Impact of Industrialization on Relative Female Survival: Evidence from Trade Policies[+]

Tanika Chakraborty

DIW, Berlin

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Washington University, St Louis

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Abstract

Previous attempts at understanding the role of industrialization on son preference are confounded by the endogeneity of the industrialization process. This paper exploits an exogenous shift in the trade policy in India to study the impact of industrialization on son preference. Using a difference-in-differences strategy, I find that households are more likely to have a male child in regions with higher trade openness relative to regions with lower degrees of trade liberalization. Moreover, trade openness seems to have intensified son preference for Hindus but not for Muslim households. I further analyze the underlying mechanisms through which industrialization might have affected son preference. I find no significant evidence to suggest that household income or a change in female labor market opportunity is affecting relative female survival following trade liberalization. On the other hand, I find a significant increase in real dowry payments in regions experiencing greater trade openness. Most interestingly, dowry inflation is concentrated in the Hindu households, but not in the Muslim households. The results are robust to falsification tests using cohorts born much before the liberalization period and are not driven by systematic migration into areas with greater trade openness.

JEL classification: J16, J18, J71, O15

1. Introduction

The imbalance in sex ratio characterizing the population of many Asian countries has caused a number of authors to argue that there is substantial excess female mortality in Asia (Sen 1990; Coale 1991; Klasen 1994). In general, they argue that neglect of female children and poor conditions for women contribute to the abysmally low proportion of fmales in the population.

Previous theoretical and empirical economic insights on gender imbalance focus on household level economic incentives affecting imbalance in sex ratios. Some of these are income (Becker 1981, Rose 1999), differential earning opportunities for women (Bardhan, 1977; Rosenzweig and Schultz, 1985, Duflo 2002, Qian 2005) and relative status of women within different kinship systems (Dyson & Moore, 1993 and Chakraborty & Kim, 2008). There are also studies that put forth biological factors, to explain sex ratio imbalance (Norberg, 2004; Oster, 2005)[1].

However, the previous literature has paid less attention to the importance of economy wide changes that might in turn affect household choices. Recent sex ratio statistics show a consistent decline in the proportion of females in many Asian countries, particularly over the past few decades. Contrary to what common sense would dictate, over the same period these economies also had rapid industrialization enabling them to come out of the traditional agricultural sector. China and India, in particular, have experienced high growth rates driven mainly by the non-agricultural sector and at the same time the concern for relatively fewer females in the population has intensified. From a public policy standpoint it is important to understand whether these counter intuitive associations are causal since it would imply that the relative scarcity of women will keep growing in the decades to come with current predictions of persistent economic growth. Few papers which do address the debate on the role of economic development in reducing gender discrimination are Foster and Rosenzweig (2001), Burgess and Zhuang (2001), Bhattacharya (2006) and Pande & Astone (2007). Foster and Rosenzweig explore the agricultural growth in India following Green Revolution and geographical marriage market of women to underline the role of economy wide changes in shaping son

1 Norberg argues that cohabitation causes male bias. Oster claims that Hepatitis B explains about 20% of the sex ratio differential in India. However, both of these hypotheses affect sex ratio only at birth, they don’t explain the differential mortality in early childhood.

preference. However, growth in agriculture says little about the impact of modernization on son preference. Bhattacharya (2006) finds that share of nonagricultural laborers or better infrastructure significantly reduced relative excess female mortality[3]. Pande & Astone, 2007, on the contrary, find that village level economic development is not enough to change gender preferences. Thus, the existing evidence on the role of economic development in changing excess female mortality, based on aggregate data and cross section studies, is mixed. One important reason for the mixed results could be the endogeneity of industrialization and economic development with parental preference for sons. Particularly, if non agricultural growth is driven by supply of educated labor and better educated households have systematically different preferences then estimates of the effect of industrial or economic development on relative female survival would be biased. In this paper, I address this endogeneity problem by using an exogenous policy shift in India, to analyze the impact of industrialization, at the time of the policy shift, on relative female survival. This paper also contributes to the growing body of literature which analyzes the consequences of trade openness on various measures of development. While previous literature focuses mainly on the effects of trade on wage inequality, gender gap in wages and schooling and child labor outcomes, there is no previous attempt at understanding how trade openness affects household preferences for sons or daughters. In this paper I discuss how trade affected the relative survival chances of female children and the mechanism driving such changes.

Following the balance of Payments crisis of 1980s, India was forced to undergo a series of trade reforms in the 90s, starting with the slashing of tariff rates in July 1991. The average effective tariff rate was reduced from about 86 per cent in 1989-90 to about 40 per cent in 1994-95, and further to about 30 per cent in 1999-2000 (Goldar, 2002). In this paper I exploit the variation in the degree of trade openness across districts to capture the effect of industrialization on son preference. Using tariff data from the World Integrated Trade System (WITS) I compute district level measures of the magnitude of tariff reduction in the industrial sector following the 1991 trade reforms. Then I link a household's probability of having a male child with the district level trade measures. The exposure of parents to the trade reform is determined both by the extent of tariff reduction in the district of their residence and by the cohort of birth of the concerned

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3 Burgess & Zhuang, find qualitatively similar evidence from China's rural industrialization

child. In other words, a child is exposed to the trade reform only if it is born close to or after the reform year of 1991. My estimation strategy is similar in spirit to a difference in difference strategy although it also estimates the effect on each birth cohort exposed to the reform. This decreases the probability that the results are confounded by the effect of other policies in the liberalization period. I find that households are more likely to have a male child in districts with higher trade liberalization as compared to their counterparts in the districts with relatively smaller changes in trade openness. My results are robust to falsification tests using cohorts born much before the liberalization period and hence not likely to be affected by preexisting differential district trends. They are also not driven by systematic migration into areas with greater trade openness.

At first glance it might appear that industrialization and development of an economy should lead to an improvement in the sex-ratio due to a positive change in all other socio-economic parameters that might affect son preference. Theoretically, however, industrialization might affect relative survival of girls through several channels. To the extent that poverty might drive households to allocate relatively fewer resources to the less valuable children, trade liberalization might lead to greater negligence of girls if poverty increases. Second, trade openness might affect relative preference for sons over daughters if the relative wage of women changes. Third, trade liberalization might affect relative marriage market value of females versus males by changing employment opportunities. For example, if trade leads to higher employment opportunities for men outside the traditional agricultural sector then marriage market value of men might increase. In a social setting characterized by marriage payments this will lead to a greater marriage price for men[4]. I find no significant evidence to suggest that household income or a change in female labor market opportunity is affecting relative female survival following trade liberalization. On the other hand, I find a significant increase in real dowry payments consequent upon trade liberalization. The practice of dowry as a marriage payment is traditionally prevalent amongst the Hindus and not Muslims. When I compare the 2 groups I find that trade openness has raised dowries only for Hindus. If dowry increases parental discrimination towards daughters then we would expect relative female survival to decline more for Hindus. Indeed, I find that trade liberalization affects

4 Anderson, 2003, provides a detailed theoretical framework explaining this mechanism

relative survival of females only for Hindus. There is no comparable effect in the Muslim population.

The rest of the paper is organized as follows. Section 2 discusses the theoretical mechanisms by which industrialization might affect relative female survival. Section 3 describes the exogeneity of the trade reforms and its impact on industrialization. Section 4 outlines the empirical framework and estimation specifications. Section 5 summarizes the data sources. Sections 6 to 10 report and discuss the results and robustness checks. Section 11 concludes.

2. Conceptual Framework

Industrialization can affect relative survival of girls versus boys through a number of channels. First, it might change the relative status of women in the labor market by changing the returns to female labor in the industrial sector. Second, industrialization can also affect relative female survival through a household income effect, by changing poverty. Becker (1981) argued that relative investment on female children responds to household income. However, previous evidence on this front is shows a weak relation between income and female survival. Qian (2005) does not find any causal link between total household income and sex specific investment on children. Lastly, industrialization, or in other words a movement away from the traditional occupational structure, might affect sex ratios by changing marriage payments for daughters. Anderson (2003) shows that in a socially segregated society, like India, industrialization leads to increasing dowry payments through increasing within-group wealth inequality due to a movement away from the traditional occupation structure. In India, social status is determined not only by wealth but there is also an inherited component to status (caste) which is patrilineal. Due to patrilineal inheritance if a girl marries down in the caste hierarchy, it leads to a loss in social status of her family. Anderson shows that in this framework where marriages are arranged by the families of the bride and the groom, positive assortative matching implies that girls are matched with grooms of similar or higher status with respect to both caste and wealth. Modernization in such a framework leads to dowry-inflation due to increasing within caste inequality. To the extent that increase in dowry would reduce the

desirability of girls, this might play an important role in determining relative investment on daughters in terms of health care or increasing the possibility pre natal sex selection.
Formally, in order to see how industrialization, through trade liberalization, might affect parental preference I develop a simple model where parental preference for daughters is reflected in their observed birth history. The theoretical framework on how trade policies might shape parental preference for sons and hence determine relative female survival is discussed in Appendix I.

