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ABSTRACT

The Effects of Trade Exposure on Marriage and Fertility Choices: Evidence from Brazil*

This paper investigates the effect of a large economic shock on marriage and fertility choices. I exploit the 1990's trade liberalization in Brazil, which created exogenous negative labor market shocks to regions most exposed to foreign competition. While trade liberalization had a positive impact on reducing the price of consumer goods in Brazil, it also negatively impacted employment in previously protected industries, affecting men more than women. I find that young women living in regions more exposed to international competition are less likely to have children. Most effects persist for 20 years after trade liberalization. I use causal mediation analysis to show that declines in the employment rate of young men is an important driver of changes in fertility outcomes of young women. Changes in women's employment opportunities are not a mediator for the effect of trade exposure on fertility. There is no evidence of changes in marriage rates across regions exposed to trade liberalization.

JEL Classification: F16, J12, J13, J21, J23
Keywords: marriage rates, fertility, trade liberalization

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1 Introduction

Marital and childbearing choices are two of the most important decisions individuals make throughout their lives. Couples typically marry when the gains from marriage exceed the gains from being single. The benefits of marriage come from household specialization, with men traditionally devoting more time to wage-earning activities, while women assume responsibility for household labor (Stevenson and Wolfers 2007). In the same way, permanent changes in income and the cost of raising a child impact fertility outcomes (Becker 1960). Neoclassical economic theory suggests that improvements in male labor market conditions should be associated with increases in fertility, while better labor market opportunities for women should have opposing income and substitution effects (Schaller 2016).

This paper investigates the impact of a large economic shock on marriage and fertility decisions of young women in Brazil. Studying the effect of trade liberalization on family life is important because of the tremendous increase of international trade over the past thirty years, and its potential to change the economic circumstances of families. For instance, the world's exports and imports, as a share of the world's GDP, rose from just under 40% in 1990 to 58% in 2015 (World Bank Development Indicators 2018). Trade liberalization can have two main offsetting effects on a family's well-being: a positive impact from reducing the price of consumer goods, as well as a negative impact on the employment rates of previously protected industries due to an increase in international competition. (Feliciano 2001, Galiani and Sanguinetti 2003, Kovak 2013, Dix-Carneiro and Kovak 2017).

The adverse effects of trade liberalization are not gender neutral. Men are disproportionately employed in the tradable sector (manufacturing and agriculture), while women are disproportionately employed in the non-tradable sector (services). In addition, international competition might generate a more favorable labor market for women by preventing taste-based discrimination (Becker et al. 1971). Finally, foreign competition may also induce technological change. If technological induced progress reduces the physical strength required for work, relative demand for female workers can increase with trade liberalization (Juhn et al. 2014). Consistent with theory, a growing number of studies have shown that, while regions with greater exposure to trade experienced worse labor market outcomes for both men and women, the effects on men are larger than the effects on women (Black and Brainerd 2004, Aguayo-Tellez et al. 2014, Gaddis and Pieters 2017, Benguria and Ederington 2017).[1]

Building upon the existing evidence that trade liberalization worsens economic opportunities for workers in more exposed regions, this paper examines the effect of a trade liberalization shock on marriage and fertility choices of young women in Brazil. Traditional economic theory suggests that there are greater gains to a marriage when men can specialize in wage-earning activities and women can specialize in household activities (Becker 1981). As a result, inferior labor market opportunities for men are expected to decrease marriage rates, while inferior labor market opportunities for women are expected to increase marriage rates (Blau et al. 2000).

In terms of fertility decisions, traditional neoclassical economic theory predicts that a substandard labor market for males likely represents a substantial decline in the family income, which is typically associated with lower fertility rates (Becker 1960). However, researchers have also identified potential trade-offs between quantity and quality of children, where families favor investing in the quality of children, rather than in the quantity of children, when income increases (Willis 1973 and Becker and Lewis 1973). Finally, changes in the female labor market conditions could have two opposite effects on fertility. On one hand, an inferior economic labor market for women is associated with a decline in family income. On the other hand, opportunity costs of bearing a child are lower when women cannot easily find work (Schaller 2016).

This paper tests these predictions in the context of Brazil's unilateral tariff reductions during the early 1990s. The empirical strategy consists of comparing outcomes in regions more or less exposed to foreign competition before and after the change in the trade policy of Brazil (Kovak 2013). Consistent with previous literature, I first show that trade liberalization had negative labor market effects for young workers in more exposed regions, impacting men more than women (Gaddis and Pieters 2017). The negative effects for male and female workers are persistent even twenty years after trade liberalization (Dix-Carneiro and Kovak 2017).

Next, I show that 20 to 35-year-old women in regions more exposed to trade liberalization are less likely to have children. I estimate that a median increase in trade exposure faced by a microregion during the early 1990s is associated with a 1.5 percentage-point increase in the region's

[1] Gaddis and Pieters (2017) find significant effects of trade liberalization on gender gap in employment *levels* but do not find changes in gender gaps in *log* employment. Employment *levels* are the most relevant statistic for this study, as I am interested in how family structure changes with household income.

share of young women with no children in 2000. Fertility rates are lower in more exposed regions even 20 years after trade liberalization, with a median increase in trade exposure being associated with a 0.11 decline in the number of children per woman in the region in 2010.

To disentangle the role of male and female employment on fertility, I use the causal mediation analysis framework developed by Imai et al. (2011). Assuming sequential ignorability, I estimate the importance of male and female employment, as mediators for the effects of trade exposure on fertility decisions. I find that changes in the employment of young men explain about 20% of the medium and long-term effects of trade exposure on fertility decisions of young women. I also find that women's employment is a weak mediator for the effect of trade liberalization on fertility. In other words, the adverse labor market outcomes for women did not translate to higher or lower fertility rates.

In the case of marriage outcomes, I find no significant differences in marriage or cohabitation rates in regions most affected by trade liberalization. While traditional theory predicts lower incentives for household specialization and, therefore, lower marriage rates in regions more affected by trade liberalization, I find no changes in marriage rates as a response to this substantial negative labor market shock. This result is consistent with recent work conducted in the United States in which marriage rates are unresponsive to significant economic changes (Kearney and Wilson 2017).

The paper is organized as follows: Section 2 discusses the existing evidence of the impact that labor market shocks have on fertility and marriage outcomes and presents the contribution of the paper. Section 3 presents a conceptual framework for the empirical results. Section 4 outlines the institutional background of trade liberalization in Brazil and the data used for estimation. Moreover, section 5 discusses the estimation strategies used, while Section 6 presents our results for labor market, marriage, and fertility outcomes. Lastly, Section 7 the mechanisms driving those findings and and Section 8 concludes.

2 Related Literature and Contribution

This paper adds to the recent literature that estimates the impact of labor market shocks on marriage and fertility decisions (Lindo 2010, Black et al. 2013, Schaller 2016, Autor et al. 2017, Kearney and Wilson 2017 and Kis-Katos et al. 2017). The overall finding in this literature is that improvements (declines) in men's labor market conditions are associated with increases (decreases) in fertility. The relationship between demand shocks and marriage rates is less conclusive. On one hand, exploiting trade shocks in the United States during the 1990 to 2010 period, Autor et al. (2017) estimate a decline in marriage rates associated with unfavorable job prospects for men and Kis-Katos et al. (2017) find that a trade-induced increase in female employment has led to reductions in marriage rates among young women in Indonesia. On the other hand, Kearney and Wilson (2017) find no evidence of increasing marriage rates associated with a positive economic shock generated by "fracking booms" throughout the United States. While these findings are significant, this literature has focused almost exclusively on developed countries. This paper explores the effect of a large economic shock on both fertility and marriage choices in a developing country. Demand shocks can entail vastly different fertility and marriage outcomes in less developed countries. This is due to the rigid social and cultural norms present, as well as limited access to contraceptive methods, which could make women's fertility decisions less responsive to economic shocks.

This paper also adds to a growing body of recent literature on the regional impacts of trade liberalization in Brazil. While trade liberalization had a potential positive impact on reducing the price of consumer goods nationwide, it also generated persistent negative labor market effects in previously protected industries (Kovak 2013 and Dix-Carneiro and Kovak 2017), impacting more negatively the employment level of men than that of women (Gaddis and Pieters 2017). The increased exposure to foreign competition is also associated with declines in racial and gender wage gaps (Hirata and Soares 2015 and Benguria and Ederington 2017). Finally, there is evidence that areas more affected by trade liberalization experience an increase in crime rates (Dix-Carneiro et al. 2018). Despite extensive literature exploring the regional effects of trade shocks in Brazil, this is the first paper, to my knowledge, to look at the impact of trade liberalization on marriage and fertility decisions in Brazil. Given the size and persistence of employment shocks, one should expect the possibility that trade liberalization could affect the marriage market and the desire to have children.

3 Conceptual Framework

The traditional neoclassical theory of fertility decisions was set by Becker (1960): children provide utility to parents in much the same way as the consumption of other goods. Utility maximizing parents make fertility decisions subject to a family budget constraint. As a result, changes in wages, income, and the cost of raising children cause income and substitution effects on fertility decisions. Because children have very few substitutes, Becker (1960) predicts that fertility increases with family income. However, parents derive utility from both the quantity and the quality of children, which can be proxied by the amount spent on each child (Doepke 2015). Economists have argued that a low-income elasticity of child quantity and a high-income elasticity of child quality can justify a weak correlation between income and fertility. In terms of substitution effects, raising a child is a time-intensive activity, especially for mothers. Labor market improvements increase the opportunity costs of raising a child which might lead to a decline in fertility rates (Black et al. 2013).

