States on school attendance and participation in waged work of women aged 10-16 years in 2006-2010 using the NSS data.

$$y_{sdi} = \beta * I[princelystates]_{sd} * female_i + \gamma_s + \phi_s * female_{sdi} + X'_{sdi}\eta + D'_{sd}\sigma + \epsilon_{sdi} \tag{2.2}$$

Where y_{sdi} is an indicator of school attendance/participation in waged work [12] as the principal activity in the 365 days before individual i in state s of district d was interviewed. $I[princelystates]_{sd}$ is a district level indicator of whether district d in state s belonged to a Princely State. The coefficient of interest is β, which is the coefficient for interaction term between $I[princelystates]_{sd}$ and $female_i$, which is a female dummy variable for person i. This captures whether females do worse in Princely States compared to direct British-ruled regions. γ_s is the state fixed effect

[11] All standard errors are clustered at district level

[12] This includes casual wage labor and not regular salaried work, and should be more relevant for the age range in our sample

for school attendance that captures state level differences in school attendance, such as different levels of provisions of schools. ϕ_s is a state-specific female fixed effect. This state-specific female fixed effect would mostly capture the different degrees of gender bias that exist in different states, which could be attributed to differences in gender norms between different ethnicities or differences in the labour market return of females. X_{sdi} is a set of socioeconomic controls that include the age of the child, the age of the head of the household, and the square of the age of the household head, an indicator for Muslim, Christian and other religions, an indicator of rural areas, and an indicator of the landownership of the households. D_{sd} is the geographic controls for district d in state s, which includes latitude and distance to the coast.

Moreover, we use the district level aggregate of the District Level Household and Facility survey to test the impact of Princely States rule on the number of girls that marry under the legal age. We estimate the following equation:

$$M_{sd} = \sigma * I[princelystates]_{sd} + X'_{sd}\Phi + \kappa_s + \tau_{sd} \tag{2.3}$$

M_{sd} is a continuous measure of the percentage of marriages under the legal age in 2006-2007 for district d in state s, σ is the coefficient of interest as it tells us whether in former Princely State regions, more marriage are carried out under the legal age. X_{sd} is the district level controls that include latitude, distance to the coast and log GDP per capita (in 2000). κ_s is the state-fixed effect which controls for systematic differences across the states.

2.5.1 Response to Sarda Act

In this paper we argue that British legal reforms affected the behaviour of the natives in British provinces by abolishing their traditional customs. To show the impact of the British legal reforms on the behavior of the natives, we begin with the study of the effects of the Sarda Act, the child marriage abolition law in 1929-1930. Figure 2.3 plot the percentage of male/female married in the age group of 5-10 and 10-15 from 1901-1931 for the whole of India. The marriage pattern were stable from 1901-1921, while in 1931 the proportion of females married increase dramatically for all young age groups, particularly among the female. This is most likely due to the anticipation effect in the six months between its announcement and implementation. (Census of India 1931) Figure 2.4 shows the geographical distribution for the proportion of female married at age 5-10 in 1921, as well as the change from 1921-1931. It is not clear that places that experienced the highest increase in child marriage in

1931 were those that traditionally practiced child marriage in most numbers . Using historical census data, we estimate the following equation to test whether historical institutions explains the change in marriage pattern from 1921 to 1931.

$$M_{pdt} = \alpha * I[Britishdirectrule]_{pd} * I[t = 1931] + \gamma_p * t + \phi_t + \sigma_d + \tau_{pdt} \quad (2.4)$$

M_{pdt} is the percentage of female who already got married at the age 5-10 in district d of pre-independent province p (i.e. political division before State-Reorganization Act) in year t between 1911-1931.[13] $I[Britishdirectrule]$ is an indicator which equals 1 if the district were under British direct rule.[14] α captures the differential changes in marriage pattern from 1921 to 1931 between former British direct rule regions and Princely States. Assuming there is no other factors that affect marriage pattern of the two regions differently between 1921-1931, α identify the effect of anticipation of actual implementation of the law. We control for the province specific trend ($\gamma_p * t$), district fixed effect (σ_d) and year fixed effect (ϕ_t).

2.5.2 Long run impact of Sarda Act

In this section we test the hypothesis that the awareness of the Sarda Act has long run impact on female marriage and education outcome, as measured in the DLHS 2002 for Hindu female aged above 18.

The equation of interest would be

$$y_{sdi} = \beta L_{sd} + \beta_2 M_{sd,1921} + X'_{sdi}\sigma + \epsilon_{sdi} \quad (2.5)$$

y_{sdi} are outcome variables measured in DLHS in 2002, for individual female i in state s, district d. It includes outcomes for marriage and education: indicator of marrying below the age of 14, marrying in the age range of 14-17 and marrying under the age of 18, indicator for any level of schooling and a continuous measure of years of schooling. L_{sd} is a measure of awareness of the Sarda Act for district d in state s since 1929. If the awareness of the Sarda Act reduces the probability of early marriage for female and increases the educational outcomes for females in the long run, we expect β to be positive. β_2 captures the accumulative effect of traditions and historical reforms before 1921 that could explains outcome in 2002.

[13] It is defined as the number of married female at age 5-10 divided by the total number of female at age 5-10, reported by the Census

[14] so $I[Britishdirectrule] = 1 - I[princelystates]$, the variables we used in the specification described earlier

Without a direct measure of L_{sd}, we use $M_{sd,1931}$ as proxy, and estimate the following equation

$$y_{sdi} = \eta M_{sd,1931} + \gamma M_{sd,1921} + X'_{sdi}\sigma + \mu_{sdi} \tag{2.6}$$

$M_{sd,1921}$ and $M_{sd,1931}$ are the percentage of female who already got married at the age 5-10 in district d of state s in year 1921 and 1931 respectively, constructed by mapping new districts with their historical counterparts.

Given the historical context, $M_{sd,1931}$ should be positively related to L_{sd}, in the sense that in the districts with higher awareness of the law, more females in age range of 5-10 would be married as measured in 1931. This implies that η in equation 2.6 would have the same sign of the effect of β in equation 2.5.[15] One can argue that there are unobserved factors that determine how much a district would react to the legislation in 1929-1930, and influence the age of marriage and education in 2002 at the same time. For example a district could be conservative to begin with,where many girls may have married in a rush in 1929-1930, while the same culture could make girls more likely to marry young in 2002. Therefore to estimate the equation 2.6 consistently, we need instruments that determines the awareness of the legislation of a district and are unrelated to unobserved factors that affect the female marriage and educational outcomes after 1921.

The abolition of the child marriage in the 1920s was mostly influenced by women and husbands of women in the All India Women's Conference (AIWC). The members of AIWC were the first ones to support the legalization of the abolition of the traditional norm of child marriage. The members of the Women's Indian Association and National Council of Women in India also participated. But the leadership of these two organizations was hesitant about state involvement in matters of marriage and traditional customs. The AIWC fiercely promoted the legislation and articulated arguments in favour of the legislation to be put forward in front of the Joshi Committee, which was evaluating the bill proposed in the Central Legislative Committee. To construct an instrument for the awareness of the act, we make use of the birthplace of the reformers who advocated this law. The underlying assumption being that the place the reformers were born is unrelated to the practice of child marriage. Child marriage was prevalent throughout India across all castes and religion. Individuals that have gone on to become feminist reformers were mostly influenced by the peers in the cities or in the colleges where they might get educated,

[15] We can assume formally that $M_{sd,1931} = \alpha L_{sd} + M_{sd,1921} + v_{sd}$, where $\alpha > 0$. Then one can rewrite $y_{sdi} = \eta\alpha L_{sd} + (\eta + \gamma)M_{sd,1921} + X'\sigma + \eta v_{sd} + \mu_{sdi}$. Therefore $\eta\alpha = \beta$, and $\eta + \gamma = \beta_2$, with the assumption that $\alpha > 0$, η and β follow the same sign. Thus, η is well-identified but α and β is only identified up to scale.

which were mostly away from their birth places. The relevance of the instrument comes from the assumption that a reformer would have family links to the place of birth and would be more likely to go and spread their propaganda against child marriage in their birth place first. So, the places (even urban places) that are far away from the places of birth of the reformer are less likely to be aware of the law and hence are less likely to respond to the legislation.

Equation 2.6 is then estimated using D_{sd}, the distance of district d to the nearest birth place of any reformers, and $I[Princelystates]_{sd}$, an indicator of Princely States as instrumental variables.

2.6 Results

2.6.1 DISE

The OLS results for equation 2.1 of school enrolment are presented in Table 2.3. Column 1 reports the estimates for equation 2.1 for the ratio of gross enrolment of boys to girls in Class 6; the coefficient suggests that on average there are 2% more boys enrolled in schools than girls in former Princely States versus British-ruled regions. The availability of larger schools measured by the number of classrooms predicts lower boy to girl enrolment ratio. We also included log GDP per capita to control for the provision of schools and household budget constraints across districts within the same state, however, it is only marginally significant. Column 2 reports the same measure for Class 5, where the coefficient is very small and insignificant; this suggests that the results in Column 1 are mainly due to the dropping out of girls in higher grades rather than being driven by the differences in the gender ratio. Girls reach the age of 12 when they reach Class 8. Older girls may be more helpful for household domestic work for which parents might take them out of school. If a community thinks that return of education of girls are low, then it is less likely for the community to invest in secondary schooling of girls.

In Table 2.4, we report equation 2.1 estimated by each class from 1 to 8. Comparing across columns, it is clear that the gender enrolment ratio only starts differing at Classes 6, 7 and 8, at which time the decision to attend school is more closely related to a human capital investment decision beyond basic literacy. The magnitudes of the coefficients across Columns 6, 7, and 8 are quite consistent at around 2-3 %, suggesting that Class 6 is a critical time when, if girls drop out of school, they may not return, whereas those that stay in education are likely to proceed with similar probability to boys.

2.6.2 NSS

Table 2.5 reports the estimates of equation 2.2 on the main activities of children aged 10-16 years from the NSS data. Columns (1)-(2) report the estimates for school attendance, Columns (3)-(4) report the estimates for participation in waged work, and Columns (5)-(6) report participation in domestic work. The estimated interaction term Princely States*female is significant for school attendance, which shows that girls in Princely States are 1.6 % less likely to attend school compared to girls in British-ruled regions within the same modern state. We do not see similar significances in other outcome variables once we include the state female fixed effect to control for gender bias at the state level. The main effect of the Princely State for paid work participation (Columns (3) and (4)) is only significant when we exclude the interaction term with females, and the magnitude is small (0.5 % difference in the probability of market participation between Princely States and British-ruled regions). This could potentially be explained either by the lower age of marriage in Princely States or a small difference in the availability of market work.

However, our estimates on school attendance cannot be solely driven by the availability of market work. We further report estimates of equation 2.2 by Hindus and Muslims, as the historical overview section has shown that there were stark differences in how Hindus and Muslims responded to the social reforms in British-ruled regions. Column (1) in Table 6 reports the estimate for Hindus only. The coefficient is highly significant with a magnitude higher than that in the sample including all religions - females among the Hindu population are 2.1 % less likely to attend school in former Princely States, greater than the equivalent estimate of 1.6 % for the whole population. Moreover, the main coefficient of the Princely States is *positive* and is marginally significant for Hindus, which means that males are *more* likely to attend school in Princely States - this supports the hypothesis that the fundamental cause of the observed difference is driven by the persistence of cultural practices rather than the availability of education. On the other hand, the same estimate for Muslims in Column (2), despite its smaller sample size, is not only statistically insignificant but the sign of Princely States*female turns positive with a very large standard error. Instead of explaining the difference by time invariant inherent cultural differences between Hindus and Muslims, we tend to associate this difference in our estimates by how cultures interact with the implementation of the law in British-ruled regions before Independence.

In Figure 2.A.2 we plot the percentage of married female at age 10-15 for districts that now belong to Madhya Pradesh - there were historically huge differences in how Hindus and Muslims responded to the Sarda Act of 1929 and Age of Consent Law of 1891.

2.6.3 Marriage under legal age

Table 2.7 reports estimates of equation 2.3 using the district level aggregate of the percentage of marriages under the legal age for females in 2006-2007. The coefficients estimated are positive, highly significant and robust upon inclusion of log GDP per capita (Column (1) and (2)). Our estimates suggest that Princely States have approximately 5 percentage points more marriages under the legal age for females. In Column (3) and (4), we report the results using the mean age of marriage in 2002-2004 as an outcome variable; it can be seen that districts formerly belonging to Princely States have a lower mean age of marriage by 0.4 years. An average of 22.66 % of all marriages in India take place under the legal age for female [16]; our estimated 5 percentage points difference between Princely States and direct British-ruled regions explains a significant number of underage marriages in India.

2.6.4 Sarda Act using Census Data 1911-1931

Above we provide a mapping of gender inequality to different political institutions. In this section we focus on the impact of the legal reforms under two different political rule that affected behaviour in the past.

The estimates for equation 2.4 are presented in Table 2.8. The coefficient of *Princely states*1931* is statistically significant in column (2) where we control for province specific trend. The coefficient estimate is 2.8, which shows that among girls aged between 5-10, there are on average 2.8 percentage more girls among natives in British provinces who got married in 1931. The natives in British provinces feared the implementation of the act in the coming months and millions of girls under the age of 14 were married off. This result is also well documented in the census reports of the British. The Sarda Act applied to only British India, however in Princely States such as Mysore and Baroda also tried to enact laws abolishing child marriages. [17] We observe a slight bunching in female child marriages in the Princely States, but in British provinces it is on average more severe.

Compare to Column (1), which we did not include province specific trend, the coefficient in column (2) is more significant and with a larger magnitude, this suggesting that provinces may have differential trends before 1931.

[16]From our district level aggregate not weighted by population share in each districts

[17]Mysore in 1894 abolished child marriage below the age of 8. Many reformers from Mysore who pushed for the legislation of the Sarda Act could not raise the age of marriage for girls in Mysore. Therefore, in Mysore there was a weak form of child marriage restraint reform.

Table 2.9 reports the estimates for equation 2.6, both in OLS and IV using distance to nearest reformers and Princely States dummy. Column (1) and (2) in Panel (a) report the estimates for the outcome of marrying below 14, where the coefficients for both $M_{sd,1931}$ and $M_{sd,1921}$ are insignificant, suggesting that the historical marriage pattern does not explain the probability of marrying below 14 in 2002. Column (3) and (4) report the estimates for the outcome of marrying between 14-17, and the coefficient for $M_{sd,1931}$ and $M_{sd,1921}$ are both statistically significant but of *opposite sign*. The estimate for $M_{sd,1921}$ in column (3) for the OLS is 0.00445, which could be interpreted as one percentage point increase in the proportion of females married at the age of 5-10 in 1921 predicts a 0.45 higher probability of a girl marrying below the legal age in 2002. The magnitude of the coefficient suggest that a significant part of child marriage in India has a very strong historical roots, going far beyond 1921.

Moreover, the estimates for $M_{sd,1931}$ is negative and also statistically significant in column (3). It shows that one percentage point increase in the proportion of female married in 1931 (between 5-10) predict a smaller probability of a girl getting married below legal age post-independence. Column (4) report the IV estimates, and the coefficients for $M_{sd,1931}$ and $M_{sd,1921}$ both become larger in magnitude and highly statistically significant, and remains in opposite sign. With the assumption that an increase in child marriage in 1931 proxy for a high awareness of the Sarda Act, the estimate provide evidence that the Sarda Act reduced child marriages in the long run. In column (1) of Panel (b), we find that 1% more girls married at the age 5-10 in 1921 predicts a 0.6% lower chance of females getting at least some education measured in 2002; and 1% more girls married at the age 5-10 in 1931 predicting a 0.44% higher chance of females getting some education in 2002.

In table 2.10 we further control for state fixed effect. The coefficient for $M_{sd,1921}$ remains significant and positive in most OLS specification, while the coefficient for $M_{sd,1931}$ remains negative but become statistically insignificant except in the OLS regression for year of schooling in column (3) of panel (b). This may be due to low power of the instrument once we control for state fixed effect.

We find regions that experienced bunching of marriages in 1931 have fewer girls marrying below legal age and are more likely to have experienced schooling post-independence, controlling for cultural variation across regions up to 1921. One interpretation of this long run effect could be that in British provinces the natives anticipated the implementation of the law in 1931 and did not wish to get affected by it. Therefore, the generation most affected by the Sarda Act are the later generations, who are more likely to conform their behaviour to any new law. This may explain why regions that got affected by legal reforms imposed by a foreign administrative body behave differently than regions that are culturally similar but did not get affected by the reform.

2.7 Discussion and robustness check

2.7.1 Robustness check - Princely States that potentially undergone reform

As discussed in the previous section, it was documented that in two of the Princely States (Mysore and the Kathiawar Agency of Baroda), there were reforms related to child marriages independent of similar reforms in the British Provinces. In the previous section, where we estimated equation 2.2 and 2.3 (school attendance in NSS and marriage under the legal age), we did not exclude Mysore and the Kathiawar Agency of Baroda because their implementation is weak from the historical description. We present the estimates for equations 2.2 and 2.3, excluding these two Princely States as a robustness check, in Table 2.A.1 and 2.A.2.

In Table 2.A.1, we report the estimates for school attendance in the NSS data excluding Mysore and Kathiawar Agency of Baroda in Column (2). The coefficient of the Princely States indicator increases slightly from 1.6 to 1.8, which implies a bigger difference in female school attendance among 10-16 years old between the Princely States and the British Provinces. Similarly, in Table 2.A.2 we report the estimates of equation 2.3 for the percentage of marriages under the legal age and the mean age of marriage, excluding Mysore and Baroda. The coefficients again increased slightly upon the exclusion of the two districts (in Columns (2) and (4)), which is what one would expect if Mysore and Baroda had weak social reforms that were similar in nature to those in the British Provinces.

2.8 Conclusion

In this paper we show that two regions that have had different legal reforms in the past behave differently when placed under the same modern institution. In particular we find that girls are more likely to go to school in regions that have had gender reforms in the past. If two regions are given the same opportunities in terms of provision of schools, we argue that the region that has had gender related legal reforms will have more females exploiting the opportunities. Our findings support policy intervention that eliminates prejudice behaviour by showing its positive long term impact. Providing infrastructure by the social planner might not be enough for economic growth, we also need to change the bottlenecks on the demand side.

2.9 Figures and Tables

Figure 2.1: Timeline of key historical events

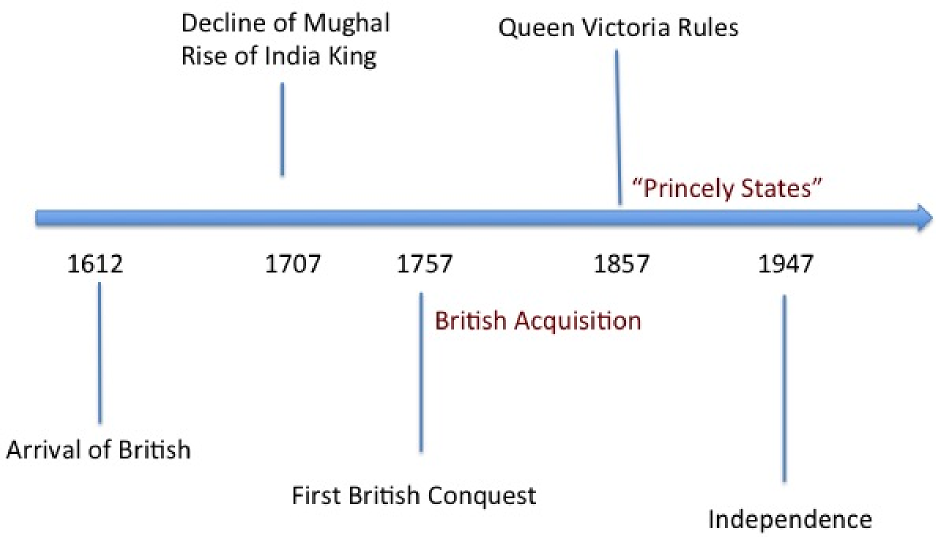
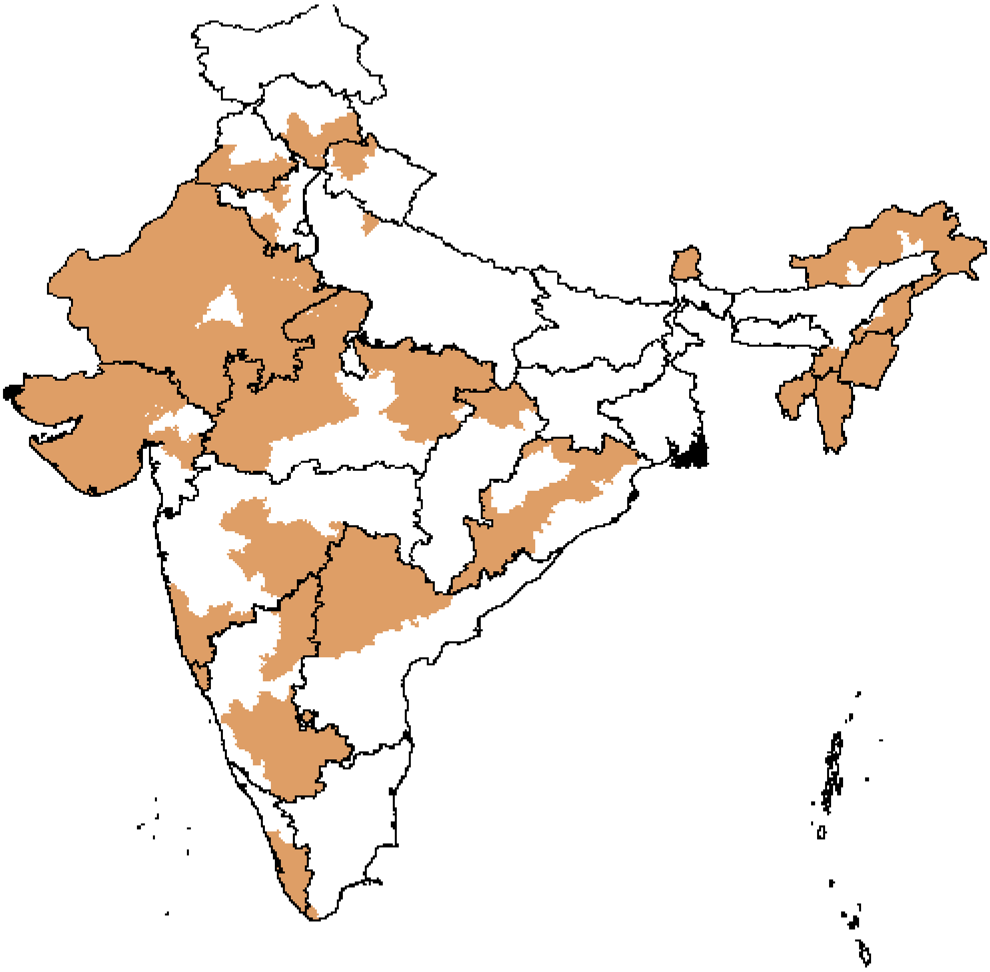
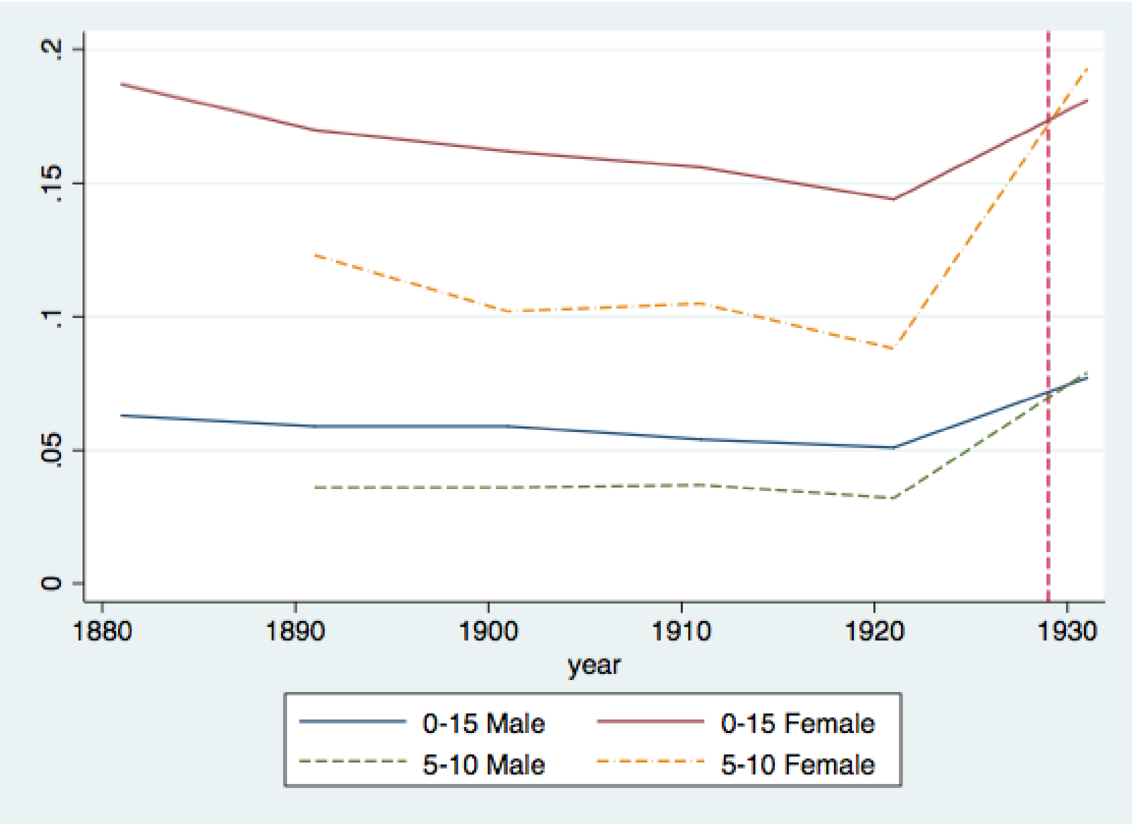


Figure 2.2: Distribution of Princely States and British direct rule regions



Note: The shaded parts were districts that belonged to Princely States and the white parts are districts that were under British direct rule.

