Trade and Intergenerational Mobility: Evidence from the Brazilian Trade Liberalization*

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Abstract

This paper examines the impact of Brazil's 1991 trade liberalization on intergenerational mobility (IGM). To measure the local economy's exposure to the trade shock, we employ a shift-share instrument that combines tariff cuts and employment across industries within a region. Our findings indicate that regions with higher exposure to the exogenous tariff reductions experienced a permanent relative decline in upward occupational mobility and a long-term decrease in absolute income and relative educational mobility compared to regions with less pronounced effects. Additionally, we analyze IGM measures separately for son-father and daughter-mother pairs to account for gender heterogeneity, where daughters suffer higher reductions in income and educational mobility. We find that the overall effect is primarily driven by negative consequences in the labor market, characterized by an increase in the proportion of low-skilled occupations and a decrease in earnings levels following the trade shock.

Keywords: Trade Shocks; Local Economy; Intergenerational Mobility; Brazil.

JEL classification: F13, I23, F14, F16, J62, D12.

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1 Introduction

Enhancing intergenerational mobility (IGM) —the ability of children to achieve a higher socioeconomic status than their parents— over time and across generations is fundamental to economic growth and poverty reduction, especially in developing countries. In this sense, understanding how economic shocks, such as those from trade liberalization, influence IGM is vital. Losses from trade liberalization are often regionally concentrated, with increased import competition leading to a weakening in labor market conditions (Autor, Dorn, & Hanson, 2013; Kovak, 2013) —such as declines in earnings and employment— which, in turn, affect the socioeconomic conditions of households. This situation could limit the opportunities for upward mobility for children from disadvantaged backgrounds.

This paper explores how socioeconomic status persists across generations following Brazil's trade liberalization in 1991. We estimate the impact of the trade reform on income, occupational, and educational mobility. We explore variations at the time of the trade reform in the employment structure across local economies. From 1991 to 2010, regions more exposed to tariff cuts also experienced a relative *decrease* in income and occupational absolute mobility levels and educational relative mobility. Moreover, the effects of these changes differed between sons and daughters.

We argue that this result is explained by shifts in local labor markets, particularly in the demand for skills and the earnings differentials among different skill levels. As has already been documented, after the trade shock in Brazil, regions that experienced significant tariff cuts in their industry specialization suffered a relative and persistent loss in labor market conditions over time (Kovak, 2013; Dix-Carneiro & Kovak, 2017, 2019). Our findings indicate that these effects varied according to skill levels.

Inspired by the literature on international trade gender-specific effects on the labor market (Berik, Rodgers, & Zveglich, 2004; Black & Brainerd, 2004; Menon & Rodgers, 2009; Autor, Dorn, & Hanson, 2019), as these factors may influence women's investments in human capital and fertility, we differentiate between son-father and daughter-mother pairs when estimating the effects of the trade shock on IGM. We acknowledge the empirical challenge of obtaining data on the labor market and educational outcomes for parents and children to measure mobility. To address this issue, we utilize survey datasets from the Brazilian Demographic Census, which provide comprehensive information on parents and children residing in the same household.

To analyze upward income mobility, we employ a linear rank-rank regression. This method estimates the relationship between the income percentile ranks of children within their birth cohort and their parents' ranks in the national income distribution. Following the recent literature (Dahl & DeLeire, 2008; Chetty, Hendren, Kline, & Saez, 2014; Acciari, Polo, & Violante, 2022), we define as our measure of absolute mobility the expected income rank for children whose parents are below the median, specifically at the 25th percentile. Our findings indicate a strong persistence of income levels over the years in Brazil, with significant regional variations. On a national scale, children from parents below the parental median income are expected to reach the 37^{th} income percentile in the children's income distribution. Notably, male children are almost 6 percentile points higher than the expected rank for female children.

This gender difference is highlighted when we consider occupational mobility. In 2010, 35 percent of the sons worked in jobs with higher educational levels than their fathers, and 38 percent of daughters were in better occupations than their mothers. When we compare this fraction across regions, we find higher levels of high-skill occupations in the Center-South than in the North-Northeast.

After calculating regional intergenerational mobility (IGM) measures, we explored how trade liberalization impacts mobility levels. Our analysis indicates that regions heavily impacted by tariff reductions experienced significant declines in occupational mobility and long-term decreases in both absolute income and relative educational mobility, compared to regions less affected. The magnitudes are sizeable. Shifting a region from the 10 to the 90 percentiles in the distribution of the regional tariff cut is related to a 7 percent decrease in absolute income mobility for son-father pairs from 1991 to 2010. For daughter-mother pairs, there was an 11 percent reduction from 1991 to 2000, although this partially recovered over time. In turn, the share of sons with better occupations than their fathers presented a more pronounced reduction through the years (10 percent), while for the daughter-mother pairs, the effect was stable (7.5 percent). Relative educational mobility presented a greater effect of tariff cuts compared to the other measures, where sons' mobility was reduced by 15% and daughters' mobility by almost 20%. Altogether, these results demonstrate that regions confronting higher exposure to the liberalization episode underwent a permanent relative decrease in IGM that was heterogeneous across genders in magnitude.

To ensure the reliability of our findings, we conducted various tests using different specifications. We accounted for potential omitted variables that could be correlated with regional tariff cuts and mobility levels by including controls for pre-liberalization socioeconomic conditions and mesoregion fixed effects. Additionally, we performed a pre-intervention placebo test. We show that regions more exposed to tariff cuts did not display pre-existing trends in mobility changes, supporting our identification assumption. Further, we also show that the co-residency endogeneity of children and parents living in the same residence is not a problem, as the trade shock did not affect the share of independent children. Moreover, we show that differences in the cost of living across areas did not affect our estimates of regional mobility.

Finally, we investigated the mechanisms driving the observed effects of the trade shock on IGM. We examined changes in skill demand within regions, assessing whether diminished mobility levels could be linked to shifts in production activities towards industries that require less skilled labor. We also analyzed the impact on average earnings and earnings differentials within these sectors. In general, the medium-skilled sectors lost space for the low-skill sector, which absorbed most of the changes in labor market shares. We find an increase of 7% in the share of low-skilled occupations and an average reduction of 9% in the share of high-skill occupations relative to low-skilled occupations. We also find that the earnings gap between medium- and low-skilled workers is slightly reduced.

We hope to contribute to the literature estimating IGM by measuring social mobility levels for a developing country marked by significant inequality and poverty. Britto, Fonseca, Pinotti, Sampaio, and Warwar (2022) stands as a pioneering analysis of income mobility in Brazil. Using rich administrative data, they document mobility patterns across different groups and regions. We follow a similar approach with survey data, yielding findings that align with those for absolute income and relative mobility, reinforcing our procedure.

The literature on how trade shocks affect IGM is still incipient. Ahsan and Chatterjee (2017) investigates the 1991 Indian trade reform, suggesting that import competition drives firms at the efficiency frontier to innovate and increase demand for skilled workers, thereby improving job prospects for sons from disadvantaged backgrounds. Similarly, Mitra, Pham, and Ural Marchand (2022) analyzes the impact of massive export expansion from Vietnam to the US subsequent to the US-Vietnam Bilateral Trade Agreement (BTA) in 2001 on intergenerational occupational mobility in Vietnam. Their finding suggests that trade exposure has facilitated considerable upward occupational mobility, enabling children to have better occupations than their parents.

Our paper also dialogues with recent studies that investigate the impact of the 1991 Brazilian trade reform on other economic outcomes, such as labor market (Kovak, 2013; Dix-Carneiro & Kovak, 2015, 2017, 2019; Ponczek & Ulyssea, 2021), crime rates (Dix-Carneiro, Soares, & Ulyssea, 2018), child labor (Figueiredo & Lima, 2022), infant health (Charris, Branco, & Carrillo, 2024), elections (Ogeda, Ornelas, & Soares, 2024) and religions expansion (Costa, Marcantonio, & Rocha, 2019).

Our paper enhances the existing research in three key aspects. Firstly, we utilize the 1991 Brazilian trade reform as an exogenous shock to examine the generational consequences of a trade reform. Secondly, our analysis extends to the field of gender and development by exploring the dynamics between daughter-mother pairs and son-father pairs. Our comparison is similar to

Mitra et al. (2022), however, we focus on the impact of competition from imports rather than the broadening of export avenues. Also, different from Ahsan and Chatterjee (2017), we compare IGM for males and females. Third, we investigate IGM across three dimensions —income, occupation, and education— offering a comprehensive view of social mobility in a developing context.

Beyond this introduction, this paper includes a brief overview of the trade liberalization shock in Brazil in Section 2. Section 3 describes the data. Section 4 discusses the IGM measures we use. Section 5 presents the empirical strategy. Section 6 contains the main results. Section 7 analyzes the potential mechanisms compelling the results. Section 8 comprises the robustness exercises of our results. Finally, Section 9 concludes.

2 The Brazilian trade reform

By the late 1980s, Brazil began dismantling the policy of import-substituting industrialization that had shielded its industries for nearly a century. This shift involved three significant phases of tariff reductions from 1988 to 1994. The initial phase between 1988 and 1989 saw nominal tariffs decrease from 57.5% to 32.1%, as some redundancies were removed (Abreu, 2004). Prior to 1990, not only nominal tariffs were high, but trade barriers were pervasive in Brazil. For instance, in 1987, 74% of imports were regulated by a special customs regime, and many goods required suspended import licenses (Carvalho, 1992). This made the actual level of protection vary significantly for different industries.

The efforts between 1987 and 1990 to reduce tariff redundancy and eliminate certain special regimes and trade-related taxes aimed at increasing trade policy transparency.[1] However, these changes did not significantly alter the level of protection for Brazilian industries or affect wholesale prices in Brazil, indicating that true liberalization began only in 1990 with the disposal of non-tariff barriers and special customs regimes (Kovak, 2013). This process became known as "tariffication" (*tarificação*, see Carvalho (1992)), marking a shift towards using tariffs as the main mechanism of trade policy, providing a clearer indicator of protection levels.

The second phase, from 1991 to 1993, focused on aligning tariffs more closely with the gap between international and domestic prices met by Brazilian producers. During that period, tariffs were reduced to 13.5% and most non-trade barriers were dismantled. This period saw a marked variation in tariff reductions across industries, with minor changes in the agriculture and mining sectors, while tariffs on apparel and rubber dropped by more than 30 percentage points (Dix-Carneiro & Kovak, 2019). Lastly, in 1994, average tariffs had a modest reduction to 11.2%.

[1] See Kume (1990) and Kume, Piani, and Souza (2003) for more details.

Post-1994, the momentum of trade liberalization in Brazil waned. Tariffs were increased, reversing earlier liberalization efforts, and new non-tariff barriers and safeguards were introduced.[2] While the initial push for liberalization was driven by the government with minimal influence from the private sector, the subsequent rollback after 1994 was heavily influenced by specific industry groups and the election of President Fernando Henrique, whose party had ties to import-substituting industries (Abreu, 2004). As a result, the liberalization phase from 1990 to 1995 stands as a distinct event, validating the exogeneity of tariff changes.