Based on this neoclassical framework, the prediction of the impact of trade exposure shocks on fertility is ambiguous. On one hand, an unfavorable labor market for men is typically associated with negative income effects. If children are a normal good, the traditional neoclassical model predicts a decline in fertility in regions more exposed to international competition. The strength of income effects is determined by the preference of families for quantity or quality of children. On the other hand, an unfavorable labor market for women is likely associated with both substitution and income effects, with women typically devoting more time to raise a child. If substitution effects are strong enough, the neoclassical model can predict increases in fertility in more exposed regions.

Becker (1973) and Becker (1974) proposed the neoclassical theory of marriage decisions: couples typically marry when the expected gains from marriage exceed the gains from being single. The decision to get married is based on "production complementarities", where the husband typically specializes in market activities and the wife typically specializes in domestic activities (Becker 1981), though changes in social norms might challenge the perceived gains from household specialization (Stevenson and Wolfers 2007).

The prediction of the neoclassical model about the impact of trade liberalization on marriage rates is also ambiguous. First, unfavorable labor market opportunities for men are expected to decrease marriage rates (Blau et al. 2000), with the negative effect of trade exposure decreasing the supply of "marriageable" men (Wilson 1987). Nevertheless, unfavorable opportunities in the labor market for women are expected to increase marriage rates, as the decline in paid work opportunities improves the relative gains from domestic work specialization for women. Finally, changes in social norms have challenged the gains of household specializations predicted by the neoclassical theory (Kearney and Wilson 2017). Family formation might become inelastic to changes in economic conditions if women do not specialize in domestic activities.

4 Institutional Background and Data Sources

During most of the 1900s, Brazil was one of the world's most heavily protected economies, with a trade policy based on deliberate import substitution. While Brazil's trade policy had historically been coincident with long periods of strong economic growth, it became clear by the 1980s that the policy was no longer sustainable (Kovak 2013). Beginning in the early 1990s, Brazil initiated a major unilateral trade liberalization process, when the administration of the newly elected President, Fernando Collor, unexpectedly eliminated virtually all non-tariff barriers and started a gradual reduction in import tariffs. Starting in 1990, non-tariff barriers and special regimes were eliminated and typically immediately replaced by equivalent import tariffs, in a process known as "tariffication". While this process left the actual protection structure unaltered, it allowed the federal government to use tariffs as the main instrument for trade policy. At the same time, the government established a timeline for the gradual reduction of tariffs, which was approved and implemented. The trade liberalization process happened quickly, and, by the end of 1993, the major phase of tariff reductions had already taken place. In a further movement toward openness, the next elected government of President Fernando Henrique Cardoso reduced some additional tariffs in 1994, as part of a broader effort towards monetary stabilization. Overall, one can see the tariffs in 1990 as accurately reflecting the historical levels of trade protection in Brazil, and the reductions in tariffs between 1990 and 1995 as capturing the main implications of the reform in terms of exposure of the domestic industry to foreign competition. These phased tariff reductions were implemented with the goal of reducing average tariff levels and reducing the dispersion of tariffs across industries, in hopes of reducing the gap between internal and external costs of production.

This paper uses data on industry-specific tariff changes between 1990 and 1995 provided by Kume et al. (2003). This data has been extensively used in the previous literature on the impact of trade liberalization and labor markets in Brazil (Kovak 2013, Dix-Carneiro and Kovak 2015, Hirata and Soares 2015, Dix-Carneiro et al. 2018 and Gaddis and Pieters 2017). Nominal tariff cuts varied significantly across industries. For instance, apparel and rubber faced tariff reductions of more than 30 percentage points, while agriculture and petroleum faced only small tariff changes (Figure 1). Because tariff cuts were greater for industries that were more protected pre-liberalization, there is little scope for endogeneity concerns that might occur if tariff cuts were driven by industry performance or political preference (Figure 2).

4.1 Exposure to Trade Liberalization

This paper explores the heterogeneous effects of trade liberalization across regions of the country. For this purpose, I use a measure of tariff exposure which effectively captures the degree to which trade liberalization affected labor demand in each microregion of the country (Kovak, 2013):

$$Trade.Shock_r = -\sum_i s_{ri} d\ln(1+t_i)$$

$$\text{with } s_{ri} = \frac{\frac{\lambda_{ri}}{\theta_i}}{\sum_{i' \in E} \frac{\lambda_{ri'}}{\theta_{i'}}}$$

where $d\ln(1+t_i)$ is the log difference of the tariff rate in industry i from 1990 to1995, λ_{ri} is the initial share of workers in region r employed in industry i, θ_i equals the wage bill share of industry i, and E is the set of all tradable industries. Different from most papers in this literature (e.g. Kovak 2013), I multiply the tariff declines by minus one to interpret coefficients as a response to a higher tariff exposure. s_{ri} is the effective weight that tradable industry i has in the total employment of all tradable sector of region r. Note that $s_{ri} > 0$ and $\sum_i s_{ri} = 1$ for every r.

One of the advantages of this approach is to exclude the non-tradable sector from the analysis, and to rescale employment shares to sum to unity over traded sectors only. Kovak (2013) shows that because non-tradable output must be consumed within the region where it is produced, non-tradable prices move together with the prices of locally produced tradable goods. As a result, the magnitude of the trade-induced regional shock depends only on how the local tradable sector is allocated.[2]

The trade shock variable shows substantial geographic variation across the country (Figure 3). To illustrate this variation, Figure 4 shows the initial industry distribution of employment for the region with the main cities of Volta Redonda and Petropolis. The industries on the x-axis are sorted from the most positive to the most negative tariff change. Both regions are in the Rio de Janeiro state and are less than 100 miles from each other, but there was a substantial difference in industry composition between these regions before trade liberalization. The Volta Redonda region produced mostly metal goods and suffered a lesser impact from trade liberalization than the Petropolis region, which produced mostly apparel products. The identification strategy will consist of comparing the marriage and fertility outcomes across the more and less exposed regions, before and after trade liberalization.

4.2 Data on Labor Market, Fertility, and Marriage Outcomes

The unit of analysis of this study is a microregion, a grouping of contiguous municipalities with similar economic characteristics within a state, resembling a local labor market or commuting zone. Following the literature that studies the impact of trade liberalization in Brazil, I use constructed microregions that are consistently identifiable from 1980 to 2010 (Dix-Carneiro et al. 2018 and Dix-Carneiro and Kovak 2017).

The final sample contains 411 microregions.[3] I obtain labor market, fertility, and marriage outcomes from the four waves of the Brazilian Demographic Census covering the years 1980, 1991, 2000 and 2010. For the marriage and fertility outcomes, I restrict the sample to women between the ages of 20 and 35, aiming to estimate the impact of trade liberalization on a population that has made fertility and marriage decisions within the past 10 years.

In terms of labor market outcomes, I look at the share of 20 to 35-year-old men and women working for pay at the microregion level, and the share working in manufacturing jobs. For marriage outcomes, I focus on the share of married, cohabiting, and never married women within the 20-

[2] Topalova (2010) suggests using nontradables as an additional sector, with tariffs being assigned zero over the entire period. The issue with this measure is that employment in the non-tradable sector at baseline is highly correlated with initial female labor force participation and therefore likely to be correlated with fertility outcomes. See Gaddis and Pieters (2017) further discussion and details.

[3] The region containing the Manaus free trade zone is not included since it was exempt from tariffs and unaffected by the tariff changes occurring during liberalization.

35 age range. For fertility outcomes, I look at the share of women with no children, the average number of children per woman and the average age of first-time mothers. While the age of first-time mothers at the time of birth is not asked in the Census, I estimate this variable by calculating the difference between the age of the mother and the age of the oldest children living in the household.[4] Finally, low-skilled individuals are defined as those with less than a high school degree.

The changes in the main outcome variables for 20 to 35 year old women during the period of analysis are shown in Table 1. Throughout the period, the share of young women who are married decreased, while the share of women cohabiting, divorced or separated increased. I also find that the share of women having never married remained fairly constant.[5] The increase of divorce rates in Brazil can be explained by social and political changes that happened during the period, such as the introduction of divorce legislation in 1977, a decline in church attendance, and the spread of access to media and information (Chong and Ferrara 2009). The table also shows an increase in the share of women with no children and a decline in the fertility rate consistent with the demographic transition the country experienced during the period (Lam and Marteleto, 2005; La Ferrara et al., 2012). I also show the geographic distribution of changes in the share of young women with no children between 1991 and 2000 in Figure 5. Finally, Brazil experienced an increase in both diploma attainment, represented by the increase in the share of young women with a high school degree or more (i.e. high skilled), and in the share of young women working for pay. The share of young women working in manufacturing jobs was low and did not change significantly during the period.

5 Estimation Strategy

The empirical strategy used in this paper follows Kovak (2013) and Autor et al. (2013). I estimate the impact of trade shocks using a difference-in-differences model:

$$y_{rs,t} - y_{rs,1991} = \beta_0^t + \beta_1^t Trade.Shock_{rs} + \beta_2^t X_{r,s1991} + \beta_3^t y_{r,s1980} + \gamma_s + \varepsilon_{rs,t} \quad (1)$$

[4] I assume that the oldest child is alive and living in the same household as the mother, which is likely given the focus on young mothers. The infant mortality rate in Brazil in 1991 is relatively low, with 62 deaths per 1,000 live births (World Bank Development Indicators 2018).