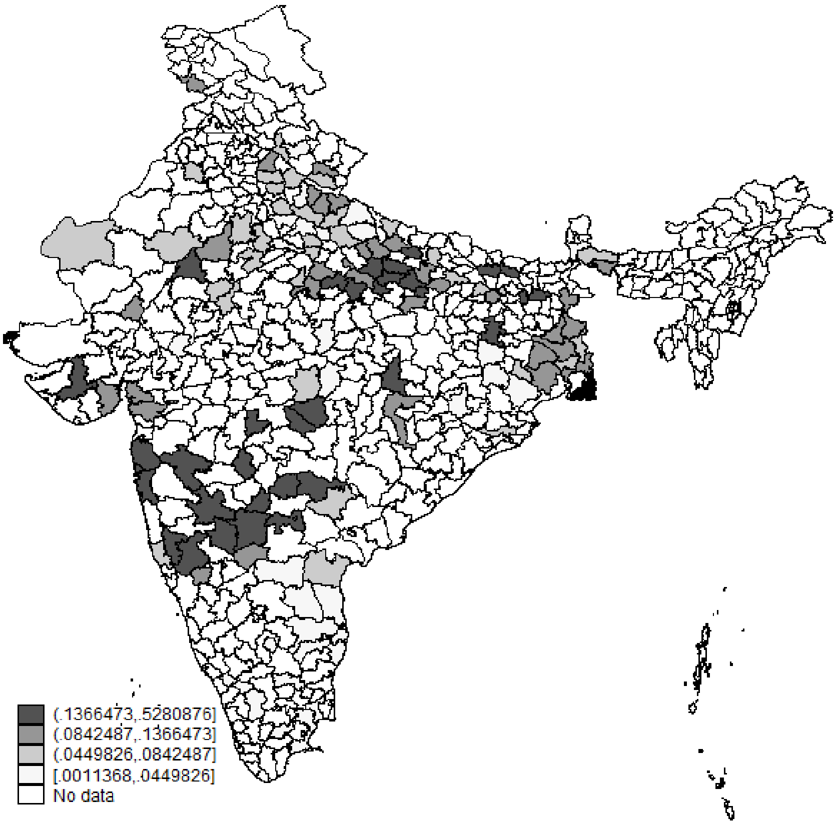
Figure 2.3: Marriage pattern in 1929-1930: time series



The graph plots the proportion of children married in each Census year, by gender and age group, for India as a whole.

Figure 2.4: Distribution of proportion of female married at 5-10

(a) Girls married 5- 10 (%): 1921



(b) Δ Girls married 5- 10 (%): 1931- 1921

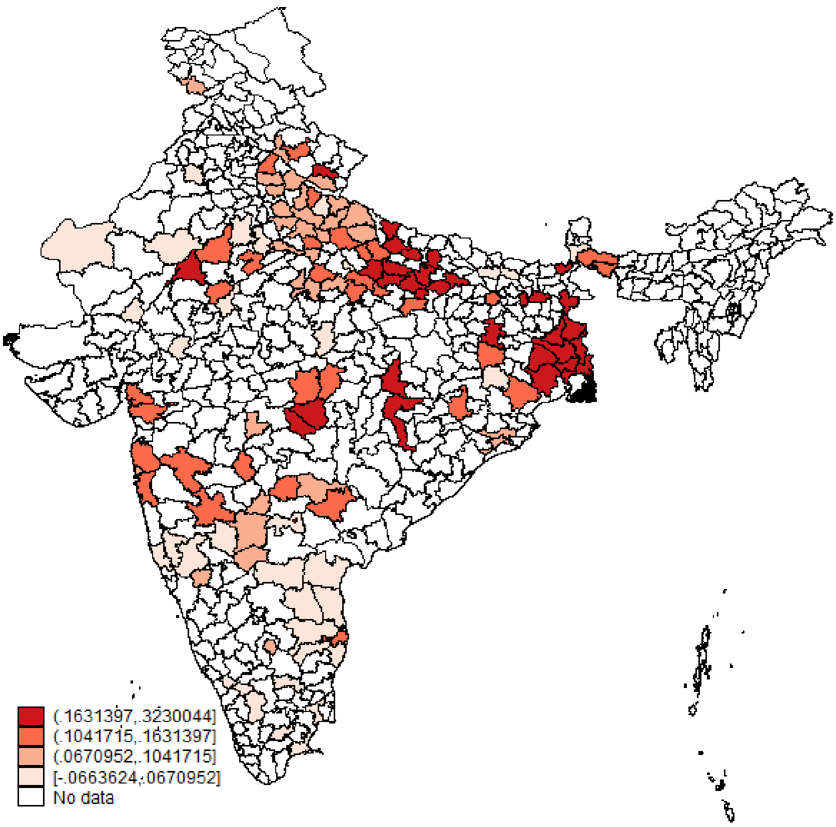


Table 2.1: Summary statistics of the DISE data

	mean	sd
Total boy / total girl enrollment in class 1	1.079	0.093
Total boy / total girl enrollment in class 2	1.070	0.099
Total boy / total girl enrollment in class 3	1.068	0.107
Total boy / total girl enrollment in class 4	1.069	0.115
Total boy / total girl enrollment in class 5	1.076	0.130
Total boy / total girl enrollment in class 6	1.088	0.157
Total boy / total girl enrollment in class 7	1.094	0.176
Total boy / total girl enrollment in class 8	1.103	0.195
Distance to coast	475.330	332.864
Proportion of rural schools	0.885	0.129
Princely states	0.261	0.439
Number of classrooms	4.418	1.778
Observations	2749	
Number of districts	433	

Note: Data aggregated at district level from DISE school records, forming a district level (unbalanced) panel for 2005-2013; Kerala not in the sample; Princely states is a {0,1} indicator. Distance to coast measured in kilometers from the centroid of each district.

Table 2.2: Summary statistics - NSS 64th and 66th round

	1983 mean
female	0.467
Age	11.799
Scheduled caste	0.120
Scheduled tribe	0.166
Head female	0.082
Head literate	0.433
Head primary	0.139
Head complete primary	0.079
Head complete secondary	0.042
Head complete higher than secondary	0.010
Observations	52711

Note: Sample includes children in the NSS 64th and 66th round who aged 10-16 at the time of interview. School attendance is an {0,1} indicator for whether the principal activities in the past year of the child were attending school. Head literate, Head complete primary, Head complete secondary, Head complete higher than secondary were indicators for the education level of the household head, with base group illiterate. Land ownership is an {0,1} indicator for whether the household owns any land.

Table 2.3: OLS Regression of Boy / Girl enrollment ratio at class 5/6: 2005-2013

Outcome: Ratio of boy/girl enrollment		
	Class 6	Class 5
	(1)	(2)
Princely states	0.0201**	0.00839
	(0.00993)	(0.00722)
Number of classrooms	-0.0201***	-0.0131***
	(0.00565)	(0.00476)
Ln GDPPC (2000)	-0.0303*	-0.0176
	(0.0166)	(0.0132)
Observations	2749	2749
State FE	Y	Y
Year FE	Y	Y

Note: Standard errors clustered at district level; The outcome variable is ratio of number of boys enrolled to the number of girls enrolled in each district, year and class, from 2005-2013; Other controls include latitude and distance to coast; Ln GDP per capita measured are district level GDP measured at 2000

Table 2.4: OLS Regression of Boy/Girl enrollment ratio at all class: 2005-2013

	Outcome: Ratio of boy/girl enrollment							
	Class 1	Class 2	Class 3	Class 4	Class 5	Class 6	Class 7	Class 8
Princely states	0.00280	-0.00142	-0.00374	0.0000208	0.00841	0.0201^{**}	0.0300^{***}	0.0256^{**}
	(0.00589)	(0.00593)	(0.00602)	(0.00656)	(0.00723)	(0.00993)	(0.0114)	(0.0123)
Proportion of rural schools	0.0552^{**}	0.0228	0.00334	-0.0205	-0.0478	-0.0814^{**}	-0.0865^{**}	-0.113^{**}
	(0.0259)	(0.0269)	(0.0290)	(0.0318)	(0.0369)	(0.0408)	(0.0432)	(0.0472)
Number of classrooms	0.00596^{**}	0.000161	-0.00387	-0.00744^{**}	-0.0132^{***}	-0.0201^{***}	-0.0234^{***}	-0.0266^{***}
	(0.00269)	(0.00287)	(0.00319)	(0.00367)	(0.00473)	(0.00558)	(0.00597)	(0.00586)
Ln GDPPC (2000)	-0.00785	-0.00905	-0.00938	-0.0141	-0.0173	-0.0303^{*}	-0.0382^{**}	-0.0511^{***}
	(0.00747)	(0.00840)	(0.00986)	(0.0114)	(0.0132)	(0.0166)	(0.0176)	(0.0185)
Observations	2749	2749	2749	2749	2749	2749	2749	2658

Note: Standard errors clustered at district level; The outcome variable is ratio of number of boys enrolled to the number of girls enrolled in each district, year and class, from 2005-2013; Other controls include latitude and distance to coast; Ln GDP per capita measured are district level GDP measured at 2000

Table 2.5: Activity of 10-16 years old: NSS 64 - 66th round

	School	School	Waged work	Waged work	DW	DW
	(1)	(2)	(3)	(4)	(5)	(6)
Princely states=1	-0.00152	0.00601	-0.00565**	-0.00432	0.000973	-0.000434
	(0.00597)	(0.00580)	(0.00264)	(0.00286)	(0.00262)	(0.00162)
Princely states=1 × Female=1		-0.0160**		-0.00284		0.00299
		(0.00630)		(0.00298)		(0.00532)
Observations	150084	150084	150084	150084	150084	150084
state*female FE	Y	Y	Y	Y	Y	Y

Note: Sample includes children in the NSS 64th and 66th round who aged 10-16 at the time of interview. The outcome variables in column (1)-(2),(3)-(4) and (5)-(6) are indicators of school attendance, casual wage labor and domestic work as the principal activity in the year before survey respectively. Controls include: age at time of interview fixed effects; indicator of urban/rural; indicators for muslim, christian sikh and other religions; age of household head; indicators for education level of household head; district distance to coast and Latitude; Indicators for scheduled caste, schedule tribe and other backward caste. All regressions control for state fixed effects; Standard error clustered at district level.

Table 2.6: School attendance - NSS - by religion

	School		Waged work		DW	
	Hindu	Muslim	Hindu	Muslim	Hindu	Muslim
	(1)	(2)	(3)	(4)	(5)	(6)
Princely states=1	0.0115*	-0.0123	-0.00559	-0.00958	-0.0000947	-0.000354
	(0.00638)	(0.0146)	(0.00340)	(0.00641)	(0.00137)	(0.00325)
Princely states=1 × Female=1	-0.0217***	0.0147	-0.00250	0.00268	0.00405	0.00986
	(0.00746)	(0.0236)	(0.00371)	(0.00873)	(0.00534)	(0.0159)
Observations	109689	23137	109689	23137	109689	23137

Note: Sample includes children in the NSS 64th and 66th round who aged 10-16 at the time of interview. The outcome variables in column (1)-(2),(3)-(4) and (5)-(6) are indicators of school attendance, casual wage labor and domestic work as the principal activity in the year before survey respectively. Controls include: age at time of interview fixed effects; indicator of urban/rural; indicators for muslim, christian sikh and other religions; age of household head; indicators for education level of household head; district distance to coast and Latitude; Indicators for scheduled caste, schedule tribe and other backward caste. All regressions control for state fixed effects; Standard error clustered at district level.

Table 2.7: Marriage under legal age and mean age of marriage

Outcome:	% under legal age		Mean age of marriage	
	(1)	(2)	(3)	(4)
Princely states	5.186***	5.018***	-0.399***	-0.329**
	(1.361)	(1.369)	(0.147)	(0.149)
Distance to coast	-0.0141***	-0.0136***	0.000601	0.000172
	(0.00518)	(0.00512)	(0.000586)	(0.000529)
Latitude	1.119**	0.967*	-0.132**	-0.0910*
	(0.523)	(0.514)	(0.0572)	(0.0514)
Ln GDPpc 2000		-11.91***		1.192***
		(1.644)		(0.164)
Observations	568	508	570	508
mean	22.66	23.74	19.44	19.21
State FE	Y	Y	Y	Y

Note: Sample include percentage of marriage under legal age for female*100 reported of each districts, recorded from District Level Household and Facility Survey in 2006-2007 (from DevInfo 3.0); Mean age of marriage are district level mean age of marriage from DLHS 2002-2004 (from DevInfo 3.0); Robust standard errors reported in parenthesis

Table 2.8: The impact of Sarda Act; Census 1911-1931; Difference-in-difference

	Outcome variable: Married female (%) 5-10					
	(1)	(2)	(3)	(4)	(5)	(6)
Sarda act/1931=1 × British Provinces=1	1.631	2.804**				
	(1.433)	(1.311)				
Sarda act/1931=1 × Year of direct control 69-100=1			2.930*	1.686		
			(1.505)	(1.733)		
Sarda act/1931=1 × Year of direct control 101-130=1			-2.156*	-2.069		
			(1.096)	(1.373)		
Sarda act/1931=1 × Year of direct control 131-160=1			2.290	1.241		
			(2.973)	(2.577)		
Sarda act/1931=1 × Year of direct control $\geq$ 160=1			6.558***	5.836***		
			(1.745)	(1.768)		
Sarda act/1931=1 × Ln dist. to reformers					-0.442	-1.489***
					(0.559)	(0.502)
Observations	694	694	694	694	525	525
District FE	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y
Province specific trend	N	Y	N	Y	N	Y

Note: Sample includes a panel that consists of 278 districts according to pre-independence administrative division, for the three Census year from 1911, 1921 and 1931. Sarda act/1931 is a dummy for the year 1931, capture the effect of Sarda act announced and implemented in 1929-1930. Base group in column (1)-(2) are districts that were princely states; Base group in column (3)-(4) are districts where year of acquisition are not reported from the map from Baden (1898) and thus were under indirect control. All columns include district and year fixed effects. Column (2), (3) and (6) control for province (pre-independence division) specific time trend. Robust standard errors reported in parenthesis.

Table 2.9: Long run impact of Sarda Act, 1931 marriage ratio instrumented

(a) Panel A: Marriage in 2002

	Probability of getting married at					
	<14		14-17		<18	
	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)
Married female 5-10 (%) 1931	0.000734	0.00126	-0.00190*	-0.0111***	-0.00116	-0.00982**
	(0.000635)	(0.00220)	(0.00110)	(0.00407)	(0.00134)	(0.00499)
Married female 5-10 (%) 1921	0.00109	0.000428	0.00445**	0.0146***	0.00554***	0.0150***
	(0.000855)	(0.00243)	(0.00172)	(0.00454)	(0.00189)	(0.00551)
Observations	86214	84614	86214	84614	86214	84614
Number of dist.	126	123	126	123	126	123
F-stat		7.906		7.906		7.906

(b) Panel B: Schooling in 2002

	Any schooling		Year of schooling	
	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)
Married female 5-10 (%) 1931	0.00440***	0.0336**	0.0105	0.0921*
	(0.00146)	(0.0134)	(0.00715)	(0.0497)
Married female 5-10 (%) 1921	-0.00619***	-0.0384**	-0.0145	-0.104*
	(0.00197)	(0.0149)	(0.00931)	(0.0555)
Observations	86214	86214	41843	41843
Number of dist.	126	126	126	126
F-stat		6.567		4.352

Note: Sample includes Hindu female 18 years old or older at the time of interview in DLHS 2002. Married female 5-10 (%) are defined by the total number of married female at age 5-10, divided by the total number of female at age 5-10 reported in the Census for each district d, scaled up by 100 to the unit of percentage point. Column (1), (3), (5) in panel (a) and column (1), (3) in panel (b) report the OLS estimates of equation 2.6. Column (2), (4) and (6) in panel (a) and column (2) and (4) in panel (b) report the IV estimates of equation 2.6 - Married female 5-10 1931 are instrumented by: an $\{0, 1\}$ indicator of whether a district belongs to Princely states and ln distance of a district to the nearest birth place of any reformers (measured in km) in the IV specifications. Other controls includes: indicators for caste, urban/rural, age at time of interview; district level distance to coast, ln GDPpc (2000), Latitude; Historical districts are mapped to modern district by taking unweighted average if more than one historical districts are belonging to one modern district; Standard errors clustered at district level.

Table 2.10: Long run impact of Sarda Act, 1931 marriage ratio instrumented

(a) Panel A: Marriage in 2002

	Probability of getting married at					
	<14		14-17		<18	
	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)
Married female 5-10 (%) 1931	0.000642	0.00249	-0.00157	-0.00479	-0.000931	-0.00230
	(0.000662)	(0.00334)	(0.00113)	(0.00336)	(0.00141)	(0.00527)
Married female 5-10 (%) 1921	0.00120	-0.00104	0.00328**	0.00703*	0.00448**	0.00599
	(0.000912)	(0.00397)	(0.00158)	(0.00391)	(0.00179)	(0.00603)
Observations	86214	84614	86214	84614	86214	84614
Number of dist.	126	123	126	123	126	123
F-stat		5.200		5.200		5.200
State FE	Y	Y	Y	Y	Y	Y

(b) Panel B: Schooling in 2002

	Any schooling		Year of schooling	
	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)
Married female 5-10 (%) 1931	0.00122	0.0321	0.0109*	-0.0218
	(0.00103)	(0.0556)	(0.00585)	(0.0748)
Married female 5-10 (%) 1921	-0.00248*	-0.0373	-0.0229***	0.0129
	(0.00143)	(0.0625)	(0.00872)	(0.0829)
Observations	86214	86214	41843	41843
Number of dist.	126	126	126	126
F-stat		0.338		0.547
State FE	Y	Y	Y	Y

Note: Sample includes Hindu female 18 years old or older at the time of interview in DLHS 2002. Married female 5-10 (%) are defined by the total number of married female at age 5-10, divided by the total number of female at age 5-10 reported in the Census for each district d, scaled up by 100 to the unit of percentage point. Column (1), (3), (5) in panel (a) and column (1), (3) in panel (b) report the OLS estimates of equation 2.6. Column (2), (4) and (6) in panel (a) and column (2) and (4) in panel (b) report the IV estimates of equation 2.6 - Married female 5-10 1931 are instrumented by: an $\{0, 1\}$ indicator of whether a district belongs to Princely states and ln distance of a district to the nearest birth place of any reformers (measured in km) in the IV specifications. Other controls includes: indicators for caste, urban/rural, age at time of interview; district level distance to coast, ln GDPpc (2000), Latitude; Historical districts are mapped to modern district by taking unweighted average if more than one historical districts are belonging to one modern district; Standard errors clustered at district level.

