

Trade Liberalization and Mortality Rates: Evidence of Pro-Cyclical Mortality from Brazil *

Vinicius Curti Cícero[†]

vinicius.cicero@colostate.edu

Lucas Corrêa-Dias [‡]

lucas.dias@fgv.edu.br

Sammy Zahran [§]

sammy.zahran@colostate.edu

Abstract

We trace the evolution of all-cause mortality rates in Brazilian regions with varying exposure to trade-induced economic shocks before, during, and after liberalization reforms in the 1990s. We find consistent evidence of pro-cyclical mortality, with areas more exposed to tariff reductions experiencing larger declines in mortality across varying time horizons. The observed decline in mortality rates is evident across sex, age groups, and for both internal and external causes of mortality. We falsify the observed relationship between mortality and tariff reductions with analyses of causes of death that are plausibly unrelated to economic activity. Concerning proximate mechanisms involved in our finding of pro-cyclical mortality, we show that healthcare infrastructure expanded in local economies more affected by the trade-induced economic shock. This expansion was characterized by the increased capital-intensity of care, facilitated by the import of diagnostic technologies that reduce mortality from internal causes. We also find supporting evidence for the idea that pro-cyclical mortality is partially caused by a decrease in transport and non-transport-related accidents. Overall, our findings highlight an underappreciated dimension of trade policy effects, namely public health.

Keywords: Health outcomes; Trade Liberalization; Local economic shocks; Pro-cyclical mortality; Healthcare infrastructure; Capital-intensity of care.

JEL codes: I15; I18; F13; F16; H51; H75.

*We benefited from fruitful discussion with and useful comments and impressions from Gilberto Tadeu Lima, Maria Dolores Montoya Diaz, Paula Pereda, Ray Miller and Timothy Komarek. We are also thankful for two anonymous referee and the participants of seminars at the University of Utah, University of Wyoming, and Colorado State University. Any remaining errors are our own.

[†]Corresponding author. Department of Economics, Colorado State University, 1771 Campus Delivery, Fort Collins CO 80523-1771, USA

[‡]São Paulo School of Economics - EESP/FGV, Brazil.

[§]Department of Economics, Colorado State University; Department of Epidemiology, Colorado School of Public Health, USA

1 Introduction and Motivation

In the late 1980s, the Brazilian economy was protected against external competition by a complex system of trade barriers (Kume et al., 2003; Kovak, 2013). With the election of Fernando Collor de Mello, the new Brazilian administration launched a series of trade liberalization reforms involving the harmonization of tariff levels across all industries of the economy. From 1990 to 1995, the average tariff decreased from 30.5 percent to 12.8 percent but varied considerably across industries. Sectors like agriculture and mining experienced negligible changes in tariffs, while others, such as apparel and rubber underwent declines of roughly 30 percentage points (Dix-Carneiro and Kovak, 2015). Given preexisting regional differences in industry mix, these trade reforms produced strikingly different regional economic impacts.

Differential exposure to trade reforms by region produced variation in labor demand shocks, that in turn caused measurable changes in labor markets and firm survival across Brazilian regions. Dix-Carneiro and Kovak (2015) note that “regions that initially specialized in industries facing larger tariff cuts experienced prolonged declines in formal sector employment and earnings relative to other regions” with these labor market effects operating across workers of varying levels of education, age, sex, and employment tenure. These trade-induced economic shocks to local labor markets also caused changes in criminogenic conditions that led to measurable increases in homicide mortality across affected regions in Brazil (Dix-Carneiro et al., 2018).

Taking inspiration from Dix-Carneiro et al. (2018), we widened the scope of the mortality analysis by tracing the evolution of all-cause and cause-specific mortality rates in Brazilian regions before, during, and after liberalization reforms.¹ Brazil’s trade reform in the 1990s provides an excellent analytical setting in which to study the causal relationship between mortality and economic dynamics. As with previous research (Ruhm, 2000, 2015), we find consistent evidence of pro-cyclical mortality in Brazil, with the local economies most affected by trade-induced shocks witnessing substantial declines in all-cause mortality rates. Charris et al. (2023) find a similar result of relative decline in infant mortality in locations facing larger tariff reductions in Brazil. Our results indicate that the mortality effect is considerably more extensive and consistently observed for all subgroups by age and sex and across cause-specific sources of mortality. Our placebo exercises confirm that pre-reform mortality trends in each region bore no correlation to the subsequent trade-induced shocks. Moreover, the relationship between mortality and trade-induced economic shock is falsified with analyses of causes of mortality unrelated to economic activity like exposure to forces of nature and poisonous animals and plants.

These observations affirm that our findings capture the causal impact of trade-induced shocks on mortality. Our baseline specification indicates that a region facing a tariff reduction of 0.1 log points - reflecting a shift from the ninetieth percentile to the tenth percentile of regional tariff changes - exhibited a corresponding reduction of about 0.1 log points in

¹For contextualization, homicide-related fatalities constituted a minor fraction accounting for less than 4% of the total deaths within the country on average between 1985 and 2010.

mortality rates. This translates to a 10 percent decrease in the all-cause mortality rate, five years subsequent to the completion of liberalization reforms, and about 0.2 log points (or 18 percent), fifteen years after the reform.

Our investigation into the mechanisms underlying pro-cyclical mortality reveals important insights, adding to the existing body of literature in health economics. First, we corroborate previous work (Ruhm, 2000, 2015), finding that the decline in mortality from external causes is attributable to a reduction in transport and non-transport-related accidents. Second, and unique to the Brazilian case, we uncover that the observed relative decline in mortality rates from internal causes over the long term in regions more exposed to tariff cuts can be attributed, at least partially, to a relative expansion of healthcare infrastructure (particularly in capital-intensive machinery) in the localities more impacted by the shock.

Following the enactment of Brazil's federal Constitution and the establishment of the *Sistema Único de Saúde* (Unified Health System - SUS) in the late 1980s, we find evidence of increased spending per hospitalization and hospital procedures, as well as a notable increase in non-basic procedures compared to basic ones (i.e., more capital-intensive procedures) within the outpatient system of SUS in the local economies more impacted by the trade-induced economic shock. Specifically, we document a substantial increase in procedures aimed at detecting malignant tumors in these regions. These findings provide empirical support for the hypothesis that the reduction in deaths from internal causes, particularly from cardiovascular diseases and neoplasms in the harder-hit local economies was due mainly to a pronounced expansion of healthcare infrastructure toward prevention and diagnostic services, surpassing the growth observed in areas with lower exposure to tariff reductions. Of pivotal importance, our study uncovers a distinctive facet of the trade liberalization episode. We provide evidence that the trade-induced economic shock impacted the accessibility and affordability of imports which, in turn, directly facilitated the expansion of capital-intensive healthcare infrastructure and “life-saving” technologies, particularly within regions bruised by tariff adjustments.

The present study contributes to the extensive literature on the relationship between economic shocks and health outcomes (see Ruhm (2012) for a comprehensive review). Existing research consistently demonstrates that death rates, particularly at regional levels (Lindo, 2015), tend to decline during economic recessions and rise during economic upturns in developed countries.² However, evidence concerning developing countries is limited and less conclusive compared to the observed patterns in high-income nations.³ We extend the literature by examining the medium and long-term dynamics of mortality associated with a lasting shock that predominantly impacted urban regional markets within a developing economy. Particularly, our analysis capitalizes on a distinctive episode of trade liberalization, which

²Various studies conducted for the United States (Ruhm, 2000, 2003, 2005, 2015; Miller et al., 2009; Stevens et al., 2015) and other developed nations (van den Berg et al., 2017; Ballester et al., 2019; Haaland and Telle, 2015) have observed this pro-cyclical pattern.

³For example, Gonzalez and Quast (2010) and Arroyave et al. (2015) document counter-cyclical mortality patterns in poorer areas of Mexico and among working-age men in Colombia, respectively. Hone et al. (2019) demonstrate that the recent Brazilian recession (2014-2016) led to increased mortality at the municipal level, although health and social protection expenditures appeared to mitigate adverse health effects.

closely resembles a once-and-for-all event, providing us with a valuable opportunity to address identification challenges commonly encountered in country-level studies, establishing a causal relationship between local economic shocks and mortality rates. Lastly, this paper also contributes to a recent body of work that examines the implications of economic shocks resulting from trade policy changes on adult health outcomes (Autor et al., 2019; Pierce and Schott, 2020; McManus and Schaur, 2016; Lang et al., 2019; Fan et al., 2020; Feng et al., 2021).

In the next section, we describe data sources and our empirical strategy in pursuit of the causal relationship between mortality and trade-induced economic shocks. In Section 3, we report results on all-cause mortality and then results on age-specific mortality rates and cause-specific sources of death. Section 4 presents a robustness check on our identification strategy. In Section 5 we examine potential mechanisms underlying the effect of the trade-induced shock on internal causes of mortality, focusing on cardiovascular diseases and neoplasms, and, lastly, Section 6 concludes this paper with a recapitulation of the results and a discussion of implications.

2 Data Description and Empirical Strategy

2.1 Trade Liberalization and Local Economic Shocks

In the era prior to liberalization, the Brazilian economy was regulated by a wide array of protective measures aimed at limiting competition from abroad. These measures encompassed both non-tariff barriers and tariffs (Kume et al., 2003). Subsequent to trade liberalization initiatives launched by the newly elected government in March 1990, there was a notable decline in the average import tariffs across various industries. From 1990 to 1995 tariffs decreased by an average of approximately 17 percentage points. The standard deviation in nominal tariffs decreased from 14.9 percent to 7.4 percent in the same period (Dix-Carneiro, 2014), pushing the country toward greater harmonization of tariff levels across industries (Dix-Carneiro et al., 2018).⁴ Figure 1 shows the percentage change in tariffs across main industries.

Because of preexisting regional variation in industry mix, these tariff reductions impacted the regions of Brazil heterogeneously. Following the literature on the regional labor market effects of foreign competition (Dix-Carneiro, 2014; Dix-Carneiro and Kovak, 2015; Ponczeck and Ulyssea, 2022; Kovak, 2013; Hirata and Soares, 2020), our measurement of trade-induced shocks to local labor demand exploits the coincidence of sector-specific tariff change and the preexisting composition of employment across sectors at the regional level. The average tariff change faced by region r weighted by the importance of each sector in regional employment - our shift-share or “Bartik” instrument (Bartik, 1991; Borusyak et al., 2022) - is defined as

⁴The correlation coefficient between tariff cuts between 1990 and 1995 and the pre-liberalization tariff levels (in 1990) is near to -0.9, as sectors with initially higher tariffs experienced larger subsequent reductions. Figure A.1 presents a simple visualization of the relationship between tariff changes and pre-liberalization levels for each industry.

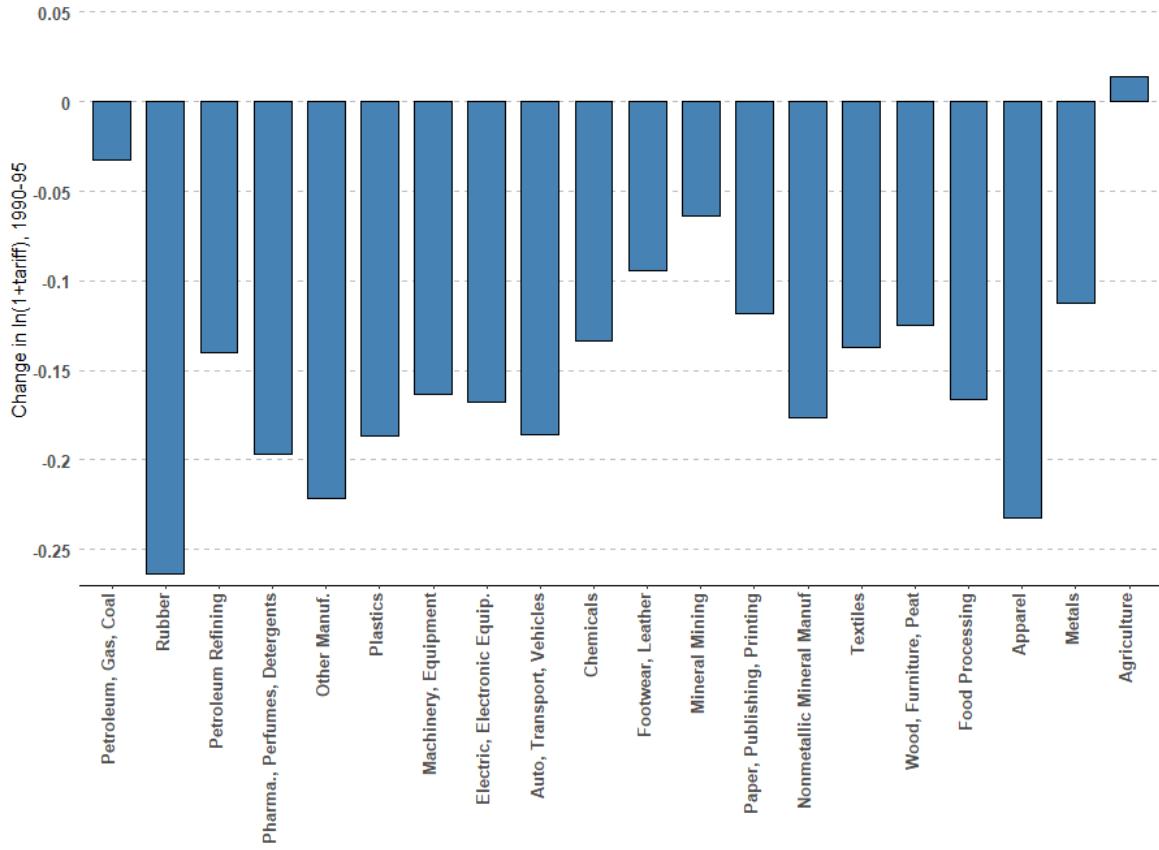


Figure 1: Nominal tariff changes, 1990-1995

Notes: Following Dix-Carneiro and Kovak (2017), we aggregate tariff data from Kume et al. (2003) to allow for a classification that is consistent with the Demographic Census data used to construct local tariff shock measures. Industries are sorted - from left to right - by increasing participation in terms of national employment in 1991.

follows:

$$RTC_r = \sum_{i \in T} \eta_{ri} \Delta \log(1 + \tau_i), \text{ with } \eta_{ri} = \frac{\frac{\lambda_{ri}}{\delta_i}}{\sum_{j \in T} \frac{\lambda_{rj}}{\delta_j}} \quad (1)$$

where τ_i is the tariff on industry i , λ_{ri} is the initial share of region r workers employed in industry i , δ_i equals one minus the wage bill share of industry i and T denote the set of all tradable industries. From Equation (1) it is evident that the magnitude of the trade-induced regional shock depends on how the local tradable sector is affected.⁵

⁵For a detailed discussion of how the non-tradable sector is incorporated in this measure, see Kovak (2013).

2.2 Data Description

Our analysis is conducted at the micro-region level, involving groupings of economically integrated municipalities with similar geographic and productive characteristics. Micro-regions are defined by the Brazilian Institute of Geography and Statistics (IBGE - *Instituto Brasileiro de Geografia e Estatística*) and are commonly used in economic literature to characterize local labor markets in Brazil (Dix-Carneiro and Kovak, 2017; Dix-Carneiro et al., 2018; Ponczek and Ulyssea, 2022; Hirata and Soares, 2020). Our analysis deploys a crosswalk between municipalities and micro-regions detailed in Dix-Carneiro and Kovak (2015), arriving at a set of 411 repeatedly observed local economies.⁶ Table A.2, in the Appendix, provides descriptive statistics at the micro-region level for the main variables used in our empirical analysis.

2.2.1 Tariff Changes

The tariff data used in this paper is provided by Kume et al. (2003), and is extensively used in the literature on trade and labor markets in Brazil (see, for instance, Kovak (2013), Dix-Carneiro and Kovak (2015, 2017)). We focus on changes in output tariffs to construct our measure of trade-induced local labor demand shocks (or regional tariff changes) described in Equation (1). Previous studies show that analyses using changes in effective rates of protection negligibly change results obtained using output tariff changes (Dix-Carneiro et al., 2018; Ponczek and Ulyssea, 2022).

2.2.2 Mortality Data

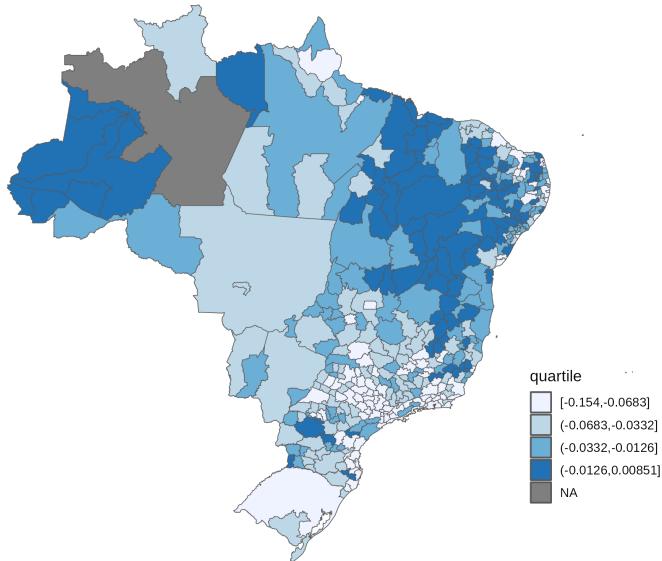
We use mortality records from DATASUS (administrative dataset from the Ministry of Health), available at the municipality level from 1979-1995 (ICD-9) and 1996-2020 (ICD-10). Although data are available since 1979, not all municipalities are observed until 1985 (Charris et al., 2023). Therefore, we exclude the 1979-1984 period from our analysis.⁷ For each specific cause of mortality, we compute the number of obits by municipality in each year and then aggregate to the micro-region level. Population data from four census waves (described in detail below), were used to calculate mortality rates per 100,000 inhabitants. Cause-specific sources of mortality examined in this paper are described in Table A.1.

Figure 2 provides a visual representation of the spatial relationship between regional tariff shocks and mortality rates across micro-regions in Brazil. In Panel (a), the spatial distribution of regional tariff shocks is depicted, with colors indicating quartiles of the regional tariff change variable. Lighter shades of blue denote higher exposure to tariff cuts, particularly evident in historically more developed regions such as the Southeast and South of Brazil.

⁶Although we systemically observe 413 micro-regions, we exclude the regions of “Manaus” and “Fernando de Noronha” due to insensitivity to the trade liberalization reform.

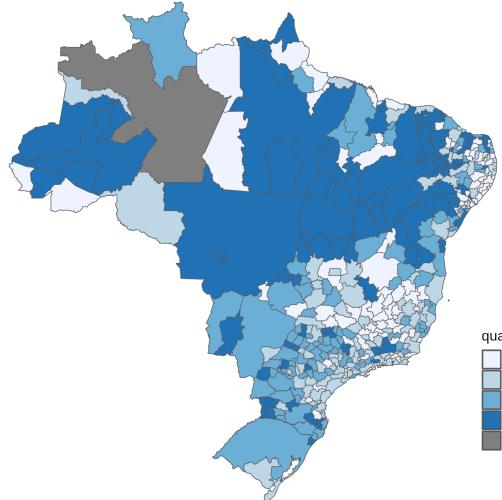
⁷Excluding mortality data from 1979 to 1984 is imperative to ensure the validity and reliability of our analysis due to significant measurement error. Importantly, including these years does not substantially alter our core findings, which remain robust both qualitatively and quantitatively. Detailed discussions on the robustness of our results and the validity of the parallel trends assumption can be found in the Online Appendix, serving as supplementary material to the article.

Regional Tariff Change (RTC)



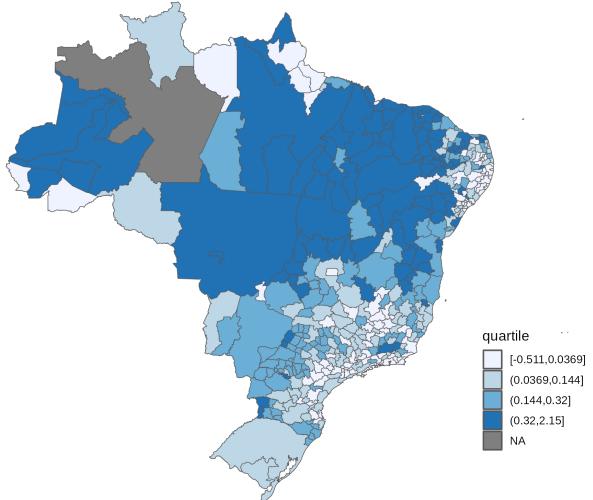
(a) Distribution of regional tariff changes

Change in log mortality rates 2000-1991



(b) Distribution of log changes in local mortality rates: 2000–1991

Change in log mortality rates 2010-1991



(c) Distribution of log changes in local mortality rates: 2010–1991

Figure 2: Pre-trends, regional tariff changes, and post-liberalization log changes in mortality rates

Source: Mortality rates per 100,000 inhabitants computed from DATASUS. Regional tariff changes, RTC_r , are computed according to Equation (1).

Panels (b) and (c) display the distribution of log changes in all-cause mortality rates during the medium-run post-liberalization period (1991-2000) and the subsequent long-run period (1991-2010), respectively. In these panels, lighter shades of blue signify smaller increases or larger decreases in regional mortality rates. Similar to the regional tariff changes, significant variation is observed in mortality rate shifts across regions in both timeframes.

An intriguing finding emerging from our analysis is the spatial coincidence between regions most affected by regional tariff shocks and those experiencing declines in mortality rates, particularly noticeable in the bottom quartile of changes in log mortality rates. This observation suggests a potential correlation between exposure to regional economic shocks and reductions in mortality rates. These initial findings align with existing research on pro-cyclical mortality, indicating complex connections between economic dynamics and health outcomes.⁸

2.2.3 Other Variables

We use four waves of the Brazilian Decennial Population Census, from IBGE, covering thirty years (from 1980-2010) to compute population sizes of micro-regions, as well as distributions by sex and age groups. Toward the investigation of mechanisms involved in pro-cyclical mortality, we computed annual government spending per category at the municipality level with data from the Ministry of Finance (*Ministério da Fazenda - Secretaria do Tesouro Nacional*), the number of health establishments from the *Pesquisa de Assistência Médico-Sanitária* (1992, 1999, 2002) and *Cadastro Nacional de Estabelecimentos de Saúde* (2005-2010), expenditures from the Brazilian Unified Health System (SUS - *Sistema Único de Saúde*) on outpatient care and procedures rates (per 100,000 inhabitants), hospital expenditures, hospitalization rates, and procedures of detection of malign tumors (measured in per 100,000 inhabitants).

2.3 Identification

Following Dix-Carneiro and Kovak (2017, 2015) and Dix-Carneiro et al. (2018) we assess the dynamic response of mortality rates to trade-induced regional economic shocks using the following specification:

$$\Delta_{91-t} \log(kMR_r) = \log(kMR_{r,t}) - \log(kMR_{r,1991}) = \beta_t RTC_r + \alpha_{s,t} + \varepsilon_{r,t} \quad (2)$$

where $kMR_{r,t}$ is the k -specific mortality rate, described in Table A.1, in region r at time $t > 1991$ and $\alpha_{s,t}$ are state-time fixed effects.

Note that the difference-in-differences specification described in Equation (2) analyzes variation in RTC_r across micro-regions within states, providing transparent treatment-

⁸Supplementary visualizations of this correlation over time, including high-level analysis of the relationship between regional tariff change and log changes in local mortality rates over both the medium and long run, are available in the Online Appendix. The visual evidence indicates that regions most profoundly impacted by the trade shock also experienced the most significant reductions in log mortality rates.

control comparisons (Dix-Carneiro et al., 2018). In all specifications, we cluster the standard errors at the meso-region (grouping of micro-regions also defined by IBGE) level to account for potential spatial correlation in outcomes.⁹

Recent research has provided a formal framework to establish the identifying assumptions for shift-share regression designs (Goldsmith-Pinkham et al., 2020; Borusyak et al., 2022; Borusyak and Hull, 2023). Building on the work of Goldsmith-Pinkham et al. (2020) and Borusyak et al. (2022), the identifying assumption in our specific context is that the trade-induced economic shock - RTC_r - is independent of local political and institutional dynamics across micro-regions.¹⁰ This assumption is corroborated by the substantial correlation between the tariff cuts and the pre-liberalization tariff levels, which, in turn, were determined by the level of protection established in previous decades. Consequently, concerns related to the political economy of the tariff reductions are alleviated, as sectoral and regional peculiarities appear to have minimal influence (Dix-Carneiro and Kovak, 2017; Dix-Carneiro et al., 2018; Goldberg and Pavcnik, 2007).

To ensure a causal interpretation of our estimates, a crucial assumption is that without the trade liberalization reform, local economies in Brazil would have experienced similar changes in mortality rates. We follow Dix-Carneiro and Kovak (2017) and incorporate pre-liberalization outcome trends in our analysis, which helps address potential confounding factors varying with trade-induced shock exposure. Unobservable shocks reflecting pre-existing long-run trajectories are accounted for by the pre-trend outcome control. To assess the validity of our research design, we provide a comprehensive set of estimates encompassing the periods prior to, during, and following the trade liberalization reform.

If our identifying assumption holds true, it is expected that areas with greater exposure to the reform and those with lower exposure would exhibit similar mortality trajectories before the reform, with divergence occurring only after its implementation. Our results broadly support this assumption. Furthermore, we conduct a falsification test to examine potential misleading correlations between declining mortality in local economies and the magnitude of tariff cuts. Our findings provide evidence that specific mortality rates, which are theoretically unrelated to economic conditions, remain unaffected by trade-induced regional economic shocks. Overall, the evidence underscores the unique nature of the Brazilian trade liberalization episode starting in the early 1990s, serving as a natural experiment facilitating the identification of the impacts of local economic shocks on mortality rates.

⁹In the Online Appendix (supplemental material to the manuscript), we show the robustness of our results by using the inference procedures recommended by Borusyak et al. (2022) to address cross-region residual correlation in shift-share designs. While Adao et al. (2019) propose an alternative method for standard error estimation in such designs, it is important to note that Ferman (2022) has raised concerns about its suitability in settings like ours, with a relatively small number of industries where it may lead to excessive over-rejection (Alvarez et al., 2022; Ogeda et al., 2024).

¹⁰Given that we employ a linear shift-share design, where the exposure shares in all micro-regions sum up to one, the identification concerns raised by Borusyak and Hull (2023) do not apply to our specific research setting.

3 Results

3.1 Pro-cyclical Mortality

Table 1 presents estimates for Equation (2), describing effects for all individuals.¹¹ We start with a specification absent controls and weighting of observations. In column 1, our results indicate that there is a significant positive relationship between changes in mortality rates and regional tariff changes. The magnitude of the coefficient decreases marginally but remains statistically significant with the incremental saturation of the model, involving the weighting of the observations by the average population between 1991 and 2000 - for the medium-run - and, 1991 and 2010, for the long-run (column 2), the inclusion of state fixed effects (column 3) and a variable capturing the pre-period trend in mortality rates (columns 4 and 5).

Following Dix-Carneiro et al. (2018), we address concerns that preexisting trends in region-specific mortality rates could be correlated with (future) trade-induced local shocks. In column 4 we include this trend variable as an additional control and estimate the equation by ordinary least squares. A potential problem with this procedure is that the log of 1991 mortality rates appears on the right- and left-hand side of the estimating equation, potentially introducing a mechanical bias in the estimators (Dix-Carneiro et al., 2018). This problem is solved by using a ratio of the number of total obits in 1990 and 1985, $\left(\frac{\text{TotalObits}_{r,1990}}{\text{TotalObits}_{r,1985}}\right)$, as an instrument for the preexisting trends of mortality rates in a 2SLS estimation. In both cases there are modest changes in the coefficients of interest, suggesting that our estimated relationship between changes in mortality rates and regional economic shocks is not driven by preexisting trends. The coefficients associated with such pre-trends are not statistically significant (at the 5% significance level) in any of the specifications in Table 1. Going forward, and for ease of exposition, we only present the results of the specification of column 3, involving the weighting of observations by the population and state fixed effects.¹²

The medium-run results in Table 1 indicate that the effect of regional tariff changes on mortality rates is substantive: a change in RTC_r equivalent to -0.1 log points is accompanied by a decrease in all-cause mortality rates of at least 0.1 log points, or 10 percent. To provide context on the effect size, a micro-region at the mean of the 1991 mortality rate distribution with an average population size in 2000, would experience a decrease of approximately 400 deaths with a tariff shock of this size. Interestingly, the effects of the trade-induced regional economic shock on all-cause mortality rates appear stronger in the longer run, with a change in RTC_r equivalent to -0.1 log points being accompanied by a decrease in all-cause mortality rates of approximately 0.2 log points, or 18 percent. Utilizing similar back-of-the-envelope calculations, it can be inferred that a micro-region characterized by the average mortality

¹¹We provide results disaggregated for males and females separately in the Online Appendix. Remarkably, the findings demonstrate substantial similarity, both quantitatively and qualitatively, to the effects observed for all individuals.

¹²We show in detail the notable stability of the coefficients associated with the impact of the trade-induced shock on all-cause mortality rates for our preferred specification in the Online Appendix (supplemental material to the manuscript).

rate from the 1991 distribution and possessing the average population in 2010 would witness a remarkable reduction of around 800 deaths in comparison to the observed average of 2700 deaths in the year 2010.

For perspective, the standard deviations of $\Delta_{91-00} \log(MR_r)$ and $\Delta_{91-10} \log(MR_r)$ across micro-regions are, respectively, 0.26 and 0.32 log points, so we document an increase of approximately 42% of a standard deviation in decadal changes in log mortality rates in the medium run and approximately 63% of a standard deviation in bi-decadal changes in log mortality rates in the long run caused by the trade-induced economic shock.¹³ Overall, our results point to a strong positive relationship between the all-cause mortality rate and regional tariff changes.