[5] The share of never married women remaining constant is consistent with small changes of age of marriage in Brazil during the period. The average age at first marriage was 22.7, 22.5, 23.1 and 29.7 for the years 1980,1991,2000, 2010 respectively (World Bank Indicators).

where $y_{rs,t} - y_{rs,1991}$ is the change in outcome y in microregion r in state s between 1991 and t=2000, 2010. For example, this expression could indicate the change of the share of women never married in the Volta Redonda microregion between 1991 and 2000. $Trade.Shock_{rs}$ is a measure of the tariff exposure shock faced by microregion r between 1991 and 1995 described in subsection 4.1. $X_{rs,1991}$ is a set of characteristics of region r in 1991. It includes educational attainment of the adult population, share of the population in rural areas and share of the population age 20 to 35. These controls have been used in literature and are likely to be related with the trade exposure (Autor et al. 2017). For example, regions more exposed to trade shocks in Brazil during the period, were more likely to be urban and educated than the regions which were less exposed. These controls account for the possibility that such regions are in different fertility and marriage trajectories during the period. I also control for $y_{r,s1980}$, which is the outcome of interest measured in 1980.[6] Finally, γ_s are state fixed effects, and I compare the effect of trade shocks across microregions within the same state. I also present robustness checks of the main findings of the paper, where I do not control for baseline characteristics, state fixed effects or pre-trade liberalization measure of the outcome. I cluster standard errors at the meso-region level to account for potential spatial correlation in outcomes across neighboring microregions and weigh the regressions by the microregion population in 1991.[7]

The identification comes from a parallel trends assumption: regions varying in the degree of exposure to trade liberalization can be inherently different in terms of the outcome of interest, but this difference cannot change over time. While I control the regressions for lagged $y_{r,s1980}$ to account for different trends in the outcome before trade liberalization, I also test the parallel trends assumption by estimating the effect of trade shocks on changes in outcomes between 1980 and 1991.

Another potential threat to this identification strategy is that families might move away from more exposed regions as a response to trade liberalization. However, there is strong evidence of imperfect interregional labor mobility in Brazil that justifies the persistent negative employment shocks associated with trade exposure (Dix-Carneiro and Kovak 2017). I corroborate this evidence

[6] Papers in the literature have used $\Delta y_{rs,1991-1981}$ to account for pre-existing trends that could be related to (future) trade shocks (Kovak 2013 and Dix-Carneiro et al. 2018). The issue with this approach is that y_{i1991} appears both in the right and left hand side of the estimating equation, potentially introducing bias and contaminating all of the remaining coefficients.

[7] Meso-regions are 91 groups of micro-regions defined by the Brazilian Statistical Agency IBGE which have been used for clustering standard errors in the literature (Dix-Carneiro et al. 2018).

in my sample of interest by confirming that young women in Brazil are not likely to migrate as a response to trade liberalization shock.

6 Results

6.1 Negative Effect of Trade Exposure on Employment

The first step of the analysis is to quantify the effect of trade liberalization on the labor market outcomes of young men and women in Brazil. Consistent with literature (Dix-Carneiro and Kovak 2017), I find persistent negative effects of trade exposure on employment outcomes of men and women (Table 2, Panel A). A median decline in tariffs during trade liberalization (7.5 percentage points) is associated with a 3.8 percentage-point decline in the share of young men and 2.9 percentage-point decline in the share of women working for pay in the year 2000. The negative effects for men are persistent even 20 years after trade liberalization, with a median decline in tariffs being associated with a 4.4 percentage point decline in the share of young men working for pay and a 2.9 decline in the share of women working for pay in 2010. This result is consistent with the findings from Gaddis and Pieters (2017), who estimate a more adverse impact of trade liberalization on the employment level of men than that of women.[8]

I also investigate the effect of trade liberalization on the share of men and women working in manufacturing jobs. As shown in Figure 1, manufacturing was the tradable sector which was the most affected by foreign competition during the period. I find a substantial decline in the share of men working in manufacturing jobs in regions more exposed to trade shocks both in the medium and long-run. A median decline in tariffs during trade liberalization is associated with a 5.1 percentage-point decline in the share of young men working in manufacturing in 2000 and 9.3 percentage-point in 2010. This result is consistent with the expectation that the sectors most exposed to foreign competition experienced the highest drop in employment after trade liberalization. I find no effect of trade liberalization on the employment of women in the manufacturing sector, however only a small share of young women worked in manufacturing at baseline.

[8]Gaddis and Pieters (2017) do not find significant employment differences by gender in log terms. Nonetheless the percentage-point gender gap in employment is more relevant for this study, as I am interested in how family structure changes with a decline in *levels* of marriageable men.

6.2 Trade Exposure is Associated with Declines in Fertility Rates

I now turn my attention to the impact of trade liberalization on the fertility choices of young women. First, I classify each microregion of the country as more exposed to trade (tariff decline higher than the median tariff decline) or less exposed to trade (tariff decline lower than the median tariff decline). I show the evolution in the share of young women with no children across regions over the Census years on Figure 6. Overall, I find that women in regions more exposed to trade are less likely to have children before trade liberalization, which is consistent with the fact that those regions are more urban and industrialized. Nonetheless, the gap between regions tends to increase in the years 2000 and 2010, suggesting that trade liberalization might be associated with a decline or delay in the likelihood of having a first child for young women. The figure also presents a visual test for the parallel trends assumption, showing no substantial change in the fertility gap across regions more and less exposed to trade liberalization between 1980 and 1991.

The medium and long-term effects of trade liberalization on fertility outcomes of young women are shown in Table 3. Across all models, I find trade liberalization increases the likelihood that young women have no children. I estimate smaller effects between 1991 and 2000, with a median decline in tariff exposure during the trade liberalization associated with a 1.6 percentage-point increase in the share of young women with no children. (column 2, panel A). This result is consistent with evidence from the United States (Schaller 2016, Autor et al. 2017 and Kearney and Wilson 2017). I estimate an even greater long-term effect of trade liberalization on fertility in column 4, where I find that a median decline in tariff exposure increases the share of women with no children by 1.9 percentage points in 2010 (column 4).

Next, I turn to the effect of trade liberalization on the number of children (panel B). I find significant negative effects of trade liberalization on the average number of children in the year 2000 (columns 1 and 2). I also find significant long-term effects of trade liberalization (columns 3 and 4). A median tariff decline during the period is associated with a decrease of 0.12 children per young woman. Finally, I investigate whether changes in fertility are driven by a mother postponing childbearing (Panel C). I do not find significant effects of trade exposure when it comes to the age of first-time mothers, both in the short and long-run, suggesting that fertility changes are permanent and not just a result of women postponing childbearing. This evidence is confirmed in Table 4, where I estimate the effect of trade exposure for women who are 20 to 27 years old and 28 to 35 years old separately. If anything, I find that the effects of trade liberalization are stronger for older women.

6.3 Marriage Rates are Not Affected by Trade Exposure

I now investigate how the unilateral trade liberalization affected marriage decisions of young women in Brazil. Figure 7 shows the evolution in the share of young women married or cohabiting across regions more or less exposed to trade over the Census years. There is an overall decline in the share of women married or cohabiting over the years, but the gap between regions has remained unchanged overtime. This result suggests that trade liberalization might have had a small effect on marriage and cohabiting decisions. I test this hypothesis in Table 5.

Panel A presents the effect of trade shocks on the share of 20-35 year old women who are married, Panel B shows the effect on the share of women who are cohabiting, and Panel C on women who were never married. I find generally small and statistically insignificant coefficients for the tariff exposure shock on marriage outcomes both in the medium and long-term. Trade shocks are only marginally significant for the share of young women never married in 2010. Overall, despite the substantial changes in fertility of young women associated with trade liberalization shocks, there is not much evidence that young women in Brazil change their marriage decisions as a response to trade shocks. In addition, there is no evidence that women are postponing their marriage and cohabitation decisions, as I find zero effects of trade exposure on marital outcomes across different age ranges (Table 6).

This result contradicts some evidence for the United States, which suggests that negative labor market demand shocks for men are expected to decrease marriage rates (Blau et al. 2000 and Autor et al. 2017). However, more recent literature has shown the importance of social norms as well as economic conditions in their effect on family formation outcomes (Kearney and Wilson 2017). In the past, a stable job prospect for the husband was a necessary condition for couples to get married, with women assuming responsibility for housework. However, the results show that marriage and cohabitation decisions are less sensitive to negative economic shocks in recent decades.

In Tables 3 and 5, I estimate the effect of trade liberalization on marriage and fertility outcomes using a difference-in-difference specification controlling for municipalities characteristics at baseline and state fixed effects. However, my main results are robust to this control choice as well as weighting the observations by microregion population in 1991. In all specifications in Table 7, I estimate that trade exposure is associated with an increase in the medium and long-run share of young women with no children. I also find nothing significant in the share of married and cohabiting women in regions more exposed to trade shocks.