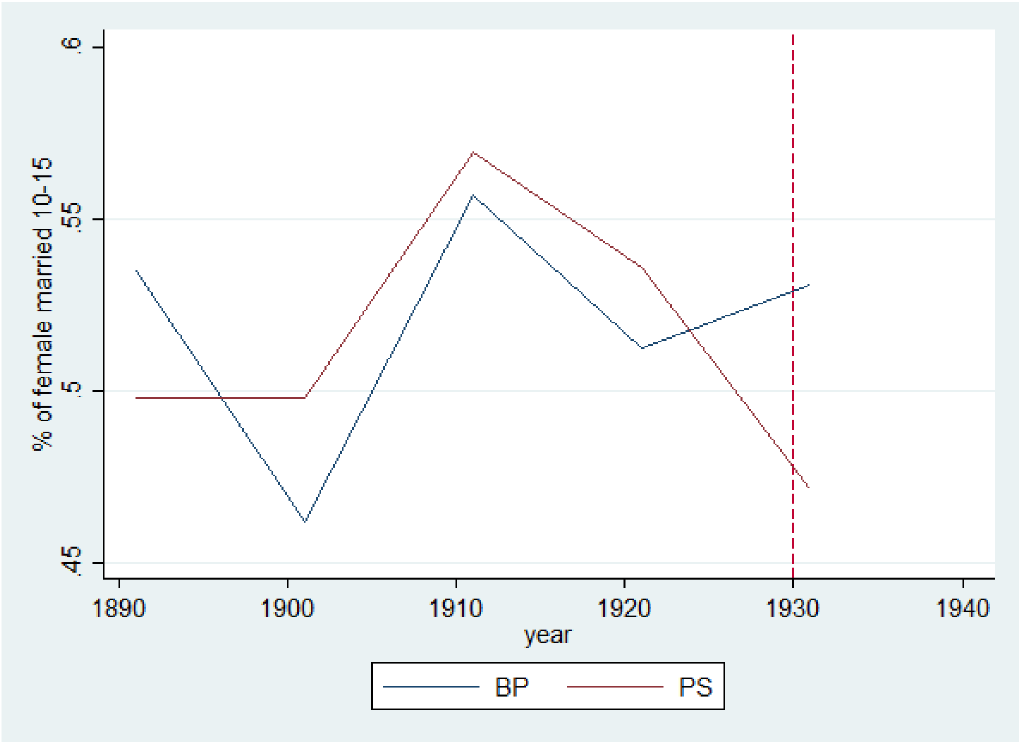
Appendix

2.A Appendix of chapter 2

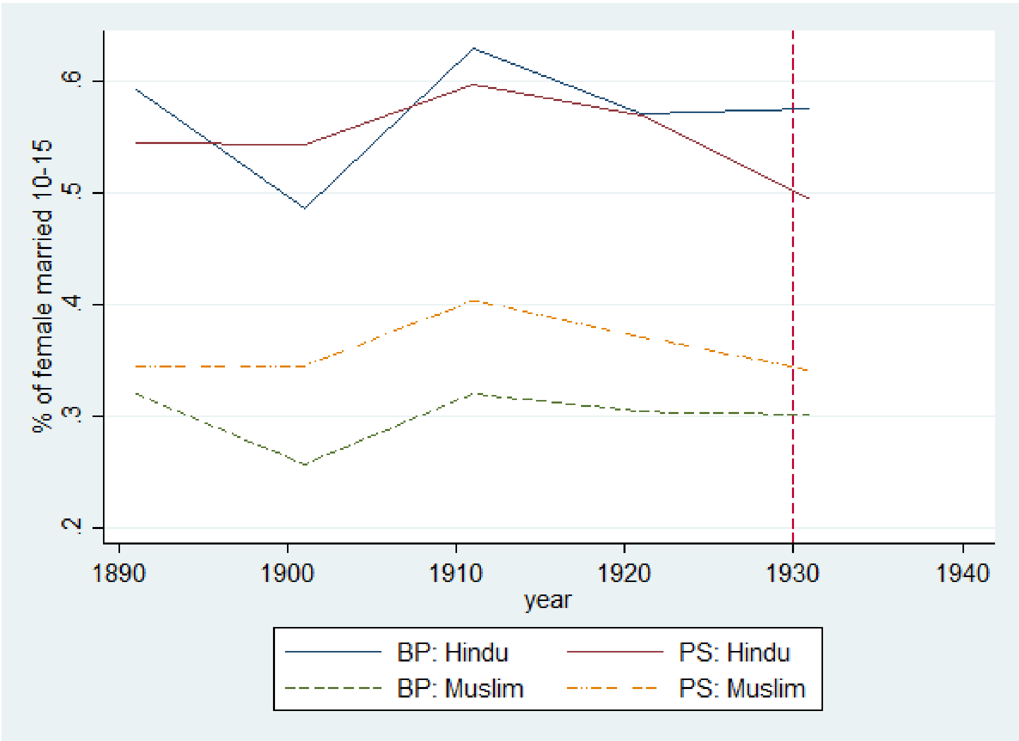
Figure 2.A.1: Geographical distribution of birth place of pro-Sarda Act reformers

Figure 2.A.2: Percentage of married female - 10-15 years old - Madhya Pradesh

(a) All religion



(b) Hindu - Muslim



Note: Data from Census of India 1891-1931 and cover Central Provinces and Central India Agencies which belongs to Madha Pradesh today; the red line denotes the enactment of the Sarda Act

Table 2.A.1: Robustness check - exclusion of Mysore and Baroda: NSS Education

	School (1)	School (2)
Princely states=1	0.00601	0.00556
	(0.00580)	(0.00580)
Princely states=1 × Female=1	-0.0160**	-0.0186***
	(0.00630)	(0.00650)
Observations	150084	146074
Include Mysore and Baroda	Y	N

Note: Sample includes children in the NSS 64th and 66th round who aged 10-16 at the time of interview; Specifiation same as in Table 2.5

Table 2.A.2: Robustness check - exclusion of Mysore and Baroda: Marriage

outcome:	% under legal age		mean age	
	(1)	(2)	(3)	(4)
Princely states	5.018***	6.067***	-0.329**	-0.375**
	(1.369)	(1.367)	(0.149)	(0.151)
Distance to coast	-0.0136***	-0.0121**	0.000172	0.000217
	(0.00512)	(0.00516)	(0.000529)	(0.000532)
Latitude	0.967*	0.633	-0.0910*	-0.0896
	(0.514)	(0.529)	(0.0514)	(0.0550)
Ln GDPpc 2000	-11.91***	-11.75***	1.192***	1.190***
	(1.644)	(1.668)	(0.164)	(0.167)
Observations	508	496	508	496
Include Mysore and Baroda	Y	N	Y	N

Note: Sample include percentage of marriage under legal age for female*100 reported of each districts, recorded from District Level Household and Facility Survey in 2006-2007 (from DevInfo 3.0); Mean age of marriage are district level mean age of marriage from DLHS 2002-2004 (from DevInfo 3.0); Robust standard errors reported in parenthesis; Specification same as in Table 2.7

Table 2.A.3: Robustness check - exclusion of Princely States in the south: Marriage

outcome:	% under legal age			
	All states	Karnataka (exclu. Hyderabad)	Kerala	Exclu. Ke and Ka
	(1)	(2)	(3)	(4)
Princely states	5.018***	-7.346**	-7.188	6.550***
	(1.369)	(3.247)	(4.657)	(1.485)
Distance to coast	-0.0136***	0.0930***	0.0961	-0.0112**
	(0.00512)	(0.0252)	(0.0727)	(0.00522)
Latitude	0.967*	3.128	-1.270	0.168
	(0.514)	(2.497)	(1.458)	(0.553)
Ln GDPpc 2000	-11.91***	-8.846	-16.24	-11.03***
	(1.644)	(9.318)	(21.00)	(1.674)
Observations	508	23	14	467

Note: Sample include percentage of marriage under legal age for female*100 reported of each districts, recorded from District Level Household and Facility Survey in 2006-2007 (from Dev-Info 3.0); Mean age of marriage are district level mean age of marriage from DLHS 2002-2004 (from DevInfo 3.0); Robust standard errors reported in parenthesis; specification same as in Table 2.7

Chapter 3

On the Quantity and Quality of Girls: New Evidence on Abortion, Fertility, and Parental Investments[1]

The introduction of prenatal sex-detection technologies in India has led to a phenomenal increase in abortion of female fetuses. We investigate their impact on son-biased fertility stopping behavior, parental investments in girls relative to boys, and the relative chances of girls surviving after birth. We find a moderation of son-biased fertility, erosion of gender gaps in breastfeeding and immunization, and convergence in the post-neonatal mortality rates of boys and girls. For every five aborted girls, we estimate that roughly one additional girl survives to age five. Our findings have implications not only for counts of missing girls but also for the later life outcomes of girls, conditioned by greater early life investments in them.

3.1 Introduction

Innovations in birth control technology have had substantial socioeconomic impacts. The contraceptive pill, for instance, gave women unprecedented control over fertility, preventing unwanted births and allowing women to determine the timing of births, with dramatic consequences for their marriage and labor market choices (Goldin and

[1] This chapter is a joint work with S Anukriti and Sonia Bhalotra.

Katz (2002), Bailey (2006)). The legislation of abortion has similarly empowered women by enabling them to eliminate unwanted births (Gruber et al. (1999)). Birth control technology has had far-reaching implications not only for women but also for children, as parental investments tend to be greater in children that are more "wanted" when they are born (Grossman and Joyce (1990), Gruber et al. (1999), Donohue and Levitt (2001), Charles and Stephens Jr. (2006), Donohue et al. (2009)).

This paper investigates the impacts on child quantity and quality of another new technology that has altered the demographic landscape in countries where sons are valued more than daughters, namely, prenatal sex detection technology (henceforth, ultrasound; see Section 3.2). An ultrasound scan can reveal fetal sex reliably at as early as 12 weeks of gestation, enabling selective abortion of unwanted girls without, in principle, risking the mother's health (Epner et al. (1998)). We focus on India, where ultrasound was introduced in the mid-1980s, before which abortion had been legalized. The low cost and the non-invasive nature of ultrasound scans has led to their widespread use for fetal sex determination, resulting in a staggering rise in sex-selective abortion, equivalent to 6 percent of potential female births during 1995-2005 (Bhalotra and Cochrane (2010)).[2]

When fetal sex determination is impossible, parents can adjust the gender composition of their children in two ways. First, they could continue childbearing till they achieve the desired number of sons. Several studies document son-biased fertility stopping behavior, which results in girls having more siblings than boys ((Clark (2000), Bhalotra and van Soest (2008), Jensen (2012), Rosenblum (2013)). The quantity-quality trade-off, driven by the budget constraint, implies that a gender gap in outcomes will emerge simply because girls on average grow up in families with fewer per capita resources, even if parents do not actively discriminate against daughters. The second option is to subject girls to deliberate neglect, culminating as excess girl mortality in early childhood (Das Gupta (1987), Pitt and Rosenzweig (1990), Sen (1990), Rose (2000), Oster (2009), Bhalotra (2010), Jayachandran and Kuziemko (2011)). In this paper, we test the hypothesis that the facility to detect and, subsequently, terminate unwanted female fetuses in the post-ultrasound era weakened both son-biased fertility stopping behavior and postnatal discrimination against girls, measured by child investments and mortality rates.

It is challenging to identify the behavioral effects of prenatal sex detection technology when there is selection into conception or abortion. Moreover, the direction of selection into conception and abortion is *a priori* ambiguous (Gruber et al. (1999),

[2] Bhalotra and Cochrane (2010) estimate that 480,000 girls—greater than the number of girls born in the United Kingdom each year—were aborted per year in India during 1995-2005.

Pop-Eleches (2006), Ananat et al. (2009)).[3] A second issue is that revelation of the sex of the child may lead parents to deliberately lower fetal investments in girls (Almond et al. (2010), Bharadwaj and Lakdawala (2013)). Any resulting gender-biased miscarriage, by creating selection into birth, can make it harder to interpret changes in post-birth investments as causal effects of availability of prenatal sex detection (through it facilitating sex-selective abortion). We attempt to address these challenges.

Our strategy combines supply-driven changes in ultrasound availability with plausibly exogenous family-level variation in the incentive to sex-select. We construct an indicator for cohorts born pre- versus post-ultrasound exploiting information on the first imports of ultrasound scanners after tariff reductions in the mid-1980s. We construct a second indicator for cohorts born after a major expansion in availability associated with initiation of local production from the mid-1990s, driven by relaxation of industrial licensing regulations. Previous work shows that these supply-side changes resulted in a phenomenal rise in sex-selective abortion (Bhalotra and Cochrane (2010)). Family-level variation in the incentive to conduct sex-selection is captured in an indicator for the sex of the firstborn child. Previous research shows that (a) the sex of the firstborn child is quasi-random and (b) that sex-selective abortion among second- and higher-order births is concentrated in families with a firstborn daughter (Almond and Edlund (2008), Abrevaya (2009), Bhalotra and Cochrane (2010)). The identifying assumption, which we verify, is that, in the absence of ultrasound technology, trends in the outcomes would have been identical across first-son and first-daughter families. We distinguish neonatal and post-neonatal child survival rates because changes in prenatal investments are more likely to exhibit in neonatal rates, while post-neonatal survival is more clearly a function of postnatal investments.[4]

We use data containing the complete fertility histories of women to construct indicators of births and deaths, as a result of which we can link biological siblings. This allows us to use mother fixed effects to account for selection. In contrast to earlier studies of abortion (Gruber et al. (1999), Pop-Eleches (2006), Ananat et al. (2009)), we analyze parental investments in and survival of girls *relative* to boys. Differencing by gender allows us to control for unobservable trends that equally affect both boys and girls. Another useful feature of the data, that we exploit, is

[3] Pop-Eleches (2006) discusses studies of abortion in developed countries, highlighting that many preceding his study do not adjust for selection. Selection into conception is a persistent issue in analyses of birth outcomes; see, for instance, Dehejia and Lleras-Muney (2004), Bhalotra (2010), Ananat and Hungerman (2012), and Brown (2016).

[4] Neonatal mortality refers to death within one month of birth and post-neonatal child mortality means death after the first month of birth but before age five.

that they contain the mother's stated preferences, including desired fertility and the desired number of sons versus daughters. We investigate the robustness of our estimates to conditioning upon changes in these preferences in a flexible manner. This allows us to more confidently attribute any changes in the outcomes to changes in (ultrasound) technology rather than to other factors such as modernization, which may have contemporaneously changed (stated) preferences.

We estimate reduced form equations for three sets of outcomes. First, we estimate changes in the relative survival of girls. Since we know from previous research that the introduction of ultrasound led to a sharp increase in female feticide, our expectation is that this substituted for postnatal discrimination against girls, leading to a reduction in girl relative to boy mortality after birth. We examine results by birth order and the socioeconomic status of parents so as to identify where in these distributions the relevant changes occur. Second, we estimate changes in parental investments in girls relative to boys, so as to tie any changes in survival to parental behavioral choices. Third, we investigate impacts on fertility defined, as is usual, as live births. In general, increased opportunities for abortion may lead to lower or higher fertility (e.g., Ananat et al. (2009), Ananat and Hungerman (2012)). Abortion mechanically reduces the number of live births conditional upon the number of pregnancies; but the knowledge that a pregnancy can be aborted may stimulate more pregnancies and the net effect on fertility is ambiguous *a priori*. However, since ultrasound led to sex-selective abortion and not abortion *per se*, we can test an unambiguous prediction, which is that there is less son-biased fertility stopping post-ultrasound.

Our main findings are as follows. We find that the pre-ultrasound gender gap in *post-neonatal* child mortality for second- and higher-order births in firstborn-girl families relative to firstborn-boy families (equal to 2.17 percentage points (p.p.)) is reduced by ultrasound technology. Similarly, there is closure of the gender gap in sibling size between firstborn-girl and firstborn-boy families (which, pre-ultrasound, was 0.14). Consonant with these results, we find a significant narrowing of gender gaps in parental investments. A simple accounting exercise suggests that increases in vaccination and breastfeeding in favor of girls explain 34 to 38 percent of the observed decline in post-neonatal excess female mortality (EFM).[5]

Nevertheless, the decline in postnatal death of girls only partially offsets the rise in female feticide. Our estimates imply that 60,879 excess postnatal female child deaths were averted each year (via ultrasound-driven declines in male-biased fertility

[5]Note that vaccination and breastfeeding are only two of the many markers of parental investments. It is possible that parents also increased other investments in girls and that can further explain the observed EFM decline; however, we do not have data on the same.

and substitution of postnatal discrimination with prenatal discrimination), but that the rise in the number of aborted girls was much higher. For every five girls that "went missing" before birth, only one girl survived after birth who otherwise would have died.

If sex-selective abortion is concentrated among families with stronger son preference then girls born in the post-ultrasound era will be more likely to be born into families with weaker son preference, and this would then be a reason that they do better. We investigate this and find no evidence of such sorting. Instead, there are marked socioeconomic status (SES) patterns in sex-selection, with wealthy and educated women being more likely to abort girls (Bhalotra and Cochrane (2010)). So the average girl born post-ultrasound (relative to pre-ultrasound) is more likely to be born into a low SES family.

The relative fertility decline in firstborn-girl families is driven by a shift from having four or more children toward having three or fewer children. The EFM decline is concentrated in first, second, and third children in firstborn-girl families. In general, the absolute decline in excess girl mortality is larger among women of low SES, while the decline in fertility among women with firstborn daughters is broadly similar across socioeconomic groups. Previous work shows that sex-selective abortion in India is more prevalent among educated and wealthy women (Bhalotra and Cochrane (2010)). So, responses to ultrasound technology in terms of sex-selective abortion, fertility, and postnatal investments in girls versus boys occur on different SES margins. These results suggest that ultrasound availability generated a shift in the distribution of girls in favor of low-SES families since more girls were aborted in high-SES families and more girls survived to age five in low-SES families.

While a number of studies have sought to identify how changes in abortion law in different countries have modified fertility and investment in children, the evidence on how prenatal sex detection technology (which made sex-selective abortion feasible) modifies fertility and investment in girls is more limited. Lin et al. (2014) show that abortion legalization in Taiwan decreased neonatal female mortality for higher-parity births conditional on SES and had no impact on post-neonatal mortality. They also find a reduction in fertility at third and higher parity, and a shift in the composition of births towards low-SES families. They do not examine health investments. Almond et al. (2010) find, in contrast to us, that ultrasound access in China *increased* neonatal EFM, which suggests that parents consciously reduced prenatal inputs in girls. They find no impact on gender gaps in post-neonatal mortality, and they do not examine fertility. Our finding of a post-ultrasound decline in excess post-neonatal girl mortality is empirically relevant, given that the gender gap in post-neonatal child mortality in India is quite large—1.5 p.p. overall and 3 p.p. among children preceded by a firstborn girl (Table 3.A.8).[6]

The paper most closely related to ours is Hu and Schlosser (2015), that seeks to answer a similar question, but using a different empirical approach. They model gender gaps in malnutrition and mortality of children as a function of state-year variation in the sex ratio at birth. However, the state level sex ratio at birth is jointly determined with the outcomes. Once fetal sex detection is feasible, parents will simultaneously decide whether to conceive, whether to use prenatal sex diagnosis, whether to abort if the fetus is a girl, and how much to invest in male versus female births that are taken to term. The joint outcomes, thus, are fertility, the sex ratio at birth, post-birth investments in girls relative to boys, and girl relative to boy survival. The identifying assumption in Hu and Schlosser (2015) is that changes in the average sex ratio at birth within a state over time are uncorrelated with unobserved factors that could differentially affect male and female outcomes. Since we model excess female mortality by firstborn sex (which varies at the household level), we can difference out all such omitted variables.

Hu and Schlosser (2015) find a narrowing of gender gaps in malnutrition but, despite this, no change in excess girl mortality. They state: "These results are somewhat at odds with our previous findings on nutritional outcomes, family size and breastfeeding duration. Particularly puzzling is why we find a differential improvement in female nutritional status but do not see any significant increase in female survival probabilities."[7] Using a different statistical approach, we find large post-ultrasound changes in mortality rates that result in a significant closure of the boy-girl mortality gap, specifically in firstborn-girl households where this gap is marked for pre-ultrasound birth cohorts. In this way, we resolve the puzzle.[8]

[6] In comparison, before the 1985 abortion legalization, the average post-neonatal infant EFM in Taiwan in 1981 was -0.11 p.p., i.e., not biased against girls (Yang et al. (1996)). Moreover, India and China differ in the age distribution of missing girls. In China, the imbalance in the sex ratio for under-5 children in primarily at birth (83 percent), while there is a more even spread across early childhood in India (Table 3.A.9). In particular, post-neonatal EFM contributes 33 percent in India but only 3 percent in China.

[7] page 1257 Hu and Schlosser (2015)

[8] This difference in findings appears to be for the following reasons. First, their sample includes few pre-ultrasound cohorts. In one case they use children born just three years before the survey date, and in the other extend this to ten years, which implies either no pre-ultrasound cohorts or, for the earliest survey round, only a few. A second important factor is that they pool firstborn-girl and firstborn-boy families. Previous research, cited above, shows that the introduction of ultrasound led to sex-selection primarily in firstborn-girl families. The summary statistics we present for EFM in the pre-ultrasound period similarly show that it was concentrated in firstborn-girl families. For instance, the girl-boy differential in post-neonatal child mortality was 0.9 p.p. for births preceded by a firstborn boy but 3 p.p. for births preceded by a firstborn girl. This is an enormous difference, so a specification that forces equal coefficients for the two groups may veil relevant changes. If we select a shorter sample like Hu and Schlosser (2015) do and if our

More generally, we take this research agenda forward. The large changes in girl relative to boy *survival* that we document contribute new evidence to current debates on biased population sex ratios in India. Moreover, we investigate SES gradients in the gender-differentiated survival and fertility responses to the introduction of ultrasound technology. The dynamic implications of recent changes in the quantity and quality of girls (relative to boys) in India depend on where in the SES distribution these changes occur, for instance, because of marital hypergamy (Edlund (1999)).

The results of this study are of considerable significance. First, we find that there is increased post-neonatal survival of girls relative to boys post-ultrasound but that this is more than offset by increased prenatal loss of girls due to feticide. So the number of missing women continues to increase, and this has implications for violence against women, prostitution, sexually-transmitted diseases, marriage market imbalances, and elderly care (Edlund (1999), Drèze and Khera (2000), Kaur (2004), Edlund et al. (2007), Ahlawat (2009), Ebenstein and Sharygin (2009), Bhaskar (2011), Amaral and Bhalotra (2016)). Second, we find that girls born in the post-ultrasound era are receiving greater early-life investments, with pre-existing boy-girl gaps in immunization and breastfeeding narrowing considerably. This not only contributes to their survival but is predictive of improvements in cognitive attainment, income, and longevity for girls (Bhalotra and Venkataramani (2013), Bhalotra and Venkataramani (2015), Currie and Rossin-Slater (2015), Bhalotra et al. (2016)), and indeed the outcomes of their offspring (Currie and Moretti (2007), Almond and Currie (2011), Bhalotra and Rawlings (2011)). Narrowing of gender differences in human capital also tends to be associated with higher growth rates and social change (Klasen (2002b), Lagerlöf (2003)). Third, we find that excess girl mortality is falling more in low-SES households. Together with the finding in Bhalotra and Cochrane (2010)) that female feticide is greater in high-SES households, this means that, over time, the share of girls in low-SES households is increasing. This result aligns with the predictions of Edlund (1999) and, as she discusses, has implications for marriage and violence. Fourth, the fertility decline (concentrated in firstborn-girl families)

specification does not distinguish firstborn-girl and firstborn-boy families, then we are able to replicate their result that there is no post-ultrasound change in EFM. But our estimates *and* the raw data confirm large reductions. Third, as discussed above, their approach may not sufficiently account for simultaneity and selection. We use information on the first imports of ultrasound scanners and the initiation of local production to generate variation in nationwide availability of prenatal sex-detection techniques, and we are able to include pre-ultrasound cohorts (our data for estimation of both survival and investments contains several pre-ultrasound cohorts). In addition, we employ a mother fixed effects estimator to account for selection. Given the longer span of our data, many mothers have fertility histories that are interrupted by the arrival of ultrasound technology.

we observe not only benefits girls through increased resources per capita, it is also potentially beneficial for the health of mothers, which is depleted by the high levels of fertility motivated by the desire to bear sons (Milazzo (2014)). More generally, fertility decline in developing countries has been shown to be associated with economic growth, human capital accumulation, and women's empowerment (Joshi and Schultz (2007), Rosenzweig and Zhang (2009), Miller (2010), Ashraf et al. (2013)).

The rest of the paper is organized as follows. Section 3.2 describes the Indian context. Sections 3.3 and 3.4 discuss the empirical strategy and the data. Section 3.5 presents results and Section 3.6 presents estimates of the implied magnitude of substitution between postnatal discrimination and sex-selective abortions due to ultrasound access. Section 3.7 concludes.

3.2 Context

While son preference has characterized parts of Indian society for centuries, the availability of affordable prenatal sex-diagnostic techniques combined with legal access to abortion is more recent. Abortion was legalized in India with the passage of the Medical Termination of Pregnancy (MTP) Act in 1971, effective in most states in 1972. The Act specifies the reasons for which an abortion can be legally performed and requires that it be performed by a registered medical practitioner in certified abortion facilities.[9] Abortion is legal if the pregnancy that it terminates endangers the woman's life, causes grave injury to her physical or mental health, is a result of rape or contraceptive failure (the latter applies only to married women), or is likely to result in the birth of a child suffering from serious physical or mental abnormalities. Consent is not required from the woman's husband or from other family members; however, a guardian's consent is required if the woman seeking an abortion is either less than 18 years old or is mentally ill. The Act allows an unintended pregnancy to be terminated up to 20 weeks' gestation; however, if the pregnancy is beyond 12 weeks, approval is required from two medical practitioners (Arnold et al. (2002)). The stated purpose of the Act was to regulate and ensure access to safe abortion, although it has been argued that the political motivation was population control (Phadke (1997)).

