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ABSTRACT

The Effect of Trade Liberalization on Marriage and Fertility: Evidence from Indian Census

Using a district-level panel constructed from five waves of decennial Indian censuses covering 1971-2011, we examine the medium-term (1991-2001) and long-term (1991-2011) impacts of the 1991 Indian trade liberalization on marriage and fertility rates among young women aged 15-34 years. We exploit the fact that countrywide tariff reductions varied across industries creating exogenous local labor market shocks based on the initial employment composition of the district. We find heterogeneous results across urban and rural areas. We find that urban areas of the districts that experienced larger tariff cuts experienced relative increase in marriage rate compared to the districts that experienced smaller tariff cuts. Moreover, tariff cuts positively affect the workforce participation among both young men and women in urban areas. However, there is no impact of tariff cuts on marriage rate or workforce participation among young for rural areas. In contrast, tariff cuts reduced fertility rate mostly in rural areas.

JEL Classification: J12, J13, O12
Keywords: marriage, fertility, trade liberalization

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1 Introduction

Nationwide exogenous economic shocks may have serious implications for firm and household dynamics, and a growing literature is probing the evolution of marriage and fertility patterns in both developed and developing countries, resulting from local labor demand shocks. Recent contributors to this literature include Black et al. (2013) and Kearney and Wilson (2018), who examine the impact of transitory local economic shocks stemming from energy booms and bust cycles in the United States on marriage and fertility. Braga (2018) exploits the Brazilian trade liberalization, a one-time event (permanent), to look at its impact on family formation decisions.

Traditionally, gains from marriage came from household specialization with women taking up domestic responsibilities, while men engaging in income earning activities by selling their labor in the market (Becker, 1981; Stevenson and Wolfers, 2007). Changes in the economic prospects of men and women emanating from exogenous economic shocks may impact their value in the marriage market. Wilson and Neckerman (1987) posited that declining economic prospects of less educated men made them "less marriageable". Autor, Dorn, and Hanson (2019) support these claims using manufacturing declines induced by trade shocks in the US. In the Indian context, Anderson (2003) suggests that better employment opportunities for men (women) may increase within-caste income inequality which in turn causes dowries to increase (decrease), thus, reducing (increasing) the value of women. In a similar vein, permanent changes in income, coupled with the expenditures associated with raising a child may affect fertility decisions (Becker, 1960; Willis, 1973; Kugler and Kumar, 2017).

In this paper, we test the above-mentioned predictions by exploiting the trade-induced local labor demand shocks stemming from India's unilateral tariff reductions during the 1990s as a natural experiment for women aged 15-34 years, across rural and urban areas of the districts, tracing out the effects over medium- and long-term. As discussed in next section, the Indian trade liberalization in 1991 was mostly unanticipated and different industries experienced differential tariff cuts over 1991-1997. The impact of these cross-industry tariff

cuts varied across districts based on the pre-liberalization industrial composition of the districts. We exploit the variation in reductions in effective tariff protection experienced by districts that stem from differential cuts across industries and differential pre-liberalization industrial composition of the workforce to identify the impact of trade liberalization on marriage and fertility outcomes of young women. Similar identification strategies are used in other contexts to identify the impact of trade liberalization (e.g., Topalova, 2007, 2010; Edmonds, Pavcnik, and Topalova, 2010).[1] The Indian trade liberalization was similar to a once-and-for-all event, with tariffs being reduced between 1991 and 1997, and remaining more or less constant thereafter. This allows us to empirically examine the evolution of marriage and fertility rates as a result of the trade shock. By comparing the variation in reduction in tariff protection across districts, we are able to measure the relative effect of trade liberalization on the change in marriage and fertility rates in districts facing relatively larger tariff cuts to those in districts facing smaller tariff cuts over time, while controlling for pre-reform trends in these outcomes. We further explore potential mechanisms through which the effects of trade liberalization on family decisions operate.

Previous research has exploited the exogenous nature of Indian trade liberalization to study the impact of trade liberalization on poverty (Topalova, 2007, 2010), firm-level productivity (Topalova and Khandelwal, 2011), unemployment (Hasan et al., 2012), industry wage premium (Kumar and Mishra, 2008), children's schooling outcomes (Edmonds, Pavcnik, and Topalova, 2010), and sex-ratio (Chakraborty, 2015). However, the issue of family formation has drawn considerably less attention in the Indian context. In contrast, there exists a considerable literature in developed and developing countries that look at the impact of economic shocks on family formation. For example, Schaller (2016) shows that improvement in men's (women's) labor market conditions are associated with increase (decrease) in fertility in the US. Black et al. (2013) examine the effect of increase in the world energy

[1]We use Indian districts as local labor markets due to very low rate of mobility across districts, and several papers (Rosenzweig and Evenson, 1977; Duflo and Pande, 2007; Topalova, 2010; Edmonds, Pavcnik, and Topalova, 2010) have used Indian districts as local labor markets.

prices during the mid-70s, which also increased men's earnings in the Appalachian region in the US, on fertility. They find a positive relationship between men's earnings and fertility, thus, labeling children as normal goods. Using a similar notion, Kearney and Wilson (2018) exploit the positive economic shock arising from the fracking boom of the 2000s in the US to observe its impact on marriage and fertility. They find that fertility increases in response to the fracking boom, but there is no impact on marriage. Kis-Katos, Pieters, and Sparrow (2018) examine the gender-specific effects of Indonesia's second wave of trade liberalization in the 1990s by looking at the impact of regional exposure to tariff on the district-level marriage and employment outcomes for men and women. They find positive effects of the trade reforms on women employment which led to a reduction in the marriage rates since the opportunity cost of performing domestic duties increases.

Keller and Utar (2022) find increased import competition from China, following the lifting of the Multi-Fibre Arrangement (MFA) quotas on Chinese exports to the European Union, led to an increase in the earnings gap between men and women in Denmark. They find that single workers exposed to import competition from China have a higher likelihood of getting married and having children. Autor, Dorn, and Hanson (2019) exploit gender-specific labor demand shocks stemming from increased international manufacturing competition in the US and find that the trade-induced labor demand shocks deferentially reduce employment and earnings of young males, thereby reducing marriage and fertility.

For the Indian context, we are aware of only one paper that looks at the impact of Indian trade liberalization on probability of birth. Using the retrospective birth histories from the surveyed married women in age 15-44 from a single cross section data collected in 2002-04, Anukriti and Kumler (2019) examine the impact of Indian trade liberalization on the probability of child birth in a given year, and the probability that these births are male. They do not look at the marriage outcomes. Our paper contributes to the existing literature in the following ways. First, ours is the first paper to examine the impact of Indian trade liberalization on the marriage outcomes of young women to the best of our knowledge. We

also explore the mechanism through which trade liberalization may have an impact on family formation outcomes. Second, unlike the majority of the existing literature on Indian trade liberalization that use survey data and focus on rural areas only, we use census data that allows us to look at urban and rural areas separately in addition to looking at for all areas.[2] The analysis of rural and urban areas separately brings out some contrasting differences in impacts of trade liberalization. Additionally, the census data allows us to look at workforce participation rate among young men and women separately. Third, though our analysis on fertility is closely related to Anukriti and Kumler (2019), our fertility analysis differs from theirs in several ways. First, our district fertility measure is based on children ever born (CEB) to women living in the district that captures the number of children born alive to women up to that point in time, whereas Anukriti and Kumler (2019) fertility measure is the probability that a woman living in the district gives birth in that year. As noted by the authors, a higher probability of birth does not necessarily imply higher completed fertility and their results may be capturing changes in the timing of births rather than changes in overall fertility levels. Second, while their paper focuses on rural areas only, we look at the fertility outcomes in both rural and urban areas, separately besides for all areas combined. Third, while Anukriti and Kumler (2019) focus on short-term impacts, we capture both medium and long terms impacts.[3]

Our findings are the following. We find heterogeneous impacts of trade liberalization across urban and rural areas. We find that the tariff cuts positively affected the marriage rates among age 15-34 women in urban areas. Urban areas of the districts that experienced relatively larger tariff cuts experienced increase in marriage rate among 15-34 age group

[2]Azam (2022) also uses census data to examine the impact of Indian liberalization on human capital accumulation.