3. Trade reforms and industrial growth

Like many developing countries, import-substituting industrialization was one of India's development strategies from the 1950s to the early 1980s. A complex regime of import licensing requirements along with other barriers to trade kept the Indian economy fairly insulated from international competition. Around the late 80s the government embarked upon an effort at economic reforms including reducing barriers to trade, especially allowing imports of capital goods. However, while a reduction in the barriers to trade and the de-licensing process in India started in the mid 1980s when Rajiv Gandhi was elected the prime minister after a series of political crisis, the major breakthrough came about only in 1991 after the assassination of Rajiv Gandhi and election of the new leader, Narasimha Rao. The 1991 reforms was a follow up of the macroeconomic crisis of the late 1980s, whereby India was conditioned to implement a structural adjustment program in return for a stand-by arrangement with the IMF.
The maximum tariff was reduced from 400 percent to 150 percent in July 1991. After the initial slashing of tariffs in 1991 there was a gradual reduction in tariffs through 1997. Subsequent reductions saw the maximum tariff down to roughly 45 percent by 1997–98. The average effective tariff rate was reduced from about 86% in 1989-90 to about 40% after the initial tariff cut in 1991, and further to about 30% in 1999-2000 (Goldar, 2002).
While the standard Heckscher-Ohlin-Stolper-Samuelson framework predicts that lowering of trade barriers would lead to an increase in demand for skilled workers in the developed countries and unskilled labor in developing countries, theoretically the effect of trade liberalization on manufacturing employment is unpredictable. Among other things it would depend on the social and physical infrastructure of the country, labor

market conditions like the excess labor or dualistic labor markets and type of specialization – import competing or export oriented[6].

However, previous evidence on the impact of trade reforms on industrialization in India provides ample evidence on the growth in productivity and employment of the manufacturing sector following trade liberalization in 1991. For e.g. Goldar (2000, 2002) and Tendulkar (2000), among others, find evidence of acceleration in employment growth in the post-reform period both at the aggregate manufacturing level and for most two-digit industries. Table-3A reports estimates of employment growth in various industrial groups and aggregate manufacturing. As has been noted by several authors, a large part of the employment growth was driven by export oriented industries due to the large growth in labor intensive manufacturing exports after trade liberalization. Estimates suggest that growth rate in aggregate manufacturing employment increased from -.12% to 2.92% per annum while in export oriented industries it increased from -0.57% to 3.36% per annum consequent upon tariff cuts in 1991 (Goldar, 2002). Table 1, Appendix 2, providing employment growth estimates across consumer, intermediate and capital goods industries also indicates a marked increase in all three industry types in '90s. Compared to a growth rate below 8% per annum between 1980 and '90, exports grew at a rate of more than 25 % after the reforms (Pandey, 2004). Along with employment, Topalova, 2004 finds that trade liberalization led to an increase in the both levels and growth of firm productivity.

While these estimates are primarily based on time series data, using the NSS data between 1987 and 1999, I separately create a panel of district level estimates of fraction of employment in the industrial sector as well as in non-agricultural occupations. When I regress these on estimates of district level net tariff I find supportive results after taking in to account district and year fixed effects. Details on the estimation of net tariff and non-agricultural employment are provided in section 5. Table 3B, column 1, shows the effect of net tariff reduction on proportion of employment in non agriculture. It shows that a 10 percentage point fall in tariff rate, increased proportion of the population employed in the non-agriculture sector by 4 percentage points. I find similar results by

6 Ghose (2000) discusses the theoretical implication of trade liberalization on employment and wages in details

using employment in non-agricultural occupational category instead of non-agricultural industrial classification, reported in column 2.

In this paper I use this exogenous shift in trade policy as a proxy to evaluate the impact of industrialization on relative female survival. Even though tariff changes continued to take place in the later phases between 1995 and 1997, I use only the initial structural break of 1991. An analysis of the subsequent reduction between 1995 and 1997 would be confounding since they cannot be taken as exogenous knowledge to the households anymore.

4. Empirical Strategy

Few studies in the past that have looked in to the effect of industrialization on the gender inequality in survival, correlate district level measures of industrialization at different points of time with sex ratio in the 0-4 age group in that district. The difficulty with these studies is that even after controlling for time and district fixed effects, there might be omitted variables that are correlated with industrialization that also affect gender inequality. For example, a bigger supply of educated labor may trigger greater degree of industrialization in a region and also affect the sex ratio. However, if growth in industrialization is an exogenous shock determined by unexpected government policies then it will be independent of the omitted variables. Since the liberalization policy led to different degrees of tariff cuts in different industries and there was a wide variation in location of specific industries across different districts in India, the effective tariff decline varied widely across different districts. I utilize this variation of tariff reduction to estimate the effect of a child's exposure to trade policies on relative survival chances of female children. Particularly, using a cross section of children born between 1980 and 1999, I linked the probability of observing a male child in a family with district-level measures of employment weighted tariff protection following the 1991 reforms. The time and district of birth of a child and the percentage decline in tariff in district of birth determine the extent to which the decision-making households are exposed to the reform. This strategy enables us to identify the effect of industrial growth around the time of the trade reform. Thus if the identification strategy is correct we would expect it to affect only the cohorts born close to and after the reform. If the identification is spurious, for e.g. it is driven by unobserved factors in regions of high tariff decline or is a part of a trend, then we would expect that the effect of a tariff decline will be reflected on all cohorts born in regions with higher tariff decline.

Fig 1 illustrates the rough idea behind the identification strategy. It plots the proportion of males in each birth cohort for the regions with high tariff decline (green) vs. the regions with low tariff decline (red)[7]. Since children in the 0-5 age group are most sensitive to nutritional and health care provision, cohorts below 5 years of age at the time of (and after) the policy change are most likely to be affected. It shows that cohorts born much before the reform, in areas with higher post reform tariff decline, had lower proportion of males. In cohorts born close to and after the reform (specifically, 1988 onwards), the proportion of males fell in both types of regions but it fell more in regions with lower tariff decline. The change in the proportion of males between the high and low tariff regions occurred for cohorts who happened to be in their early childhood (0-3) during the reform and for cohorts born immediately after the reform supports the identification strategy used in this paper.

4.1. Estimation Specification

Formally, in a strategy similar to Duflo (2001), I estimate in a difference-in-difference framework the effect of trade reforms on the relative survival chances of female children who are exposed to the reforms[8]. Since children in the below 4 age group are most sensitive to health care and nutrition deficiencies, children born after 1987 would be exposed to differential treatment from parents due to trade reforms[9]. Moreover, the degree of exposure varies with the intensity of reforms, as measured by the percentage of tariff decline in a region, discussed above. The following equation estimates the effect of trade reforms on female survival.

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7 High decline regions comprise districts above the median in the distribution of percentage tariff decline. Low tariff decline regions comprise districts below median.

8 This methodology has also been used by other authors, including Card and Krueger (1994), to estimate the impact of school quality on educational achievement and labor market outcomes. A more recent adoption of the strategy is by Chen an Zhou(2007) to estimate the impact of famine on the health of the affected.

