Why does the sex ratio at birth rise? Evidence from Vietnam*

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Abstract

Vietnam's male-to-female sex ratio at birth jumped from 107 to 112 between 2001 and 2014. Using the 2001 US-Vietnam Bilateral Trade Agreement as an exogenous shock, we investigate the link between Vietnam's market access and the rise in the sex ratio at birth. Based on a difference-in-difference framework, we find that mothers exposed to 10.6 percentage points of tariff cuts are 3.58% more likely to have boys than their counterparts with lower exposure. As a result of this exposure, mothers are also 8.6-9.9% less likely to give birth in a given year and work up to 10% more hours. These findings are consistent with the mechanism by which trade policy heightens the trade-off between work and children for mothers, causing them to reduce fertility and sex-select more intensely to fulfill their son's preference. Moreover, the results hold up when we account for other competing mechanisms, including changes in fathers' wages, daughters' economic returns, and increases in non-labor income.

Keywords: Sex ratio at birth, Abortion, Trade agreement, Vietnam

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1 Introduction

Sex imbalance at birth is a concerning demographic phenomenon in a number of Asian and Eastern European countries. The sex ratio at birth (SRB), defined as the number of boys per 100 girls at birth, usually stabilizes around a natural level of 105. However, in countries with known son preference such as China, India, and Vietnam, the SRB is much higher than this biological level. In Figure 1a, SRB in these three countries has risen well above 110, foreshadowing a serious shortage of women in decades to come. The case for Vietnam is particularly interesting, because the SRB in this country only picked up in the 2000s, much delayed compared to China and India. Coincidentally around this time, Vietnam also started its integration to the world economy. Notably, in 2001, this country signed a major trade agreement with the US, which has been shown to greatly impact the Vietnamese economy (McCaig, 2011; McCaig & Pavcnik, 2018). Figure 1a shows that, soon after the enactment of the 2001 US-Vietnam Bilateral Trade Agreement (BTA), the SRB started accelerating in this country. In this paper, we link these seemingly unrelated events together in a way that allows us to test existing hypotheses for the rise in SRB.

Several causes for birth masculinity have been suggested in the literature. On the demand side, cultural norms that emphasize male roles in society drive the desire for sons. With such preexisting son preference, easy access to prenatal diagnosis enables sex-selective abortion. In addition to these preconditions, sex selection can further accelerate at lower fertility (Jayachandran (2017)). Declined fertility exerts additional pressure on parents to seek sex-selection technology to guarantee a son at fewer births.[1] Besides these explanations, economic forces can also influence sex selection behavior. For eaxmaple, with rising income, parents can afford the technology to screen and select the sex of their children (Almond et al., 2019). The SRB can also become skewed in favor of boys if the relative returns of daughters falls (Bhalotra et al., 2019; Chakraborty, 2015; Qian, 2008). Our context particularly highlights the role of declining desired fertility for women in explaining the masculinization of SRB.

Our study explores whether the work-childcare trade-off faced exclusively by women affects the sex of their newborn children. We answer this question by exploiting the BTA as an exogenous shock, which particularly affects female-dominated industries and potentially intensifies the work-childcare trade-off for Vietnamese women. The industry-specific export tariff cuts following the trade shock can raise wages, thereby raising the opportunity cost of childcare particularly for mothers. This substitution effect is faced only by mothers

[1] China's "One-Child" policy is an example of forced reduction in fertility, leading to severe sex imbalance at birth in this country.

(but not fathers) when they are the main caretakers of children. Although an increase in their wage rate also has an income effect, if it is dominated by the aforementioned substitution effect, mothers with preexisting son preference will opt for few children but prioritize having sons over daughters. To test this hypothesis, we adopt a difference-in-difference approach and compare the birth outcomes of mothers exposed to greater tariff cuts with that of mothers experiencing lower tariff cuts, before and after the enactment of the trade agreement. The outcomes of our interests are the gender of their infants and their probability of giving birth.

An economy-wide shock such as the US-Vietnam BTA may also affect SRB through other channels. We extend a simple quality-quantity (Q-Q) model to augment the mother's work-childcare tradeoff with other mechanisms that can affect parental choice on sex selection and fertility. This guides us on how to differentiate these competing mechanisms from the mother's substitution channel.

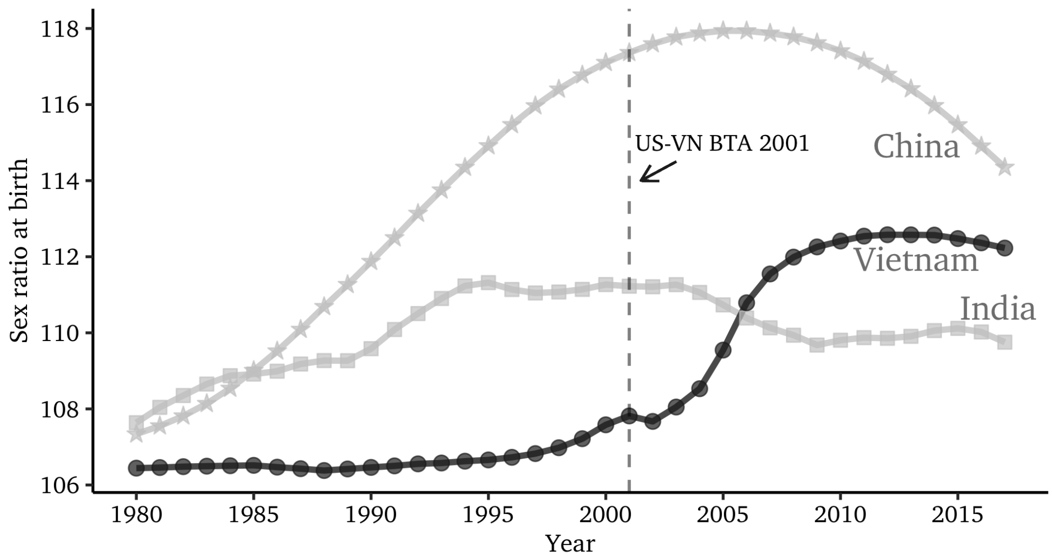
First, there may be a pure income effect of father's wage. This channel echoes the findings in Almond et al. (2019), whose framework predicts that the income effect from wage increase could allow parents to afford sex selection technology. While Almond et al. (2019) do not allow the time cost of children to change differently between mothers and fathers, our framework takes this into account. When father is not involved in childcare, an increase in father's wage induced by the trade shock would only induce an income effect on fertility, raising the desired number of children without effecting the demand for sons. Although our theoretical prediction counts out the income effect of father's wage on sex ratio at birth, we nevertheless acknowledge the mechanism suggested by Almond et al. (2019) and allow for this possibility in our empirical test. Similar to our strategy to test the impact of mother's wage on the birth outcomes above, we test the effect of father's wage with his industry tariff cuts.

Second, the BTA may also raise non-labor income, such as land value. This generates a type of pure income effect similar to that of father's wage above. In our setup, an increase in non-labor income encourages higher fertility without impacting sex selection behavior. Again, we still carefully test for the possibility that rising non-labor income may affect sex selection (à la Almond et al. (2019)). Too do so, we construct a shift-share variable measuring the exposure of a province to the trade shock. Conditioning on mother's and father's direct exposure through their industries, we interpret this measure as a proxy for changes in non-labor income. Effectively, we compare infants born in provinces with more exposure with those born in provinces with less exposure to the trade shock, again before and after the trade shock.

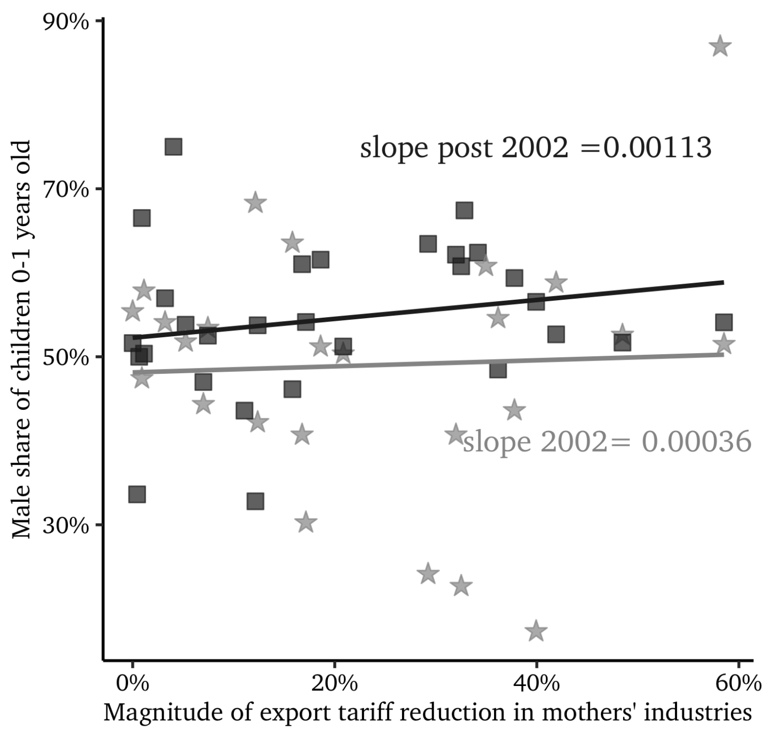
Third, the trade shock can lower the relative returns to girls versus boys, thereby en-

Figure 1: Motivating evidence

(a) Sex ratio at birth by country, 1980-2017



(b) Male share among 0-1 years old children and their mothers' treatment



Notes: (a) Source: Chao et al. (2019). (b) Source: Authors' calculation based on the VHLSS and tariff changes followed the BTA.

couraging parents to opt for boys (Bhalotra et al., 2019; Chakraborty, 2015; Qian, 2008). This explanation is unlikely to hold in this context, since the BTA benefits female workers more than male workers. As a result, the future returns of daughters is more likely to increase than to decrease. Nonetheless, we still test for this possibility in the data. To proxy for changes in relative returns to girls, we construct a female-specific shift share variable that captures the relative exposure to the trade shock of female workers in the province. In a similar difference-in-difference spirit, we compare the differential in birth outcomes over time for infants in provinces with higher relative exposure to the same differential among infants in provinces with lower female exposure. We test simultaneously four channels – the mother's work-childcare trade-off channel and these three alternative mechanisms – in the same regression framework.