3 Data and summary statistics

To document the degree of intergenerational mobility, we utilize four waves of the Brazilian Demographic Census (1980, 1991, 2000, and 2010). We match the father (mother) and son (daughter) within each household to form pairs where each son (daughter) represents a unit of observation. In line with the literature (Ahsan & Chatterjee, 2017; Mitra et al., 2022), our sample is specified to cohorts of children aged between 15 and 30 from each census year.

Our sample of children considers all individuals who (i) are aged from 15 to 30 years, (ii) the parents are the household head or co-head, (iii) the family is responsible for the residence. We impose this upper age limit because of the tradeoff of the coresidence. The chance that a child's parents have either retired or deceased increases with the child's age. In such cases, it would be impossible to identify whether or not the child is in a better situation than her parents.[3] Hence, the upper age cap of 30 is chosen to reduce the chances of including child-parent pairs where the parent is either retired or no longer living.

We define the sample of parents as all individuals who (i) are aged from 30 to 64 years, (ii) are the household head or co-head, (iii) are married or live in a civil union, (iv) the family is responsible for the residence. Figure B.1 in Appendix B shows the age distribution for male and female household heads or co-heads. It is possible to observe that most part of the distribution is concentrated over individuals aged 30 or older. In Figure B.2 we plot the age distribution of children and parents pairs, noting a strikingly similar pattern between son-father pairs and daughter-mother pairs.

Children and parents are linked through the household number provided in the census. This process leads to four samples for each census round that are described in Table 1. This is our

[2]For example, the average level of tariffs for all sectors rose from 13.6% in 1994 to 17.1% in 1995 and 18.7% in 1999, with automotive industry experiencing the highest reversal along with a more limited temporary reversion for agricultural products, sugar processing, textiles and others (Abreu, 2004; Kume et al., 2003).

[3]The lower age limit is more flexible. The legal starting age to work in Brazil is 14 years. However, the child must be hired as an apprentice – which requires several requirements to be observed by the employer. Also, 14 is the limited age to finish secondary education. So, the lower limit of 15 would comprise individuals that have already finished at least secondary education, where the chances of getting a job should be higher.

Table 1. Census Sample: Number of Households and Individuals

Year	1980	1991	2000	2010
No. of Households	1,393,222	743,851	1,312,780	1,317,113
Children	2,931,656	1,414,665	2,302,655	2,018,085
Male	1,630,308	818,577	1,334,255	1,183,944
	(55.61)	(57.86)	(57.94)	(58.67)
Female	1,301,348	596,088	968,400	834,141
	(44.39)	(42.14)	(42.06)	(41.33)
Household Head/Cohead				
Father	1,297,079	691,368	1,228,993	1,240,449
Mother	1,367,137	735,817	1,304,696	1,310,049

Source: Four rounds of the Brazilian Demographic Census.

core sample, including children for whom we are capable of identifying parents. We document each child with their respective parents to form pairs, where we use the son-father pairs and daughter-mother pairs to construct our measures of intergenerational mobility. In each round of the census, about 1,300,000 households are included. The exception is the 1991 census, where 743,000 households were interviewed. However, the pattern of information does not alter. For each round, there is a greater proportion of male (57%) relative to female (42%) children. [4]

We follow the previous literature (Kovak, 2013; Dix-Carneiro & Kovak, 2015; Costa, Garred, & Pessoa, 2016; Hirata & Soares, 2016; Dix-Carneiro & Kovak, 2017) and use the microregional level of aggregation as the unit of analysis. The Brazilian Statistical Agency IBGE (*Instituto Brasileiro de Geografia e Estatística*) identifies a micro-region as a group of neighboring municipalities that are economically cohesive and share similar geographic and productive features. To adjust for the change in the boundaries, creation, and extinction of municipalities during the period of analysis, based on the approach by Reis, Pimentel, Alvarenga, and Santos (2008), we group municipalities to create 411 distinct micro-regions that are consistently recognized from 1980 to 2010.

3.1 Tariff data

The local regional tariff reduction is constructed as in Kovak (2013). Tariff data to construct the local economic shock are provided by Kume et al. (2003), which includes data at the Nível 50 industrial classification level (similar to two-digit SIC). Since the shift-share strategy requires merging data on industry-specific and individual levels to generate the regional tariff changes, wage and employment information are extracted from the Brazilian Populational Censuses. The

[4]According to the full census data, the proportion of males and females is 49% and 51%, respectively. Between sons and daughters —identified as those being a son(or daughter) to the household head— this proportion is 52% and 48%. For a sample of children 15 to 30 years old, it is 56% to 44%. Due to the need to link sons and daughters to their parents, we only consider nuclear and single families, which draw a ratio of 57 to 42.

basic idea is to combine variation across both observations (micro-regions) and some other common dimension (tariff shocks) to construct an instrumental variable (Autor et al., 2013).

4 Intergenerational Mobility Measures

IGM denotes the extent to which a child's socioeconomic characteristics depend on his parent's status. Mobility measures can be observed in relative or absolute terms. Prior literature estimates the intergenerational elasticity of income by conducting a regression of the logarithm of child income against the logarithm of parent income. However, this method's results might be misleading as there is no linearity in the relationship between log child income and log parent income, and the presence of children with negligible or no income requires a treatment that challenges the estimates (Chetty et al., 2014). Measuring absolute mobility may be more adequate given that enhancements in relative mobility can lead to ambiguous interpretations, as it could result from deteriorating conditions for individuals of higher economic status rather than improvements for those less affluent.

Since we are using census data that only contains information for parental income or education for dependent children residing within the household, there is an endogeneity concern about the choice of grown-up sons/daughters to co-reside with parents and not live independently. In section 8.1, we will alleviate this problem by finding no evidence that the Brazilian trade liberalization has significantly affected the share of working-age sons and daughters living independently from their parents.[5]

We calculate mobility measures for income, occupation and education. As relative and absolute mobility have different normative implications, we measure income and occupational mobility in absolute terms and educational mobility in relative terms. In this section, we define the statistics used to calculate mobility measures and discuss their conceptual properties.

4.1 Absolute Income Mobility

We follow the recent literature and explore the linear relationship between children and parents' income ranks (Chetty et al., 2014; Dahl & DeLeire, 2008). The rank-rank relationship compares children's income positions relative to their peers against their parents' income standings relative to other parents, providing a direct comparison of socioeconomic statuses across

[5]Hilger (2017) uses a similar assumption in cross-sectional decennial census data to estimate educational mobility in the USA. The necessary assumptions are as follows: first, a "parallel trends" assuming that an additional year of parental education predicts equally higher educational attainment of dependent and independent young adult children, and a "smooth cohorts", requiring that the population shares of different SES groups do not change fertility in differential ways.

generations. The linear relationship in the rank-rank comparison allows us to summarize intergenerational income mobility from statistical figures that are easily comparable between different locations.

We characterize variation in intergenerational mobility across microregions by associating children with the microregion where they reside. We determine the linear parameters of the rank-rank regression for each microregion c by conducting separate regressions of children's ranks (at the national level) from the same age cohort against their parents' ranks for each region:

$$R_{ic} = \alpha_c + \beta_c P_{ic} + \varepsilon_{ic} \tag{1}$$

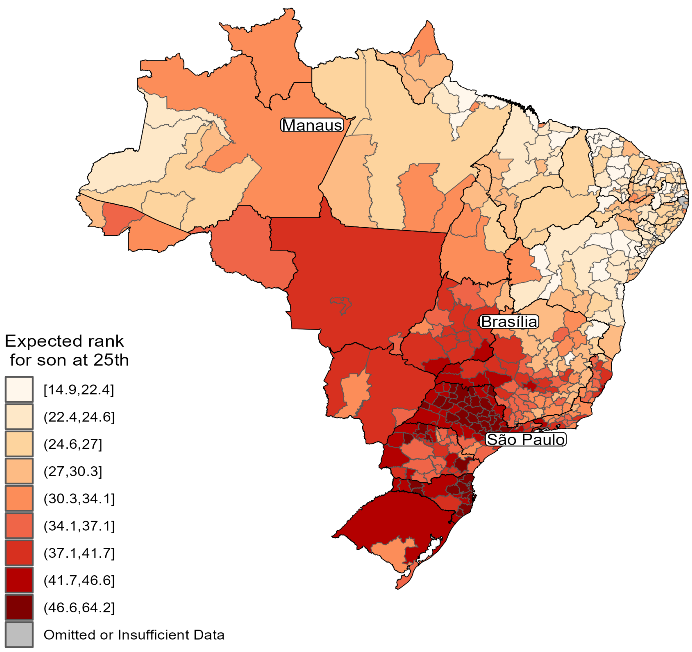
Here, the outcome variable R_i denotes the percentile rank of the child i in the national income distribution for children from the same age cohort, and the independent variable P_i is the percentile rank of her parent. This regression results in two IGM measures. β_c quantifies the relative mobility within microregion c, showing the mobility gap for children from parents with a one-percentile difference in income. Thus, the slope β_c indicates the impact of parental income on the economic outcomes of their children within a specific area. α_c predicts the average rank of children from the lowest parental income bracket. By using rank-rank regression, we account for zero incomes among children and capture spatial differences in mobility trends, aligning children's ranks from various regions against a fixed national scale of parental income. We prefer to use the national benchmark over local metrics for parental income to ensure consistent cross-regional comparisons, unaffected by local income distribution disparities.

Using the regression estimates, we obtain the expected rank of children within any percentile p of the national parental income scale. We define absolute income mobility as $r_{25} = \alpha_c + 25 \times \beta_c$, representing the expected rank for a child based on their parents being in the 25^{th} income percentile. Hence, a higher r_{25} means greater income mobility. We choose to focus on the 25^{th} percentile because outcomes of disadvantaged youth are a central focus of policy interest and the observation that there is a broader range of outcomes among children from lower-income backgrounds compared to those from wealthier families across different locations.

Figure 1 depicts spatial variation in absolute mobility at $p = 25$ for father-son pairs across microregions in 2010. The map highlights some important aspects. First, there is substantial heterogeneity in absolute mobility across microregions, where the expected rank of below-median income sons varies from the 14^{th} to the 64^{th} percentiles. We can notice that the Center-South region presents significantly higher upward mobility than the North and Northeast regions. This result is not surprising, as these regions present a higher incidence of poverty, inequality and unemployment. In Appendix Table A.1, we report these socioeconomic indicators for the states of the country.

The second aspect is the concentration of the top 10% areas in respect of income mobility in three states: Santa Catarina, Paraná and São Paulo. On the other hand, the bottom 10% areas are mainly concentrated in states from the Northeast region: Alagoas, Bahia, Ceará, Maranhão, Paraíba, Pernambuco, Piauí, Rio Grande do Norte, and Sergipe. Also, Appendix Figure B.4 shows that areas with higher absolute mobility might not present higher relative mobility. In the Center-South regions, places with high absolute mobility experience higher relative mobility. However, in the Northern regions, this correlation is too small.

Figure 1. Absolute Income Mobility: Expected Rank for Sons at the 25^{th} Percentile



Notes: Expected rank across regions for sons with father income at the 25^{th} percentile.