9 The literature on gender inequality in survival ubiquitously focuses on the sex ratio in the 0-4/0-5 age group as discrimination is likely to be most influential in affecting relative mortality during these early ages

$$M_{ihdt} = \alpha_1 + (T_d * Young_i)\, \beta_1 + X_h\, \delta_1 + \gamma_{1d} + \varphi_{1t} + \varepsilon_{ihdt} \qquad \textbf{(1)}$$

where, M_{ihdt} is the event that child i born to household h in district d and belonging to birth cohort t is male. T_d is the percentage tariff decline in district d after the 1991 reform, calculated from the base pre- reform tariff rate. $Young_i$ indicates a dummy for whether child i belong to the younger cohort (i.e. born after 1987) which would be affected by any change in parental behavior due to the reform. γ_{1d} is the district fixed effect that differences out time invariant district characteristics; and φ_{1t} is the cohort of birth fixed effect that differences out any trend effect that are invariant across all districts. X_h represents household specific demographic controls (dummies for household head's education, whether household belongs to schedule caste/tribe and household standard of living index). In the analysis I include only those children for whom the first sibling was born in 1980 or later enabling me to observe complete families. The coefficient on the interaction between child i's cohort of birth and the intensity of tariff decline in the corresponding district of birth measures the effect of trade reforms on relative survival of the females. In particular, a positive β_1 implies that the probability of female survival, in a cohort exposed to the reform, decreased due to a greater fall in tariff. In this case trade policy serves a proxy for industrialization.

While the above strategy measures the overall impact of tariff decline on the younger cohort who is likely to be affected by the policy change, it cannot outline the differential effect that the policy might have on each cohort. Specifically, in accordance with the earlier literature on son preference which emphasizes sex discrimination in the 0-4 age group, we are assuming that the cohorts born after the reform and those in the below 4 age group at the time of the reform would be affected. However, it does not say anything about which cohorts were affected by the policy change. It is quite possible that the effect of trade reforms varied over time (in other words it varied by cohort). To measure differential impact of trade reforms on different cohorts, I estimate an unrestricted coefficient model allowing the effect of trade policy on relative female survival to vary by year of birth. Specifically, I estimate the following equation:

$$M_{ihdt} = \alpha_1 + \sum^{19}_{q=2} (T_d * c_{iq}) \beta_{1q} + X_h \delta_1 + \gamma_{1d} + \varphi_{1t} + \varepsilon_{idt} \qquad \textbf{(2)}$$

Where M_{ihdt} is a dummy variable if the living child born in household h district d, cohort t is a function of: the interaction term between T_d, the extent of decline in effective district tariff rate in district d and Ciq, a variable which indicates if child i is born in year q. γ_{1d} again is the district fixed effect and φ_{1t} is the cohort of birth fixed effect. The dummy variable for the 1980 cohort and all of its interactions are dropped. Each coefficient β_{1q} can be interpreted as an estimate of the impact of the trade reform on the probability of being male for cohort q.

Equation (1) assumes that relative survival chances of male children compared to female children should have been affected for cohorts born close to and after the reform. The exact timing of the response in relative survival will depend on the nature of sex discrimination. If sex selection was conducted by infanticide or pre natal selection, then the reform should only affect sex ratios of cohorts born after the reform. However, if sex selection is conducted by negligence in childhood, then the reform can also affect sex ratios of children who were born a few years before it. To see if the latter effect is at work, in other words, if reform increased early childhood discrimination against girls, we need mortality data by gender. DHS data does provide detailed birth history of each woman interviewed. From the birth history records I can construct a subsample of death records to see if female children exposed to the reform have higher probability of death in areas with higher tariff decline.

However, due to the limited number of observations on death records of children, I cannot estimate the unrestricted model in equation (2). Instead, I pool the records of children born to the interviewed mothers and who died in the early years of childhood and estimate a specification similar to equation 1.

$$F_{ihdt} = \alpha_1 + (T_d * young_i) \beta_1 + X_h \delta_1 + \gamma_{1d} + \varphi_{1t} + \varepsilon_{ihdt} \qquad \textbf{(3)}$$

where, F_{ihdt} is the probability of being female in the subsample of children who died within the first few years of life. $Young_i$ is a dummy indicating whether the dead child belonged to the younger cohort, which would have been affected by the trade policy.

Moreover, the biological literature on gender specific mortality of children suggests, that in infancy girls have a relative biological advantage over boys, so that we usually see an excess of male deaths. While this biological advantage continues for the entire phase of childhood, the biological difference subsides with age[10]. Since male mortality is generally much higher in infancy (below 1 year of age) and hence may confound the results, I restrict my analysis to children who died after infancy, at ages more than 1 year.

5. Data

The empirical estimation in this paper primarily relies on the 1999 round of Demographic and Health Survey (also known as National Family and Health Survey) conducted in India with aid from USAID**.** It provides complete birth history of 90,303 ever-married women in the age group of 15-49. I focus on all children born in and after 1980, which accounts for 95% of the data. This is primarily because cohorts born before 1980 do not have sufficient observations. To observe complete families in terms of children ever born, I keep only those observations where the first child is born in and after 1980 and their corresponding siblings. Using information on the district of birth of each child, I matched the individual survey data with district level estimates of percentage tariff decline during 1991. Table-2, Panel A, outlines the average demographic and occupational characteristic of the population in this sample. Following Rose (1999), I construct my dependent variable as the probability of being male or female conditional on survival during the time of the survey. This means that my main dependent variable reflects a measure of female deficiency due to both prenatal sex selection and higher relative mortality. Since the transition from traditional to industrial economy is relevant mainly for the rural sector, I limit the analysis to the rural households.

For constructing the tariff measure I use custom duty data obtained from the World Integrated Trade System (WITS) database which provides the data at the 6-digit HS (Indian Trade Classification Harmonized System) code level. These HS codes are then matched with the 3 digit NIC codes using the concordance of Debroy and Santhanam (1993). I use the tariff reduction in the in the manufacturing sector.

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10 See Waldron (1985) for a discussion of the biological causes of higher male mortality in childhood. Bhuiya and Streatfield(1992) &Muhuri and Preston(1991) also discuss the prevalence of sex discrimination after infancy.

The tariff measure used in this paper is in line with Topalova (2004). Since the fall in protection was identical across all regions in India, but varied across industries, the tariff measure exploits district level variation in industrial composition prior to the 1991 reforms. Specifically, district d's tariff at time t is measured by weighting the tariff in industry j and time t by the pre-reform employment share of industry i in district d. Thus tariff in district d, time t is given by

$$\textbf{Tariff}_{\textbf{dt}} = \textstyle\sum_{\textbf{j}} \textbf{w}_{\textbf{jd}} * \textbf{tariff}_{\textbf{jt}}$$

Where,

$$w_{jd} = \frac{employment_{jd}}{\Sigma_j employment_{jd}}$$

Then I compute weighted tariff in each district d in 1991 by using tariff data from 1991. For tariff measure prior to the reform I use tariff duty data for year 1990, the earliest year for which tariff data is available.

Finally I take the percentage change in tariff between pre reform and 1991 as a measure of the intensity of trade reform in a district in 1991. The average tariff rates in the manufacturing sector are outlined in Table 2, Panel B. On an average tariff declined by almost 30% after the first phase of liberalization in 1991. It is this phase of the liberalization shock that we use in our analysis. Since this measure of tariff uses district specific employment weights determined prior to the trade reform, changes in industrial composition after the policy change that are result of the tariff changes do not affect the measure of exposure to trade reforms (Topalova 2003).

Sex specific wage data comes from 3 rounds of NSS covering the time period before and after the trade reforms. The NSS is a repeated cross section at the household level which collects a broad set of data on sampled households every 5 years. It's a nationally representative sample covering all states. I use the Employment and Unemployment rounds of the survey available for the years 1983, 1987-88, 1993-94 and 1999-2000. Apart from wage data, I use them to construct estimates of sex ratio and non-agricultural employment share at the district level.

Finally, I use the Rural Economic and Demographic Survey of NCAER to obtain Dowry data. It is a nationally representative longitudinal survey covering demographic, economic and health information for approximately 7500 households distributed over 100 districts spread over the 16 major states of India. Using retrospective marital information I create a panel of individuals married over various points of time and link the individual marriage payments data to the district level measure of trade liberalization.

6. Results

6.1 Industrialization and Son Preference

To investigate the relationship between industrialization and gender composition of the population I conduct basic OLS regressions using the NSS district estimates for sex ratio and fraction of non-agricultural employment from 1987-1999. Table-1A lists the results. Column (1) presents the OLS estimates of the effect of share of non-agricultural sector in a district on the proportion of males in the 0–4 age group using. Column (2-5) control for the fraction of illiterate, concentration of backward social groups and female labor force participation rate respectively. All the regressions imply that a higher share of employment in the non agricultural sector is associated with a higher fraction of male children in a district.