We find supportive evidence of the mother's work-childcare trade-off. In particular, mothers of infants on average are exposed to 10 percentage point in tariff cuts over the study period, resulting in a 3.58% increase in the male share among the infants. Consistent with the proposed trade-off between children and work faced by women, women are also less likely to give birth as a result of the trade shock. The average 10 percentage point in tariff cuts that women experience between 2002 and 2016 translates to 8.6-9.9% reduction in their chance of giving birth in a given year. We also find corroborating evidence on the labor supply of mothers: mothers exposed to greater tariff reduction increase their work hours as compared to those getting less tariff cuts. The trade shock could raise their work hours by up to 10% for mothers working in garment and textile manufacturing— the industries with the deepest tariff cuts.

Tested against the mother's exposure channel in a horse race, the other three channels are not likely to drive the skewed sex ratio at birth in Vietnam. In particular, although father's exposure (proxying for father's wage rate) and provincial exposure (proxying for non-labor income) has some positive effects on fertility, their effects on sex ratio are indistinguishable from zero. The increase in the relative female returns proxied by relative female exposure at the province level does not have significant effect birth outcomes, when tested against other forces.

We check the robustness of our results with a number of alternative specifications. In our analysis with repeated cross-sections, we add time-variant observable characteristics of parents and households to control for changes in the composition of the data over time. To improve the identification of the shift-share variables, we also control for the impact of initial "share" over time, as suggested by Goldsmith-Pinkham et al. (2019). We cluster the standard errors at the province level in the main analysis, but we also check with clustering at the mother's industry level; this level of cluster is defined by the treatment unit of the treatment of our interest (mother's exposure). Our results remain stable throughout these checks.

Finally, we exploit the rotating panel structure of our data to complement our cross-sectional results with panel results. Panel data allows us to control for unobservables such as mother's preferences on fertility and sex composition of her children. It also allows us to test the impact of tariff reduction through her *initial* industry prior to the trade shock, thereby addressing concerns about selection bias that could arise from mother switching industries. Although the small size of the panel lowers the precision of some estimates, their sign and magnitude corroborate with cross-sectional findings.

Literature Our evidence for the quality-quantity tradeoff contributes to the literature that aims to explain the causes of the skewed sex ratio at birth. While this literature dated back to the missing women phenomenon observed by Sen (1990), the evidence on the main cause of the rise in SRB remains mixed as recently reviewed by Dasgupta and Sharma (2022).

Earlier work has provided evidence for the fertility-stopping mechanism that we argue here, including Jayachandran (2017) and studies on the One-Child policy in China such as Ebenstein (2010) and Li et al. (2011). However, recent evidence challenges this explanation.

First, the relative returns story by Qian (2008) argues that improvement in the female returns in the labor market should improve the survival of girls. Similarly, Rosenzweig and Schultz (1982) and Carranza (2014) argue that by increasing the opportunity costs of having sons, returns to women's work reduces son preference. Despite including the provincial variation in the relative returns to women through the trade shock, we still find evidence for the fertility story because we can directly observe mothers and fathers' labor market returns using household data.

Furthermore, Almond et al. (2019) argues that land value matters more to sex selection than fertility pressure, even in the context of the One-Child policy in China, as land reform increases household's income and enables families to afford sex-selective abortion. Our Q-Q model, which differentiates the burden of child rearing between mothers and fathers, predicts that land returns or non-labor income only increases fertility and should be independent of sex-selective abortion, after controlling for the mother and father's private returns to labor market. Exploiting the provincial variation in exposure to the trade shock à la Topalova (2010) and Kovak (2013), we find evidence consistent with the model.

We also contribute to the literature on the impact of trade policy that aims to promote economic development on gender discrimination. Prior literature demonstrates that the US-VN BTA increases wages, formality, and reduces poverty (Fukase, 2013; McCaig, 2011;

McCaig & Pavcnik, 2018). We show that this policy may exercabate the sex ratio at birth in the presence of son preference and heightened time constraint for women. Anukriti and Kumler (2019) and Chakraborty (2015) have provided some evidence on the impact of trade liberalization on SRB in India using census data. However, they do not observe the mothers' and fathers' industries and locations at the time of births due to the long gap between censuses. Thanks to our frequent and large household data, not only can we observe the parents at the time of births, but we also account for regional variations that may impact sex selection.

Our study also complements the literature that explainsthe rise in the SRB in Vietnam. Bongaarts and Guilmoto (2015), Chao et al. (2019), and Guilmoto et al. (2009) observe the rise in the sex ratio at birth in Vietnam. These studies stop at providing the facts and point to the rise in the adoption of ultrasound technology which is is associated with sex-selective abortion. However, this observation cannot explain why the SRB in Vietnam started rising in 2001, since abortion was already widespread as a means to reduce fertility prior to 2000. In fact, Guilmoto (2012) and Becquet and Guilmoto (2018) hypothesize that declining total fertility rate from 2.6 in 1999 to 2.0 in 2010 contributes to the rise in sex ratio at birth in Vietnam. They argue the main cause, however, is changes in the intensity of the demand for sons. Son preference (as a cultural norm) tends to be deep-rooted and slow-moving, thus it is unlikely to change in just a few years following the BTA. We account for these time-invariant characteristics in individual panel data and still find that the fertility channel dominates.

Finally, we acknowledge alternative mechanisms that can explain the rise in the SRB in the literature such as the access to ultrasound technology (Chen et al., 2013; Lin et al., 2014), the inheritance motive (Bhalotra et al., 2019), divorce law (Sun & Zhao, 2016), pension program (Ebenstein & Leung, 2010), marriage market (Borker et al., 2020; Wang & Chang, 2002), and changes in policies that directly target the rise in SRB (Anukriti, 2018). Nonetheless, we believe our model plays a more important role in our context than these alternative stories, as we find evidence that are largely consistent with the model's twelve predictions based on variations in both parents' industries and provincial exposures to the BTA.

The remainder of the paper is organized as followed. Section 2 outlines the simple Q-Q model with testable predictions. Section 3 provides the institutional background on the 2001 US-Vietnam BTA and the data employed in our empirical analysis. The end of this section also presents descriptive statistics that hint at our proposed hypothesis. Section 4 lays out the empirical strategy for our tests of the model predictions. Section 5 discusses the results, first for the repeated cross-section sample, then for the panel sample. At the end of this section, we brief through some checks to ensure the robustness of our results. Lastly, Section 6 concludes.

2 Model

We develop a quality-quantity model in which quality is the sex of the child. The child's sex depends on many economic factors including wages of moms and dads, the household's non-labor income, and the relative returns to female-to-male jobs.

Based on the Vietnamese context, we make two assumptions. First, households demand at least one son to carry on the family line. Second, the responsibility of child-rearing falls solely on the mothers, while they also participate in the labor market.[2]

We incorporate these assumptions in a standard Q-Q model in Section 6 of Jones et al. (2010), which features mothers as the primary caregivers. In addition to interpreting quality as the sex of the children, we further impose that the household only needs at least one son.

In particular, consider a household consisting of a woman of reproductive age and her husband. Since they will have children, we call them mom and dad. They decide their private consumption c_g, leisure ℓ_g, the number of children n, and the quality of their children q. Let n be the number of children, and q denote the quality per child, where $q = 1$ means that the child is a son and 0 otherwise.

As households prefer at least one son, we define $Q \equiv qn$ as the effective number of sons. Under the log-utility assumption, the model implies that a household would prefer one son, $Q = 1$, over three daughters, $n = 3$ and $q = 0$. The gender-specific utility function is given by:

$$U_g = \alpha_c \log(c_g) + \alpha_\ell \log(\ell_g) + \alpha_n \log(n) + \alpha_q \log(Q)$$

where $g \in \{m, d\}$, which stands for mom and dad, α_i is the weight on component i among consumption, leisure, fertility, and the effective number of sons.

While each parent has a unit amount of time to allocate between labor and leisure, only women take care of them, which thus costs her γ amount of time per child. In addition to the mother's time cost, the household must pay p_q for the effective quality of the children Q. This price captures the net cost of sex selection, including all the costs and the future

[2] According to a time-use study by ActionAid-Vietnam (2016), women spend about 5.24 hours per day on unpaid care work, while men's time allocation on such activities is 3.17 hours per day. This gender gap in housework widens as the number of children increases. For example, at two children, the wife spends 7.41 hours while the husband spends 4.34 hours hours on unpaid care work. Meanwhile, the female employment rate in Vietnam is high, around 77% among prime-aged women (20-64 years old) in VHLSS 2002-2016.

benefits of sex selection. An example of this cost is the relative returns to women's work to men's, a la Qian (2008) as p_q would increase.