Overall, mobility in Brazil has slowly increased through the years. Figure B.3 exhibits the connection between parent and child income ranks for the 1991 and 2010 census samples. For each parental income percentile, resulted from the average of the father and mother incomes, it displays the mean rank for children aged 15-30. We can observe that the mobility curve in Brazil has become slightly more horizontal through the years. This reflects that parents' socioeconomic status has diminished the influence in predicting children's economic outcomes. While the ran-rank slope decreased from 0.64 in 1991 to 0.53 in 2010, the expected rank for children with parents at the 25th percentile slightly increased (from 36th in 1991 to the 37th in 2010). However, mobility in Brazil shows to remain much lower than in other countries with similar measures. For example, the rank-rank slope in the US is calculated at 0.34 (Chetty et al., 2014), 0.22 in Italy (Acciari et al., 2022), 0.245 in Germany, 0.223 in Norway, and 0.215 in

Sweden (Bratberg et al., 2017).

Furthermore, mobility curves are also distinct across genders, as depicted in Figure B.6. Male children present both higher absolute and relative mobility relative to female children. The expected rank for male children at the 25th in the parental income distribution is almost 6 percentile points higher than the expected rank for female. This highlights the importance of our approach in differentiating the mobility measures across genders.

4.2 Occupational Mobility

Our measure of occupational mobility follows Ahsan and Chatterjee (2017) and Mitra et al. (2022). We distinguish the jobs depending on their skill intensity. A job is better compared to another if it requires more skill than the other. Thus, our measure of skill intensity by occupation depends on the worker's average educational attainment within each occupation.

We proceed as follows. First, we construct a ranking of occupations according to the average educational attainment of workers in each occupation o, EI_o:

$$EI_o = \frac{\sum_{i \in o} Educ_i}{L_o} \tag{2}$$

where L_o is the total of individuals employed in occupation o and $Educ_i$ is the years of study completed by the individual i. This index is calculated using data from all workers aged 15 to 64, based on information prior to the 1991 reform to guarantee that our occupational rankings are not influenced by the trade liberalization in Brazil. For individuals in the survey years of 1980, 2000, and 2010 surveys, we align each individual's job with the skill intensity of that occupation as determined in 1991. Consequently, individuals can alter the skill intensity associated with their occupation by changing jobs, though the skill intensity assigned to each occupation cannot change over time. An education index and the occupation rankings are detailed in Table A.2 in Appendix 9.

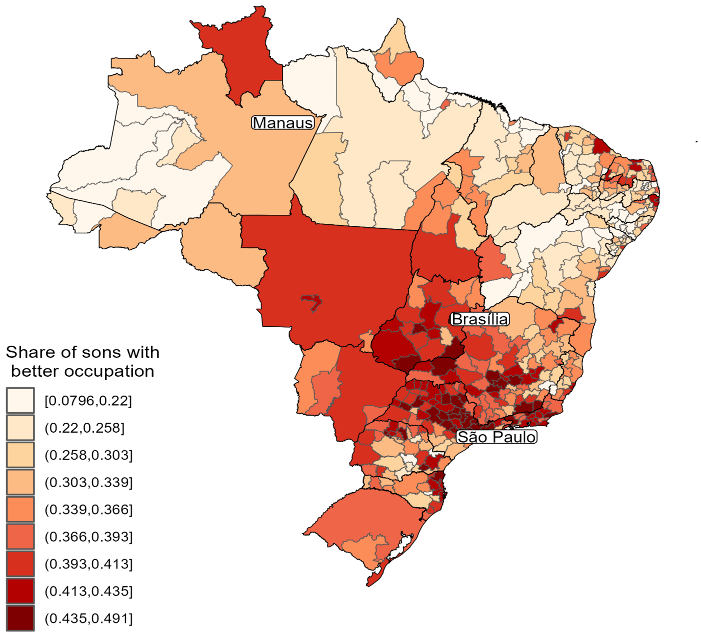
We define upward occupational mobility as when a child works in a higher-ranked occupation than that of their parents. Utilizing the occupational rankings, we create an indicator variable of upward occupational mobility, $Upward_i$, where this indicator equals one if the son(daughter) i works in a job ranked higher than that of their parents, and equals zero otherwise.

Then, our measure of upward occupational mobility in region c is the share of children that have better occupations than their parents. We calculate the share of sons(daughters) with higher-ranked occupations as follows:

$$\text{UOM}_c = \frac{1}{N_c} \sum_i \mathbb{1}\{Upward_i\}$$

where N_c is the total number of sons or daughters in the microregion c that co-reside with their parents. UOM_c represents the share of sons (daughters) working in a higher-skilled occupation than their fathers(mothers) in microregion c. Here, we considered only employed individuals aged between 15 and 30. We used each survey to measure upward occupational mobility through the years for a microregion.

Figure 2. Upward Occupational Mobility across regions



Notes: Share of sons employed in higher ranked occupations compared to his father across regions.

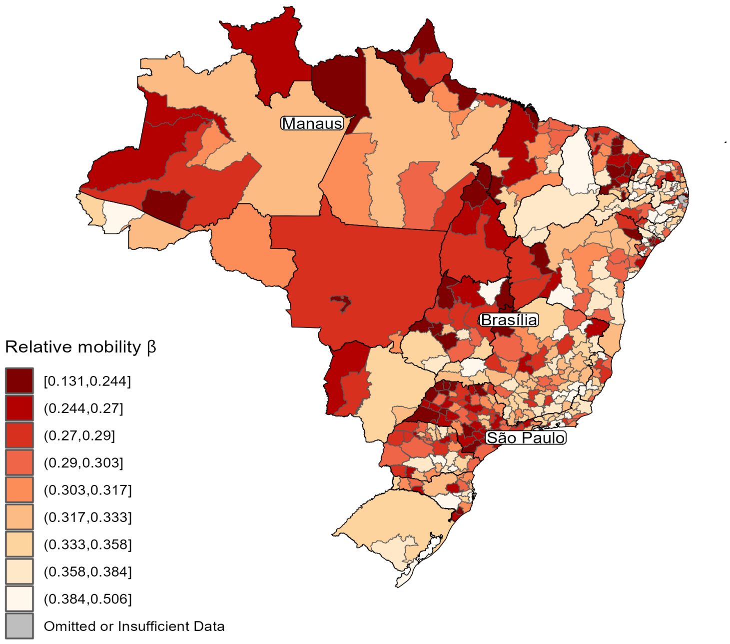
Figure 2 shows the spatial variation in upward occupational mobility for son-father pairs across regions in 2010. The pattern is remarkably similar to the one observed in Figure 1 for absolute income mobility. The share of sons with better occupations than their fathers varies substantially across regions, where this share ranges between 7% and 50%. Center-South regions concentrate most regions with higher mobility, as Northern regions present lower mobility.

4.3 Educational Mobility

We run the same specification as in equation (1) in parent and child education ranks rather than levels. However, instead of calculating the expected absolute mobility (intercept of the regression), we use the β_c of the rank-rank regression(the slope coefficient) of the child's rank on the parents' rank to measure educational relative mobility. One should notice that relative mobility tells us the degree to which parents determine children's education; hence, what we see is how this dependency changes after the trade shock. Importantly, the rank-rank measure of mobility has the advantage of being invariant to changes in educational inequality across generations Hilger (2017).

We could also use the absolute mobility measure (α_c), as in the case of income mobility.

Figure 3. Relative Educational Mobility across regions



Notes: Regional relative mobility β estimated as in equation 1, measuring the inverse relative mobility between children whose parents are 1 percentile of difference in the parental income distribution.

However, the relative mobility in the educational aspect is more elucidative. Children can aim to achieve the same socioeconomic status as their parents. Still, the income level depends on other circumstances rather than personal components, where the parents can highly influence education.

Figure 3 presents the relative educational mobility across regions for son-father pairs in 2010. As for income and occupational mobility, the map shows a substantial heterogeneity across regions. However, different from the absolute mobility measures, educational relative mobility is especially higher in the Center and in the state of São Paulo. Also, there are more areas in the Northeast region presenting high educational mobility, different from income and occupational mobility.

5 Empirical strategy

5.1 Regional Tariff Cut

The definition of regional economic shocks is derived from the empirical literature on trade shock exposure consequences on local labor markets. This approach was initially introduced by Topalova (2010) and then formalized by Kovak (2013) within the context of a specific-factors model for regional economies. Intuitively, regions within a country often focus on the production of different goods. Thus, trade shocks affect industries to varying degrees making them differently susceptible to trade shocks. Consequently, regions with key industries that face more significant tariff reductions tend to experience more pronounced adverse effects on their labor market dynamics.

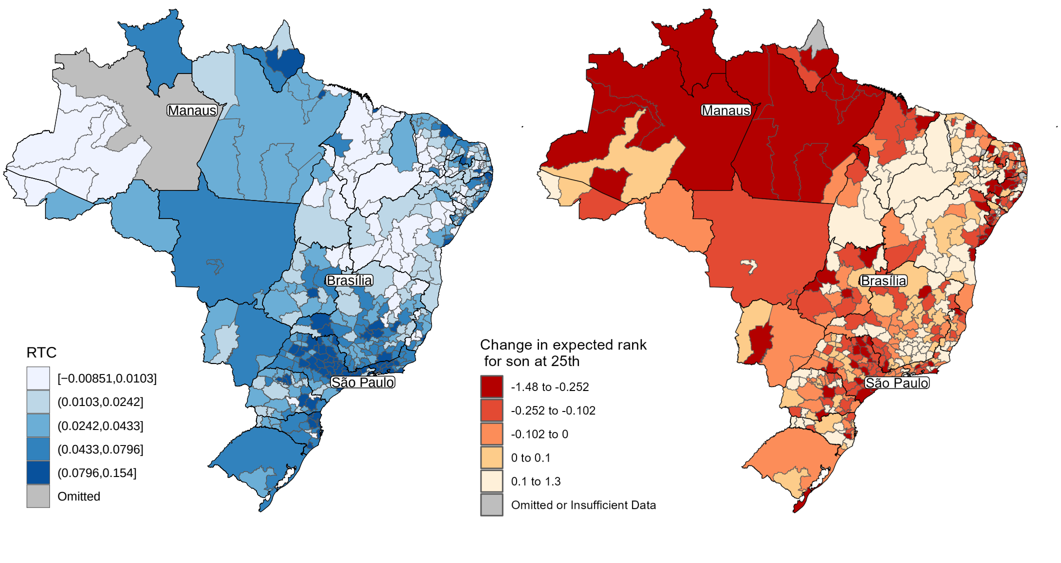
Following Kovak (2013), the local economic shock is defined as the "Regional Tariff Cut" (RTC) in region c, which measures the regional industries' exposition to tariff reductions. Thus, the RTC_c is constructed as the weighted mean of the tariff reductions encountered by the industries, where the weighting reflects the significance of each industry within the region's economy. This relationship is formally expressed as:

$$RTC_c = -\sum_{i \in T} \beta_{ci} \Delta \log\left(1 + \tau_i\right), \quad \text{where} \quad \beta_{ci} = \frac{\lambda_{ci}}{\sum_{j \in T} \lambda_{cj}} \tag{3}$$

such that τ_i is the tariff on sector i, Δ symbolizes the time-difference from 1990 to 1995, $\lambda_{ci} = \frac{L_{ci}}{L_c}$ is the initial share of region c workers employed in the industry i. T denotes the set of all tradable sectors. Thus, although all regions face the same aggregate tariff modifications, distinctions in the regional industry combination yield regional variation in shocks to the labor market.