Fetal sex determination first became possible in India with the advent of amniocentesis in the 1970s. This technology was introduced to detect genetic abnormalities but was soon being used to detect fetal sex. As early as 1976, the government banned the use of these tests for sex determination in government facilities (Arnold et al.

[9]More information on the certification criteria is available in Stillman et al. (2014).

(2002)). The private sector remained unregulated but widespread use was limited by the high direct cost and the invasiveness of amniocentesis. Fetal sex-selection only really became feasible after 1980—becoming evident at the population level after 1985 and widespread by 1995—with the arrival of ultrasound scanners. Ultrasound availability during the early diffusion period was driven by the liberalization of India's import sector. The first ultrasound scanner was imported in 1987 (Mahal et al. (2006)). Thereafter, the quantity of imports rapidly increased (Figure 3.A.1) as import duties on medical equipment were gradually lowered. The import tariff on medical devices declined from 40 to 60 percent in the 1980s to 25 percent in the late 1990s to 12.5 percent in 2003-04, and then to the currently uniform rate of 5 percent. Domestic production of ultrasound machines grew 15-fold between 1988 and 2003 (George (2006), Grover and Vijayvergia (2006)), following relaxation of industrial licensing regulations. The bottom graph in Figure 3.A.1 shows that, once domestic production began, it was orders of magnitude larger than imports.

Demand for ultrasound scans proliferated as a result of the technology being non-invasive and its wide affordability at about $10-$20 for a scan or an abortion (Arnold et al. (2002)).[10] The trend in ultrasound use (also in Figure 3.A.1) closely tracks the supply of ultrasound machines. Additionally, Figure 3.A.2 shows that the officially reported number of abortions (that includes both sex-selective and other abortions) follows a similar trend and is positively correlated with self-reported ultrasound use during pregnancy. Clinics and portable facilities have mushroomed, advertising availability of ultrasound with slogans conveying that the cost of a scan is much lower than the future costs of dowry.[11] Additional amendments to the MTP Act in 2002 and 2003 increased public sector provision and made abortion safer (Stillman et al. (2014)). Other things equal, this could have contributed to a further increase in feticide since 2002.

Since the late 1980s, sex-selection has become the dominant concern amongst women's and human rights organizations.[12] Their campaigns led to the central

[10] These may be significant costs in a country where many live under the $1.25-a-day line. The costs cumulate if repeated scans and abortions are needed before a boy is conceived and vary with distance of the household from the clinic and with the safety of the procedures.

[11] Dowry is a ubiquitous feature of the Indian marriage market and payments from the bride's family to the groom's family at the time of marriage can amount to several multiples of annual household income (Anukriti et al. (2016)), and can motivate parents to eliminate female births (Bhalotra et al. (2016)).

[12] Feminist and socialist groups in the United States and other richer countries have hotly defended a pro-choice stance against a pro-life stance on abortion. The focus of public discussion is on benefits for women rather than on benefits for children. For instance, http://www.theguardian.com/commentisfree/2014/oct/14/abortion-right-to-privacy-women-right-to-equality and https://socialistworker.org/2013/11/01/abortion-every-womans-right. Indian feminists, on the other hand, have been divided by the seeming contradiction of supporting a

government passing the Prenatal Sex Diagnostic Techniques (Regulation and Prevention of Misuse) (PNDT) Act in 1994. This act was effective from January 1, 1996. The PNDT Act made it illegal to use prenatal sex-diagnostic techniques (like ultrasound) to reveal the sex of a fetus. Following the revelation in the 2001 Census of a continuing deterioration in the sex ratio, the PNDT Act was strengthened by a 2002 Amendment (effective 2003) incorporating a ban on advertising prenatal sex determination and increased penalties for violations.[13] It is widely believed that these regulations have made little difference (Visaria (2005)), although Nandi and Deolalikar (2013) find that they did have some impact. These bans are difficult to enforce because ultrasound (or alternatives like amniocentesis) are also used for medical purposes and in routine prenatal care, making it easy to cover up sex determination as a motive.

In general, the fetal environment has improved in India. The growth in income and the decline in poverty since the early 1980s has been widely documented; fertility decline set in from 1981 (Bhalotra and van Soest (2008)); and neonatal mortality rates have been decreasing. Maternal mortality is estimated to have declined (Bhat (2002)) and maternal age at birth has risen. Improvements in fetal health tend to favor boys, whereas our hypothesis is that the trend, driven by the availability of ultrasound scanners, has been in favor of girls.

3.3 Empirical Methodology

The hypothesis we test in this paper is that the availability of prenatal sex-detection technology (and, in particular, ultrasound) simultaneously modified the decisions to conceive, to use prenatal sex detection and abort if the fetus is female, and to invest, possibly differently, in surviving girls and boys. Observable outcomes of these decisions are live birth, the probability that a birth is female (or, the sex ratio at birth), gender-differentiated investments (like breastfeeding and vaccination) and girl relative to boy mortality rates.

As mentioned in Section 3.1, we exploit exogenous variation in the supply of ultrasound scanners, interacted with quasi-random variation in the family-level proclivity to commit sex-selection (conditional on underlying preferences) to estimate the impacts of ultrasound availability. Since we are interested in gender gaps in parental investments, survival, and sibship size, we use the sex of the child as a third interaction in a triple difference-in-differences (DDD) regression specification.

woman's right to abortion while opposing sex-selective abortion (Kumar (1983), Gangoli (1998)).

[13] More details on the PNDT Act are available in Retherford and Roy (2003) and Visaria (2005).

Since the first imports of ultrasound scanners and the initiation of local production constitute structural breaks in supply at the national level, a simple pre-post comparison of outcomes that relies upon differential exposure to ultrasound technology is at risk of reflecting correlated macro-events.[14] The wave of economic liberalization in India that lowered import tariffs and relaxed licensing of domestic production was, for instance, also associated with greater exposure of women to Western media and with rising incomes for large sections of the population. We address this problem by interacting cohort variation in exposure to the new technology with the sex of the firstborn child of the mother, which is a quasi-random variable indicating the mother's "willingness" to conduct sex-selective abortion.

The assumption that the sex of the first child is randomly determined is supported by the data. The top left graph in Figure 3.A.3 shows that the sex ratio at first birth in India lies within the normal range during our sample period, and there is no tendency to increase over time.[15] Additionally, Table 3.A.11 shows that there are no significant socioeconomic differences between families with a firstborn son and a firstborn daughter. Exogeneity of first-born sex has also been previously defended (Das Gupta and Bhat (1997), Visaria (2005), Bhalotra and Cochrane (2010)) and lines up with recent survey data that suggest that parents do not always prefer having a son over a daughter. Jayachandran (2016) finds that although the vast majority of families want to have a son if they can only have one child, at a family size of two they prefer having one daughter and one son over having two sons. As desired and actual fertility in India are well above one (Table 3.A.10), it is reasonable to assume that parents are not averse to having one daughter, despite a strong desire for at least one son.

Previous studies have established that parents randomly exposed to a "firstborn girl treatment" are more likely to practice sex-selection at higher-parity births (Bhalotra and Cochrane (2010), Rosenblum (2013)) consistent with a documented desire for at least one son. Figure 3.A.3 clearly depicts this pattern: after ultrasound technology became available, second, third, and fourth births became increasingly more male but only for families without a son. So, the interaction with first child's sex captures the differential incentives to sex-select among otherwise similar families.

If sex of the firstborn is random, excess female mortality (EFM) among families with a firstborn girl and a firstborn boy should follow a similar trend before

[14] We do not use measures of state-specific adoption since it is endogenous; see Jayachandran et al. (2010) for a similar argument pertaining to the introduction of antibiotics in the United States. Bhalotra and Cochrane (2010) document that access to ultrasound in India was widespread and that the costs of ultrasound and abortion were not prohibitive even for relatively poor households.

[15] Figure 3.A.3 reproduces Figures 1-4 from Bhalotra and Cochrane (2010).

the availability of prenatal sex-selection technology. Figure 3.A.4 plots the 5-year moving average of EFM for firstborn-boy and firstborn-girl families for our sample period and the differential trend for these two groups. Although EFM is significantly higher for births preceded by a firstborn girl during the pre-ultrasound period, the gap remained constant until 1985, providing support for the identifying assumption of parallel pre-reform trends.

Even if our specification addresses potential concerns about omitted trends, one might still worry that we identify compositional rather than causal effects. For example, if higher SES women were more likely to change their behavior in response to ultrasound availability, then the post-ultrasound composition of births will be lower-risk than pre-ultrasound. If, in addition, the gender gap in outcomes among high SES births is narrower (we discuss the evidence later), any causal effects of ultrasound availability on post-birth gender gaps will be conflated with this compositional effect. More generally, if there was selection into conception, sex-selection, and abortion after prenatal sex detection became feasible, then our estimates of child mortality are subject to bias. We believe that this is not a first-order problem since we effectively difference between first-daughter and first-son families conditional upon SES. We nevertheless address this concern by introducing mother fixed-effects, comparing outcomes for children born to the *same* mother but differentially exposed to ultrasound technology.[16] We also investigate sensitivity of our estimates to conditioning upon stated preferences including desired fertility and the desired number of sons versus daughters.

3.3.1 Regression Specifications

To capture the time variation in ultrasound availability, we split our sample into three broad time-periods, defining 1973-1984 as the pre-ultrasound period, 1985-1994 as the early diffusion period, and 1995-2005 as the late diffusion period when ultrasound supply and use became widespread. Bhalotra and Cochrane (2010) identify 1985 as a break-point in the trend of the average sex ratio at birth using non-parametric plots and flexible parametric specifications. As an imbalanced sex ratio at birth captures sex-selective abortions, the break-point also serves as a proxy for the trend-break in ultrasound availability. Bhalotra and Cochrane (2010) identify 1995 as a second break-point based on the sharp increase in supply of ultrasound scanners following the acceleration of trade liberalization in the early and mid-1990s; this trend-break is clearly visible in Figure 3.A.1. Our results are similar if we vary

[16] This involves selecting a sample of families with at least two children of opposite sex. We therefore present estimates with and without fixed effects to assess coefficient stability.

the precise thresholds used to define the three time periods.

For child i of birth order b born to mother j in year t and state s, we estimate the following specification:[17]

$$\begin{aligned} Y_{ijt} =& \alpha + \beta_1 G_j * F_i * Post_t^1 + \beta_2 G_j * F_i * Post_t^2 \\ &+ \gamma G_j * F_i + \omega_t G_j + \sigma_t F_i \\ &+ \mathbf{X'_{ijt}} \tau + \delta_s F_i + \nu_s G_j + \psi_b F_i + \xi_b G_j \\ &+ \rho_{bt} + \eta_{bs} + \phi_{st} + \epsilon_{ijt} \end{aligned} \tag{3.1}$$

The dependent variable, Y_{ijt}, is either a mortality indicator for child i or a measure of parental investments in children, including breastfeeding and immunization.[18] The indicator variable G_j equals one if the first child of mother j is a girl. The variable F_i equals one if child i is female. $Post_t^1$ indicates that t belongs to the early diffusion period (1985-1994) and $Post_t^2$ indicates that t belongs to the late diffusion period (1995-2005). Attached to the two triple interaction terms are the coefficients of interest.

A vector of socioeconomic and demographic characteristics, $\mathbf{X'_{ijt}}$, comprises indicators for household wealth quintiles, educational attainment of child's parents, mother's birth cohort, mother's age at birth, caste, religion, and residence in a rural area. We also control for the main effects of G_j and F_i and fixed effects for state, birth year (or cohort, of the child), and birth order. We allow birth cohort fixed effects to vary by firstborn sex ($\omega_t G_j$), by child gender ($\sigma_t F_i$), and by birth order (ρ_{bt}), which flexibly control for potential omitted trends. We also allow state and birth order fixed effects to vary by firstborn sex and by child gender ($\delta_s F_i$, $\nu_s G_j$, $\psi_b F_i$, $\xi_b G_j$). In addition, we include state fixed effects specific to birth order (η_{bs}) and to birth year (ϕ_{st}).

Child gender-specific cohort fixed effects ($\sigma_t F_i$) account for any nationwide changes that may influence gender gaps in the outcomes, including improvements in maternal health or prenatal care which we expect benefit male fetuses more than female fetuses given the evidence on greater sensitivity of males to prenatal inputs (Low (2000)). They also account for any trends in son preference associated with modernization. Cohort fixed effects varying by sex of the firstborn child in the family ($\omega_t G_j$) control for nationwide trends that may have differentially affected mortality of children in firstborn-girl versus firstborn-boy families. For instance, the "Trivers-

[17]The variable state refers to the mother's state of residence at the time of survey and may differ from the child's state of birth. Restricting the sample to women who have not migrated between their first birth and the survey date does not substantively change the estimates.

[18]More details on the variables used in the regression analysis are available in Appendix 1.

Willards hypothesis" implies that firstborn-boy families are more often of higher SES than firstborn-girl families, and it is plausible that trends in the outcomes differ by SES and that the SES-observables we control for do not capture every relevant expression of SES.

Allowing the state fixed effects to vary with both child gender ($\delta_s F_i$) and sex of the firstborn child ($\nu_s G_j$) allows state-level time-invariant factors, such as soil quality (Carranza (2015)), to have gender-specific effects and ensures that we absorb any cross-sectional heterogeneity that may be correlated with firstborn sex. We interact indicators for child gender and sex of the firstborn with birth order ($\psi_b F_i$, $\xi_b G_j$) given previous evidence that son preference varies with birth order and that sex of the firstborn child influences the exercise of son preference.

Lastly, ϕ_{st}, η_{bs}, and ρ_{bt} control non-parametrically for, respectively, state-specific time effects (e.g., differential growth rates of state GDP or availability of abortion and other health services), state-specific birth order effects, and birth order specific time effects. This rich set of fixed effects enables us to rule out a wide range of confounding variables and trends that can interfere with a causal interpretation of our findings.

We include first births in our sample and set G_j equal to zero for them.[19] The "control" group comprises pre-ultrasound births, second- and higher-order births to mothers whose firstborn is a boy, and first births. The coefficient γ measures the difference in excess female mortality (EFM) or the gender gap in health investments between the treatment and control groups during the pre-ultrasound period. The coefficients β_1 and β_2 capture how these gaps evolved over the early and late diffusion periods relative to the pre-ultrasound period. Standard errors are clustered by state.

Since our data comprise multiple births per woman, we also estimate specification (1) including mother fixed effects, exploiting the differential exposure of siblings to ultrasound technology. This addresses concerns pertaining to selection on underlying preferences or socioeconomic characteristics correlated with both access to or uptake of ultrasound technology and investments in girls versus boys. This approach also addresses any potential bias arising from a preponderance of male births in higher SES families ("Trivers-Willards hypothesis") even in the absence of sex-selection. Since we do not observe large differences in the socioeconomic characteristics of firstborn-boy and firstborn-girl families (in Table 3.A.11), we expect that selection is of limited concern, but we nevertheless allow for it.

We examine the impact of availability of prenatal sex detection on gender gaps in fertility in two ways. First, we test if ultrasound altered the male-bias in the

[19] For higher-order births, G_j equals one if the first child of mother j is a girl and zero otherwise.

hazard of birth in a given year for firstborn-girl versus firstborn-boy mothers.[20] For this specification, we utilize a retrospective mother-year panel in which a woman enters the panel in her year of marriage and exits in the year of survey. For mother i from state s of age a in year t, who has given birth to $b-1$ children by t and whose last birth took place r years ago, we estimate the following logistic regression:

$$\begin{aligned} Birth_{it} =& \alpha + \beta_1 G_i * Post_t^1 + \beta_2 G_i * Post_t^2 + \gamma G_i + \omega_t \\ & + \mathbf{X_i'}\tau + \phi_a + \psi_b + \sigma_r + \delta_s + \nu_s G_i + \theta_{st} + \epsilon_{it} \end{aligned} \tag{3.2}$$

The outcome variable, $Birth_{it}$ equals one if the mother gives birth in year t and is zero otherwise. $Post_t^1$, $Post_t^2$, and G_j are defined as earlier. The vector $\mathbf{X_i'}$ comprises indicators for household wealth quintiles, educational attainment of the mother and her husband, caste, religion, residence in a rural area, and mother's year of birth. We include fixed effects for year (ω_t), state (δ_s), mother's age (ϕ_a), parity (ψ_b), and years since last birth (σ_r), state-specific firstborn-girl fixed effects ($\nu_s G_j$), and state-specific year fixed effects (θ_{st}).[21]

Additionally, we attempt to estimate the effect on the "stock" of children a woman has at the time of survey.[22] Specifically, we estimate the following specification for woman j in state s who has N_{jt} children in the year of survey, t:

$$\begin{aligned} N_{jt} =& \alpha + \beta_1 G_j * Post_t^1 + \beta_2 G_j * Post_t^2 + \gamma G_j + \sigma Post_t^1 + \psi Post_t^2 \\ & + \mathbf{X_j'}\tau + \delta_s + \nu_s G_j + \theta_s Post_t^1 + \omega_s Post_t^2 + \epsilon_{jt} \end{aligned} \tag{3.3}$$

We restrict the sample to mothers who either were always exposed or never exposed to ultrasound for the year-span of their births. In other words, we retain women who had all their births strictly within one of the three time-periods—pre-ultrasound, early diffusion, or late diffusion. $Post_t^1$ and $Post_t^2$ indicate that a woman began and completed childbearing respectively during 1985-1995 and after 1995. The variable G_j is, as before, an indicator for the firstborn being a girl. The vector $\mathbf{X_j'}$ comprises indicators for household wealth quintiles, educational attainment of the woman and her husband, caste, religion, residence in a rural area, woman's birth year, and woman's age at the time of survey. We include fixed effects for the woman's

[20]Since our empirical strategy relies on the sex of the first birth, the sample excludes the 11 percent of women in the data who had never given birth by the time of the survey.

[21]We also modify this specification by including mother fixed effects to test if ultrasound availability delays time to the next birth for a given mother, conditional upon the time since last birth, in the post-ultrasound (relative to the pre-ultrasound) period and whether this delay is on average greater in firstborn-girl (relative to firstborn-boy) families. The results from this specification are available upon request.

[22]Like specification (2), here too we exclude women who had never given birth by the time of the survey.

birth year as fertility is right-censored for some women. Moreover, we include state fixed effects (δ_s), state-specific firstborn-girl fixed effects ($\nu_s G_j$), and allow the effects of the post-ultrasound indicators to vary by state ($\theta_s Post_t^1$ and $\omega_s Post_t^2$). The coefficient γ provides an indication of the extent to which the pre-ultrasound period was characterized by son-biased fertility stopping. The coefficients β_1 and β_2 test our hypothesis that after fetal sex determination became feasible, there was less son-bias in fertility decisions, i.e., the fertility difference between families with a firstborn girl and families with a firstborn boy narrowed.[23]

3.4 Data and Descriptive Statistics

The mortality and fertility equations are estimated using three pooled rounds of the National Family Health Survey (NFHS) conducted in 1992-93, 1998-99, and 2005-06. These nationwide, repeated, cross-sectional surveys are representative at the state level and report complete birth histories for all interviewed women, including children's month, year, and order of birth, mother's age at birth, and age at death of deceased children.[24] The sample comprises 503,316 births of 232,259 mothers that occurred in 1973-2005.

For our mortality and postnatal health specifications, we pool all births of the surveyed women to create a child-level dataset. The hazard of birth specification, on the other hand, is a retrospective mother-year panel that is created from women's retrospective birth histories. Our third specification, that examines the fertility stock, simply pools all surveyed women to form a woman-level dataset.

In the pre-ultrasound era, there were large gender gaps in post-neonatal child mortality. Trends in EFM by the sex of the first child are depicted in Figure 3.A.4 and Table 3.A.1. During 1973-1984, girls preceded by a firstborn girl were 3 p.p. more likely than boys to suffer post-neonatal child mortality. After 1985, following introduction of ultrasound, there is a clear decline in the gender gap in mortality in this group, to 1.78 p.p. during 1985-1994 and to 1.18 p.p. during 1995-2005. There

[23]In principle, the excluded mothers, whose fertility spans more than one period should be similar to the included mothers. We checked for balance and found that in fact the excluded mothers are, on average, of lower SES and are older. However, we always control for SES (education, wealth, urban) and flexibly control for age, and, within the sample, our estimates identify differences by firstborn sex. In any case, we also present fertility results separately for each SES group (within our sample) and, as we show later, for most SES indicators, differences in the coefficients of interest are small. Lastly, our results are robust to the inclusion of fixed effects for the years of first and last birth of every women and fixed effects for their interactions with each other, with state fixed effects, and with the firstborn sex indicator.

[24]Since the state of Sikkim changed its border during the period of analysis, we exclude it from our sample.

was also a marked decline in the gender gap in mortality in families with a firstborn son, but from a smaller initial level. Among pre-ultrasound cohorts, the gap was 0.88 p.p. and this declined to 0.50 p.p. in the early diffusion period and 0.15 p.p. during the late diffusion period.

Table 3.A.11 reports summary statistics for the main variables used in the analysis, by the sex of the firstborn child for pre- and post-ultrasound periods. The fraction of female births, a marker of sex-selective abortion, declined from 0.48 in the pre-ultrasound period to 0.47 in the late diffusion period for births preceded by a firstborn girl, but not for births preceded by a firstborn boy. Fertility is consistently higher for mothers with a firstborn daughter than for mothers with a firstborn son, illustrating son-biased fertility stopping. This difference narrows in the post-ultrasound period, consistent with our prediction that sex-selective abortion substituted son-biased fertility behavior as a way of achieving the desired sex composition of births. Mother's age at birth has increased over time, and more so for mothers with a firstborn girl, consistent with reduced fertility.

Firstborn-boy and firstborn-girl families appear well-matched in each period on most relevant characteristics including rural residence, religion, caste, wealth, and father's education. More than 70 percent of the births occur in Hindu families and close to 70 percent in rural areas. However, over time, births preceded by a firstborn girl are increasingly born to less educated mothers relative to births preceded by a firstborn boy. We investigate this more formally in exploring heterogeneity in fertility responses to ultrasound by education of the mother and, as we shall see, our results confirm that literate women exhibit greater reductions in son-biased fertility stopping than illiterate women, consistent with the summary statistics. In the main regressions, we consistently control for various markers of SES and, importantly, specifications with mother fixed effects are free from any potential bias introduced by differences in SES between firstborn-boy and firstborn-girl families.

Although the NFHS contain fairly rich data on investments in children, these questions are asked only for children born in a few years preceding each survey and there is no pre-ultrasound data on investments.[25] For this reason, we utilize the 1999 round of the Rural Economic and Demographic Survey (REDS). REDS is restricted to rural women from 16 major states and reports age at death for deceased children but does not record the year of birth for children who did not survive to the date of survey. To the extent that deceased children are likely to have received lower health investments than surviving children and (as we show) excess girl mortality

[25] NFHS-1, 2, and 3 collected health investments for, respectively, the last three children born after January 1988, the last two children born after January 1995, and all children born after January 2001.

was higher during the pre-ultrasound period than in the post-ultrasound years, the exclusion of deceased children will tend to bias the estimated effects downward. In this sense, the estimates we present are conservative. We nevertheless also present estimates using the NFHS data—which is not biased by the exclusion of deceased children—by exploiting increasing penetration of ultrasound during post-ultrasound years.