In addition to labor incomes, the household also receives non-labor incomes I, which captures transfers from other household members or their land's value as in Almond et al. (2019). Finally, children and their quality are the public goods of the household. The household's optimization problem is given by:

$$\max_{\{c_m, c_d, n, q\}} \lambda_d U_d + \lambda_m U_m$$
$$\text{s.t. } c_m + c_d + p_q Q \leq I + w_d(1 - \ell_d) + w_m (1 - \ell_m - \gamma n)$$

where λ_m is the bargaining weight of the mother, and λ_d the father's. We assume that $\lambda_m + \lambda_d = 1$.

Defining the household's total income as $W = I + w_m + w_d$, we get the following solutions:

$$\ell_g = \frac{\lambda_g \alpha_\ell W}{w_g}, \text{ for each } g \in \{d, m\},$$
$$n = \frac{\alpha_n}{\gamma} \left(\frac{W}{w_m} \right),$$
$$q = \frac{\alpha_q}{p_q} \frac{W}{n} = \frac{\alpha_q}{\alpha_n} \frac{w_m \gamma}{p_q}.$$

Based on simple comparative statics, we state the following seven predictions on the sex of the children, the mother's labor supply, and the number of children. These decisions depend on four key factors: the mother's wage w_m, her spouse's wage w_h, the household's non-labor income I, and the relative returns to the female's work p_q.

Testable predictions:

1. The probability of having a son increases with mother's wage, conditional on the relative returns to having a son p_q and preferences $\frac{\alpha_q}{\alpha_n}$.

2. The probability of having a son decreases with the net costs to having sons p_q, conditional on mother's wage w_f. Or, it decreases with the relative returns to having daughters.

3. Conditional on mother's wage and the relative returns (or prices) to having daughters, sex selection is independent of all other channels such as father's wage and household's non-labor income.

4. Labor supply of moms increase with their own wages, conditional on household's

non-labor income and husband's wage

5. Controlling for husband's wages and non-labor income, fertility is decreasing in female wage.

6. Fertility is increasing in the wage of the husband, conditional on wife's wages and non-labor income.

7. Fertility is increasing in non-labor income, conditional on mom's wages and father's wages.

We also summarized all of these predictions in Table 1 to help keep track of the various effects predicted by the model. Each row of the table is a statement, "if the factor increases, ceteris paribus, then ...," while the column is the outcome of interest. For instance, the upper leftmost cell reads, "If the mom's wage w_m increases, ceteris paribus, then the probability that the newborn child is a son increases." We proceed to test the seven predictions above using the BTA shock and household data.

3 The Trade Shock and Data

3.1 The 2001 US-Vietnam BTA

The identification strategy in this paper relies on the variation in the exposure to tariff reduction following the US-VN BTA. The agreement was implemented on December 10, 2001. While the trade agreement made little change to Vietnam's tariff rates to US imports, it significantly reduced tariff rates for Vietnamese exports to the US market. Specifically, the US changed Vietnam's trading status from the preexisting Column 2 tariff schedule, which was reserved to former communist countries, to the preexisting Most Favored Nation (MFN) tariff schedule. This led to a substantial immediate reduction in import tariff for Vietnamese products to the US market across all traded industries (Figure A1). While agriculture and mining (code 1-14) received modest reduction, manufacturing industries (code 15-36) enjoyed steep tax cuts – a reduction of 30.2 percentage point on average. The industries receiving the highest tariff reduction are tobacco, textile, and garment productions (code 16, 17, 18), whose tariff cuts range up to 48.5-58.2 percentage points. A few studies have documented that the BTA led to a significant surge in Vietnamese exports to the US starting since 2002 (Fukase, 2013; McCaig, 2011; McCaig & Pavcnik, 2018).

In contrast to tariff cuts committed by the US, the commitment by Vietnam to the bilateral agreement was more focused on regulations and legal reforms. In terms of trade

Table 1: Model and Estimands

Panel A: Model Predictions

If ..., then ...	Prob(son), q	Mom's labor supply, $(1-\ell_m)$	Fertility, n
Mom's wage $w_m \uparrow$	+	+	–
Father's wage $w_d \uparrow$	$\sim$	–	+
Non labor income $I \uparrow$	$\sim$	–	+
Opportunity cost of sons $p_q \uparrow$	–	+	$\sim$

Panel B: Hypotheses

	Prob(son), q	Mom's labor supply, $(1-\ell_m)$	Fertility, n
β_M	+	+	–
β_D	$\sim$	–	+
β_I	$\sim$	–	+
β_F	–	+	$\sim$

Notes: Each cell in Panel A represents a prediction of the model, "If [row] increases, then [column] [cell]. The "+" symbol means increases, "–" decreases, and "$\sim$" does not change. All predictions assume ceteris paribus.

policies, Vietnam was required to cut tariffs mainly in agricultural and food products. Yet the magnitude of Vietnam's tariff cuts for US imports are very small compared to that of the US on Vietnamese exports, ranging between 0.03 and 2.7 percentage points. In addition, Vietnam also eliminated several import quotas under the BTA. However, by the end of 2002, almost all of these had been removed (STAR-Vietnam, 2003). These details make tariff changes from the US side the major factor in the US-VN BTA that affects the Vietnamese economy.

Previous work by McCaig (2011) and McCaig and Pavcnik (2018) have shown that the US-VN BTA has induced dramatic structural change to the Vietnamese economy. Such structural change has reduced poverty rate and increased the formality of the labor market. These papers also show ample evidence for the exogeneity of the BTA tariff cuts committed by the US, including non-correlation between those tariff changes and preexisting economic conditions in Vietnam, which may correlate with prexisting preference for son and fertility. Their work provide a persuasive rationale to consider the US-VN BTA a natural experiment to study its impacts on fertility behaviors of Vietnamese parents.

The swift implementation of the BTA provides the empirical analogs to the exogenous parameters in our model that could potentially affect fertility and sex selection behaviors. These parameters are w_m, w_d, I and p_q, representing the mom's wage, dad's wage, non labor income and the relative returns to daughters. The substantial tariff cuts pose a positive demand shock for Vietnamese workers, especially women who are the dominant workforce in textile and apparel manufacturing. As such, the wage rate in industries with relatively higher tariff cuts will increase compared to wage rate in industries with lower tariff cuts. Thus, we use the variation in the magnitude of the industry-specific tariff reduction to proxy for the change in the wage rate that mothers and fathers (w_m and w_d) experience through their industries.

In addition to potentially changing the wage rate, the BTA can have direct income effect by raising the local average non-labor income (I). For example, it is common in Vietnam for members of an extended family to live in the same household. This suggests that an adult couple can receive transfers from other household members. Consequently, couples living in a province that are more exposed to the BTA could get a larger (non-labor) income effect than those living in a less exposed province. Another channel through which the BTA may affect household's non-labor income is an increase in the local returns to land holdings, which has been shown by Almond et al. (2019) to contribute to the skewed sex ratio in China from 1978 to 1986. To proxy for these potential changes in the non-labor income induced by the BTA, we construct a shift-share measure at the province level.

Lastly, as the BTA induces the largest tarif reduction among industries that primarily employ women, it can increase the female relative wage (captured by p_q in the theoretical model). Parents can internalize this trend as an increase in the relative returns of girls compared to boys, thus decide to have more girls. Using a similar shift-share construction, we generate a female-specific BTA exposure at the province level to proxy for the change in the female relative returns introduced by the trade deal. We elaborate on the construction of these shift-share exposure variables in subsection 3.2 below.

3.2 Data and descriptive statistics

In this subsection, we describe the data used in our empirical analysis. The two main sources of data include (i) household data from the Vietnam Household Living Standards Survey (VHLSS), and (ii) tariff exposure data from McCaig and Pavcnik (2018) in combination with the Vietnamese Census. Our source of household data, VHLSS, is particularly suited for the task at hand, because it captures the workers at the moment of their jobs, thus we do not need to extrapolate this information over a long period of time as in other approaches in the literature. We document descriptive statistics that suggest the trade shock may play a role in the rise of the sex ratio and the increase in labor supply of affected mothers. We also discuss how province-wide exposure to the trade shock that proxies for changes in non-labor income and the relative returns to daughters can be captured with shift-share (Bartik) variables.

Household data

Our primary source of data comes from the Vietnam Household Living Standards Surveys (VHLSS). This household survey is conducted biennially by the General Statistics Office (GSO), with technical support from the World Bank. VHLSS is nationally representative and has a 12-month recall period.

We utilize two main sets of variables in VHLSS: employment and family interrelationship. The employment module provides us the industry code to generate dependent variables that measure each parent's exposure to the trade shock through their industry. This module also provides use information about parents' work hours.

We obtain family interrelationship data from the household roster module. We extract outcome variables that measure fertility and sex selection, including whether a women gave birth recently and the gender of her infant. To make appropriate inference about fertility and sex selection behaviors, it is critical to correctly locate children of individuals. This task in the Vietnamese context is challenging, because extended family cohabitation is common. As a result, VHLSS's default household relation classifier often fail to distinguish one's children from their nieces or nephews. To enhance the family interrelationship, we write an algorithm to identify the spouse, father and mother for each individual. In Appendix A, we document the construction of our algorithm and how it fares in comparison to the true parental locators that are available only for 2014 and 2016 waves. The success rate of our algorithm is 95.27% for mothers, 90.99% for fathers, and 88.95% if both parents are considered. Since we also counts stepmothers and stepfathers, the correction rate for both biological parents is lower than that for each biological parent. These results reassure that our algorithm does a good job at identifying family interrelationship.