6.4 Pre-Trend Tests

One important concern in any difference-in-difference estimation strategy is the existence of pre-treatment trends in the outcome of interest. In the framework of this paper, the issue is whether more exposed regions experienced a rapid decline in fertility before trade liberalization than less exposed regions. In Tables 3 and 5, I control for 1980 measures of the outcome of interest to rule out the possibility that the estimated effects of trade exposure were driven by a correlation between pre-existing trends and future regional tariff changes. I also directly test the parallel trend hypothesis by estimating the effect of future trade liberalization on past changes on the outcomes of interest between 1980 and 1991 in Table 8. In these specifications, I measure the controls used in this regression at baseline (1980) as well as weight the observations by the microregion population in 1980. Overall, I find little evidence that future trade exposure is associated with past fertility and marriage outcome changes. I estimate non-significant effects of future trade exposure on all six outcomes used in this paper. This result prevents pre-trends in fertility and marriage outcomes from driving the main results of this paper.

7 Mechanisms

In this section, I investigate the reasons trade shocks cause changes in the fertility decisions of young women. Specifically, I turn to unpacking the causal relations between trade exposure, male and female employment changes, and fertility. To answer this question, I use the causal mediation analysis developed in Imai et al. (2011). While very popular among political scientists, causal mediation analysis has been increasingly used in Economics (e.g. De Mel et al. 2013, De Mel et al. 2014, Dippel et al. 2015 and Dippel et al. 2017).

The mediator effects of employment changes on fertility are outlined in Figure 8. This simple graphical representation shows the decomposition of the causal effects of trade shocks on fertility through employment changes (mediator) or some other factors, such changes in the provision of local public goods. In the figure, our object of interest is $(i) \times (ii)$, which represents the effect of trade exposure on fertility that works through observed male and female employment adjustments.

To estimate mediation effects, the first step is to distinguish direct and indirect effects by estimating the following two linear regressions:

$$\Delta M_{rst} = \alpha_0^t + \alpha_1^t Trade.Shock_{rs} + \alpha_2^t X_{r,s1991} + \gamma_s + u_{rs,t} \tag{2}$$

$$\Delta Y_{rst} = \delta_0^t + \delta_1^t Trade.Shock_{rs} + \delta_2^t \Delta M_{rst} + \delta_3^t X_{r,s1991} + \gamma_s + \epsilon_{rs,t} \tag{3}$$

where ΔY_{rst} is the change in fertility outcomes between 1991 and the reference year t, and ΔM_{rst} is the mediator. In this setup, the average causal mediation effect (ACME) is calculated as $\alpha_1^t \delta_2^t$. Imai et al. (2010) show that this ACME can be non-parametrically identified without functional form or distributional assumptions under a sequential ignorability assumption. Using a 'Holland-Rubin potential outcomes' notation of causal inference, the formal sequential ignorability conditions are:

$$\{\Delta Y_i(t', m), \Delta M_i(t)\} \perp\!\!\!\perp Trade.Shock_i | X_i = x, \gamma_s \tag{4}$$

$$\Delta Y_i(t', m) \perp\!\!\!\perp \Delta M_i(t) | Trade.Shock_i = t, X_i = x, \gamma_s \tag{5}$$

The first assumption is that, given the observed pretreatment confounders X_i, the treatment assignment is assumed to be statistically independent of potential outcomes and potential mediators. This part of the assumption is often called no-omitted-variable bias, exogeneity, or unconfoundedness. It means that the treatment $Trade.Shock$ is exogenous conditional on controls X_i. The empirical strategy presented in section 5 and the results discussed in section 6 rely on this assumption, and I apply the same estimation strategy in the mediation analysis.

The second part of sequential ignorability is a new assumption. It implies that the observed mediator is ignorable given the trade exposure and X_i. In other words, there are no unobserved variables that affect both employment and fertility outcomes after conditioning on the trade exposure and controls. An example where the sequential ignorability assumption is not satisfied is given in Figure 9, where the relation (iii) implies that there are factors affecting both employment changes and fertility decision. For example, if the federal government decides to compensate areas more exposed to trade liberalization with higher investment in infrastructure, those investments could potentially affect labor market outcomes but also fertility decisions of young women in the region.

To address the possibility that sequential ignorability assumption may not hold, Imai et al. (2010) develop a sensitivity analysis under the framework of the structural linear equations model. The propose analysis estimate different mediation effects under an unobserved confounder of various magnitudes. I also present the results of this sensitive analysis below.

7.1 Male and Female Employment Mediators

In this section, I decompose the causal effects of trade shocks on fertility through different mediation effects. The first column of Table 9 investigates the importance of male employment as a mediator of trade exposure on fertility decisions between 1991 and 2000. In Panel A, the first stage coefficient is the direct effect of trade shocks on the change in the share of male employment between 1991 and 2000, as described in equation 2. Note that this coefficient was presented in Table 2 and discussed in section 6. In Panel B, I find that trade exposure increased the probability that a woman has no children by 3.5 percentage points because of changes in the employment of young men during this period. From this estimation, one can conclude that about 18% of the total effect of tariff exposure on the share of women with no children was mediated through male employment. When looking at the changes in the average number of children, I estimate that 24% of the effect of trade shocks on this outcome can be explained by changes in male employment during the period.

The change in employment of young men is also a strong mediator for the long-term effects of trade liberalization on the fertility choices of young women. In the third column of Table 9, I estimate that 22% of the effect of trade liberalization on changes of the share of women with no children between 1991 and 2010 is mediated through changes in the employment of young men. I also find that the trade exposure decreased the average number of children by -0.263 because of changes in male employment between 1991 and 2010, which means that 18% of the effect of trade shocks on this dependent variable is mediated through this channel.

I find that changes in the employment of young women is a weaker mediator of the impact of trade liberalization on fertility both in the medium and long term. While trade exposure had a substantial effect on the employment of young women, I estimate that only 4% of the total effect of tariff exposure on the share of women with no children between 1991 and 2000 can be explained by changes in the employment of female workers (in the second column of Table 9). Income effects on the demand for children were high enough to compensate for the increase in fertility associated with declines in young women's opportunity cost to have children.

These findings are rationalized by neoclassical fertility decision models. The strong and persistent negative labor shock for men is likely to substantially decrease the household income. The weaker negative labor market shock for women had opposite substitution and income effects, which likely cancel each other. As a result, changes in the labor market opportunities for men is the main driver of changes in the fertility of women in more exposed regions.

7.2 Sensitivity Analysis

The previous results indicate that male employment is likely a mediator of the effect that trade exposure has on fertility choices. However, these findings are obtained under the sequential ignorability assumption, which implies that we have fully accounted for any confounding factors that might have an effect on the mediator or outcome. Therefore, one must ask whether regions experiencing declines in male employment due to trade exposure have unobserved characteristics that also influence fertility decisions. For example, if the federal government decides to compensate areas more exposed to trade liberalization with higher public spending, those investments could potentially affect that region's labor market outcomes as well as fertility decisions of young women.

I assess the sensitivity of the estimated ACME to unmeasured confounding factors (Imai et al. 2010). The method relies on the fact that under equation (2) and (3), one can summarize the degree the ignorability assumption was violated by the importance of an unobserved confounder in explaining the observed variation in the mediator and outcome variables. The sensitivity analysis is based on how much the omitted variable would alter the coefficients of determination (R-squared) of the mediator and outcome models. For example, if federal investment is important in determining both male employment and fertility decisions, then the model excluding federal investment will have a much smaller R-squared value compared to the full model including the variable. However, if federal investment is not a strong determinant of the mediator or the outcome, the model excluding federal investments will have a R-squared value that is similar to the full model. Thus, the relative change in R-squared can be used as a sensitivity parameter for the estimated ACME.

The results are presented in Figure 10, where the true ACME is shown as contours with respect to the proportions of the variance in the mediator (horizontal axis) and in the outcome (vertical axis), each explained by the unobserved confounder in the true regression models. In the Figure, I explore the case where the unobserved confounder affects the male employment and fertility in opposite directions, which is what we would expect to upward bias the ACME. Panel A shows the sensitivity analysis where the outcome is changes in the share of women with no children and Panel B looks at changes in the average number of children. In both panels, the graphs on the left display changes between 1991 and 2000 and the graphs on the right display changes between 1991 and 2010. In all graphs, the mediator is change in male employment.

Looking at changes in the share of women with no child between 1991 and 2000 (Panel A, left graph), I find that the ACME can be estimated zero if the product of these two proportion is equal to 0.036. For example, the real ACME is zero if the federal investments explain about 20% of the change in employment between 1991 and 2000 and 17.8% of the change in the share of women with no children during the same time period. When looking at changes in the total number of children between 1991 and 2000, (Panle B, left graph), I find that ACME is zero if the unobserved confounder explains about 20% of the variation of male employment and 14.7% of the variation of change in the total number of children per women (or any product of R-squared equal to 0.029). Similar proportions are found when looking at changes in the outcomes between 1991 and 2010 (right graphs). In conclusion, for the true ACME to be equal to zero, one must assume an unobserved confounding factor that explains a substantial variation in male employment and fertility decisions of young women. In all cases, if the confounder were to affect the mediator and fertility in the same direction, then the effects of the mediator would be even greater than the ones estimated in Table 9.

8 Conclusion

The Brazilian unilateral trade liberalization led to sizable declines in employment in regions more exposed to trade, with more negative effects for men than women. This study uses this event to investigate how worse economic opportunities for men and women affected their marriage and fertility outcomes. My analysis suggests that women in areas more affected by trade are less likely to have children even twenty years after trade liberalization. Using a causal mediation analysis, I show that declines in the employment of young men is an important mechanism driving changes in the fertility decisions of young women. Changes in the employment of men explain only about 20% of the medium and long-term effects of trade exposure on the fertility decisions of young women. I also find that changes in female employment are not an important mediator for the effect of trade exposure on fertility outcomes.