[3]Anukriti and Kumler (2019) use yearly tariff between 1987 to 1997 to explain the yearly probability of birth given by a woman. Goldberg and Pavcnik (2007) note that the research on liberalization has mainly focused on short run or medium run impacts due to the inconsistent designs of household surveys. Census data and the fact that the Indian trade liberalization was a discrete shock where tariffs were reduced dramatically between 1991-1997 allows us to look at long term impacts. As discussed in Autor et al. (2014) and Dix-Carneiro and Kovak (2017), studying shocks that continually evolve over time is much more challenging.

women compared to urban areas of districts that experienced lower reduction in tariff in both medium and long term. Additionally, longer-term impact, which captures the cumulative effect, is stronger on marriage outcome than the medium-term impact. However, we cannot statistically rule out zero impact of tariff cuts on marriage rates in rural areas. We also find that districts that were more exposed to tariff cuts saw a relatively larger reduction in fertility rate. However, the negative impact of tariff cut on fertility is mostly driven by rural areas, while we cannot rule out zero impact for urban areas in long-run. Further exploring the mechanism, we find that that tariff cuts have positive impact on workforce participation of young individuals (15-34) of both genders in urban areas, and longer-term impact seems stronger than the medium term. We do not find evidence of effect on workforce participation for rural areas. While our estimation strategy account for pre-reforms trends in outcomes, we also implement a separate placebo test to show that the pre-reforms trends (1981-1991) in marriage, fertility, and workforce participation rates were not correlated with future (1991-1997) tariff reductions. This supports the causal interpretation of the impact of trade liberalization on the outcomes of interest.

The rest of the paper proceeds as follows. In the next section, we provide a brief background on the Indian Trade Liberalization. Section 3 outlines the empirical strategy. Section 4 discusses the data sources and construction of the variables used for the analysis. Section 5 summarizes the main findings, and Section 6 concludes.

2 Indian Trade Liberalization

The Indian economy adopted a highly regulated development strategy post-independence, and embarked on the Five-Year Plans, developed, implemented and supervised by the Planning Commission. Being a developing country, India chose to follow a policy of protectionism through high nominal tariffs, complex industrial licensing requirements, and extensive non-trade barriers in order to develop its domestic industrial sector, and by the end of 1980s, only

12% of manufactured products were allowed to be imported under an open general license (Cerra and Saxena, 2002). To integrate with the global economy, India turned to export-led growth strategy in a phased manner during the Seventh Five-Year Plan (1985-1990). However, the average tariff was still very high in the early 1990s averaging more than 90%, one of the highest in Asia. Several factors like, rising oil prices, drop in remittances from Indian workers in the Middle East, lack in the confidence of investors due to weakened credibility, lead to a major balance of payments crisis in 1991 (Topalova, 2010). India turned to the IMF for bailout and was asked to introduce macroeconomic stabilization policies and structural reforms as a condition for help with the external payments. One such major reform was a reduction in tariff levels, with the maximum average tariff reducing to about 47% by 1997, and a removal of quantitative restrictions on imported inputs and capital goods for export production (Chopra et al., 1995). Prior to the trade liberalization, ad-valorem tariffs averaged over 80% in India making international trade extremely restrictive. The tariff reduction was an exogenous trade shock to the economy as it was not initially a part of the Five-Year Plans, leaving less room for political manipulations.

In his seminal paper, Trefler (1993) posited that level of trade protection is endogenous since increased import competition is accompanied by an increase in lobbying by private domestic industries. The Indian trade liberalization provides a perfect setting to study the economic and social impacts of trade openness. Some features that contribute toward the identification strategy and mitigate potential endogeneity concerns are as follows. The tariff reductions across products were not systematic or planned. The timing and implementation of the level of tariff cuts across industries varied significantly (Topalova, 2010). Since the reforms were unannounced, there was minimal possibility of political influence or lobbying by private domestic interests and adjustment in consumption or production decisions. Topalova and Khandelwal (2011) find uniform movements in tariffs until 1997 that were uncorrelated with productivity, after which the tariff movements were non-uniform suggesting selective manipulation to protect less efficient industries. However, one potential shortcoming of

analyzing outcomes based solely on tariff changes is that they may be correlated with non-tariff barriers (NTBs). Though NTBs, such as, import licensing were crucial to the Indian trade policy historically, they were gradually removed during the 1990s, and about 64% of the imports were free of import licenses by 1997 (Nouroz, 2001). Tariffs are easier to measure across industries and time, and NTBs data are not available at a detailed industry level. Hence, we do not include non-trade barriers as a measure of protection in our analysis.[4]

3 Empirical Strategy

3.1 District level reduction in tariff protection

Our empirical strategy is based on two sources of exogenous variation. First, although tariffs were reduced nationally, there was substantial variation across industries. Second, the impacts of those tariff reductions were non-uniform across districts based on their pre-reform distribution of workers across industries. Similar identification strategy has been used in the context of India by others (e.g., Topalova, 2010; Edmonds, Pavcnik, and Topalova, 2010; Topalova and Khandelwal, 2011; Hasan et al., 2012). The tariff exposure measure is created following Dix-Carneiro and Kovak (2017) who define the tariff reductions as a source of regional labor demand shock which are calculated using long-differences in the log of 1 plus the average industry-level tariff from 1991 to 1997. We define these local trade shocks as:

$$TR_d = \sum_i \beta_{di} \Delta ln(1 + \tau_i) \tag{1}$$

$$\beta_{di} = \frac{\gamma_{di} \frac{1}{\theta_i}}{\sum_j \gamma_{dj} \frac{1}{\theta_j}}$$

[4]See Edmonds, Pavcnik, and Topalova (2010) for more details. They argue that trade volumes increased in spite of the existence of NTBs in 1997, which suggests a positive correlation between declining tariffs and NTBs.

where tariff reduction, TR_d, is the effective reduction in tariff protection faced by district d, τ_i is the nominal tariff for industry i. γ_{di} is the total workers employed in industry i in district d; θ_i is the cost share of nonlabor factors. TR_d is the long-difference in tariffs from 1991-1997, and hence a positive change implies tariff cuts. Thus, our tariff measure is interpreted as a weighted industry tariff change at district-level that captures the effect of trade openness on labor demand in district d.[5] The cross-sectional variation in the trade exposure can be attributed to the 1991 employment shares that act as weights for the industry-level tariff changes. The tariff measure excludes the non-tradable industries since non-tradable output is consumed within the districts where they are produced indicating that their prices move together with those of the tradable sector goods produced locally (Kovak, 2013). Based on this intuition, we construct our tariff measure based on tradable industries.[6]

3.2 Difference-in-Differences

To measure the medium- and long-term effects of trade liberalization on marriage and fertility outcomes, we use the following specification:

$$Y_{d,t} - Y_{d,1991} = \alpha + \delta_t TR_d + \eta_s + \Gamma_t(Y_{d,1991} - Y_{d,1981}) + \varepsilon_{dt} \tag{2}$$

where $Y_{d,t}$ is the district-level outcome of interest in district d at time t ($t = 2001$ or 2011). $Y_{d,1991}$ is the initial district-level outcome in district d in 1991. TR_d captures the reduction in tariff protection enjoyed by district. First difference accounts for time-invariant unob-

[5]The trade shock measure is similar in notion to the Bartik (1991) labor demand shocks that predict changes in local labor demand constructed by interacting a region's initial industrial structure with national changes in industry-specific employment and wages, except that now the source of the shock is known providing a better source of exogenous variation (Dix-Carneiro, Soares, and Ulyssea, 2018). Topalova (2010) uses a tariff measure similar to that of Bartik (1991) where she does not consider the labor demand elasticity represented by the cost of non-labor factors (Kovak, 2013).