VHLSS is particularly suitable for studying fertility and sex selection behaviors. Our main results employ eight repeated cross sections from 2002 to 2016 to estimate the effects following the enactment of the US-VN BTA in December 2001. These high-frequency representative cross-sections enable us to make timely inference about the effect of parents' current jobs on their fertility decisions. Existing approaches in the literature often use infrequent surveys, relying on the assuption that workers are unlikely to migrate or switch jobs. VHLSS overcomes this (unnecessarily true) assumption, while still has reasonably large sample size to tackle the tasks at hand.

Furthermore, we take advantage of the rotating panel feature of VHLSS to form a three-wave panel of women to complement our cross-sectional results. The 2002-2004-2006 women panel helps us control for unobservables such as preferences for fertility and sex ratio. In addition, the panel allows us to distinguish the direct effect of *within* industry exposure from the compositional effect resulting from workers sorting across industries.

Although the first wave of available data is in 2002, the 12-month recall feature still makes it the appropriate data point for the pre-BTA period. Specifically, the BTA enacts on December 2001, but the information (on parents' employment) from 2002 wave mostly applies to 2001 due to the 12-month recall period. Furthermore, our sample selection of infants aged 0-1 year old. This implies that, for children selected from the 2002 wave, the BTA was likely to start late during their gestation, thus outcomes for these children reflects the fertility decision of parents prior to the BTA. The data from 2002 thus constitute our pre-BTA observations. Although one may be concerned about the short pre-period in our data, this may not be a big problem because past evidence has shown that BTA is random (McCaig, 2011; McCaig & Pavcnik, 2018). In addition, we use birth-year to define pre/post period to get more observations in the pre-group, including children born in 1999, 2000, 2001, and 2002. Overall, we get reasonably large sample of roughly 5,000 infants from each wave.[3]

Exposure data to the US-VN BTA

We use two sources of data to construct various measures of exposure to the BTA. First, we make use of industry-specific tariff data for Vietnamese imports to the US at the 2-digit ISIC2 industry level aggregated by McCaig and Pavcnik (2018). The authors obtained the original tariff schedules from the US International Trade Commission. We link industry-specific tariff cuts to each parent's industry code to capture their direct individual exposure to the BTA, which presumably changes their wage rate.

As discussed above, we also construct shift-share measures at the province level to proxy for indirect exposure to the trade agreement that could affect (i) their non-income labor, and (ii) the relative returns to daughters. As for the former, we measure it with

[3]For 2002 only, we have 8,234 infants in the sample because the sample size in this initial round is larger than that of later rounds.

an overall provincial exposure, with the "shifts" being the industry-specific tariff cuts and the "shares" being the pre-BTA industry-specific employment shares at the province level. The "share" component comes from the Population and Housing Census 1999. For each province p, the province-specific exposure to BTA is given by:

$$\tau_p \equiv \sum_j \frac{L_{pj,1999}}{L_{p,1999}} \Delta\tau_j$$

where $\Delta\tau_j$ is the change in tariff imposed by the US on industry j, $L_{p,1999}$ is the labor force of province p based on Census 1999, and $L_{pj,1999}$ is the number of workers in industry j living in province p. Since McCaig (2011) already computed this variable, we extract it from the author's replication dataset.

To capture the relative returns to gender of the child, we also construct a gender-specific shift-share variable in the spirit of Autor et al. (2019). We define the female exposure of province p as:

$$\tau_p^F \equiv \sum_j \frac{W_{j,1999}}{L_{j,1999}} \frac{L_{pj,1999}}{L_{p,1999}} \Delta\tau_j$$

where the only difference between τ_p^F and τ_p is the female share in industry j. Again, using Census 1999, we compute this female share using the industry j's prime-aged work force ($L_{j,1999}$) and the number of female workers in the same industries ($W_{j,1999}$). Unlike Autor et al. (2019), we define this female share, not at the industry-province level, but at the industry level to avoid the selection problem of female-intensive industries locating in female-biased provinces.

Descriptive Statistics

We start with some descriptive statistics on the trends in sex ratio at birth over our study period. Using the parental locators generated by our algorithm above, we link children to their parents and restrict the sample to children who are between 0 and 1 year old at time of the survey. Panel A of Table A4 presents the summary statistics on these children. A little less than half of the infants are girls. Although this (unconditional) fraction seems quite stable over time, below in Figure A2 we show that, conditioning on survey-year fixed effects, the fraction of girls among newborn children declined in the 2000s. This confirms that our data corroborates with the trend in SRB that has documented in the literature with other sources of data (Becquet & Guilmoto, 2018; Chao et al., 2019; Guilmoto et al., 2009). By 2010, this probability has reduced by 5 percentage point. This effect is equivalent to an increase in SRB from 105 to 115, consistent with the pattern documented in Figure 1a.

Mothers of the infants are in their mid twenties when giving birth, while fathers are about three years older than mothers. Both parents completed some or all grades in middle school, reflecting the relatively low skills of the Vietnamese labor force. At the onset of the BTA, mothers and fathers on average received tariff cuts of 8.81 and 7.79 percentage points through their industries, respectively. Over time, mothers' tax cuts deepened by three additional percentage points, while fathers' tax cuts deepened by less than 1 percentage point. This suggests women were gaining more exposure to export-oriented industries than men were over the study period. The work hours of mothers also increase consistently over the years, hinting at the positive link between their commitment to the work force and their increased exposure to greater tariff cuts.

In terms of exposure at the province level, children in a given province were exposed to an average tariff cut of 9 percentage points, and a "female-specific" tariff cut of about 4.15 percentage points.[4] The female-specific provincial tariff cut here is a little less than half of the gender-neutral counterpart, because prior to the BTA women were more likely to work in industries that later received lower tax cuts, such as sales, hospitality, and farming.

Panel B of Table A4 reports the statistics for the data used in the panel analysis. This is a panel of 20-40 year old women who are followed for three consecutive rounds 2002, 2004, and 2006. To study how their sex selection behavior changes over time, we select women who gave birth within one year prior to the BTA and also gave birth once after the BTA was enacted. As a result, the sample here consists of women who are at least one child when entering the panel, thus they are slightly older (early 30s) than the mothers of newborn infants in the cross-section sample. As we follow these women over time, their chance of giving birth increases because they are progressing on their fertility path. Other statistics, including the women's own exposure to the BTA through her industry, their husband's exposure, and the province-level exposure yields similar values as in Panel A.

Lastly, we probe the potential link between the mother's industry-level exposure to the BTA and her sex selection behavior in Figure 1b. Each dot in the scatterplot represents an industry, with its y-axis value as the percentage reduction in export tariffs following the BTA, and its x-axis value being the male share of infants born to mothers working in this industry. We plot the scatter points and fit a linear line separately for infants born by 2002 (and conceived prior to the BTA) and those born after that. While the fitted line for the

[4]These provincial-level exposures were stable through out the entire time frame of our study, because they are time-invariant by definition. The small fluctuations from year to year in these variables simply reflects changes over time in the sample size of children 0-1 year old.

infants conceived prior to the BTA is flat, it is positive for the infants born after this event. This pattern suggests that the mother's exposure to tariff cuts may have a positive impact on the chance of giving birth to sons. In the next section, we formalize the empirical strategy to causally identify this hypothesis.

4 Empirical Strategy

We exploit the trade shock's time, industry, and spatial variations to test the model's predictions in Table 1 . Furthermore, we leverage cross-sectional and panel data to alleviate concerns about each approach.

4.1 Baseline specification

Motivated by the model in Section 2, we specify the following difference-in-difference (DID) regression and estimate it by OLS:

$$\begin{aligned} y_{imdpt} = {} & \beta_M \cdot \tau_{j(m)} \cdot \text{Post}_t + \beta_D \cdot \tau_{j(d)} \cdot \text{Post}_t + \beta_I \cdot \tau_p \cdot \text{Post}_t + \beta_R \cdot \tau_p^F \cdot \text{Post}_t \\ & + \theta_t + \xi_p + \psi_{j(m)} + \phi_{j(d)} + X'_{imdpt} \cdot \lambda + \varepsilon_{imdpt}, \end{aligned} \tag{1}$$

where each variable y_{imdpt} is an outcome for child i who is less than two years old at the time of survey and born to mom m and dad d in province p and in year t. Using cross-sectional data, we observe two outcomes including whether the child is male, denoted by the dummy $\text{Male}_{imdpt} = 1$ if he is, and mom m's reported working hours in the main job.

We focus on four coefficients β_M, β_D, β_I, and β_R. Variations in tariff reduction across mom's industry $j(m)$ and the time of birth relative to the enactment of the trade shock allow us to identify the first coefficient β_M. Consider, for instance, outcome Male_{imdpt}. The interaction between the time-invariant tariff change in moms' industry $\tau_{j(m)}$ and the indicator Post_t, which is true if child i's birth year is after 2002, captures the changes in the sex selection behaviors of mothers working in industries with higher tariff reduction after compared to before the trade policy takes place.

Similarly, identification of β_D follows by replacing $\tau_{j(m)}$ with tariff reductions in dads' industries $\tau_{j(d)}$. Furthermore, the interaction between the province exposure to the trade shock τ_p and the time dummy identifies β_I, which examines whether households in more exposed provinces adjust their behaviors after the trade shock. Finally, estimates of β_R let us test if the relative returns to female's work, captured at the province level by τ_p^F, impacts sex selection and moms' labor supply.

We include the year fixed effects θ_t (which can be birth year or survey year or both depending on outcomes) to capture any aggregate shocks that equally impact all households in year t, province fixed effects ξ_p to capture any time-invariant differences across provinces, moms' industries fixed effects $\psi_{j(m)}$ any time-invariant differences across moms' industries, and dad's industry fixed effects $\phi_{j(d)}$ any time-invariant differences across dads' industries. The time-varying controls X_{imdpt} includes moms' and dads' birth age, their education, household size, urban dummy, and minority dummy.