Figure 4 shows the kind of variation we examine. Regions with darker colors reflect more exposition to the trade shock and greater reductions in absolute income mobility. The figures show that the reductions in the expected rank for sons with fathers in the 25th percentile were more concentrated in the Center-South and North regions, coinciding with the places facing more competition from imports following the tariff cuts.

Figure 4. Regional Tariff Cuts and changes in absolute income mobility



(a) Regional Tariff Cuts (b) Changes in Absolute Income Mobility for son-father pairs (1991-2010)

Notes: Figure (a) depicts the weighted average of tariff changes. Figure (b) represents the change in the expected rank for children given parent income at the 25^{th} percentile. See text for details.

5.2 Local Trade Shocks and changes in intergenerational mobility

The following specification compares the evolution of IGM in regions facing larger tariff reductions to those in regions facing smaller tariff declines. We estimate separate regressions for each time difference between 1991-2000 and 1991-2010:

$$\log\left(IGM_{c,t}\right) - \log\left(IGM_{c,1991}\right) = \theta_t RTC_c + \alpha_{m,t} + X_c + \epsilon_{c,t} \tag{4}$$

where $IGM_{c,t}$ is the mobility measure in region c at time t and $\alpha_{m,t}$ are meso-region time-fixed effects. RTC_c does not vary with the year t under consideration, so each year's θ_t describes the cumulative local effects of liberalization as of each post-liberalization year. We include meso-region fixed effects to control for any time-invariant regional characteristics that could be correlated with tariffs and mobility levels. Also, we cluster the standard errors at the meso-region level to account for potential spatial correlation in outcomes across neighboring regions. Since we are using the equation 1 estimates of the rank-rank regression for mobility measures as dependent variables (the first-stage), we weight the second-stage equation (4) regressions with the inverse of the standard errors of the equation (1) (e.g., (Kovak, 2013; Dix-Carneiro & Kovak, 2017; Dix-Carneiro et al., 2018)).

Importantly, we also consider a vector of controls X_c, controlling for the share of each microregion's urban population, white population, and government spending per capita in 1991. Thus, from Equation (4), our results rely on comparisons before and after the trade shock between regions with similar socioeconomic structures but specialized in different industry sectors.

The identification condition is that our main parameter of interest θ_t is identified if $\epsilon_{r,t}$ is uncorrelated with RTC_r, conditional on the meso-region fixed effects. That is, the estimated causal effect of the trade shock is unbiased if there are no omitted variables correlated with RTC_r that affected mobility levels and were not captured by the controls. One way to assess the credibility of this exposure design is to test for parallel pre-trends. We run pre-intervention placebo tests using a pre-period year before the trade reform (1980) to check if there were previous differences in the outcomes changes that could invalidate our identifying assumption. This empirical strategy is equivalent to difference-in-differences, where we take advantage of level differences in the shares to assess if differential exposure to common socks caused different changes in the outcome's changes.

6 Results

The consequence of regional tariff cuts on IGM is noticeable. From 1991 to 2010, regions exposed to larger tariff reductions underwent a persistent relative *decrease* in intergenerational occupational mobility and a long-run decrease in absolute income and relative educational mobility compared to less strongly affected regions. In Table 2, we exploit the trade shock effects on absolute income mobility changes. In Column 1, we regress pre-liberalization mobility measures on trade shocks as pre-intervention placebo tests and find no statistically significant effect of the trade reform on pre-liberalization changes, suggesting that preexisting trends are not a threat to our analysis. For son-father pairs, there is a lower expected rank for children whose parents are in the 25^{th} percentile of the income distribution, r_{25}. Pushing a region from the 10 to the 90 percentiles in the distribution of the regional tariff cut (a decrease in tariffs of 0.1 log points) reduces by approximately 7% absolute income mobility for son-father pairs in the long run.

Table 2. Regional tariff cuts effects on absolute upward income mobility

	(1)	(2)	(3)	(4)	(5)	(6)
		Son-Father			Daugther-Mother	
	1980-1991	1991-2000	1991-2010	1980-1991	1991-2000	1991-2010
RTC_c	-0.366	-0.280	-0.696**	-0.321	-1.120*	-0.920
	(0.328)	(0.352)	(0.347)	(1.071)	(0.595)	(0.610)
Observations	408	408	408	399	400	400
R-squared	0.521	0.566	0.591	0.343	0.511	0.480

Notes: Regression of changes in the expected rank for children from the 25^{th} percentile of parental income distribution on regional tariff cuts. All regressions control for meso-region r time-fixed effects and 1991 government spending per capita, 1991 share of urban population and share of white population. Observations are weighted by the inverse of the squared standard error of the estimated change in log microregion absolute upward income mobility, calculated using the procedure in haisken1997interindustry. Standard errors are clustered at the mesoregion level. Level of significance: *** at 1%, ** at 5%, * at 10%.

On the other hand, the exposure to more import competition reduced the absolute upward mobility of daughter-mother pairs in the medium run. Reducing 0.1 log points in average tariff would decrease the expected rank for daughters in the 25^{th} percentile in 11%. In the long run, this negative effect slightly reduces to 9%, but still being a more significant reduction compared to males. This weakening of the reduction in the long run could reflect the increase of women entering the labor market and better job positions. Indeed, we show in Section 7 that the share of high-skill positions increased compared to low and medium-skill positions, together with a reduction in the medium-to-low earnings gap.

A potential concern one could have with our result is how the variation in living costs across different regions impacts our regional mobility estimates. In Appendix Figure B.8, we demonstrate that adjusting for local price levels does not significantly change the overall trends observed in our map. Regions in the Center-South part of Brazil still present a higher expected rank for IGM. To adjust for real income, we create a regional price index using the 2009 edition of the POF (*Pesquisa de Orçamentos Familiares*), a household budget survey conducted by IBGE, which collects data on consumption and expenditure of products in a detailed geographic level. We distinguish the capital cities and countryside for each state and use this data to calculate the average cost of a standard basket of goods, which is also used to compute the main Brazilian consumer inflation index (IPCA). With the calculated prices, we adjust the incomes of parents and children and re-estimate the rank-rank regression from Equation (1) to obtain the price-adjusted regional mobility measures.

In Appendix Table B.8 column (1), we run the effect of the tariff cut without controlling for sociodemographic characteristics to test the robustness of our baseline results. From columns (4) to (6), we replicate the exercise using the price-adjusted regional mobility measures. In both cases, there is no signal of pretrends in absolute upward mobility changes in areas more exposed to trade liberalization, and the post-liberalization effects remain negative and statistically significant. In terms of magnitude, there is a higher reduction in mobility for both genders, suggesting that the results could be even higher.

Upward occupational mobility was also reduced in areas more exposed to import competition following the tariff cuts, as shown in Table 3. When considering a shift in 0.1 log points in the tariff exposure level, the share of sons with higher-ranked occupations than their fathers reduced by 5.6% in the medium run and 10% in the long run, representing a persistent relative decrease in the high-skilled job opportunities for sons compared to their fathers. For daughter-mother pairs, there is a signal of a persistent but stable reduction induced by trade liberalization. It is expected that the share of daughters with better occupations than their mothers reduced by 7.5% almost twenty years subsequent the trade liberalization episode.

Corroborating the results of the tariff reductions on income and occupational mobility, Table 4 denotes that the relative educational mobility decreased after the trade reform. Here, we compare the changes in β_c between periods, which measures relative educational mobility, whereby a greater βc imply lower mobility levels. The result shows a positive and statistically significant effect of tariff reductions on educational mobility in the long run. It is expected that the exposure to trade competition implied a decrease of 15% in relative mobility for son-father pairs in the long run. In contrast, daughter-mother pairs' relative mobility was reduced by 20%.

Table 3. Regional tariff cuts effects on upward occupational mobility

	(1)	(2)	(3)	(4)	(5)	(6)
		Son-Father			Daugther-Mother	
	1980-1991	1991-2000	1991-2010	1980-1991	1991-2000	1991-2010
RTC_c	0.212	-0.559***	-0.988***	-0.0522	-0.747***	-0.751***
	(0.141)	(0.126)	(0.165)	(0.265)	(0.204)	(0.263)
Observations	407	407	407	407	407	407
R-squared	0.426	0.731	0.838	0.267	0.432	0.393

Notes: Regression of changes in the share of children with better occupations than their parents on regional tariff cuts. All regressions control for meso-region r time-fixed effects and 1991 government spending per capita, 1991 share of urban population and share of white population. The unit of analysis is a micro-region. Observations are weighted by the total population aged from 15 to 30 years. Standard errors are clustered at the mesoregion level.
Level of significance: *** at 1%, ** at 5%, * at 10%.

Table 4. Regional tariff cuts effects on relative educational mobility

	(1)	(2)	(3)	(4)	(5)	(6)
		Son-Father			Daugther-Mother	
	1980-1991	1991-2000	1991-2010	1980-1991	1991-2000	1991-2010
RTC_c	-0.553	0.480	1.539*	-0.646	-0.185	1.961**
	(0.601)	(0.769)	(0.867)	(0.708)	(0.634)	(0.906)
Observations	405	405	405	407	407	408
R-squared	0.415	0.534	0.640	0.445	0.565	0.657

Notes: Regression of changes in the relative educational mobility (β_c from the rank-rank regression of child rank on parents rank in their respective educational distribution) on regional tariff cuts. All regressions control for meso-region r time-fixed effects and 1991 government spending per capita, 1991 share of urban population and share of white population. The unit of analysis is a micro-region. Weighted by the inverse of the squared standard error of the estimated change in log microregion relative educational mobility, calculated using the procedure in haisken1997interindustry. Standard errors are clustered at the mesoregion level.
Level of significance: *** at 1%, ** at 5%, * at 10%.

7 Potential Mechanisms

This section explores the relationship between labor market conditions and IGM. It has been already documented by the empirical literature that regions facing more competition from imports tend to experience deterioration in wages and earnings (Autor et al., 2013; Kovak, 2013; Dix-Carneiro & Kovak, 2017). The results found in the last section suggest that the participation of young workers in higher-ranked occupations diminished following the trade shock. This could result from changes in skill demand from higher-skilled to lower-skilled occupations. Also, this could indicate a change in the earnings gap across the skilled level of occupations.

In Figure B.7 we show the relationship between Tariff Changes and Preliberalization Sectors' Average Years of Education. The correlation is almost zero between the tariff reduction sector skill levels. This shows the trade liberalization in Brazil has impacted more the medium-skilled sectors than high skilled or low-skilled sectors. This is contrary to what (Ahsan & Chatterjee, 2017) finds, where they predict that import competition induces firms at or close to the efficiency frontier to innovate, resulting in an increase in their demand for skilled labor.