[6]Cereal/oilseed production which is tradable did not see any change in policy. As a result, the change in tariff is zero while its employment share is used in the calculation. Topalova (2007, 2010) remove cereal and oilseed from tradable industries arguing that those were canalized through government. Given the large agriculture share in employment and particularly in cereal production, our weights are closer to the weights used by Topalova "unscaled measures" where she takes all employment into account and non-traded (including cereals/oilseeds) industries are assigned zero tariffs. She also constructs weights called "scaled measure" using employment from tradable industries only excluding the cereal and oilseeds.

served districts heterogeneity. Since different states follow different policies, to account for differential state-level policies we control for state fixed effects denoted by η_s, hence, our δ_t is identified from variation in TR_d across districts within the same state.[7]

To address the possibility of omitted variables being correlated with the TR_d and outcome of interest, we also control for the pre-reforms changes in the outcome variable denoted by $(Y_{d,1991} - Y_{d,1981})$. The pre-reform change should capture any time variant district observed or unobserved characteristics that drive the change in the outcome. While the change in the outcome varies across years, the liberalization measure reflects the district-level reduction in tariff protection from 1991-1997. All the regressions are weighted by the average population of females in relevant age groups in the concerned years. Standard errors are clustered at the state-region level which are contiguous districts with similar socioeconomic characteristics to avoid spatial correlation across neighboring districts.[8] Recall that our TR_d measure is the change between 1991 and 1997, hence a positive coefficient on δ_t will be interpreted as positive effect of tariff cuts or liberalization. Our setup captures the effect of reduction in tariff protection by comparing districts that experienced a larger tariff reduction vs. districts that experienced relatively smaller tariff reduction. Migration from districts more affected by liberalization to those less affected potentially could dampen the effect, however, as discussed earlier the inter-district migration remains very low in India.

A potential issue with controlling for pre-reform trends in Equation (2) is that presence of $Y_{d,1991}$ introduces a mechanical correlation between $(Y_{d,1991} - Y_{d,1981})$ and error term (Nickell, 1981). Hence, we also instrument the pre-existing trend variable $(Y_{d,1991} - Y_{d,1981})$ with the lags (1981 and/or 1971) of different variables for different outcomes (discussed in results section).

[7]State fixed effects in first-differenced model is equivalent to state-specific trends in analysis with levels. We control for 18 major states that include 16 major states covering 97.2 percent of Indian population. Smaller Northeastern states are combined together. The Union Territories (each having 1-3 districts in the 1991 Census) are combined with the neighboring states following planning commission of India's poverty line assignment where they assign neighboring state poverty line to nearby Union Territory.

[8]The state-region is constructed by National Sample Survey (NSS) by grouping similar contiguous districts together. Our sample consists 74 state-regions based on the NSS 1987 definition.

4 Data

Our main data source is five waves of decennial census collected by Census of India. We construct a district-level panel using the decennial census covering four decades from 1971-2011. A district is an administrative division of a state or territory in India. Overtime new districts are created either by dividing or reorganizing existing districts. We trace the 2001 and 2011 districts back to the 1991 Census districts by combining the broken-up districts, and our full sample comprises of 452 1991 Census districts. There were 462 districts in 1991 Census, but the 1991 Census did not cover the state of Jammu and Kashmir (14 districts).[9] In order to disentangle the effects across rural and urban areas, we also carry out the analysis by considering the rural and urban areas separately.[10]

We use the Indian Census data to construct district level outcome variables: marriage rate, fertility rate, and workforce participation rate. Since our interest lies in the impact of liberalization on family formation, we focus on young women aged 15-34 years, since these women have made family formation decisions within the past 10 years. Marriage rate is calculated as the ratio of 'married females' to 'total population of females' in age-group 15-34.[11] For the fertility rate, we take the ratio of 'children ever born' to 'total ever-married women'.[12] For pre-reform industrial composition, we use district-level employment compo-

[9]There were 640 districts in 2011 Census and 593 districts in 2001 Census. The 1981 Census which we use to control for pre-reforms trends did not collect data from the state of Assam (23 1991 districts). The data for 1991, 2001, and 2011 Census is assembled from digital tables provided by Census of India website. The 1971, 1981 Census is provided by Reeve and Barnes (2000) which is supplemented by Census of India publication GOI (1986). Few districts were broken up between 1971 and 1991, we construct the 1991 districts by assuming that all the absolute counts were divided in same proportion as total population where the population percentages going to breakaway districts is used from Kumar and Somanathan (2009).

[10]Few districts are completely rural, while few are completely urban. Hence, the number of districts is different for the rural and urban analysis compared to the combined analysis.

[11]Indian Census provides aggregate information based on age groups. The age groups provided in data are: 15-19, 20-24, 25-29, 30-34. We do not consider ever-married women as it may include women who are divorced, separated or widowed in addition to married women.

[12]To construct the fertility outcome, we use children ever born to measure fertility as it captures total completed fertility of a woman up to that age. We use total population of women to calculate fertility rate as well and there was negligible difference with the ratios calculated using total ever-married women. India being a very conservative country, has very few childbirths out of wedlock. The institution of marriage remains strong in India. Mostly, women get married between 17-19 years (Desai and Andrist, 2010), and the rate of divorce or separation is very low during the period considered here. According to NFHS (2006) less than 1% of the women are divorced, separated or deserted by their husband, which leaves less room for

sition at 3-digit National Industrial Classification (NIC) available in 1991 census. The share of labor force in total output at the NIC level is computed from the 1990 Annual Survey of Industries (ASI). The tariff data is obtained from the UNCTAD-Trade Analysis and Information System (TRAINS) at the 6-digit level of the Indian Trade Classification Harmonized System for around 5000 product lines. Similar to Topalova (2010), we match these product lines to NIC codes using the concordance table provided in Debroy and Santhanam (1993), to get a relatively precise measure of average industry-level tariffs.

As shown in Figure 1, the average ad-valorem tariff declined sharply between 1991 and 1997. The tariffs averaged over 83% in 1991, despite some relaxation in the import-led growth strategy followed by the Indian economy in late 1980s. By 1997, the average ad-valorem tariff had declined to about 29.89%. Figure 2 maps the spatial variation in the tariff measure across districts. As evidenced from the map, there is considerable variation in tariff reduction experienced by different districts. The average tariff reduction between 1991 and 1997 was 4.4 percentage points for all areas. The district at the 10th percentile of tariff reduction distribution experienced only a marginal tariff cut of 0.7 percentage points, whereas the district at the 90th percentile of tariff cut distribution faced a tariff reduction of 10.8 percentage points. Hence, there exists a 10.1 percentage points difference in reduction in tariff protection faced by the districts at 90th and 10th percentile of tariff reduction distribution. While interpreting the subsequent regression coefficients, we will be comparing the districts that faced 10 percentage points larger tariff reductions than the rest thus capturing the tariff reduction gap faced by 90th vs 10th percentile district. There exists differences across urban and rural areas. The urban areas of districts experienced an average tariff reduction of 12 percentage points whereas rural areas of districts experienced an average tariff reduction of only 3.1 percentage points. This is largely a reflection of higher share of urban workforce in manufacturing sector which witnessed considerable decline in tariff.

Table 1 provides the summary statistics for the main variables used in the analysis.

unmarried births. So, we use total ever-married women to get more precise measures.

Average marriage rate among women aged 15-34 years has been declining over the years, and similar declining patterns are observed in both urban and rural areas. Fertility rate has gone up marginally in 2001 but declined by a relatively larger proportion in 2011. Although, a similar pattern is observed in the rural areas, the decline in fertility rate is monotonous in urban areas. The workforce participation rate among young men (15-34) has declined overtime. Although same pattern of decline in observed in rural areas, in urban areas workforce participation rate among young men remain similar between 2001 and 2011 after declining between 1991 and 2001. Following the literature on marriage and fertility dynamics in India, we also consider labor market outcomes for men aged 20-39, since there's a mean spousal age difference of approximately 5 years among married couples in India (Dommaraju, 2008; Bergstrom and Bagnoli, 1993; Edlund, 1999). The workforce participation rate among 20-39 year old men also exhibits similar trend in workforce participation rate. For young women, the workforce participation increased marginally in 2001, but declined in 2011 across all districts. The workforce participation rate is quite low in urban areas compared with the rural areas, however has improved over time monotonously whereas rural areas saw a decline in workforce participation between 2001 and 2011.

5 Results

5.1 Labor market impacts of trade exposure

Previous studies suggest that family formation decisions are affected by trade shocks through their impact on aggregate labor market outcomes. Before addressing the issue of impact of liberalization on marriage and fertility, we examine how the employment prospects of young men and women changed as a result of trade liberalization. Table 2 presents the impacts of reduction in tariff protection on the workforce participation rate (WFPR) of young men in age group 15-34. Panel A of Table 2 presents the estimates for all areas. We start with columns (1) and (5) where a univariate regression relates medium and long-run change in

workforce participation rate to district tariff reduction between 1991 and 1997. Recall that positive values of tariff change imply a tariff cut, and hence a positive coefficient on tariff change imply an increase in participation rate as a result of tariff cut. There is a statistically significant positive relation between tariff cut and change in workforce participation rate for young men both in medium and long-run. The magnitude of the relation is stronger in long-run compared with the medium-run.