Table 3.A.12 reports summary statistics for the REDS sample. The definition of pre-ultrasound and early diffusion period is the same as in Table 3.A.11 but the late diffusion period is shorter (1995-1999). Vaccination rates have increased over time for all children from about 70 percent probability of receiving at least one vaccine in the pre-ultrasound period to above 90 percent likelihood in the late diffusion period. In the pre-ultrasound period, children preceded by a firstborn boy were 3 p.p. more likely to be immunized than children preceded by a firstborn girl and this gap was closed post-ultrasound. Breastfeeding is nearly universal for girls and boys in India. The mean duration of breastfeeding was about 19 months in the pre-ultrasound period and nearly 90 percent of boys and girls (who survived to the survey date) were breastfed for at least 12 months. The gender gaps manifest in terms of breastfeeding duration, after age one, though they are not apparent in the sample means in Table 3.A.12, possibly because the mean breastfeeding duration is mechanically lower in the 1995-1999 period because some of the children in that group were of breastfeeding age at the time of survey. This is another weakness of the REDS that we redress by showing results estimated from the NFHS data as well.

3.5 Results

3.5.1 Excess Female Mortality

In Table 3.A.2, we present estimates of the impacts of the introduction and diffusion of ultrasound technology on neonatal and post-neonatal child mortality.[26] We add controls as we move across columns, with column (4) being the richest specification. The coefficient of *Firstborn girl * Female* confirms that, during the pre-ultrasound period, girls were significantly more likely to die neonatally (by 0.9 p.p.) and during early childhood (by 1.5 p.p.) among children preceded by a firstborn sister relative to a firstborn brother.[27] The triple-interaction coefficients, *Firstborn girl*

[26]Table 3.A.13 reports estimates for infant and child mortality.

[27]The probability of post-neonatal child mortality for boys in the pre-ultrasound period was 6.2 percent in firstborn-boy families and 5.7 percent in firstborn-girl families.

* *Female* * *Post*, indicate a 33 percent decline in the neonatal EFM gap between firstborn-girl and firstborn-boy families, relative to the 1.765 p.p. baseline gap. For post-neonatal child mortality, the EFM gap between firstborn-girl and firstborn-boy families reduced by 2.167 p.p. once ultrasound technology became available, which is very large relative to the baseline gap. These results are robust to the inclusion of mother fixed effects.

We can decisively reject an increase in relative girl neonatal mortality, which could arise if, having detected child sex, parents made smaller fetal investments in girls. Larger reductions in post-neonatal than in neonatal mortality are consistent with the increases in postnatal investments that we document below, given that neonatal survival is less dependent upon postnatal investments and more closely linked to maternal health and delivery conditions. Our findings are also congruous with previous work; for instance, Almond et al. (2006) show that hospital de-segregation after the Civil Rights Act led to a narrowing of the racial gap in post-neonatal but not neonatal mortality.[28] The significance of our results is enhanced by the fact that reductions in post-neonatal mortality also improve later life circumstances, predicting adult height, a marker of health (Bozzoli et al. (2009)) and cognitive performance (Chay et al. (2009)).

Estimates by birth order are in Table 3.A.14. The coefficients indicate reductions in post-neonatal mortality among girls across birth order, but the largest and only statistically significant coefficients are among second births. This may reflect the common finding that parents are particularly averse to having more than two girls (which applies at order three and above),[29] so there remains a girl-boy differential in post-neonatal mortality at higher orders.

We also estimate specification (1) for a sample of *first births*, re-defining $Post_t^1$ and $Post_t^2$ as indicators for the second child being born during the early and late ultrasound diffusion periods, respectively (see Table 3.A.15). If, in violation of our assumption, ultrasound access also led to sex-selection for first births, we would also expect to see a decline in EFM for first births. The results confirm no pre-ultrasound difference in the risk of death by gender, and no post-ultrasound change. This validates our assumption that parents do not manipulate the sex of first births.

[28] Similarly, The Million Deaths Study (2010) in India shows that only 3.2 percent of neonatal deaths were caused by diarrhea—a function of clean water and nutrition—in contrast to 22.2 percent of post-neonatal deaths.

[29] Almond and Edlund (2008) show that Indian, Chinese, and Korean families with no previous sons exhibit male-biased sex ratios at third parity but not before in the 2000 US Census and Bhalotra and Cochrane (2010)) show that the male-bias in the sex ratio at birth in India is increasing in birth order.

Changes in Preferences versus Technology

We utilize women's self-reported preferences for fertility (i.e., ideal number of children) and sons (i.e., ideal sex ratio of children) to test if trends in other factors, such as women's education, led to changes in preferences, which in turn influenced sex differences in child investments and mortality. Our main specification is robust to this source of bias as our identification strategy teases out structural breaks in the outcomes that coincide with sharp changes in the availability of ultrasound technology. We nevertheless assess the potential role of changes in preferences as follows. First, we include them as controls in specification (1) and, to strengthen this test, we interact the preference variables with the $Female$ dummy and the $Firstborn$ $girl$ dummy. The coefficients of interest are not changed by these additions (Table 3.A.16). We then control for another measure of gender biased preferences—the state-year gender enrollment ratio at ages 6-11 and 11-14 and, again, the results are stable. The coefficients of these additional terms are interesting in their own right, and are consistent with our expectations. The coefficients of *Ideal Sex Ratio * Female* and *Enrollment Gender Gap * Female* are positive, implying that EFM is higher in families with a stronger preference of sons. The coefficient on *Ideal Fertility * Female* is negative, implying that, conditional on son preference, a declining trend in fertility leads to higher EFM, a finding that is consistent with Jayachandran (2016).

Second, we check if, post-ultrasound, girls are more likely to be born into families with weaker son preference, as this would then be a reason that they do better. This pattern would arise if families that commit female feticide have stronger son preference than families that take their girl conceptions to term. We focus on the behavior of firstborn-girl families because the data show that these are the families that abort girls (Figure 3.A.3). We test for this channel by estimating specification (1) with mother's self-reported ideal fertility and ideal proportion of sons as dependent variables (Table 3.A.17). The triple-interaction coefficients indicate that female children born post-ultrasound in a firstborn-girl family relative to a firstborn-boy family are *no more likely* post-ultrasound than pre-ultrasound to be born to mothers with lower son preference. This suggests that feticide depends more upon characteristics like wealth and education (this is demonstrated in Bhalotra and Cochrane (2010), who also show that these characteristics are negatively correlated with desired fertility) than on stated son preference. Indeed, stated son preference is lower in wealthy and educated families even though they commit more female feticide; a vivid portrayal of this is in Figures 3.A.5 and 3.A.6.[30]

[30]The finding that girls in India are increasingly being born in (Bhalotra and Cochrane (2010))

Exploiting Variation in Self-reported Ultrasound Use

Thus far, we have chosen not to use the available individual data on ultrasound use because uptake is potentially endogenous and the data, being self-reported, may be inaccurate. However, we now use these data to corroborate the previous results which are based on changes in the availability of ultrasound driven by different import and licensing regimes.[31] We average individual-level data to obtain the percentage of births in a state and a year for which the mother reports having a prenatal ultrasound scan at some point during pregnancy ($Ultra_{st}$).[32] The earliest year for which this variable can be defined is 1996 so now, instead of comparing pre- and post-ultrasound births, we compare births in states with varying intensity of ultrasound use. The results in Table 3.A.18 suggest that ultrasound scan usage is associated with decline in EFM, confirming the results in Table 3.A.2.

Other Robustness Checks

In the main analysis, we incorporate pre-ultrasound cohorts in the sample and control for underlying trends in the outcomes. We nevertheless confirm that our findings are not being driven by any pre-trends by restricting the sample to the pre-ultrasound period and then re-estimating specification (1) with a single *Post* indicator, one each for "placebo" (or fake) treatment years 1977 through 1982 (Table 3.A.19). We find no evidence of an underlying convergence in mortality outcomes for boys and girls that is unrelated to ultrasound availability.

Earlier we referenced previous research which verifies that the sex of the first-born child was not manipulated upon the introduction of ultrasound (Bhalotra and Cochrane (2010)). We nevertheless investigate this differently, restricting the sample to women whose first child was born in the pre-ultrasound period. We then re-estimate equation (1) (Table 3.A.20) and although there are fewer observations

and surviving in (our results, discussed in detail shortly) low-SES families aligns with the predictions of Edlund (1999). However, if the Trivers-Willards mechanism, which accounts for parents caring not only about the sex of their offspring but also about the chances that their offspring reproduce, were a key driver for this pattern, then we may expect this to express in stated son preference. Our finding that high-SES parents report lower son preference even though they engage in stronger son-preferring behaviors can be explained either by (i) high-SES parents being more socialized into under-stating their true son preference or (ii) sex-selective behaviors being driven by economic and social factors (lower desired fertility or greater affordability of safe ultrasound and abortion facilities among high-SES women) rather than by the Trivers-Willards mechanism, even if the latter is present.

[31] Notably, self-reported ultrasound use is larger among mothers with a firstborn girl relative to a firstborn boy, consistent with the former being more likely to practice sex-selection at higher parities.

[32] The denominator equals the number of births with a non-missing response on the question pertaining to ultrasound use during pregnancy.

(especially for the late diffusion period), the coefficients of interest remain negative and statistically significant.

3.5.2 Postnatal Health Investments

The declining post-neonatal mortality among girls signals increased parental investments - we directly test for this in Table 3.A.3 with REDS data. The outcomes are the number of months a child is breastfed, indicators for breastfeeding duration being at least 12, 24, and 36 months;[33] a dummy variable indicating that the child has received at least one vaccine; and medical expenditure (in Rupees) on the child in the year prior to the survey.[34] Pre-ultrasound, in families with a firstborn girl, boys were breastfed for a longer duration and were more likely to be vaccinated and to receive expenditure during illness. The estimates show that the gender gaps in breastfeeding and vaccination were virtually eliminated post-ultrasound in firstborn-girl families.[35] Previous literature estimates that breastfeeding differences explain about 9 percent of the gender gap in post-neonatal child mortality in India (Jayachandran and Kuziemko (2011)) and that sex differences in vaccinations explain between 20 - 30 percent (Oster (2009)). We estimate that the contributions of breastfeeding and vaccination to the ultrasound-led decline in EFM are, respectively, 25 percent and 9 to 13 percent (details in Appendix 2).

We also present results using the NFHS data in Table 3.A.21. The NFHS contains information not only on immunization and breastfeeding, but also reports the number of antenatal checks during pregnancy. The results are broadly similar.[36] Estimates by birth order (available upon request) are also congruous with the birth

[33] Here, the sample is restricted to children who are at least 12, 24, and 36 months old, respectively.

[34] Medical expenditure is conditional upon illness and includes doctor's fees, medicines, and costs of special diets during the illness. The specifications are similar to those for mortality except that we drop the urban indicator since REDS covers only rural households, the wealth quintiles (not reported in REDS) and, since we have a smaller sample, we drop ρ_{bt}, $\xi_b G_j$, and $\nu_s G_j$ and replace ϕ_{st} with state-specific linear time trends.

[35] There is a significant improvement in the total duration of breastfeeding and in breastfeeding during the second year of life. Since most Indian children are breastfed through the first year of life, it makes sense that gender gaps in breastfeeding duration emerge after age one. The coefficients on breastfeeding for at least 36 months and for medical expenditure during sickness are positive but imprecise.

[36] The estimates of gender gaps during the early diffusion period inTable 3.A.21 are similar to the corresponding numbers (i.e., sums of the coefficients in the first two rows) in Table 3.A.3. The coefficients in the second row of Table 3.A.21 capture the differences in gender gaps during the two post-ultrasound periods, and are also similar to the corresponding estimates (i.e., differences of the coefficients in the second and third rows) in Table 3.A.3. For instance, column (2) of Table 3.A.21 implies that the gender gap in the likelihood of receiving at least one vaccine was 2.5 p.p. lower in the late diffusion period relative to the early diffusion period, while the corresponding magnitude is 3.7 p.p. in column (6) of panel A in Table 3.A.3.

order specific results for survival.

Consistent with there being no pre-ultrasound EFM for first births and no post-ultrasound change, we find no significant pre-ultrasound gender gaps in breastfeeding and immunization rates among first births and no differences in these outcomes for first births whose younger siblings were born before versus after ultrasound access. To the extent that the bulk of immunization and breastfeeding investments in the first child are made before the second birth, these findings are not surprising.[37] However, we find that pre-ultrasound, firstborn girls received lower medical and educational expenditure than firstborn boys and that these gaps narrowed after ultrasound became available (see Table 3.A.22).[38] Overall, these result are consistent with sex-selection driven fertility decline in firstborn-girl families raising per capita resources that benefit all children.

3.5.3 Fertility

For reasons detailed earlier, in Tables 3.A.4 and 3.A.5 we investigate if son-biased fertility stopping behavior changed subsequent to the availability of sex-selection technology. The coefficient of *Firstborn Girl* is positive and significant in both tables confirming that, pre-ultrasound, women whose first child was a girl were more likely to give birth in a given year and had 0.155 more births than women with a firstborn son. Our estimates show that these differentials were eliminated once ultrasound technology became available.[39] Controlling for mother's fertility preference and son preference does not significantly alter these effects, which suggests that the estimated coefficients reflect changes in behavior in response to changes in technology rather than changes in preferences.[40]

Since sex-selective abortion allows parents to avoid unwanted children, we also test whether availability of ultrasound drove actual fertility closer to desired fertility. Our estimates in Table 3.A.5 imply a significant drop of undesired fertility of 0.093 births, compared to the baseline difference at 0.117 in first-girl relative to first-boy

[37] However, breastfeeding acts as a natural contraceptive, and consequently girls relative to boys at lower birth orders are breastfed for shorter period so that parents can start trying to have a son (Jayachandran and Kuziemko (2011)); but the ability to sex-select is useful only post-conception, so this channel is unlikely to get significantly affected by ultrasound access.

[38] Pre-ultrasound, even among firstborns, girls were 12 p.p. less likely to be provided allopathic treatment if they were ill. Similarly, annual education and medical expenditure on firstborn girls was lower by, respectively, Rs. 562 and Rs. 69 relative to firstborn boys.

[39] Estimates conditional upon mother fixed effects are not statistically significantly different, and are available on request.

[40] As a robustness check, like for EFM, we also ran a specification where $Firstborn\ girl * Post1$ and $Firstborn\ girl * Post2$ are interacted with the preference variables; the results remain the same.

families. To put this in a wider perspective, the presence of HIV reduces the average number of births a woman has during her life-cycle by 0.15 (Shapira (2013)). The coefficients of interest (those on the triple-interaction terms) are similar for actual fertility (in column 2) and for actual minus desired fertility (in column 3). This confirms that the decline in actual fertility in first-girl families that we document is not driven by a decline in desired fertility.

The coefficients of the stated preference terms reveal that actual fertility is, as we may expect, increasing in desired fertility and in the desired ratio of sons to daughters. However, excess fertility is decreasing in the ideal ratio of sons to daughters, suggesting that desired fertility rises more steeply with son preference than actual fertility.[41] Since actual fertility is not fully in the control of parents, it will tend to rise less than proportionately with desired fertility, given that the desire to have sons leads to the widely documented phenomenon of son-biased fertility stopping (which is evident from the baseline statistics in the first row).

Table 3.A.23 shows that the relative fertility decline in firstborn-girl families is driven by a shift from having four or more children toward having two or three children.

3.5.4 Heterogeneity

Although stated son preference is weaker among urban, educated, and wealthy women (Figure 3.A.6), they exhibit higher rates of prenatal sex selection (Figure 3.A.5). This is consistent with their lower reported desired fertility (Figure 3.A.7), with educated individuals being more likely to adopt a new technology (Lleras-Muney and Lichtenberg (2005)), and with their being more efficacious in achieving their targets (Rosenzweig and Schultz (1989)). Also, wealth may matter at the margin for affordability of ultrasound scans and (safe) abortion, especially if a woman engages in multiple events. If there were a strict substitution of prenatal for post-natal girl mortality, we may expect the reductions in mortality and fertility that we document in this paper to be concentrated among educated and wealthy mothers. However, it is possible that these responses occur at different margins.

We examine if our results differ by mother's educational attainment (illiterate versus literate), household wealth (bottom 40 versus top 20 percent), mother's employment status (paid employment versus rest),[42] household caste (scheduled caste

[41] Regressions of actual fertility and ideal fertility on the ideal sex ratio variable confirm this. However, note that unobserved shocks that increase, say the (measured) desired number of boys, would drive up both the ideal sex ratio and ideal fertility creating a positive upward bias on the latter coefficient.

[42] The results are robust to using alternative comparisons, including employed (paid or unpaid)

(SC) versus other),[43] and rural versus urban residence. Tables 3.A.6 and 3.A.7 respectively present estimates for post-neonatal child EFM (using the mother fixed-effects specification) and gender gaps in sibling size.[44] In each regression, we continue to control for all SES variables, except the one being used to examine heterogeneity.

The tables show the baseline gender gaps in firstborn-girl families (first rows) and test to what extent they narrowed during the post-ultrasound period (second and third rows). In the pre-ultrasound era, among firstborn-girl families, excess girl post-neonatal mortality was greater in low SES groups (illiterate, poor, unemployed, rural), but not significantly different for high versus low caste families. However, baseline gender gaps in sibling size were larger among literate, rich, urban, and non-SC women.[45]

Post-ultrasound EFM decline was steeper in poor and rural households, and households in which women were in unpaid employment.[46] This is consistent with larger increases in investment but also with a given change in investment having larger survival impacts in low-SES households where other causes of child mortality, such as infection rates, are higher. However, the absolute decline in EFM was broadly similar among literate versus illiterate mothers (and greater in relative terms in the literate group given their lower baseline rates). The decline in son-biased fertility stopping is greater among high-caste, literate, rich, urban women and women in unpaid employment.

We also split each SES-subsample by caste.[47] Unlike other dimensions of SES, caste is exogenous in that an individual is born into a caste and remains in it. The caste hierarchy has been preserved by the low prevalence of inter-caste marriages.[48] The upper-castes have historically laid greater emphasis on ritual purity and adherence to religious texts, which often compromises the position of women (Das Gupta et al. (2003), Das Gupta (2010)). In accordance with this, pre-ultrasound excesses in mortality and family size in girl-led families and post-ultrasound declines were, on

versus unemployed.

[43] We pool high castes and other backward classes (OBC) because the first survey round does not distinguish them. Since OBC are better-off than SC households, the categories we use preserve the caste hierarchy in India. We also pool scheduled tribes (ST) with the higher caste group based upon finding that they take similar coefficients when included as a separate category.

[44] Table 3.A.24 presents the heterogeneity results for effects on the number of children.

[45] This is not the case for women's employment but, in India, on average, low-SES women are more likely to be employed, driven to work by poverty.

[46] Women's labor force participation is hockey-stick shaped in India, being most common among the poor (Das and Desai (2003).

[47] We include households of all religions and use the self-reported caste of the household for our analysis while using religion as a control variable. These results are available upon request.

[48] According to the 2005 India Human Development Survey, only 4.4 percent of women were married to a spouse from a different caste.

average, larger in higher caste households. Interacting caste with other indicators of SES shows that at the low-end of the SES distribution, low castes are more gender-equal but at the high-end, low caste behavior is similar to that of high-castes. This is consistent with the process of *Sanskritization*, wherein lower castes emulate the upper castes in seeking upward mobility (Srinivas (1962)).

3.6 Estimates of Substitution

To assess the extent of substitution between postnatal and prenatal discrimination due to ultrasound technology, we use our estimates to compute the number of female child deaths that have been averted and compare them with the number of girls who are missing, both due to prenatal sex-detection. These calculations are described in Appendix 3. We find that for every girl that survived due to ultrasound technology, five girls were aborted before birth.[49]

We calculate the proportion of discriminated births for which parents substituted postnatal discrimination with prenatal discrimination as the decline in the number of girls missing due to EFM (= -60,879) divided by the total number of missing girls in the pre-ultrasound period (= 196,667). This calculation implies that, for nearly 31 percent of the births, parents who were practicing postnatal discrimination switched to prenatal discrimination. This is much larger than the estimates in Lin et al. (2014) who find that 4 percent of parents of second-parity births and 8 percent of parents of third- and higher-parity births made the switch in Taiwan.

3.7 Conclusion

As ultrasound technology became increasingly available, the global annual number of sex-selective abortions increased from nearly zero in the late 1970s to 1.6 million per year in 2005-2010 (Bongaarts and Guilmoto (2015)), with India and China being the biggest contributors. The stark growth in female feticide has garnered a lot of attention from academics, policymakers, and popular media. While most public attention has focused upon the increasing deficit of girl children, it has also been noted that a large share of sex-selective abortions in India are conducted in unsafe environments. Complications due to unsafe abortion account for an estimated 9

[49]These estimates take into account the endogenous changes in fertility, ignoring which the substitution is 5.5 rather than 5. Lower average fertility implies that the share of all births that are lower parity is increasing and since sex ratios are closer to the biological norm at lower parities (and consistent with the norm for first births), this will contribute (*through a compositional effect*) to the average sex ratio being less male-biased than otherwise.

percent of all maternal deaths in India (Stillman et al. (2014)).[50] Moral arguments can be made both in favor of parents' right to choose the sex of their offspring as well as against selective abortion of girls (Kumar (1983)). Abstracting from these ethical dilemmas, there are several reasons why a significantly male-biased sex ratio at birth is undesirable. The resulting scarcity of women on the marriage market can substantially increase the number of unmarried and childless men,[51] who may face destitution in old age since children through marriage are the most important source of support for the elderly in countries like India that lack institutional social security (Das Gupta et al. (2010)). Rising sex ratios can lead to increased trafficking of women,[52] higher prevalence of sexually-transmitted diseases (Ebenstein and Sharygin (2009)), and more crime (Edlund et al. (2007), Drèze and Khera (2000)). Sex-selection may also result in girls being consistently born to lower-status parents, thereby relegating women to lower social strata (Edlund (1999), Bhalotra and Cochrane (2010)). On the other hand, a shortage of women on the marriage market may increase their bargaining power and welfare.[53] It has also been argued that sex-selective abortions might be preferable to infanticide or postnatal discrimination (Goodkind (1996)).

Our analysis shows that the increase in sex-selective abortions fueled by ultrasound technology substantially decreased postnatal gender discrimination against girl children in India. Relative to available studies, we contribute new evidence and present a more comprehensive analysis. We find that sex-selection eliminated gender gaps in post-neonatal child mortality, postnatal health investments, and sibling size among second- and higher-parity births in households with a firstborn daughter relative to households with a firstborn son. Although fewer girls were born, those that survived to birth were treated more equally and were more likely to survive to age five. On account of higher investments in for instance breastfeeding and immunization, we can project that they are more likely to do well as adults. However, our evidence suggests that surviving girls in the post-ultrasound regime are more likely to be in low-SES households, and for every additional girl that survived after birth, five girls were aborted.

[50] The maternal mortality ratio in India was 178 maternal deaths per 100,000 live births in 2010-12.

[51] Bhaskar (2011) estimates that one in five boys born in recent cohorts in China will be unable to find female partners.

[52] Recent evidence shows that a shortage of women in north Indian states has led to the import of brides from other poorer states in India (Kaur (2004), Ahlawat (2009)).

[53] See Chiappori et al. (2002) and related papers for the large literature on household bargaining in developed countries. Stopnitzky (2012) shows that a relative scarcity of women in Haryana has increased their bargaining power on the marriage market and they are able to secure improved sanitation facilities at home as a result.