This result supports the hypothesis that trade liberalization produces a significant income shock to the families in more exposed regions and couples respond to trade shocks by having fewer children. This paper demonstrates that even in a developing country like Brazil, where women have limited access to contraceptive methods, children are a normal good. The declines in the opportunity costs of bearing a child for women associated with their worse employment prospects were not significant enough to increase their fertility decisions.

I also find evidence that marriage decisions were not sensitive to this significant economic shock. There is no systematic evidence that young women were less likely to be married or cohabiting after trade liberalization in regions more affected by the trade shock. This result contradicts neoclassical theory predictions that the worse labor market opportunities for men are expected to decrease marriage rates (Becker 1981). My interpretation of this finding is consistent with Kearney and Wilson (2017), who state that changes in social norms have challenged the perceived gains of household specialization.

Finally, it is important to keep in mind that our results are obtained using the trade liberalization episode in Brazil. Although it is a developing country, Brazil is generally considered socially liberal compared to other low- and middle-income countries (Stern et al. 2017). Marriage rates might be more elastic with respect to economic activity in more conservative places where there are still perceived benefits from household specialization (Kis-Katos et al. 2017) and fertility rates might be less elastic to economic shocks in regions where women do not have access to contraceptive methods.

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Tables

Table 1: Descriptive Statistics - Women 20-35 years old

	Census Years			
Variable	**1980**	**1991**	**2000**	**2010**
Marital Status				
Never Married	29.6%	29.1%	28.7%	30.3%
Married	56.8%	49.4%	37.5%	28.9%
Cohabitating	8.7%	14.3%	23.5%	29.0%
Separate & Divorced	2.9%	5.4%	9.9%	11.5%
Fertility				
No Children	33.1%	31.1%	34.0%	40.7%
Total Children	1.95	1.68	1.43	1.14
Age First Child	21.6	21.3	21.1	21.1
Socio Economic Status				
High Skilled	16.6%	24.7%	33.4%	53.4%
Working for Pay	34.1%	41.2%	46.2%	56.5%
Working for Pay in Manufacturing	6.1%	6.4%	5.6%	6.3%

Note: High-skilled are women with a high school degree or more. Source: Brazilian Census

Table 2: The Effect of Trade Liberalization on Employment

Panel A	**Dependent Variable**			
	Change Share Working for Pay, 2000-1991		Change Share Working for Pay, 2010-1991	
Sample	Male	Female	Male	Female
Tariff Exposure Shock	-0.503	-0.386	-0.585	-0.394
	(0.149)***	(0.113)***	(0.193)***	(0.143)***
R-squared	0.621	0.697	0.717	0.750
Panel B	**Dependent Variable**			
	Change Share Working for Pay in Manufacturing, 2000-1991		Change Share Working for Pay in Manufacturing, 2010-1991	
Sample	Male	Female	Male	Female
Tariff Exposure Shock	-0.682	-0.194	-1.242	-0.334
	(0.194)***	(0.143)	(0.300)***	(0.194)*
R-squared	0.679	0.720	0.765	0.792
Observations	411	411	411	411

Sample: Men and Women Age 20-35 years old. Additional Controls: State Fixed-Effects, Share of Adults at each Education Attainment Level, Share Rural Population, Share Population Age 20-35 at baseline (1991). Observations are weighted by the microregion population at baseline (1991). Robust standard errors reported in parentheses are clustered at the mesoregion level. ***<0.01, **<0.05, *<0.10.

Table 3: The Effects of Tariff Exposure Shock on Fertility Outcomes

Panel A	**Dependent Variable**			
	Change Share no Children, 2000-1991		Change Share no Children, 2010-1991	
Tariff Exposure Shock	0.194	0.214	0.193	0.246
	(0.057)***	(0.059)***	(0.081)**	(0.074)***
Share No Children, 1980		-0.078		-0.211
		(0.044)*		(0.064)***
R-squared	0.589	0.597	0.762	0.780
Panel B	**Dependent Variable**			
	Change Number of Children, 2000-1991		Change Number of Children, 2010-1991	
Tariff Exposure Shock	-0.489	-0.530	-1.484	-1.568
	(0.256)*	(0.225)**	(0.370)***	(0.265)***
Number of Children, 1980		-0.148		-0.300
		(0.021)***		(0.036)***
R-squared	0.809	0.830	0.854	0.892
Panel C	**Dependent Variable**			
	Change Age at First Child, 2000-1991		Change Age at First Child, 2010-1991	
Tariff Exposure Shock	0.677	0.705	0.092	0.286
	(0.519)	(0.504)	(0.799)	(0.766)
Age at First Child, , 1980		-0.012		-0.093
		(0.022)		(0.034)***
R-squared	0.448	0.449	0.564	0.581
Observations	411	411	411	411

Sample: Women Age 20-35. Additional Controls: State Fixed Effects, Share of Adults at each Education Attainment Level, Share Rural Population, and Share Population Age 20-35 at baseline (1991). Observations are weighted by the microregion population at baseline (1991). Robust standard errors reported in parentheses are clustered at the mesoregion level. ***<0.01, **<0.05, *<0.10.

Table 4: The Effects of Tariff Exposure Shock on Fertility Outcomes by Women's Age

Panel A	**Dependent Variable**			
	Change Share No Children, 2000-1991		Change Share No Children, 2010-1991	
Sample	20-27 year old	28-35 year old	20-27 year old	28-35 year old
Tariff Exposure Shock	0.139	0.146	0.027	0.360
	(0.092)	(0.055)***	(0.108)	(0.073)***
R-squared	0.535	0.430	0.699	0.772
Panel B	**Dependent Variable**			
	Change Number of Children, 2000-1991		Change Number of Children, 2010-1991	
Sample	20-27 year old	28-35 year old	20-27 year old	28-35 year old
Tariff Exposure Shock	-0.382	-0.439	-0.569	-2.805
	(0.234)	(0.297)	(0.261)**	(0.338)***
R-squared	0.603	0.885	0.840	0.597
Observations	411	411	411	411

Additional Controls: State Fixed Effects, Share of Adults at each Education Attainment Level, Share Rural Population, and Share Population Age 20-35 at baseline (1991). Observations are weighted by the microregion population at baseline (1991). Robust standard errors reported in parentheses are clustered at the mesoregion level. ***<0.01, **<0.05, *<0.10.

Table 5: The Effects of Tariff Exposure Shock on Marriage Outcomes

Panel A	**Dependent Variable**			
	Change Share Married, 2000-1991		Change Share Married, 2010-1991	
Tariff Exposure Shock	-0.037	-0.053	0.135	0.065
	(0.086)	(0.081)	(0.165)	(0.139)
Share Married, 1980		-0.076		-0.318
		(0.026)***		(0.044)***
R-squared	0.580	0.591	0.760	0.808
Panel B	**Dependent Variable**			
	Change Share Cohabiting, 2000-1991		Change Share Cohabiting, 2010-1991	
Tariff Exposure Shock	0.081	0.082	-0.080	-0.072
	(0.095)	(0.093)	(0.181)	(0.155)
Share Cohabiting, 1980		-0.054		-0.340
		(0.043)		(0.039)***
R-squared	0.535	0.541	0.752	0.804
Panel C	**Dependent Variable**			
	Change Share Never Married, 2000-1991		Change Share Never Married, 2000-1991	
Tariff Exposure Shock	0.027	0.047	0.068	0.102
	(0.050)	(0.052)	(0.073)	(0.069)
Share Never Married, 1980		-0.106		-0.175
		(0.032)***		(0.046)***
R-squared	0.543	0.562	0.680	0.701
Observations	411	411	411	411

Sample: Men and Women Age 20-35 years old. Additional Controls: State Fixed-Effects, Share of Adults at each Education Attainment Level, Share Rural Population, and Share Population Age 20-35 at baseline (1991). Observations are weighted by the microregion population at baseline (1991). Robust standard errors reported in parentheses are clustered at the mesoregion level. ***<0.01, **<0.05, *<0.10.

Table 6: The Effects of Tariff Exposure Shock on Marriage Outcomes by Women's Age

Panel A	**Dependent Variable**			
	Change Share Married, 2000-1991		Change Share Married, 2000-1991	
Sample	20-27 year old	28-35 year old	20-27 year old	28-35 year old
Tariff Exposure Shock	0.103	-0.113	0.181	0.015
	(0.103)	(0.095)	(0.145)	(0.173)
R-squared	0.622	0.450	0.805	0.734
Panel B	**Dependent Variable**			
	Change Share of Cohabiting, 2000-1991		Change Share of Cohabiting, 2010-1991	
Sample	20-27 year old	28-35 year old	20-27 year old	28-35 year old
Tariff Exposure Shock	-0.009	0.182	-0.120	0.004
	(0.101)	(0.095)*	(0.143)	(0.178)
R-squared	0.564	0.486	0.812	0.754
Observations	411	411	411	411

Additional Controls: State Fixed Effects, Share of Adults at each Education Attainment Level, Share Rural Population, and Share Population Age 20-35 at baseline (1991). Observations are weighted by the microregion population at baseline (1991). Robust standard errors reported in parentheses are clustered at the mesoregion level. ***<0.01, **<0.05, *<0.10.