In column (2) and (6) of Table 2, we introduce state fixed effects to account for differential state policies over time. Introduction of state fixed effects reduces the magnitude of the coefficient in the medium run but only have a marginal effect on long run coefficient. This suggests that some states that witnessed larger tariff cuts following liberalization also displayed other time varying characteristics that increased the WFP among young men in medium run. In column (3) and (7) of Table 2, we control for pre-reform trends in workforce participation rate (all ages) between 1991 and 1981. We do not have the age wise workers information for the 1981 and 1971 Census, however, the work participation rate is available for all ages. It is highly likely that the trend in workforce participation among 15-34 follows the trend of workforce participation all ages in the same district. Controlling for pre-reform trends reduces the magnitude of coefficient for medium run but has no effect on long run coefficient. It is important that both medium and long run coefficients are positive and statistically significant. In column (9) of Table 2, we run a regression of pre-reform trend in workforce participation rate between 1991 and 1981 on district tariff reduction and state fixed effects. The coefficient on district tariff reduction is relatively small and statistically insignificant. This establishes that there were no differential trends in districts that experienced larger tariff cuts vs. districts that experienced smaller tariff cuts. Hence, one could take estimates from column (3) and column (7) as causal effects of tariff reductions. A district which experienced an average level of tariff cuts (4.4 percentage points) saw 0.3 and 1.1 percentage points higher workforce participation among young men in medium and long-run, respectively, compared to a district that did not experienced any tariff cut. The

long-run effect which captures the cumulative effect of tariff reduction is almost four times of the medium-term impact.

To address the possible correlation between pre-reform trend and error term, we instrument the pre-reform trends (all ages) with the district wise men workforce participation (all ages) in the 1971 Census. The IV estimates are reported in column (4) and (8) of Table 2. The IV estimates are similar in medium run, however, marginally smaller in long run. Importantly, the conclusions drawn remain similar. The districts that experienced larger tariff cuts saw a relative increase in workforce participation among young men compared to districts that experienced smaller tariff cuts, and long run impact is larger than the medium run impact. In Panel B and Panel C of Table 3, we present the results for rural and urban areas, separately. We find similar results in urban areas; however, we find no impact in rural areas. Thus, the positive effect of tariff reduction on workforce participation among young men is concentrated in urban areas only.[13]

In Table 3, we present similar results for young women. The overall conclusions are similar with few differences. For urban areas, districts that witnessed larger tariff cuts also experienced relatively larger increase in WFP among young women, and the long-run impact is much larger compared to medium-run impact. Urban areas of districts that experienced an average level of tariff cuts (4.4 percentage points) saw 0.5 and 1.4 percentage points higher workforce participation among young women in medium and long-run, respectively, compared to urban areas of districts that did not experienced any tariff cut. Given the low WFPR among young women, these impacts are much larger in percent. In column (9) of Table 3, the coefficient on district tariff reduction is relatively much smaller although statistically significant at 10% level. This suggests that there was a small divergence in WFPR between districts that were exposed more to tariff cuts compared to districts that

[13]We also look at the impact on workforce participation rate among men aged 20-39 years to allow for a spousal age gap of about 5 years between men and women in India as discussed earlier. The results are reported in Appendix Table A1. While the medium-term impact on workforce participation rate among men in 20-39 age group is not statistically significant, the long-term impact is positive and statistically significant for all areas and urban areas, while we fail to reject no effect of tariff cut on workforce participation rate among young men in rural areas.

were less exposed to tariff cuts. However, the main coefficients for urban areas (column (4) and (8)) are relative much larger than the placebo coefficient. For rural areas, although we could not reject the null of no impact of tariff cuts on WFP among young women, the placebo presented in column (9) of Table 3 suggests that WFP among women was relatively declining in districts that experienced larger tariff cuts compared to districts that experienced smaller tariff cuts. This indicates that perhaps trade liberalization stopped the relative decline of WFP among young women in districts that witnessed larger tariff cuts.

To summarize, we find trade-liberalization positively effected the WFP among young men and women in urban areas, while it has no impact on WFP among young men in rural areas. One can also speculate that probably trade liberalization stopped the declining trend in WFP among women in rural areas of more exposed districts. It is not surprising that employment effects of tariff liberalization among young men/women are concentrated in urban areas. The tariff reductions were largest in manufacturing (while agriculture remained relatively closed to trade liberalization), which is largely located in urban areas. However, the employment effects of trade liberalization in India differ from the employment effects for the Brazil trade liberalization. Dix-Carneiro and Kovak (2017) find negative effects on employment in Brazil. Braga (2018) confirm negative employment effects of Brazilian trade liberalization among young men and women in age group (20-35). Our findings of increased employment prospects of young men/women in India is consistent with some related evidence presented by others for India. Using a database of publicly listed companies in India, Topalova and Khandelwal (2011) show that reductions in trade protection led to higher levels of productivity resulting from a) lower output tariffs caused firms to increase their efficiency b) lower tariffs on inputs because of trade liberalization increased firms access to more and cheaper imported inputs. Using wage information from urban India, Kumar and Mishra (2008) find that the industries with the largest reductions in tariff protection saw wages increased relative to the economy-wide average. Hasan et al. (2011) uses state level data, and find that overall (rural plus urban) unemployment on average does not have any relationship with average protection over

time and across states. They also find that reductions in protection reduce unemployment in the urban sectors of states with large employment shares in net exporter industries. Using industry-level protection, Hasan et al. (2011) find no evidence that workers in industries that experienced larger reductions in protection were more likely to be unemployed.

5.2 Effects of trade exposure on marriage rates

As discussed in the previous section, the trade liberalization in India improved the employment prospects of young women, mostly in urban area. While the traditional neoclassical theory predicts an inverse relationship between improvement in labor market opportunities for women and their marriage rates. Cherlin, Ribar, and Yasutake (2016) argue that if the economic prospects of women improve, it may lead to an increase (decrease) in marriage rates and stabilize (destabilize) existing marriages, which is the income effect (independence effect). Given the traditional nature of the Indian society, being single or getting a divorce still remains a taboo in India. So, it is likely that independence effect, if any, remains weak. Moreover, improved employment prospects of young men especially in urban areas make them more marriageable as suggested by Wilson and Neckerman (1987) and Wilson (1987).

For the Indian context, the custom of dowry payment by bride's family adds another dimension in the decision making. Existing literature suggests that on the extensive margin, dowry payments have gone up over the years, while on the intensive margin, they rose across all parts of the income distribution. Improved workforce participation by young women in India may increase their chances of getting married since they will have more resources to pay for the dowry and probably they have to pay less dowry. For example, Anderson (2003) finds that in caste-based societies like India, increased dispersion in wealth due to modernization leads to increase in dowry payments. So, if employment opportunities of women improve, they will have more resources to pay for their dowry, thus reducing dowry inflation. Similarly, Srinivasan and Lee (2004) finds that if women are involved in unpaid work, such as family enterprise, then their families are likely to pay high dowries compared

to those where women are engaged in paid work. Given the improved employment prospects of young men/women combined with the dowry system, one may expect improved marriage rates in young women resulting from reduction in tariff protection, especially in urban areas where the employment effects are stronger.

Table 4 presents the results of impact of tariff cuts on the marriage rates of young women aged 15-34. Panel A of Table 4 presents results for all areas, while Panel B and Panel C presents results for rural and urban areas, respectively. Recall that a positive coefficient indicates that districts facing relatively larger tariff cuts experienced a relative increase in the marriage rates. Columns (1) and (4) presents the medium and long-term results of naïve model that does not control for either state fixed effects or pre-reforms trend in marriage rate. The tariff reduction has a positive impact on marriage rate in both medium and long-term suggesting that the districts that experienced larger tariff cuts experienced a relative increase in marriage rate. The long-term impact, that captures cumulative effect of tariff cut, is larger compared to medium-term. Next, we control for state fixed effects in column (2) and column (4) of Table 4. Controlling for state fixed effects reduces the magnitude of point estimates in both medium and long-term, however, the coefficients remain positive and statistically significant. In column (3) and column (6) of Table 4, we add pre-reform trends in marriage rate to our model. It is worth noting that introduction of pre-reform trend in marriage rate does not have any impact on the coefficients suggesting absence of omitted variable bias. In column (4) and column (8), we instrument the pre-existing trend in marriage rate, $(Y_{d,1991} - Y_{d,1981})$, with the 1981 marriage rate and workforce participation rate of male in 1981 and 1971. The magnitude of coefficient on tariff variable declines but remains positive, however, the statistical significance is lost.