In addition to repeated cross-sectional data, we can modify specification 1 to estimate a similar set of coefficients using panel data. Our panel sample consists of women between 20 and 40 years old who have given birth at least once before and after 2002. For each woman i in this sample, we can observe the tariff change in her industry τ_i and her husband industry τ_h, as well as whether her last-born child is a boy and her labor supply. Thus, the DID specification becomes

$$\begin{aligned} y_{iht} = \beta_M \cdot \tau_{j(i)} \cdot \text{Post}_t + \beta_D \cdot \tau_{j(h)} \cdot \text{Post}_t + \beta_I \cdot \tau_p \cdot \text{Post}_t + \beta_R \cdot \tau_p^F \cdot \text{Post}_t \\ + \theta_t + \psi_i + X'_{iht} \cdot \lambda + \varepsilon_{iht}, \end{aligned} \tag{2}$$

where y_{iht} is an outcome in survey year t of woman i who has children with husband h. The dummy $\text{Post}_t = 1$ if survey year is larger than 2002. We include individual fixed effects ψ_i and the same set of controls in X_{iht} as X_{imdpt} but replacing a mom and a dad's variables by the individual and her husband's ones. When investigating the effects on whether the last born child is male, we also include the last born child's birth year fixed effects.

All regressions are weighted by sampling weights. Standard errors are clustered at the province level. Our hypothesis for outcomes are summarized in Panel B of Table 1.

4.2 Fertility outcomes

To investigate the effects on fertility, we follow two complementary approaches. First, we construct a retrospective panel from the cross-sectional data in which an observation is at the birth-mom-year level. For each woman, we construct an observation whether she gives birth to a child at the women-year level, from the year she turns 20 years old to the year she is observed in our data. The outcome of interest is whether there is a birth i to woman w and dad d. The specification is similar to (1), except the inclusion of mom's fixed effects instead of the province and industries fixed effects. This approach helps us observe her fertility path (up to the year of survey) and control for mother's fertility and sex selection preferences when including mom's fixed effects in the specification. However, since we can only observe both parents' employment data at the time of survey, we must assume that their employment data, which we can only observe once at the survey year, are unchanged across time. Although this assumption may fail to reflect their true employment at the time of birth, it remains informative for families that rarely changed their industries.

Our second approach relies on the *true* panel data. We follow women between 20 and 40 years old across three waves 2002, 2004, and 2006, and examine the effects of trade shocks on whether they have a new birth within the year of the survey according to specification 2. This approach enables us to test, for instance, whether higher tariff reduction in a woman's industry reduces the probability that the mother gives birth after versus before the implementation of the policy. Due to the short panel, we can only observe the short-run impact of the policy.

5 Results

5.1 Repeated Cross-sections

We first report results estimated with repeated cross-sectional data on the sex ratio at birth of newborn children and on the labor supply of their mothers. After that, we also discuss fertility results on the restrospective panel constructed from cross-sectional data.

Sex ratio at birth

Table 2 reports the results of the regression specified in Equation 1 in the first three columns. Column (1) just includes the four measures of exposure to the BTA and their interaction with the $Post_t$ dummy. Column (2) adds controls for differences in socio-economic backgrounds across the children. Column (3) includes the fixed effects without the socio-economic controls. Column (4) add both the controls and fixed effects in the previous two columns. The rows report the coefficients of interests, β_I , β_R, β_M, and β_D. They capture the different channels through which the BTA could affect the gender of newborn children, as suggested by Predictions 1, 2, and 3.

First, Prediction 1, which states that the probability of a male birth increases in the mother's wage, is supported by the data. The positive and significant coefficient on the interaction between mother's exposure and the post-2002 dummy (capturing β_M) suggests that mother's direct exposure has a positive impact on the probability that her infant is a boy, conditioning on a measure of relative returns to daughters (represented by β_R). In fact, across the four specifications in columns (1)-(4), β_M is the only statistically significant effect and its size remains stable. Specifically, each percentage point reduction in tariff that

Table 2: Effects on whether a new birth is male and labor supply

	1(Male)				Mom's Hrs Worked	1(Recent Birth)		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Mom Exp x Post	0.0020**	0.0019**	0.0019**	0.0019**	0.2727***	-0.0001***	-0.0001**	-0.0001**
	(0.0008)	(0.0008)	(0.0008)	(0.0008)	(0.1021)	(0.0000)	(0.0000)	(0.0000)
Dad Exp x Post	-0.0005	-0.0005	-0.0003	-0.0003	0.1547	0.0002***	0.0002***	0.0002***
	(0.0008)	(0.0008)	(0.0008)	(0.0009)	(0.1197)	(0.0000)	(0.0000)	(0.0000)
Prov Exp x Post	-0.0032	-0.0036	0.0025	0.0029	-14.8158***	0.0018	0.0013	0.0042***
	(0.0077)	(0.0077)	(0.0151)	(0.0150)	(2.2548)	(0.0015)	(0.0014)	(0.0011)
Female Exp x Post	0.0016	0.0028	-0.0038	-0.0044	28.0347***	-0.0016	-0.0002	-0.0006
	(0.0170)	(0.0170)	(0.0245)	(0.0244)	(4.9418)	(0.0023)	(0.0021)	(0.0020)
R^2	0.000	0.001	0.008	0.009	0.333	0.286	0.290	0.290
Observations	31,176	31,175	30,771	30,770	28,303	2,038,111	2,038,100	2,038,100
Control Group Mean	0.6820	0.6820	0.4898	0.4898	134.1040	0.0186	0.0186	0.0186
Controls		✓		✓	✓		✓	✓
Province FE			✓	✓	✓			
Mom Industry FE			✓	✓	✓			
Dad Industry FE			✓	✓	✓			
Survey Year FE			✓	✓	✓	✓	✓	✓
Birth Year FE			✓	✓		✓	✓	✓
Mom FE						✓	✓	✓

Notes: Sample of children 0-1 years old, column (1) no control, (2) add controls which include household size, mom birth age, mom education, dad birth age, dad education, urban, and minority (3) add fixed effects, (4) adds both household controls and fixed effects. Column (5) the number of working hours per month of the first or main job of these mothers. It includes controls the woman's age, her husband's age, her education, her husband's education, urban, and minority. Columns (6)-(8) use the constructed panel, as explained in the main text. The outcome variable is a dummy for whether the woman has given birth in the past 2 years. The controls include mom's birth age, number of elder brothers, and number of elder sisters. All regressions include survey weights and are estimated by OLS. Standard errors, clustered at the province level, are reported in the parentheses.

the mother experiences through her industry is associated with a 0.19 percentage point increase in the likelihood that the child is male.

Prediction 2 states that a rise in the relative returns to daughters leads to lower male share among newborns. This prediction can be verified with the coefficient on the local female tariff cuts interacted with the post-2002 dummy (β_R). This prediction has some weak support from the empirical results, especially with the full specifications described by 1 in column (4). Conditioning on mother's exposure to the tariff cuts, a percentage point reduction in the local female tariff rates is associated with a 0.045 percentage point increase in the probability that the infant is female. However, this estimated effect is imprecise.

Finally, Prediction 3 states that changes in father's wage and non-labor income has no effect on sex selection. Consistent with this prediction, the father's exposure (β_D) to the BTA through his industry, as well as the provincial exposure capturing the effect from changes in non-labor income (β_I) are statistically insignificant once we control for mother's exposure and a term capturing the relative returns of daughters.

These results highlight the mother's direct exposure as the main driver of the skewed sex ratio at birth induced by the BTA. While the trade shock may potentially affect sex selection through other channels (returns to daughter or income effect), our horse-race here helps us eliminate these alternative explanations from our hypothesis on the quantity-quality trade-off faced by mothers. It is useful to put the effect of mother's exposure estimated above in perspective. Over our study period, mothers of these children were exposed to a tariff cut of 10.6 percentage point on average. Our estimate above implies that the 10.6 percentage point reduction in mothers' tariff rates over the study period would lead to a 1.9 percentage point or 3.58% increase in the newborn male share. The magnitude of this effect is very close to the fall in the probability of being female among newborn children documented in Figure A2. In fact, an increase in SRB from 105 (the natural level) to 113 (roughly the level reported by GSO towards the end of the study period) implies a increase of 3.76% in the infant male share - a remarkably close number to the effect size implied by our estimation.[5] The magnitude of this estimate are quite in line with findings from a few other studies. Anukriti (2018) finds an Indian program intended to resolve fertility-sex ratio trade-off ends up increasing the probability of first birth to be male by 1-2.3%. Almond et al. (2019) find land reform in China between 1978 and 1984 increased the fraction of males following a first girl by 3 percentage point or 5.6%.[6] Note that estimates from these studies are specific to birth parity and gender composition of the first sibling, while our findings are estimated across all birth orders.

Mother's work hours

The results on sex selection at birth above point toward the hightened quality-quantity tradeoff that mothers faces as the a result of the BTA. We further look into this work-childcare tradeoff by examining the impact of the BTA on mothers' work hours. We run the same regression as in Equation 1, but with the mother's work hours per month in her

[5] GSO annual statistics on sex ratio at birth can be found at: https://www.gso.gov.vn/dan-so/

[6] from a baseline sex ratio of about 112 for this group of children

main job as the dependent variable. The sample here includes the mothers of the infants who appear in the previous sex selection analysis. The estimates are reported in Column (5) of Table 2.