Table 7: The Effects of Tariff Exposure Shock on Fertility and Marriage Outcomes - Robustness Checks

Panel A - Fertility					Dependent Variable	
Regression Specification	State Fixed Effects	Additional Controls	Share No Children, 1980	Weights	Change Share No Children, 2000-1991	Change Share No Children, 2010-1991
(1)	No	No	No	Yes	0.222 (0.052)***	0.474 (0.064)***
(2)	Yes	No	No	Yes	0.171 (0.031)***	0.390 (0.040)***
(3)	Yes	Yes	No	No	0.169 (0.075)**	0.211 (0.075)***
(4)	Yes	Yes	Yes	No	0.206 (0.074)***	0.285 (0.062)***

Panel B - Marriage&Cohabiting					Dependent Variable	
Regression Specification	State Fixed Effects	Additional Controls	Share Married or Cohabiting, 1980	Weights	Change Share Married or Cohabiting, 2000-1991	Change Share Married or Cohabiting, 2010-1991
(1)	No	No	No	Yes	-0.044 (0.038)	-0.114 (0.063)*
(2)	Yes	No	No	Yes	0.020 (0.024)	0.037 (0.033)
(3)	Yes	Yes	No	No	0.017 (0.064)	-0.011 (0.093)
(4)	Yes	Yes	Yes	No	-0.032 (0.064)	-0.135 (0.093)

Sample: Women Age 20-35. Additional Controls: Share of Adults at each Education Attainment Level, Share Rural Population, and Share Population Age 20-35 at baseline (1991). Weights are defined by the microregion population at baseline (1991). Robust standard errors reported in parentheses are clustered at the mesoregion level. ***<0.01, **<0.05, *<0.10.

Table 8: Pre-Trend Tests

Panel A - Fertility Outcomes	**Dependent Variable**		
	Change Share no Children, 1991-1980	Change Number of Children 1991-1980	Change Age at First Child, 1991-1980
Tariff Exposure Shock	0.090	-0.359	0.607
	(0.084)	(0.531)	(1.155)
R-squared	0.539	0.745	0.500
Panel B - Marriage Outcomes	**Dependent Variable**		
	Change Share Married, 1991-1980	Change Share Cohabiting 1991-1980	Change Share Never Married, 1991-1980
Tariff Exposure Shock	-0.051	0.086	0.000
	(0.132)	(0.085)	(0.081)
R-squared	0.565	0.568	0.589
Observations	411	411	411

Sample: Women Age 20-35. Additional Controls: State Fixed-Effects, Share of Adults at each Education Attainment Level, Share Rural Population, and Share Population Age 20-35 at baseline (1980). Observations are weighted by the microregion population at baseline (1980). Robust standard errors reported in parentheses are clustered at the mesoregion level. ***<0.01, **<0.05, *<0.10.

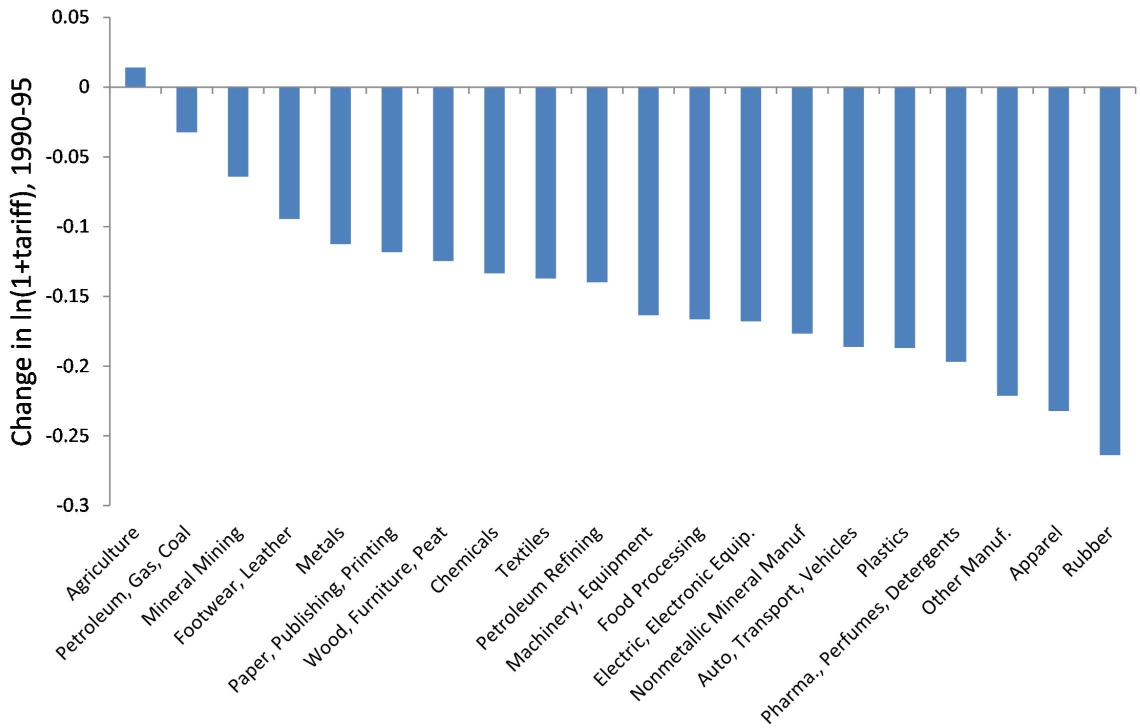
Table 9: Mediation Analysis - Male and Female Employment Changes

Panel A **Mediator**	Change Share Male Working for Pay, 2000-1991	Change Share Female Working for Pay, 2000-1991	Change Share Male Working for Pay, 2010-1991	Change Share Female Working for Pay, 2010-1991
First Stage Coefficient	-0.503	-0.386	-0.585	-0.394
	(0.149)***	(0.113)***	(0.193)***	(0.143)***
Panel B				
Dependent Variable:	Change Share no Children, 2000-1991		Change Share no Children, 2010-1991	
Tot. Eff. of the Tariff Exp. Shock	0.194		0.193	
ACME of Mediator	0.035	0.008	0.043	0.001
% of Tot. Eff. Mediated	17.8%	4.1%	22.1%	0.3%
Panel C				
Dependent Variable:	Change Number of Children, 2000-1991		Change Number of Children, 2010-1991	
Tot. Eff. of the Tariff Exp. Shock	-0.489		-1.484	
ACME of Mediator	-0.120	0.012	-0.263	-0.077
% of Tot. Eff. Mediated	24.5%	-2.5%	17.7%	5.2%

The ACME is calculated as the product of the effect of the exogenous regressor on the mediator and the effect of the mediator on the outcome. The percentage of the total effect that is mediated equals the ACME divided by the total effect. All regressions include State Fixed Effects, Share of Adults at each Education Attainment Level, Share Rural Population, and Share Population Age 20-35 at baseline (1991). Observations are weighted by the microregion population at baseline (1991). Robust standard errors reported in parentheses are clustered at the mesoregion level. ***<0.01, **<0.05, *<0.10.