First, we investigate the skill demand as a mechanism driving the diminishing impact of trade reform on intergenerational mobility. In Table 5 we assess if there were changes in skill demand within regions, to evaluate if mobility levels diminished due to the reallocation of production activities across industries towards the sectors that demand less skilled labor.[6] Our results suggest that the medium-skilled share of employment persistently reduced in areas more exposed to import competition. This reduction is accompanied by an increase in the share of high and low-skilled employment in the medium run, but only the low-skilled share increase persisted through time.

Table 5. Effects of tariff cuts on skill demand within regions

	(1)	(2)	(3)	(4)	(5)	(6)
	High-skilled Share		Medium-skilled Share		Low-skilled Share	
	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
RTC_c	0.934***	-0.192	-2.299***	-3.969***	0.829***	0.698***
	(0.189)	(0.260)	(0.192)	(0.273)	(0.095)	(0.120)
Observations	411	411	411	411	411	411
R-squared	0.699	0.670	0.799	0.875	0.753	0.649

Notes: Regression of changes in the employment shares of each skill level (high, medium, and low skill occupations) from 1991 to 2000 and 2010 on regional tariff cuts. Unit of analysis is a micro-region. All regressions control for meso-region r time-fixed effects. Observations are weighted by the total population. Standard errors are clustered at the mesoregion level.
Level of significance: *** at 1%, ** at 5%, * at 10%.

In Table 6, we assess if there were changes in skill demand inequality within regions. In general, the demand for high-skilled compared to low-skilled workers diminishes due to the tariff cuts. A trade shock moving a region from the 10 to the 90 percentiles in the distribution of the regional tariff cut (a reduction in tariffs of 0.1 log points) is estimated to reduce 9% the share of high-skill occupations relative to low-skilled occupations. As expected, we also find that the share of medium-skill occupations relative to high- and low-skilled occupations decreased after the trade shock.

We also investigate the skill demand across gender to asses if the negative consequences in the

[6]We classify the skill level according to Table A.2 in the appendix. High-skill occupations codes: 94, 95, 98, 91, 10. Medium skilled occupations codes: 15, 8, 17, 93, 12, 5, 32, 96. Low skill: 97, 25, 22, 4, 2, 14, 92, 1.

labor market are heterogeneous across genders. Table A.3 in Appendix A shows the changes in skill demand inequality across genders. We find that the tariff cuts diminished the share of male workers in high and medium occupations relative to low-skill occupations. In contrast, we find the opposite for female workers, where the share of low-skilled occupations decreased compared to high and medium-skill occupations.

Overall, these results indicate that the lower levels of mobility caused by the tariff reductions could result from higher-ranked occupations diminishing following the trade shock. However, it could be that the earnings from these occupations remained stable or increased and thus alleviated the problem of the reallocation of production activities across industries towards less skilled sectors.

To compare observationally similar individuals and net out compositional effects, such that changes in these variables have resulted from changes in the local labor market, we estimate the following Mincer regression to obtain region- and year-specific log earnings:

$$\begin{aligned} \log\left(w_{ict}\right) = & \omega_{ct} + \sum_{k} \varphi_{kt}^{w} I\left(Educ_i = y\right) + \gamma_t^w \text{ Female } + \\ & \delta_{1t}^{w}\left(age_{it} - 18\right) + \delta_{2t}^{w}\left(age_{it} - 18\right)^2 + \varepsilon_{ict}^{w} \end{aligned} \tag{5}$$

Here, the outcomes variables $Wage_{irt}$ is the total monthly earnings for an individual i in region r in year t. The term ω_{rt} is the region and time-adjusted average log earnings. This is our parameters of interest from the regression above. We save $\widehat{\omega_{rt}}$ estimates and use as measures of average log-earnings in region r and year t. To adjust for compositional effects, we include: an indicator variable for each level of education, $Educ$; a categorical variable for gender, $Female$; and age and experience measures, age and age^2, respectively.

Table 6. Effects of tariff cuts on skill-demand inequality

	(1)	(2)	(3)	(4)	(5)	(6)
	High/Low		High/Medium		Medium/Low	
	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
RTC_c	0.105	-0.890**	3.233***	3.777***	-3.129***	-4.667***
	(0.249)	(0.361)	(0.309)	(0.297)	(0.231)	(0.354)
Observations	411	411	411	411	411	411
R-squared	0.584	0.626	0.843	0.864	0.817	0.861

Notes: Regression of changes in the difference of the share of employment accounted for each skill level (high, medium and low skill occupations) on regional tariff cuts. The unit of analysis is a micro-region. All regressions control for meso-region r time-fixed effects. Observations are weighted by the total population. Standard errors are clustered at the mesoregion level.
Level of significance: *** at 1%, ** at 5%, * at 10%.

In Table 7, we run equations similar to (4) with the earnings estimates $\widehat{w_{rt}}$ from the exercise described above, on the left-hand side as outcome variables. We analyze the effects of tariff reductions on earnings across skill groups. While the high-skilled occupations were unaffected by the increased import competition, medium- and low-skilled workers suffered persistent and relative decreases in their earnings. We find that a reduction in tariffs of 0.1 log points leads to a decrease of 13% average earnings from medium-skilled occupations and a decrease of 9% in low-skill occupations.

We also assess if there were changes in earnings differentials across skill groups. Table 8 shows relatively small effects of tariff cuts on local average earnings log differentials. The earning gap between high occupations and low occupations remained unaltered. However, the workers' earnings from medium occupations diminished relative to the other sectors, accompanying the fact that the sector suffered the most with the tariff declines. In the long run, the gap between high and medium occupations increased, and the gap between medium and low decreased.

Finally, we find small effects of the trade liberalization on earnings differentials across skill groups by gender. The results show that male workers from the high-skill sector benefited with a slight increase relative to the medium-skilled sector, which diminished their earnings gap compared to workers from low-skilled occupations. For female workers, there was an increase in the earnings from low-skilled occupations compared to medium-skilled occupations.

Table 7. Effects of tariff cuts on earnings according to occupation skill groups

	(1)	(2)	(3)	(4)	(5)	(6)
	High-skilled		Medium-skilled		Low-skilled	
	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
RTC_c	-0.415	-0.398	-0.717***	-1.353***	-0.487**	-0.914***
	(0.253)	(0.257)	(0.207)	(0.212)	(0.245)	(0.250)
Observations	411	411	411	411	411	411
R-squared	0.490	0.511	0.692	0.848	0.742	0.831

Notes: Regression of changes in the average earnings for workers according to each skill level (high, medium and low skill occupations) on regional tariff cuts. Unit of analysis is a micro-region. All regressions control for meso-region r time-fixed effects. Observations are weighted by the total population. Standard errors are clustered at the mesoregion level.
Level of significance: *** at 1%, ** at 5%, * at 10%.

8 Robustness Checks

8.1 Selection Bias

In Section 3, we described the data used to document intergenerational mobility. Using household survey data, we measure mobility for father–son and daughter-mother pairs living in the same household. This poses a selection bias to our selected sample that could result in inaccurate mobility measures. We try to remedy this potential threat to our identification by performing two robustness exercises.

Table 8. Effects of tariff cuts on earnings differentials across skill groups

	(1)	(2)	(3)	(4)	(5)	(6)
	High/Low		High/Medium		Medium/Low	
	2000	2010	2000	2010	2000	2010
RTC_c	0.0170	0.0863	0.106	0.199**	-0.00105	-0.116**
	(0.0763)	(0.0829)	(0.103)	(0.0792)	(0.0725)	(0.0532)
Observations	411	411	410	411	411	411
R-squared	0.237	0.162	0.355	0.246	0.403	0.316

Notes: Regression of changes in the difference of workers' average earnings for each skill level (high, medium, and low skill occupations) on regional tariff cuts. The unit of analysis is a micro-region. All regressions control for meso-region r time-fixed effects. Observations are weighted by the total population. Standard errors are clustered at the mesoregion level.
Level of significance: *** at 1%, ** at 5%, * at 10%.

First, a child's decision to not co-reside with his parents and therefore live independently, establishing their own household, might correlate with a microregion's exposure to trade liberalization, or regions more exposed to tariff cuts have a greater/lower fraction of co-residency. Such endogeneity problems could lead to biased estimates of the impact of trade liberalization on IGM. We explore this point by examining whether trade shock has had any influence on the share of adult children who are household heads or co-heads.

Table A.5 in the Appendix A details the influence of tariff reductions on the share of sons and daughters living independently. We identify households that are headed or co-headed by independent sons(daughters) as follows. To identify households led or co-led by such independent individuals, we initially limit our analysis to those aged 15 to 30. We then sift through the census data to find households that are headed or co-headed by young adults and check if these individuals are employed.

The results show no effect of the trade liberalization on the share of independent sons. Nonetheless, we notice that the coefficients' sign is negative and statistically significant for daughters. However, that effect is accompanied by a pretend, indicating that the share of independent daughters was already diminishing before the trade shock in the areas more exposed. This indicates that the share of young adult females living with their parents is increasing, so we capture this pattern using census data.[7]

Our second strategy to deal with the source of selection bias involves the application of propensity score weighting (PSW), recommended by Francesconi and Nicoletti (2006) and Ahsan and Chatterjee (2017), to estimate the effect of the trade liberalization on mobility. We procedure

[7] Overall, This result is probably due to labor market changes for women, increasing their participation in the labor market and increasing the average marriage age. As we are considering young adults, this could indicate that women are leaving home later (*to do*)

as follows.

Initially, we estimate the following selection model via probit regression:

$$s_{ic}^{*} = \Theta Z_{ic} + u_{ic} \tag{6}$$

where s_{ic}^{*} represents a latent variable associated to a binary indicator s_{ic}, which equals one if child i from region c lives with their parents, and zero if not; hence, whether the child is in our sample. Z_{ic} encompasses individual characteristics that are determinants of the likelihood of a child residing with their parents, namely, the child age, if she lives in an urban area, ethnicity and microregion fixed effects. Last, u_{ic} is a random error term.

The PSW procedure operates under the assumption of selection on observables. Thus, it assumes that the set Z_{ic} correctly predicts the likelihood of a child living with their parents. The predicted values resulting from the estimation of Equation (6) are the propensity scores, they reflect the likelihood of a child residing with their parents based on observable characteristics, with higher scores indicating a greater probability.

In the second step, we use the inverse of the propensity scores from the first step as weights to assess the impact of trade liberalization on intergenerational mobility. We run the difference-in-differences (DID) specification below:

$$\mathrm{y}_{ict} = \theta \times \mathrm{PostRTC}_t + X'_{ict}\beta + \alpha_c + \alpha_t + \varepsilon_{ict} \tag{7}$$

where y_{ict} is an indicator for a child i in microregion c and year t employed in a higher ranked position. $\mathrm{PostRTC}_t$ is an indicator variable for post-liberalization years. X_{ict} is a vector of individual controls such as age, age squared, ethnicity, lives in an urban area, father/mother's age and education. Finally, α_c and α_t are microregion and year fixed effects, respectively. Additionally, we cluster standard errors at the microregion-by-year level, which is the trade shock variation level.