In column (9) of Table 4, we present the placebo specification where the change in marriage rates between 1991 and 1981 are correlated with the future change in tariff. We find no evidence on pre-existing trend in marriage rate for all, rural, and urban areas. Panel B and Panel C presents impacts in rural and urban areas, separately. For rural areas, though

the point estimates are positive in both medium and long-term, they are not statistically significant (column (4) and (8) in Table 4, Panel B). For urban areas (Panel C of Table 4), the estimates of our preferred specification suggest a positive and statistically significant impacts both in medium and long-term with long-term impact being larger.[14] The coefficients of 0.067 in column (4) of Panel C and 0.103 in column (8) of Panel C imply that urban areas of districts facing a 10-percentage point larger decline in tariffs saw a 0.67 and 1 percentage point more increase in the marriage rates in the medium- and long-term, respectively. So, the impact on marriage rates are mostly concentrated in urban areas, and are stronger in long-term. Recall that our employment results presented earlier suggest that employment prospects of both young men and women increased in urban areas in districts that experienced more reduction in tariffs, while there was no effect of liberalization on employment prospects in rural areas. Moreover, the employment improvement among young men/women is stronger in long-term compared to short-term. We get similar findings for improvement in marriage rates, where the improvement in marriage rates is concentrated in urban areas and stronger in long-term compared to medium-term.

Since the employment prospects for both young women and men improved in urban areas of the districts that witnessed larger tariff cuts, we conjecture the presence of dowry channel that reinforces the improvement in marriage rates together with the improvement of marriageability for young men because of improved employment prospects for young men. Our findings are although similar to Keller and Utar (2022) who also observe higher marriage rates for women due to the Chinese import shock, the channels are entirely different. They find improvement in marriage rate due to substitution from labor market to family activities because of negative labor market shocks from Chinese imports. In our case, young women did not face negative labor market consequences as a result of liberalization but their employment prospects improved increasing their value in the marriage market. Similarly, though our results of improved employment prospects for young women is similar to Kis-Katos, Pieters,

[14]The long-run effect is significant at 6% (*p-value* 0.052). Note that given the small sample size, it is reasonable to use a 10% significance level criterion.

and Sparrow (2018) findings of positive effects on women employment, however, they find a reduction in the marriage rates as the opportunity cost of performing domestic duties increases. It is noteworthy to point out the a custom of mahr (dowry) is followed in Indonesia too, however, it is just opposite of Indian dowry system. While in the Indian dowry system, bride's family pays dowry to the groom's family, mahr in Indonesia is paid by the groom's family to bride's family. Hence, we may expect the dowry system in these two countries (India and Indonesia) to have very different effect on the marriage rates.

5.3 Effects of trade exposure on fertility rates

We next move to the findings of the impact of trade liberalization on the evolution of fertility rates of young women aged 15-34 years.[15] Table 5 reports the results on fertility rates from our analysis. Overall, our naïve estimates suggest a relatively larger decline in fertility in more exposed districts. Adding state fixed effects reduces the magnitude of the coefficient on tariff measure in both medium and long-term (Column (2) and Column (4) of Table 5, Panel A). In column (3) and (6) of Table 5 we add the pre-reforms trend in the fertility outcome to our specification. Addition of pre-reforms trend in fertility outcome increases the magnitude of point estimate for both medium and long-term.[16] In column (4) and (8) of Table 5, we instrument pre-reform trend in fertility rate with the 1981 fertility rate. The coefficients on the tariff cut measure remain similar. So, in medium term a district that experienced 10 percentage points more reduction in tariff experienced a relatively larger decline in fertility rate by 0.07 child, while in log-run more exposed district will experience relatively larger decline by 0.1 child per women. In column (9) of Table 5, we report the result of placebo specification where change in pre-reform fertility rate is regressed on future (post reform)

[15]The link between globalization and fertility has been explored both at the macro level (Ahn and Mira, 2002; Adsera, 2004) as well as at the individual level (Adsera, 2005, Lindo, 2010) with labor force participation rates being the main channel.

[16]The 1981 Census does not have information on children ever born age wise, so we use the district wise total fertility rate difference between 1991 and 1981 as pre-trends. It is highly likely that total fertility trend among 15-34 age group is similar to fertility trend among 15-44 (used for total fertility rate calculation) in same district.

change in tariff. We find a statistically insignificant (also very small magnitude) coefficient for tariff cut measure suggesting that there was no prevailing pre-existing trend in fertility before the reform.

Panel B and Panel C of Table 5 report results for rural and urban areas, respectively.[17] The coefficients capturing the medium term and long-term impacts on fertility in rural areas are statistically significant for rural areas (Column (4) and (8) of Table 5, Panel B). For urban areas, while the long-run impact remain statistically insignificant, the medium term impact is negative and statistically significant at 10% significance level. Nevertheless, the medium-term impact in urban areas is relatively small suggesting a 10 percentage points reduction in tariff will contribute to fertility decline by 0.02 child per women. Our findings of relative decline in fertility rate in rural areas of relatively more exposed districts do not conform with Anukriti and Kumler (2019) findings. They find that probability of birth declined less in districts that experienced larger tariff cuts. It is worth pointing that variation in tariff protection in Anukriti and Kumler (2019) is yearly and they consider probability of birth in each year, while in our case we look at the liberalization as one-time event and outcome is captured by the decadal change in fertility rate. We carry out similar analysis for women aged 15-49 years, and the results are qualitatively similar (reported in Appendix Table A2), which shows selection of the age-group for our main results is not an issue.

So as mentioned in above paragraph, there is no differential impact on fertility in long-run (while marginal negative impact in medium-run) in urban areas of districts more exposed to tariff cuts. At the same time, urban areas of more exposed districts saw a relative improvement in employment prospects for both young men and women in long-run and relative increase in marriage rates. All three findings together suggests presence of quantity-quality trade-off. Beginning with the seminal work of Becker (1960) and Becker and Lewis (1973) on the relationship between fertility decision and investment in child quality several

[17]Because of non-availability of areas/age wise data in the 1981 census, we do not have fertility rates for rural and urban areas, separately. So, we control for district wise (combined rural and urban areas) pre-existing trends in total fertility for rural and urban analysis.

papers have empirically documented the relationship in different settings. Kugler and Kumar (2017) find evidence of quantity-quality trade-off in India. They also find the trade-off is larger in rural area. We conjecture that the limited impact on fertility in urban areas may also be contributed by increased opportunity cost of raising child due to the improved labor market prospects for young women potentially negating impact of income which may result from relatively better employment prospects of both young men and women.

What is puzzling in our case is that the fertility declined more in rural areas of district that experienced larger tariff cuts, however, the employment prospects and marriage rates do not explain the negative relation between tariff cuts and fertility in rural areas as neither workforce participation among young men/women and nor marriage rate seem relatively more affected in rural areas of district that experienced larger tariff cuts. In this context, it important to point out that Edmonds, Pavcnik, and Topalova (2010) find that schooling cost in rural areas increased more in more exposed districts. Topalova (2010) find that rural areas of districts that experienced large reduction in tariff saw slower poverty reduction relative to national trend. Hence, a slower reduction in poverty and higher schooling costs in more exposed districts coupled with the desire of parents to invest in child may be contributing to the relatively more decline in fertility rate in rural areas even in the absence of differential impact on labor market prospects and marriage rates for young women.

6 Conclusion

In this paper, we look at the medium- and long-run impact of Indian trade liberalization introduced in 1991 on marriage rate and fertility rate among young women (15-34) using district-level panel constructed from Indian Censuses covering 1971-2011. We exploit the heterogeneous reduction in tariff protection experienced by Indian districts based on their 1991 industrial distribution of workers to implement a difference-in-difference strategy. The use of census allows us to look at the impact in rural and urban areas separately while most

of the trade liberalization literature in India that use survey data focus on rural areas only.