Appendix

3.A Appendix of chapter 3

3.A.1 Variable Descriptions

- Excess Female Mortality: Female mortality - Male mortality
- Neonatal mortality: Death within one month of birth
- Post-neonatal child mortality: Death after the first month of birth but before age five
- $Post_t^1$: indicator variable for $t \in 1985 - 1994$
- $Post_t^2$: indicator variable for $t \in 1995 - 2005$
- F_i: child i is female
- G_j: first child of mother j is female
- $Ultra_{st}$: percentage of births in state s and year t for which the mother reports getting an ultrasound test at some point during the pregnancy
- Allopathic treatment: an indicator variable that equals 1 if a sick child received medical help from an allopathic doctor during the past year, and 0 otherwise
- Exp on education: amount (in Rupees) spent on fees, books, uniform, hostel pocket money, transportation, and private coaching during last year
- Doctors' fees: amount (in Rupees) spent on doctors' fees last year
- Medicine and special food: amount (in Rupees) spent on medicine and special food last year
- Medical exp: amount (in Rupees) spent on doctors' fees, medicine, and special food last year

- Ideal sex ratio: Ideal number of sons/ Ideal number of daughters (as reported by the mother)

- Ideal number of children: Mother's self-reported ideal number of children of any sex

- Education categories: no education, incomplete secondary education, and secondary or higher education

- Categories for mother's birth cohort: 1942-1960, 1961-1970, and 1971-1987

- Categories for Mother's age at birth: 12-15 years, 16-18 years, 19-24 years, 25-30 years, and 31-49 years

- Caste categories: Scheduled Castes (SC), Scheduled Tribes (ST), and Others

- Religion categories: Hindus, Muslims, and Others

Appendix 2: Contributions of Breastfeeding and Vaccination

Our estimates suggest that ultrasound access reduced post-neonatal child mortality by 1.490 p.p. (column (4) in panel B of Table 3.A.2) and increased the likelihood of being breastfed for at least 24 months by 27.5 p.p. (triple-interaction coefficient in column (5) of Panel B in Table 3.A.3). Since there was no significant effect on breastfeeding during the first year of birth, we assume that the 27.5 p.p. increase took place between 12 and 24 months from birth. According to the World Health Organization (2000), breastfeeding between the ages one and two decreases mortality by 50 percent relative to no breastfeeding. Applying this factor to the share of children who are being breastfed and the mortality rate during 12-24 months,[54] the implied mortality rate for breastfed children is 1.35 percent[55] and is 2.7 percent (1.35 * 2 = 2.7) for non-breastfed children in the 12-24 months range. This implies that not being breastfed during the 12-24 months age range increases the risk of mortality by 1.35 p.p. (2.7 - 1.35 = 1.35). If breastfeeding disparities (during 12-24 months) were the only cause of post-neonatal child mortality differences by gender, the EFM decline due to improvements in the breastfeeding gender gap would be

[54]We assume that the share of children who are being breastfed and the mortality rate during 12-24 months are the same as those for 12-36 months used in Jayachandran and Kuziemko (2011).

[55]Solving $0.481x + 2(1 - 0.481)x = 2.05$ yields $x = 1.35$, where 0.481 is the fraction of children aged 12 to 36 months that are being breastfed in the sample analyzed by Jayachandran and Kuziemko (2011).

0.16 p.p. (0.275 * 1.35 = 0.371). Thus breastfeeding explains about 25 percent (0.371/1.490 = 0.249) of the estimated EFM decline.

Moreover, we find that ultrasound availability increased the probability of a child receiving at least one vaccination by 0.075 to 0.112 (triple-interaction coefficient in column (6) of Panel A in Table 3.A.3). The average number of vaccinations (conditional on receiving at least one vaccination) for girls preceded by a firstborn girl during the early diffusion period is 6.64 in NFHS data. Thus the estimated effects for at least one vaccination translate into an average increase in the number of vaccinations of 0.498 (6.64 * 0.075 = 0.498) to 0.744 (6.64 * 0.112 = 0.744). Oster (2009) suggests that each vaccination reduces mortality during ages 1 to 4 by 0.26 p.p.. Thus the implied effect on EFM through vaccination is 0.129 p.p. (0.26 * 0.498 = 0.129) to 0.193 p.p. (0.26 * 0.744 = 0.193), which translates into 8.7 percent (0.129/1.490 = 0.087) to 13 percent (0.193/1.490 = 0.130) of the decline in EFM.

Note that the mortality measure is 12-36 months in Jayachandran and Kuziemko (2011) and is 1-4 years in Oster (2009). However, any exogenous change in these mortality measures would generate an almost one-to-one change in the mortality measure we use, i.e., death during 1 month to 5 years of birth.

Appendix 3: Substitution of Prenatal for Postnatal Girl Mortality

Here we use our regression estimates to calculate the magnitude of substitution from postnatal EFM to prenatal sex-selection as the ratio of the number of girls selectively aborted and the number of girls who survived due to ultrasound (and would have otherwise died postnatally). Let:

- N: annual number of births in India
- N_{FG}: annual number of births in India that are preceded by a firstborn girl
- M_{FG}: fraction of births that are male among N_{FG}
- F_{FG}: fraction of births that are female among N_{FG}
- Δ: the pre-post (counterfactual) difference

Then, the number of "missing girls" each year, i.e., the difference between the expected number of female births (given the observed number of male births and the natural sex ratio at birth) and the observed number of female births is given by:

$$\underbrace{\frac{0.49}{0.51}(N_{FG} * M_{FG})}_{\text{Expected \#female births given the observed \#male births}} - \underbrace{N_{FG} * F_{FG}}_{\text{Observed \#female births}} = N_{FG}(\frac{0.49}{0.51} - \frac{F_{FG}}{0.51})$$

The change in the annual number of missing girls due to ultrasound access is then calculated as the difference between the number of missing girls during the pre- and the post-ultrasound periods:

$$\begin{aligned}
&N_{FG,post}(\frac{0.49}{0.51} - \frac{F_{FG,post}}{0.51}) - N_{FG,pre}(\frac{0.49}{0.51} - \frac{F_{FG,pre}}{0.51}) \\
=&N_{FG,post} * \frac{0.49}{0.51} - N_{FG,post} * \frac{F_{FG,post}}{0.51} - N_{FG,pre} * \frac{0.49}{0.51} + N_{FG,pre} * \frac{F_{FG,pre}}{0.51} \\
=&\Delta N_{FG} * \frac{0.49}{0.51} - N_{FG,post} * \frac{F_{FG,post}}{0.51} + N_{FG,pre} * \frac{F_{FG,pre}}{0.51}
\end{aligned}$$

(Adding and subtracting $N_{FG,post} * \frac{F_{FG,pre}}{0.51}$)

$$\begin{aligned}
=&\Delta N_{FG} * \frac{0.49}{0.51} - N_{FG,post} * \frac{F_{FG,post}}{0.51} + N_{FG,post} * \frac{F_{FG,pre}}{0.51} - N_{FG,post} * \frac{F_{FG,pre}}{0.51} + N_{FG,pre} * \frac{F_{FG,pre}}{0.51} \\
=&\Delta N_{FG} * \frac{0.49}{0.51} + N_{FG,post} * \Delta F_{FG} * \frac{-1}{0.51} - \Delta N_{FG} * \frac{F_{FG,pre}}{0.51} \\
=&\underbrace{\Delta N_{FG} * (\frac{0.49}{0.51} - \frac{F_{FG,pre}}{0.51})}_{\text{(conception effect)}} + \underbrace{N_{FG,post} * \Delta F_{FG} * \frac{-1}{0.51}}_{\text{(sex-selective abortions effect)}}
\end{aligned}$$

The change in the number of missing girls due to ultrasound can be decomposed into a "conception effect" and a "sex-selective abortion effect," as defined above. The sex-selective abortion effect refers to the fact that a smaller $fraction$ of *post-ultrasound* births preceded by a firstborn girl are now female, where as the conception effect is driven by the change in the number of births that are preceded by a firstborn girl.

The number of excess postnatal female deaths in a year equals:

$$N_{FG} * F_{FG} * EFM_{FG}$$

where EFM_{FG} refers to the difference between the probability of death by age 5 among children preceded by a firstborn girl and a firstborn boy.

The change in the number of excess postnatal female deaths can thus be written as:

$$
\begin{aligned}
&N_{FG,post} * F_{FG,post} * EFM_{FG,post} - N_{FG,pre} * F_{FG,pre} * EFM_{FG,pre} \\
&= \underbrace{N_{FG,post} * F_{FG,post} * (\Delta EFM_{FG})}_{\text{(behavioral effect)}} + \underbrace{N_{FG,post} * (\Delta F_{FG}) * EFM_{FG,pre}}_{\text{(mechanical sex-selection effect)}} \\
&+ \underbrace{(\Delta N_{FG}) * F_{FG,pre} * EFM_{FG,pre}}_{\text{(conception effect)}}
\end{aligned}
$$

The change in the number of excess postnatal female deaths can also be decomposed into three components: the "behavioral effect" refers to the EFM decline due to, say, better postnatal health investments in girls in the post-ultrasound period; the "mechanical sex-selection effect" is driven by the decrease in the *fraction* of *post-ultrasound* births that are preceded by a firstborn girl; and the "conception effect" that reflects the change in the *number* of births that are preceded by a firstborn girl itself.

The table below uses the aforementioned formulae to calculate the number of births preceded by a firstborn girl during the pre-ultrasound (in column (1)), during the late diffusion period (in column (2)), and the change in the number of missing girls as the difference between columns (1) and (2) in column (3). EFM_{FG} in column (1) is the coefficient of *Firstborn girl * Female* and in column (2) is the coefficient of *Firstborn girl * Female * Post2* from column (4) in panel B of Table 3.A.2. During the pre-ultrasound period, the fraction of females in births preceded by a firstborn girl was 47.9 percent. Ultrasound access decreased this fraction by 1.8 p.p..[56] We calculate $\frac{N_{FG}}{N}$ from the entire sample since the pre-ultrasound sample of births is likely to be underreported due to recall bias. The number of births in column (1), N, is the 1995 figure from UN Statistics Division. The number of mothers, m, is obtained as $\frac{N}{\text{General Fertility Rate}}$ using the general fertility rate estimated from NFHS-2 by Retherford and Mishra (2001) as 131.53 per 1000 women. ΔN_{FG} is estimated to be equal to -0.0063 as is obtained from the fertility equation using the conversion method described below.

Thus, the number of "missing girls" increased by 305,496 per year due to ultrasound access. The estimated decline in EFM, on the other hand, implies that the annual number of postnatal female deaths by age 5 fell by 60,879. The ratio of increase in sex-selective abortions and the decrease in EFM is 5.02, i.e., for every 5.02 girls aborted, one girl survived due to access to ultrasound technology.

[56]This number is derived from a regression specification similar to the one estimated by Bhalotra and Cochrane (2010). These results are reported in Table 3.A.25.

Table: Decomposition and Simulation

	Pre (Pre-ultrasound) (1)	Post (Late diffusion) (2)	Δ (Post - Pre) (3)
A. Regression Estimates			
EFM_{FG}	0.0176	0.0040	-0.0136
F_{FG}	0.4790	0.4605	-0.0185
$\frac{N_{FG}}{N}$	0.3340		
B. Decomposition			
N	27,300,000		
m (#mothers)	$\approx$ 207,557,211		
m_{FG} (#mothers with a firstborn girl)	$\approx$ 0.3340*m		
N_{FG}	0.3340*(N)	0.3340*(N)-0.0063*m_{FG}	-0.0063*m_{FG}
	= 9,118,200	= 8,681,458	= -436,742
Δ number of excess postnatal female deaths:			
(1) $N_{FG,post} * F_{FG,post} * (\Delta EFM_{FG})$		$\approx$ -54,370	
(2) $N_{FG,post} * (\Delta F_{FG}) * EFM_{FG,pre}$		$\approx$ -2,827	
(3) $(\Delta N_{FG}) * F_{FG,pre} * EFM_{FG,pre}$		$\approx$ -3,682	
(4) = (1)+(2)+(3)		= -60,879	
Δ number of missing girls:			
(5) $\Delta N_{FG} * (\frac{0.49}{0.51} - \frac{F_{FG,pre}}{0.51})$		$\approx$ -9,420	
(6) $N_{FG,post} * \Delta F_{FG} * \frac{-1}{0.51}$		$\approx$ 314,916	
(7) = (5)+(6)		= 305,496	
Δ number of missing girls / Δ EFM:			
(8) = (7)/(4)		= -5.02	

– Conversion of treatment effect from the fertility specification (2) to expected number of children born in a year.

Let N to be the number of total children born in a single year. Then

$$N = N_1 + N_{FG} + N_{FB}$$

where N_1 is the number of firstborn children and N_{FB} and N_{FG} are the number of second or higher order births respectively preceded by a firstborn girl and a firstborn boy.

- m_{FG}: number of women in the 15-49 age-group with a firstborn girl
- $S_{a,FG}$ and $S_{a,FB}$: share of women aged a who have a firstborn girl and firstborn boy, respectively
- $P_{a,FG}$ and $P_{a,FB}$: probability that a women who has a firstborn girl (firstborn boy) and who is of age a in a given year gives birth in that year

$$\implies N_{FG} = (\textstyle\sum_{a=15}^{a=49} P_{a,FG} S_{a,FG}) * m_{FG}$$

$$
\begin{aligned}
&\Delta N_{FG} = N_{FG,post} - N_{FG,pre} \\
=&(\sum_{a=15}^{a=49} P_{a,FG,post}S_{a,FG,post} - \sum_{a=15}^{a=49} P_{a,FG,pre}S_{a,FG,pre}) * m_{FG} \\
=&(\sum_{a=15}^{a=49}(P_{a,FG,post} - P_{a,FG,pre})S_{a,FG}) * m_{FG}
\end{aligned}
$$

assuming that $S_{a,FG,pre} = S_{a,FG,post} = S_{a,FG}$. This assumption is reasonable since ultrasound access had no effect on the decision to have a first birth or its sex ratio.

The treatment-on-the-treated effect we estimated in equation (2) equals

$$
\begin{aligned}
&E[N|post, FG] - E[N|pre, FG] \\
=&\sum_{a=15}^{a=49}(P_{a,FG,post} - P_{a,FG,pre}) * S_{49,FG} \\
+&\sum_{a=15}^{a=48}(P_{a,FG,post} - P_{a,FG,pre}) * S_{48,FG} \\
+&\sum_{a=15}^{a=47}(P_{a,FG,post} - P_{a,FG,pre}) * S_{47,FG} \\
+&...
\end{aligned}
$$

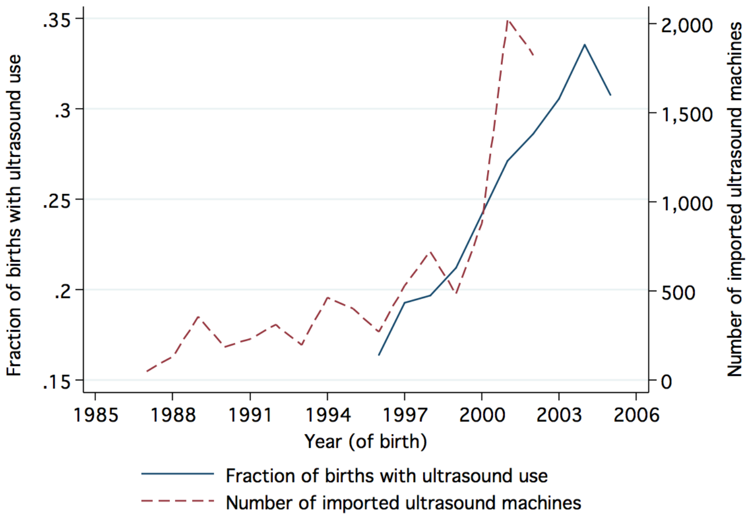
If we assume uniform impact of ultrasound access on the probability of birth across age-groups, i.e., $P_{a,FG,post} - P_{a,FG,pre} = P_{FG,post} - P_{FG,pre}$ and assuming $S_{a,FG} = S_{FG}$:

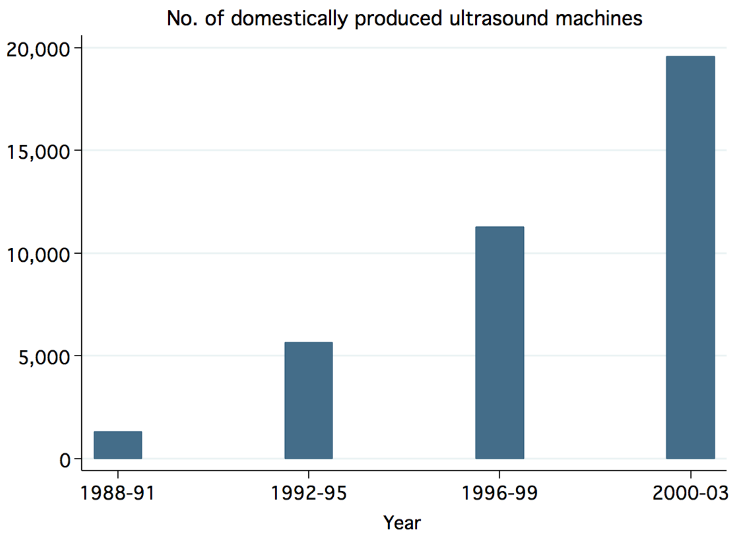
$$
E[N|post, FG] - E[N|pre, FG] = S_{FG} * (P_{FG,post} - P_{FG,pre}) * (1 + 2 + 3 + ... + 35)
$$

$$
\begin{aligned}
&\sum_{a=15}^{a=49}(P_{a,FG,post} - P_{a,FG,pre})S_{a,FG} \\
=&S_{FG} * (P_{FG,post} - P_{FG,pre}) * 35 \\
=&\frac{35}{(1 + ... + 35)} * (E[N|post, FG] - E[N|pre, FG]) \\
=&\frac{35}{630} * (E[N|post, FG] - E[N|pre, FG]) \\
\approx&0.056 * (E[N|post, FG] - E[N|pre, FG]) \\
\approx&0.056 * (-0.112) = -0.0063
\end{aligned}
$$

Figures

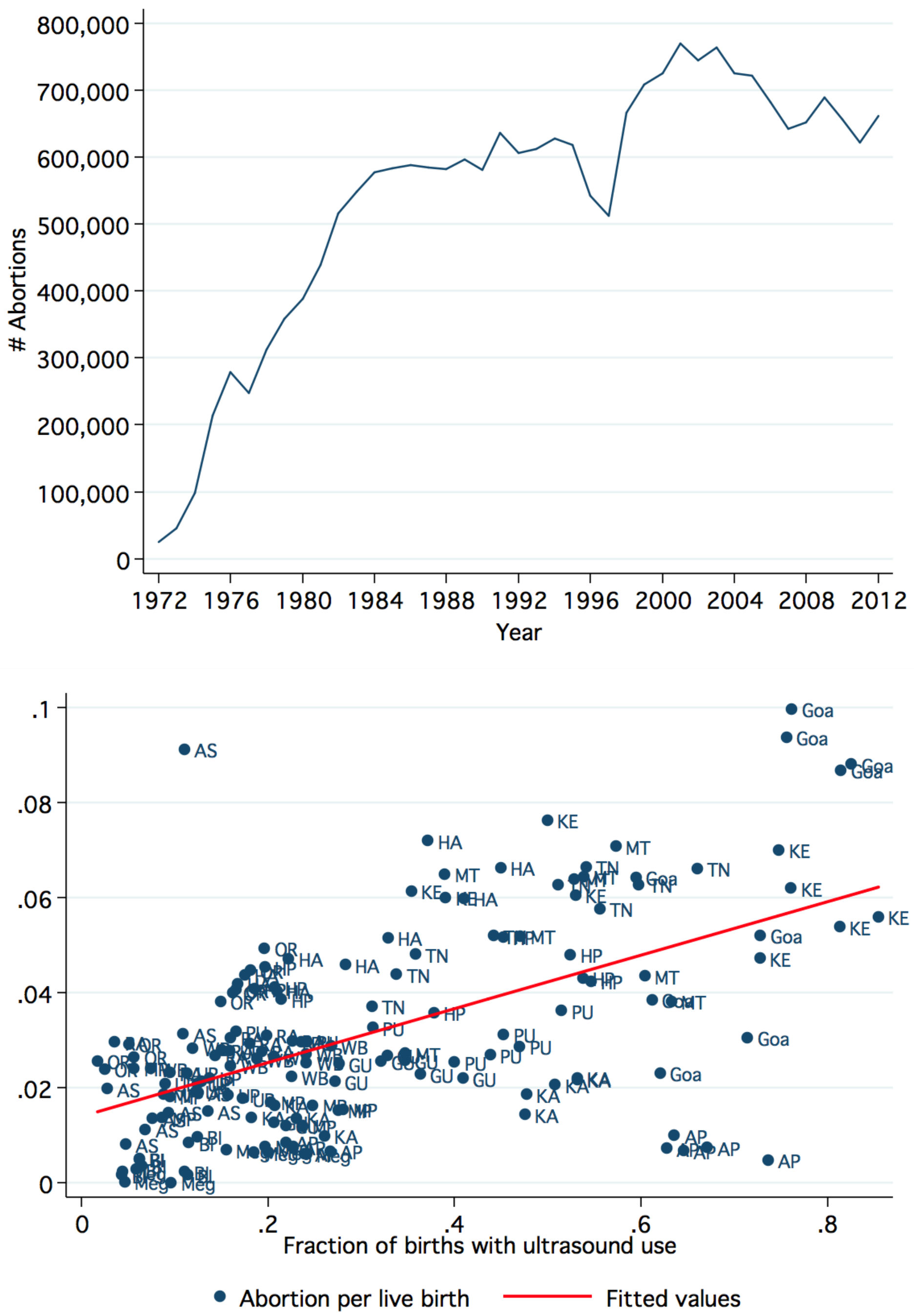
Figure 3.A.1: Ultrasound use by mothers and supply of ultrasound scanners in India





NOTES: (1) The solid line in the top graph plots the fraction of births in a year for which the mother reports getting an ultrasound test at some point during the pregnancy (the denominator equals the number of births with a non-missing response on ultrasound use). The relevant question was not asked in NFHS-1 but in NFHS-2 and NFHS-3, data on ultrasound use was collected for births since January 1995 and January 2001, respectively. The years 1995 and 2000 have been dropped due to extremely small sample sizes. (2) The dashed line in the top graph plots the number of ultrasound scanners imported at the national level, the first records of which appear in the import data in 1987; indeed, there was no category coding these scanners before then (Source: Mahal et al. (2006)). (3) The bars in the bottom graph plot the number of ultrasound machines produced domestically in India. Data source: George (2006).