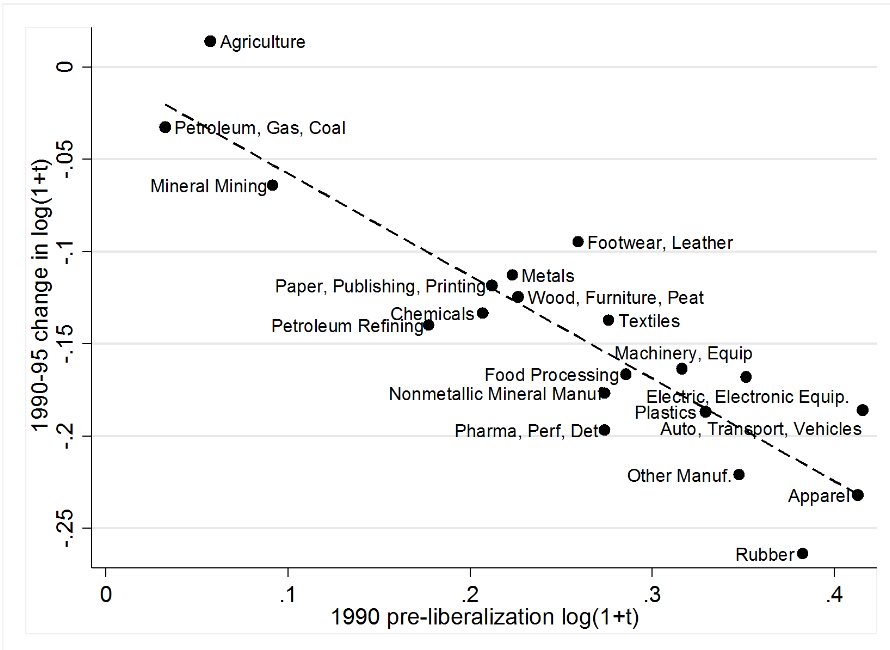
Figures

Figure 1: Tariff Changes by Industry



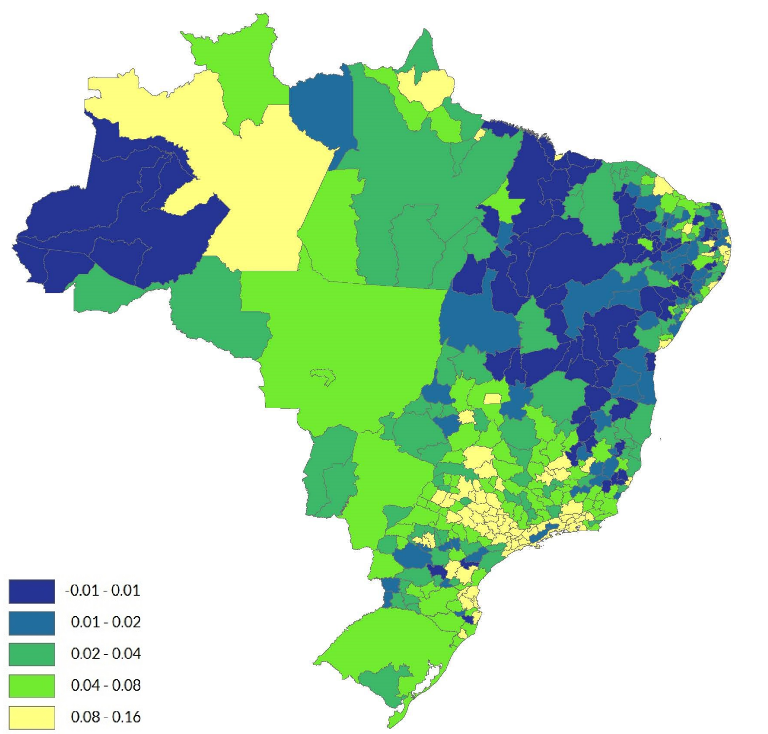
Note: Changes in log(1 + tariff), 1990-1995. Source: Kume et al. (2003)

Figure 2: Most Protected Industries Suffered Major Tariff Cuts



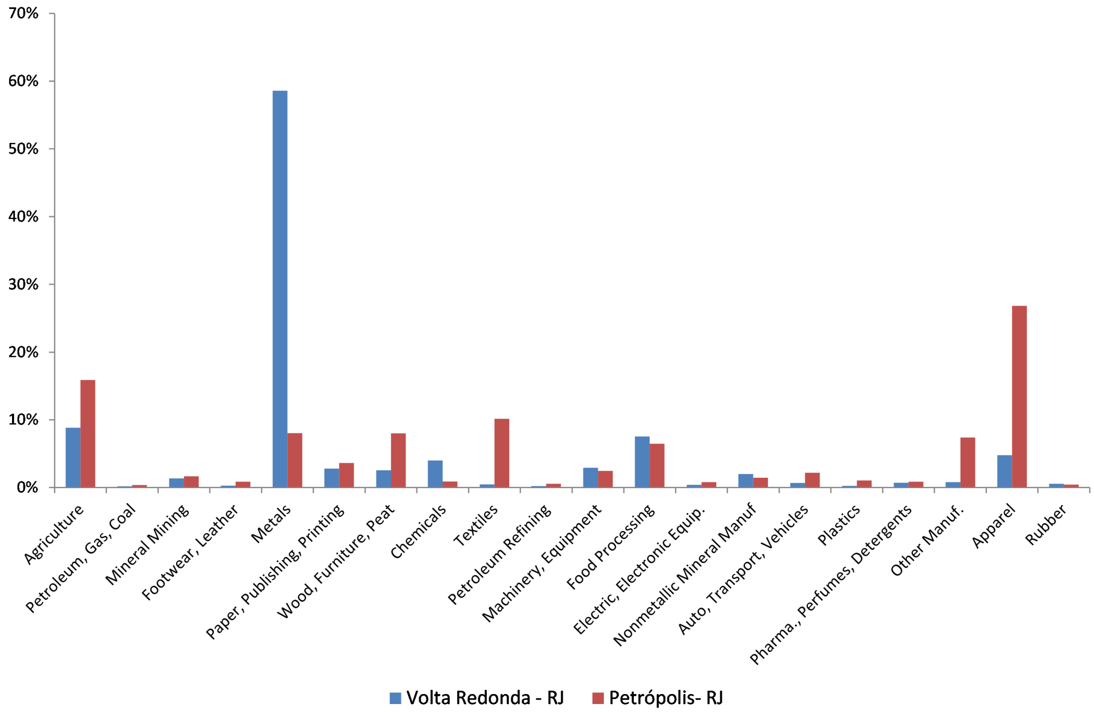
Source: Kovak (2013)

Figure 3: Distribution of Regional Tariff Exposure



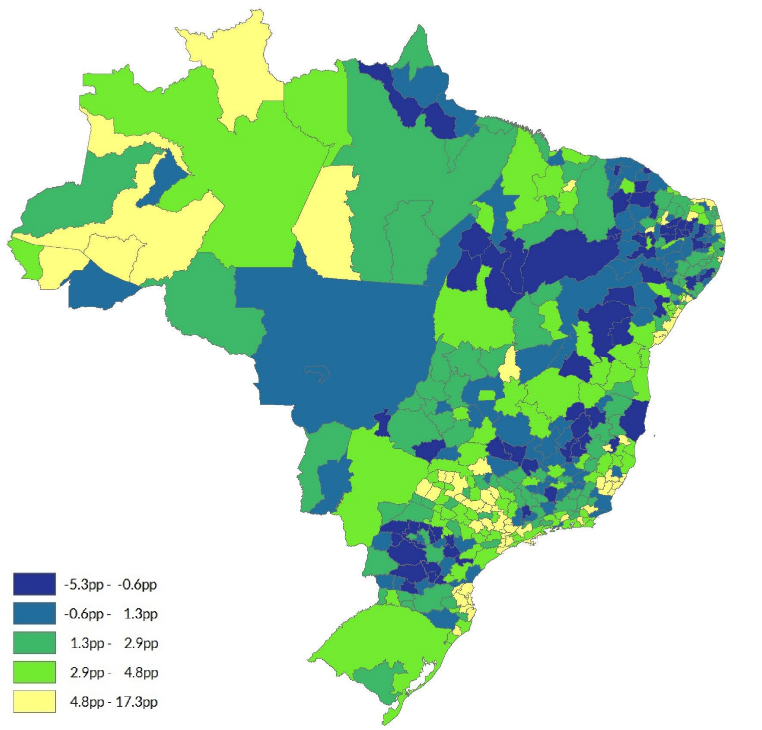
Note: Regional Tariff Exposure computed according to the expression in Section 4.1.

Figure 4: Variation Underlying Regional Tariff Change



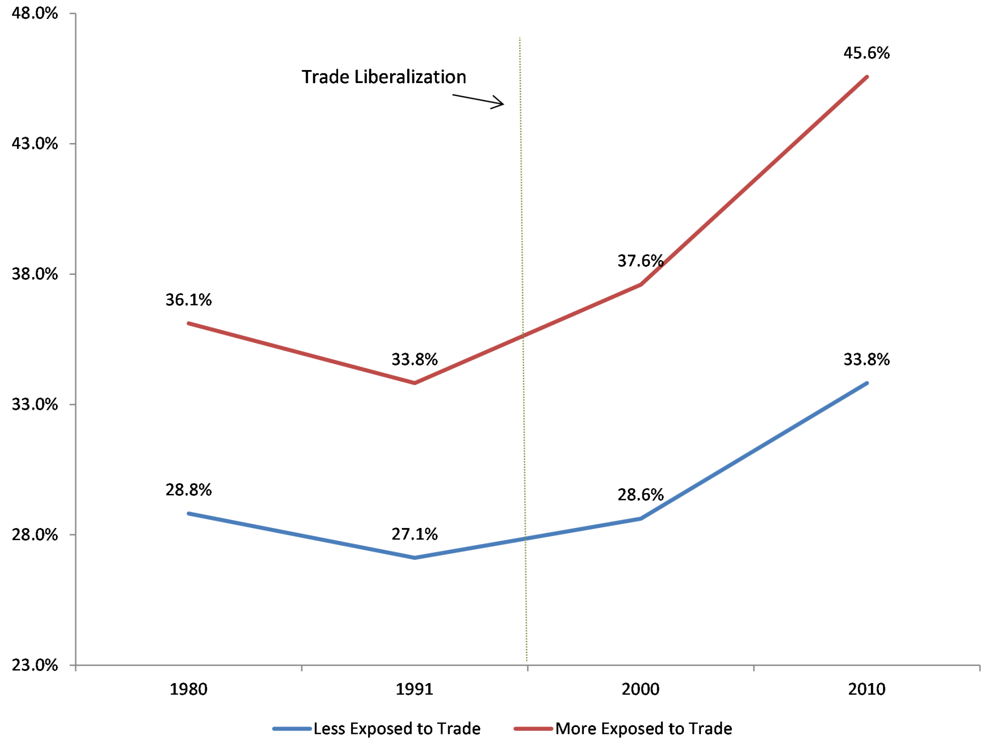
Note: Industry distribution of 1991 employment in the Volta Redonda and Petropolis regional tariff changes. Industries sorted by the tariff change, shown in Figure 1. More weight on the left side of the figure leads to a more negative regional tariff change, and more weight on the right side leads to a more positive regional tariff change

Figure 5: Distribution Changes in the Share of Young Women with no Children, 1991-2000