We find that urban areas of districts that experienced larger tariff cuts witnessed relatively increased (compared to districts that experienced smaller tariff cuts) work participation rates among young men and women, and increased marriage rates among young women. However, we fail to reject the null of no differential impact on fertility in urban areas of more exposed districts in long-run. While the improved marriage rates in urban areas of more exposed districts can be a result improved marriageability of young men because of relatively better workforce participation, we conjecture the importance of dowry channel in the Indian context which negates relatively larger opportunity cost of time because of relatively better workforce participation among young women in more exposed districts. This leads to relatively larger increase in marriage rates in more exposed districts. For fertility, in the absence of dowry channel, increased opportunity cost of women's time combined with the desire to invest in quality of child may have negated any positive income effect through better workforce participation among young men/women leading to no differential effect on fertility in urban areas of more exposed districts.

We find quite different result for rural areas. Although, we do not find differential impacts on workforce participation rate of young men/women in rural areas of more exposed districts. Similarly, there was no differential impacts on the marriage rate of young women. However, we find that the fertility rate declined more in rural areas of more exposed districts. We conjecture that the relatively slower reduction in poverty and relative increased higher cost of schooling in rural areas of districts that experienced relatively larger tariff cuts as documented in the literature combined with preference for quality may potentially explain relative decline in fertility rate in rural areas of more exposed districts.

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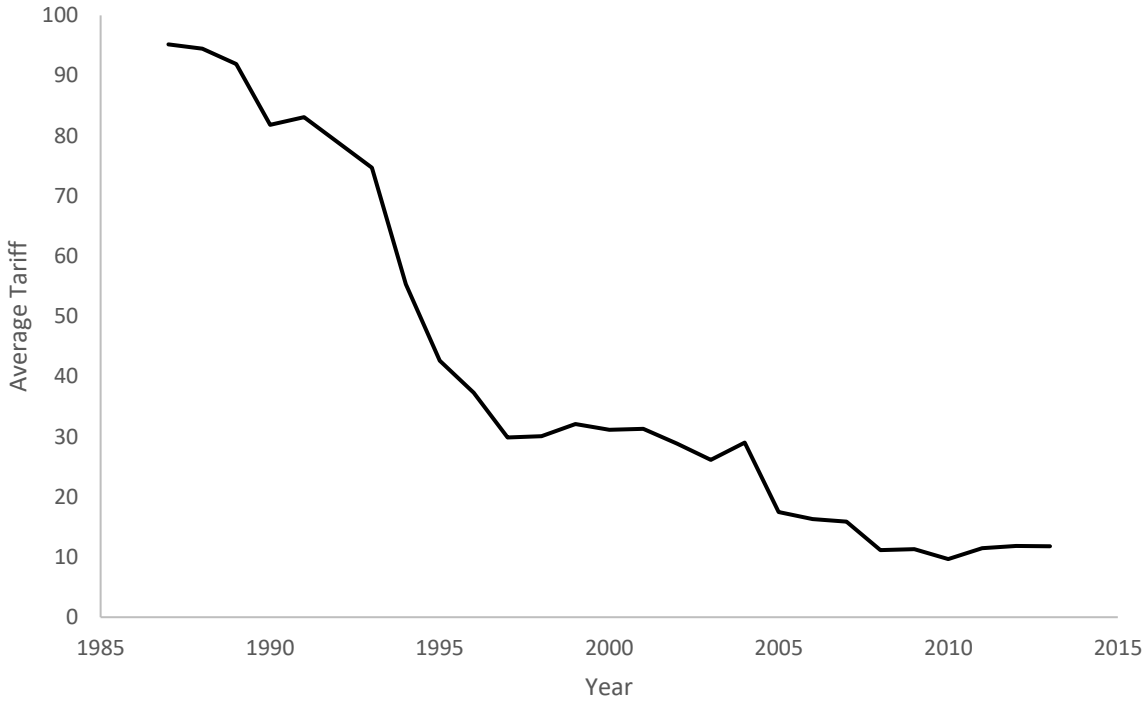
Figure 1. Evolution of average tariff overtime



Note: Author's computations using data from the UNCTAD- Trade Analysis and Information System (TRAINS). Average nominal ad-valorem tariff is plotted between over time

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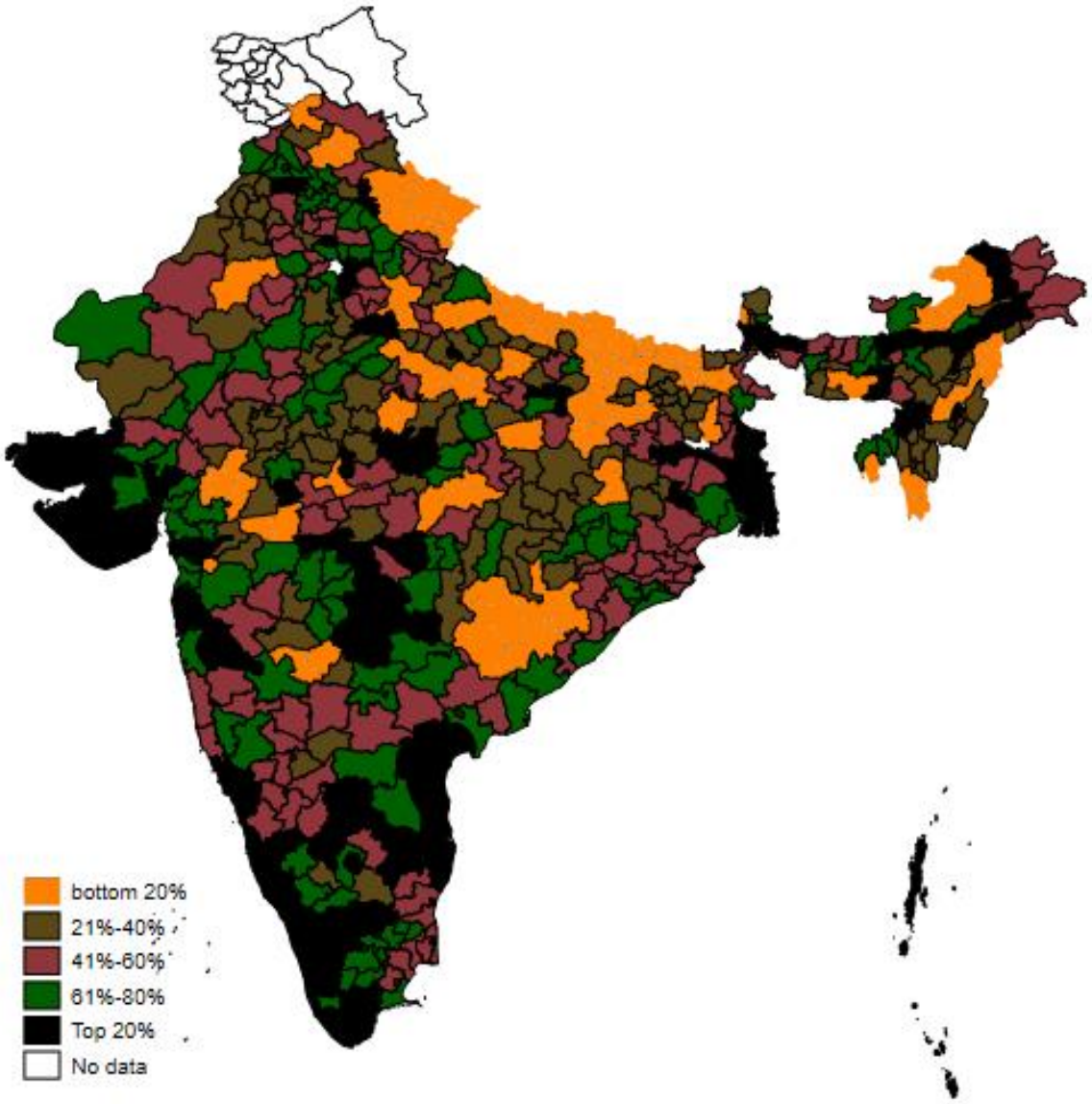
Figure 2. District-wise Tariff Reductions



			Percentile				
	Districts	Mean	10	25	50	75	90
All Areas	452	0.044	0.007	0.010	0.019	0.044	0.108
Rural Areas	446	0.031	0.005	0.007	0.013	0.027	0.077
Urban Areas	442	0.120	0.042	0.066	0.108	0.161	0.214

Notes: There were total 462 districts in 1991 Census, but 1991 Census did not cover the state of Jammu and Kashmir.