Figure 3.A.2: Number of abortions and ultrasound use by mothers in India

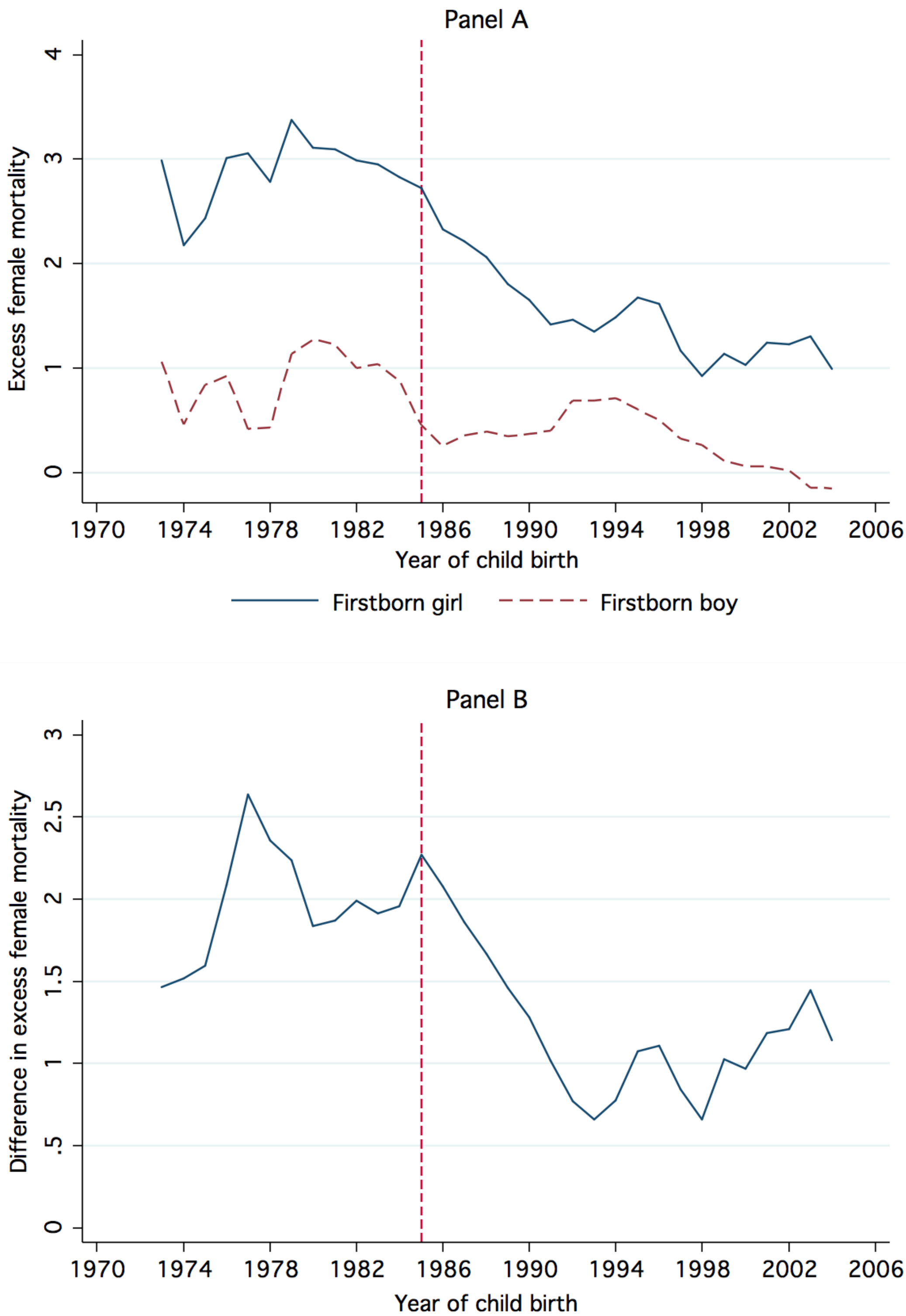


NOTES: The top graph plots the officially reported number of abortions (both sex-selective and other abortions) by state and year using data from http://www.johnstonsarchive.net/policy/abortion/india/ab-indias.html. The bottom graph plots the state-year variation in the officially reported number of abortions and the fraction of births in a state-year for which the mother reports getting an ultrasound test at some point during the pregnancy (other details in Figure 3.A.1 notes).

Figure 3.A.3: Trends in proportion of females at birth by birth order and sex composition of older siblings

NOTES: Panel A shows the evolution of percent female among first births over time. In panels B, C, and D the trend in percentage of births that are female is plotted respectively for second, third, and fourth births separately for families that have at least one son and families with no sons at the time of the respective birth. In all cases, the y-axis shows the 5-year moving average of percentage of births that are female. The figures show that, despite ultrasound availability, the sex ratio of first births has remained normal. It also shows that the sex ratio at birth in families without sons diverges from the sex ratio in families with sons after the introduction of ultrasound.

Figure 3.A.4: Post-neonatal excess female child mortality, by firstborn's sex



NOTES: Post-neonatal excess female child mortality equals the percentage of female births that die minus the percentage of male births that die after the first month of birth but before age five. The graph in Panel A plots the 5-year moving average of EFM among families with a firstborn girl and families with a firstborn boy; Panel B plots the EFM for births in firstborn girl families minus the EFM for births in firstborn boy families. The vertical line splits the years into pre- and post-ultrasound periods.

Tables

Table 3.A.1: Unadjusted mortality rates for treatment and control groups (%)

	Firstborn boy families			Firsborn girl families		
	(1) Male	(2) Female	(3) EFM (2)-(1)	(4) Male	(5) Female	(6) EFM (5)-(4)
1. Pre-ultrasound: 1973-1984						
Neonatal	7.326	5.536	-1.79	5.060	5.035	-0.025
Post-neonatal Child	6.169	7.052	0.883	5.721	8.771	3.050
N	43,833	40,425		17,252	15,908	
2. Early diffusion period: 1985-1994						
Neonatal	5.692	4.528	-1.164	4.439	4.352	-0.087
Post-neonatal Child	4.380	4.880	0.500	4.246	6.021	1.775
N	82,579	77,699		43,547	39,519	
3. Late diffusion period: 1995-2005						
Neonatal	4.434	3.627	-0.807	3.450	3.554	0.104
Post-neonatal Child	2.507	2.658	0.151	2.523	3.702	1.179
N	46,504	44,196		27,713	24,141	

NOTES: This table reports the percentage of second- and higher-order children, by firstborn sex and child gender, who suffered neonatal or post-neonatal child mortality over the three time-periods in the sample. Columns (3) and (4) report the difference between the mortality numbers in columns (1) - (2) and (4) - (5), respectively.

Table 3.A.2: Excess female mortality

A. Neonatal Mortality	(1)	(2)	(3)	(4)
*Firstborn girl * Female*	1.754***	1.705***	1.350***	0.862**
	(0.379)	(0.373)	(0.382)	(0.417)
*Firstborn girl * Female * Post1*	-0.698	-0.665	-0.681	-0.213
	(0.426)	(0.422)	(0.422)	(0.460)
*Firstborn girl * Female * Post2*	-0.872**	-0.875**	-0.895**	-0.369
	(0.353)	(0.348)	(0.357)	(0.429)
N	503,316			
Baseline mean	1.765			
B. Post-Neonatal Child Mortality	(1)	(2)	(3)	(4)
*Firstborn girl * Female*	2.171***	2.123***	1.758***	1.476***
	(0.333)	(0.332)	(0.313)	(0.383)
*Firstborn girl * Female * Post1*	-0.919**	-0.872**	-0.954**	-0.773
	(0.402)	(0.401)	(0.404)	(0.474)
*Firstborn girl * Female * Post2*	-1.191***	-1.229***	-1.355***	-1.490**
	(0.397)	(0.380)	(0.362)	(0.598)
N	478,843			
Baseline mean	2.167			
*Female * Post1* and *Female * Post2*	x	x		
*Firstborn Girl * Post1* and *Firstborn Girl * Post2*	x	x		
X_{ijt}		x	x	x
Firstborn Girl x Birth year FE			x	x
Female x Birth year FE			x	x
Female x State FE			x	x
Female x Birth order FE			x	x
Birth order x Birth year FE			x	x
Birth order x State FE			x	x
State x Birth year FE			x	x
Firstborn girl x State FE			x	x
Firstborn girl x Birth order FE			x	x
Mother FE				x

NOTES: This table reports estimates of specification (1). Each column within a panel is a separate regression. The outcomes are mortality rates, as % of births that do not survive. We always control for *Female* and fixed effects (FE) for birth year and birth order and for *Firstborn girl* and state FE in columns (1)-(3). The vector X_{ijt} comprises mother's age at birth and, except in column (4), household wealth quintiles, caste, religion, residence in a rural area, educational attainment of child's parents, and mother's birth cohort. Standard errors in parentheses are clustered by state. Baseline mean refers to the pre-ultrasound difference between EFM in firstborn-girl and firstborn-boy families. *** 1%, ** 5%, * 10%.

Table 3.A.3: Postnatal health investments

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Panel A:	Months breastfed			Received $\geq$ one vaccine			Rupees spent on illness last year		
*First girl*Female*	-5.124**	-4.451**	-3.851**	-0.130**	-0.091**	-0.067	-27.150	-15.583	-87.599
	(1.879)	(1.846)	(1.747)	(0.051)	(0.038)	(0.048)	(50.961)	(50.910)	(56.788)
*First girl*Female*Post1*	4.739**	4.040*	2.396	0.112**	0.096*	0.075*	17.108	6.186	70.589
	(2.012)	(2.042)	(2.068)	(0.049)	(0.048)	(0.043)	(68.645)	(65.100)	(80.108)
*First girl*Female*Post2*	4.162	3.735	1.536	0.094*	0.068*	0.112**	44.955	31.177	106.444
	(3.140)	(2.594)	(2.499)	(0.051)	(0.033)	(0.041)	(50.283)	(51.044)	(64.065)
N		13,084			20,561			15,156	
Baseline mean		0.17			-0.035			-111.58	
Panel B:	Breastfed $\geq$ 12 months			Breastfed $\geq$ 24 months			Breastfed $\geq$ 36 months		
*First girl*Female*	-0.096	-0.083	-0.027	-0.269***	-0.269***	-0.226**	-0.103**	-0.083**	-0.075
	(0.071)	(0.090)	(0.084)	(0.090)	(0.079)	(0.079)	(0.044)	(0.038)	(0.044)
*First girl*Female*Post1*	0.085	0.092	-0.0002	0.321***	0.275**	0.185	0.087	0.071	0.029
	(0.110)	(0.087)	(0.080)	(0.104)	(0.114)	(0.120)	(0.069)	(0.065)	(0.076)
*First girl*Female*Post2*	0.096	0.047	-0.046	0.178	0.222	0.141	0.053	0.033	0.042
	(0.101)	(0.094)	(0.099)	(0.190)	(0.161)	(0.143)	(0.082)	(0.073)	(0.068)
N		12,447			11,760			10,942	
Baseline mean		0.006			0.003			0.008	
*Female*Post1* & *Female*Post2*	x	x		x	x		x	x	
*1st Girl*Post1* & *1st Girl*Post2*	x	x		x	x		x	x	
X_{ijt}		x	x		x	x		x	x
1st Girl x Birth year FE			x			x			x
Female x Birth year FE			x			x			x
Female x State FE			x			x			x
Female x Birth order FE			x			x			x
Birth order x State FE			x			x			x
State-specific time trends			x			x			x
Firstborn girl x State FE			x			x			x
Firstborn girl x Birth order FE			x			x			x

NOTES: This table reports the coefficients from specification (1) estimated for investments in children using the REDS data. Each column within a panel is from a separate regression. We control for *Female*, *Firstborn girl*, and fixed effects for birth year, birth order, and state in all columns. The vector X_{ijt} comprises mother's age at birth, caste, religion, educational attainment of child's parents, and mother's birth cohort. Breastfeeding results are based on the last two surviving births of a mother. Vaccination and health expenditure results are based on all surviving children of a mother. Standard errors in parentheses are clustered by state. Baseline mean refers to the pre-ultrasound difference between the gender gaps in investments in firstborn-girl and firstborn-boy families. *** 1%, ** 5%, * 10%. Further investment results using the NFHS data are in Table 3.A.21.

Table 3.A.4: Fertility: Hazard of Birth

	(1)	(2)	(3)	(4)
Firstborn girl	0.173***	0.185***	0.155***	0.168***
	(0.011)	(0.012)	(0.038)	(0.038)
*Firstborn girl * Post1*	-0.073***	-0.084***	-0.092***	-0.102***
	(0.006)	(0.007)	(0.009)	(0.010)
*Firstborn girl * Post2*	-0.190***	-0.211***	-0.220***	-0.233***
	(0.016)	(0.018)	(0.017)	(0.017)
Ideal sex ratio				0.076***
				(0.022)
Ideal no. of children				0.161***
				(0.010)
N	2,455,630	2,455,553	2,455,553	2,276,192
Baseline mean	0.285			
Year FE	x	x	x	x
State FE	x	x	x	x
Age FE	x	x	x	x
Parity FE	x	x	x	x
State x Year FE	x	x	x	x
Years since last birth FE		x	x	x
Firstborn girl x Year FE			x	x

NOTES: Estimates of specification (2). The dependent variable is an indicator for birth in a given year. Sample includes all mothers who have ever given birth, for each year from their year of marriage to the year of interview. Standard errors in parentheses are clustered by state. Baseline mean is the mean probability of birth in a given year during the pre-ultrasound period. *** 1%, ** 5%, * 10%.

Table 3.A.5: Fertility: Number of Children

	Number of births		Excess Fertility
	(1)	(2)	(3)
Firstborn girl	0.155***	0.141***	0.117***
	(0.012)	(0.015)	(0.018)
*Firstborn girl * Post1*	-0.088***	-0.079***	-0.085***
	(0.016)	(0.019)	(0.024)
*Firstborn girl * Post2*	-0.112***	-0.100***	-0.093***
	(0.018)	(0.023)	(0.025)
Ideal no. of children		0.315***	
		(0.018)	
Ideal sex ratio		0.052***	-0.345***
		(0.018)	(0.024)
N	118,663	88,475	88,475
Baseline mean	3.001	3.001	0.451

NOTES: Estimates of specification (3). The dependent variable in columns (1)-(2) is the number of births at the time of interview and in column (3) is excess fertility which equals number of births minus ideal number of children. Baseline means are average of the outcome variable in each column for mothers who had both their first and last birth within 1972-1984. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 3.A.6: Heterogeneity by socioeconomic status in effects on post-neonatal child mortality

	Mother's Education		**Household Wealth**		**Mother's Employment**	
	Illiterate	Literate	Bottom 40%	Top 20%	Paid employment = 0	Paid employment = 1
	(1)	(2)	(3)	(4)	(5)	(6)
*Firstborn girl * Female*	1.901***	0.961*	1.846**	0.655	1.737***	0.569
	(0.626)	(0.538)	(0.854)	(0.635)	(0.374)	(1.023)
*Firstborn girl * Female * Post1*	-0.726	-1.057	-0.113	-1.056*	-1.136*	0.441
	(0.714)	(0.627)	(1.093)	(0.614)	(0.579)	(1.322)
*Firstborn girl * Female * Post2*	-1.579*	-1.315*	-1.918**	-0.391	-1.653**	-0.752
	(0.862)	(0.649)	(0.922)	(0.640)	(0.594)	(1.164)
N	262,539	216,304	194,658	96,819	340,073	138,426
Baseline mean	8.94	3.44	9.93	2.86	6.46	7.63
	SC	General/OBC/ST	Rural	Urban		
	(1)	(2)	(3)	(4)		
*Firstborn girl * Female*	2.114	1.326***	2.108***	-0.042		
	(1.445)	(0.400)	(0.473)	(0.476)		
*Firstborn girl * Female * Post1*	-0.732	-0.779	-1.021	-0.186		
	(1.533)	(0.493)	(0.633)	(0.635)		
*Firstborn girl * Female * Post2*	-2.001	-1.391**	-2.333***	0.470		
	(1.944)	(0.646)	(0.635)	(0.671)		
N	77,471	401,372	322,945	155,898		
Baseline mean	9.32	6.36	7.92	4.19		

NOTES: Estimates of specification (1) with mother fixed effects for different sub-samples. Each column within a panel is a separate regression. SC, ST, OBC, and General respectively denote scheduled castes, scheduled tribes, other backward classes, and upper castes. The wealth categories are based on the national household wealth distribution. Standard errors in parentheses are clustered by state. Baseline mean refers to post-neonatal child mortality for children born in pre-ultrasound period 1972-1984, by mothers with the specific socioeconomic status. *** 1%, ** 5%, * 10%.

Table 3.A.7: Heterogeneity by socioeconomic status in effects on the hazard of birth

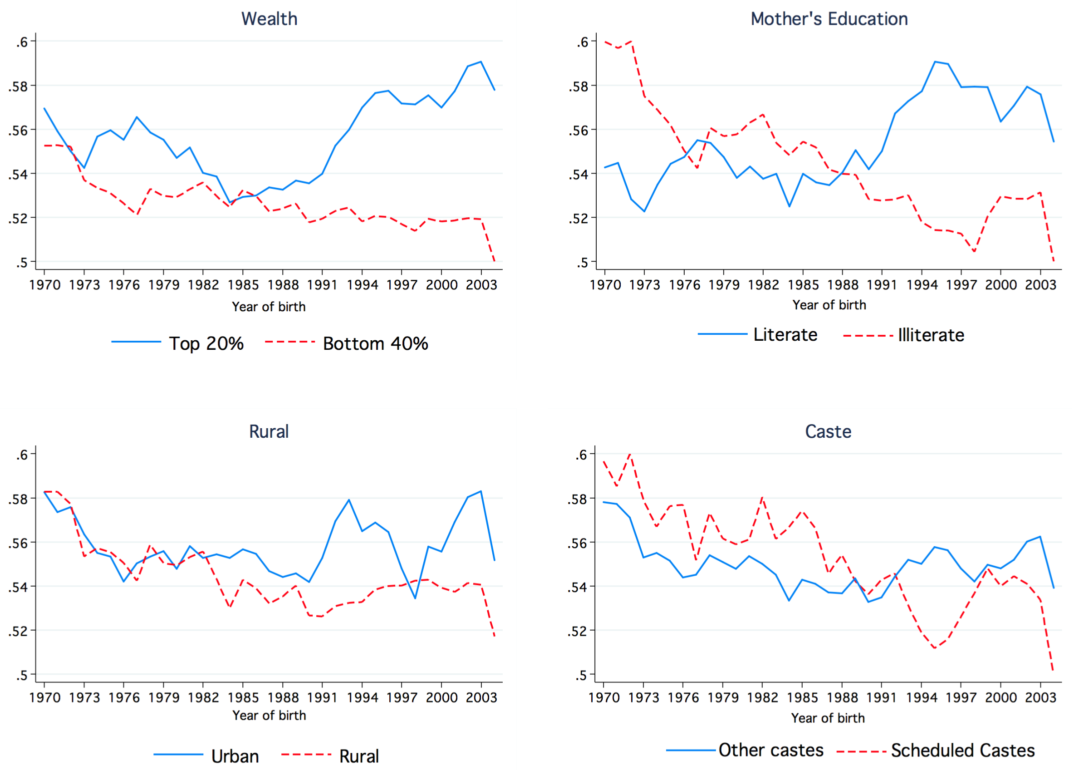
	Mother's Education		**Wealth**		**Mother's Employment**	
	Illiterate (1)	Literate (2)	Bottom 40% (3)	Top 20% (4)	Paid employment = 0 (5)	Paid employment = 1 (6)
Firstborn girl	0.060**	0.324***	0.091**	0.388***	0.177***	0.140***
	(0.029)	(0.037)	(0.043)	(0.036)	(0.046)	(0.033)
*Firstborn girl * Post1*	-0.059***	-0.133***	-0.093***	-0.114***	-0.114***	-0.062***
	(0.009)	(0.015)	(0.006)	(0.021)	(0.012)	(0.013)
*Firstborn girl * Post2*	-0.127***	-0.312***	-0.172***	-0.317***	-0.253***	-0.174***
	(0.016)	(0.017)	(0.021)	(0.020)	(0.022)	(0.020)
N	1,351,731	1,102,413	971,697	599,531	1,494,545	958,583
Baseline mean	0.289	0.279	0.288	0.263	0.287	0.283

	Caste		**Rurality**	
	SC (1)	General/OBC/ST (2)	Rural (3)	Urban (4)
Firstborn girl	0.078	0.172***	0.092**	0.310***
	(0.048)	(0.037)	(0.040)	(0.038)
*Firstborn girl * Post1*	-0.040**	-0.101***	-0.085***	-0.112***
	(0.020)	(0.009)	(0.010)	(0.017)
*Firstborn girl * Post2*	-0.160***	-0.234***	-0.184***	-0.296***
	(0.038)	(0.016)	(0.020)	(0.023)
N	374,215	2,080,707	1,598,030	856,144
Baseline mean	0.295	0.284	0.289	0.278

NOTES: Estimates of specification (3). The dependent variable is an indicator for birth in a given year. Sample includes all mothers who have ever given birth, for each year from their year of marriage to the year of interview. Standard errors in parentheses are clustered by state. Baseline mean is the mean probability of birth in a given year during the pre-ultrasound period. *** 1%, ** 5%, * 10%.

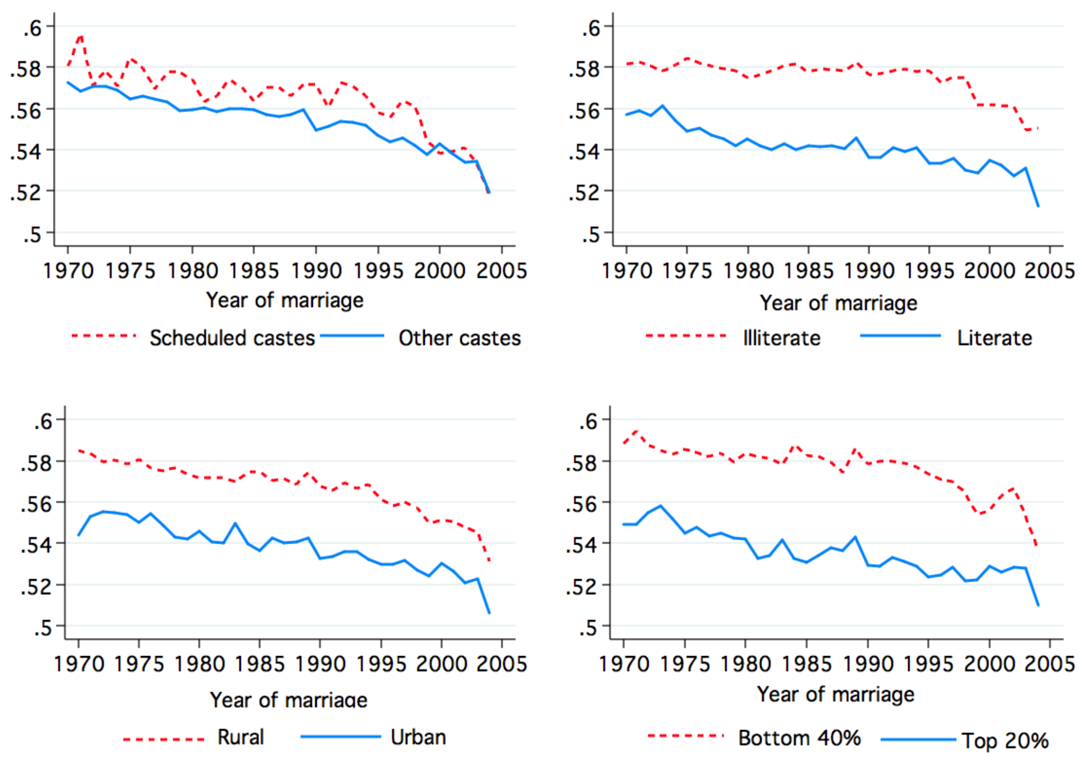
3.A.2 Additional Figures and Tables

Figure 3.A.5: Trends in Actual Fraction of Sons



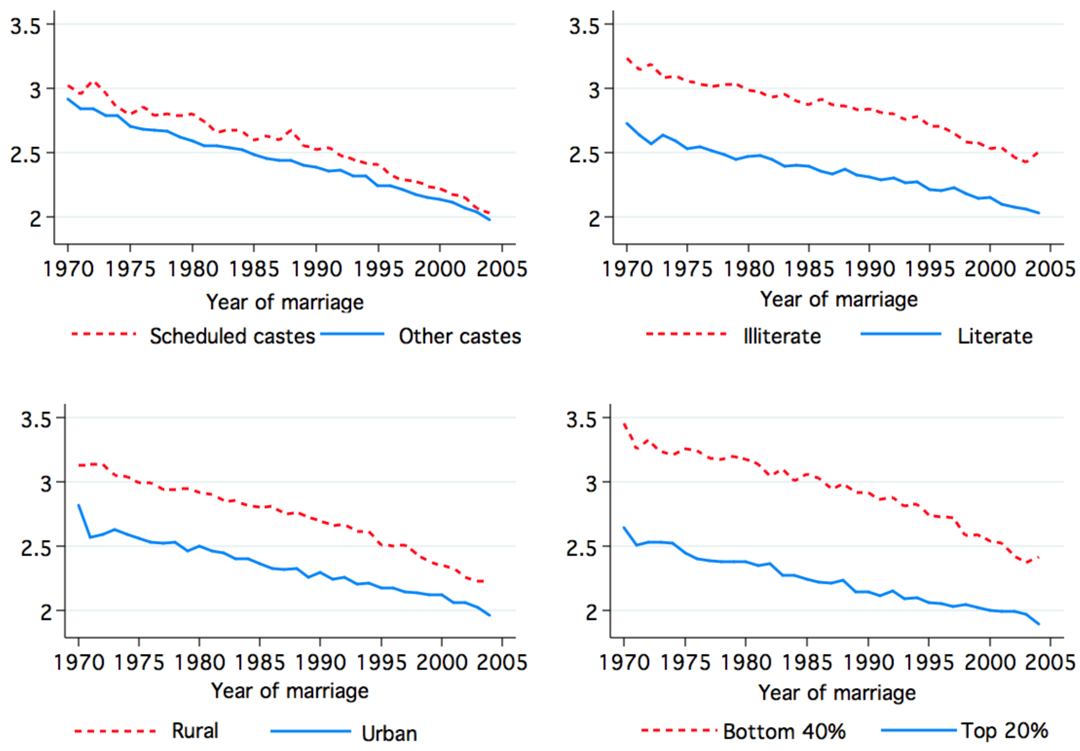
NOTES: The graph plots the trend in the 5-year moving average of actual fraction of male births among all births in a year, separately for various SES groups.