The PSW procedure implies that the weighting in Equation (7) places a higher weight on children with a low probability of co-residing with their parents. Thus, the weighting simulates a representative sample that contains all children. Together with our identification hypothesis in this setting -which is obtained by comparing across microregions the changes in the probability of children who presented absolute upward mobility before and after the trade shock with those exposed differentially to the shock due to differences in the initial industrial composition of regions- this exercise reinforces confidence in the unbiased results estimated in section 6.

Table A.6 presents the results from using this method, separating the son-father sample in Panel A and the daughter-mother sample in Panel B. Column (1) shows the baseline results. Column (2) considers individual controls. In column (3), we introduce pre-shock microregion characteristics, including the white and urban population shares, illiteracy rate, government spending per capita and income per capita. This specification controls for differences in initial conditions that could influence intergenerational mobility trends over time.

The PSW procedure in column (1) Panel A with upward occupational mobility as the outcome variable denotes a negative and statistically significant coefficient for trade liberalization close to our baseline result in magnitude. We find that a 0.1 log point decline in tariff levels is expected to decrease 6,5% the probability of a son working in a higher ranked occupation compared to his father relative to sons from microregions less exposed to the trade liberalization. In columns (2) and (3), the coefficient size decreases with the introduction of individual controls but remains statistically significant. Looking to Panel B, we observe a stable, negative and significant effect of tariff reductions on the probability of daughters succeeding their mothers.

We proceed with the analysis looking to absolute income mobility for children whose parents are below the median income. The outcome variable is an indicator for the children's percentile among the children income distribution being higher than her father's rank in the parental income distribution. The results in columns (4)-(6) show a reduction of approximately 5% in the probability of sons succeeding their fathers relative to less exposed regions. For daughter-mother pairs, exhibited in Panel B, there is also a reduction (2%) in this probability. If anything, this result differs from our baseline result, where there is no evidence between trade liberalization and changes in the expected rank of daughters at the 25^{th} percentile.

Finally, from columns (7)-(9), we look for the educational attainment of children and parents. Here, the outcome variable is an indicator of whether a son(daughter) has higher years of schooling compared to his(her) father(mother). Column (9) denotes a significant reduction of 16% in the probability of sons succeeding their fathers in years of schooling, thus very close to our baseline result in the relative mobility using rank-rank estimates of intergenerational educational mobility. For daughter-mother pairs, there is a reduction of around 5% in the likelihood of upward mobility and below the baseline estimate.

Together, the results described in Table A.5 and Table A.6 support our identification hypothesis alleviating concerns over the endogeneity induced by the co-residency in our data. We showed that exposure to liberalization did not affect the share of independent children. If anything, the potential selection bias from co-residency unlikely affected our baseline estimates in Section 6 as our estimates for upward mobility using the PSW procedure are very close to the baseline results.

Another factor that might introduce selection bias is the exclusion of individuals over the age of 30 from the analysis. This decision aimed to reduce the likelihood of including children whose parents are retired, thereby lacking current occupational information. In Table A.7 of our robustness checks, we explore the impact of trade liberalization using alternative age thresholds—specifically, 25, 35, and 40 years—to determine if our initial age limit affects the outcomes. In Panel A, we assess father-son pairs, while Panel B focuses on mother-daughter pairs. We evaluate the changes in the share of children with higher-ranked occupations from 1991 to 2000 from columns (1) to (3) and from 1991 to 2010 from columns (4) to (6). In all cases, the coefficient of the tariff cuts is negative and statistically significant, aligning closely with our primary findings in magnitude.

8.2 Additional robustness checks

An example of a potential concern with our identification strategy is the selective migration of individuals. If trade liberalization is inducing migration across regions as workers' response to the negative economic shock in their local labor market, then our estimated impact of tariff changes in IGM is not entirely driven by the impact of tariff cuts on the local labor market, but is biased by the migration of individuals from regions more affected to microregions less impacted. To verify this confound in our results, we re-estimate our baseline exercise for upward occupational IGM in columns (1) and (2) of Table A.8, using a sample excluding children that were not born in the municipality of residency. The estimated coefficient remains negative and significant for both sons and daughters, even using this restricted sample.

Alternatively, migration could be driven by the self-selected individuals. Our results show that IGM is relatively lower in microregions more affected by import competition. It could be that individuals with higher chances to exhibit upward IGM (i.e., enterprising individuals) tend to leave these regions to search for better opportunities in less affected areas, such as microregions that are focused on sectors where tariff cuts were shorter. If so, we would expect the trade shock to correlate with migration rates across regions. In columns (3) and (4) from Table A.8, we investigate the relationship between in-migration to a microregion and its exposition to trade liberalization where the outcome is an indicator variable equal to one if a child has migrated into the microregion and zero otherwise. All coefficients are statically insignificant and small in terms of magnitude compared to the long-run response of in-migration. In general, the results presented in Table A.8 indicate that migration is hardly a threat to our identification, as it is not determined by tariff reductions and excluding migrants in our sample does not affect the baseline results.

In Section 4.2, we defined our measure of upward occupational mobility as when a child works in a higher-ranked occupation than their parents, where the ranking of occupations was defined according to the average educational attainment of the individuals working in each occupation.

As an alternative ranking measure, we use a wage and earnings-based occupation ranking to assess our previous estimates' robustness.

We calculate each occupation's average wage and earnings to account for the difference between the formal and informal sectors and self-employed workers. It could be that regions facing higher liberalization also coincided with high participation of informal workers in those more relevant occupations. In this case, we would observe a difference in the estimated impact of the tariff cuts using the two measures. Table A.9 shows the effect of tariff cuts considering both measures of ranking. It is possible to observe that the estimated coefficient remains negative and statistically significant, similar to baseline estimations in Section 6 in terms of magnitude. This exercise demonstrates that our main results are robust to these alternate rankings.

9 Conclusion

In this paper, we leverage the Brazilian trade liberalization episode in 1991, which provided an external source of variation in import tariffs, to investigate the causal link between international trade and intergenerational mobility. Our method involves classifying individuals by occupation using an extensive survey dataset and correlating this with industry-specific tariff reductions across Brazilian regions. We then explore how these regional variations in exposure to trade liberalization influence upward mobility.

Our findings indicate that the liberalization shock resulted in lower levels of intergenerational mobility and had gender-distinct effects. Regions more impacted by the tariff reductions experienced a persistent relative decrease in occupational mobility and a long-run decline in absolute income and relative educational mobility compared to less affected regions. Moving a region from the 10 to the 90 percentiles in the distribution of the regional tariff cut correlates with a decrease in the absolute income mobility of approximately 7 percent for son-father pairs in the long run and 11 percent for daughter-mother pairs in the medium run. In turn, the proportion of sons with better occupations than their fathers presented a persistent reduction through the years (10 percent in the long run). In comparison, daughter-mother pairs presented a stable but permanent reduction of 7.5 percent.

We delve into the mechanisms behind these effects, first examining whether trade-induced declines in skill demand due to shifts in production towards less skilled sectors could explain our results. We show that the share of low-skilled occupations in the labor market increased 7% in areas more exposed to tariff cuts, indicating a decrease in the representation of high- and medium-skilled jobs relative to low-skilled ones. Furthermore, we also show differential effects on earnings across skill groups. While the high-skilled occupations were unaffected by the increased import competition, medium- and low-skilled workers suffered persistent and relative decreases in their earnings.

We contribute to the literature on the interplay between trade and intergenerational mobility (Ahsan & Chatterjee, 2017; Mitra et al., 2022) by investigating income, occupational, and educational mobility, considering both sons and daughters and demonstrating that the role played by trade shocks on mobility and its mechanisms differ across genders. Furthermore, our findings deviate from prior studies by highlighting a negative correlation between trade liberalization and social mobility levels within the specific context of Brazilian trade liberalization.

Our research underscores the adverse effects that trade can have on intergenerational mobility, particularly in widening the gap of socioeconomic differences between genders. This highlights the need for policymakers to design complementarity policies, such as redistributive policies and initiatives to promote more equality of opportunities for younger generations. These measures are crucial to counteract the potential frictions that trade can introduce into the social structure.

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Appendix - Trade Liberalization and Intergenerational Mobility

Appendix A - Additional Tables

Table A.1. Socieconomic indicators across states in Brazil: Unemployment, Poverty vulnerability and Gini index

Region	State	Unemployment	Poverty vulnerability	Gini coefficient
North	Acre	7.6	50.97	0.613
Northeast	Alagoas	12.1	59.76	0.572
North	Amazonas	12.4	51.78	0.509
North	Amapá	13.8	45.22	0.519
Northeast	Bahia	10.7	52.71	0.556
Northeast	Ceará	7.9	54.85	0.545
Midwest	Distrito Federal	11.5	16.00	0.624
Southeast	Espírito Santo	8.5	26.82	0.532
Midwest	Goiás	8.5	24.22	0.510
Northeast	Maranhão	8.4	63.58	0.538
Southeast	Minas Gerais	8.1	28.85	0.513
Midwest	Mato Grosso do Sul	7.2	26.83	0.521
Midwest	Mato Grosso	6.9	27.00	0.504
North	Pará	10.2	55.99	0.509
Northeast	Paraíba	9.8	53.65	0.591
Northeast	Pernambuco	13.1	51.86	0.553
Northeast	Piauí	5.2	58.13	0.555
South	Paraná	6.8	19.70	0.497
Southeast	Rio de Janeiro	9.2	22.26	0.542
Northeast	Rio Grande do Norte	11.7	47.70	0.559
North	Rondônia	9.1	33.33	0.509
North	Roraima	11.3	45.72	0.521
South	Rio Grande do Sul	6.7	18.65	0.500
South	Santa Catarina	5.6	12.36	0.460
Northeast	Sergipe	12.0	52.13	0.576
Southeast	São Paulo	9.8	16.13	0.489
North	Tocantins	7.9	44.71	0.523

Notes: The unemployment rate is the percentage of people who looked for, but did not find, a paid job among all those considered "active" in the labor market, a group that includes all people aged 10 years and over who were looking for a job or working in the reference week of the Pesquisa Nacional por Amostra de Domicílios (Pnad) - IBGE. Poverty vulnerability is calculated as the proportion of individuals with per capita household income equal to or less than R$ 255.00 per month, in reais in August 2010, equivalent to 1/2 the minimum wage on that date. The universe of individuals is limited to those living in permanent private households. Data is from the Brazilian Demographic Census - IBGE.
Source: Instituto de Pesquisa Econômica Aplicada (IPEA)

Table A.2. Education Indices (EIs) and Ranking of Occupations

Rank	Occupation EI	Code
1	11.082829	94
2	9.2163734	95
3	9.1868849	98
4	8.5965137	91
5	8.3719339	10
6	7.6996527	15
7	7.4031687	8
8	7.1902432	17
9	7.016593	93
10	6.8915944	12
11	6.7440915	5
12	6.0850735	32
13	6.0621581	96
14	5.8788533	97
15	5.7197461	25
16	5.6915159	22
17	4.8217802	4
18	4.7860026	2
19	4.7613087	14
20	4.2526064	92
21	2.4131267	1

Notes: Education indices are based on Brazilian Demographic Census - IBGE (1991) and computed using Equation X.