When we try to look at the long term relationship between proportion of non-agriculture and female to male sex ratio from the Census data (1961 – 1991) a similar relationship is corroborated after correcting for district level clustering effect. The results are shown in Table – 1B. Female to male sex ratio in the age group 0-5 (and 5-9) decreases with proportion of agricultural employment in a district after controlling for the district literacy rate, proportion of Scheduled castes/tribes, proportion of population who are landless, and the rate of urbanization. This means that in infancy, the biological reasons might be more dominant than parental preference.

However, in all these results the causal interpretation depends on the assumption that there is no unobserved variation or omitted variables in these regressions that might be correlated with the error term. However, there might be an omitted variable bias in these OLS estimates. First, there might be family background variables like education which determine the supply of educated or skilled labor in a district and hence, to an extent, industrialization, and at the same time exercise less/more son preference[11]. Therefore if

11 Shastry, 2008, in her working paper discusses how districts with a more elastic supply of English skills

we cannot fully control for the parental skill, we would underestimate/overestimate the effect of industrialization on the female to male sex ratio. Previous literature analyzing the effect of economic development on relative female survival controls for a variety of factors to overcome this problem. However, there might still be unobserved variables driving the bias.

Since the trade liberalization polices of 1991 potentially had a large impact on industrialization and at the same time were unexpected, I use it as an exogenous shock to analyze the impact of trade policies on the relative survival of females.

6.2. Basic Results

Table 4 shows the estimation results from regression equation (1). The previous literature points to birth order as an important determinant of male biased sex ratio. This means that if we take a sample of children born in year t and onwards, we would be selecting in the group some children who are of higher birth order and this may bias our results. Hence, I include only those children for whom the first sibling was born in 1980 or later enabling me to observe complete families.[12] Column (1) shows the results for the entire sample of living children (as of 1999). “Young” refers to the fact that the child under observation belongs to the younger cohort (born post 1987, so that they are in the 0-4 age group in 1992) which is exposed to the trade reforms. Specifically, I define “Young” as an indicator taking value 1 when the individual was born between 1988 and 1996 and 0 when the person was born between 1980 and 1988[13]. All regressions control for a dummy indicating whether the household belongs to disadvantaged social group (Scheduled Caste or Scheduled Tribe). It shows that for children born in the younger cohort and hence exposed to the reform period, those born in regions of higher trade openness (higher absolute tariff decline) are more likely to be male. The underlying assumption here is that individuals born between 1988 & 1996 are affected by the reform. To see

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experienced greater growth jobs requiring such skills

12 I checked other years prior to 1980 for which data on birth records are available. However, after retaining only those children where the oldest sibling is born in year *x*, records for all *x* prior to 1980 are insufficient to be included in the analysis.

13 I do not include the cohorts born after 1996 since in 1997 there was a second phase of tariff cut, which I do not include in my analysis. It cannot be expected to be exogenous with households forming expectations about future tariff cuts. Hence, pooling later cohorts in the “Young” group might confound the exogenous effect of tariff change in the 1st phase. The second phase of tariff reduction reduced across industry variability in tariff compared to the first phase.

which cohorts are indeed affected by the policy change of 1991, I estimate equation (2). The unrestricted coefficient estimates are presented in column (1) of Table 5 for the whole sample. The estimates suggest for children born close to and after the policy change (1987 onwards) the male bias increases significantly with the extent of percent tariff reduction. This indicates that apart from prenatal selection (that would affect cohorts born after the reform) there is possibly differential treatment of females who are still in their early childhood during the first tariff cut and hence more susceptible to differential child care and nutrition. On the other hand, tariff decline did not have any significant effect on the cohorts born between 1981 and 1987. In other words the tariff decline did not affect cohorts that were born much before the first round of liberalization and hence not exposed to the reform. This finding lends support to the causal effect of the change in trade policy and reaffirms that the estimates in Table 4 are not driven by pre-existing differential trends between high and low tariff-change regions. It affected only those cohorts that would have been exposed to the reform.

6.3. Absence of Elder Male Sibling

The literature on sex ratio emphasizes on the role of siblings in son preference. In particular, it is well established that preference against a daughter is strongest in the absence of a son. In an exercise similar to regression equation (1), I estimate the effect of the trade reform on the probability of having a male child in the absence of an older male sibling[14].

The results are presented in column 2 of Table 5 for the whole sample from estimation of equation (2). However, the results are not significantly different from zero except for the cohort of 1991.

7. Bias from selective migration

Any analysis with spatial variation is fraught with possibilities of selective migration. The analysis above does not account for the possibility of increased migration that might have followed trade liberalization. The underlying assumption is that a woman interviewed at her current place of residence has stayed there ever since her children were

14 A similar effect of not having an elder male sibling is also estimated in Bhat & Zavier (2007)

born so that the region of birth of the children is same as the region of the current residence of their mother. However, if migration patterns differed systematically between high and low trade districts then the OLS estimates could be capturing the effects of migration due to trade liberalization. In particular, families with relatively more male children might have moved to areas which experienced greater industrialization due to tariff reduction with an expectation of greater child-labor avenues for male children in the industrial sector or with an expectation of better future earnings when the children join market work as adults. In that case the OLS estimates would be upward biased - capturing the effect of selective migration.

To address this issue I need information on the region of birth of the children, and hence history of the woman's earlier place of residence. However, in the DHS data there is no information about the district of past residence. The only relevant question asked is about the length of stay at the current place of residence. Since the survey year is 1999 and the liberalization policy was introduced in 1991, the concern is mainly about selective migration that took place between these years as a response to the policy change. So I re-estimate equation 1 over a subsample of interviewed ever-married females who have stayed in their current place of residence (i.e. in 1999) for more than 9 years. This implies, we are eliminating all those birth records where the mother might have migrated to (from) areas that experienced greater (lower) tariff cut. Table-7 column 1 provides the details for each cohort for the whole sample as well. The coefficients in this subsample, rid of migration effects, suggest that the OLS estimates without migration correction are not upward biased. In fact the estimates in the migration cleaned sample are higher on average.

8. Mechanisms

As noted before there are several channels through which trade policies could have affected household preferences. In particular, changes in poverty or standard of living in areas where industries lost tariff protection change in female wages or work force participation in response to increased competition or increasing demand for dowry payments in more industrialized areas might be the factors that affected parental preference. I present some evidence on the relative movement of these factors in areas of high tariff reduction compared to areas that experienced lower trade openness.

8.1. Poverty

Theoretically, the link between poverty and sex ratio suggests that households facing tighter income constraints might prefer to have sons than daughters where sons are more likely to participate in market work. Topalova (2005) and Topalova et al (2007), using National Sample Survey data from 1983-1997, finds that districts which were more exposed to trade reforms experienced smaller poverty reduction than the national average. So it is possible that in districts that experience a relatively greater decline in tariff, households are relatively more credit constrained and observe a relatively greater degree of son preference. However, on the empirical front there is mixed evidence regarding the relationship between poverty and son preference. Data from India suggests that the richest states of North India are also the ones that always had the lowest proportion of females in the 0-4 age group[15]. While macro level studies find an inverse correlation between poverty and relative female survival across states, household level investigation, by Agnihotri et al, also confirm a trend of more masculine child sex ratio as one ascends the income distribution curve. However, more careful analysis of the causal relationship between household income and son preference, by Qian (2007), suggests that total household income does not affect relative female survival. Drèze and Sen (1998) also conclude that there is little evidence that income has a significant relationship to gender differentials.

Although the relationship between poverty and son preference is weak or uncertain, I control for household standard of living index in the equation (1) and (2). DHS data reports whether a household is in the low, middle or high standard of living group. I include dummies for the low and high index groups. The results, outlined in column #2 of Table 4 and column # 3 of Table 5, show that the addition of controls for poverty does not alter the results in column (1). The coefficients are not significantly different from each other.

8.2. Wages

15 At the same time, Kerala, one of the poorest states in India, performs the best in terms of relative female survival. All related studies in India, from Sen (1990) to Bhat & Zavier (2007) acknowledge this notable feature.