Figure 3.A.6: Trends in Ideal Fraction of Sons



NOTES: The above graphs plot the trend in the 5-year moving average of the reported ideal fraction of sons by year of first marriage for mothers who had at least one child at the time of survey. Ideal fraction of sons = $\frac{ideal_{boys}+(0.5*ideal_{either})}{ideal_{kids}}$ if $ideal_{kids} > 0$, where $ideal_{boys}$ is the ideal number of boys, $ideal_{either}$ is the ideal number of children of either sex, and $ideal_{kids}$ is the ideal number of total children, as reported by a woman.

Figure 3.A.7: Trends in Ideal Number of Children



NOTES: The above graphs plot the trend in average reported ideal number of children by the year of first marriage for mothers who had at least one child at the time of survey.

Table 3.A.8: Sample means: Mortality rates by period

	(1) Male	(2) Female	(3) EFM
1. Pre-ultrasound: 1973-1984			
Neonatal	6.686	5.395	-1.291
Post-neonatal Child	6.040	7.539	1.499
N	61,085	56,333	
2. Post-ultrasound: 1985-1994			
Neonatal	5.259	4.469	-0.790
Post-neonatal Child	4.333	5.265	0.932
N	126,126	117,218	
3. Post-ultrasound: 1995-2005			
Neonatal	4.066	3.601	-0.465
Post-neonatal Child	2.513	3.027	0.514
N	74,217	68,337	

NOTES: This table reports the percentage of second- and higher-order children, by child's gender, who suffered from neonatal, post-neonatal infant, or post-neonatal child mortality over the three time-periods in our sample. Column (3) reports the difference between the mortality numbers in columns (1) and (2).

Table 3.A.9: Age distribution of missing girls in 2000

Age group	No. missing (in 000s) India (1)	China (2)	(3)	% of all missing women India (4)	China (5)	(6)	% of under-5 missing girls India	China
At birth	184	644		11%	37%		37%	83%
0-1	146	109		9%	6%		30%	14%
1-4	164	23		10%	1%		33%	3%
< 5	494	776		29%	45%			
5-14	93	2		5%	0%			
≥ 15	1125	947		66%	55%			
Total	1712	1727						

NOTES: This table is based on Anderson and Ray (2010) and reports the number of missing girls for various age groups (columns (1) and (2), missing girls as a percentage of all missing women (columns (3) and (4)), and missing girls across age groups as a percentage of all under-5 missing girls (columns (5) and (6)), separately for India and China.

Table 3.A.10: Sample means: Gaps between actual and ideal fertility and fraction of sons

	(1) Actual	(2) Ideal	(3) Actual - Ideal
1. Pre-ultrasound: 1973-1984			
Fertility	2.96	2.21	0.75
Fraction of Sons	0.58	0.57	0.01
N			
2. Post-ultrasound: 1985-1994			
Fertility	2.16	2.17	-0.01
Fraction of Sons	0.55	0.56	-0.01
N			
3. Post-ultrasound: 1995-2005			
Fertility	1.91	1.93	-0.03
Fraction of Sons	0.53	0.55	-0.02
N			

NOTES: This table reports the percentage of second- and higher-order children, by child's gender, who suffered from neonatal, post-neonatal infant, or post-neonatal child mortality over the three time-periods in our sample. Column (3) reports the difference between the mortality numbers in columns (1) and (2).

Table 3.A.11: Sample means, all variables

	1973-1984		1985-1994		1995-2005	
	First boy	First girl	First boy	First girl	First boy	First girl
	(1)	(2)	(3)	(4)	(5)	(6)
Number of children	2.91	3.12	2.13	2.22	1.85	1.96
N (Mothers)	14,095	10,690	25,857	22,169	23,480	22,372
N (Births)	84,258	33,160	160,278	83,066	90,700	51,854
Female	0.48	0.48	0.48	0.48	0.49	0.47
Rural	0.69	0.70	0.68	0.71	0.63	0.67
Hindu	0.77	0.77	0.75	0.75	0.71	0.71
Muslim	0.11	0.12	0.13	0.14	0.15	0.16
SC	0.14	0.14	0.16	0.17	0.18	0.19
ST	0.12	0.13	0.14	0.14	0.16	0.16
Mother birth cohort:						
1942-1960	0.60	0.73	0.11	0.19	0.01	0.01
1961-1970	0.40	0.27	0.62	0.67	0.17	0.28
1971-1987	0.00	0.00	0.28	0.14	0.82	0.71
Mother's age at birth:						
12-15	0.09	0.01	0.06	0.01	0.03	0.00
16-18	0.26	0.12	0.20	0.08	0.15	0.06
19-24	0.49	0.56	0.49	0.49	0.49	0.44
25-30	0.15	0.27	0.20	0.33	0.25	0.37
31-49	0.02	0.04	0.05	0.09	0.08	0.13
Mother's education:						
No education	0.61	0.63	0.56	0.62	0.43	0.53
Incomplete secondary	0.34	0.34	0.36	0.33	0.43	0.38
Secondary or higher	0.05	0.03	0.08	0.06	0.14	0.08
Father's education:						
No education	0.33	0.34	0.31	0.33	0.25	0.30
Incomplete secondary	0.51	0.52	0.51	0.50	0.56	0.55
Secondary or higher	0.17	0.14	0.19	0.16	0.19	0.15
Household wealth:						
2nd quintile	0.15	0.14	0.17	0.17	0.18	0.21
3rd quintile	0.18	0.19	0.17	0.18	0.18	0.19
4th quintile	0.21	0.22	0.21	0.21	0.21	0.20
Richest quintile	0.22	0.21	0.20	0.17	0.21	0.16
Child can read and write	0.80	0.79	0.81	0.79	0.73	0.71
Child still in school	0.52	0.62	0.77	0.79	0.70	0.69
N	51,484	20,531	67,794	32,749	26,379	14,475
Full immunization			0.39	0.35	0.51	0.45
N			19,154	11,398	31,127	18,130
Number of months breastfed			13.04	13.34	13.64	14.10
N			19,871	12,057	32,324	19,228
Number of antenatal checks			2.62	2.11	3.75	3.07
N			23,571	13,865	34,197	20,892

NOTES: SC and ST denote Scheduled Castes and Scheduled Tribes. Number of children refers to fertility at the time of survey. Data are NFHS, three rounds.

Table 3.A.12: Sample means, postnatal health investments

Variable	1973-1984 First boy	1973-1984 First girl	1985-1994 First boy	1985-1994 First girl	1995-1999 First boy	1995-1999 First girl
Female	0.44	0.47	0.46	0.47	0.47	0.45
SC	0.13	0.14	0.14	0.16	0.15	0.16
ST	0.07	0.06	0.08	0.07	0.09	0.08
OBC	0.35	0.39	0.35	0.34	0.35	0.34
Hindu	0.88	0.90	0.87	0.91	0.89	0.90
Muslim	0.08	0.06	0.09	0.06	0.07	0.07
Sikh	0.03	0.03	0.03	0.02	0.03	0.02
Christian	0.01	0.00	0.01	0.00	0.01	0.01
Mother's age at birth	24.12	26.31	23.62	26.03	23.72	26.21
Mother is literate	1.52	1.49	1.47	1.47	1.36	1.43
Father's years of education	5.23	5.54	5.83	5.71	6.93	6.15
At least one vaccine	0.73	0.70	0.88	0.89	0.93	0.94
N (Births)	5,788	1,962	6,440	2,744	2,960	1,268
Breastfed	0.99	0.99	0.99	0.99	0.77	0.79
No. of months breastfed	18.88	19.07	18.34	18.32	12.66	13.39
N (Births)	2,674	1,021	3,969	1,846	2,765	1,209
Breastfed at least 12 months	0.90	0.90	0.88	0.89	0.71	0.75
N (Births)	2,674	1,021	3,969	1,846	2,332	976
Breastfed at least 24 months	0.40	0.40	0.36	0.35	0.26	0.29
N (Births)	2,674	1,021	3,969	1,846	1,820	780
Breastfed at least 36 months	0.07	0.08	0.07	0.05	0.03	0.05
N (Births)	2,674	1,021	3,969	1,846	1,217	527
Expense on illness (last year)	160.78	227.01	168.16	152.01	216.64	184.66
N (Births)	192	643	6,253	2,665	2,877	1,240

NOTES: Sample from 1999 REDS data. SC, ST, and OBC denote Scheduled Castes, Scheduled Tribes, and Other Backward Classes. The sample is restricted to children who were alive at the time of survey. Breastfeeding information is available only for a woman's last two children who were alive at the time of survey. Sample is restricted to children at least 12, 24, and 36 months old while calculating the proportion of children who were breastfed for 12, 24, and 36 months, respectively. Expenditure on illness is in Rupees.

Table 3.A.13: Further Results: Different age-exposures for childhood mortality

	Neonatal (1)	Infant (2)	Post-neo Child (3)	Child (4)
*Firstborn girl * Female*	1.350***	2.067***	1.758***	2.893***
	(0.382)	(0.479)	(0.313)	(0.481)
*Firstborn girl * Female * Post1*	-0.681	-0.984*	-0.954**	-1.490**
	(0.422)	(0.536)	(0.404)	(0.550)
*Firstborn girl * Female * Post2*	-0.895**	-1.175***	-1.355***	-2.059***
	(0.357)	(0.408)	(0.362)	(0.468)
N	503,316	503,316	478,843	503,316

NOTES: Estimates of specification (1) for different measures of mortality. The explanatory variables are the same as those in Column (3) of Table 3.A.2. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 3.A.14: Heterogeneity in effects on post-neonatal child mortality by birth order

Birth order	2nd (1)	3rd (2)	4th (3)	5th (4)	6th (5)	7th (6)
*Firstborn girl * Female*	2.198***	1.578**	1.384	1.104	0.843	7.630
	(0.551)	(0.741)	(1.068)	(2.213)	(3.206)	(10.26)
*Firstborn girl * Female * Post1*	-2.242***	-0.251	-0.141	2.004	-0.903	-7.292
	(0.788)	(0.782)	(1.218)	(2.617)	(3.395)	(11.94)
*Firstborn girl * Female * Post2*	-1.585**	-1.458	-0.393	-0.927	-0.912	-10.69
	(0.651)	(0.994)	(1.301)	(2.759)	(3.327)	(10.35)
N	133,635	86,551	48,491	25,363	12,408	5,648
Baseline mean	6.97	7.10	7.97	9.43	9.15	13.00

NOTES: Estimates corresponding to the specification in column (3) of Table 3.A.2 estimated separately for various birth orders. Each column is a separate regression. The outcome measureS mortality as % of births that do not survive. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 3.A.15: Post-neonatal child mortality: Estimates for first births

Dependent Variable: Post-neonatal child mortality	
Female	-0.404
	(0.855)
*Female * Post1*	0.0264
	(0.615)
*Female * Post2*	-0.204
	(0.731)
N	158,194

NOTES: This table reports the coefficients from a specification similar to (1). The sample is restricted to first births. *Post*1 and *Post*2 indicate that the second birth took place during early and later diffusion period, respectively. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 3.A.16: Placebo test: Ideal household structure and gender gaps in school enrollment

	Neonatal (1)	Post-neo Child (2)	Neonatal (3)	Post-neo Child (4)
*Firstborn girl * Female*	1.051*	0.521	1.261**	1.926***
	(0.560)	(0.926)	(0.455)	(0.366)
Firstborn girl Female * Post1*	-0.193	-0.685	-0.516	-0.945*
	(0.397)	(0.518)	(0.521)	(0.493)
*Firstborn girl * Female * Post2*	-0.474	-1.192***	-0.656	-1.643***
	(0.336)	(0.426)	(0.394)	(0.544)
*Ideal number of children * Female*	-0.465***	-0.525***		
	(0.095)	(0.151)		
*Ideal Sex Ratio * Female*	1.222***	1.382***		
	(0.112)	(0.267)		
*Ideal number of children * Female * Firstborn girl*	0.084	0.410**		
	(0.186)	(0.190)		
*Ideal Sex Ratio * Female * Firstborn girl*	0.043	0.009		
	(0.268)	(0.396)		
*Girl-boy enrollment rate 6-11 * Female*			0.163***	0.460***
			(0.0520)	(0.0734)
*Girl-boy enrollment rate 14-17 * Female*			3.285***	0.225
			(0.912)	(0.710)
N	400,035	380,605	364,989	345,078

NOTES: Estimates of specification (1) with additional controls for son preference. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 3.A.17: Mother's fertility preference and son preference

	Ideal No. of Children (1)	Ideal Fraction of Sons (2)
*Firstborn girl * Female*	-0.026	0.011***
	(0.017)	(0.003)
*Firstborn girl * Female * Post1*	0.013	-0.002
	(0.017)	(0.002)
*Firstborn girl * Female * Post2*	0.028*	0.001
	(0.015)	(0.003)
N	473,195	471,559

NOTES: Estimates of specification (1) with dependent variables changed to stated preferences. Ideal fraction of sons = $\frac{ideal_{boys}+(0.5*ideal_{either})}{ideal_{kids}}$ if $ideal_{kids} > 0$, where $ideal_{boys}$ is the ideal number of boys, $ideal_{either}$ is the ideal number of children of either sex, and $ideal_{kids}$ is the ideal number of total children, as reported by a woman. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 3.A.18: Using variation in self-reported ultrasound use

	Neonatal (1)	Post-neonatal Infant (2)	Post-neo natal Child (3)
*Female * Firstborn girl*	0.930**	0.734**	0.760*
	(0.335)	(0.282)	(0.376)
*Female * Firstborn girl * $Ultra_{st}$*	-1.302	-1.219	-1.006
	(0.864)	(0.855)	(1.031)
N	113,170	108,837	108,837

NOTES: Estimates of a specification similar to (1) for childhood mortality rates, except that $Post_t^1$ and $Post_t^2$ have been replaced with $Ultra_{st}$. So, rather than use the timing of imports and production of ultrasound scanners, we use self-reported use of ultrasound scans in the NFHS data. Each column is from a separate regression. The sample is restricted to years after 1995. *** 1%, ** 5%, * 10%.

Table 3.A.19: Placebo test: Fake treatment years in the pre-ultrasound era

Dependent variable: Post-neonatal Child Mortality						
Treatment Year	1977 (1)	1978 (2)	1979 (3)	1980 (4)	1981 (5)	1982 (6)
*Firstborn girl * Female*	2.070** (0.954)	2.004** (0.836)	2.125*** (0.559)	2.334*** (0.504)	2.460*** (0.490)	2.087*** (0.405)
*Firstborn girl * Female * Post*	0.016 (0.967)	0.094 (0.898)	-0.053 (0.659)	-0.357 (0.521)	-0.642 (0.587)	-0.007 (0.438)
N			110,295			

NOTES: This table presents coefficients from a specification similar to (1), except that a single *Post* indicator is used. The treatment year used to define *Post* varies across columns. Each column represents a different regression. The sample is restricted to the 1973-1984 period (i.e., pre-Ultrasound period). Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 3.A.20: Testing whether first births were subject to sex selective abortion

	Neonatal (1)	Post-neo Infant (2)	Post-neo Child (3)
*Firstborn girl * Female*	1.182*** (0.381)	0.730** (0.301)	1.790*** (0.400)
*Firstborn girl * Female * Post1*	-0.969* (0.550)	0.0238 (0.350)	-0.485 (0.517)
*Firstborn girl * Female * Post2*	-1.270 (1.550)	-1.338 (0.942)	-2.139* (1.124)
N	196,374	185,894	185,894

NOTES: This table reports the coefficients from specification (1) for the sample of women whose first child was born before 1985. The explanatory variables are the same as those in Column (3) of Table 3.A.2. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 3.A.21: Parental health investments: Prenatal and postnatal

	# Antenatal Checks	At least 1 vaccination	Full immunization	# Months breastfed
	(1)	(2)	(3)	(4)
*Firsborn girl * Female*	-0.041	-0.029**	-0.032***	-0.422**
	(0.056)	(0.012)	(0.011)	(0.155)
*Firsborn girl * Female * Post2*	0.002	0.025**	0.017	0.129
	(0.060)	(0.012)	(0.017)	(0.258)
N	92,525	79,809	79,809	83,480

NOTES: NFHS data. Investments are only queried for cohorts born in a few years before each survey so there are no pre-ultrasound cohorts. The comparison here is therefore across the two post-ultrasound periods. The estimates are from the following specification for child i of birth order b born to mother j in year t: $I_{ibjt} = \alpha + \beta G_j * F_i * Post_t^2 + \gamma G_j * F_i + \omega_t G_j + \sigma_t F_i + \psi_b F_i + X_{ijt}^{'} \tau + \delta_s F_i + \rho_{bt} + \eta_{bs} + \phi_{st} + \epsilon_{ibjt}$. Among children who were at least 12 months old at the time of survey, we define a child to be fully immunized if he or she had received the eight most common vaccines by that time. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 3.A.22: Parental investments: Estimates for first births

	Allopathic treatment	Exp on education	Doctors' fees	Medicine and special food	Medical exp
	(1)	(2)	(4)	(5)	(3)
$Female$	-0.122**	-562.307*	-18.251**	-50.795*	-69.045*
	(0.043)	(319.263)	(8.138)	(26.306)	(32.630)
$Female * Post1$	0.140***	317.710	24.433**	62.443**	86.876**
	(0.037)	(329.164)	(9.610)	(26.887)	(33.986)
$Female * Post2$	0.187***	615.970*	16.469	57.778**	74.248*
	(0.057)	(333.976)	(10.737)	(27.141)	(35.897)
N	3,018	3,541	3,541	3,541	3,541

NOTES: This table reports the coefficients from a specification similar to (1). The sample is restricted to first births. $Post1$ and $Post2$ indicate that the second birth took place during early and later diffusion period, respectively. Standard errors in parentheses are clustered by state. The dependent variables are defined in Appendix A. *** 1%, ** 5%, * 10%.

Table 3.A.23: Fertility: Investigating the margin of response by number of children

	Number of children			
	$= 1$	$= 2$	$= 3$	≥ 4
	(1)	(2)	(3)	(4)
Firstborn girl	0.012**	-0.067***	-0.015*	0.069***
	(0.005)	(0.006)	(0.007)	(0.006)
*Firstborn girl * post1*	-0.010	0.021**	0.030***	-0.042***
	(0.007)	(0.008)	(0.009)	(0.007)
*Firstborn girl * post2*	-0.017**	0.040***	0.035***	-0.058***
	(0.007)	(0.009)	(0.010)	(0.009)
N	118663	118663	118663	118663
Baseline mean	0.114	0.271	0.300	0.315

NOTES: This table presents estimates from specification (3) using indicators for the mother having, respectively, 1, 2, 3, and ≥ 4 children at the time of survey in columns (1)-(4). Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

Table 3.A.24: Heterogeneity by socioeconomic status in effects on fertility, cross-sectional approach

	Mother's Education		**Wealth**		**Mother's Employment**	
	Illiterate (1)	Literate (2)	Bottom 40% (3)	Top 20% (4)	Paid employment = 0 (5)	Paid employment = 1 (6)
Firstborn girl	0.113***	0.198***	0.126***	0.183***	0.172***	0.142***
	(0.020)	(0.018)	(0.031)	(0.019)	(0.014)	(0.021)
*Firstborn girl * Post1*	-0.063*	-0.110***	-0.060	-0.099***	-0.114***	-0.054*
	(0.031)	(0.023)	(0.042)	(0.029)	(0.018)	(0.030)
*Firstborn girl * Post2*	-0.059**	-0.148***	-0.069*	-0.121***	-0.132***	-0.088***
	(0.028)	(0.024)	(0.036)	(0.027)	(0.021)	(0.029)
N	46,597	72,066	36,831	36,352	78,717	39,840
Baseline mean	3.230	2.812	3.101	2.769	3.026	2.964

	Caste		**Rurality**	
	SC (1)	General/OBC/ST (2)	Rural (3)	Urban (4)
Firstborn girl	0.146***	0.154***	0.134***	0.190***
	(0.039)	(0.012)	(0.017)	(0.015)
*Firstborn girl * Post1*	-0.014	-0.098***	-0.095***	-0.079***
	(0.056)	(0.016)	(0.023)	(0.023)
*Firstborn girl * Post2*	-0.032	-0.124***	-0.108***	-0.122***
	(0.041)	(0.018)	(0.026)	(0.018)
N	17,043	101,620	71,155	47,508
Baseline mean	3.243	2.966	3.109	2.850

NOTES: Estimates of specification (3). The dependent variable is the number of births at the time of interview. Standard errors in parentheses are clustered by state. Baseline mean refers to the mean fertility of mother at time of interview, with the specific socioeconomic status who had their first and last birth in the pre-ultrasound period 1972-1984. *** 1%, ** 5%, * 10%.

Table 3.A.25: Impacts of ultrasound availability on the sex ratio at birth

Dependent Variable: Child is female	
Firstborn girl	-0.002
	(0.004)
*Firstborn girl * Post1*	-0.008*
	(0.005)
*Firstborn girl * Post2*	-0.018***
	(0.006)
N	503,316

NOTES: This table reports the coefficients from the following specification: $G_{ibjt} = \alpha + \beta_1 F_i + \beta_2 F_i * Post_t^1 + \beta_3 F_i * Post_t^2 + X_{ijt}^{'}\tau + \Pi'\gamma + \epsilon_{ibjt}$, where G_{ibjt} indicates that child i of birth order b born to mother j in year t is female; F_i indicates that the firstborn is a girl; $X_i'\tau$ is a vector of household characteristics and $\Pi'\gamma$ is a vector of fixed effects that are analogous to equation (1), except that all interactions with F_i are omitted. Standard errors in parentheses are clustered by state. *** 1%, ** 5%, * 10%.

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