The trajectory of the trade liberalization episode reveals a consistent pattern of increasing effects on mortality rates over time, mirroring the longer-run dynamics of employment documented in Dix-Carneiro and Kovak (2017).¹⁴ Indeed, the consistency in the dynamic impacts of the trade reform on mortality and employment outcomes suggests that changes in local economic conditions – as highlighted by the pro-cyclical mortality literature – play a crucial role in explaining our findings.¹⁵

Existing literature has extensively documented adverse employment and earnings effects in regions more heavily impacted by tariff reductions in Brazil (Dix-Carneiro and Kovak, 2017; Dix-Carneiro et al., 2018; Kovak, 2013; Gaddis and Pieters , 2017; Dix-Carneiro and Kovak, 2015). Although the relative deterioration of local labor market conditions in the medium run is evident in the regions most impacted by the trade shock, the long-term results presented in the literature vary depending on the measurement of employment rates and wages. Of particular note is the nuanced interplay between adjustments in formal and informal employment post-trade reform, a facet necessitating closer scrutiny.¹⁶

Dix-Carneiro and Kovak (2017) document that formal sector employment rates were consistently and increasingly impacted over time but, importantly, the authors also document that the long-run recovery in employment rates experienced by harder-hit regions reflects relative increases in informal employment, while formal employment keeps falling. Further analysis by Charris et al. (2023) underscores this pattern, revealing a significant increase

¹³In view of results presented by Dix-Carneiro et al. (2018), the intensification of pro-cyclical mortality conforms to the subsidence of the initial increase in the homicide rate, a component of all-cause mortality during the period analyzed in Brazil.

¹⁴Notably, both employment reductions and mortality rate decreases show substantial growth from 2000 to 2010, indicating a continued divergence for regions facing different exposures to the tariff reform. A comprehensive analysis of these concomitant dynamics is provided in the Online Appendix.

¹⁵We provide suggestive evidence, in the Online Appendix, that the significant decrease in mortality rates post-liberalization is primarily concentrated in regions characterized by higher levels of economic activity, specifically in terms of employment rates prior to liberalization. This observation further supports the pro-cyclical pattern observed in mortality rates.

¹⁶We present evidence in the Online Appendix that echoes the findings of the literature. Specifically, our findings reveal a substantial relative decline in formal employment among both men and women over the medium and long term in regions more impacted by trade-induced regional economic shock. However, the overall employment rate (that accounts for informality) exhibits a different trajectory, with its significant impact fading away in the long term.

in self-employment among males alongside a meaningful and permanent drop in overall employment rates for women. The authors also find that greater exposure to the tariff cuts is associated with lower aggregate household income in the medium run, with an intensification of the result in the long run – reflecting the persistent and amplifying deterioration in economic conditions of the regions more impacted by the tariff cuts.

The correspondence in the dynamic impacts of the trade reform on mortality and employment underscores the role of changes in local economic conditions — and hence, the pro-cyclical mortality story — as a key mechanism underlying our main findings, particularly in the medium run. We further explore the temporary nature of the effects of the aggregate employment rate in the broader context of the mortality findings, with disaggregations by cause-specific mortality rates, in the subsequent sections.

A possible concern with the results above is that RTC_r may be correlated with pre-existing trends in the outcome of interest. Besides the inclusion of preexisting trend variables as additional controls, Column 6 of Table 1 presents evidence that regional tariff changes are uncorrelated with pre-trends by regressing pre-liberalization changes in mortality directly against (future) trade shocks (that is, using $\Delta_{85-91}\log(MR_r)$ as the dependent variable). The non-significance of the coefficients in our placebo test corroborates the previous evidence obtained with the inclusion of trend variables in the estimations.

Table 1: Regional tariff changes and log changes in mortality rates

	OLS (1)		OLS (2)		OLS (3)		OLS (4)		2SLS (5)		Placebo (6)
Dep. var.: $\Delta\log(MR_r)$	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1985-1991
All											
RTC_r	1.599*** (0.449)	3.162*** (0.567)	1.164*** (0.275)	2.496*** (0.346)	1.099*** (0.240)	1.957*** (0.278)	1.099*** (0.276)	1.958*** (0.310)	1.099*** (0.249)	1.958*** (0.296)	-0.00567 (0.312)
$\Delta_{85-91}\log(MR_r)$							-0.188 (0.162)	-0.256 (0.171)	-0.104 (0.194)	-0.239 (0.211)	

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. In column 1, observations are not weighted; in column 2, observations are weighted by population; column 3 adds state fixed effects to column 2; column 4 adds pre-trends to column 3; column 5 shows two-stage least squares, with an instrument for $\Delta_{85-91}\log(MR_r)$. Column 6 presents a placebo test, with observations weighted by population and considering state fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Figure 3 presents a graphical representation of the dynamic effects of tariff reductions on all-cause mortality rates. Importantly, the trade-induced shock starts to affect the mortality rate only after the end of the trade liberalization episode, with all-cause mortality increasingly reducing over time.

The model specified in Equation (2), the baseline for our main results, represents a first-difference specification. Alternatively, we can estimate a standard dynamic difference-in-differences model instead of the first-difference model where our measure of exposure to the trade-induced regional economic shock is interacted with year indicators and using the (log of) mortality rates in levels instead of relative changes. Importantly, this event-study design aligned with the recent literature (Roth et al., 2023; Borusyak et al., 2024). In this case, the equivalent dynamic differences-in-differences specification to Equation (2) is given

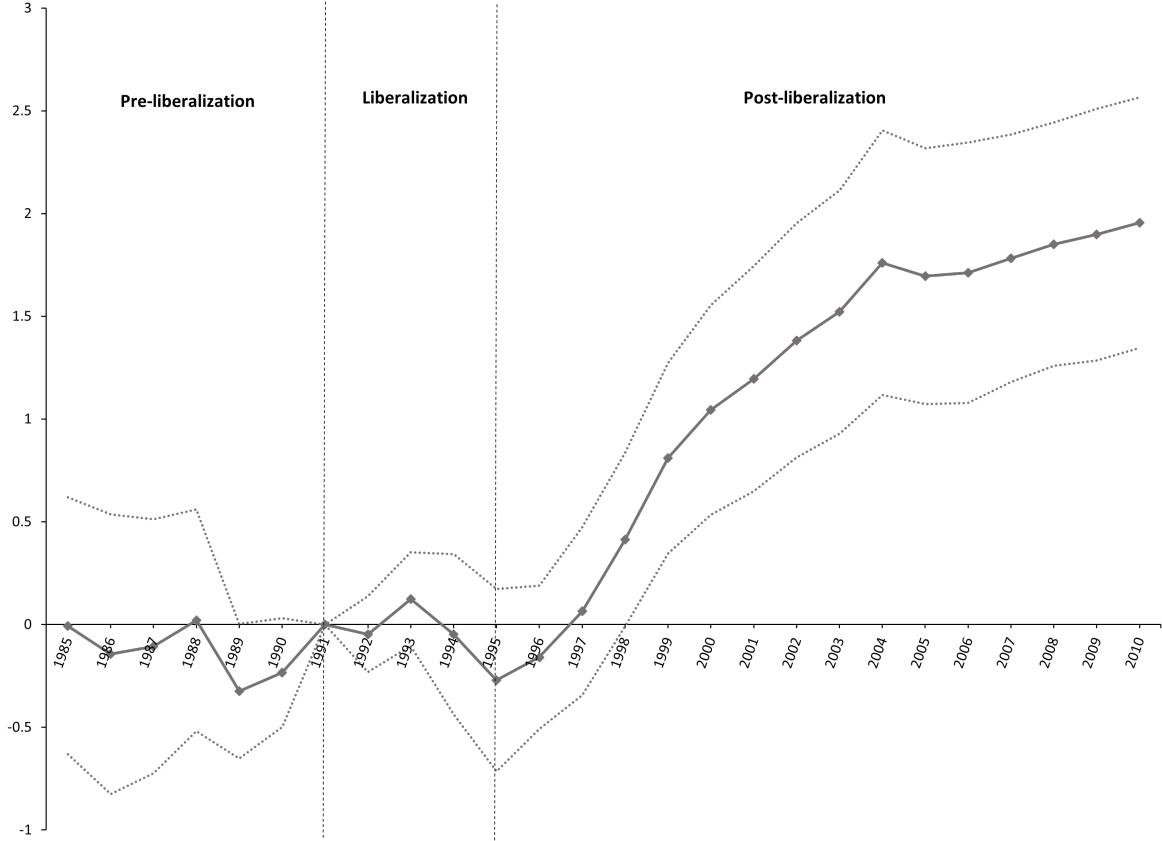


Figure 3: Dynamic effects of regional tariff changes on log changes in mortality rates

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (2), where the dependent variable is the change in regional log all-cause mortality rates and the independent variable is the RTC_r in $t = 1985, \dots, 2010$. All regressions include state fixed effects. Dashed lines show 95 percent confidence intervals. Standard errors are adjusted for 91 meso-region clusters.

by:

$$y_{r,t} = c + \sum_{t=1985}^{2010} \beta_t \mathbb{1}\{\tau = t\} RTC_r + \sum_{t=1985}^{2010} \gamma_t \mathbb{1}\{t > 1991\} (y_{r,1991} - y_{r,1985}) + \mu_r + \alpha_{s,t} + \varepsilon_{r,t} \quad (3)$$

In the specification described in Equation (3), we set 1991 as the baseline treatment year and, as before, RTC_r is our treatment variable. μ_r now represent the micro-region fixed effects and the terms $\mathbb{1}$ are years indicators. It is important to note that, since all micro-regions were treated at the same time by federal legislation in 1991, this empirical design does not suffer from the recent methodological criticisms of the difference-in-differences literature (Callaway and Sant'Anna, 2021; De Chaisemartin and d'Haultfoeuille, 2020, 2022; Goodman-Bacon, 2021).

Our estimation of this equivalent difference-in-differences specification yields results that closely mirror those derived from the first-difference model, demonstrating the consistency and robustness of our primary findings. Figure 4 displays the coefficients β_t and their respective 95 percent confidence intervals obtained from estimating Equation (3). We include, in Figure 4, the coefficients associated with the years 2000 and 2010 for comparison with our main results presented in Table 1.¹⁷

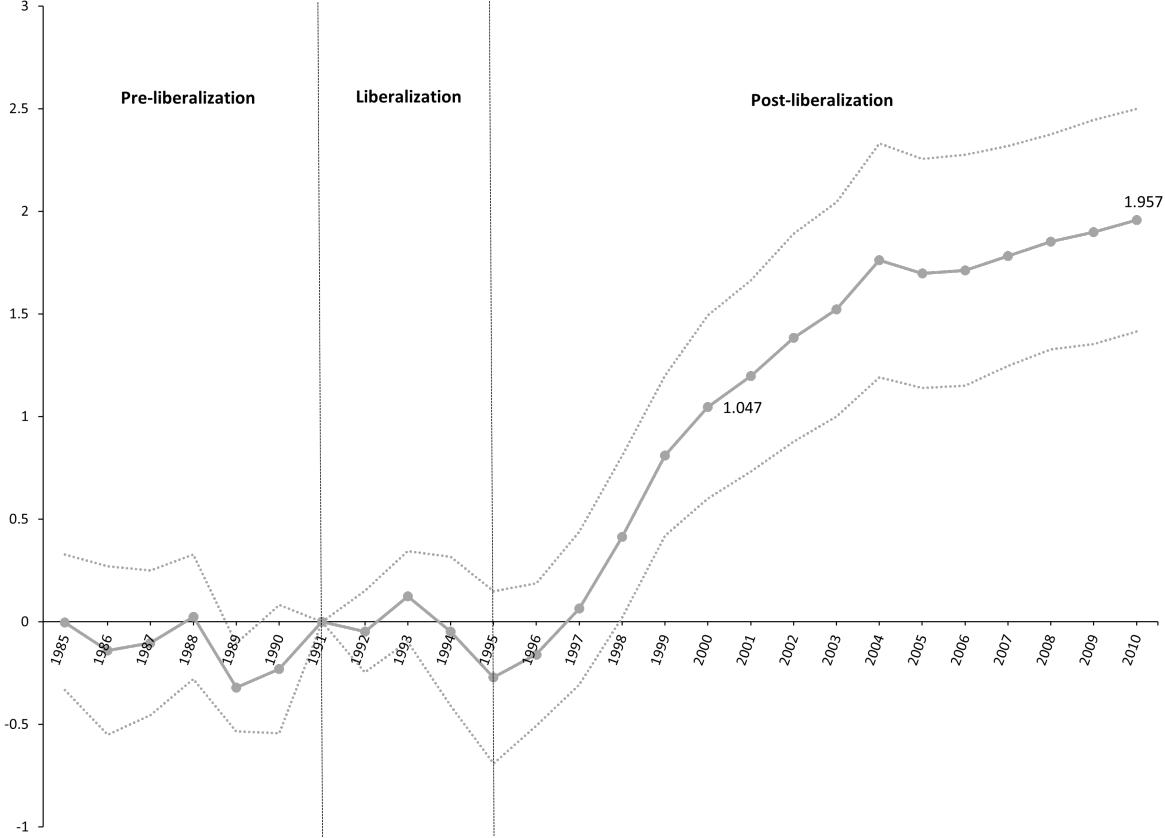


Figure 4: Dynamic effects of regional tariff changes on log changes in mortality rates - Difference-in-Difference

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (3), where the dependent variable is the regional log all-cause mortality rate in year $t = 1985, \dots, 2010$. The regression includes micro-regions fixed effects and state-year fixed effects. Standard errors are adjusted for 91 meso-region clusters and the observations are weighted by population. Dashed lines show 95 percent confidence intervals.

¹⁷For enhanced clarity regarding the magnitude of all estimated coefficients, we present the estimation results of the dynamic difference-in-differences model in the Online Appendix.

3.2 Age-Specific Effects

Because the risk of mortality increases with age, the declining trend in mortality might reflect changes in the age structure of micro-regions as opposed to shifting economic conditions. During our period of analysis, birth rates increased roughly three times faster than death rates, decreasing the average age of residents across micro-regions. To address this issue, we recapitulate our analysis for six different age groups. Table 2 presents the results from the estimation of the effect of local economic shocks on mortality rates for each age group. In the second column of Table 2, we include the share of deaths out of the total (average from 1985 - our initial data point - to 2010) for each group to help discern the economic relevance of estimated effects.

Table 2: Regional tariff changes and log changes in group-specific mortality rates

Type of Mortality	Share of deaths Average 1985-2010	Estimated coefficients	
		1991-2000	1991-2010
All deaths	1.000 (0.240)	1.099*** (0.240)	1.957*** (0.278)
Sex-specific			
Males	0.579	1.095*** (0.246)	2.103*** (0.245)
Females	0.421	1.141** (0.276)	1.788*** (0.368)
Age-specific			
0-14	0.131	2.859*** (0.657)	3.242*** (0.907)
15-24	0.041	1.348 (0.884)	2.157*** (0.727)
25-44	0.126	0.990*** (0.292)	2.653*** (0.283)
45-64	0.238	1.452*** (0.227)	2.764*** (0.335)
65-74	0.178	0.576*** (0.264)	1.749*** (0.315)
75+	0.286	0.355 (0.270)	1.131*** (0.296)

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. In all regressions, observations are weighted and state fixed effects are added. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

The results are quantitatively similar to those presented in column 3 of Table 1. The strongest impact of the local economic shock on mortality rates is observed in the first age group, from 0 to 14 years, in both the medium and long run. Infant mortality accounts for the largest share of deaths in this age interval of 0 to 14. This result is compatible with Charris

et al. (2023), showing a reduction in infant mortality at the municipality level in Brazil after the trade liberalization shock. The authors indicate that public policies pertaining to healthcare access focused on prenatal and newborn care had a significant impact on reducing infant mortality in the country, especially in the second half of the 2000s.

Concerning the medium run, the results for persons between 24 and 64 years of age are statistically significant and similar to the average effect of the trade-induced economic shock.¹⁸ The result is less pronounced for elderly populations.¹⁹ In the longer run, we observe an intensification of estimated effects for all age groups. In particular, the effect of tariff reductions amplifies for working-age populations, as expected from the direct relationship with the unemployment rate. Overall, the results in Table 2 suggest that the results reported in Table 1 are not driven by changes in age structure.

3.3 Cause-Specific Sources of Mortality

Next, we investigate the impacts of the traded-induced economic shocks on cause-specific sources of mortality. Following Ruhm (2015), we separately examine three diseases and four external causes of mortality. The three disease categories are: cardiovascular, cancer, and other diseases²⁰, accounting for 28%, 12%, and 49% of all deaths over the 1985–2010 period (average), respectively. The four external sources are: transport accidents, other (non-transport) accidents, homicides, and suicides. Those causes were responsible for 3.5%, 2.8%, 3.9%, and 0.7% of all deaths, respectively. Additionally, we decompose non-transport accidents into four specific types – falls, drowning/submersion, smoke/fires/flames, and poisoning/exposure to noxious substances. Table 3 presents the results for each source of death in both the medium and long run.

As shown in Table 3, the reduction in mortality from external causes is greater than that from diseases in regions most affected by the economic shock. In particular, deaths from transport accidents are most impacted by the tariff reduction - a change in RTC_r equivalent to -0.1 log points is accompanied by a large decrease in the transport accidents mortality rate of almost 0.6 log points (45 percent) in the medium run and of 0.8 log points (55 percent) after 20 years. This result is consistent with Ruhm (2000) and Miller et al. (2009) who argue that an increase in the unemployment rate reduces motor vehicle miles traveled and therefore number of fatal traffic accidents.²¹ Results also indicate that the

¹⁸The apparent reduction in the effect for persons 15–24 years of age in the medium run may be related to the increase in the homicide rate documented in Dix-Carneiro et al. (2018), which disproportionately affects this age group.

¹⁹The population over 75 years of age in Brazil was remarkably small in the early decades of our analysis. For example, according to census data, less than 2.5% of the country’s total population in 2000 was over 75 years old.

²⁰The “other diseases” category is defined as encompassing all internal causes of mortality excluding those attributed to cardiovascular diseases or malignant neoplasms. As such, it covers a wide range of internal causes including infectious and parasitic diseases, digestive system issues, congenital malformations, and other abnormalities.

²¹Another contributing factor to this outcome could stem from trade liberalization’s influence on import affordability. This, in turn, could have facilitated the revitalization of the country’s automotive fleet, po-

economic shock arising from trade liberalization efforts significantly reduced the mortality rate from non-transport accidents in Brazil. Intuitively, because of the observed increase in the unemployment rate, fatal non-transport accidents in work settings decreased. Besides, as documented in the literature, increased household care due to higher unemployment could also contribute to a reduction in accidents at home (e.g. accidental falls as in Ruhm (2000)).

Table 3: Regional tariff changes and log changes in cause-specific mortality rates

Source of death	Share of deaths	Estimated coefficients	
		Average 1985-2010	1991-2000 1991-2010
Diseases	0.891	0.935*** (0.202)	1.703*** (0.276)
Cardiovascular disease	0.282	2.822*** (0.772)	6.226*** (0.968)
Cancer	0.120	2.695*** (0.676)	6.042*** (0.876)
Other diseases	0.489	0.607** (0.255)	0.107 (0.355)
External causes	0.109	2.557** (0.985)	4.830*** (0.678)
Transport accidents	0.035	5.950*** (1.080)	8.093*** (1.671)
Other accidents	0.028	3.033*** (1.015)	4.705*** (1.123)
Suicides	0.007	1.560 (1.140)	2.177 (2.023)
Homicides	0.039	-3.855*** (1.445)	-1.311 (2.462)
Other accidents	0.028	3.033*** (1.015)	4.705*** (1.123)
Falls	0.006	2.620* (1.349)	2.398*** (0.694)
Drowning/submersion	0.007	5.209*** (0.725)	9.070*** (0.990)
Smoke/fire/flames	0.001	4.619*** (1.159)	3.982*** (0.949)
Poisoning/noxious	0.001	6.434** (3.108)	2.108 (1.803)

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. In all regressions, observations are weighted and state fixed effects are added. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

tentially reducing accidents. We show, in the Online Appendix, a notable surge in imports of automotive vehicles and parts in aggregate for Brazil, especially during the 2000s.

Examining these findings alongside our prior discussion on employment adjustments in regions most affected by tariff reductions prompts intriguing inquiries into the interplay between labor market dynamics and mortality trends. Why might the reintegration of a segment of the workforce into paid activities, even within self-employment or informal sectors, not impact mortality trends from external causes? One plausible hypothesis is that informal employment, often entailing reduced commuting (potentially associated with lower household income) and potentially leading to increased time spent at home, could correlate with a sustained decline in mortality rates, encompassing both traffic accidents and other incidents, whether occupational or otherwise. For example, a partial return to self-employment in informal sectors might lead adults to spend more time at home, thereby bolstering supervision of children and the elderly and potentially diminishing the risk of fatal accidents. Noteworthy is the marked decrease in drowning-related deaths, which appears to influence other accident categories in the long term, such as those linked to smoke/fire/flames, falls, and poisoning. Despite these shifts, it remains crucial to recognize that a significant portion of the working-age population, particularly women, continues to encounter restricted opportunities in the local labor market over the long term. These hypotheses suggest the potential continuation of the pro-cyclical trend in the long run for these causes of death.

Concerning deaths from disease, mortality rates from cardiovascular diseases and cancer are greatly reduced by trade-induced economic shocks in the medium run, and with the magnification of this effect in the long run - a change in RTC_r equivalent to -0.1 log points are accompanied by decreases in the mortality rates from cardiovascular disease and cancer of more than 0.26 log points (roughly 23 percent) in the medium run and of more than 0.6 log points, or 45 percent, after 20 years. By contrast, deaths from other diseases are less markedly reduced in the medium run, with the coefficient on the tariff change losing statistical significance in the long run. These results suggest that, although in the medium term the reduction in mortality from diseases may be explained by a common mechanism – the deterioration of local labor market conditions –, in the long run it may be governed by other forces. We analyze in more detail the potential mechanisms behind the results for internal causes of mortality in Section 5.

Importantly, this divergence in long-term effects across disease categories may be influenced by the relative enhancement in local labor markets, possibly driven by the reintegration of prime working-age men into informal employment. This resurgence in employment, albeit informal, might limit access to medical care for diagnosis or treatment and could lead to a resurgence in unhealthy habits (Ruhm, 2005; Cawley and Ruhm, 2011). Thus, while not directly impacting mortality from external causes, the rise in informal employment could mitigate the long-term impact of mortality reduction by improving local economic conditions. This apparent smoothing effect on mortality decline, possibly due to the pro-cyclical nature of mortality, is evidenced by reductions in deaths from other diseases.²²

²²To establish the causal effects of trade liberalization on regional cause-specific mortality rates, we evaluate the validity of the parallel trends assumption by analyzing the dynamic impacts of trade liberalization on cause-specific mortality rates in the Online Appendix. Our findings reveal that the significant impact of the trade-induced shock on both external and internal mortality rates becomes evident only after the trade

4 Falsification

In the previous section, we presented extensive empirical evidence demonstrating that the Brazilian trade liberalization episode - which largely impacted the economic conditions of regional economies - led to notable decreases in mortality rates in the areas with greater exposure to the trade-induced economic shock. These findings support the notion that mortality rates are influenced by economic conditions and follow a pro-cyclical pattern during the analyzed period. Moreover, our investigation revealed a consistent downward trend in mortality across various age groups following the tariff reduction shock, attenuating doubt that our results merely reflect changes in the age distribution of the Brazilian population.

In addition to using instruments for pre-trends in the variables of interest and conducting placebo tests across our main specifications, we implement a falsification test to further assess the robustness of our findings. The logic is simple: if an omitted variable affects all mortality rates, then we would anticipate significant effects of regional tariff changes on causes of mortality unrelated to the economic conditions of local areas or economic cycles in general. For this falsification test, we examine mortality causes not linked to economic activity, specifically deaths due to natural forces or exposure to toxic animals and plants.

The results are summarized in Table 4. It is direct to note that, from the specifications in which we weight the observations by the population of each micro-region and consider state fixed effects (from column 2 to column 6), there is no statistical significance of the coefficients associated with the economic shock in the specific mortality rate in either time window.

Table 4: Regional tariff changes and log changes in “nature” mortality rates

	OLS (1)		OLS (2)		OLS (3)		OLS (4)		2SLS (5)		Placebo (6)
Dep. var.: $\Delta \log(NMR_r)$	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1985-1991
RTC_r	2.389*** (0.831)	1.280 (1.486)	0.986 (1.046)	-7.350 (4.616)	1.780 (1.281)	-2.393 (2.119)	0.802 (1.169)	-2.836 (2.129)	1.698 (1.289)	-1.364 (2.015)	-1.513* (0.879)
$\Delta_{85-91} \log(NMR_r)$						-0.626*** (0.103)	-0.275** (0.134)	-0.0525 (0.162)	0.629 (0.369)		

Notes: There are 408 micro-region - the unit of analysis - observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. In column 1, observations are not weighted; in column 2, observations are weighted by population; column 3 adds state fixed effects to column 2; column 4 adds pre-trends to column 3; column 5 shows two-stage least squares, with an instrument for $\Delta_{85-91} \log(NMR_r)$. Column 6 presents a placebo test, with observations weighted by population and considering state fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

These results indicate that there are likely no omitted factors or variables in our previous estimates that would have a universal impact on all mortality rates during the analyzed period. Our findings suggest that specific mortality rates, which are theoretically independent of economic conditions, remain unaffected by the trade-induced economic shock. This falsification test serves as a robustness check for our empirical approach and supports the credibility of our identification strategy.

liberalization episode concludes, affirming the parallel trend assumption. Additionally, our results support the hypothesis that while the medium-term decline in mortality rates from both external and internal causes may stem from the same mechanism – deterioration in local labor market condition – the long-term trajectories diverge significantly.

5 Potential Mechanisms for Internal Causes

Next, we explore possible mechanisms underlying the effect of the trade-induced shock on internal causes of death.²³

5.1 Healthcare Infrastructure

The enactment of Brazil's new federal Constitution in 1988, a few years before trade liberalization reforms, brought about a substantial increase in social spending aimed at fostering social development across the country. This constitutional reform had a profound impact on the country's institutional framework, particularly within the health sector. Brazil introduced the *Sistema Único de Saúde*, which now stands as the largest publicly-funded healthcare system globally. The legislative process leading to the creation of this health system ensured mandatory government spending on the health sector to guarantee universal access to healthcare as a right, resulting in the expansion of healthcare infrastructure across different regions of the country. Moreover, health spending underwent decentralization from the 1990s onwards, aligning with a constitutional principle of decentralization in the administration of healthcare (Paim et al., 2011).²⁴

Figure 5 depicts the evolution of public spending on “health and sanitation” (henceforth, H&S), as well as total expenditures by state and municipal governments in Brazil by year from 1990 to 2010. Panel (a) highlights the remarkable growth of H&S spending relative to total state expenditures, particularly from the late 1990s onwards. In 2010, Brazilian states collectively spent nearly five times more on H&S than they did in 2000. A similar trend is observed for municipal health expenditures, as shown in panel (b), with a four-fold increase between 2000 and 2010. To provide perspective, panel (c) presents the ratio of health and sanitation expenditures to total expenditures for both levels of government. The impact of decentralization in health management is evident, particularly for municipalities, with a sharp increase in health spending relative to total expenditures. As noted earlier, the most significant expansion in H&S spending occurred in the late 1990s or early 2000s, coinciding with the largest impacts of the trade-induced regional economic shock on internal causes of mortality.

The precise reasons behind the uneven decline in mortality rates from cardiovascular disease and neoplasms in regions that experienced the greatest impact of the trade shock in the long run remain unclear. One plausible hypothesis suggests that the documented surge in healthcare expenditures has been allocated to the development of medium and high-complexity healthcare infrastructure, enhancing the accessibility and effectiveness of

²³It is worth noting that Dix-Carneiro and Kovak (2017) and Charris et al. (2023) find no evidence of shifts in migration patterns across local economies in response to the trade-induced economic shock. Consequently, the evidence suggests that selective migration is unlikely to be a key factor driving our results.

²⁴Article 198 on the Brazilian Constitution indicates that, among other things, within a regionalized and hierarchical network, public health actions and services should function as a unified system, guided by the principle of decentralization. It also describes the allocation of obligatory resources for public health actions and services for all spheres of government.

prevention, diagnosis, and treatment for these particular diseases. Notably, evidence of this phenomenon appears to have emerged in Brazil since the late 1990s.

Table 5 presents an overview of the supply and evolution of selected diagnostic imaging equipment - more directly related to these specific diseases - in Brazil from 1999 to 2005. In terms of the whole country, the number of units increased 20% compared to 1999, with the most significant change occurring from 2002 to 2005. Notably, magnetic resonance imaging (MRI), mammography, color Doppler ultrasound, and X-ray machines for hemodynamics experienced the most substantial changes during this period. The growth of other types of X-ray machines remained below the average, as did other ultrasound machines. This suggests that while simpler equipment, though more abundant in number, exhibited modest growth, the more advanced equipment demonstrated relatively greater expansion (IBGE, 2009).

Table 5: Diagnostic imaging equipment per selected type in Brazil: 1999-2005

Equipment	Total			Variation (%)	
	1999	2002	2005	Δ_{05-99}	Δ_{02-99}
Mammography	2065	2498	3245	57,1	21,0
X-Ray for Hemodynamics	355	451	537	51,3	27,0
Other X-Ray	17069	18538	18720	9,7	1,0
MRI	285	433	549	92,6	51,9
Tomography	1515	1617	1961	29,4	6,7
Ultrasound	11500	11849	14242	23,8	3,0
- <i>Doppler</i>	3921	4638	6185	57,7	18,3
- <i>Others</i>	7579	7211	8057	6,3	-4,9
Total	32789	35386	39254	19,7	7,9

Source: Data from IBGE (2009) and *Pesquisa de Assistência Médico-Sanitária*; 1999, 2002 and 2005.