If tariff declines are associated with changes in women's economic contribution then the value of female children to parents will also change. However, daughters do not contribute to their parental households once they are married[16] and since average age at marriage is about 18 years[17] for women, it is unlikely that changes in female labor market income will be an important factor in affecting parental preference for sons. In any case, I use wage data from National Sample Survey Rounds of 1987-88 and 1997 to in order to check how wages moved in response to tariff reduction[18]. If tariff reduction increased/decreased productivity of female labor, then we would expect to see an improvement/decline in female wages in areas of high tariff compared to areas of low tariff in years after the policy reform.

$$logwage_{idt} = \alpha_2 + Tariff_{dt}\ \beta_2 + X\ \varphi_2 + \gamma_d + \tau_t + \eta_{idt} \qquad \textbf{(4)}$$

Where, $logwage_{idt}$ represents log of wage of individual i in district d and time t. $Tariff_{dt}$ represents the district tariff rate in time t. X is a vector of household and individual controls like religion, caste and age of the individual. γ_d is a district fixed effect and τ_t is a year fixed effect. η_{idt} is an idiosyncratic error term. Since trade openness is expected to affect skilled and unskilled jobs differently, I further run the regression for subsamples of literates and illiterates. Table (9) shows the effect of tariff on wages. Results in column (1) imply that lower tariff is associated with higher wages for skilled male labor. However, I do not find any effect of tariff on wages of unskilled labor. On the other hand for female wages, I do not find any significant effect of tariff in either the skilled or the unskilled group. Thus, while the effect of tariff reduction did improve wages for skilled male workers, it did not change the returns to female employment. Overall, no change in returns to female labor coupled with the fact that daughters do not contribute to their

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16 Moreover, Foster & Rosengweig (2001) note "the common inference from these findings (effect of women's labor market outcome on 0-4 sex ratio) that mother's earnings measure the returns to the investments in their daughters neglects the fact that the daughters when adults do not reside in the same village as the mothers."

17 I calculated this using REDS 1999 survey for rural India

18 I could not include Round 1993-94 since it does not provide district identifiers. However, it does provide stratum which can be mapped to districts for some of the rural regions. Since my aim here is to look for any change in wages across the district and not just rural regions (the wage effect would work even if people migrate and work in urban regions temporarily or permanently) I show results without including Round 1993-93. However, when I do the regression only for the rural sample including 1993-94, I find similar results

parental households once they are married, it is less probable that tariff change would have changed parental preference away from having daughters.

8.3. Dowry

Anderson (2003) argues, in a analytical framework, that in a socially segregated economy like India where marriages are restricted within each group, modernization of the economy away from traditional to industrial might put an upward pressure on dowries (groom price). The mechanism is mainly driven by increasing within group inequality with the most skilled people from each group joining the industrial sector. This leads to an increasing demand for better quality grooms within each group, thus pushing up the groom price. Recent evidence from India suggests that there has been noticeable dowry inflation in the past few decades[19]. Most interestingly, apart from the general inflation in dowries, Rahman and Rao (2004) dispute the idea that the regional patterns (North versus South) in the existence of dowry identified by Dyson and Moore (1983) still hold. They note that in modern India, Southern brides are as likely to pay dowry (and pay as much) as Northern brides[20]

If regions with higher tariff reduction experience higher relatively dowry inflation then dowry might be a driving factor behind stronger parental preference for sons in these regions. If tariff reduction affected dowry payments then we would expect dowries to change for marriages that. In a regression similar to that of relative survival we can estimate how dowry payments changed in marriages that happened after the policy change, and hence exposed to the reform, compared to the marriages that took place before the change in policies, and hence unexposed to the reform. In other words, I compare dowries between marriages that happened strictly before and after the reform and also across districts with different intensity of tariff reduction. In particular, I estimate the following model.

$$\text{Dowry}_{mhdt} = \alpha_3 + (T_d * \text{Post}_{mh})\, \beta_3 + X\, \varphi_3 + \gamma_h + \tau_t + \upsilon_{mhdt} \qquad \textbf{(5)}$$

Where, Dowry_{mhdt} represents dowry paid (dowry received) in marriage m, of a daughter (son), of household h in district d and time t. T_d as before represents the percentage tariff

19 See Rao (1993), Srinivasan (2005)
2 Quoted from Pande & Astone, 2007

reduction in district d. $Post_{mt}$ is a dummy indicating whether marriage m in household h took place before or after 1992. If dowry increased with tariff reduction then we would expect β_3 to be positive. X is a vector of household controls. γ_h and τ_t are household and year of marriage fixed effects. υ_{mhdt} is an idiosyncratic error term.

Since DHS does not ask questions related to dowry, I use a different sample to do this estimation. I use the Rural Economic & Demographic Survey (REDS) longitudinal data set which tracks 4500 households, representative of India[21]. I use annual CPI data at the state level for industrial workers from Central Statistical Organization of India to adjust for nominal dowry inflation. Moreover since the practice of dowry as a marriage payment is traditionally prevalent amongst the Hindus and not Muslims, I do the analysis separately for the 2 religious groups. Lack of sufficient data doesn't allow me to restrict the sample to Muslims only; instead I take all other religion. However, Muslims comprise 60% of this group.

Table (10) presents the results from the estimation of equation (4). Column (1) shows that after controlling for years of education, year of marriage fixed effects and household fixed effects, real dowries have increased with higher trade openness amongst Hindus. However, the results cease to be significant after clustering at the district level, in column (1b). When I restrict to the sub-sample of high caste Hindus, who are more likely to practice dowry traditionally, I find significantly bigger estimates. The coefficients, shown in column 3a of Table 10, are precise even with this small subsample and even after clustering at the district level. The estimates of equation (4) for the non-Hindu subsample are presented in column 2a-2b. Tariff cut has had no significant impact on dowry payments in marriages of other religious groups (with or without clustering).

9. Heterogeneity of Impact

In the perspective of the above results, if dowry increase is responsible for a part of the parental discrimination towards daughters then we would expect relative female survival to decline more for Hindus compared to Muslims. Thus more insight into why trade policies might have worsened relative female survival could be obtained by examining its effect on different religious groups. In order to see if trade liberalization had differently

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21 REDS has 3 consecutive rounds between 1969-1971 followed by 1982 and finally 1999. However, the dowry questions are asked only in the 1999 round of REDS

affected these religious groups, Table 4 reports results from the interaction of the linear treatment variable with a set of dummy variables indicating a household's religion.

In column 1b-1c of Table-4, I compare the effect of tariff cut on the probability of having a male child in Hindu and Muslim households respectively. Strikingly, I find that the results are strongly significant for Hindus compared to all other religion but the relative probability of having a male child remains unaffected for the Muslim households. To see if this is true for all specifications discussed in sections 6 and 7 I rerun on the Hindu and Muslim sub population. For the unrestricted estimates the results are shown in Column (1a-1b) of Table 6. Once again, the results indicate a higher male bias for the cohorts exposed to reform but only in the Hindu subsample. The results are not significant for Muslims, indicating that the growing sex bias, as a fall out of the liberalization policies, was a particularly Hindu phenomenon[22]. As before the coefficients remain unaffected compared to the base specifications when standard of living indices are introduced as controls. The results for the subsample of Hindus are presented in column # 4 of Table 4 and column # 3 of Table 6 for the restricted and unrestricted estimations respectively. The results for the subsample of non migrants among Hindu households are presented in Table-4, column 6 for the base specification and in column 2 of Table 7 for the unrestricted specification. Again, the estimates are not statistically different from the estimates including both possible migrants and non-migrants.

I do a further check to see if the relative female survival in Hindu households responds to the sex selection situation discussed before. Table 4, column # 4, shows the estimates for those children who are do not have elder male siblings and hence are more likely to be discriminated against according to gender. They are in fact greater than the estimates in column 3, the base specification for Hindu households; implying that the effect of trade policies in reducing relative female survival chances is stronger when the child is in a sex selection situation. Table 6, column 2a, shows the effect in case of individual cohorts and again the coefficients are larger when the sample is restricted to children with no prior male sibling (compare with column 1a). Column 2b shows that the bias does not exist for Muslim children in a similar situation. This implies that children born to households with elder female siblings are more likely to be male in districts with higher tariff decline compared to their counterparts in districts with lower average tariff cut. It is in

22 For the other religions the data is too small to add any valuable insights.

accordance with our argument that parents observe a higher level of preference for sons in areas that face greater trade liberalization– parental selection of a child's gender is stronger in areas with high tariff cut we compare children who are in a sex selection situation.