Table 1: Descriptive Statistics

	All Districts			Rural Districts			Urban Districts		
	1991	**2001**	**2011**	**1991**	**2001**	**2011**	**1991**	**2001**	**2011**
Women aged 15-34 years									
Marriage Rate	0.752	0.700	0.655	0.778	0.732	0.677	0.673	0.624	0.610
Fertility Rate	1.985	2.011	1.736	1.996	2.079	1.828	1.953	1.822	1.533
Workforce Participation Rate									
Women aged 15-34 years	0.335	0.363	0.326	0.417	0.457	0.398	0.123	0.143	0.180
Men aged 15-34 years	0.740	0.704	0.665	0.779	0.743	0.686	0.646	0.619	0.622
Men aged 20-39 years	0.889	0.869	0.845	0.915	0.897	0.866	0.826	0.808	0.806
Observations	452	452	452	446	446	446	442	442	442

Table 2: Medium- and Long-term impacts of Tariff Exposure on Workforce Participation Rate (WPR), Men 15-34

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	1991-2001				**1991-2011**				**1981-1991**
Panel A: All Areas									All Ages
District level tariff cuts	0.107**	0.075**	0.062*	0.059*	0.258***	0.254***	0.255***	0.184**	0.034
	(0.045)	(0.029)	(0.033)	(0.035)	(0.087)	(0.063)	(0.067)	(0.078)	(0.021)
Formal WPR (All Ages) trend, 1981-1991			0.399***	0.485			0.186	2.021***	
			(0.107)	(0.320)			(0.143)	(0.478)	
Observations	452	452	427	427	452	452	427	427	427
Kleibergen Paap Wald rk F-statistics				31.40				29.18	
Panel B: Rural Areas									
District level tariff cuts	0.059	0.009	-0.007	-0.015	0.185*	0.023	0.014	-0.016	0.022
	(0.054)	(0.024)	(0.020)	(0.024)	(0.099)	(0.047)	(0.048)	(0.066)	(0.035)
Formal WPR (All Ages) trend, 1981-1991			0.454***	0.831*			0.288**	1.768***	
			(0.113)	(0.432)			(0.135)	(0.432)	
Observations	445	445	420	420	445	445	420	420	420
Kleibergen_Paap Wald rk F-statistics				18.93				17.18	
Panel C: Urban Areas									
District level tariff cuts	0.153***	0.065**	0.062**	0.060**	0.117	0.130***	0.127***	0.103**	0.015
	(0.050)	(0.030)	(0.030)	(0.029)	(0.098)	(0.041)	(0.042)	(0.047)	(0.017)
Formal WPR (All Ages) trend, 1981-1991			0.209	0.449			-0.048	1.618***	
			(0.134)	(0.335)			(0.187)	(0.611)	
Observations	441	441	414	414	441	441	414	414	414
Kleibergen_Paap Wald rk F-statistics				19.46				20.20	
State fixed effects	NO	YES	YES	YES	NO	YES	YES	YES	YES
IV	NO	NO	NO	YES	NO	NO	NO	YES	NO

Notes: * p<0.01; ** p<0.05; *** p<0.01. Standard errors (in parenthesis) are clustered at the state-region level (74 clusters). Positive coefficients for district tariff reduction imply that districts facing larger tariff cuts witnessed a relative increase in workforce participation. All regressions are weighted by the average population of men aged 15-34 years in the concerned years. The pre-reform trend for men WPR is instrumented by the 1971 men WPR. The number of observations with pretends omits one state because of non-coverage by the 1981 Census. The number of observations differ across rural and urban areas as few districts are either complete urban or rural.

Table 3: Medium- and Long-term impacts of Tariff Exposure on Workforce Participation Rate (WPR), Women 15-34

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	1991-2001				**1991-2011**				**1981-1991**
Panel A: All Areas									All Ages
District level tariff cuts	-0.178**	0.082	0.098	0.124	0.088	0.473***	0.476***	0.539***	-0.108***
	(0.072)	(0.075)	(0.078)	(0.077)	(0.124)	(0.129)	(0.130)	(0.123)	(0.032)
Formal WPR (All Ages) trend, 1981-1991			0.083	0.321			-0.261**	0.282	
			(0.092)	(0.210)			(0.127)	(0.286)	
Observations	452	452	427	427	452	452	427	427	427
Kleibergen_Paap Wald rk F-statistics				14.75				14.61	
Panel B: Rural Areas									
District level tariff cuts	-0.483***	-0.039	-0.039	-0.035	-0.540***	-0.048	-0.037	-0.001	-0.106**
	(0.152)	(0.062)	(0.071)	(0.072)	(0.144)	(0.081)	(0.080)	(0.076)	(0.041)
Formal WPR (All Ages) trend, 1981-1991			0.000	0.037			-0.272*	0.066	
			(0.107)	(0.255)			(0.145)	(0.327)	
Observations	445	445	420	420	445	445	420	420	420
Kleibergen_Paap Wald rk F-statistics				10.02				9.82	
Panel C: Urban Areas									
District level tariff cuts	0.104**	0.118***	0.122***	0.110***	0.220***	0.316***	0.328***	0.313***	0.027*
	(0.047)	(0.027)	(0.027)	(0.024)	(0.067)	(0.049)	(0.053)	(0.046)	(0.015)
Formal WPR (All Ages) trend , 1981-1991			-0.122	0.375*			-0.427**	0.179	
			(0.104)	(0.199)			(0.168)	(0.338)	
Observations	441	441	414	414	441	441	414	414	414
Kleibergen_Paap Wald rk F-statistics				11.61				11.07	
State fixed effects	NO	YES	YES	YES	NO	YES	YES	YES	YES
IV	NO	NO	NO	YES	NO	NO	NO	YES	NO

Notes: * p<0.01; ** p<0.05; *** p<0.01. Standard errors (in parenthesis) are clustered at the state-region level (74 clusters). Positive coefficients for district tariff reduction imply that districts facing larger tariff cuts witnessed a relative increase in workforce participation. All regressions are weighted by the average population of women aged 15-34 years in the concerned years. The pre-reform trend for women WPR is instrumented by the 1991 and 1981 men WPR. The number of observations with pretends omits one state because of non-coverage by the 1981 Census. The number of observations differ across rural and urban areas as few districts are either complete urban or rural.

Table 4. Medium- and Long-run impacts of Tariff Exposure on Marriage Rates of young women (15-34)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	1991-2001				**1991-2011**				**1981-1991**
Panel A: All Areas									
District level tariff cuts	0.188***	0.069***	0.070***	0.042	0.433***	0.191***	0.200***	0.134	0.038
	(0.051)	(0.022)	(0.023)	(0.034)	(0.114)	(0.049)	(0.052)	(0.087)	(0.023)
Formal marriage rate pre-trend, 1981-1991			0.142**	0.943***			0.068	2.130***	
			(0.066)	(0.259)			(0.069)	(0.663)	
Observations	452	452	427	427	452	452	452	427	427
Kleibergen Paap Wald rk F-statistics				25.82				22.99	
Panel B: Rural Areas									
District level tariff cuts	0.281***	0.041	0.058*	0.057	0.631***	0.098	0.126*	0.125	0.002
	(0.084)	(0.029)	(0.031)	(0.035)	(0.177)	(0.064)	(0.072)	(0.077)	(0.042)
Formal marriage rate pre-trend, 1981-1991			0.199**	0.602***			0.056	0.669	
			(0.081)	(0.211)			(0.089)	(0.507)	
Observations	445	445	421	421	446	446	446	422	421
Kleibergen Paap Wald rk F-statistics				21.33				17.56	
Panel C: Urban Areas									
District level tariff cuts	0.172***	0.081***	0.081***	0.067**	0.253***	0.137***	0.133***	0.103*	0.021
	(0.039)	(0.027)	(0.027)	(0.031)	(0.081)	(0.045)	(0.045)	(0.053)	(0.023)
Formal marriage rate pre-trend, 1981-1991			0.032	0.618**			0.077	1.143***	
			(0.062)	(0.250)			(0.087)	(0.443)	
Observations	441	441	420	420	441	441	420	413	420
Kleibergen Paap Wald rk F-statistics				28.08				19.69	
State Fixed Effect	No	Yes	Yes	Yes	No	Yes	Yes	Yes	Yes
IV	No	No	No	Yes	No	No	No	Yes	NA

Notes: * p<0.01; ** p<0.05; *** p<0.01. Standard errors (in parenthesis) are clustered at the state-region level (74 clusters). Positive coefficients for district tariff reduction imply that districts facing larger tariff cuts witnessed a relative increase in marriage rate. All regressions are weighted by the average population of women aged 15-34 years in the concerned years. The pre-reform trend for marriage rate is instrumented by the 1981 marriage rate, and WFPR among men in 1971 and 1981. The number of observations with pretends omits one state because of non-coverage by the 1981 Census. The number of observations differ across rural and urban areas as few districts are either complete urban or rural.