Furthermore, the expansion of the supply of diagnostic equipment associated with the specific diseases that showed the largest reductions in mortality rates was even more pronounced when comparing 1999 to 2009. For instance, data from the 2009 Survey of Medical-Sanitary Assistance indicates that the number of MRI machines grew by 320% in a decade, while tomography equipment grew by nearly 100%. In terms of basic equipment for diagnosing cardiovascular diseases, the number of electrocardiographs increased by more than 60% over the same period.

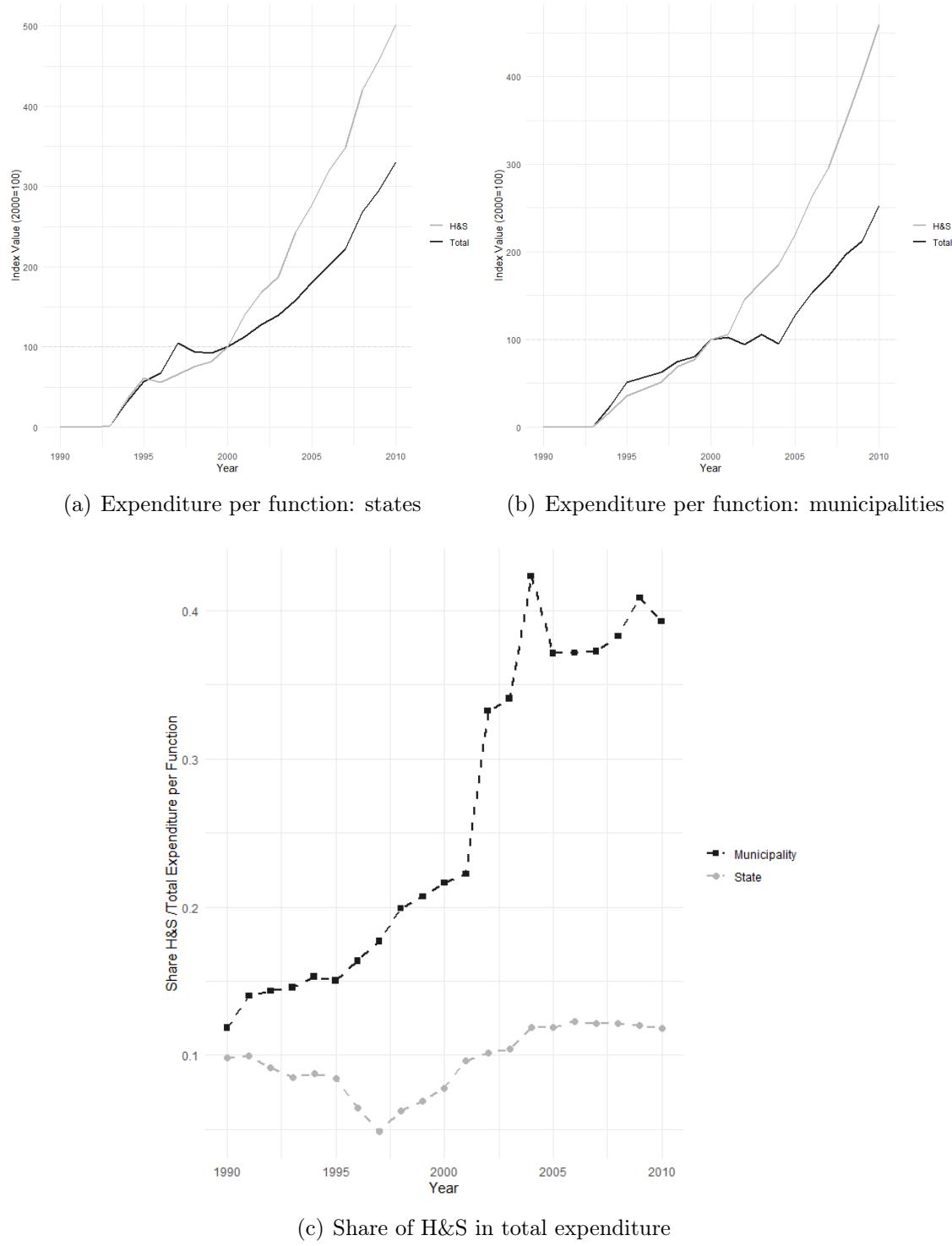


Figure 5: Trends in public expenditure for states and municipalities: 1990-2010

Source: Annual government spending per category both at the municipality and the state level is from the Ministry of Finance (STN - *Secretaria do Tesouro Nacional*).

According to IBGE (2009), the distribution of healthcare services in Brazil was characterized by significant inequality, a situation that was exacerbated when it came to the availability of selected diagnostic imaging equipment. Importantly, this unequal distribution pattern mirrors the regional distribution of the economic shocks resulting from trade liberalization. Regions such as the North and Northeast, which experienced relatively smaller changes in regional tariffs during the liberalization episode, faced significant delays in terms of the availability of advanced diagnostic imaging equipment compared to the national average. In contrast, the Southeast and South regions, which were more heavily impacted by tariff changes, witnessed a high concentration of such equipment throughout the 2000s. To provide context, the Southeast region alone accounted for approximately 58% of the CT scanners and over 61% of the MRI scanners in health establishments across the country, despite representing only 42% of the total population.

The available evidence thus far suggests some notable trends, particularly from the late 1990s onwards, which are of significant relevance to understanding the mechanisms underlying the impacts of the trade-induced economic shock on internal causes of mortality. First, there has been a substantial rise in health expenditures by municipal and state governments in Brazil, leading to the expansion of healthcare infrastructure throughout different regions of the country. Secondly, this expansion of healthcare infrastructure is closely linked to a marked increase in the availability of medium and high-complexity exams and procedures, at least concerning the necessary equipment associated with these procedures. Furthermore, it is worth noting that the increase in diagnostic and treatment equipment, particularly useful for cardiovascular diseases and neoplasms, appears to be concentrated in the more affluent regions of the country.²⁵ Intriguingly, this expansion seems to coincide geographically with exposure to tariff reductions during the trade liberalization episode.

5.2 Impacts of Trade Liberalization

To examine the plausibility of this hypothesis, our investigation delves into evaluating the impacts of the trade-induced regional economic shock on various indicators related to healthcare infrastructure.²⁶ These variables include government expenditure dedicated to healthcare, the proportion of total spending allocated to health, the production output of the Unified Health System (SUS), and the increase in advanced diagnostic services and procedures within the micro-regions of Brazil. Consistent with the empirical analyses previously presented in this study, we treat these variables as dependent variables and regress on our measure of the

²⁵Sub-sample analyses are available in the Online Appendix, showing that the majority of the estimated impact of tariff reductions on lowering regional mortality rates is associated with regions with larger populations, higher average labor remuneration, greater per capita income, and a more robust local labor market before liberalization.

²⁶In the Online Appendix, we explore additional channels through which the trade-induced economic shock might have impacted mortality rates from internal causes following liberalization. These factors encompass shifts in air quality, the opportunity cost of medical care, the availability of healthcare personnel, and regional educational outcomes. Nonetheless, the findings suggest the limited significance of these mechanisms in ameliorating regional mortality outcomes.

economic shock arising from the trade liberalization episode, RTC_r . The specification aligns with Equation (2). The outcomes of these analyses are summarized in Table 6.²⁷

Table 6: Potential mechanisms: regional tariff change and outcomes

Panel A. Local Government Spending		Expenditures per function		Expenditures H&S		Share H&S / Total	
		1995-2000 (1)	1995-2010 (2)	1995-2000 (1)	1995-2010 (2)	1995-2000 (1)	1995-2010 (2)
RTC_r		1,774 (1,553)	3,310*** (569.3)	-161.8 (240.2)	-236.5 (200.5)	-0.405*** (0.115)	-1.002*** (0.145)
R^2		0.019	0.496	0.091	0.336	0.337	0.379

Panel B. SUS Expenditures and Health Establishments		Expenditures (SIH) Hospital System		Expenditures (SIA) Outpatient System		Total Medical Establishments	
		1995-2000 (1)	1995-2010 (2)	1995-2000 (1)	1995-2007 (2)	1992-1999 (1)	1992-2010 (2)
RTC_r		-22.48** (10.69)	-338.1*** (43.37)	-87.25*** (25.71)	-153.5*** (34.28)	-0.157 (0.686)	-3.262*** (0.583)
R^2		0.455	0.455	0.415	0.340	0.462	0.573

Panel C. Capital-intensity of Healthcare Infrastructure		Hospitalization Rate		Outpatient Procedures Non-Basic / Basic Ratio		Diagnoses Rate Neoplasm Detection	
		1992-2000 (1)	1992-2010 (2)	1995-2000 (1)	1995-2007 (2)	1995-2000 (1)	1995-2005 (2)
RTC_r		-0.222 (0.632)	-1.273 (0.822)	-2.428*** (0.879)	-2.030* (1.230)	-10.60*** (1.960)	-9.514*** (2.023)
R^2		0.219	0.192	0.204	0.305	0.421	0.35

Notes: The expenditure variables in Panels A and B are measured in per capita changes. Hospitalization and procedures rates, in Panel C, as well as the total medical establishments in Panel B are given by the changes of logs of the variables measured in per capita terms over the indicated period. Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. There are 411 micro-region observations in the estimations of Panels A and B, except for three to four missing values in government spending. In Panel C, there are 386 micro-region observations for the diagnoses rate and 411 observation for the other regressions. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

In Panel A of Table 6, we describe the impact of the tariff cuts on government spending in the micro-regions (aggregating from municipality-level data). Looking at spending on H&S, it is notable that, while the coefficients are not statistically significant, there appears to be an increase in spending on health services in the regions more exposed to the tariff shock. Looking at the ratio of expenditures on H&S to total expenditures, we observe statistically significant increases both in the medium and in the long run.

Next, in Panel B, we examine the expenditures of both the hospital (SIH - *Sistema de Informações Hospitalares do SUS*) and outpatient (SIA - *Sistema de Informações Ambulatoriais do SUS*) systems of SUS. In the first column, note that expenditures on hospitalizations and hospital procedures increased significantly - economically and statistically - with the tariff shock in the medium and long run, with the latter effect being much larger. Similarly,

²⁷Due to data constraints, we are unable to assess the effects of variables representing potential mechanisms with starting years prior to treatment (pre-treatment). This limitation also prevents the performance of placebo tests for these mechanisms. Detailed explanations of these data limitations for each analyzed category in Table 6 are provided in the Online Appendix.

per capita expenditure in the outpatient system increased both in the medium and long run after the trade liberalization episode. Moreover, we estimate the effect of the trade-induced economic shock on the number of health establishments to more directly assess the expansion of health infrastructure. In the medium run, the average number of establishments does not seem to change in the regions most affected by the tariff cuts. However, the total effect becomes statistically significant in the long run - that is, the local economies more exposed to the shock show a relative increase in the number of health establishments after almost two decades following liberalization.

These findings provide support for the hypothesis that regions most affected by the trade-induced economic shock experienced an expansion in their healthcare infrastructure. While Figure 5 demonstrates a substantial overall increase in health expenditure at the aggregate level for Brazilian municipalities, our analysis reveals that urban and industrial regional economies exhibited a proportionally higher budget allocation for healthcare services.²⁸ This trend manifested in increased expenses within both the hospital and outpatient systems of SUS, as well as a relative rise in the number of healthcare establishments in the micro-regions more exposed to tariff cuts. Building on this evidence, we posit that the observed reduction in mortality rates from various diseases, including cardiovascular ailments, cancer, and other conditions, can be partially attributed to the expansion of healthcare infrastructure in these regions.

An intriguing aspect that warrants exploration is whether the expansion of healthcare infrastructure in the aforementioned regions entailed an augmentation in the complexity of available medical examinations and procedures or their capital-intensity, particularly within the context of the observed substantial reduction in mortality rates from cardiovascular diseases and neoplasms over the long term. To investigate this further, Panel C of Table 6 focuses on assessing the effects of the regional economic shock on the production outcomes within both the hospital and outpatient systems. Initially, we assess the effects of tariff cuts on the hospitalization rate per 100,000 population. Our findings reveal that the increase in the hospitalization rate in the regions most impacted by the trade liberalization shock does not exhibit statistical significance in either the medium or long term. However, the significant increase in expenditures on hospitalizations and procedures suggests a potential rise in expenditure per hospitalization, which could be indicative of the utilization of more advanced and costly methods for diagnostic and treatment purposes. Furthermore, we document significant increases in the ratio of non-basic procedures to basic procedures within the outpatient system, both in the medium and long run. These findings indicate a notable shift towards more advanced and intricate procedures within the outpatient healthcare setting.

Moreover, in an endeavor to establish a connection between the increasing complexity of medical exams and procedures and the reduction in mortality rates from cardiovascular diseases and cancer, in the last two columns of Panel C we investigate the impact of the

²⁸In the Online Appendix, we provide evidence suggesting a correlation between regional tariff changes and per capita government expenditure, particularly in health. Regions with higher relative health spending prior to the trade shock saw the most significant long-term reduction in mortality rates, supporting our hypothesis that post-reform healthcare infrastructure enhancement played a crucial role in mitigating mortality in heavily impacted regions post-trade liberalization.

regional economic shock on the rate of procedures aimed at diagnosing malignant tumors within the hospital system, expressed as the number of procedures per 100,000 inhabitants.²⁹ Our findings provide evidence indicating a substantial increase in procedures for detecting malignant neoplasms in the regions most affected by the tariff cuts. This effect is statistically significant in the medium run and retains a qualitatively similar magnitude in the longer run. It is worth emphasizing that the increase in tumor detection in the medium term may be directly linked to both the reduction in cancer mortality rates a decade following the tariff shock and, notably, the subsequent magnification of this reduction over the long term. The underlying rationale is straightforward: heightened detection of malignant tumors, particularly when associated with early detection, leads to a diminished number of deaths caused by the condition over the ensuing years.

In summary, our investigation into the potential mechanisms underlying the pro-cyclical patterns observed in mortality rates from internal causes documents, in addition to the previously outlined increases in government spending on health infrastructure, a relative rise in spending per hospitalization and hospital procedures and an upsurge in the number of non-basic procedures compared to basic procedures within the outpatient system of the Unified Health System (SUS) in regions that were more significantly impacted by the trade liberalization episode. Furthermore, our findings highlight a substantial increase in procedures aimed at detecting malignant tumors in regional economies exposed to more substantial tariff reductions. Collectively, these results provide empirical support for the hypothesis that the reduction in deaths from internal causes in the more severely affected regions can be partially attributed to the more than proportional expansion of healthcare infrastructure within these micro-regions, in comparison to areas with lower exposure to the trade-induced economic shock.³⁰

5.3 Imports and Access to Diagnostic Machinery

Lastly, we examine the facilitation of imports of diagnostic-related machinery as a potential catalyst for the reduction in mortality rates from internal causes associated with trade liberalization. We investigate how enhanced access to foreign markets might have played a role in improving healthcare infrastructure. As alluded to previously, the years following liberalization witnessed a remarkable upswing in the prevalence of diagnostic imaging equipment across Brazil. This phenomenon prompts consideration of the plausible conjecture that the trade reform, in its essence, actively contributed to the expansion of these resources, both in terms of affordability and availability.

²⁹Regrettably, we encountered challenges in identifying hospital procedures specifically related to the detection of cardiovascular diseases that maintained consistent classification throughout the analysis period in Brazil.

³⁰Additional insights into the mechanisms underlying the impact of trade-induced regional economic shocks on mortality rates, including attenuation regressions and direct impacts of mechanism proxies, are available in the Online Appendix. In short, the findings underscore the importance of expanding capital-intensive healthcare infrastructure, evidenced by regional-level relative health expenditure and diagnosis rates, in driving long-term reductions in mortality rates in heavily affected regions.

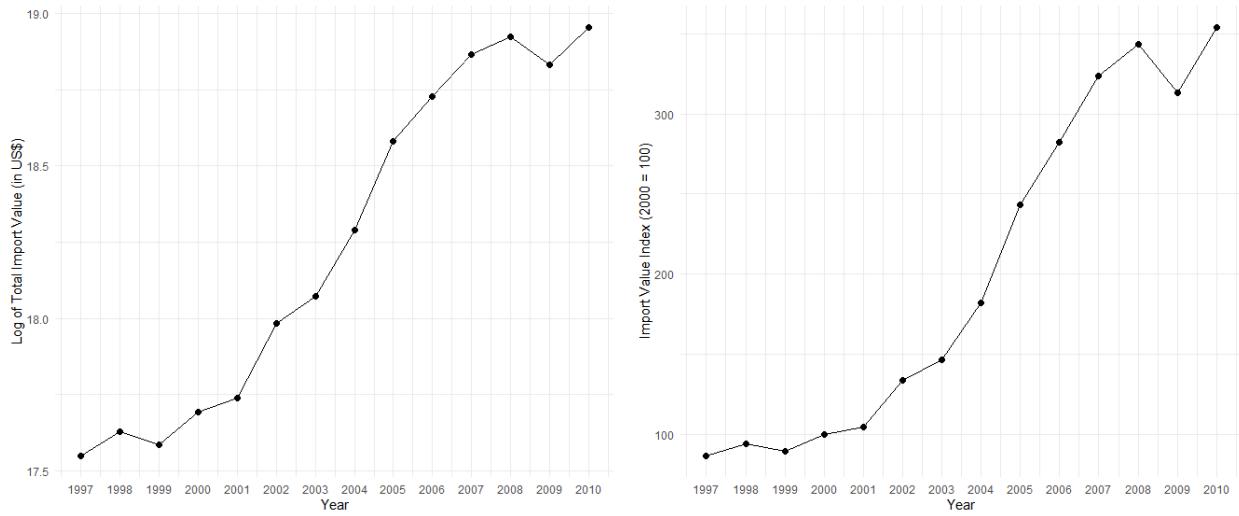
Our objective is to evaluate whether such imported diagnostic machinery influenced the decline in mortality from internal causes in regions experiencing significant tariff reductions. While one might initially expect the relative affordability of imported machinery to be uniform across the country due to the nature of the trade liberalization episode, our prior discussions underscore the disproportionate concentration of machinery expansion in regions more exposed to the trade-induced shock. Furthermore, we documented earlier that these harder-hit regions also showed a relative upswing in healthcare expenditure. Coupled with the higher pre-reform income per capita, these patterns suggest the possibility of import facilitation exerting a more accentuated impact on these particularly affected regions.

To explore this potential mechanism, we rely on detailed import data from the *Sistemas de Comércio Exterior* (SISCOMEX) provided by the Ministry of Industry, Foreign Trade, and Services. In our analysis, we specifically harness aggregated metrics detailing the total import value (measured in US dollars) of selected machinery across individual Brazilian states per annum.³¹ In particular, our focus is directed towards two distinct Harmonized System header codes - 9018 and 9022 - which encompass: i) instruments and appliances utilized in medical and surgical procedures, encompassing items like scintigraphic and other electro-medical apparatus, and ii) apparatuses operating on the basis of X-rays or utilizing alpha, beta, gamma, or other ionizing radiations. These categories effectively encapsulate the selected diagnostic imaging equipment highlighted in Table 5.

Figure 6 sheds light on this mechanism by indicating the pronounced increases in total imports of selected healthcare machinery during the 2000s. Notably, panel (b) displays a growth of nearly 250% in total imports between 2000 and 2010. These discernible trends suggest that, indeed, the trade liberalization episode played a role in facilitating the influx of specific (“life-saving”) machinery imports. Consequently, this exerted a direct influence on the availability of diagnostic imaging tools, a pivotal resource for the identification of cardiovascular ailments and neoplasms.

Importantly, around 70% of these imports were concentrated in the Southeast region of Brazil, where the state of *São Paulo* singularly contributed to 55% of the nation’s total imports within the span of ten years. This geographic concentration closely aligns with the exposure to tariff reductions during the trade liberalization episode. While the increased affordability and accessibility of machinery could theoretically have permeated broadly across the nation due to the trade reform, the marked accumulation of such equipment within the regions more exposed to the tariff cuts indicates that this mechanism actively contributed to the discernible reduction in mortality rates arising from internal causes.

³¹It is important to note that while there exists available data regarding imports at the municipality level, it is highly improbable that such finely disaggregated measures accurately represent the actual locations of machinery utilization. This is due to the fact that such goods are distributed among various micro-regions of the country after the import process.



(a) Total imports of selected machinery: Log value

(b) Total imports of selected machinery: Index

Figure 6: Trends in total imports of machinery related to diagnostic: 1997-2010

Source: Brazilian foreign trade data based on the declaration of importers (SISCOMEX from the Ministry of Industry, Foreign Trade and Services).

6 Discussion

This paper draws inspiration from the influential work of [Ruhm \(2000\)](#) and subsequent literature, investigating the effects of regional economic shocks on mortality rates in Brazil using the country's trade liberalization as a natural experiment. By providing causal evidence of the impact of trade-induced economic shocks on health outcomes, we address identification challenges commonly encountered in the literature that has predominantly focused on developed countries with pro-cyclical mortality patterns. Unlike prior studies that often report ambiguous or counter-cyclical effects in developing countries, our research demonstrates consistent pro-cyclical impacts within a major developing economy. We reveal that mortality rates not only fluctuate in the short term but also adjust significantly over the long term following trade liberalization. Furthermore, our work contributes to the growing body of literature examining the broader effects of economic shocks resulting from trade policy changes on adult health outcomes, offering new insights into how such policies shape health dynamics in relatively underexplored developing country contexts.

Our findings indicate a clear pro-cyclical pattern in mortality during the analyzed periods, with regions more exposed to trade-induced economic shocks having higher reductions in mortality rates in both the medium and long run. Specifically, our baseline specification reveals that a region facing a tariff reduction reflecting a shift from the tenth to the ninetieth percentile of regional tariff changes witnessed a consequential reduction of over 10 percent in all-cause mortality rates five years post-reform and a remarkable 18 percent reduction fifteen years thereafter. These results align with previous research by [Dix-Carneiro and Kovak](#)

(2017) on employment dynamics within these local economies.

Additionally, our investigation addressed concerns regarding changes in the age structure of the population by showing that the decline in mortality rates among regions exposed to the trade shock is observed across different age groups and specific causes of death. A falsification test confirms the robustness of the findings, as it demonstrates that cause-specific mortality rates unrelated to macroeconomic conditions did not show significant changes with tariff cuts. These results collectively reinforce the causal link between deteriorating local labor markets and declining mortality rates. Moreover, they substantiate our use of the Brazilian trade liberalization as a quasi-natural experiment, providing robust support for our identification strategy.

Central to our analysis was the exploration of mechanisms through which the trade-induced shock impacted mortality rates. While it is challenging to attribute the observed changes in mortality to a single explanatory factor, our study identifies several key channels influencing these outcomes. Some of these mechanisms corroborate findings from existing literature, while others present novel insights into how economic shocks affect health outcomes. Concerning external causes of mortality, our results are similar to Ruhm (2000) and Miller et al. (2009), pointing to the pro-cyclical nature of transport accidents and other accidents since individuals use relatively fewer means of transport due to higher unemployment as well as practice less hazardous activities. This mechanism appears to persist both in the medium and long term of our analysis.

Importantly, we argue that the expansion of healthcare infrastructure played a fundamental role in reducing mortality rates from internal causes in the regions more exposed to the trade-induced economic shock in the longer run. The enactment of Brazil's federal Constitution in 1988 brought significant social spending increases and the establishment of the *Sistema Único de Saúde*, the world's largest publicly-funded healthcare system. The expansion of health-related government spending and decentralization of health management were observed across different regions, although the availability of medium and high-complexity exams and procedures, particularly for cardiovascular diseases and neoplasms, saw a pronounced growth in more affluent localities, coinciding geographically with the magnitude of the regional tariff shock.

We then investigated the direct impacts of the trade-induced economic shock on healthcare infrastructure indicators in the country's local economies. Our findings suggest that regions most affected by the tariff reductions experienced significant increases in the share of government spending on health and sanitation and in the number of healthcare establishments compared to regions less impacted by tariff cuts. Moreover, we find significant increases in hospitalization and outpatient expenditures, as well as potential rises in the expenditure per hospitalization and a shift towards more complex procedures within the outpatient system in these harder-hit regions. We also document a substantial increase in procedures for detecting malignant tumors in the micro-regions more impacted by the shock.

Lastly, we explore the potential role of trade reform in directly fostering the expansion of specific diagnostic imaging machinery through imports and cost reduction. Our analysis reveals a concentrated expansion of total imports of these "life-saving" or "death-minimizing"

goods in regions more significantly affected by the trade-induced shock, suggesting that import-driven expansion of capital-intensive healthcare infrastructure played a role in the reduction of mortality rates stemming from internal causes.

Overall, our in-depth exploration of the mechanisms underlying pro-cyclical mortality yields novel insights that advance the existing body of literature in health economics. By delineating the intricate ways in which economic shocks influence health outcomes, particularly in the context of a developing economy, our study contributes to a deeper understanding of the broad impacts of trade liberalization on public health.

In summary, these findings underscore the pivotal role of healthcare infrastructure in ameliorating the ramifications of economic shocks on mortality rates. Interestingly, the existence of a universal public-funded healthcare system in Brazil since the late 1980s and its expansion post-liberalization appears to mitigate fluctuations in access to capital-intensive medical care during economic downturns. The implications of these insights are profound. Policymakers should prioritize healthcare infrastructure investment, especially in vulnerable regions, to enhance health outcomes and curtail mortality rates. Furthermore, efforts aimed at enhancing healthcare management decentralization emerge as enablers of healthcare infrastructure expansion. Importantly, policymakers should account for potential health repercussions arising from trade policy shifts and adopt measures ensuring healthcare access.

By way of conclusion, it is important to highlight that there is still much to explore to fully comprehend the intricate connections between macroeconomic conditions and mortality rates in Brazil. However, this study is a relevant step in this research agenda. It not only contributes to the health economics literature by shedding light on the pro-cyclical nature of mortality but also enhances our understanding of the impacts of trade liberalization experiences worldwide. While it is important to note that our analysis captures only a partial equilibrium effect of the Brazilian trade liberalization episode, as previously pursued, for instance, in Dix-Carneiro et al. (2018) and Charris et al. (2023), our empirical findings provide valuable evidence that expands upon the existing literature on developing countries and can indicate an intriguing avenue for future research.

References

- R. Adao, M. Kolesár, and E. Morales. Shift-share designs: Theory and inference. *The Quarterly Journal of Economics*, 134(4):1949–2010, 2019.
- L. Alvarez, B. Ferman, and R. Oliveira. Randomization inference tests for shift-share designs. Technical report, arXiv:2206.00999, 2022. URL <https://arxiv.org/pdf/2206.00999.pdf>.
- I. Arroyave, P. Hessel, A. Burdorf, J. Rodriguez-Garcia, D. Cardona, and M. Avendaño. The public health impact of economic fluctuations in a latin american country: mortality and the business cycle in colombia in the period 1980–2010. *International Journal for Equity in Health*, 14(1):1–14, 2015.

- D. Autor, D. Dorn, and G. Hanson. When work disappears: Manufacturing decline and the falling marriage market value of young men. *American Economic Review: Insights*, 1(2):161–178, 2019.
- J. Ballester, J.-M. Robine, F. R. Herrmann, and X. Rodó. Effect of the great recession on regional mortality trends in europe. *Nature Communications*, 10(1):679, 2019.
- T. J. Bartik. *Who benefits from state and local economic development policies?* WE Upjohn Institute for Employment Research, 1991.
- K. Borusyak and P. Hull. Non-random exposure to exogenous shocks. Technical report, National Bureau of Economic Research, 2023.
- K. Borusyak, P. Hull, and X. Jaravel. Quasi-experimental shift-share research designs. *The Review of Economic Studies*, 89(1):181–213, 2022.
- K. Borusyak, X. Jaravel, and J. Spiess. Revisiting event study designs: Robust and efficient estimation. *Review of Economic Studies*, Forthcoming, 2024.
- B. Callaway and P. H. C. Sant'Anna. Difference-in-differences with multiple time periods. *Journal of Econometrics*, 225(2):200–230, 2021
- J. Cawley and C. J. Ruhm. The economics of risky health behaviors. In *Handbook of health economics*, volume 2, pages 95–199. Elsevier, 2011.
- C. Charris, D. Branco, and B. Carrillo. Economic shocks and infant health: Evidence from a trade reform in brazil. *Journal of Development Economics*, 166, 2023.
- C. De Chaisemartin and X. d'Haultfoeuille. Two-way fixed effects estimators with heterogeneous treatment effects. *American Economic Review*, 110(9):2964–2996, 2020.
- C. De Chaisemartin and X. d'Haultfoeuille. Two-way fixed effects and differences-in-differences with heterogeneous treatment effects: A survey. *National Bureau of Economic Research*, 2022.
- R. Dix-Carneiro. Trade liberalization and labor market dynamics. *Econometrica*, 82(3):825–885, 2014.
- R. Dix-Carneiro and B. K. Kovak. Trade liberalization and the skill premium: A local labor markets approach. *American Economic Review*, 105(5):551–57, 2015.
- R. Dix-Carneiro and B. K. Kovak. Trade liberalization and regional dynamics. *American Economic Review*, 107(10):2908–46, 2017.
- R. Dix-Carneiro, R. R. Soares, and G. Ulyssea. Economic shocks and crime: Evidence from the brazilian trade liberalization. *American Economic Journal: Applied Economics*, 10(4):158–95, 2018.