Trade liberalization is expected to affect the entire population irrespective of religious groups though poverty and wage increases. However, dowry is known to be a particularly Hindu phenomenon (compared to Muslims) and hence is likely to affect son preference only for the Hindu households. The above results are suggestive of dowry being an important channel through which trade policies might have affected parental preference for sons.

10. Some further robustness check: Early Childhood Mortality

As mentioned earlier, the exact timing of the response in relative survival will depend on the nature of sex discrimination. The estimates in Table 5 & 6 indicates that the cohorts born close to but before the reform were also affected by the policy change and we assumed that it would have been due to higher childhood mortality amongst these cohorts. However, industrialization might affect relative female survival due to greater availability of sex selection technology in the more developed regions and hence only through greater pre natal selection without affecting the underlying preferences. In that case, the results for cohorts born before the policy change might actually be spurious. Thus it remains to be investigated whether trade reform actually led to higher childhood mortality of children, due, may be, to negligence.

Due to the higher mortality of males in infancy it is widely believed that the effect of sex discrimination due to negligence starts to show only in periods after infancy[23]. However, it is believed that the biological advantage, even though smaller, continues to persist in childhood so that under equal treatment females have higher chances of survival than male children. Thus any excess mortality of female children is considered to be an indication of discrimination. Using Census data between 1981 and 1991 Bhattacharyya (2006) finds that female disadvantage in child survival is strongest in the age range of 2 to 5.

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23 Female children are believed to have a relative biological advantage in survival due to better immunity systems

Thus I restrict my analysis to children who died after their first birthday. Table 8 shows the results from regression equation (3) for Hindu households, who are the ones affected by the policy change. The results in column (1) indicate that among children who died in the first few years of life, girls had a higher probability of death compared to boys in regions exposed to higher percentage tariff reduction. This suggests that growth in the availability of sex selection technology was not the only reason why the sex ratios diverged between regions with different degrees of trade openness. Negligence of female children was also one of the possible mechanisms driving the lower survival chances of female children due to the policy change.

11. Conclusion

The above analysis suggests that industrialization played a significant role in shaping or aggravating already existing parental preference over children's gender. The small previous literature addressing this question does not account for the endogeneity of industrialization. In this paper I exploit an exogenous shock to industrialization in India in 1991 to retrieve the causal link between industrialization and the growing female deficit. I find that trade liberalization reduced the survival chances of female relative to male children. Another interesting result that emerges from the analysis is the difference in the impact of trade liberalization on different religious groups. Specifically, while relative female survival changes significantly in the Hindu households, there is no change in the behavior of Muslims. One possible explanation is the effect of trade liberalization on Dowry inflation. The results indicate, there was a relative increase in real Dowries paid in Hindu marriages, while the coefficients were opposite in sign and insignificant for other religions. Thus although it is difficult to point to any specific cause that might have triggered a stronger son preference as a result of the industrialization process, supplementary evidence suggests that change in the relative value of women in the marriage market might have played an important role in determining the effect of industrialization. In contrast, economic factors probably were less instrumental in affecting relative female survival through trade openness.

Whatever the mechanism, it is likely that the continued process of industrialization in many of these traditional economies, like India or China, might aggravate the mortality differential between males and females. It is thus important to incorporate these indirect adverse effects while doing a cost benefit analysis and to design policies to offset the unwanted by products of trade liberalization.

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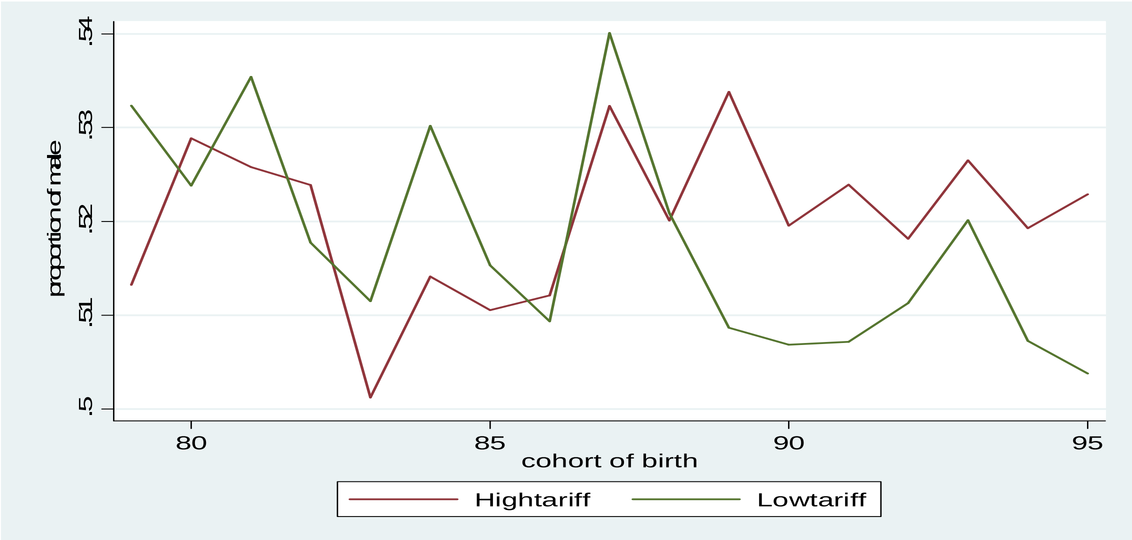
Fig 1



x-axis: different birth cohorts ranging from 1980 to 1995; y-axis: average proportion of male

High tariff decline = above median tariff cut, Low tariff decline = below median tariff cut

Table 1A

Dependent variable : Proportion of males in the 0-4 age group (1987-1999)				
	(1)	(2)	(3)	(4)
share of non-agriculture	0.082 **	0.08 **	0.079 **	0.073 *
	(0.037)	(0.038)	(0.038)	(0.039)
proportion of illiterate	-	-0.013	-0.020	-0.014
		(0.053)	(0.057)	(0.058)
proportion of backward social groups	-	-	0.020	0.021
			(0.042)	(0.042)
female labor participation	-	-	-	-0.015
				(0.022)
Number of obs	1089	1089	1089	1089
R^2	0.365	0.365	0.366	0.366

Note: All specifications include district and year fixed effect. Standard errors in parentheses are clustered at the district level.

TABLE 1B

Dependent Variable: Sex Ratio(Male/Female) in different age categories (1961 1991)		
	Sex ratio (0-5)	Sex ratio (5-9)
Non agricultire	0.052 *****	0.104 *******
	(0.029)	(0.038)
Schedule Caste	-0.004	-0.010*
	(0.004)	(0.006)
Urbanization	0.059	0.051
	(0.034)	(0.044)
Landless	-0.016	-0.002
	(0.023)	(0.030)
Literacy	-0.007	0.011
	(0.004)	(0.005)
Constant	-0.978	-0.9849
	(0.023)	(0.030)
R square	0.818	0.817
Obs	1266	1266

Note: Standard errors are clustered at the district level. Includes year and District FE.

Table 2: PANEL A

DHS Sample			
	Mean		
Schedule caste/tribe	0.327778		
Hindu	0.779675		
# of children	3.324748		
Proportion low SOLI	0.313471		
Proportion med SOLI	0.464961		
Household head characteristics			
Years of education	5.12395		
Industry of Occupation		**high tariff fall**	**low tariff fall**
Agriculture	0.420588	0.3589809	0.4850511
Education	0.077742	0.0877641	0.0672551
Industry	0.367172	0.416411	0.3156503
Menial Services	0.134498	0.1368439	0.1320435

Table 2: PANEL B

District Average Tariff Measures and its variation

	Net tariff in manufacturing sector	SD	Average decline in tariff	SD
before 91	87.43782	(6.33)		
1991	61.10069	(2.55)	0.298	(0.037)
1997	34.25083	(2.55)	0.439	(0.027)

Table 3A

Percentage growth rate per year in manufacturing employment, 1973-74 to 1997-98, by industry category:

	Food, beverages, tobacco	Petroleum products	Export oriented	Import competing	All manufacturing Industries
1973-89	1.41	5.26	-0.57	2.8	1.6
1990-97	2.56	6.09	3.36	2.67	3.09

Table 3B

Effect of Tariff Rates on proportion of employment in non-agriculture

	Employment in nonagri industry	Employment in nonagri occupation
Net Tariff	-0.004**	-0.003**
	(0.001)	(0.001)
constant	0.446***	0.503***
	(0.129)	(0.053)
District FE	Yes	Yes
Year FE	Yes	Yes
R-square	0.7396	0.803
Obs	1222	1222

Note: Standard errors in parentheses are clustered at the district level. The results imply a fall in tariff, or more trade openness increases the share of non agricultural employment.