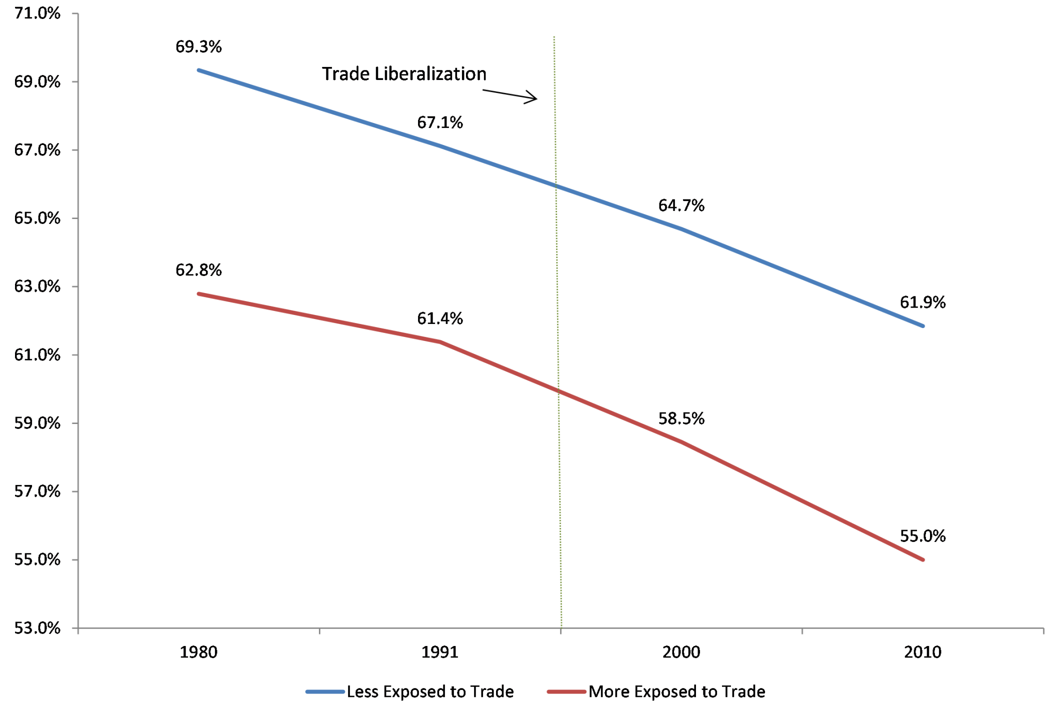
Note: Share of Women 20-35 years old with no children by miroregion from Census.

Figure 6: Share of Women with no Children by Census Year



Note: Sample is restricted to Women 20-35 years old. Regions more (less) exposed to trade are defined as those who experience tariff exposure shock greater (lower) than the median tariff exposure shock.

Figure 7: Share of Women Married or Cohabiting



Note: Sample is restricted to Women 20-35 years old. Regions more (less) exposed to trade are defined as those who experience tariff exposure shock greater (lower) than the median tariff exposure shock.

Figure 8: Mediation Effects

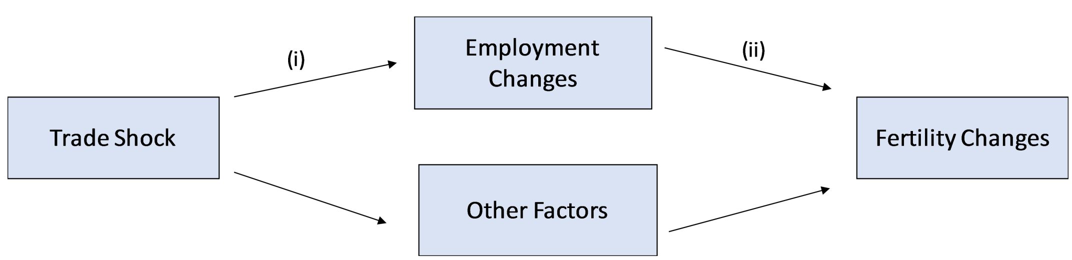


Figure 9: Sequential Ignorability Assumption Failure

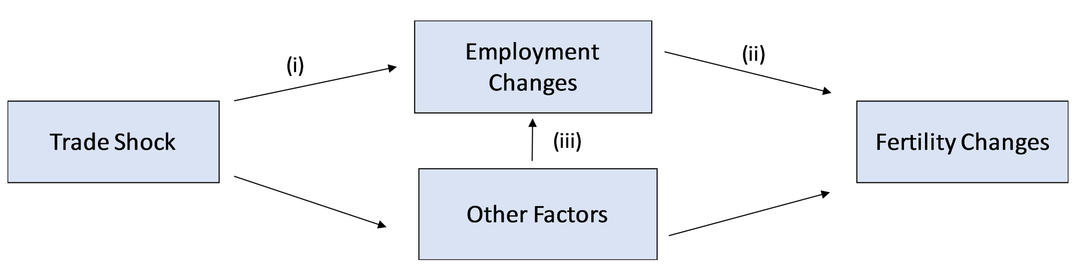
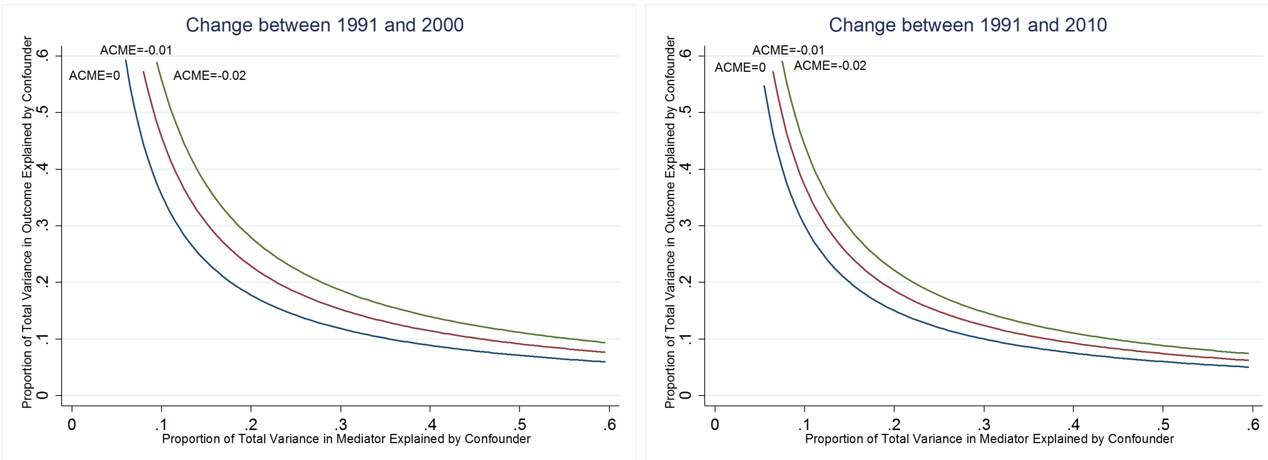
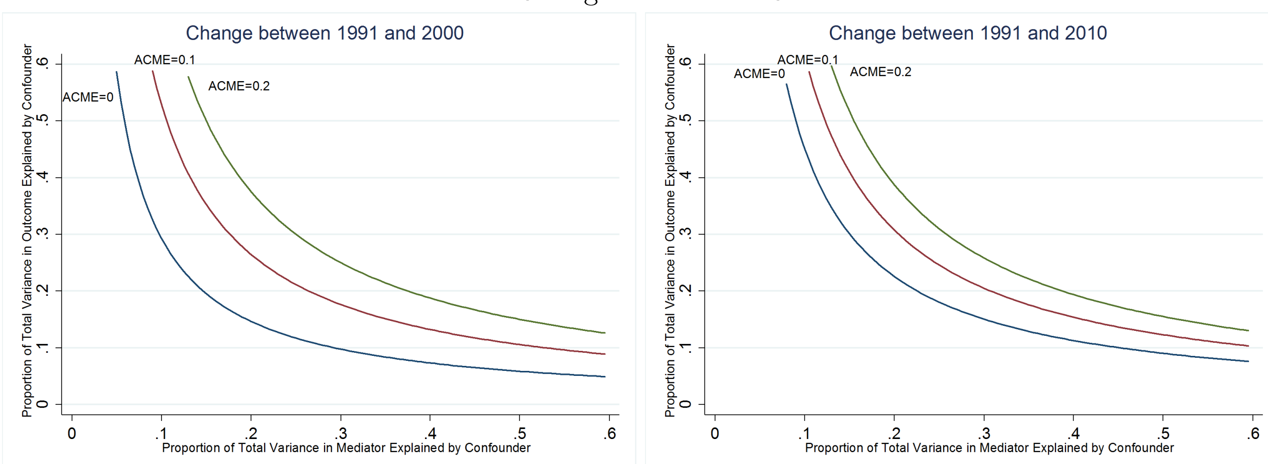


Figure 10: Sensitivity Test

Panel A - Change Share no Children



Panel B - Change Number of Children



Note: The plot presents the results of the sensitivity analysis described in Section 7.2. Each plot contains various mediation effects under an unobserved confounder of various magnitudes. The contours represent the true ACME plotted as a function of the proportion of the total mediator variance (horizontal axis) and the total outcome variance (vertical axis), that are each explained by the unobserved confounder included in the corresponding regression models. The mediator in all graphs is the change in the share of 20-35 year old men employed in the microregion between 1991 and the year of reference. The unobserved confounder is assumed to affect the mediator and fertility opposite directions.