- H. Fan, F. Lin, and S. Lin. The hidden cost of trade liberalization: Input tariff shocks and worker health in China. *Journal of International Economics*, 126:103349, 2020. ISSN 0022-1996.
- J. Feng, Q. Xie, and X. Zhang. Trade liberalization and the health of working-age adults: Evidence from China. *World Development*, 139:105344, 2021. ISSN 0305-750X.
- B. Ferman. Assessing inference methods. Technical report, arXiv:1912.08772, 2022. URL <https://arxiv.org/pdf/1912.08772.pdf>.
- I. Gaddis and J. Pieters. The gendered labor market impacts of trade liberalization: Evidence from Brazil. *Journal of Human Resources*, 52(2):457–490, 2017.
- P. K. Goldberg and N. Pavcnik. Distributional effects of globalization in developing countries. *Journal of Economic Literature*, 45(1):39–82, 2007.
- P. Goldsmith-Pinkham, I. Sorkin, and H. Swift. Bartik instruments: What, when, why, and how. *American Economic Review*, 110(8):2586–2624, 08 2020. ISSN 0002-8282.
- F. Gonzalez and T. Quast. Mortality and business cycles by level of development: evidence from mexico. *Social Science & Medicine*, 71(12):2066–2073, 2010.
- A. Goodman-Bacon. Difference-in-differences with variation in treatment timing. *Journal of Econometrics*, 225(2):254–277, 2021
- V. F. Haaland and K. Telle. Pro-cyclical mortality across socioeconomic groups and health status. *Journal of Health Economics*, 39:248–258, 2015.
- G. Hirata and R. R. Soares. Competition and the racial wage gap: Evidence from brazil. *Journal of Development Economics*, 146:102519, 2020.
- T. Hone, A. J. Mirelman, D. Rasella, R. Paes-Sousa, M. L. Barreto, R. Rocha, and C. Millett. Effect of economic recession and impact of health and social protection expenditures on adult mortality: a longitudinal analysis of 5565 brazilian municipalities. *The Lancet Global Health*, 7(11):e1575–e1583, 2019.
- IBGE. *Indicadores sociodemográficos e de saúde no Brasil: 2009*. Instituto Brasileiro de Geografia e Estatística - Coordenação de População e Indicadores Sociais, 2009.
- B. K. Kovak. Regional effects of trade reform: What is the correct measure of liberalization? *American Economic Review*, 103(5):1960–76, 2013.
- H. Kume, G. Piani, and C. F. B. Souza. A política brasileira de importação no período 1987-98: Descrição e avaliação. In K. Corseuil and H. Kume, editors, *A Abertura Comercial Brasileira nos Anos 1990: Impactos sobre Emprego e Salário*. 2003.
- M. Lang, T. C. McManus, and G. Schaur. The effects of import competition on health in the local economy. *Health Economics*, 28(1):44–56, 2019.

- J. M. Lindo. Aggregation and the estimated effects of economic conditions on health. *Journal of Health Economics*, 40:83–96, 2015.
- T. C. McManus and G. Schaur. The effects of import competition on worker health. *Journal of International Economics*, 102:160–172, 2016.
- D. L. Miller, M. E. Page, A. H. Stevens, and M. Filipski. Why are recessions good for your health? *American Economic Review*, 99(2):122–27, 2009.
- P. M. Ogeda, E. Ornelas, and R. R. Soares. Labor unions and the electoral consequences of trade liberalization. *Journal of the European Economic Association*, jvae020, 2024.
- J. Paim, C. Travassos, C. Almeida, L. Bahia, and J. Macinko. The Brazilian health system: history, advances, and challenges. *The Lancet*, 377(9779):1778–1797, 2011.
- J. R. Pierce and P. K. Schott. Trade liberalization and mortality: evidence from us counties. *American Economic Review: Insights*, 2(1):47–64, 2020.
- V. Ponczek and G. Ulyssea. Enforcement of labour regulation and the labour market effects of trade: Evidence from brazil. *The Economic Journal*, 132(641):361–390, 2022.
- J. Roth, P. H. C. Sant’Anna, A. Bilinski, and J. Poe. What’s trending in difference-in-differences? A synthesis of the recent econometrics literature. *Journal of Econometrics*, 235(2), 2023.
- C. Ruhm. Are recessions good for your health? *The Quarterly Journal of Economics*, 115(2):617–650, 2000.
- C. Ruhm. Good times make you sick. *Journal of Health Economics*, 22(4):637–658, 2003.
- C. Ruhm. Healthy living in hard times. *Journal of Health Economics*, 24(2):341–363, 2005.
- C. Ruhm. Understanding the relationship between macroeconomic conditions and health. In *The Elgar Companion to Health Economics, Second Edition*, chapter 1. Edward Elgar Publishing, 2012.
- C. Ruhm. Recessions, healthy no more? *Journal of Health Economics*, 42:17–28, 2015.
- A. H. Stevens, D. L. Miller, M. E. Page, and M. Filipski. The best of times, the worst of times: understanding pro-cyclical mortality. *American Economic Journal: Economic Policy*, 7(4):279–311, 2015.
- G. J. van den Berg, U.-G. Gerdtham, S. von Hinke, M. Lindeboom, J. Lissdaniels, J. Sundquist, and K. Sundquist. Mortality and the business cycle: Evidence from individual and aggregated data. *Journal of Health Economics*, 56:61–70, 2017. ISSN 0167-6296.

A Appendix

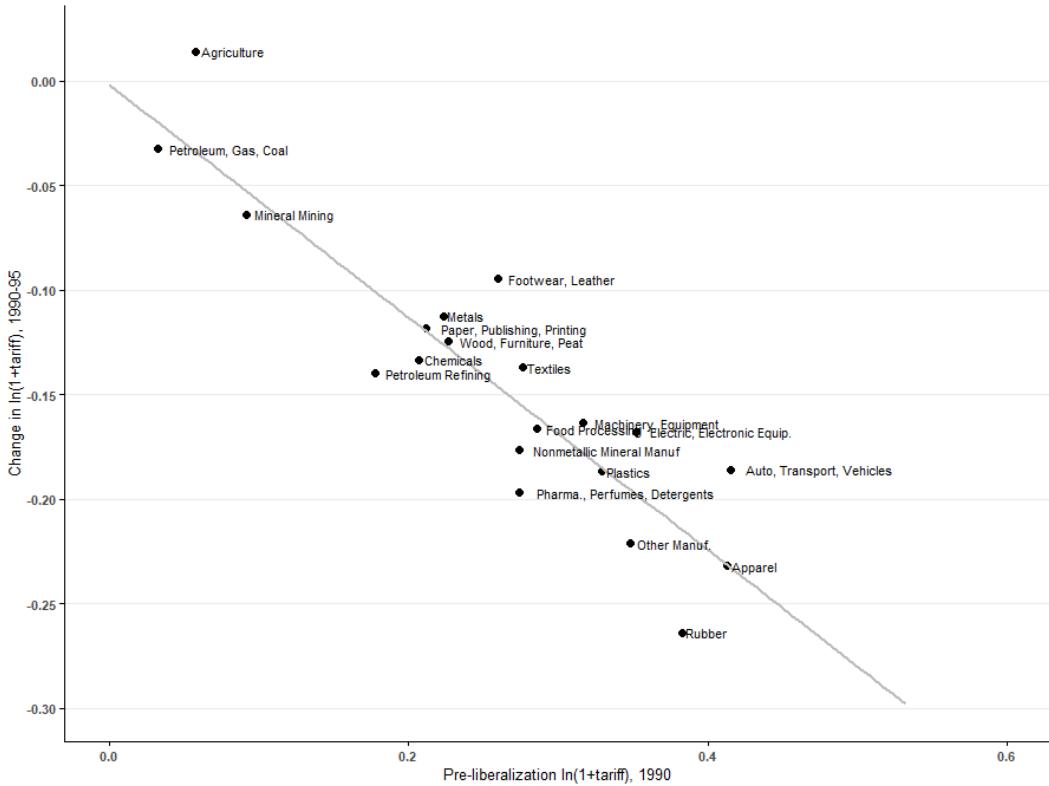


Figure A.1: Relationship between tariff changes and pre-liberalization tariff Levels

Table A.1: Definition of specific causes of mortality

Variables	Description	ICD-9 (1985-1995)	ICD-10 (1996-2010)
General	General mortality: all causes	001-E999	A00-Y99
Infant	General mortality of infants (less than 1 year old)	001-E999	A00-Y99
Endocrine	Endocrine, nutritional and metabolic diseases	240-279	E00-E89
CVD	Major cardiovascular diseases	390-448	I00-I78
Cancer	Malignant neoplasms	140-208	C00-C97
Transport	Transport accidents	800-848, 929.0, 929.1	V02-V99, Y85
Other Ac	Other (non-transport) accidents	850-928, 929.2-949	W00-X59, Y86
Medical	Misadventures to patients during surgical and medical care ^a	870-879	Y62-Y69, Y83-Y84
Falls	Accidents: falls	880-888	W00-W19
Drowning	Accidents: drowning/submersion	910	W65-W74
Fires	Accidents: smoke/fire/flames	890-899	X00-X09
Poison	Accidents: poisoning/noxious substances	850-869, 924.1	X40-X49
Suicide	Suicide (intentional self-harm)	950-959	X60-X84, Y87.0
Homicide	Homicide and legal intervention	960-978	X85-Y09, Y87.1, Y35, Y89.0
Nature	Accidents due to natural and environmental factors ^b	900-909	T63, X30-X39

^a Including surgical and other medical procedures as the cause of abnormal reaction of the patient, or of later complication, without mention of misadventure at the time of the procedure.

^b Including contact with venomous animals and plants to correctly crosswalk between ICD-9 and ICD-10.

Table A.2: Descriptive statistics at the micro-region level

Variables	Source	Mean	SD	Min	Max	Observations
Overall Mortality Rates						
Mortality rate (per 100,000 inhabitants) - 1985	DataSUS	556.13	183.79	17.46	1304.80	411
Mortality rate (per 100,000 inhabitants) - 1991	DataSUS	497.81	170.37	44.22	1014.97	411
Mortality rate (per 100,000 inhabitants) - 2000	DataSUS	517.06	140.22	43.85	856.56	411
Mortality rate (per 100,000 inhabitants) - 2010	DataSUS	586.71	127.53	148.87	889.09	411
Expenditure per Function - Municipality						
Expenditure - 1995 (annual, 2010 R\$)	Ministry of Finance	507.91	259.19	0.00	1899.63	411
Expenditure - 2000 (annual, 2010 R\$)	Ministry of Finance	930.42	2248.65	0.00	45909.62	411
Expenditure - 2010 (annual, 2010 R\$)	Ministry of Finance	1112.01	326.54	0.00	3498.28	411
Expenditure H&S - 1995 (annual, 2010 R\$)	Ministry of Finance	75.15	49.58	0.00	405.70	411
Expenditure H&S - 2000 (annual, 2010 R\$)	Ministry of Finance	176.02	266.53	0.00	5279.31	411
Expenditure H&S - 2010 (annual, 2010 R\$)	Ministry of Finance	393.13	146.88	0.00	1605.14	411
Number of Health Establishments						
Total - 1992	AMS - IBGE	40.07	14.25	6.17	103.73	411
Total - 1999	AMS - IBGE	41.97	15.14	9.96	159.14	411
Total - 2006	CNES - DataSUS	93.42	44.36	15.47	283.35	411
Total - 2010	CNES - DataSUS	118.99	55.56	20.97	327.43	411
Hospitalization Rates and Hospital Procedures						
Hospitalization Rate - 1992	SIH - DataSUS	10113.53	5080.35	1.91	29329.57	411
Hospitalization Rate - 2000	SIH - DataSUS	6846.78	2552.50	1.70	25883.81	411
Hospitalization Rate - 2010	SIH - DataSUS	5497.35	2245.28	9.66	18908.80	411
Neoplasm detection - 1995	SIH - DataSUS	2.12	3.37	0.08	51.38	386
Neoplasm detection - 2000	SIH - DataSUS	4.41	6.79	0.28	88.40	386
Neoplasm detection - 2005	SIH - DataSUS	4.14	5.39	0.11	37.80	386
Outpatient Procedures						
Basic - 1995	SIA - DataSUS	660572	338365	41097	1938314	411
Basic - 2000	SIA - DataSUS	681947	231265	13116	1965240	411
Basic - 2007	SIA - DataSUS	844912	257051	45816	1712669	411
Non-basic - 1995	SIA - DataSUS	106937	82667	1	542790	411
Non-basic - 2000	SIA - DataSUS	233076	170008	241	1484847	411
Non-basic - 2007	SIA - DataSUS	479284	441460	42211	3757104	411
Expenditure per capita - SUS systems						
SIH - 1995	SIH - DataSUS	6.91	8.13	0.00	56.60	411
SIH - 2000	SIH - DataSUS	9.57	9.81	0.00	63.22	411
SIH - 2007	SIH - DataSUS	39.92	32.32	0.04	267.88	411
SIA - 1995	SIA - DataSUS	39.10	24.26	1.50	180.14	411
SIA - 2000	SIA - DataSUS	44.65	31.44	0.04	305.45	411
SIA - 2007	SIA - DataSUS	41.79	45.05	1.70	610.60	411

Online Appendix

For Review

“Trade Liberalization and Mortality Rates: Evidence of Pro-Cyclical Mortality from Brazil”

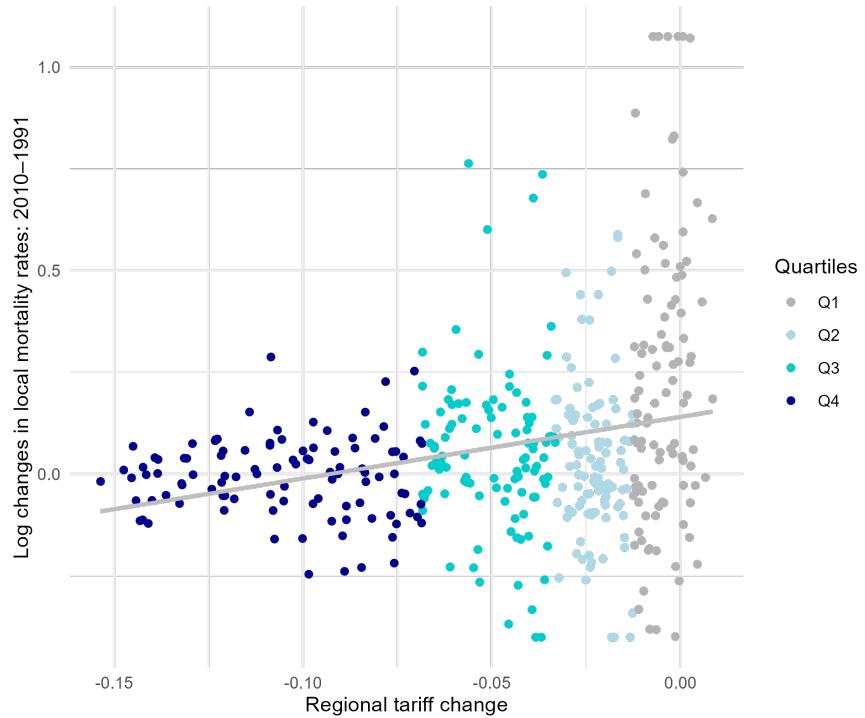
A Additional Visualizations

Initially, Figure 1 depicts the correlation between regional tariff change and log changes in local mortality rates for both the medium and long term. In this representation, the upper quartile, Q4, signifies greater exposure to the regional economic shock and consequently a more negative regional tariff change. Gradually lighter shades of blue to gray denote, in sequence, lower quartiles of exposure to the trade-induced economic shock. Panel 1 illustrates the correlation between the change in log mortality rates between 1991 and 2000 (medium run in the manuscript), while Panel 2 presents the long-run difference between the log mortality rates between 1991 and 2010. In both temporal contexts, there are clear positive correlations between the variables. Therefore, supplementing the evidence from the maps, it is evident that regions most economically affected by the trade shock also witnessed a reduction in log mortality rates.

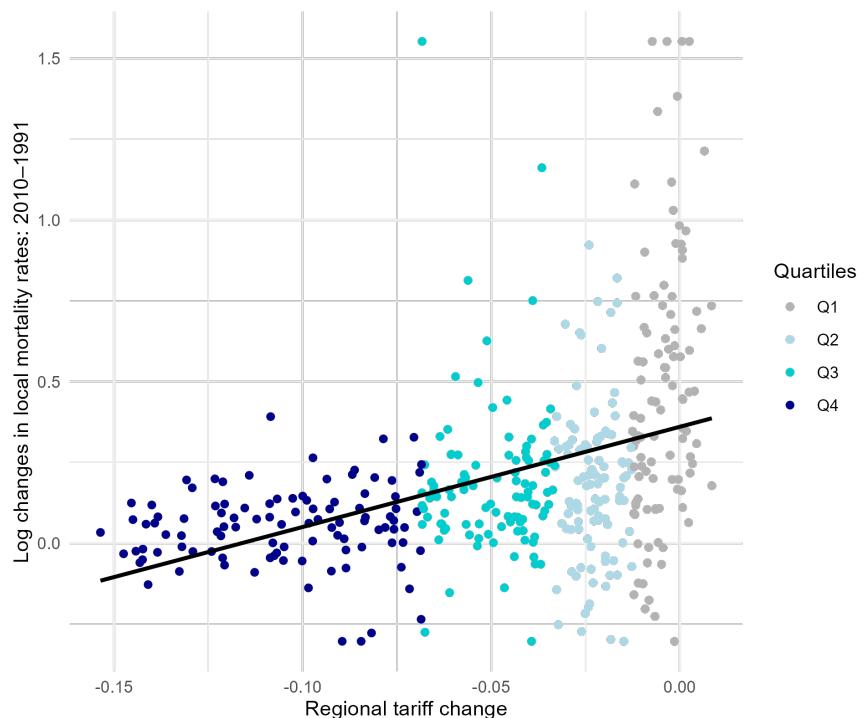
It is important to note that we applied winsorization to the changes in log mortality rates in Figure 1 to avoid visual clutter caused by outliers potentially driving the trend. Specifically, we set the lower and upper outliers to the values associated with the 1st and 99th percentiles of the variable distribution, respectively.

We also included the same visualizations without the winsorization process. Figure 2 illustrates the correlation for both the medium and long run separately, while Figures 3 and 4 visually depicts changes in mortality rates over time per exposure to the trade-induced regional economic shock with and without winsorization, respectively.

Figure 1: Log changes in local mortality rates and regional tariff change - correlation

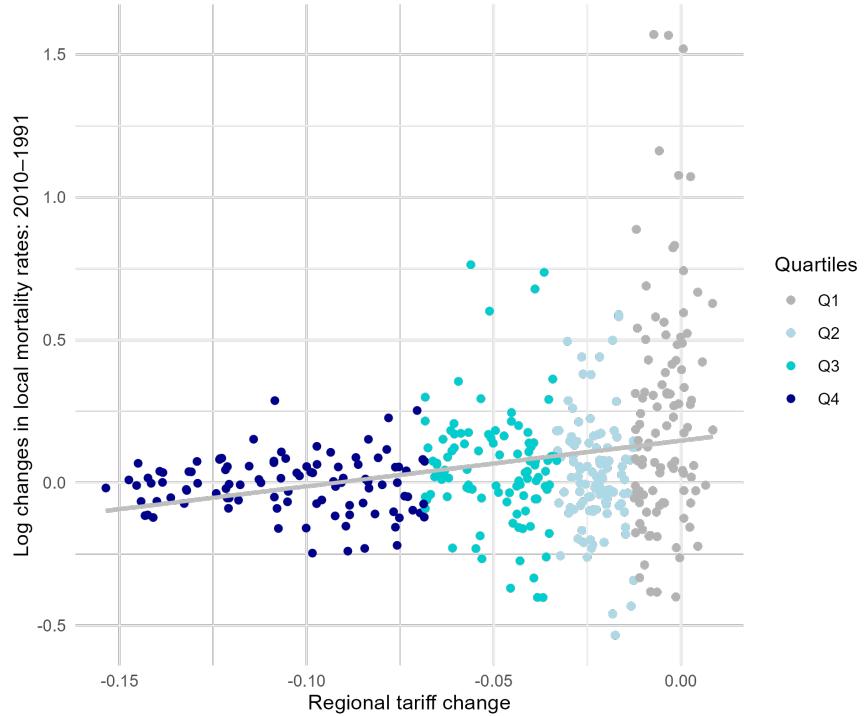


(a) Medium-run: 1991-2000

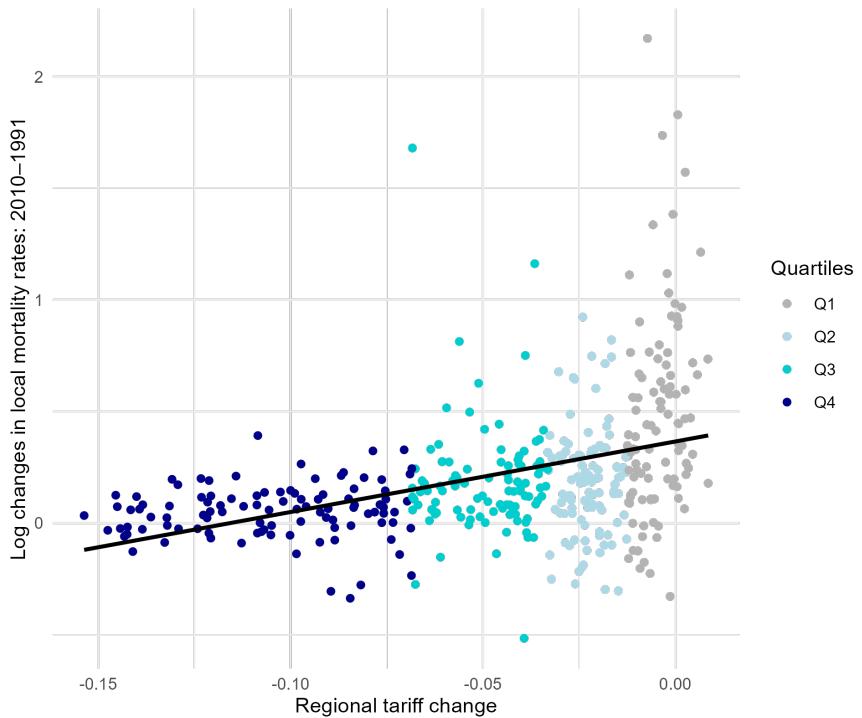


(b) Long-run: 1991-2010

Figure 2: Log changes in local mortality rates (not winsorized)



(a) Medium-run: 1991-2000



(b) Long-run: 1991-2010

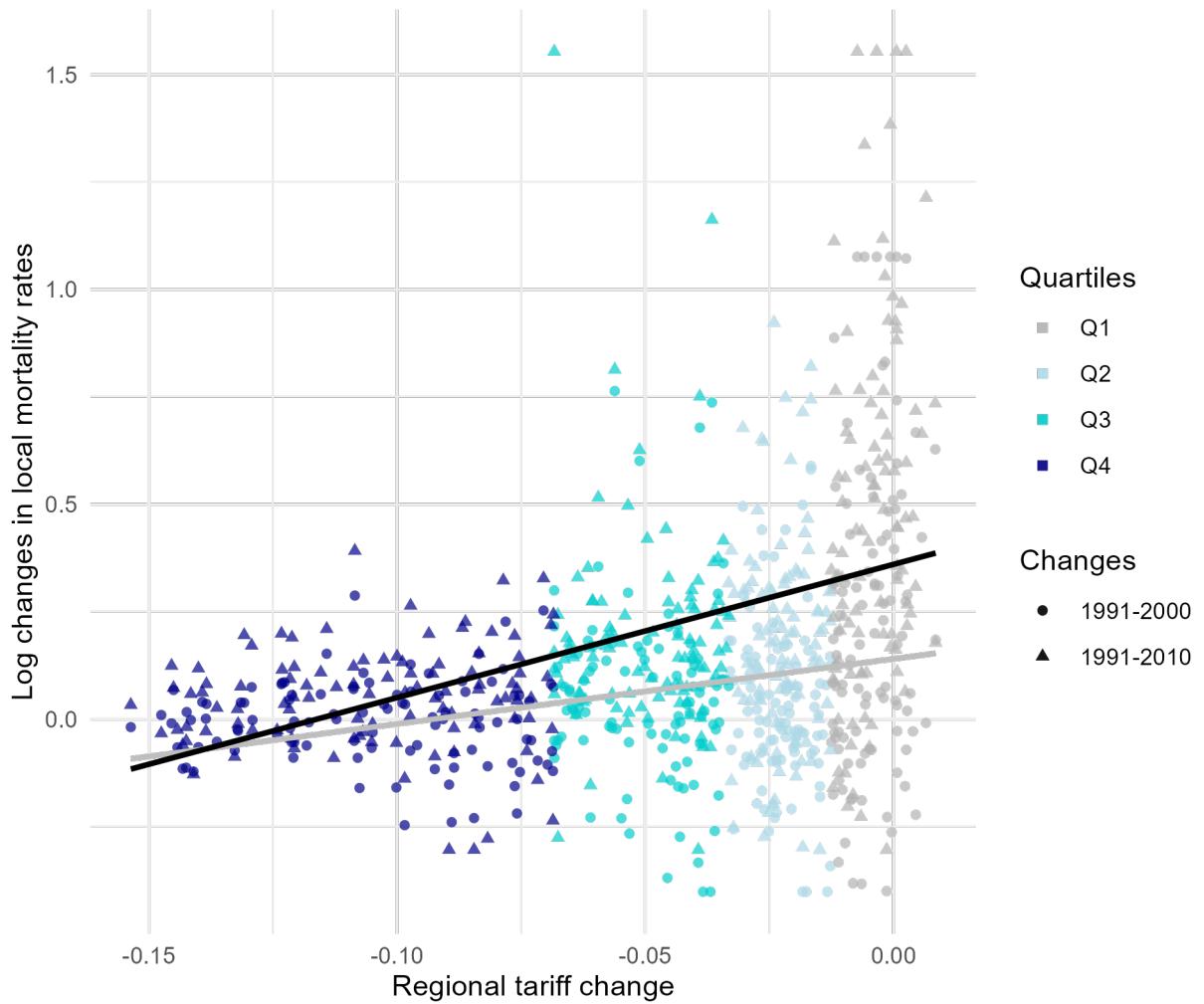


Figure 3: Log changes in local mortality rates per exposure to regional tariff changes (winsorized)

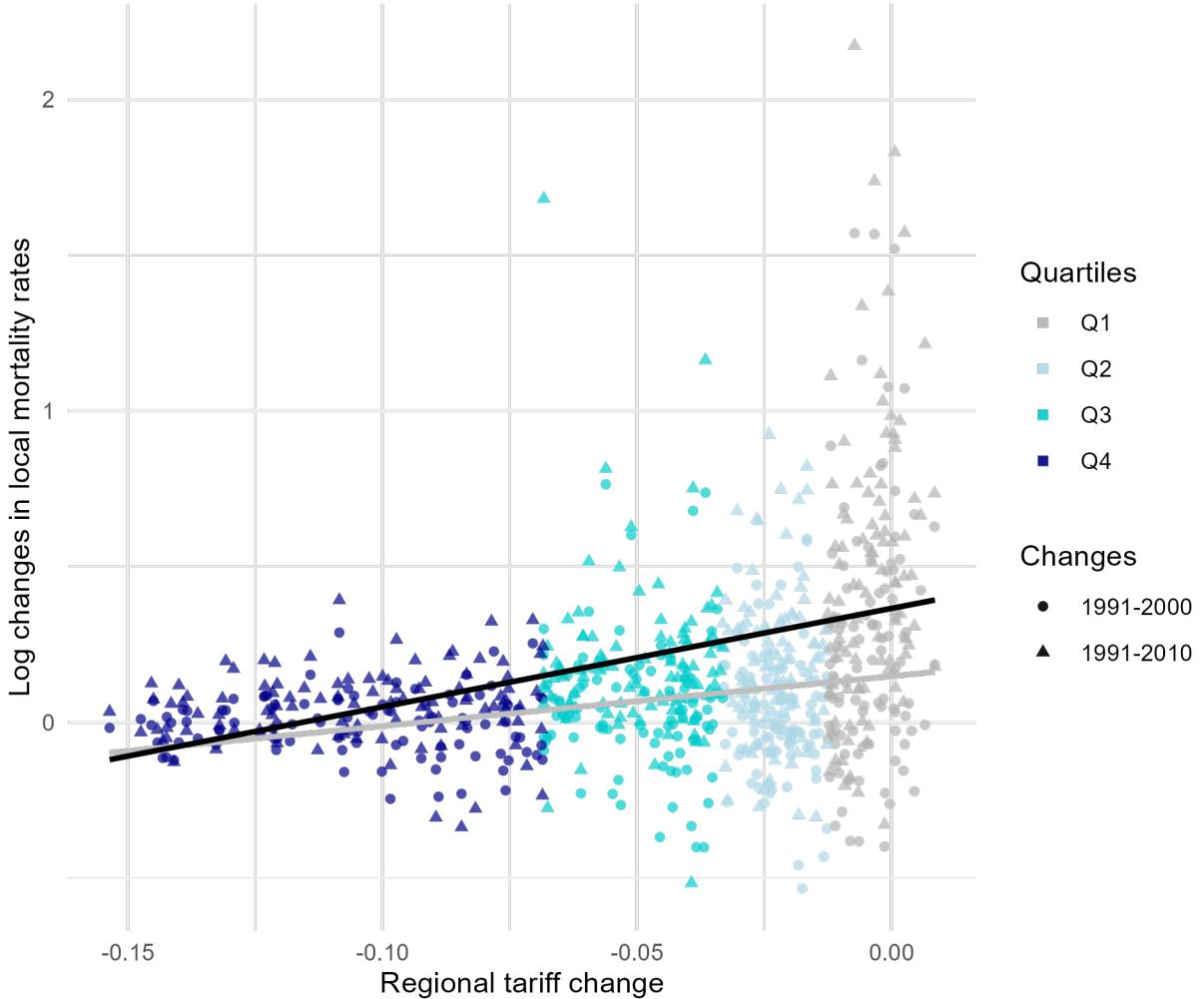


Figure 4: Log changes in local mortality rates per exposure to regional tariff changes (not winsorized)

B Heterogeneity of Baseline Results by Sex

Table 1 presents estimates from our main specification, describing effects for all individuals in the first row and then disaggregating for males and females separately.

Besides, Figure 5 presents a graphical representation of the dynamic effects of tariff reductions on all-cause mortality rates for all individuals and the disaggregations by males and females.