Table 5: Medium- and Long-run impacts of Tariff Exposure on Fertility Rates of young women (15-34)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	1991-2001				**1991-2011**				**1981-1991**
Panel A: All Areas									
District level tariff cuts	-1.559***	-0.571***	-0.678***	-0.679***	-1.681***	-0.684***	-0.998***	-0.998***	0.033
	(0.326)	(0.156)	(0.131)	(0.125)	(0.373)	(0.204)	(0.203)	(0.197)	(1.100)
Total fertility rate, 1981-1991			-0.004	-0.013			0.006	0.001	
			(0.012)	(0.012)			(0.012)	(0.011)	
Observations	447	447	416	416	447	447	416	416	416
Kleibergen Paap Wald rk F-statistics				2914.90				2911.84	
Panel B: Rural Areas									
District level tariff cuts	-2.078***	-0.478***	-0.511***	-0.527***	-2.390***	-0.783***	-0.875***	-0.885***	
	(0.538)	(0.100)	(0.112)	(0.109)	(0.640)	(0.198)	(0.215)	(0.209)	
Total fertility rate, 1981-1991			-0.010	-0.019			-0.004	-0.010	
			(0.014)	(0.014)			(0.012)	(0.012)	
Observations	441	441	414	414	441	441	414	414	
Kleibergen Paap Wald rk F-statistics				2675.99				2684.30	
Panel C: Urban Areas									
District level tariff cuts	-0.522**	-0.279	-0.289*	-0.274*	-0.365**	-0.171	-0.185	-0.176	
	(0.212)	(0.172)	(0.155)	(0.151)	(0.179)	(0.181)	(0.171)	(0.166)	
Total fertility rate, 1981-1991			0.009	0.003			0.012	0.009	
			(0.010)	(0.010)			(0.013)	(0.012)	
Observations	437	437	407	407	437	437	407	407	
Kleibergen Paap Wald rk F-statistics				2310.00				2079.06	
State Fixed Effect	No	Yes	Yes	Yes	No	Yes	Yes	Yes	Yes
IV	No	No	No	Yes	No	No	No	Yes	NA

Notes: * p<0.01; ** p<0.05; *** p<0.01. Standard errors (in parenthesis) are clustered at the state-region level (74 clusters). Positive coefficients for district tariff reduction imply that districts facing larger tariff cuts witnessed a relative increase in fertility rate. All regressions are weighted by the average population of women aged 15-34 years in the concerned years. The pre-reform trend for total fertility rate is instrumented by the 1981 total fertility rate. The number of observations with pretends omits one state because of non-coverage by the 1981 Census. The number of observations differ across rural and urban areas as few districts are either complete urban or rural.

Table A1: Medium and Long-term impacts of Tariff Exposure on Workforce Participation Rate among men (age group 20-39)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	1991-2001				**1991-2011**			
Panel 1a: All Areas								
District level tariff cuts	0.065*	0.017	0.010	0.021	0.144**	0.124***	0.122***	0.114***
	(0.033)	(0.016)	(0.018)	(0.017)	(0.063)	(0.031)	(0.032)	(0.032)
Formal WPR (All Ages) trend , 1981-1991			0.235***	-0.117			0.191*	0.390*
			(0.064)	(0.266)			(0.102)	(0.213)
Observations	452	452	427	427	452	452	427	427
Kleibergen Paap Wald rk F-statistics								
Panel 1b: Rural Areas								
District level tariff cuts	0.065	0.006	-0.000	0.008	0.143**	0.012	0.009	0.015
	(0.040)	(0.014)	(0.014)	(0.020)	(0.071)	(0.028)	(0.028)	(0.033)
Formal WPR (All Ages) trend , 1981-1991			0.214***	-0.173			0.191	-0.072
			(0.077)	(0.376)			(0.127)	(0.339)
Observations	445	445	420	420	445	445	420	420
Kleibergen_Paap Wald rk F-statistics				18.93				17.18
Panel 1b: Urban Areas								
District level tariff cuts	0.118***	0.023	0.019	0.018	0.096	0.080***	0.073***	0.063**
	(0.039)	(0.021)	(0.022)	(0.021)	(0.073)	(0.024)	(0.025)	(0.026)
Formal WPR (All Ages) trend , 1981-1991			0.290***	0.452*			0.266*	0.994**
			(0.095)	(0.249)			(0.142)	(0.387)
Observations	441	441	414	414	441	441	414	414
Kleibergen_Paap Wald rk F-statistics				19.46				20.20
State fixed effects	NO	YES	YES	YES	NO	YES	YES	YES
IV	NO	NO	NO	YES	NO	NO	NO	YES

Notes: * p<0.01; ** p<0.05; *** p<0.01. Standard errors (in parenthesis) are clustered at the state-region level (74 clusters). Positive coefficients for district tariff reduction imply that districts facing larger tariff cuts witnessed a relative increase in workforce participation. All regressions are weighted by the average population of men aged 20-39 years in the concerned years. The pre-reform trend for men WPR is instrumented by the 1971 men WPR. The number of observations with pretends omits one state because of non-coverage by the 1981 Census. The number of observations differ across rural and urban areas as few districts are either complete urban or rural.

Table A2. Medium- and Long-run impacts of Tariff Exposure on Fertility Rates of women (15-49 years)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	1991-2001				**1991-2011**			
Panel A: All Areas								
District level tariff cuts	-1.825***	-0.487**	-0.711***	-0.713***	-2.015***	-0.590**	-1.076***	-1.076***
	(0.458)	(0.211)	(0.161)	(0.153)	(0.530)	(0.291)	(0.262)	(0.253)
Total fertility rate, 1981-1991			0.003	-0.009			0.021	0.012
			(0.016)	(0.015)			(0.017)	(0.015)
Observations	447	447	416	416	447	447	416	416
Kleibergen Paap Wald rk F-statistics				2951.21				2952.49
Panel B: Rural Areas								
District level tariff cuts	-2.605***	-0.515***	-0.536***	-0.560***	-3.060***	-0.827***	-0.903***	-0.922***
	(0.707)	(0.149)	(0.166)	(0.160)	(0.841)	(0.297)	(0.328)	(0.319)
Total fertility rate, 1981-1991			-0.002	-0.015			0.011	0.001
			(0.018)	(0.018)			(0.018)	(0.017)
Observations	441	441	414	414	441	441	414	414
Kleibergen Paap Wald rk F-statistics				2710.64				2725.44
Panel C: Urban Areas								
District level tariff cuts	-0.626**	-0.204	-0.262	-0.242	-0.391	-0.025	-0.089	-0.076
	(0.286)	(0.212)	(0.195)	(0.189)	(0.242)	(0.245)	(0.225)	(0.219)
Total fertility rate, 1981-1991			0.010	0.002			0.018	0.013
			(0.013)	(0.012)			(0.015)	(0.014)
Observations	437	437	407	407	437	437	407	407
Kleibergen Paap Wald rk F-statistics				2267.97				2054.77
State Fixed Effect	No	Yes	Yes	Yes	No	Yes	Yes	Yes
IV	No	No	No	Yes	No	No	No	Yes

Notes: * p<0.01; ** p<0.05; *** p<0.01. Standard errors (in parenthesis) are clustered at the state-region level (74 clusters). Positive coefficients for district tariff reduction imply that districts facing larger tariff cuts witnessed a relative increase in fertility rate. All regressions are weighted by the average population of women aged 15-34 years in the concerned years. The pre-reform trend for total fertility rate is instrumented by the 1981 total fertility rate. The number of observations with pretends omits one state because of non-coverage by the 1981 Census. The number of observations differ across rural and urban areas as few districts are either complete urban or rural.