Table A.3. Effects of the Trade Liberalization on Skill Demand gap across Gender

	(1)	(2)	(3)	(4)	(5)	(6)
	High/Low		High/Medium		Medium/Low	
	2000	2010	2000	2010	2000	2010
Panel A: Male Workers						
RTC_c	-1.417***	-3.733***	1.121***	1.276***	-3.103***	-5.009***
	(0.243)	(0.477)	(0.361)	(0.375)	(0.291)	(0.411)
Observations	411	411	411	411	411	411
R-squared	0.523	0.709	0.647	0.693	0.822	0.853
Panel B: Female Workers						
RTC_c	4.252***	5.186***	5.819***	7.199***	-1.130***	-2.013***
	(0.290)	(0.559)	(0.462)	(0.531)	(0.356)	(0.404)
Observations	411	411	411	411	411	411
R-squared	0.864	0.872	0.849	0.891	0.446	0.597

Notes: Standard errors are clustered at the mesoregion level. Level of significance: *** at 1%, ** at 5%, * at 10%. The analysis proceeds at the micro-region level r. All regressions control for meso-region r time-fixed effects.

Table A.4. Effects of the Trade Liberalization on Earnings gap across Skill Groups by Gender

	(1)	(2)	(3)	(4)	(5)	(6)
	High/Low		High/Medium		Medium/Low	
	2000	2010	2000	2010	2000	2010
Panel A: Male Workers						
RTC_c	0.0370	0.0547	0.113	0.169*	-0.0366	-0.0955*
	(0.106)	(0.0942)	(0.122)	(0.101)	(0.0723)	(0.0548)
Observations	411	411	410	411	411	411
R-squared	0.275	0.179	0.386	0.225	0.397	0.304
Panel B:Female Workers						
RTC_c	0.929	-0.158	0.838	-0.110	-0.166	-0.433*
	(0.848)	(0.241)	(0.760)	(0.240)	(0.266)	(0.230)
Observations	406	408	402	406	408	408
R-squared	0.972	0.934	0.969	0.959	0.497	0.218

Notes: Standard errors are clustered at the mesoregion level. Level of significance: *** at 1%, ** at 5%, * at 10%. The analysis proceeds at the micro-region level r. All regressions control for meso-region r time-fixed effects.

Table A.5. Effects of the Trade Liberalization on the share of Independent Childs Head (Co-head) of Households

	(1)	(2)	(3)	(4)	(5)	(6)
		Son			Daugther	
	1980-1991	1991-2000	1991-2010	1980-1991	1991-2000	1991-2010
RTC_c	-0.145	-0.227	0.0105	-0.980*	-2.396***	-2.723***
	(0.146)	(0.139)	(0.168)	(0.556)	(0.332)	(0.546)
Observations	411	411	411	411	411	411
R-squared	0.683	0.523	0.459	0.656	0.757	0.712

Notes: Standard errors are clustered at the mesoregion level. Level of significance: *** at 1%, ** at 5%, * at 10%. The analysis proceeds at the micro-region level r. All regressions control for meso-region r time-fixed effects.

Table A.6. Selection bias

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
		Occupation			Income ($P \leq 50^{th}$)			Education	
Panel A: Son-Father									
RTC_c	-0.675***	-0.386***	-0.386***	-0.485***	-0.476***	-0.476***	-2.292***	-1.660***	-1.660***
	(0.053)	(0.040)	(0.040)	(0.046)	(0.046)	(0.046)	(0.095)	(0.069)	(0.069)
Observations	2,388,017	2,388,017	2,388,017	1,336,485	1,336,485	1,336,485	4,559,678	4,559,678	4,559,678
R-squared	0.055	0.103	0.103	0.027	0.062	0.062	0.034	0.224	0.224
Panel B: Daughter-Mother									
RTC_c	-0.666***	-0.574***	-0.574***	-0.217***	-0.264***	-0.264***	-0.869***	-0.470***	-0.470***
	(0.051)	(0.046)	(0.046)	(0.074)	(0.078)	(0.078)	(0.080)	(0.072)	(0.072)
Observations	539,451	539,451	539,451	246,315	246,315	246,315	3,612,333	3,612,333	3,612,333
R-squared	0.022	0.044	0.044	0.020	0.040	0.040	0.520	0.635	0.635
Individual Controls	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
Microregion Controls	No	No	Yes	No	No	Yes	No	Yes	Yes
Microregion Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Standard errors are clustered at the mesoregion level. Level of significance: *** at 1%, ** at 5%, * at 10%. The analysis proceeds at the micro-region level r. All regressions control for meso-region r time-fixed effects.

Table A.7. Selection bias: Different upper bounds for age cutoffs

	(1)	(2)	(3)	(4)	(5)	(6)
		1991 - 2000			1991 - 2010	
	25	35	40	25	35	40
Panel A: Son-Father						
RTC_c	-0.604***	-0.524***	-0.513***	-1.038***	-0.946***	-0.941***
	(0.132)	(0.127)	(0.127)	(0.174)	(0.165)	(0.164)
Observations	409	409	409	409	409	409
R-squared	0.729	0.734	0.736	0.825	0.838	0.838
Panel B: Daughter-Mother						
RTC_c	-0.755***	-0.679***	-0.684***	-0.793***	-0.731***	-0.745***
	(0.214)	(0.195)	(0.192)	(0.261)	(0.259)	(0.248)
Observations	408	408	408	408	408	408
R-squared	0.418	0.431	0.431	0.396	0.392	0.393

Notes: Cluster robust standard errors in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Unit of analysis is a micro-region. All regressions control for meso-region r time-fixed effects.

Table A.8. Robustness checks: accounting for migration

	(1)	(2)	(3)	(4)
	No immigrants		Migrated	
	2000	2010	2000	2010
Panel A: Son-Father				
RTC_c	-1.177***	-1.124***	-0.734	-0.122
	(0.282)	(0.153)	(0.676)	(0.235)
Observations	405	409	409	409
R-squared	0.583	0.813	0.739	0.684
Panel B: Daughter-Mother				
RTC_c	-1.179***	-0.570**	-0.568	-0.0312
	(0.408)	(0.255)	(0.646)	(0.229)
Observations	394	407	409	409
R-squared	0.292	0.383	0.738	0.660

Notes: Cluster robust standard errors in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Unit of analysis is a micro-region. All regressions control for meso-region r time-fixed effects.

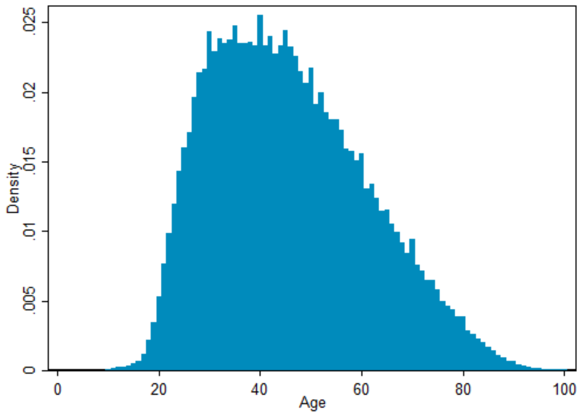
Table A.9. Robustness checks: wage and earnings-based ranking of occupations

	(1)	(2)	(3)	(4)
	Wage		Earnings	
	2000	2010	2000	2010
Panel A: Son-Father				
RTC_c	-0.419***	-0.798***	-0.433***	-0.786***
	(0.0786)	(0.112)	(0.0789)	(0.111)
Observations	409	409	409	409
R-squared	0.732	0.808	0.721	0.799
Panel B: Daughter-Mother				
RTC_c	-0.632***	-0.629***	-0.658***	-0.625**
	(0.187)	(0.213)	(0.203)	(0.244)
Observations	408	408	408	408
R-squared	0.502	0.451	0.423	0.381

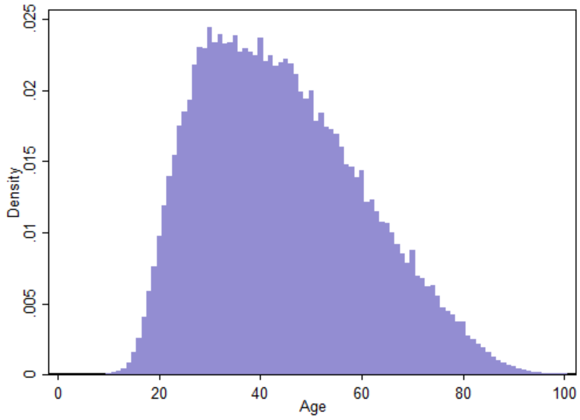
Notes: Cluster robust standard errors in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Unit of analysis is a micro-region. All regressions control for meso-region r time-fixed effects.

Appendix B - Additional Figures

Figure B.1. Households Head (or Co-Head) age distribution: Males and Females



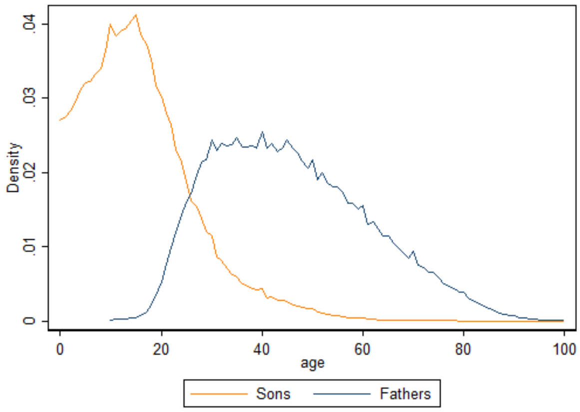
(a) Males



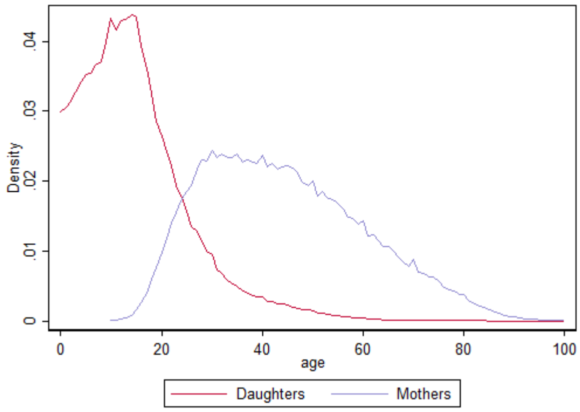
(b) Females

Note: The figures exhibit a histogram counting the age distribution of household head or co-head by gender.