Table 1: Regional tariff changes and log changes in mortality rates

	OLS (1)		OLS (2)		OLS (3)		OLS (4)		2SLS (5)		Placebo (6)
Dep. var.: $\Delta \log(MR_r)$	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1985-1991
All											
RTC_r	1.599*** (0.449)	3.162*** (0.567)	1.164*** (0.275)	2.496*** (0.346)	1.099*** (0.240)	1.957*** (0.278)	1.099*** (0.276)	1.958*** (0.310)	1.099*** (0.249)	1.958*** (0.296)	-0.00567 (0.312)
$\Delta_{85-91} \log(MR_r)$							-0.188 (0.162)	-0.256 (0.171)	-0.104 (0.194)	-0.239 (0.211)	
Male											
RTC_r	1.426*** (0.472)	3.368*** (0.549)	1.059*** (0.298)	2.857*** (0.327)	1.095*** (0.246)	2.103*** (0.245)	1.051*** (0.292)	2.051*** (0.287)	1.075*** (0.259)	2.062*** (0.20)	-0.229 (0.321)
$\Delta_{85-91} \log(MR_r)$							-0.199 (0.161)	-0.241 (0.168)	-0.0867 (0.185)	-0.189 (0.207)	
Female											
RTC_r	1.873*** (0.479)	2.876*** (0.645)	1.340*** (0.338)	2.053*** (0.497)	1.141** (0.276)	1.788*** (0.368)	1.197*** (0.300)	1.874*** (0.396)	1.172*** (0.272)	1.868*** (0.383)	0.306 (0.308)
$\Delta_{85-91} \log(MR_r)$							-0.181 (0.141)	-0.270* (0.148)	-0.0989 (0.180)	-0.250 (0.180)	

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. In column 1, observations are not weighted; in column 2, observations are weighted by population; column 3 adds state fixed effects to column 2; column 4 adds pre-trends to column 3; column 5 shows two-stage least squares, with an instrument for $\Delta_{85-91} \log(MR_r)$. Column 6 presents a placebo test, with observations weighted by population and considering state fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

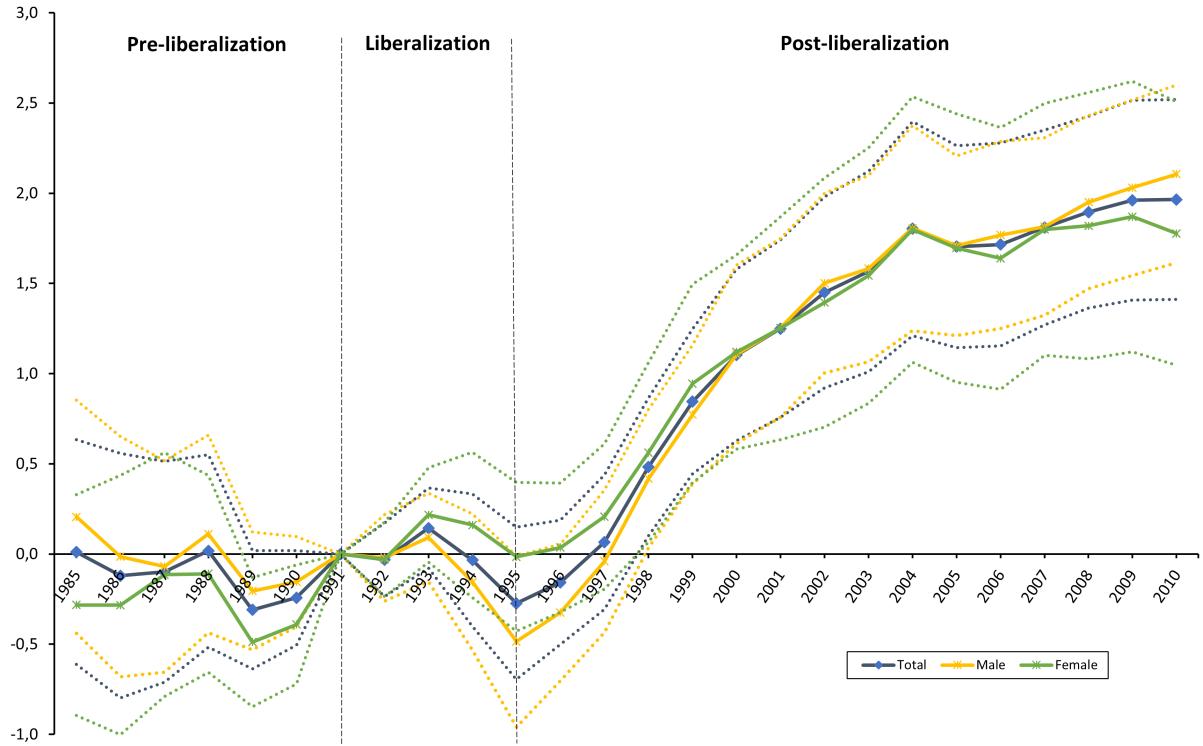


Figure 5: Dynamic effects of regional tariff changes on log changes in mortality rates

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (2) in the article, where the dependent variable is the change in regional log all-cause mortality rates - total, male and female - and the independent variable is the \hat{RTC}_r in $t = 1985, \dots, 2010$. All regressions include state fixed effects. Dashed lines show 95 percent confidence intervals. Standard errors are adjusted for 91 meso-region clusters.

C Robustness to Alternative Inference Procedures

In this section, we show that our baseline results are very similar if we use the inference procedures proposed by Borusyak et al. (2022), which address cross-region correlation in residuals in shift-share designs. Table 2 indicates that our baseline results, presented in the first four columns of Table 1 in the manuscript, are not altered when following alternative inference procedures.

Table 2: Regional tariff changes and log changes in mortality rates (Borusyak et al. (2022) robust standard errors)

	Unweighted		Observations weighted by population		+ State fixed effects		+ Pre-trend	
	(1)	(2)	(3)	(4)				
Dep. var: $\Delta \log(MR_i)$	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
All								
RTC_i	1.593*** [0.0877]	3.128*** [0.153]	1.155*** [0.0325]	2.525*** [0.0539]	1.079*** [0.0444]	1.952*** [0.0460]	1.078*** [0.0464]	1.952*** [0.0389]
Male								
RTC_i	1.430*** [0.0868]	3.359*** [0.163]	1.058*** [0.0297]	2.930*** [0.0682]	1.082*** [0.0476]	2.117*** [0.0568]	1.036*** [0.0479]	2.062*** [0.0463]
Female								
RTC_i	1.844*** [0.0956]	2.805*** [0.148]	1.313*** [0.0424]	2.014*** [0.0689]	1.109*** [0.0462]	1.758*** [0.0530]	1.165*** [0.0499]	1.844*** [0.0529]

Notes: This table presents an alternative approach to inference on the baseline results from columns 1 to 4 in Table 1 of the manuscript. There are 20 industry observations in each regression (industry-level regressions). Borusyak et al. (2022) robust standard errors are reported in brackets. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

This suggests that, in practice, much of the relevant cross-regional correlation in residuals occurs within states. It is worth mentioning that this robustness of the results to potential cross-region correlation in the residuals is also observed in Charris et al. (2023).

D Sensitivity Analysis

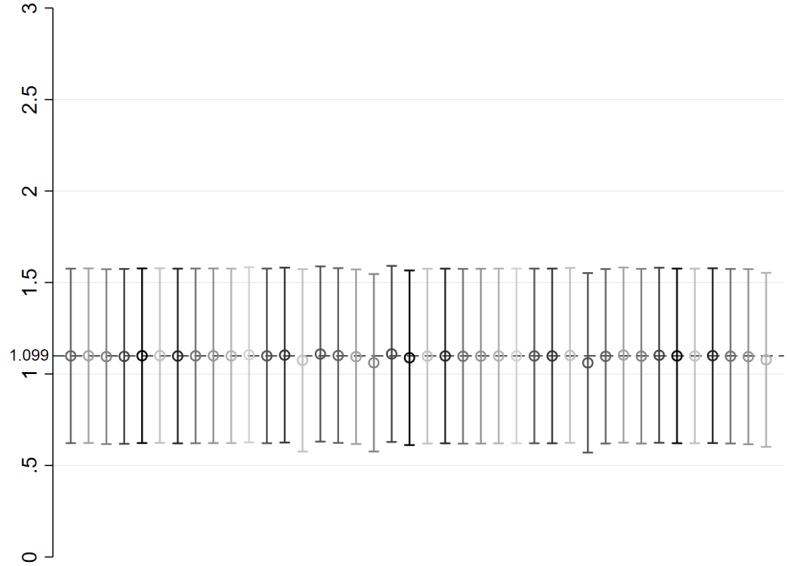
To further evaluate the robustness of our baseline results, presented in Table 1 of the manuscript, we perform a simple sensitivity analysis for the coefficients associated with the impacts of the regional tariff change on the all-cause mortality rates.

First, we estimated the preferred specification - weighted observations and with the inclusion of fixed effects - by removing, in each regression, one of the top 40 micro-regions in

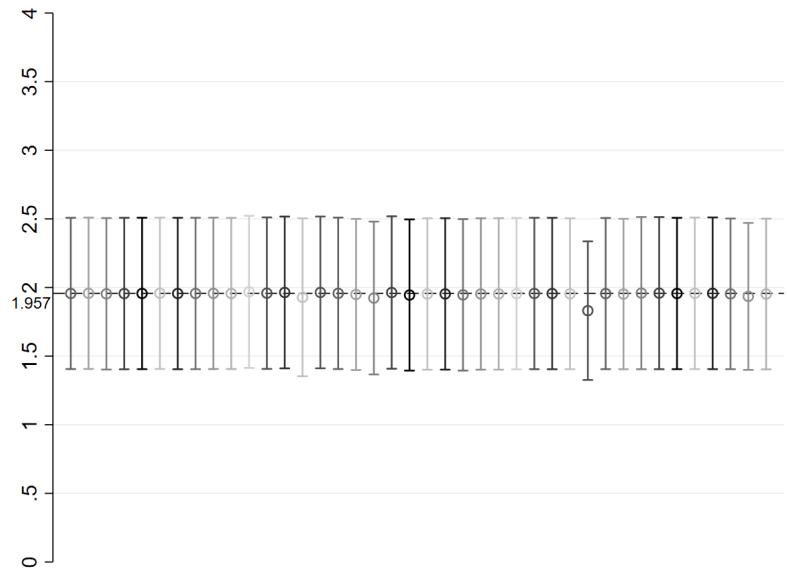
terms of i) the magnitude of the trade-induced economic shock, and ii) the number of obits in 1991.

Figures 6 and 7 graphically present the results of the estimations removing the top-ranked micro-region in terms of RTC_c and total death in 1991 respectively, emphasizing the coefficient of interest and the 95% confidence interval for each one of the regressions. We start from the highest RTC_r or number of deaths in 1991 on the left. The dotted black line describes the coefficient obtained in our baseline regression.

Similarly, Figures 8 and 9 present the results of estimation removing the lowest-ranked micro-region in terms of the trade-induced shock and total deaths in 1991 respectively. Notice that both qualitatively and quantitatively the additional results are very close to the estimation presented in the paper.



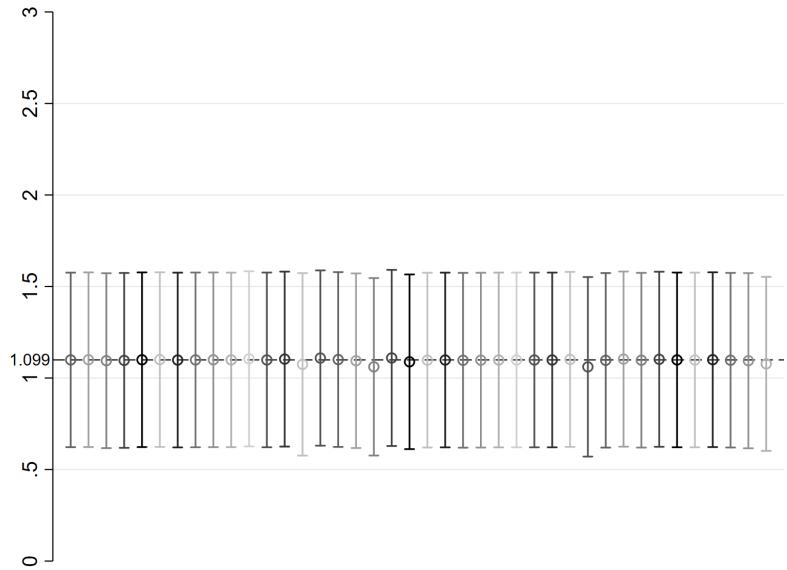
(a) 1991-2000 differences



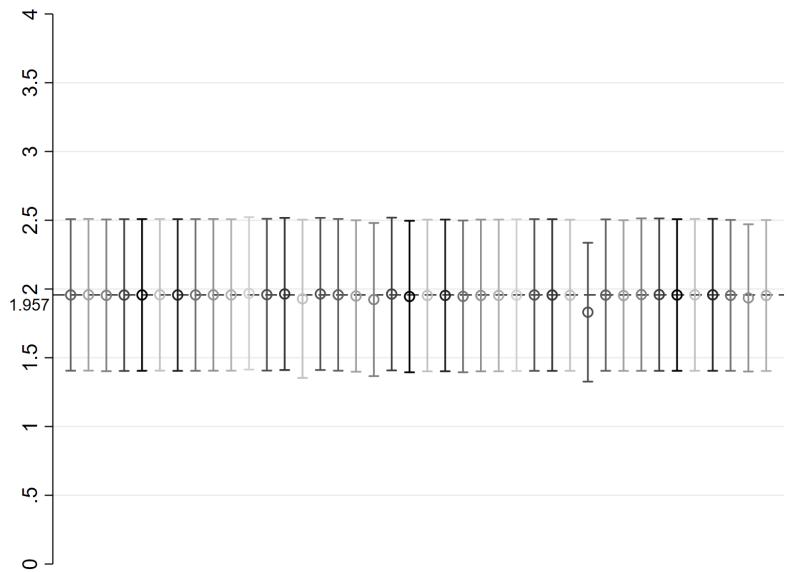
(b) 1991-2010 differences

Figure 6: Regional tariff change and log-changes in all-cause mortality rate
(Exclusion of micro-regions - Top 40 based on RTC_r value)

Notes: The figure shows the robustness of the results to excluding, one by one, the top 40 micro-regions with the highest RTC_r . The estimated coefficients and confidence intervals at 95 percent are reported. Each estimated coefficient and confidence interval emanate from a single estimation. Micro-regions ranked from the left to the right - that is, higher RTC_r exposure is the first observation.



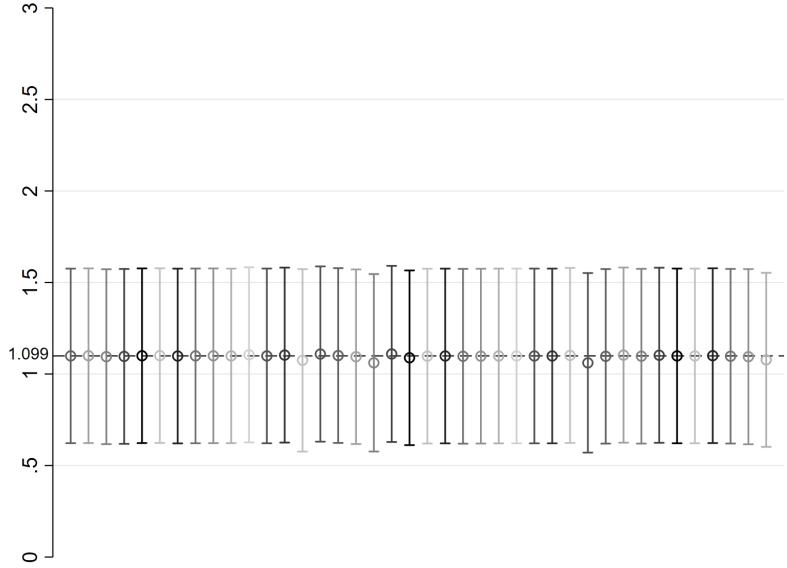
(a) 1991-2000 differences



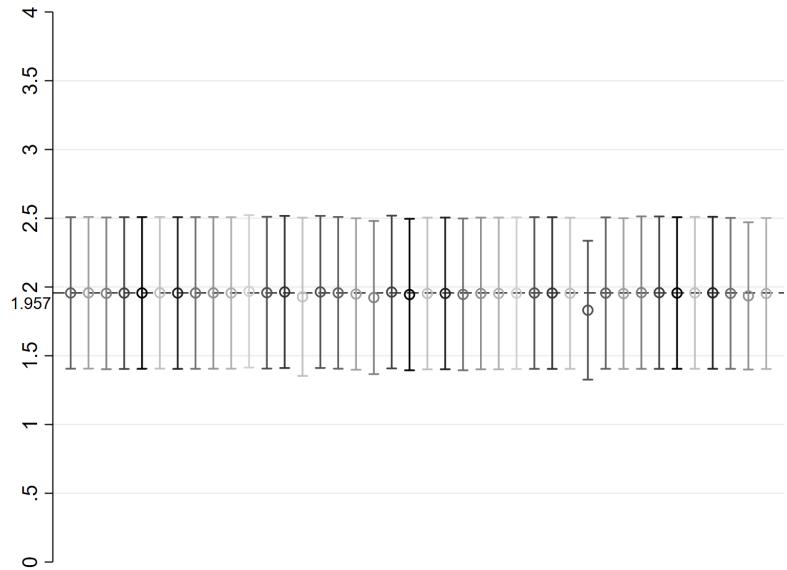
(b) 1991-2010 differences

Figure 7: Regional tariff change and log-changes in all-cause mortality rate
(Exclusion of micro-regions - Top 40 based on obits in 1991)

Notes: The figure shows the robustness of the results to excluding, one by one, the top 40 micro-regions with the highest number of deaths in 1991. The estimated coefficients and confidence intervals at 95 percent are reported. Each estimated coefficient and confidence interval emanate from a single estimation. Micro-regions ranked from the left to the right - that is, the first coefficient on the left is associated with the micro-region with the largest number of obits in 1991.



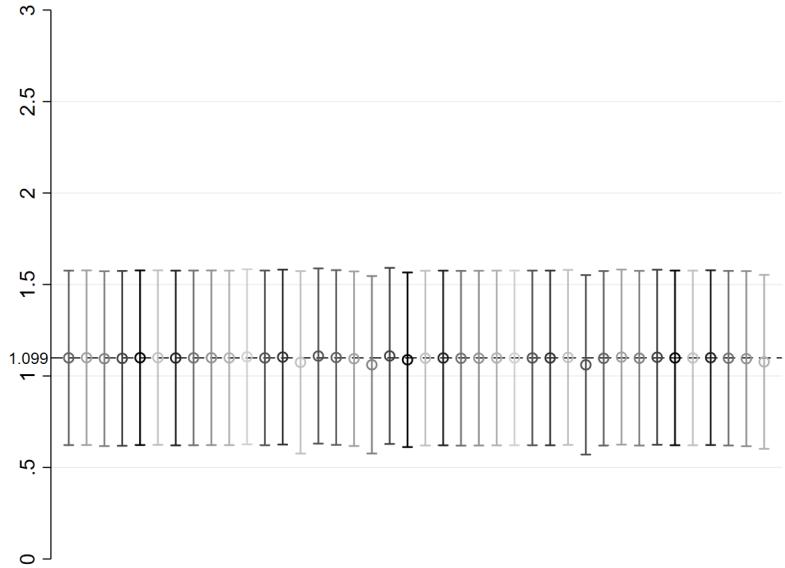
(a) 1991-2000 differences



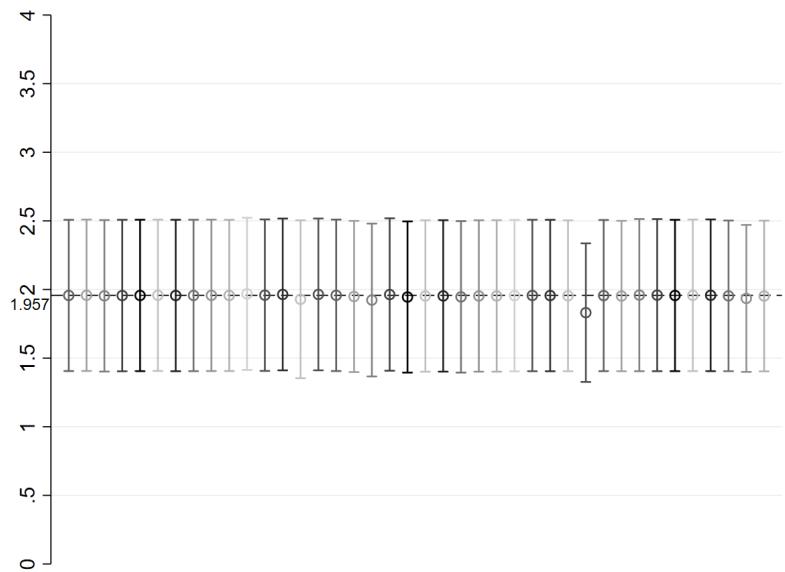
(b) 1991-2010 differences

Figure 8: Regional tariff change and log-changes in all-cause mortality rate
(Exclusion of micro-regions - Last 40 based on RTC_r value)

Notes: The figure shows the robustness of the results to excluding, one by one, the last 40 micro-regions with the lowest RTC_r . The estimated coefficients and confidence intervals at 95 percent are reported. Each estimated coefficient and confidence interval emanate from a single estimation. Micro-regions ranked from the left to the right - that is, the lowest RTC_r exposure is the first observation.



(a) 1991-2000 differences



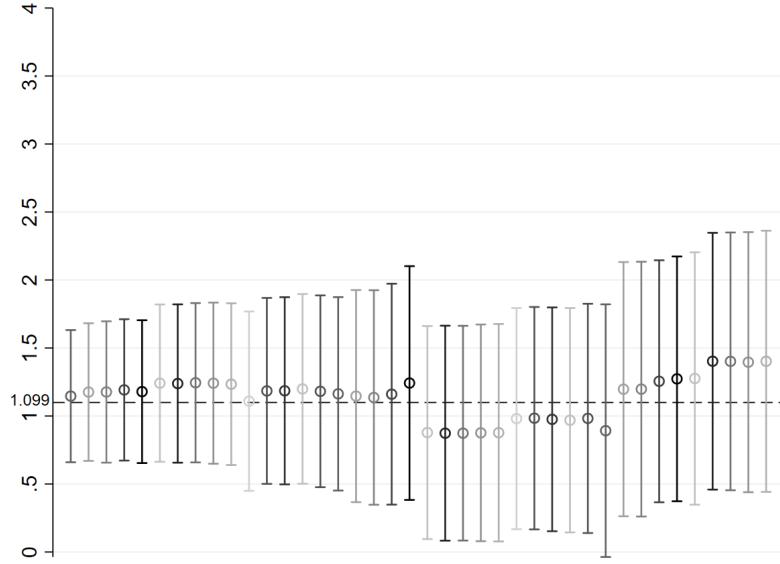
(b) 1991-2010 differences

Figure 9: Regional tariff change and log-changes in all-cause mortality rate
(Exclusion of micro-regions - Last 40 based on obits in 1991)

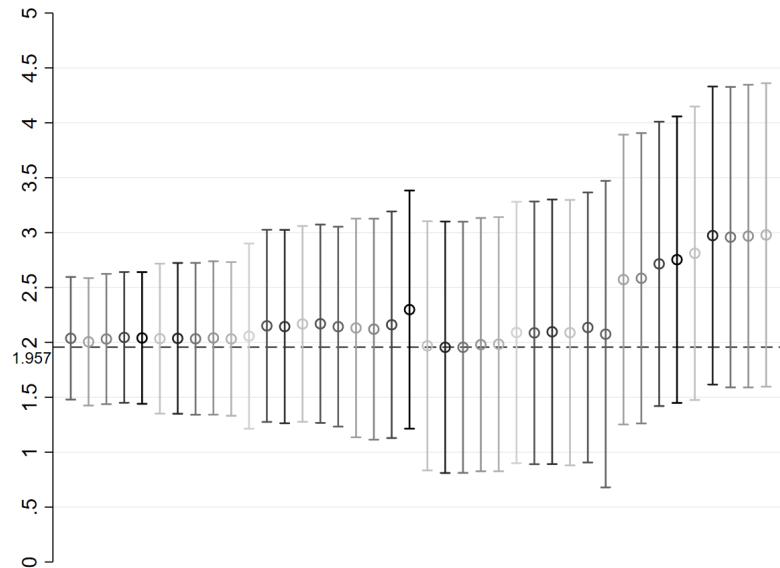
Notes: The figure shows the robustness of the results to excluding, one by one, the last 40 micro-regions with the lowest number of deaths in 1991. The estimated coefficients and confidence intervals at 95 percent are reported. Each estimated coefficient and confidence interval emanate from a single estimation. Micro-regions ranked from left to right - that is, the first coefficient on the left is associated with the micro-region with the lowest number of obits in 1991.

Next, we perform a similar sensitivity analysis, but now we drop the top-ranked micro-regions following the same criteria one to many. That is, we sequentially drop the top-ranked local economies from the estimation up to roughly 10% of our total sample (40 micro-regions). Figures 10 and 11 present, respectively, the results for excluding regions classified by RTC_r and total deaths in 1991. Again, we emphasize the coefficient of interest and the 95% confidence interval for each one of the regressions. We start from the highest RTC_r or number of deaths in 1991 on the left. The dotted black line describes the coefficient obtained in our baseline regression.

Similarly, Figures 12 and 13 present, respectively, the results for excluding regions classified from the lowest exposure to the trade-induced economic shock and total deaths in 1991.



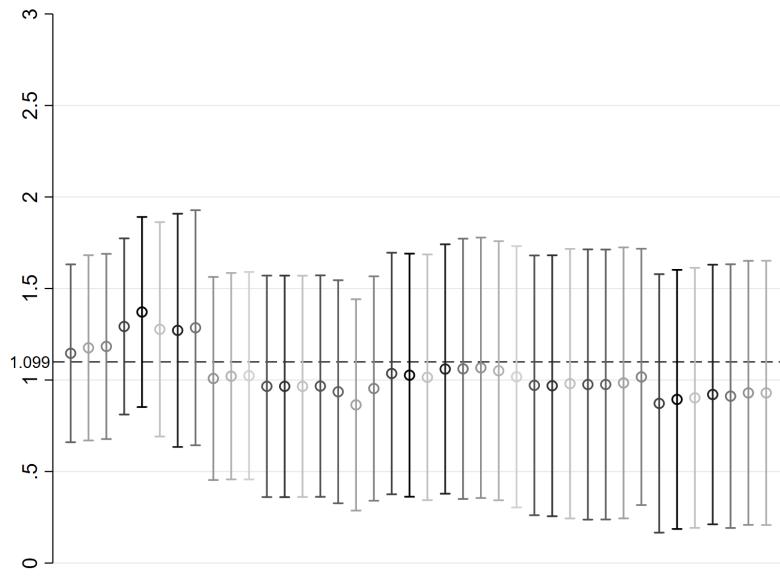
(a) 1991-2000 differences



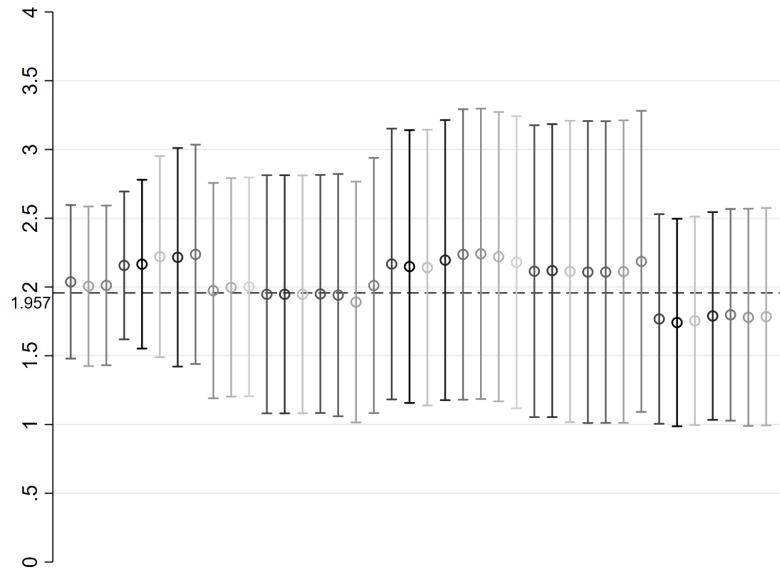
(b) 1991-2010 differences

Figure 10: Regional tariff change and log-changes in all-cause mortality rate
(Exclusion of micro-regions one to many - Top 40 based on RTC_r value)

Notes: The figure shows the robustness of the results to excluding the top 40 micro-regions with the highest RTC_r . The estimated coefficients and confidence intervals at 95 percent are reported. Each estimated coefficient and confidence interval emanate from a single estimation. Micro-regions ranked from the left to the right - that is, higher RTC_r exposure is the first observation.