Table 4:

	Dependent variable: Indicator whether observed child is male							
	1a	1b	1c	2	3	4	5	6
	Full Sample	Full Sample	Full Sample	Full Sample	Hindu	Hindu	Hindu	Hindu
							No elder male sibling	Non Migrant
ΔTariff*Young	0.205*			0.212*	0.284**	0.280**	0.349***	0.344***
	(0.11)			(0.11)	(0.12)	(0.12)	0.13	(0.12)
ΔTariff*Young *Hindu		0.0423**						
		(0.021)						
ΔTariff*Young *Muslim			-0.0218					
			(0.028)					
Controls for SOLI	No	No	No	Yes	No	Yes	No	No
Constant	0.522***	0.522***	0.522***	0.528***	0.528***	0.523***	0.512***	0.520***
	(0.003)	(0.003)	(0.003)	(0.0034)	(0.004)	(0.006)	(0.004)	(0.004)
Observations	91731	91630	91630	89801	71922	71157	41106	59685
R-squared	0.01	0.01	0.01	0.01	0.01	0.01	0.01	0.01

*** p<0.01, ** p<0.05, * p<0.1 (convention in all specifications)
Note: Sample consists of all children reported to be living by interviewed mother. ΔTariff is the percentage change in tariff between pre (1989) and post reform (1991) tariff rates. All regressions have control for SCST. Standard errors are clustered at the district level. Regression includes District and Time fixed effects.

Table 5: Unrestricted estimates

Dependent variable: indicator whether observed child is male						
Born in year	Whole Sample		No elder male sib		Whole sample	
1981	0.151	(0.28)	-0.0554	(0.48)	0.144	(0.28)
1982	0.229	(0.30)	0.399	(0.41)	0.291	(0.30)
1983	0.119	(0.33)	0.112	(0.51)	0.134	(0.33)
1984	0.203	(0.31)	0.290	(0.49)	0.235	(0.31)
1985	0.366	(0.26)	0.429	(0.42)	0.370	(0.26)
1986	0.344	(0.24)	0.199	(0.40)	0.363	(0.24)
1987	0.389	(0.27)	0.297	(0.39)	0.427	(0.27)
1988	0.435 *****	(0.25)	0.430	(0.44)	0.445 *****	(0.25)
1989	0.576 ******	(0.25)	0.457	(0.41)	0.584 ******	(0.25)
1990	0.379	(0.30)	0.319	(0.53)	0.423	(0.30)
1991	0.319	(0.27)	0.767 *****	(0.41)	0.361	(0.27)
1992	0.199	(0.24)	0.0401	(0.39)	0.211	(0.24)
1993	0.611 *******	(0.22)	0.377	(0.36)	0.619 *******	(0.22)
1994	0.413 *****	(0.25)	0.556	(0.38)	0.417 *****	(0.24)
1995	0.599 ******	(0.24)	0.476	(0.33)	0.598 ******	(0.24)
1996	0.0699	(0.28)	0.172	(0.40)	0.0934	(0.28)
1997	0.147	(0.25)	0.287	(0.42)	0.137	(0.25)
1998	0.316	(0.24)	0.0432	(0.43)	0.330	(0.24)
1999	0.163	(0.36)	1.410 *****	(0.78)	0.208	(0.36)
dshsli1					0.00183	(0.0036)
dshsli3					0.00486	(0.0051)
SC/ST					-0.00260	(0.0035)
Constant	0.474 ***	(0.11)	0.0952	(0.23)	0.457 ***	(0.11)
Observations	106803		59284		105662	
R-squared	0.01		0.01		0.01	

All specifications include district and cohort fixed effect. Standard errors are clustered at the district level.

Table 6: Unrestricted estimates for religious sub-samples

Dependent variable: Whether Observed child is male					
	1a	1b	2a	2b	3
Born in year	Hindu	Muslim	Hindu	Muslim	Hindu
			No elder male sib	No elder male sib	
1981	0.213	-1.619	0.106	-3.132*	0.217
	(0.34)	(1.11)	(0.40)	(1.62)	(0.34)
1982	0.552	-0.598	0.436	-2.845	0.575
	(0.43)	(1.37)	(0.46)	(1.97)	(0.43)
1983	0.217	0.599	0.421	-0.651	0.179
	(0.37)	(1.47)	(0.44)	(1.80)	(0.38)
1984	0.386	-1.662	0.476	-1.808	0.426
	(0.35)	(1.17)	(0.45)	(1.86)	(0.35)
1985	0.482	0.213	0.569	-0.921	0.486
	(0.31)	(1.28)	(0.37)	(1.55)	(0.31)
1986	0.405	0.806	0.213	-0.410	0.421
	(0.28)	(1.12)	(0.37)	(1.67)	(0.27)
1987	0.578 **	-1.514	0.651 *	-1.032	0.612 **
	(0.29)	(1.45)	(0.36)	(1.82)	(0.29)
1988	0.641 **	-0.418	1.102 ***	-2.954	0.666 **
	(0.27)	(1.23)	(0.35)	(1.84)	(0.27)
1989	0.949 ***	-0.465	0.811 **	-2.084	0.917 ***
	(0.27)	(1.07)	(0.39)	(1.77)	(0.27)
1990	0.725 **	-1.574	0.907 **	-2.904 *	0.774 **
	(0.33)	(1.18)	(0.41)	(1.54)	(0.33)
1991	0.491 *	-0.438	0.627 *	-2.632	0.529 *
	(0.30)	(1.15)	(0.36)	(1.71)	(0.29)
1992	0.260	-0.147	0.230	-0.128	0.257
	(0.28)	(1.09)	(0.33)	(1.63)	(0.28)
1993	0.583 **	1.291	0.409	0.794	0.581 **
	(0.26)	(1.00)	(0.33)	(1.58)	(0.26)
1994	0.672 ***	-0.0766	0.624 *	-1.982	0.703 ***
	(0.26)	(1.08)	(0.34)	(1.64)	(0.26)
1995	0.771 ***	-0.308	1.021 ***	-2.187	0.774 ***
	(0.29)	(1.04)	(0.34)	(1.65)	(0.29)
1996	0.310	-1.015	0.247	-3.031 *	0.282
	(0.32)	(1.07)	(0.41)	(1.57)	(0.32)
1997	0.274	0.0772	0.0243	-1.768	0.275
	(0.27)	(1.28)	(0.34)	(1.66)	(0.27)
1998	0.334	0.701	0.446	-1.720	0.346
	(0.28)	(1.10)	(0.34)	(1.76)	(0.28)
1999	0.501	0.281	1.554**	-0.870	0.548
	(0.45)	(1.78)	(0.70)	(2.64)	(0.44)
dshsli1					0.00102
					(0.0041)
dshsli3					0.00544
					(0.0059)
scst			0.00204	-0.0237	-0.000447
			(0.0050)	(0.020)	(0.0040)
Constant	0.468***	1.082***	0.495***	1.529***	0.364***
	(0.098)	(0.33)	(0.12)	(0.49)	(0.13)
Observations	86029	10098	48158	4730	84395
R-squared	0.01	0.03	0.01	0.06	0.01

Note: Standard errors are clustered at the district level. All regressions have district and cohort of birth FE