The mother's own exposure to tariff cuts (representing changes in her own wage in Prediction 4) increases her work hours. Mothers who were more exposed to the BTA by one percentage point through her main job on average work an additional 0.27 hour per month, compared to less exposed mothers. This estimate is statistically significant, but seems small in magnitude. However, this effect is observed among women who recently gave birth and might be on maternity leave, possibly explaining why the increase in hours for them is modest.[7] If we consider a mother working in the apparel industries, this result implies an increase of 15.66 hours per month, equivalent to a 10% increase from a baseline of 157.11 hours per month among mothers in such industries.

Her husband's exposure (representing changes in his wage) is associated with an increase in the mother's work hours. Although the direction of the effect fits Prediction 4, the estimate is imprecise.

Prediction 4 does not involve the returns for daughters, which we proxy with the female relative provincial exposure in the previous subsection. In this regression, however, the female provincial exposure may partially capture the relative wage of mother compared to father. Thus it is no surprise that it has a positive impact on mother's work hours. However, compare to mother's industry exposure effect, the effect from the female provincial exposure is much larger, indicating that mothers living in a province with more advantageous female exposure work substantially more hours than mothers living in provinces with poorer female exposure. One percentage point increase in the female provincial exposure is associated with 15.5 additional work hours. Here, the female province exposure captures the labor demand for women in the local economy. Not only does such demand directly affect the traded industries in the province, but it could also spill over to the non-traded service industries. Thus this measure is likely to pick up a much larger effect than the industry-specific exposure does.

Conditioning on the mother's direct exposure and female provincial exposure, the gender-neutral provincial exposure represents non-labor income. The result here shows that it has a negative impact on mother's work hours. Taken together, the findings here are rather consistent with Prediction 4.

[7]The sample of mother here is smaller than the sample of children in the sex selection analysis. This most likely reflects maternity leave, because since 2010 wave, VHLSS asks about the number of work hours within the last 30 days, resulting in missing values for work hours. If the mother is on maternity, she will drop out of the sample, although her child appears in the infant sample.

Fertility

Columns (6), (7) and (8) in Table 2 examines the fertility side of the work-childcare trade-off faced by mothers. These three columns are estimated with Equation 1 on the retrospective panel constructed from cross-sectional data. We report the same coefficient of interests as before. Three estimates here helps us check Predictions 5, 6 and 7.

First, Prediction 5 states that, all else being equal, an increase in mother's wage leads to a smaller number of children. The mother's direct exposure through her industry, which proxies for her wage change induced by the BTA, has a negative impact on the chance of giving birth. The magnitude of the estimate is stable across specifications: each percentage increase in tariff cuts in the mother's industry is associated with a 0.07-0.08 percentage point reduction in her probability of giving birth in a given year. Given mothers experience on average an increase of 10.6 percentage point in tariff cuts between 2002 and 2016, this translates to a drop of 0.7-0.8 percentage point, or a 8.6-9.9% reduction, from a baseline of 8.06% chance that a woman gives birth in a given year.

On the other hand, the father's wage change, is likely to increase the probability of having an additional child, as stated by Prediction 6. This reflects the pure income effect of the husband wage because he faces little work-childcare trade-off. The (gender-neutral) provincial exposure which proxies for other non-labor income is also predicted to have a positive impact on fertility, according to Prediction 7. However, the estimate here is inconclusive, and mostly insignificant in most specifications.

Lastly, the female relative return, proxied by the female provincial exposure, does not have a significant role in fertility decision. This is also consistent with the absence of this variable in the model's solution for optimal fertility choice. Overall, the the results on fertility seems to support Predictions 5 and 6, but inconclusive about Prediction 7.

5.2 True Panel

In this subsection, we discuss the complementary results estimated with the 2002-2004-2006 panel of women. We estimate the same effects in a panel-data specification (in Equation 2) that potentially improves upon the previous identification strategy using repeated cross-sections. First, we can control for mother's unobserved preferences on fertility and sex composition of children. Second, this specification allows us to estimate the causal effect of the BTA on mothers and fathers using their *initial* industry prior to the BTA. This helps relieve concerns about the compositional bias that could result from parents sorting into industries with greater tariff cuts.

Restricting the sample to 20-40 year old women who recently gave birth in both the pre

Table 3: Effect on whether the last birth is male, labor supply, and fertility. Panel data

	1(Last born is male			Hrs Worked	1(Recent Birth)
	(1)	(2)	(3)	(4)	(5)
Own Exp x Post	0.0074**	0.0035	0.0091***	0.6105	-0.0004
	(0.0031)	(0.0031)	(0.0031)	(0.5342)	(0.0012)
Husband Exp x Post	0.0067	0.0062	0.0068	-0.6714	0.0000
	(0.0042)	(0.0037)	(0.0043)	(0.7210)	(0.0011)
Prov Exp x Post	-0.0617	-0.0285	-0.1168	9.8204	-0.0219
	(0.0756)	(0.0795)	(0.0716)	(20.9042)	(0.0257)
Female Exp x Post	0.0619	0.0424	0.0829	3.2277	0.0142
	(0.1272)	(0.1235)	(0.1110)	(35.8795)	(0.0348)
Had a son before		-0.5672***			
		(0.0636)			
R^2	0.82	0.85	0.84	0.76	0.48
Observations	2,596	2,596	2,596	2,514	11,045
Control Group Mean	0.2857	0.2857	0.2857	108.9048	0.0741
Controls	✓	✓	✓	✓	✓
Individual FE	✓	✓	✓	✓	✓
Last Child Birth Year FE	✓	✓	✓		
Survey Year FE	✓	✓	✓	✓	✓

Notes: Panel of women between 20-40 years old from 2002-2006 and who have given birth once before and after 2002: Column (1) includes, as controls, household size, woman's age, her husband's age, her education, her husband's education, urban, and minority dummies. In addition to these controls, columns (2) adds whether having a son before the last born. Column (3) includes 1999 agriculture and manufacturing shares interacting with the birth years of the last born. Column (4) reports the effect on first job working hours per month including the same controls and interactions between survey years and shares of agriculture and manufacturing. Column (5) uses the panel of women 20-40 years old between 2002-2006 and the dependent variable is whether having a new born in the last year newchl1y. The regression is the same as the labor supply one with all controls. All regressions include survey weights and are estimated by OLS. Standard errors, clustered at the province level, are reported in the parentheses.

and post periods of the 2002-2004-2006 panel, we have a small sample of 1,443 women and a total of 2,596 women-year observations. By construction, the sample requires the women to have at least one child in 2002, so the effect of the BTA post 2002 will be identified on the second birth and above. We report results on sex selection in columns (1)-(3), on mother's work hours in column (4), and on fertility decision in column (5). Note that for fertility results, we do not requires the women to give birth in both the pre and post periods, so the sample size are larger.

Sex ratio at birth. Column (1) estimates the equation specified in Equation 2, while controlling for parents' demographic information, household characteristics, and the birth order of the child born in the survey year. According to Predictions (1)-(3), two variables may have an impact on the sex of a newborn child: mother's exposure through her industry (capturing mother's wage) and female provincial exposure (capturing returns to daughters).

Like cross-sectional results, only mother's exposure has statistically significant impacts in the gender of her infant. Compared to mothers with less exposure to the trade shock, mothers with more exposure are more likely to give birth to a boy after the enactment of the BTA. The size of the effect here more than triple the same estimate in the cross-section regression, possibly because it is conditioned on higher birth order. Notice that the R^2 here is much larger than than the R^2 in cross-sectional regressions (columns (1)-(4) in Table 2) because we can control for mother fixed effects.

In column (2), we further control for whether the mother has a son before the current birth. The estimate on the interaction between mom's exposure and post-2002 dummy drops by half and becomes insignificant. This result aptly points to the preference for at least one son assumed in our model: as long as they have at least one son, the BTA no longer imposes a quality-quantity constraint on the women.

Column (3) gets back to the basic specification in column (1), but further controls for the effect of initial economic conditions over time to ensure the robustness of the shift-share independent variables. This slightly increases the size of the effect of mother's industry exposure.

Mother's work hours. Column (4) regresses the mother's number of work hours on the same independent variables as in column (3). The sample is also restricted to the same mothers as in column (3). As in our sample selection for the cross-sectional results for mother's labor supply, we focus on mothers who recently give births because the work-childcare constraint is most binding for them. None of the reported estimate is statistically significant, perhaps because of the small sample size, but their sign and magnitude are quite in line with the cross-sectional estimates. Notwithstanding their lack of precision, these panel estimates cement our earlier findings that support Prediction 4.

Fertility. Column (5) looks at the effects of the BTA on the probability of mothers giving birth. The sample here no longer imposes that the women must give birth once in 2002 and at least once in 2004 or 2006. As a result, the sample size increases substantially compared to the previous for four columns. However, this is still very small relative to the sample size of the fertility retrospective panel regression in Table (2). Perhaps this explains why the fertility estimates are underpowered. Despite this, again, the direction and magnitude of these panel estimates are also similar to those in the retrospective panel approach. Overall, the result in column (5) also corroborates with Prediction 5 and 6. That is, the mother's wage, induced by her direct industry exposure, reduces the probability that she gives birth, but the father's wage has the opposite effect. Prediction 7 predicts that non-labor income, captured by the gender-neutral provincial exposure, increases fertility. However, this is not consistently supported in the data (including the results here and in the last three columns of Table 2).

5.3 Robustness

This subsection aims to address potential concerns to identification and alternative explanations to our main results in Table 2. Table A5 shows that our results remain stable throughout.

To alleviate concerns with identifying of the shift-share variables, we include interactions between survey years and the pre-shock agriculture and manufacturing shares at the province level using the population census in 1999 in the spirit of Goldsmith-Pinkham et al. (2019). Furthermore, as argued by Ebenstein (2014), agriculture intensity predicts the strength of patrilocality, which correlates with son preference. Thus, having the initial agriculture shares $\times$ year fixed effects allows us to account for potential changes associated with the initial levels of patrilocality that might affect the intensity of son preference at the provincial level.