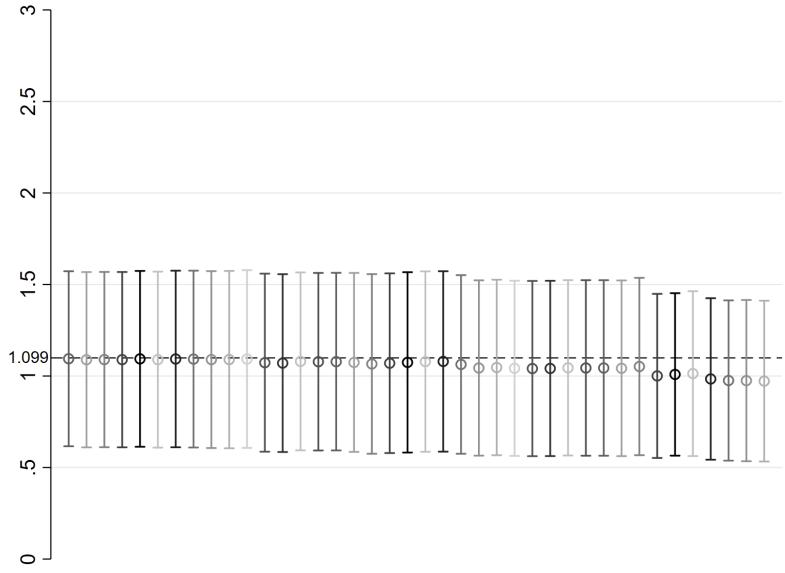
(a) 1991-2000 differences



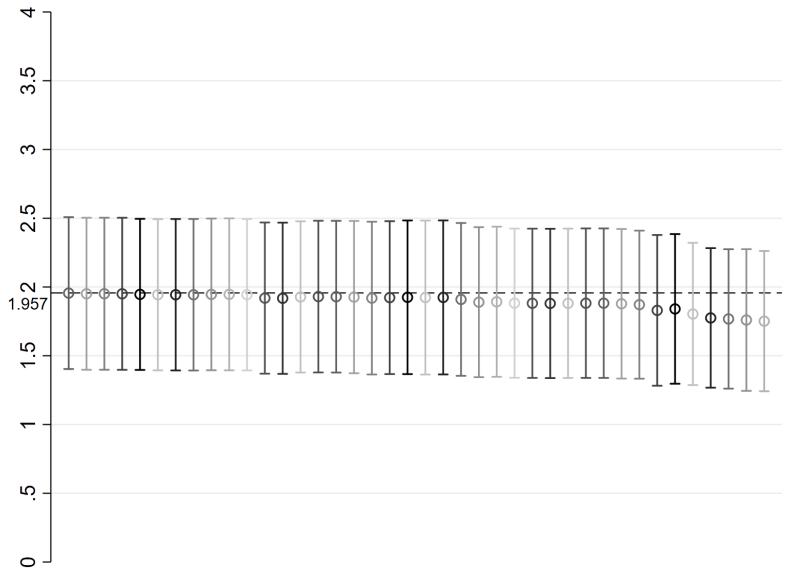
(b) 1991-2010 differences

Figure 11: Regional tariff change and log-changes in all-cause mortality rate
(Exclusion of micro-regions one to many - Top 40 based on obits in 1991)

Notes: The figure shows the robustness of the results to excluding the top 40 micro-regions with the highest number of deaths in 1991. The estimated coefficients and confidence intervals at 95 percent are reported. Each estimated coefficient and confidence interval emanate from a single estimation. Micro-regions ranked from the left to the right - that is, the first coefficient on the left is associated with the micro-region with the largest number of obits in 1991.



(a) 1991-2000 differences



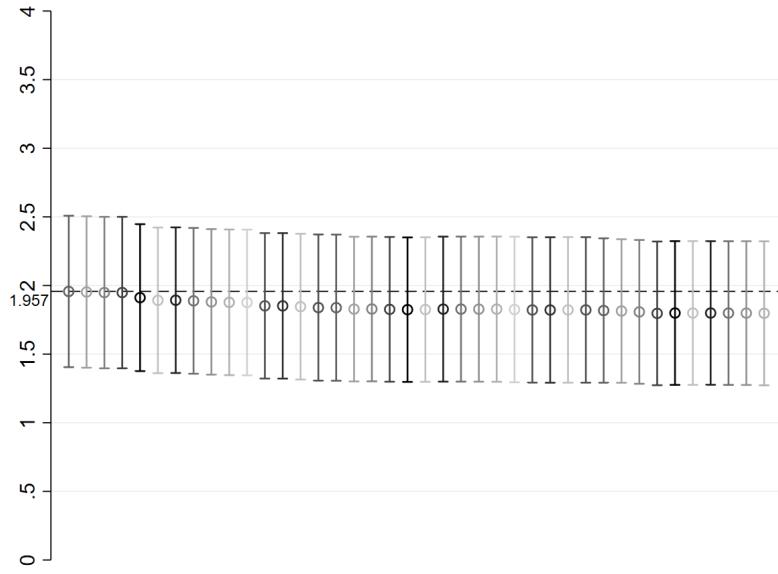
(b) 1991-2010 differences

Figure 12: Regional tariff change and log-changes in all-cause mortality rate
(Exclusion of micro-regions one to many - Last 40 based on RTC_r value)

Notes: The figure shows the robustness of the results to excluding the last 40 micro-regions with the lowest RTC_r . The estimated coefficients and confidence intervals at 95 percent are reported. Each estimated coefficient and confidence interval emanate from a single estimation. Micro-regions ranked from the left to the right - that is, the lowest RTC_r exposure is the first observation.



(a) 1991-2000 differences



(b) 1991-2010 differences

Figure 13: Regional tariff change and log-changes in all-cause mortality rate
(Exclusion of micro-regions one to many - Last 40 based on obits in 1991)

Notes: The figure shows the robustness of the results to excluding the last 40 micro-regions with the lowest number of deaths in 1991. The estimated coefficients and confidence intervals at 95 percent are reported. Each estimated coefficient and confidence interval emanate from a single estimation. Micro-regions ranked from the left to the right - that is, the first coefficient on the left is associated with the micro-region with the lowest number of obits in 1991.

E Robustness to Different Time Frames

The choice of the analysis period warrants a detailed explanation. As highlighted in the article, Charris et al. (2023) – who similarly utilize mortality records, although focusing solely on infant mortality – note the limitations of using data before 1985 for longitudinal comparisons. This limitation stems from the incomplete coverage of mortality records across all the municipalities that were consistently observed from the 1980s to 2010. In their study, in fact, the exclusion of observations from the period between 1979 and 1985 reflects a forced balancing of the panel. Importantly, Charris et al. (2023) use minimum comparable areas (municipalities that are consistently observed over the period of interest) as the unit of analysis in the empirical strategy. However, it is imperative to note that this approach differs from the methodology employed in our case.

Remember that we are using micro-regions as a unit of analysis. These local economies encompass municipalities with similar socioeconomic characteristics and locations. Importantly, for every year since 1980, we have mortality records for all the 411 local economies analyzed in the original manuscript. The exclusion of these years from our analysis stems from a significant issue that could substantially compromise the integrity of our econometric specifications: the introduction of substantial measurement error.

This problem occurs for quite straightforward reasons. As we are aggregating mortality and population records at the municipality level to construct mortality rates by micro-region, the absence of mortality records for certain municipalities during this period poses significant challenges in accurately constructing regional mortality rates. It is worth mentioning that population estimates data are more representative of the sample as they rely on census data.¹ For each period before 1985, the existence of missing values for mortality records at the municipality level indicates that we are computing populations of all municipalities that compose a particular micro-region but not necessarily the “correct” number of deaths in those localities.

This discrepancy in data availability across municipalities results in an underestimation of mortality rates for regions with missing mortality records. Given that the omitted municipalities before 1985 were predominantly located in relatively less developed states (in terms of socioeconomic conditions) during the period under analysis, and considering that the tariff cuts predominantly impacted more industrialized and wealthier regions of the country, this introduces a specific bias in our estimates. Consequently, we might measure disproportio-

¹Demographic censuses have been carried out decennially in Brazil, with some exceptions, since 1890.

ately lower (higher) mortality rates for regions subsequently relatively less (more) affected by the trade-induced regional economic shock.

Moreover, the comparison of mortality rates for the same micro-regions over time becomes problematic. This is particularly due to the staggered expansion of mortality record reporting coverage by the Ministry of Health between 1979 and 1984 (Charris et al., 2023). If several municipalities within a micro-region begin reporting mortality records at different points before 1985, achieving consistent measurements of regional mortality rates becomes problematic, even for the same geographic location.

To better illustrate the limitation of mortality records in the years between 1979 and 1984, Figure 14 presents the number of missing values in terms of the numbers of deaths for the minimum comparable areas (municipalities consistently observed between the end of the 1970s and 2010) per year compared to the initial year of our sample – and previously used in the literature – 1985. For better comparison, the minimum number of comparable areas in this period is just over 3,000 municipalities. Notably, there is a substantial expansion in the number of localities reporting mortality records in the years preceding 1985. Particularly noteworthy is the significant reduction in missing values between 1979 and 1980 – from more than 270 localities (almost 10% of the sample) to approximately 170 –, as well as between 1982 and 1983 – from almost 130 to less than 50. Although the decline continues until reaching zero in 1985, the number of observations becomes relatively comparable from 1983 onwards.

For these reasons, we opted not to include years before 1985 in our sample, given concerns about measurement error in estimating regional mortality rates. However, to avoid concerns for readers, we conducted an empirical exploration of the dynamic impacts of trade liberalization on regional mortality rates, including the years 1980-1984 in the sample. The estimated specification is quite similar to Equation (2) of the article, but now we use, as a pre-trend, the delta in log mortality rates between 1991 and 1980. Figure 15 presents the results of such empirical exploration, separating the pre-liberalization period between i) the years with potential measurement error in measuring regional mortality rates (between 1980 and 1984), and, ii) the years originally used in the manuscript (between 1985 and 1990).

First, it is worth highlighting that the general results are very similar to those presented in the article. Particularly, comparing the dynamic impacts of regional tariff changes on all-cause log mortality rates presented in Figures 15 and Figure 4 of the article, it is evident that the results from including the years 1980-1984 exhibit similar trends. Nevertheless, it is crucial to highlight that, as discussed previously, we are capturing potential problems

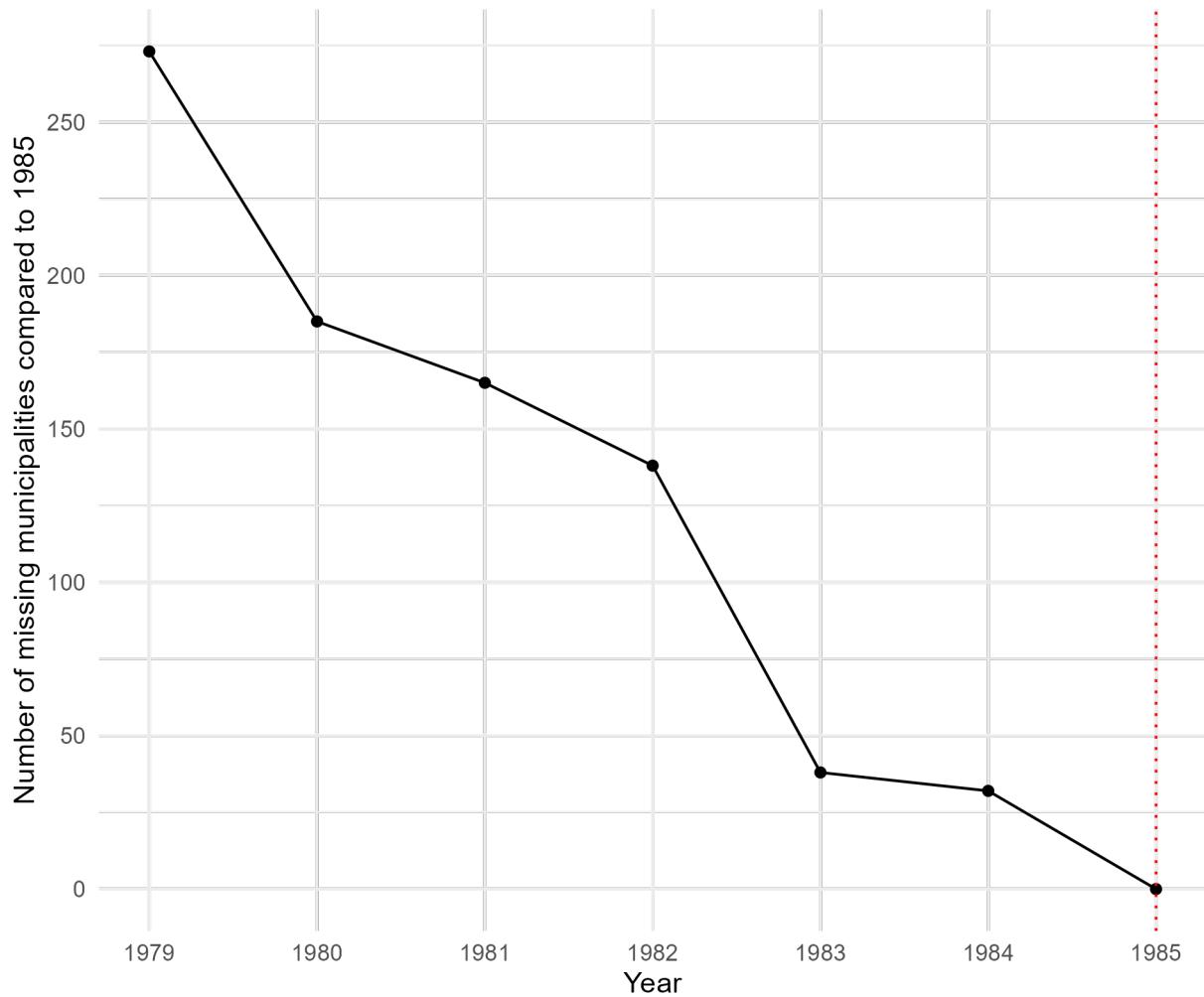


Figure 14: Missing observations per year - Mortality data

Source: Mortality records per place of residency (municipality level) of individuals are from DATA-SUS (administrative dataset from the Ministry of Health).

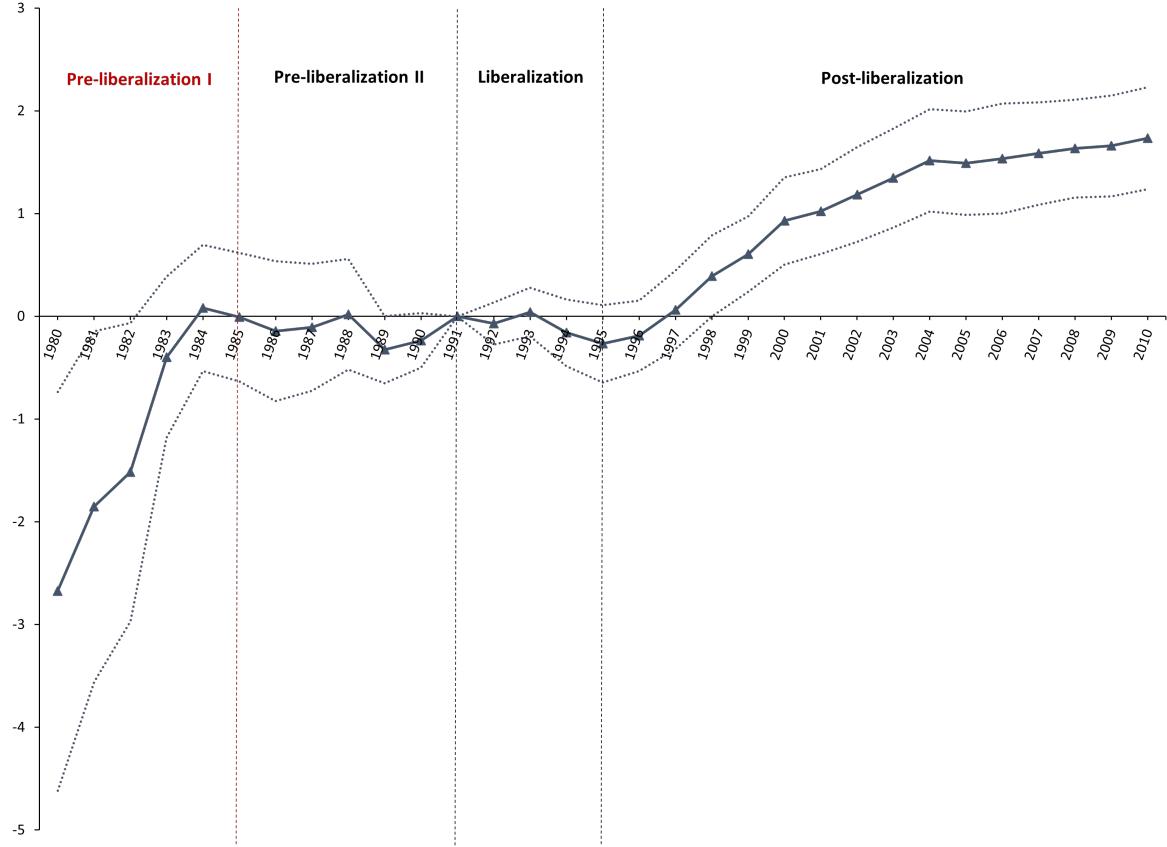


Figure 15: Dynamic effects of regional tariff changes on log changes in mortality rates - 1980-2010

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (2) of the article, where the dependent variable is the change in regional log all-cause mortality rates and the independent variable is the RTC_r in $t = 1980, \dots, 2010$. All regressions include state fixed effects. Standard errors are adjusted for 91 meso-region clusters and the observations are weighted by population. Dashed lines show 95 percent confidence intervals.

in the data, mainly before 1984. Notably, in addition to estimating negative effects that are in line with our hypothesis that the measurement error would be negatively biasing the coefficients – thus, the comparison between regions with different exposures to the shock is compromised – we are also capturing significant noise in the estimates in these initial years. This can be seen by the larger relative size of the standard errors of the estimated coefficients before 1983. These standard errors appear to stabilize in a less noisy range only after 1983. This noise likely stems from the measurement error inherent in the data before 1984, further

supporting our decision to exclude this period from our analysis.

To address the question of whether the data completeness in certain regions might bias our results for the pre-liberalization period, we undertook a detailed examination. Initially, we identified all municipalities that exhibited any missing values in the annual number of deaths from 1979 to 1984. We then flagged and excluded from our dataset any micro-regions that included these municipalities. This process resulted in the exclusion of just over 27% of our micro-regions (115 observations), ensuring that our analysis retained only regions with consistently available data starting in 1979.

We then employed a dynamic model similar to the previous for a refined subset that excluded micro-regions containing municipalities potentially compromised by missing mortality data (i.e. excluding flagged observations).

The outcomes of this meticulous approach are depicted in Figure 16. Comparing with Figure 15, it is direct to note that the observed increase in mortality during the years with limited data is predominantly attributable to micro-regions with incomplete records. Importantly, this result confirms the consistency of the parallel trend assumption across the pre-treatment periods, with statistically significant coefficients emerging only post-liberalization.

In light of these considerations, the exclusion of mortality observations from the period 1979-1984 is a methodological necessity aimed at preserving the integrity and reliability of our analysis.

F Explicit Link Between Economic Changes and Employment

In this section, we examine the repercussions of the trade liberalization episode on local labor markets.

To sharpen our analysis, and following the existing literature, we examine the trade liberalization's implications on local labor markets, explicitly linking the economic impacts of tariff reform to employment changes in Brazilian regional economies. Using Census data, we construct a key employment outcome metric — namely, the share of 20 to 45-year-old men and women employed outside the home for pay — at the municipality level, later aggregated to the micro-region level. According to Charris et al. (2023), this measure largely includes formal employment in tradable industries and is an alternative to the formal employment measures analyzed in Dix-Carneiro and Kovak (2017)². Additionally, we incorporate the

²The authors use data from the *Relação Anual de Informações Sociais* (RAIS), an administrative dataset assembled yearly by the Brazilian Ministry of Labor providing formal labor market measures.

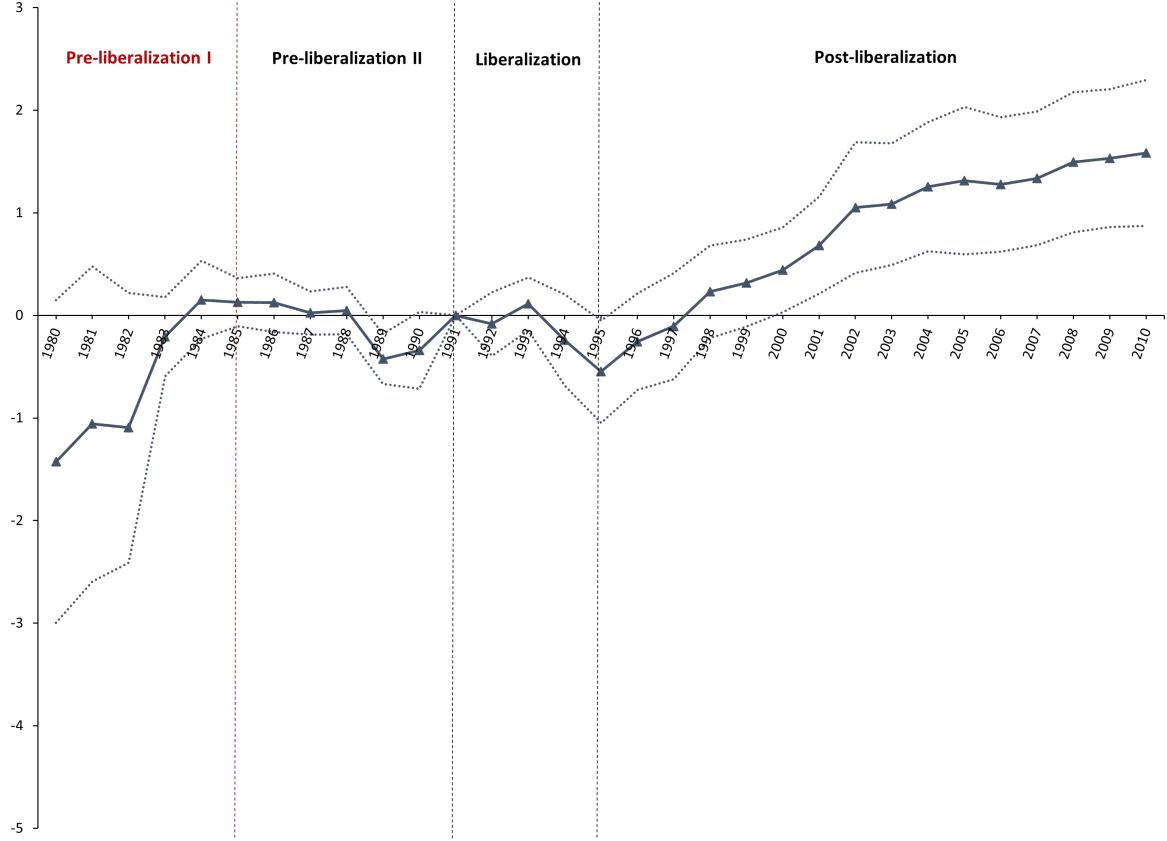


Figure 16: Dynamic effects of regional tariff changes on log changes in mortality rates (excluding flagged micro-regions) - 1980-2010

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (2) of the article, where the dependent variable is the change in regional log all-cause mortality rates and the independent variable is the RTC_r in $t = 1980, \dots, 2010$. All regressions include state fixed effects. Standard errors are adjusted for 91 meso-region clusters and the observations are weighted by population. Dashed lines show 95 percent confidence intervals.

broad employment measure utilized in Dix-Carneiro et al. (2018), which examines individuals' employment status (employed or not employed), also sourced from Census data. Table 3 presents a summary of our findings.

Table 3: Regional tariff changes and alternative employment outcomes

Dependent variables:	Share of men working outside for money		Share of women working outside for money		Employment	
	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
	(1)	(2)	(3)	(4)	(5)	(6)
RTC_r	0.503*** (0.149)	0.585*** (0.193)	0.386*** (0.113)	0.394*** (0.143)	0.643*** (0.06)	-0.051 (0.102)

Notes: Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. There are 411 micro-region observations. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

From the first four columns of Table 3, we document that the trade-induced regional economic shocks led to a statistically significant and negative decline in employment, consistent with the literature, for both men and women (Braga, 2018; Charris et al., 2023; Dix-Carneiro & Kovak, 2017). Note that, comparing columns (1) and (2) to (3) and (4), there appears to be a stronger reduction in formal employment for men. This result is also consistent with Gaddis and Pieters (2017), who document a more adverse impact of trade liberalization on the employment level of men than that of women, potentially reflecting the fact that males are more likely to work in industries that experienced larger import tariff cuts (e.g. manufacturing).

Nevertheless, note that when we consider the overall employment measure, in columns (5) and (6), there seems to be a reduction of employment rates in the regions more impacted by the trade-induced economic shock in the medium run but this effect vanishes in the long run. Dix-Carneiro and Kovak (2017) document that formal sector employment rates were consistently and increasingly impacted over time but, importantly, the authors also document that the long-run recovery in employment rates experienced by harder-hit regions reflects relative increases in informal employment, while formal employment keeps falling. Further analysis by Charris et al. (2023) underscores this pattern, revealing a significant increase in self-employment among males alongside a meaningful and permanent drop in overall employment rates for women. The authors also find that greater exposure to the tariff cuts is associated with lower aggregate household income in the medium run, with an intensification of the result in the long run – reflecting the persistent and amplifying deterioration in economic conditions of the regions more impacted by the tariff cuts.

The consistency in the dynamic impacts of the trade reform on mortality and employment outcomes suggests that changes in local economic conditions — as highlighted by the

pro-cyclical mortality literature -- play a crucial role in explaining our findings. To contextualize these results, we reproduce Figure 4 from Dix-Carneiro and Kovak (2017) for better visualization, alongside the dynamic effect of the regional tariff change on all-cause mortality rates in Figure 18. It is worth noting that Dix-Carneiro and Kovak (2017) use the regional tariff reduction as the shift-share instrumental variable – the negative of our measure of regional tariff *change*, so their coefficients are compared with the negative values of ours. Both employment reductions and mortality rate decreases show substantial growth from 2000 to 2010, as highlighted, respectively, in Figures 17 and 18, indicating a continued divergence for regions facing different exposures to the tariff reform. Importantly, for both cases, note that most of the divergence in employment and mortality growth for regions facing different regional tariff changes “was complete by 2004, after which the estimates level off.” (Dix-Carneiro & Kovak, 2017, p. 2923).

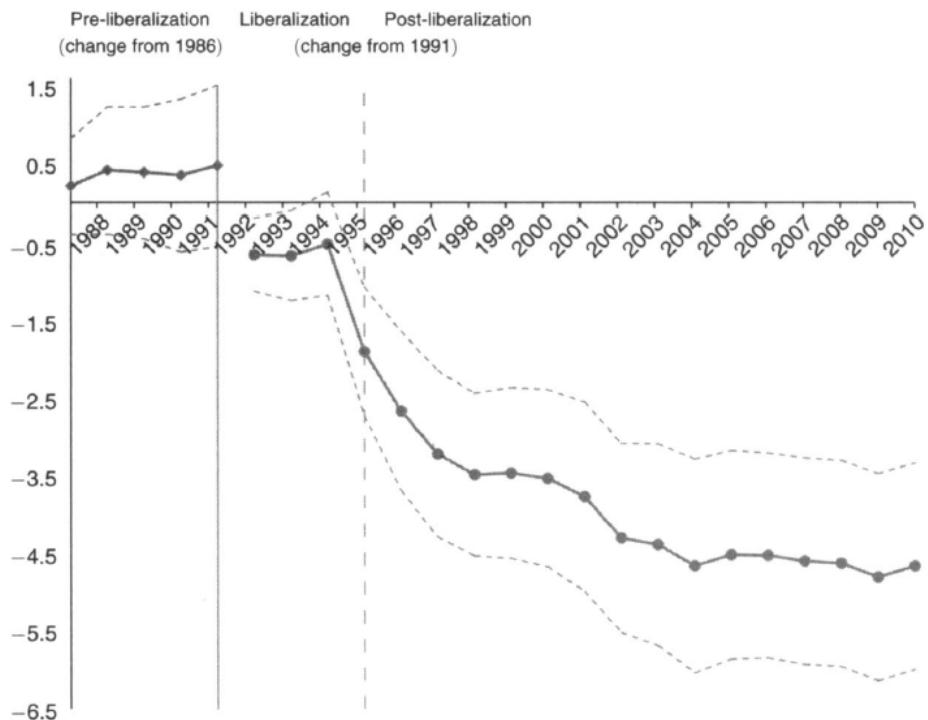


Figure 17: Dynamic effects of regional tariff reductions on log changes in formal employment

Source: Figure 4 from Dix-Carneiro and Kovak (2017)

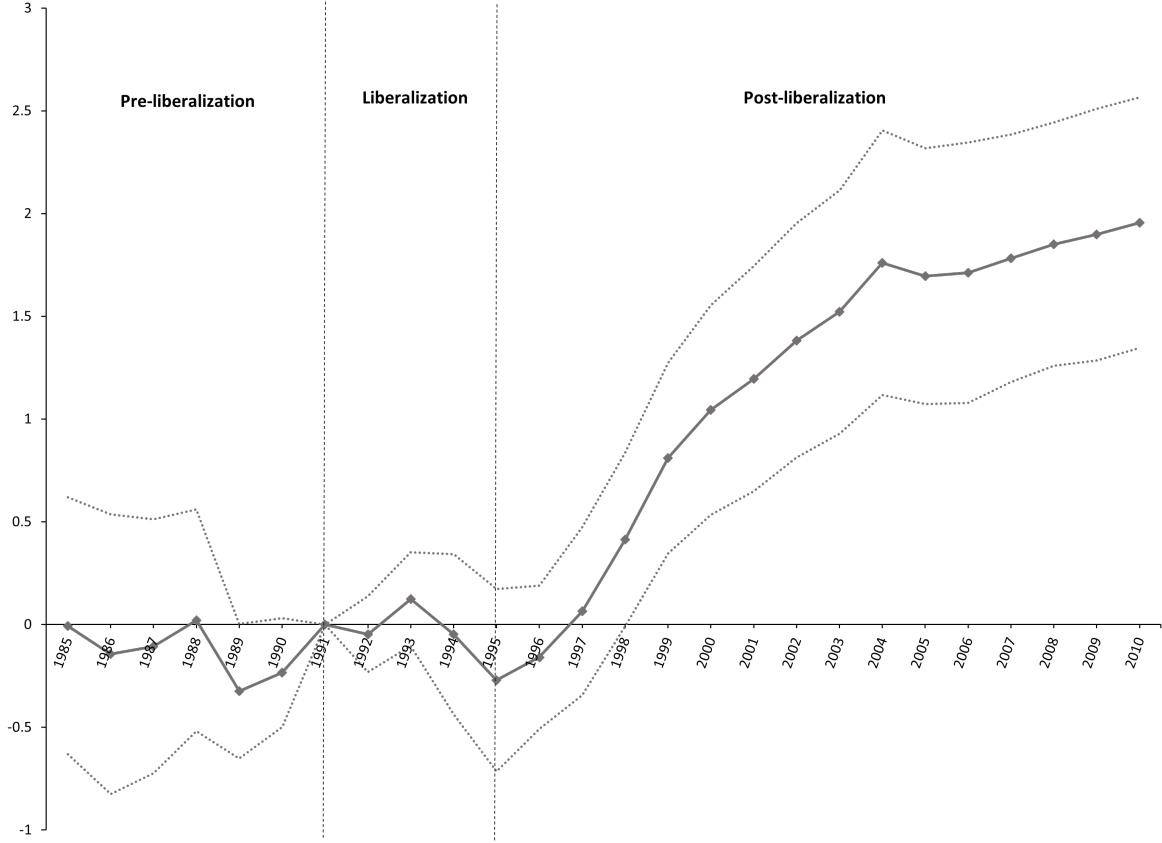


Figure 18: Dynamic effects of regional tariff changes on log changes in mortality rates

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (3) in Manuscript (2023), where the dependent variable is the change in regional log all-cause mortality rates - total, male and female - and the independent variable is the RTC_r in $t = 1985, \dots, 2010$. All regressions include state fixed effects. Dashed lines show 95 percent confidence intervals. Standard errors are adjusted for 91 meso-region clusters.