Table 7: Migration

Dependent variable: Whether Observed child is male		
	1	2
Born in year	All religion Non Migrant	Hindu Non Migrant
1981	0.289	0.341
	(0.29)	(0.34)
1982	0.228	0.547
	(0.31)	(0.34)
1983	0.0509	0.0653
	(0.36)	(0.40)
1984	0.186	0.327
	(0.32)	(0.38)
1985	0.373	0.436
	(0.29)	(0.34)
1986	0.390	0.440
	(0.25)	(0.29)
1987	0.301	0.477
	(0.27)	(0.31)
1988	0.455 *	0.686 **
	(0.28)	(0.30)
1989	0.705 **	1.047 ***
	(0.27)	(0.30)
1990	0.587 *	0.967 ***
	(0.33)	(0.37)
1991	0.186	0.378
	(0.28)	(0.30)
1992	0.190	0.246
	(0.26)	(0.30)
1993	0.741 ***	0.675 **
	(0.26)	(0.30)
1994	0.549 *	0.746 **
	(0.31)	(0.31)
1995	0.421	0.599
	(0.32)	(0.38)
1996	0.00519	0.132
	(0.32)	(0.37)
1997	0.0285	0.207
	(0.33)	(0.41)
1998	0.526*	0.564
	(0.30)	(0.35)
1999	0.263	0.597
	(0.54)	(0.72)
SC/ST	-0.00446	0.000239
	(0.0042)	(0.0047)
Constant	0.458***	0.335
	(0.085)	(0.21)
Observations	81223	64951
R-squared	0.01	0.01

Note: Standard errors are clustered at the district level. All regressions have district and cohort of birth FE

Table 8: Relative Mortality of female children in early ages

Dependent variable : D(female=1)		
	1	2
DelTariff*Young	1.224 *****	1.229 *****
	(0.68)	(0.68)
Stndrd living (low)		-0.0432 *****
		(0.024)
Stndrd living (high)		-0.0768
		(0.059)
SC/ST (low caste)	-0.0166	-0.00772
	(0.029)	(0.031)
Constant	0.544***	0.562***
	(0.024)	(0.025)
Observations	2835	2822
R-squared	0.14	0.14

Note: Sample of children who died between their 1st and 5th birthday. Analysis restricted to Hindu households. The control group is comprised of children who were born before 1987. Note: Standard errors are clustered at the district level. All regressions have district and cohort of birth FE

Table 9: Wages

Dependent Variable: log of real wage				
	male		female	
	Literate	Illiterate	Literate	Illiterate
tariff	-0.0119 ******	-0.00190	-0.00160	0.00203
	(0.0057)	(0.0057)	(0.0077)	(0.0038)
age	0.0334***	0.00130**	0.0292***	0.00139**
	(0.00091)	(0.00060)	(0.0019)	(0.00061)
scst	-0.315***	-0.0615***	-0.283***	0.0108
	(0.021)	(0.016)	(0.051)	(0.019)
Constant	-0.594***	-0.533***	-1.065***	-1.264***
	(0.21)	(0.19)	(0.28)	(0.14)
Observations	54546	17725	9658	17503
R-squared	0.35	0.28	0.38	0.27

Robust standard errors in parentheses, Sample contains individuals in the age range 16<age<60. All regressions include control for religion, year and district fixed effect. *** p<0.01, ** p<0.05, * p<0.1

Table 10: Dowry

	Dependent variable: Real Dowry					
	Hindu	Hindu	Other Religion	Other Religion	Hindu-Upper caste	Hindu-Other castes
	1a	1b	2a	2b	3a	3b
DelTariff*Post	210.7 **	210.7	-89.01	-89.01	727.3 *	191.5
	(102)	(139)	(144)	(391)	(424)	(142)
Household FE	Yes	Yes	Yes	Yes	Yes	Yes
Year of Marriage FE	Yes	Yes	Yes	Yes	Yes	Yes
Constant	58.23***	58.23***	140.4***	140.4***	77.58***	56.66***
	(3.67)	(3.45)	(13.7)	(20.8)	(13.4)	(3.64)
Obs	3365	3365	403	403	235	3130
R-squared	0.80	0.80	0.90	0.90	0.94	0.79
cluster		yes		yes	yes	yes

Note: Standard errors are in parenthesis. Post indicates years of marriage 1992 and up. All regressions control for household and year of marriage fixed effects.

Appendix I:

Consider a household as a single family decision maker. Suppose y_0 is the household's income when it has no child and y_d is the household's net income when they have a daughter. y_d is net of direct dowry cost of marrying the daughter, $y_d = y_0 - D$. Suppose the family associates an positive emotional attachment α to having a daughter.
The family's decision to have a daughter or not would be then depend upon whether the utility from having a girl child is higher than the utility from not having a child. i.e. the parents would have a daughter if

$$u(y_d, D) + \alpha + e_d \geq u(y_0, 0) + e_0 \qquad (1)$$

where e_k, $k \in \{d, 0\}$, is an additively separable, mean zero, i.i.d,stochastic term.
The weight attached to a daughter depends on the social environment, which changes alongside other social indicators with socio-economic development and modernization.

The utility from having a daughter is then,

$$u(y_d, D) = v(y_0 - D, p) + \alpha \qquad (2)$$

where v(.) is the indirect utility associated with income y_d at the vector of consumer prices p.

Then the probability that a daughter is born to a family is :

$$\Pr(d = 1) = \Pr(v(y_0 - D, p) + \alpha + e_d \geq v(y_0, p) + e_0)$$

$$= \Pr(e_0 - e_d \leq v(y_0 - D, p) + \alpha - v(y_0, p)) \qquad (3)$$

Define $u = e_0 - e_d$ which is mean zero with cdf *F(u)* and strictly positive density *f(u)*.

Equation (2) can be written as:

$$\Pr(d = 1) = F\big(v(y_0 - D, p) + \alpha - v(y_0, p)\big) \qquad (4)$$

To analyze the determinants of changes in preference for a daughter, totally differentiate equation (4).

$$dPr(d = 1) = f(u)\{\left[\frac{\partial v_d}{\partial y} - \frac{\partial v_0}{\partial y}\right] dy_0 - \frac{\partial v_d}{\partial y} dD + d\alpha\} \qquad (5)$$

where $v_d = v(y_0 - D, p)\, and\ v_0 = (y_0, p)$. For simplicity, I assume that p is a constant unit vector. Thus a tariff decline (*dt*) enhancing the non-agricultural sector influences the preference for daughters through changes in standards of living i.e. increase in family income. Greater employment opportunities outside the traditional sector might affect the dowry, groom prices, as has been documented by Anderson (2003). Opening up of the economy to trade and improving standards of living might lead to improvement in social values, increasing the preference for daughters in a formerly traditional society. Thus we can rewrite (5) to incorporate the tariff decline (*dt*):

$$dPr(d=1) = f(u)\left\{\left[\frac{\partial v_d}{\partial y} - \frac{\partial v_0}{\partial y}\right]\frac{\partial y_0}{\partial t}dt - \frac{\partial v_d}{\partial y}\frac{\partial D}{\partial t}dt + \frac{\partial \alpha}{\partial t}dt\right\} \qquad (6)$$

Thus preference for daughters can be affected by any of the following in the face of declining trade protection:

(i) Since diminishing marginal utility of income implies $((\partial v_{d})/(\partial y))>((\partial v_0)/(\partial y))>0$,if tariff declines affects living standards, then it would affect parental preference for daughters to the extent that income affects the ability of parents to provide nutrition. It could also affect preference for daughters in a traditional society where economic status dictates the preference for daughters - commonly called the "sanskritization effect".

(ii) If declining tariff and hence higher competition in the non agricultural sector leads to greater employment opportunities outside the traditional sector, then income inequalities might lead to increasing Dowry payments in a caste segregated society, Anderson(2003). This in turn will reduce the preference for daughters.

(iii) If parents associates a certain value to having a female child, then such parameters might be affected by socio economic changes that are brought about by a more competitive and modernizing economy. Dasgupta (2007) discusses how the economic growth in Korea led to a declining son preference as a result of improving socio economic indicators.

The role of trade liberalization, and the consequent growth of nontraditional sector , in influencing the parental preference for daughters is thus an empirical issue.

Appendix 2:

Table 1:

Growth in Employment – percent per annum

	1980-1989	1988-1997
Consumer goods	0.15	4.35
Intermediate Goods	-1.38	6.54
Capital Goods	0.75	1.40

Source: Pandey, 2004