We interact time-varying controls with a dummy for Post 2002 to account for potential changes in the impact of our control variables on outcomes after the trade shock.

We further account for changes in infrastructure projects such as road network over time that could affect access to sex-selective abortions as argued by Almond et al. (2019) by including region-year fixed effects. Inter-provincial development projects in Vietnam are often clustered into eight regions: Northeast, Northwest, Red River Delta, North Central Coast, South Central Coast, Central Highlands, Southeast, and Mekong River Delta. After this inclusion, the effects of mothers' exposure on our main outcomes change little.

By clustering our standard errors at the industry's level instead of the province level, the statistical significance of our result on sex selection increases to 5%. Since we use both variations at the parents' industry and provincial shift-shares, we keep our clustering at the province level in our main results. Overall, our results are robust to these alternate specifications.

6 Conclusion

In this paper, we explore the causal link between mothers' work-childcare trade-off and their decision on the sex of their infants. Mothers who face a higher time cost of children are often pressured to have fewer children. Yet in a son-favored society with easy access to sex-selection technology such as Vietnam, the decision to lower fertility may also couple with an intensified demand for boys. We exploit the 2001 US-Vietnam BTA, a major trade agreement which coincides with the rise in SRB in Vietnam, to test this hypothesis.

We find that mothers who are more exposed to the BTA through their industries are less likely to give birth, and conditional on giving birth, their infants are more likely to be boys. We also find that more exposed mothers also work more hours, suggesting that the opportunity cost for children has increased for these mothers. These results are robust when tested against other potential channels through which the BTA may affect fertility and sex selection decisions.

Our findings highlight a troubling consequence of economic growth on demographic imbalance in fast-growing Asian economies with preexisting son preference. Development policies strive for better economic opportunities for women, yet female engagement in the labor force may pressure for a greater demand for sons. Given that women play an important role in economic development, the resulting son-favored bias in fertility outcomes are likely to sustain in these contexts. Therefore, in order to address the sex imbalance in these countries, awareness and sensitization campaigns to direct the norm towards gender equality will be much needed.

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A Appendix: Family interrelationship data

Variables categorizing family interrelationship within the household are crucial in this paper. In particular, we need "pointer" variables identifying each person's mother, father, and spouse. VHLSS has a variable that classifies each individual's relationship to household head. However, this variable is ambiguous and erroneous for the following reasons. First, extended family is common in Vietnam: roughly 34.91% of VHLSS sample live in the same household with relatives other than their nuclear family members. Because VHLSS does not identify subfamilies within the extended household, the "relate-to-head" variable will miss out the interrelationship of subfamilies other than the head's. Second, with the exception of VHLSS 2002, all other waves do not identify "children-in-law." An immediate consequence is missing spousal links and therefore parental links in many cases. Thus, we write an algorithm to improve the precision of parent-child and spousal pairings in VHLSS.

We adapt a similar algorithm that has been used by the Minnesota Population Center (Sobek & Kennedy, 2009) to generate "pointer" variables for IPUMS–International Census Data. We utilize the four variables, relationship to household head, age, and marital status, in combination with the relative position of household members in the roster listing, to infer relationships. We first establish spousal links (generate variable SPLOC), then we find mother and father for each individual (generate variables MOMLOC and POPLOC). Regarding parental linkage, we look for the mother before looking for the father; once we find the mother for a child, we assign her husband as the father of this child. Only when we cannot a child's mother, we locate its father (and assign his wife as the mother).

Tables A1 and A2 document the rules we apply to these matching tasks. Our algorithm applies the rules sequentially: if the first rule finds a match (spousal match or parental matches) for a given individual, the second rule no longer applies to this person, and so on. Whenever there are ambiguous multiple potential spouses or multiple potential parents, we drop the entire household from the sample.

Fortunately, VHLSS 2014 and 2016 do have two variables to locate the biological fathers and mothers of children under 16 years old. We use this information to test our algorithm and report the results in Table A3. We find the correct mother for 95.27% of these children and the correct father for 90.99% of them. The algorithm also finds both parents correctly for 88.95% of the children. The correction rate for both biological parents is lower than that for each biological parent because our algorithm also counts stepmothers and stepfathers. With these results, we are very confident that our algorithm does a good job at identifying family interrelationship.

Table A1: Rules for SPLOC construction

Rule	Individual's relationship to head	Partner's relationship to head	Age difference	Both Married	Require adjacency	Only applicable to 2002	Notes
Strong couple pairing, couple adjacency prefered							
	Head	Spouse	No	No	No		1
	Parent	Parent	No	Yes	Yes		1
	Grandparent	Grandparent	No	Yes	Yes		1
	Child	Child-in-law	No	Yes	Yes	Yes	1
Weak couple pairing, couple adjacent							
	Grandchild	Grandchild	Yes	Yes	Yes		1, 2
	Other relationship	Other relationship	Yes	Yes	Yes		1, 2
	Sibling	Sibling	Yes	Yes	Yes	Yes	1, 2
	Grandchild	Other relationship	Yes	Yes	Yes		1, 2
	Sibling	Other relationship	Yes	Yes	Yes	Yes	1, 2
	Child	Other relationship	Yes	Yes	Yes		1, 2
Weak couple pairing, special type child-child							
	Child	Child	Yes	Yes	Yes		1, 2, 6
Weak couple pairing, couple not adjacent							
	Child	Child-in-law	No	Yes	Closest proximity	Yes	5
	Child	Other relationship	Yes	Yes	No		2, 3
	Child	Child	Yes	Yes	No		1, 2, 4, 6

Notes:

1. Drop the entire household if there is any person that could be assigned to 2 couples by the adjacency rule.

2. A woman can be no more than 20 years older or 35 years younger than a potential male partner.

3. For non-adjacent couple pairing, among the potential spouses who satisfy the age and marital requirements, select the person who is closest in age and impose that the husband is older than the wife.

4. For child-child non-adjacent couple pairing, drop all households where this rule yields multiple potential spouses.

5. These non-adjacent couples are matched based on having closest proximity to each other.

6. Once the couple is identified, the in-law is disinguished. For child-child couples, assume the first listed spouse in the biological child to the household head, hence the second listed spouse is the child-in-law.

Table A2: Rules for MOMLOC and POPLOC construction

Rule	Child's relationship to head (updated after SPLOC is generated)	Parent's relationship to head (updated after SPLOC is generated)	Age difference	Proximity requirement	Only applicable to 2002	Notes
Links involving Head, Spouse, and Grandparent (unambiguous)						
	Child	Head, spouse	No	No		
	Child-in-law	Head, spouse	No	No	Yes	
	Head	Parent	No	No		
	Spouse	Parent	No	No		
	Sibling	Parent	No	No	Yes	
	Parent	Grandparent	No	No		
Links between grandchildren and children						
	Grandchild	Child, child-in-law	15-44	Weak		1
	Grandchild	Other relationship	15-44	Weak		1, 2
Links involving other relatives						
	Other relationship	Grandchild	15-44	Weak		1
	Other relationship	Other relative	15-44	Weak		1
	Other relationship	Sibling	15-44	Weak		1
	Other relationship	Child, child-in-law	15-44	Weak		1, 3

Notes:

1. Weak proximity requires that the child must be listed after its potential mothers (or potential fathers if it has no potential mothers); among them, its mother is the one listed closest to the child.

2. Impose that no person with code ”Child” is present in the household. The mother of the grandchildren in these cases tend to be listed as ”Others” since there is no category for ”Child-in-law.”

3. Impose that no person with code ”Grandchild” is present in the household. These cases tend to mix up the numerical code for ”Grandchild” (code 6) and that for ”Other relationship” (code 7).

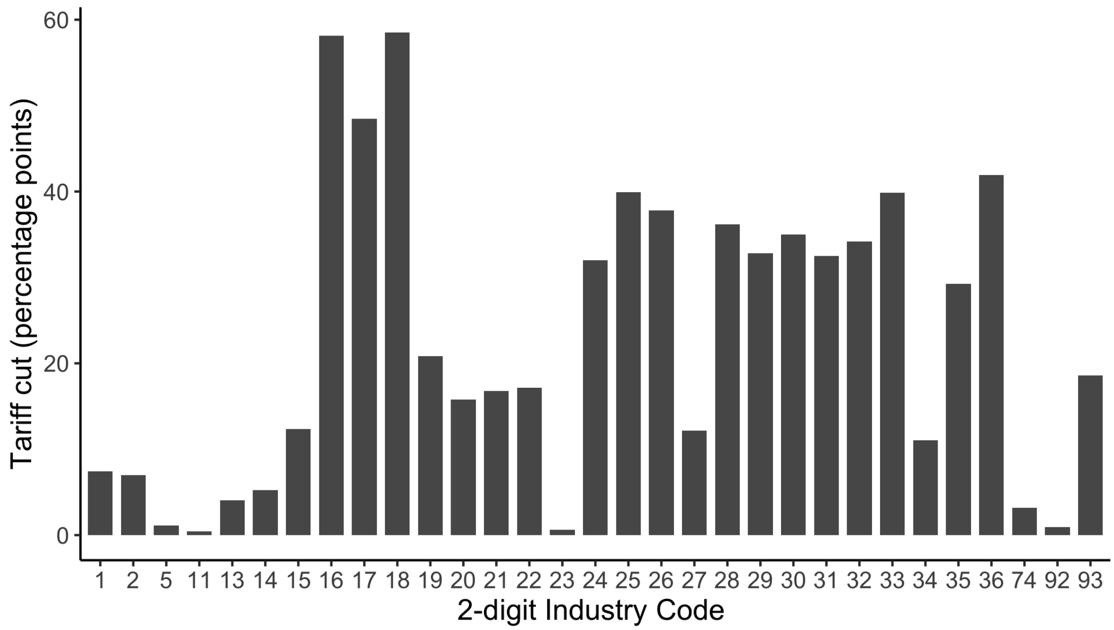
Table A3: Performance of our algorithm in locating parents

	Freq.	Percent	Cum.
Both mom and dad correct	80021	88.95	88.95
Dad correct, mom incorrect	1831	2.04	90.99
Dad incorrect, mom correct	3853	4.28	95.27
Both mom and dad incorrect	4253	4.73	100.00
Total	89958	100.00	

Notes: This table summarizes the comparison between the parent locators generated by our algorithm with the true parents locators provided by VHLSS 2014-2016 for children under 16 years old.