G Sub-sample Analysis

In this section, we delve deeper into the relationship between employment dynamics, health-care expenditure, and mortality reductions stemming from the regional economic shock induced by trade liberalization. Our analysis focuses on sub-samples drawn from regions with varying initial conditions, particularly those characterized by distinct levels of employment and healthcare spending. To do so, we use two sources of data to construct initial conditions for Brazilian regions before or during the episode of trade liberalization. First, we use the Brazilian Decennial Population Census of 1991, from the Brazilian Institute of Geography

and Statistics (*IBGE - Instituto Brasileiro de Geografia e Estatística*), to compute population sizes of micro-regions, as well as employment, earnings, and income measures at the regional level. Second, we computed the baseline government spending per category at the municipality level, in 1995, with data from the Ministry of Finance (*Ministério da Fazenda - Secretaria do Tesouro Nacional*).

Our analysis unfolds along two main avenues. Firstly, we consider stratifying micro-regions into distinct quartiles based on pertinent variables, effectively reducing the sample size for deeper examination. Alternatively, we maintain the sample size and explore interactions between the treatment variable — regional tariff change — and binary indicators denoting whether a region falls above or below the median for specific initial conditions. Adopting both strategies, we endeavor to provide a comprehensive understanding of how variations in employment, population, income, and government expenditures influenced mortality outcomes post-trade liberalization.

G.1 Differing initial population, employment and income conditions

We initiate our analysis by scrutinizing the impact of regional tariff changes on mortality rate fluctuations in micro-regions positioned in the first and fourth quartiles of population, employment, remuneration, and per capita income distributions in 1991. Table 4 presents the outcomes of this approach.

Table 4: Sub-sample analysis: regional tariff change and mortality rates per population, employment, labor earnings and income per capita

Sub-sampling variable - base 1991	Population				Employment				Earnings				Income per capita			
	1st quartile		4th quartile		1st quartile		4th quartile		1st quartile		4th quartile		1st quartile		4th quartile	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
Dep. var: $\Delta \log(MR_r)$	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
RTC_r	4.993*** (1.704)	6.181*** (1.871)	1.112*** (0.341)	1.842** (0.394)	0.402 (0.727)	1.372 (1.176)	0.746** (0.247)	1.265*** (0.285)	5.022*** (1.484)	8.173*** (1.771)	1.043*** (0.301)	2.030*** (0.383)	6.911** (2.628)	11.02*** (2.939)	0.976*** (0.236)	2.073** (0.361)
Observations	102	102	103	103	102	102	103	102	102	103	103	102	102	103	102	103

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Note that, for all variables, we observed the pattern of magnification of the reduction in the mortality rate from all causes in the long term compared to the medium term. Furthermore, for population, remuneration, and income per capita, we observed a similar pattern, with the coefficients associated with micro-regions in the first quartile of the distribution being greater in magnitude than those in the fourth quartile. Notably, the magnitude of the average coefficient presented, for example, in Table 1 of the manuscript is much closer to the coefficients estimated for the micro-regions in the upper quartile of the distribution of all

base variables. This is a preliminary indication that there is a correlation between exposure to regional tariff change and population, employment rate, earnings, and per capita income. This point is discussed in the article, but we further explore such correlations below.

However, the most relevant result for our analysis which corroborates the pro-cyclical nature of the mortality rate occurs in relation to the variation in the employment rate. Note that the reduction in the mortality rate from all causes is not statistically significant either in the medium or long term in regions that were in the lower quartile of the employment rate distribution before the shock. The average result presented in the manuscript appears to be driven by the most economically active regions, at least in terms of the local labor market.

The results are qualitatively similar when we follow the alternative strategy of maintaining the number of observations in the regressions and interacting the treatment with a binary variable that indicates whether the region is above or below the median of the distributions associated with the baseline characteristics of the micro-regions. Table 5 summarizes our findings.

Table 5: Regional tariff change and mortality rates - above/below median for baseline population and employment characteristics

<i>Sub-sampling variable - base 1991</i>	Population		Employment		Earnings		Income per capita	
<i>Dep. var: $\Delta \log(MR_r)$</i>	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
$RTC_r * Belowmedian$	1.663*** (0.386)	2.610*** (0.472)	1.085*** (0.276)	1.839*** (0.323)	2.594*** (0.620)	3.629*** (0.674)	2.809*** (0.880)	4.066*** (0.880)
$RTC_r * Abovemedian$	1.145*** (0.240)	2.011*** (0.280)	1.112*** (0.256)	2.072*** (0.268)	1.050*** (0.215)	1.901*** (0.248)	1.201*** (0.253)	2.085*** (0.275)

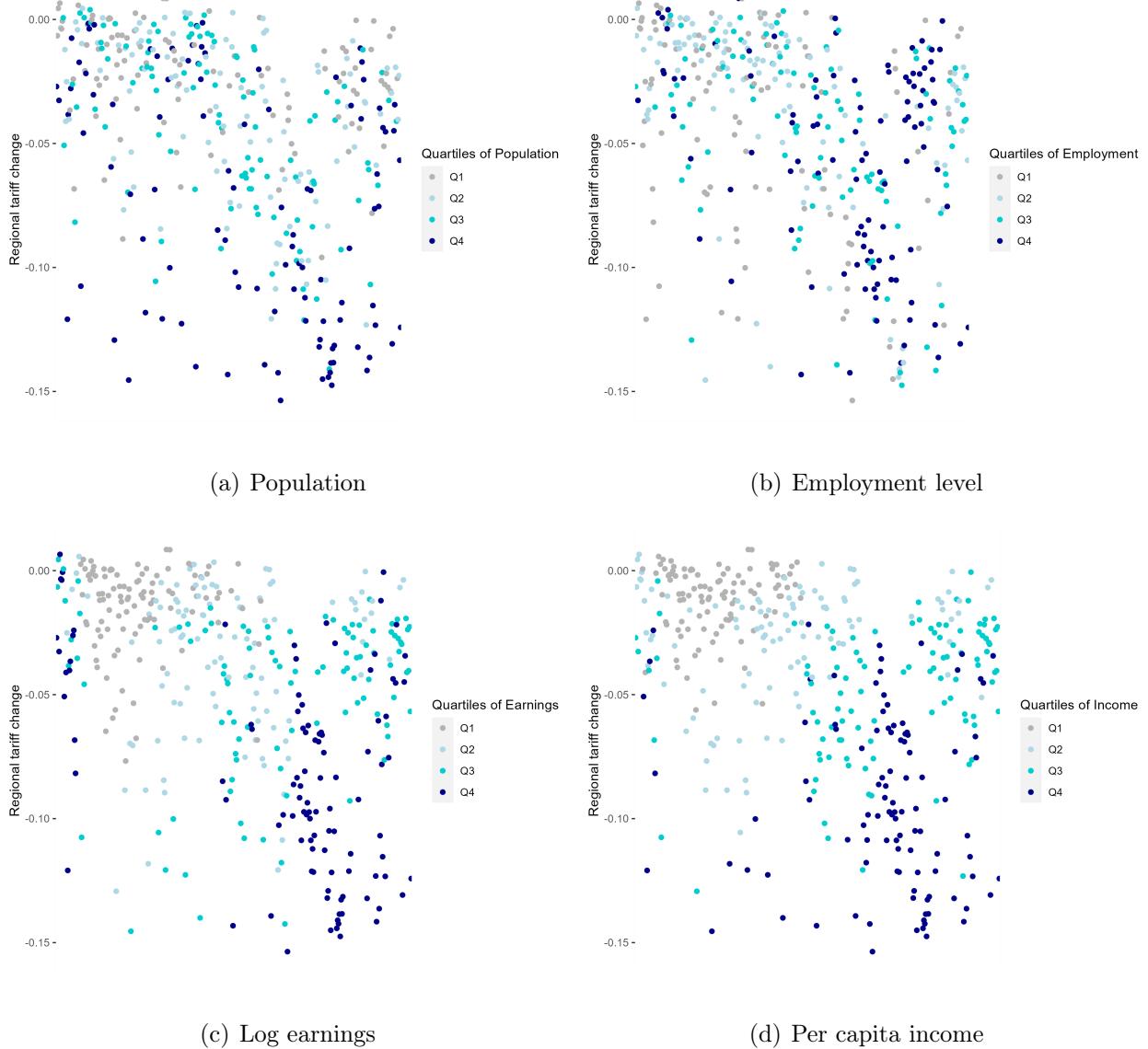
Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

To visually elucidate the correlation between exposure to regional tariff change and pre-treatment regional characteristics, we presented a simple visualization in Figure 19. In all panels, each point represents a micro-region and we distribute them according to exposure to tariff cuts, captured by the regional tariff change variable in the y-axis. The color - from gray to dark blue - indicates, in ascending order, the quartile of the distribution of the variables of interest that such observations occupied in 1991 (at the beginning of the treatment).

Note that, for all the variables in question, there is a clear concentration of the regions most affected by the trade-induced regional economic shock in the upper quartile of the distribution, while the regions that are in the first quartile of such variables appear to have been less affected by the tariff cuts.

Lastly, we disaggregated mortality rates into internal and external causes and interacted with the treatment with binary variables representing above or below median values of relevant base variables. The results are summarized in Table 6. Note that, in both Panel A and Panel B, the results are qualitatively similar to those presented previously. However, it is important to highlight that in Panel B, referring to external causes of mortality, the amplification of the reduction in the mortality rate in the regions most affected by the shock in the long term is notably greater in micro-regions characterized by more heated economic activity (regarding the labor market) than in those that are less vigorous in terms of employment. This evidence goes in the same direction as the argument made previously in this appendix that, especially concerning external causes of mortality, the relative improvement in the conditions of local labor markets in the long term, due to the return of working-age men to informal occupations, does not seem to profoundly affect the functioning of the pro-cyclic mortality mechanism.

Figure 19: Correlation: regional tariff change and census variables in 1991



Source: The regional tariff change is constructed, as described in Manuscript (2023), from tariff data from Kume et al. (2003). The variables of interest are computed from the Brazilian Population Census of 1991, obtained from IBGE.

Table 6: Regional tariff changes and log changes in mortality rates - Diseases and external causes for above/below median of baseline population and employment characteristics

Sub-sampling variable - base 1991	Population		Employment		Earnings		Income per capita	
Dep. var: $\Delta \log(MR_r)$	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
<i>Panel A. Diseases</i>								
<i>RTC_r * Belowmedian</i>								
	1.431*** (0.335)	2.423*** (0.474)	0.863*** (0.239)	1.594*** (0.320)	2.413*** (0.595)	3.438*** (0.659)	2.847*** (0.865)	4.078*** (0.890)
<i>RTC_r * Abovemedian</i>								
	0.976*** (0.201)	1.763*** (0.277)	1.005*** (0.234)	1.811*** (0.275)	0.887*** (0.171)	1.646*** (0.252)	1.049*** (0.212)	1.848*** (0.269)
<i>Panel B. External</i>								
<i>RTC_r * Belowmedian</i>								
	3.583** (1.656)	5.355*** (0.990)	3.093*** (0.993)	4.712*** (0.581)	5.402*** (1.455)	7.470*** (1.364)	2.568 (1.996)	5.202** (2.110)
<i>RTC_r * Abovemedian</i>								
	2.640** (1.021)	4.874*** (0.680)	2.038** (0.948)	4.946*** (0.867)	2.465** (0.959)	4.743*** (0.594)	2.558** (0.973)	4.853*** (0.696)

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

G.2 Differing initial government expenditures

Next, we examine the impact of regional tariff changes on mortality rate across micro-regions categorized into the first and fourth quartiles of government expenditure per function, health-care and sanitation expenditures (henceforth, H&S), and the share of H&S expenditures over total expenditures in 1995. Table 7 encapsulates the outcomes of this endeavor.

Table 7: Sub-sample analysis: regional tariff change and mortality rates per expenditure quartile

Sub-sampling variable - base 1995	Total expenditures per function								Expenditures H&S				Share H&S / Total			
	1st quartile				4th quartile				1st quartile		4th quartile		1st quartile		4th quartile	
	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
Dep. var: $\Delta \log(MR_r)$	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
RTC_r	2.254* (1.294)	3.836** (1.679)	1.038*** (0.196)	2.074*** (0.247)	1.507* (0.797)	2.246** (1.030)	0.905*** (0.181)	1.936*** (0.289)	0.603* (0.318)	0.894*** (0.318)	1.369*** (0.226)	1.369*** (0.485)	2.566*** (0.440)			
Observations	102	102	103	103	102	102	103	103	102	102	103	102	103			

Notes: The expenditure variables, used from determining the sub-samples, are measured in per capita terms. Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Consistent with our overarching findings, we observe a notable amplification of the reduction in mortality rates from all causes in the long term compared to the medium term across all expenditure variables. Intriguingly, regions with lower levels of per capita spending, both in general and specifically on health, prior to the tariff reform experienced the most significant reductions in mortality rates. This aligns with our earlier observations indicating

that regions with lower population densities, remuneration levels, and per capita incomes exhibited greater mortality reductions during the analysis period. However, the magnitude of the average coefficient presented, for example, in Table 1 of the article is much closer to the coefficients estimated for the micro-regions in the upper quartile of the distribution of base government expenditure variables. This is indicative that there is a correlation between exposure to regional tariff change and the levels of per capita government expenditure in all categories and in H&S. Again, we provide evidence of that correlation below.

Notwithstanding, it is important to highlight that the results of the columns associated with the share of H&S in total government spending do not present a similar pattern to that of the other spending variables. Notably, regions with higher relative spending on health compared to total government spending before the trade shock exhibited the greatest reduction in mortality rates in the long term. This result lends support to our argument that the expansion of healthcare infrastructure following the reform played a pivotal role in mitigating mortality in the regions most impacted by the economic shock induced by trade liberalization.

Similar to the previous case, our findings remain robust when employing alternative analytical approaches, including interactions between the treatment variable and binary indicators based on median values of relevant expenditure variables. Table 8 succinctly encapsulates these results.

Table 8: Regional tariff change and mortality rates - above/below median for baseline expenditures

<i>Sub-sampling variable - base 1995</i>	Total expenditures per function		Expenditures H&S		Share H&S / Total	
<i>Dep. var: $\Delta \log(MR_r)$</i>	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
<i>RTC_r * Belowmedian</i>	1.514*** (0.501)	2.235*** (0.576)	1.345*** (0.407)	1.934*** (0.433)	1.202*** (0.294)	1.903*** (0.317)
<i>RTC_r * Abovemedian</i>	0.971*** (0.207)	1.871*** (0.242)	1.039*** (0.230)	1.962*** (0.279)	1.037*** (0.238)	1.989*** (0.297)

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

To enhance our understanding of the correlation between pre-treatment government expenditures and exposure to tariff changes, we provide a visual representation in Figure 20. In all panels, each point represents a micro-region and we distribute them according to exposure to tariff cuts, captured by the regional tariff change variable in the y-axis. The color

- from gray to dark blue - indicates, in ascending order, the quartile of the distribution of the variables of interest that such observations occupied in 1995 (during the treatment). For both per capita expenditures in all categories and in H&S, there is a clear concentration of the regions most affected by the trade-induced regional economic shock in the upper quartile of the distribution, while the regions that are in the first quartile of such variables appear to have been less affected by the shock. Nevertheless, there is no clear correlation between the distribution of the share of H&S expenditures over the total before the liberalization and exposure to the trade shock.

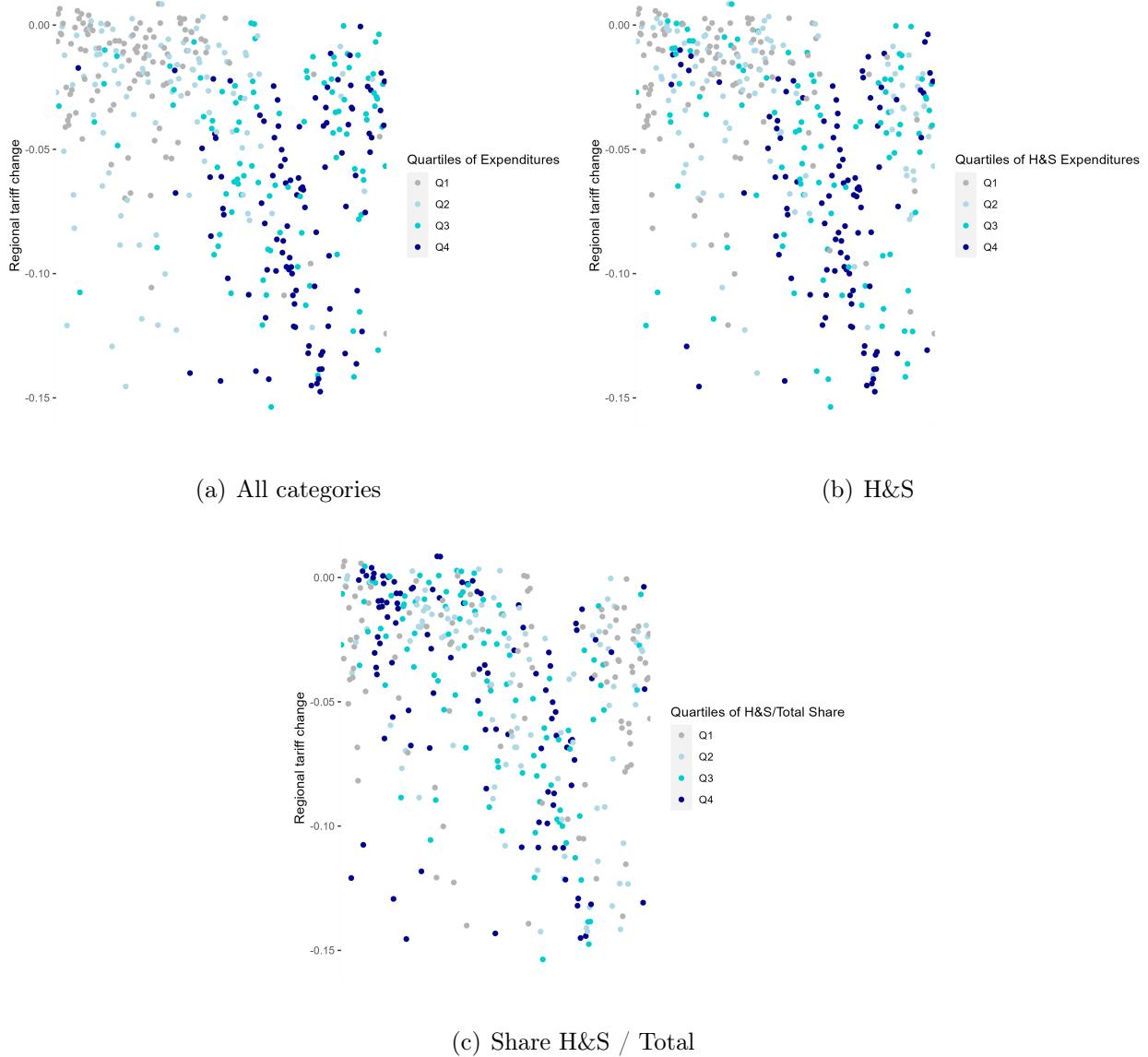
Lastly, disaggregating mortality rates by internal and external causes reaffirms the consistent magnitude patterns observed previously. The results, summarized in Table 9, underscore that the reduction in mortality rates was relatively stronger in regions with lower per capita government expenditures before treatment.

Table 9: Regional tariff changes and log changes in mortality rates - Diseases and external causes above/below median of baseline expenditures

<i>Sub-sampling variable - base 1995</i>	Total expenditures per function		Expenditures H&S		Share H&S / Total	
<i>Dep. var: $\Delta \log(MR_r)$</i>	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010
<i>Panel A. Diseases</i>						
$RTC_r * Belowmedian$	1.300*** (0.486)	2.035*** (0.590)	1.067*** (0.343)	1.716*** (0.453)	0.996*** (0.250)	1.665*** (0.330)
$RTC_r * Abovemedian$	0.823*** (0.150)	1.601*** (0.240)	0.903*** (0.207)	1.700*** (0.278)	0.899*** (0.214)	1.726*** (0.298)
<i>Panel B. External</i>						
$RTC_r * Belowmedian$	4.269*** (1.089)	5.442*** (1.045)	3.838*** (1.395)	4.698*** (0.609)	3.110*** (1.066)	4.600*** (0.583)
$RTC_r * Abovemedian$	2.029* (1.122)	4.641*** (0.696)	2.243** (0.882)	4.862*** (0.740)	2.225** (0.982)	4.967*** (0.780)

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Figure 20: Correlation: regional tariff change and per capita expenditures in 1995



Source: The regional tariff change is constructed, as described in Manuscript (2023), from tariff data from Kume et al. (2003). Annual government spending per category at the municipality level is from the Ministry of Finance (*STN - Secretaria do Tesouro Nacional*).

H Transport and Other Accidents

Table 10 presents the complete set of regressions for both transport and other accidents.

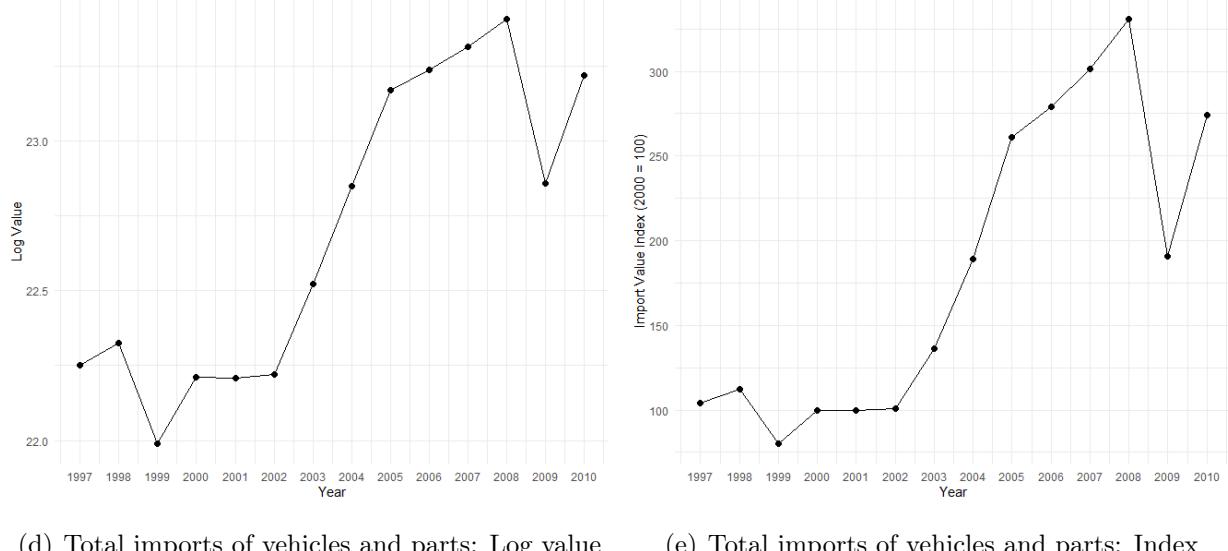
Table 10: Regional tariff changes and log changes in accident-related mortality rates

	OLS (1)		OLS (2)		OLS (3)		OLS (4)		2SLS (5)		Placebo (6)
Dep. var.: $\Delta \log(MR_r)$	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1985-1991
Transport Accidents											
RTC_r	5.992*** (0.771)	11.11*** (0.944)	6.774*** (1.227)	10.26*** (0.907)	5.950*** (1.080)	8.093*** (1.671)	6.170*** (0.788)	8.433*** (0.948)	5.874*** (0.839)	8.055*** (1.112)	0.633 (1.568)
$\Delta_{85-91} \log(MR_r)$							-0.349*** (0.046)	-0.559*** (0.065)	-0.158* (0.089)	-0.336*** (0.110)	
Other Accidents											
RTC_r	3.661*** (0.886)	6.820*** (0.960)	1.935*** (0.863)	3.771*** (1.340)	3.033*** (1.015)	4.705*** (1.123)	2.622*** (0.901)	4.320*** (1.056)	2.912*** (0.948)	4.605*** (1.082)	-0.866 (0.701)
$\Delta_{85-91} \log(MR_r)$							-0.488** (0.082)	-0.458*** (0.079)	-0.138 (0.101)	-0.0056 (0.111)	

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. In column 1, observations are not weighted; in column 2, observations are weighted by population; column 3 adds state fixed effects to column 2; column 4 adds pre-trends to column 3; column 5 shows two-stage least squares, with an instrument for $\Delta_{85-91} \log(MR_r)$. Column 6 presents a placebo test, with observations weighted by population and considering state fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

I Imports of Automotive Vehicles and Parts

Figure 21 describes the pronounced increase in imports of automotive vehicles and their parts in Brazil during the 2000s.



(d) Total imports of vehicles and parts: Log value

(e) Total imports of vehicles and parts: Index

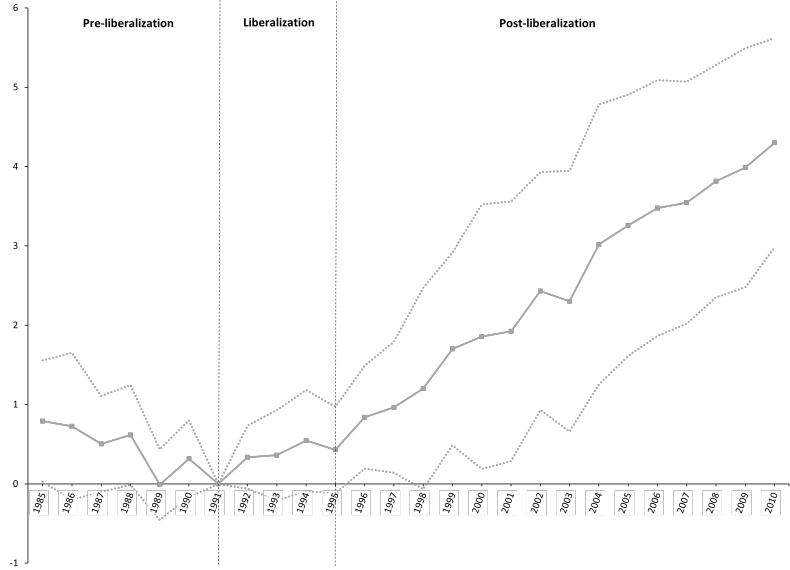
Figure 21: Trends in total imports of vehicles and parts: 1997-2010

Source: Brazilian foreign trade data based on the declaration of importers (SISCOMEX from the Ministry of Industry, Foreign Trade and Services) - Chapter 87 in the Harmonized System.

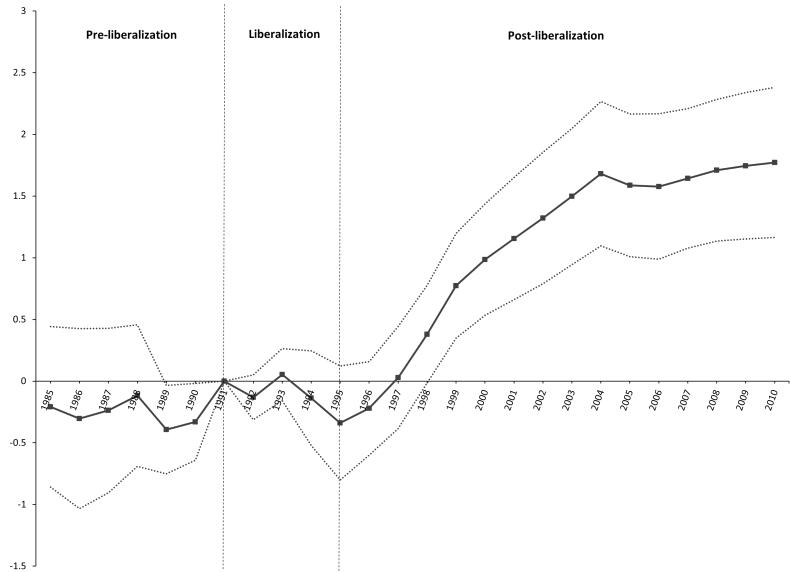
J Dynamic Impacts of Regional Tariff Changes in Cause-Specific Mortality Rates

In this section, we present additional visualizations of the dynamic effects of regional tariff shock on cause-specific mortality rates.

We disaggregate the dynamic impact of the trade-induced regional economic shock on specific causes of mortality in each of the large groups in Table 3 of the article – i.e. external and internal causes. First, Figure 22 offers graphical insights into the dynamic impact of tariff reductions on mortality rates attributed to external (Panel (a)) and internal causes (Panel (b)). Second, Figure 23 describes the dynamic effect of trade reform on cause-specific causes related to internal causes of mortality and Figure 24 disaggregates the impact on external causes of mortality into smaller cause-specific groups.