Figure B.2. Sons-Fathers and Daughters-Mothers age distribution



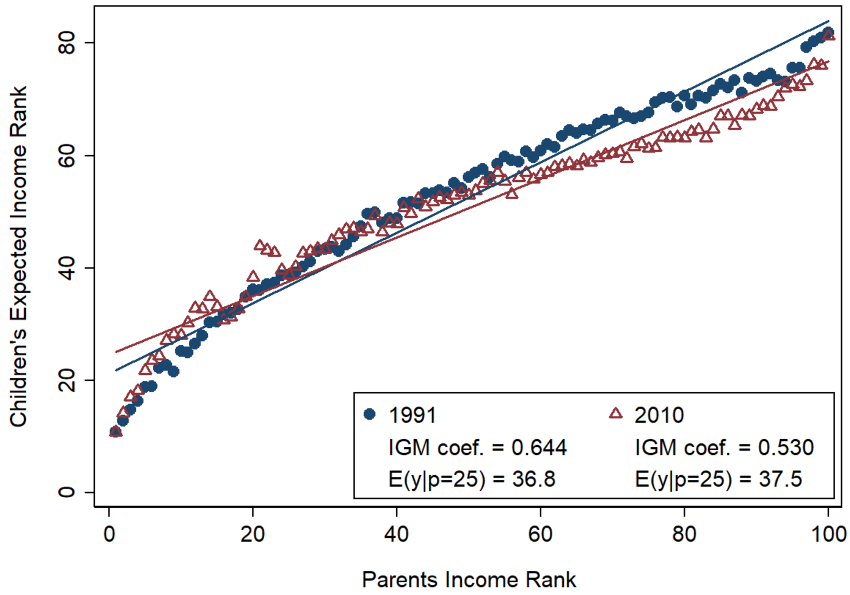
(a) Sons/Fathers



(b) Daughters/Mothers

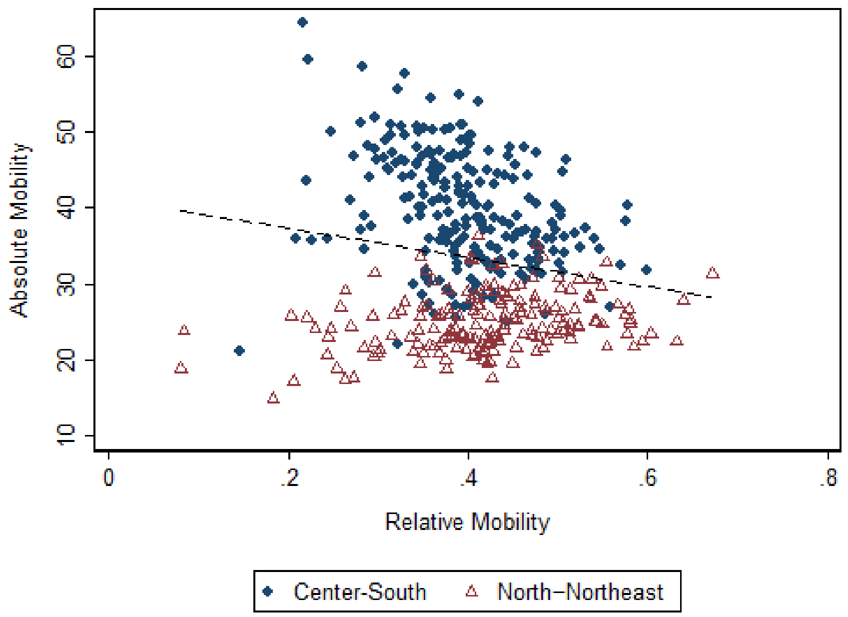
Note: The figures exhibit a histogram of Households counting the age distribution of children's and parents by gender. Daughters/Mothers.

Figure B.3. Mobility Curve in Brazil (1991 - 2010)



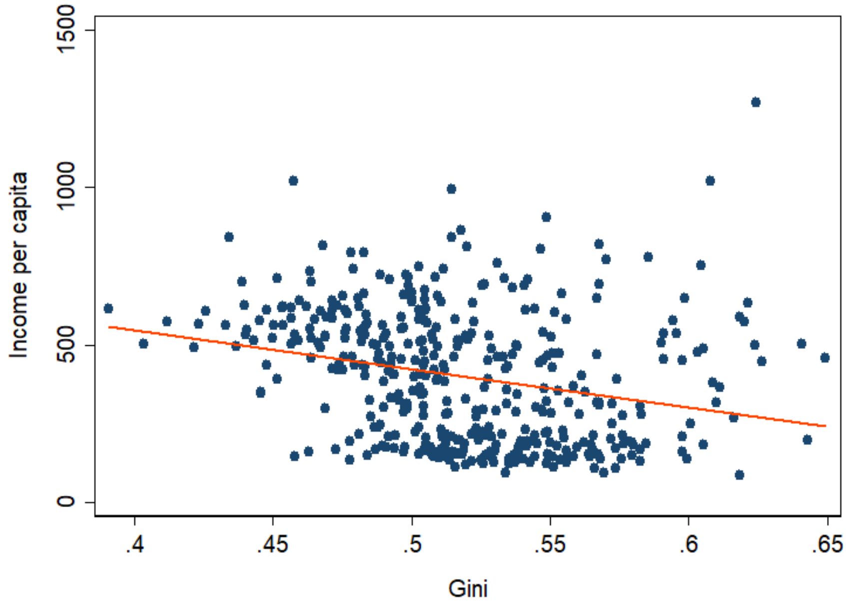
Note: The figure shows the association between parent and child income ranks at the national level, for the 1991 and 2010 census samples and children aged 15-30. The mean of child rank is computed according to each parental income percentile (calculated from the average between the father's and mother's income).

Figure B.4. Differences in Absolute vs. relative mobility across areas



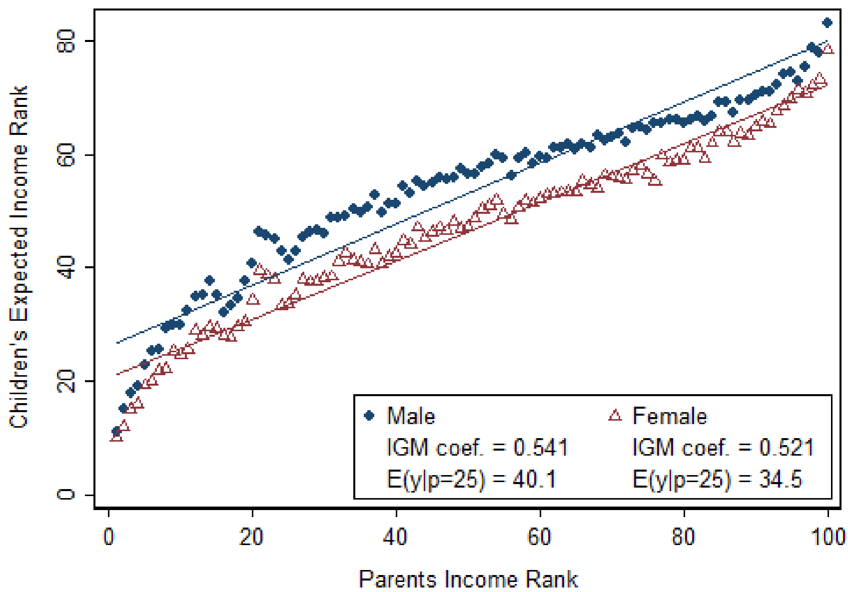
Note: The figure shows the relationship between relative mobility measured by the rank-rank slope and absolute mobility across Brazilian regions. Regions in the Center-South of the country are displayed in blue, and regions in the North-Northeast are displayed in red.

Figure B.5. The Great Gatsby Curve



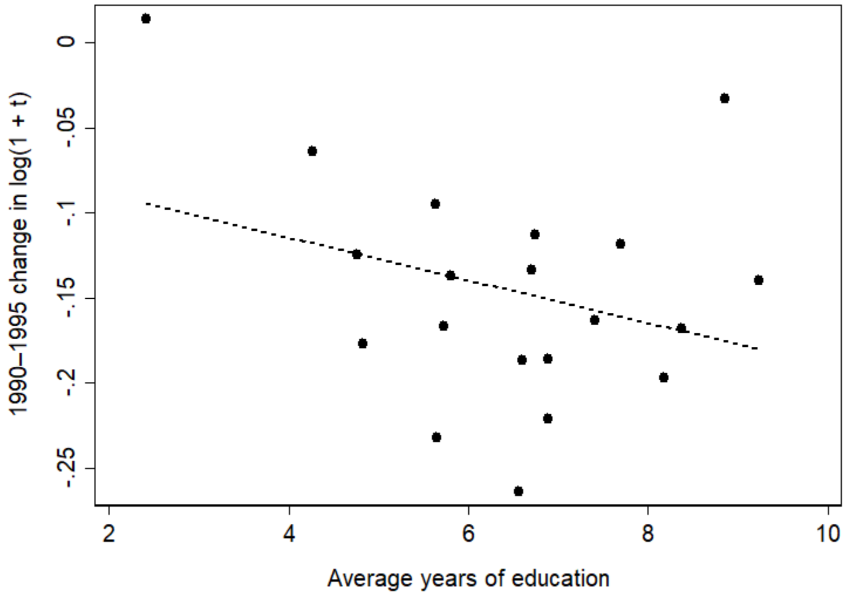
Note: The figure displays the relationship between income inequality (Gini index) and absolute mobility across Brazilian regions in 2010.

Figure B.6. Mobility Curves by Gender



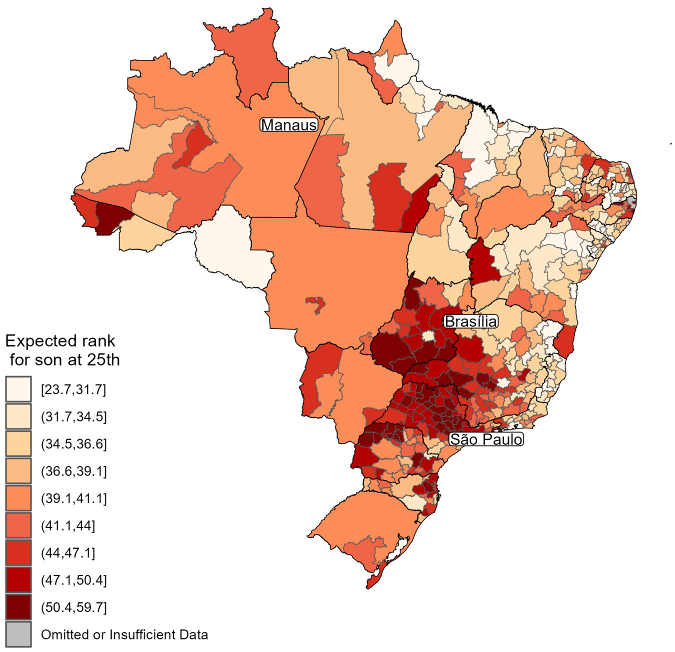
Note: This figure displays mobility curves by male and female children for the 1991 sample cohort. The mean of child rank is computed relative to the full sample (not separately by gender) and according to each parental income percentile (calculated from the average between the father's and mother's income).

Figure B.7. Relationship between Tariff Changes and Preliberalization Sectors' Average Years of Education



Note: This figure plots the average years of education of the workers in each sector and the change in tariffs from 1990 to 1995.

Figure B.8. Price-Adjusted Absolute Upward Mobility across regions



Note: The map represented in the figure depicts price-adjusted absolute income mobility for son-father pairs in 2010. Darker tones indicate higher absolute mobility. Local price indexes were constructed using POF survey data to deflate parent and child incomes. Children's ranks are calculated from the national income distribution relative to other children from the same age cohort.