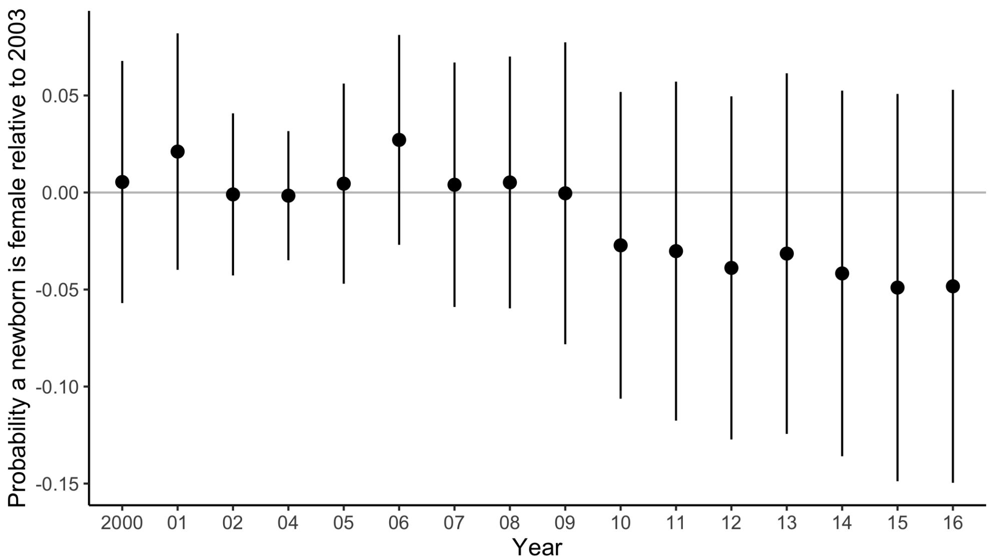
B Appendix: Additional Figures

Figure A1: US tariff cuts by 2-digit ISIC industry codes



Notes: Data come from McCaig and Pavcnik (2018)

Figure A2: Sex ratio trend in VHLSS



Notes: This graph plots the birth year fixed-effect coefficients from a regression of female dummy on birth year fixed effects and survey year fixed effects. The omitted group is children born in 2003. The sample of this regression consists of newborn children between 0-1 years old.

C Appendix: Additional Tables

Table A4: Descriptive Statistics

Panel A: Kids stats

Characteristic	**2002**	**2004**	**2006**	**2008**	**2010**	**2012**	**2014**	**2016**
Male	53%	53%	53%	51%	53%	53%	53%	52%
Mom Exposure	8.82 (11.40)	9.43 (12.85)	10.19 (14.12)	10.03 (14.67)	10.66 (15.36)	10.97 (15.76)	12.27 (17.33)	11.80 (16.91)
Dad Exp.	7.80 (10.52)	7.75 (10.99)	7.95 (11.61)	7.49 (11.78)	8.27 (12.77)	7.68 (12.30)	8.11 (13.00)	8.19 (13.08)
Province Exp.	8.90 (1.75)	9.14 (2.04)	8.97 (1.89)	9.15 (2.03)	9.09 (1.93)	9.04 (1.86)	8.86 (1.49)	8.94 (1.77)
Female Exp.	4.10 (1.04)	4.24 (1.23)	4.15 (1.13)	4.24 (1.21)	4.19 (1.14)	4.15 (1.10)	4.02 (0.85)	4.08 (1.05)
Household size	6.25 (2.29)	6.49 (2.33)	6.31 (2.22)	6.31 (2.29)	5.68 (1.84)	5.81 (1.79)	5.84 (1.73)	5.79 (1.66)
Mom birth age	27.10 (6.43)	27.25 (6.53)	26.85 (6.42)	27.02 (6.25)	26.78 (5.99)	26.80 (5.95)	26.96 (5.98)	27.12 (6.08)
Mom edu	6.57 (3.59)	7.06 (3.70)	7.32 (3.71)	7.94 (3.66)	8.19 (3.65)	8.42 (3.65)	9.07 (3.37)	9.26 (3.41)
Dad birth age	30.38 (6.72)	30.81 (6.87)	30.38 (6.69)	30.79 (6.74)	30.23 (6.45)	30.17 (6.36)	30.71 (6.67)	30.41 (6.64)
Pop edu	7.07 (3.51)	7.48 (3.50)	7.66 (3.56)	8.22 (3.48)	8.35 (3.52)	8.59 (3.44)	9.10 (3.25)	9.20 (3.28)
Mom's Main Job (hr/month)	110.42 (76.34)	122.05 (67.33)	125.26 (67.96)	127.32 (69.88)	150.29 (70.56)	149.68 (70.35)	151.43 (68.88)	153.86 (69.67)
Urban	21%	24%	26%	30%	28%	27%	28%	27%
Minority	20%	20%	21%	19%	20%	20%	21%	24%
Num. Obs.	8,234	4,569	4,482	5,826	5,365	5,089	6,538	4,826

Panel B: Panel stats

Characteristic	**2002**	**2004**	**2006**
Had a child within the past 2 years	17%	23%	24%
Main job (hr/month)	132.64 (74.26)	140.12 (69.40)	144.04 (68.15)
Own Exposure	8.62 (11.30)	9.43 (12.94)	9.91 (13.87)
Husband Exp.	7.14 (9.33)	7.19 (9.82)	7.16 (10.03)
Province Exp.	9.04 (1.87)	9.13 (1.99)	9.11 (1.98)
Female Exp.	4.16 (1.12)	4.23 (1.20)	4.21 (1.18)
Age	31.49 (5.74)	31.13 (6.21)	30.40 (6.47)
Education	7.07 (3.48)	7.43 (3.54)	7.75 (3.61)
Urban	23%	24%	27%
Minority	15%	15%	16%
Num. Obs.	5,707	5,524	5,486

Table A5: Robustness: Effects on whether a new birth is male and labor supply

	1(male)	Mom's Hrs Worked	1(male)	Mom's Hrs Worked	1(male)	Mom's Hrs Worked	1(male)	Mom's Hrs Worked
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Mom Exp x Post	0.0019**	0.2732***	0.0019**	0.2792***	0.0019**	0.2561**	0.0019***	0.2732*
	(0.0009)	(0.1022)	(0.0009)	(0.1015)	(0.0008)	(0.1038)	(0.0005)	(0.1495)
Dad Exp x Post	-0.0003	0.1648	-0.0003	0.1800	-0.0002	0.1968	-0.0003	0.1648*
	(0.0009)	(0.1215)	(0.0009)	(0.1224)	(0.0009)	(0.1210)	(0.0005)	(0.0935)
Prov Exp x Post	-0.0251	-9.4960***	-0.0295	-9.4021***	-0.0166	-0.4450	-0.0251	-9.4960***
	(0.0191)	(3.0085)	(0.0211)	(2.5067)	(0.0258)	(2.3605)	(0.0175)	(3.1083)
Female Exp x Post	-0.0492*	15.5537**	-0.0379	12.5827**	-0.0576	-4.1198	-0.0492	15.5537**
	(0.0273)	(6.9000)	(0.0283)	(6.1951)	(0.0360)	(5.2211)	(0.0306)	(7.5672)
Standard-Errors	province						momisic2	
R^2	0.011	0.337	0.011	0.338	0.013	0.344	0.011	0.337
Observations	30,770	28,303	30,770	28,303	30,770	28,303	30,770	28,303
Control Group Mean	0.4898	134.1040	0.4898	134.1040	0.4898	134.1040	0.4898	134.1040
Controls	✓	✓	✓	✓	✓	✓	✓	✓
ControlsPost			✓	✓				
svyear fixed effects	✓	✓	✓	✓	✓	✓	✓	✓
region8-svyear fixed effects					✓	✓		
Province FE	✓	✓	✓	✓	✓	✓	✓	✓
Mom Industry FE	✓	✓	✓	✓	✓	✓	✓	✓
Dad Industry FE	✓	✓	✓	✓	✓	✓	✓	✓
Birth Year FE	✓		✓				✓	

Notes: Sample of children 0-1 years old, column (1) is similar to Column (4) in Table 2 but adds 1999 shares interacting with birth years. Column (2) the number of working hours per month of the first or main job of these mothers. It includes all controls of Column (5) in Table 2 plus the interactions between survey years and the agriculture and manufacturing shares at the provinces in 1999. Columns (3) and (4) add time-varying controls interacting with dummy for Post 2002. Column (5) and (6) add region8-svyear fixed effects to (1) and (2). All regressions include survey weights and are estimated by OLS. Clustering of the standard errors are reported above the R^2 row.