(a) External causes of mortality



(b) Internal causes of mortality

Figure 22: Dynamic effects of regional tariff changes on log changes of specific-causes mortality rates

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (2) in the article, where the dependent variable is the change in regional log cause-specific mortality rates and the independent variable is the RTC_r in $t = 1985, \dots, 2010$. All regressions include state fixed effects. Dashed lines show 95 percent confidence intervals. Standard errors are adjusted for 91 meso-region clusters.

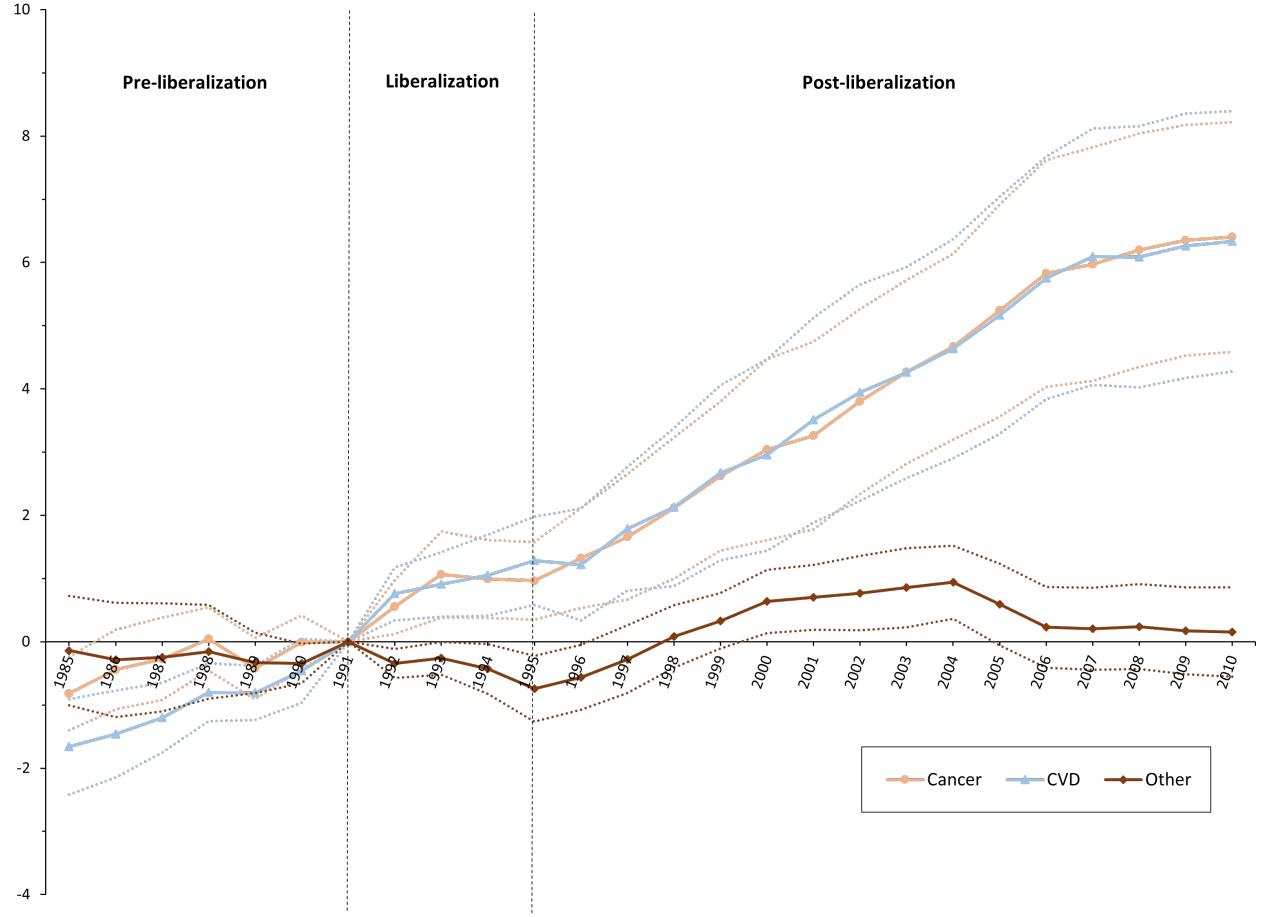


Figure 23: Dynamic effects of regional tariff changes on log changes in mortality rates from internal causes

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (2) in Manuscript (2023), where the dependent variable is the change in regional log (internal) cause-specific mortality rates and the independent variable is the RTC_r in $t = 1985, \dots, 2010$. All regressions include state fixed effects. Dashed lines show 95 percent confidence intervals. Standard errors are adjusted for 91 meso-region clusters.

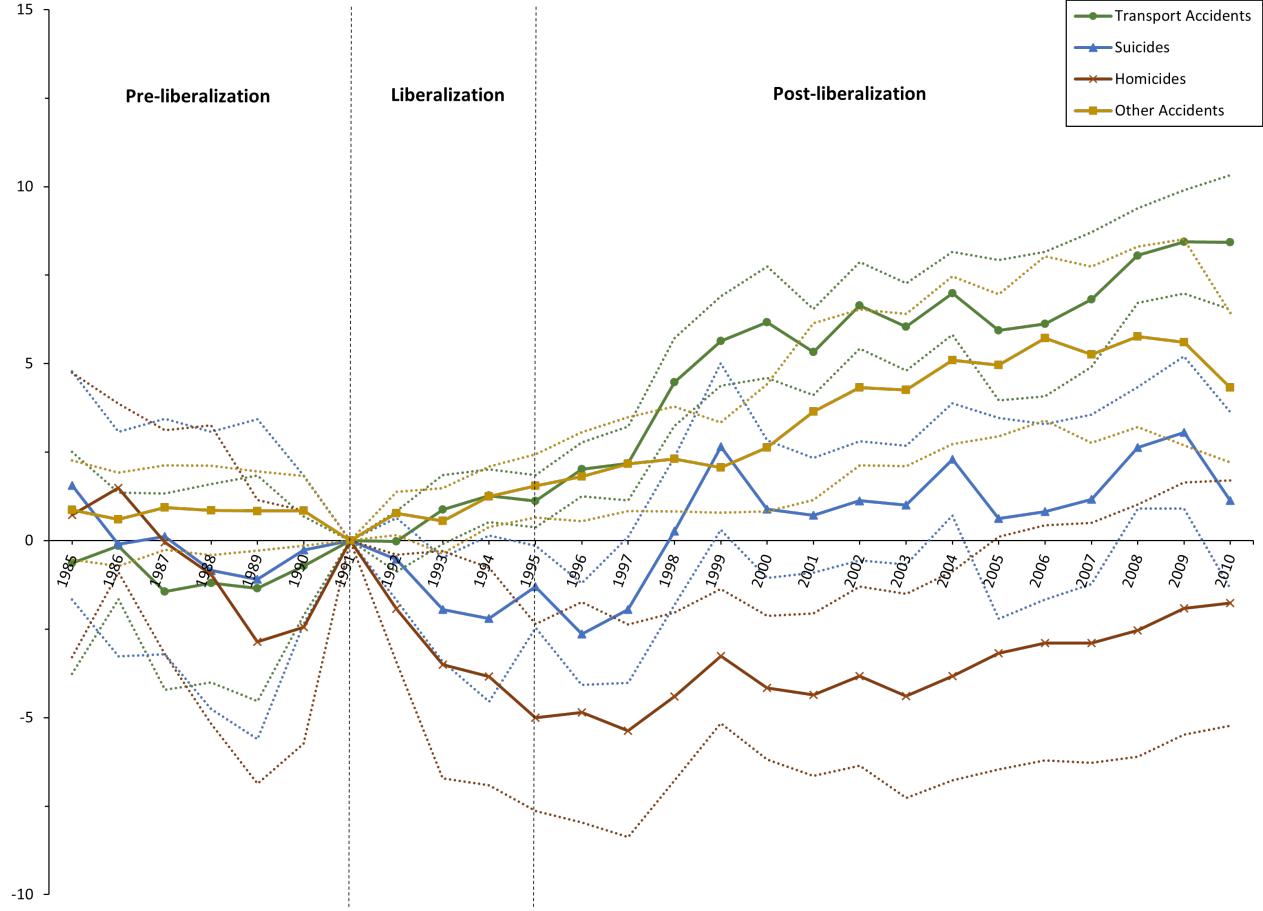


Figure 24: Dynamic effects of regional tariff changes on log changes in mortality rates from external causes

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (2) in Manuscript (2023), where the dependent variable is the change in regional log (external) cause-specific mortality rates and the independent variable is the RTC_r in $t = 1985, \dots, 2010$. All regressions include state fixed effects. Dashed lines show 95 percent confidence intervals. Standard errors are adjusted for 91 meso-region clusters.

K Attenuation and Direct Impacts

In this section, we delve into additional insights regarding potential mitigations of the shock's impact on mortality.

In particular, we begin by examining whether the inclusion of variables used in Table 6 of the article substantially alters the impact of regional tariff changes on mortality rates.

Table 11 presents the results focusing on the medium-run repercussions of the trade-induced regional economic shock on mortality rates. We evaluate the mechanism variables also in medium-run variations, between 1992 and 2000 (with the exceptions described in Table 6 of the article). In column (1), we estimate the main model outlined in the article and progressively introduce variables of interest as they appear in Table 6. Notably, the estimated coefficients demonstrate remarkable consistency across the board. However, it is crucial to highlight instances of coefficient attenuation observed under specific circumstances: i) when incorporating the variation of the share of H&S expenditures over total, ii) when factoring in the variation in expenditures within the outpatient system (SIA), and iii) when considering the variation in the ratio between non-basic and basic procedures within the outpatient system.

Table 11: Regional tariff changes and log changes in mortality rates - First-difference main results

<i>Dep. Variable: $\Delta \log(MR)$</i>	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
RTC	1.099*** (0.240)	1.104*** (0.240)	1.143*** (0.249)	1.137*** (0.248)	1.252*** (0.227)	1.242*** (0.220)	1.256*** (0.227)	1.249*** (0.229)	1.193*** (0.225)	1.234*** (0.251)
Exp. per function - Total	-2.59e-06** (1.14e-06)	-1.50e-05 (1.01e-05)	-1.82e-05 (2.66e-05)	-1.43e-05 (2.49e-05)	-1.55e-05 (2.76e-05)	-1.82e-05 (2.75e-05)	-1.74e-05 (2.79e-05)	-1.07e-05 (2.79e-05)	-7.09e-06 (2.74e-05)	
Exp. H&S		0.000107 (9.02e-05)	0.000134 (0.000227)	0.000111 (0.000213)	0.000122 (0.000238)	0.000146 (0.000236)	0.000137 (0.000240)	8.40e-05 (0.000240)	5.17e-05 (0.000235)	
Share H&S/Total		-0.0363 (0.327)	0.00576 (0.307)	0.00193 (0.315)	-0.00593 (0.319)	-0.0220 (0.325)	0.0257 (0.325)	0.0681 (0.329)		
Exp. SIH			0.00486*** (0.00184)	0.00491*** (0.00185)	0.00522*** (0.00194)	0.00565*** (0.00186)	0.00573*** (0.00192)	0.00558*** (0.00198)		
Exp. SIA				-0.000152 (0.000759)	-0.000255 (0.000775)	-0.000294 (0.000776)	-8.82e-05 (0.000743)	-0.000105 (0.000720)		
Medical Est.					0.0659 (0.0515)	0.0682 (0.0537)	0.0908* (0.0526)	0.0922 (0.0607)		
Hosp. Rate						-0.0187 (0.0192)	-0.0148 (0.0186)	-0.0166 (0.0212)		
Non-Basic/Basic Procedures							-0.0328*** (0.00985)	-0.0284** (0.0139)		
Outpatient Procedures Rate								0.00289 (0.00788)		
Diagnoses Rate - Neoplasm Detection									0.00289 (0.00788)	

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. All observations are weighted by population and state-year fixed effects are considered. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Moving to Table 12, we undertake a similar analysis but with a focus on long-term variations in regional log mortality rates, incorporating the mechanism variables of interest as long-term variations (between 1992 and 2010, with the exceptions highlighted in Table 6 of the article). Despite the sustained consistency of coefficients, a notable reduction in

the magnitude of the estimated coefficient associated with RTC_r is discernible between columns (3) and (4) upon including the long difference in the share of H&S expenditure over total. This underscores the significance of such a measure as a potential mechanism driving observed mortality reductions, particularly in the long run. Furthermore, we observe a minor reduction or attenuation in the magnitude of the coefficient of interest upon including the diagnosis rate associated with neoplasm detection, another pivotal metric capturing the relative expansion of capital-intensive healthcare infrastructure in regions most affected by tariff shocks over the long term.

Table 12: Attenuation results - Long-run

<i>Dep. Variable:</i> $\Delta \log(MR)$	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
RTC	1.957*** (0.278)	2.039*** (0.313)	2.555*** (0.458)	1.988*** (0.325)	2.152*** (0.296)	2.153*** (0.296)	2.203*** (0.321)	2.210*** (0.322)	2.217*** (0.317)	2.124*** (0.340)
Exp. per function - Total	-2.47e-05 (3.61e-05)	-2.02e-05 (5.82e-05)	4.79e-05 (4.76e-05)	3.68e-05 (4.40e-05)	3.62e-05 (4.43e-05)	3.72e-05 (4.45e-05)	4.58e-05 (4.45e-05)	4.53e-05 (4.62e-05)	4.53e-05 (4.59e-05)	4.95e-05 (4.82e-05)
Exp. H&S		-1.06e-05 (0.000138)	-0.000273 (0.000186)	-0.000273 (0.000181)	-0.000265 (0.000178)	-0.000280 (0.000182)	-0.000283 (0.000183)	-0.000283 (0.000182)	-0.000283 (0.000182)	-0.000311* (0.000174)
Share H&S/Total			0.253 (0.243)	0.183 (0.244)	0.179 (0.242)	0.177 (0.239)	0.195 (0.243)	0.197 (0.241)	0.226 (0.233)	
Exp. SIH				0.000584* (0.000342)	0.000498 (0.000429)	0.000394 (0.000507)	0.000357 (0.000505)	0.000366 (0.000513)	0.000399 (0.000506)	
Exp. SIA					0.000204 (0.000430)	0.000189 (0.000406)	0.000181 (0.000404)	0.000148 (0.000438)	0.000165 (0.000479)	
Medical Est.						0.0296 (0.0451)	0.0328 (0.0449)	0.0323 (0.0451)	0.0258 (0.0465)	
Hosp. Rate							0.0161 (0.0196)	0.0154 (0.0190)	0.0229 (0.0218)	
Non-Basic/Basic Procedures								0.00384 (0.0105)	-0.00121 (0.0150)	
Outpatient Procedures Rate										-0.00911 (0.0112)
Diagnoses Rate - Neoplasm Detection										

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. All observations are weighted by population and state-year fixed effects are considered. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Finally, while Table 6 of the article elucidates the relationship between RTC and variables capturing the operation of our mechanism, we now pivot to examining the direct relationship between such variables and variations in regional mortality rates, both in the medium and long term, in Table 13. In column (1), we consider the medium-run variations in such variables as the independent variables and the medium-run variation in regional log mortality rates as the dependent variable. Column (2) presents a similar analysis for the long-run variations in all variables.

In short, our analysis reveals significant insights into the long-term dynamics, particularly regarding the pivotal role of capital-intensive healthcare infrastructure expansion in mitigating mortality rates attributed to specific internal causes. Notably, our findings underscore the negative coefficients associated with various healthcare metrics, including health expenditure, the proportion of expenditure on H&S/total, expenditures in the hospital system, the number of medical establishments, the ratio between basic and non-basic procedures within the outpatient system, and the diagnosis rate associated with neoplasm detection in the SUS hospital system. These results bolster the argument that these variables serve as potential mechanisms driving the observed relative reduction in mortality rates in the regions more impacted by the trade shock in the long run. Of particular significance are the statistically significant negative coefficients linked to health expenditure and the diagnosis rate, further illuminating their roles in influencing long-term mortality trends (particularly from internal causes).

L Additional Mechanisms

L.1 Air Pollution

In this section, we assess the potential impact of the regional tariff changes on air quality. Extensive research has shown a strong association between air pollution and adverse health outcomes, particularly increased mortality risks from cardiovascular and respiratory diseases (Clancy et al., 2002; Graff Zivin & Neidell, 2013; Heutel & Ruhm, 2016; Mustafić et al., 2012). Infants are particularly vulnerable to the effects of air pollution, and numerous studies have documented its impact on infant mortality (Arceo et al., 2016; Chay & Greenstone, 2003; Currie & Neidell, 2005; Currie et al., 2009, 2014; Knittel et al., 2016).³

While the existing literature regarding the effect of air quality on the pro-cyclical nature of overall mortality rates is relatively limited, a noteworthy exception is the study conducted by Heutel and Ruhm (2016), which identifies a significant positive correlation between pollution concentrations and overall mortality rates. Their research demonstrates that controlling for carbon monoxide, particulate matter, and ozone considerably weakens the relationship between overall mortality and the unemployment rate.

Motivated by these insights, we proceed to examine the impact of the regional tariff shock

³However, it is worth mentioning that Charris et al. (2023) only find a minimal impact of tariff cuts on carbon monoxide emissions two decades after liberalization, suggesting that air quality may have played a minor role in reducing infant mortality in the regions more exposed to the trade-induced economic shock.

Table 13: Direct effects - mechanism variables and mortality rates

Dep. Variable: $\Delta \log(MR)$	Medium-run	Long-run
	(1)	(2)
Exp. per function -		
Total	-1.84e-05 (3.29e-05)	0.000213** (0.000103)
Exp. H&S	0.000165 (0.000284)	-0.000477* (0.000262)
Share H&S/Total	-0.260 (0.363)	-0.203 (0.283)
Exp. SIH	0.00349 (0.00215)	-0.000412 (0.000617)
Exp. SIA	-0.00106 (0.000925)	0.000262 (0.000566)
Medical Est.	0.0838 (0.0733)	-0.0207 (0.0490)
Hosp. Rate	-0.0173 (0.0207)	0.0214 (0.0275)
Non-Basic/Basic Procedures	-0.0403* (0.0216)	-0.0127 (0.0205)
Outpatient Procedures Rate		
Diagnoses Rate -		
Neoplasm Detection	-0.0167* (0.00888)	-0.0247* (0.0136)

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. All observations are weighted by population and state-year fixed effects are considered. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

on air quality in the local economies of Brazil, with a specific focus on the concentration of particulate matter. We employ data on global air pollutant estimates, specifically particulate matter (PM_{10}), from the Emissions Database for Global Atmospheric Research (EDGAR). We construct municipality-year concentration levels of PM_{10} , aggregating up from grid-level estimates to examine the potential role of air pollution in explaining the observed mortality patterns. Our analysis involves estimating Equation 2 of the manuscript, utilizing the log changes in PM_{10} during the periods of 1991-2000 and 1991-2010 as dependent variables. Table 14 summarizes our results.

We find statistically significant reductions in particulate matter concentration both in

Table 14: Other factors

Panel A. Air pollution			
	PM ₁₀		
	1991-2000	1991-2010	Placebo 1985-1991
	(1)	(2)	(3)
RTC_r	0.817*** (0.173)	1.298*** (0.401)	0.0148 (0.300)
R^2	0.282	0.279	0.422

Notes: Data from the Emissions Database for Global Atmospheric Research (EDGAR). All left-hand-side variables are given by the changes of logs over the indicated period. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. There are 411 micro-region observation. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

the medium and long term, although the estimated effects have very limited magnitudes. A decrease of -0.1 log points in RTC_r corresponds to an approximately 0.8% reduction in particulate matter between 1991 and 2000. The influence of trade-induced economic shock on air quality shows a slight increase over time, as shown in column 2. In short, while improvements in air quality seem to contribute to the effects of trade-induced economic shock on mortality rates from internal causes, the evidence suggests that these effects are relatively modest and not the primary driving force behind the overall impacts.

L.2 Opportunity Cost of Medical Care

Next, we turn our attention to another potential mechanism identified in the health economics literature: the opportunity cost of engaging in health-improving activities. Ruhm (2000) proposes that during periods of economic expansion, leisure time decreases, leading to a rise in the opportunity cost of engaging in activities such as physical exercise. Ruhm (2005) provides empirical evidence supporting this mechanism. Unfortunately, data specifically pertaining to this mechanism in Brazil during the period of interest are not available. However, we posit that the logic of opportunity cost of time may play a role in health care utilization within the Brazilian context.

In Brazil, where healthcare coverage is nearly universal, the influence of employer-provided

insurance or coverage rates on healthcare utilization during economic downturns is not as pronounced.⁴ Instead, the cost of health care utilization is primarily associated with the time required to schedule medical appointments and procedures. While specific municipality-level data on healthcare utilization are lacking for the early 1990s, we utilize mortality rates from medical complications as a proxy for the demand for medical care. Intuitively, if the demand for medical care and surgical procedures increases during economic downturns, due to a lower opportunity cost, we expect an increase in medical care-related mortality. Our findings, summarized in Table 15, suggest that medical care-related mortality rates increased in the regions more exposed to the tariff cuts, partially supporting the notion of higher healthcare utilization in the local economies more negatively impacted by the liberalization episode.

Table 15: Regional Tariff Changes and log changes in “medical” mortality rates

	OLS (1)		OLS (2)		OLS (3)		OLS (4)		2SLS (5)		Placebo (6)
Dep. var.: $\Delta \log(MMR_r)$	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	1985-1991
RTC_r	-3.257*** (0.944)	-1.279 (1.281)	-2.930* (1.624)	-3.083 (2.018)	-6.520*** (1.543)	-4.389** (2.126)	-7.162*** (1.613)	-4.768** (2.070)	-6.075*** (1.558)	-4.189** (2.001)	-1.668 (1.316)
$\Delta_{85-91} \log(MR_r)$							-0.383*** (0.121)	-0.225** (0.107)	0.265 (0.301)	0.119 (0.208)	

Notes: There are 408 micro-region - the unit of analysis - observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. In column 1, observations are not weighted; in column 2, observations are weighted by population; column 3 adds state fixed effects to column 2; column 4 adds pre-trends to column 3; column 5 shows two-stage least squares, with an instrument for $\Delta_{85-91} \log(MMR_r)$. Column 6 presents a placebo test, with observations weighted by population and considering state fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

L.3 Healthcare personnel

In this subsection, we explore a potential mechanism related to the proportion of workers employed in medical care jobs. In Table 16, we examine the potential impact of the economic shock on the proportion of health professionals as a total of all workers in each micro-region (per 100,000 inhabitants).

⁴This mechanism is explored, for instance, in Ruhm (2003) and Lang et al. (2019) in the US context.

Table 16: Regional tariff changes and health personnel

	Dependent variable: Δ Share of Medical Care Personnel					
	Health Personnel		Only Physicians		Only Nurses	
	1991-2000 (1)	1991-2010 (2)	1991-2000 (1)	1991-2010 (2)	1991-2000 (1)	1991-2010 (2)
RTC_r	-1.452** (0.650)	1.141** (0.481)	1.283*** (0.444)	5.097*** (0.502)	-1.838** (0.850)	1.082 (0.632)

Notes: All left-hand-side variables in Panels A and B are given by the changes of logs of the variables measured in per capita terms over the indicated period. The expenditure variables in Panel C are measured in per capita changes. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. There are 411 micro-region observations. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

The regions that were relatively more impacted by the tariff reductions observed a relative increase in the share of health professionals in the medium run, although this effect is reversed in the long run. Breaking down this effect, our results indicate that the increase in the share of health professionals is due to an increase in the number of nurses in the medium run, representing a substitution relative to physicians that faced a statistically significant relative reduction in the regions more impacted by the regional economic shock. In the long term, this effect of increasing the number of nurses in those regions loses statistical significance, although the effect of the reduction in the number of physicians intensifies.

Overall, we observe a relative substitution of physicians for nurses in harder-hit regions in the decades after the trade shock, which, alongside the mortality reduction findings, might indicate greater efficiency in the production function of healthcare. This result is partially in line with Carrillo and Feres (2019) for a recent period in Brazil, evaluating the limited impacts of the expanded supply of primary care physicians (*More Physicians Program*) in infant mortality.

L.4 Educational outcomes

Next, we delve into the relationship between trade liberalization, regional education outcomes, and mortality rates in the following paragraphs.

In fact, one might consider a situation in which the regions most affected by the trade shock may have experienced a relative increase in education expenditures (also due to the partially obligatory nature of the expenditure in this category) which, in turn, may have led to better educational levels on average for the population of such regions comparatively to those least affected by the regional economic shock. This higher educational level over time may be related to the reduction in mortality observed in the hardest hit regions through two

main channels: i) reduction in mortality from internal causes due to less harmful consumption habits (for example, with lower incidence smoking and obesity)⁵ and, ii) through a more direct search for medical help and a greater understanding of the steps and needs of possible treatments (Clark & Royer, 2013; Conti et al., 2010; Cutler & Lleras-Muney, 2010; Davies et al., 2018; Lager & Torssander, 2012; Lleras-Muney, 2005).

To investigate this avenue, we utilize comprehensive datasets on government spending and educational attainment at the regional level. With regard to government spending, as discussed earlier in this response letter, we constructed measures of total regional per capita spending (across all categories) and on education and culture (Henceforth, E&C) by computing annual government spending per category at the municipality level with data from the Ministry of Finance (*Ministério da Fazenda - Secretaria do Tesouro Nacional*). We use government expenditures by function since 1995 and aggregate them to the micro-region level. Besides, we use data from the Atlas of Human Development, constructed from the Brazilian Decennial Population Census microdata, for variables related to educational attainment at the municipality level. In particular, we use measures for the share of 25+ year-old individuals that have completed middle school, high school, and college education, the illiteracy rate for individuals above 15 years old, and the expected years of education of individuals at 18 years old. Similar to Dix-Carneiro et al. (2018), we also compute the share of youth aged 14–18 out of school (high school dropouts) from census data.

After describing the data, we now delve into evaluating the impacts of the trade-induced regional economic shock on various indicators related to regional educational expenditures and education outcomes. The outcomes of these analyses are summarized in Table 17.

⁵Notwithstanding, it is worth indicating that the causal evidence in the literature regarding this channel is mixed. Galama et al. (2018) indicate that there is no convincing evidence of an effect of education on obesity, and the effects on smoking are only apparent when schooling reforms affect individuals' track or their peer group, but not when they simply increase the duration of schooling.

Table 17: Potential mechanisms: regional tariff change and education outcomes

Panel A. Educational expenditure							
	All categories		E&C		Share E&C / Total		
	1995-2000	1995-2010	1995-2000	1995-2010	1995-2000	1995-2010	
	(1)	(2)	(1)	(2)	(1)	(2)	
RTC_r	1,774	3,310***	898.6**	1,162***	0.512***	0.0688	
	(1,553)	(569.3)	(372.4)	(182.1)	(0.0806)	(0.183)	

Panel B. Educational attainment							
	Middle-school		High-school		College		
	1995-2000	1995-2010	1995-2000	1995-2010	1995-2000	1995-2010	
	(1)	(2)	(1)	(2)	(1)	(2)	
RTC_r	0.982***	2.805***	0.538***	1.833***	0.975***	4.213***	
	(0.185)	(0.397)	(0.164)	(0.419)	(0.261)	(0.742)	

Panel C. Additional measures							
	Illiteracy rate		Expected years of education		High school dropouts		
	1991-2000	1991-2010	1991-2000	1991-2010	1991-2000	1991-2010	
	(1)	(2)	(1)	(2)	(1)	(2)	
RTC_r	0.792***	1.822***	0.184	1.007***	-0.354*	-2.397***	
	(0.0678)	(0.154)	(0.160)	(0.196)	(0.200)	(0.291)	

Notes: The expenditure variables in Panel A measured in per capita changes. In Panels B and C, the variations are given by the changes of logs of the variables measured in per capita terms over the indicated period. Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. There are 411 micro-region observations in the estimations of all panels, except for three to four missing values in government spending. Observations are weighted by population. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

In Panel A of Table 17, we describe the impact of the tariff cuts on government spending in the micro-regions. Similar to the findings in the article, we document an overall reduction in government expenditures in the regions more impacted by the trade-induced shock in the long run. Looking at spending on E&C, there also appears to be a clear decrease in spending on education in the regions more exposed to the tariff shock compared to lesser hit regions, with statistically significant impacts both in the medium and long run. Regarding the ratio of expenditures on E&C to total expenditures, we observe a statistically significant decrease in the medium run - indicating a relative reduction in education expenditures as a share of total expenditures in harder-hit locations -, with no significant impact in the long run. Interestingly, the relative increase in health expenditure documented in the original

manuscript may be the flip side of this relative decrease observed for education spending.

Turning to educational attainment, Panel B highlights a concerning trend of diminishing shares of the population obtaining middle school, high school, and college degrees comparatively in heavily impacted regions. While the magnitude of these reductions may appear limited in the short term, there is a notable amplification of these effects in the long term, indicating persistent challenges in educational attainment post-trade liberalization. Taken together, this evidence indicates contrary to the hypothesis that the regions most affected by the shock are obtaining relatively better educational outcomes than the local economies less affected after trade liberalization.

Finally, in Panel C, we examine the impact of the trade-induced regional economic shock on variables that represent additional education outcomes at the regional level. Note that, similar to the results in Panel B, our findings indicate a relative worsening in the illiteracy rate (the portion of the population aged 15+ that is illiterate) and in the expected years of education at age 18 in the regions most affected by the shock, especially in the long term. The last column, similarly to Dix-Carneiro et al. (2018), shows that the share of youth aged 14–18 out of school (high school dropouts) experience relative deterioration in regions facing larger tariff shocks in both the medium and long run.

It is important to highlight that there is evidence in the literature (although still limited) pointing in the same direction as the findings presented in Table 17. For instance, Mariano (2023) presents evidence that the micro-regions more affected by the tariff cuts experienced a net increase in dropout rates from higher education relative to less strongly affected regions. The author shows that a one-standard-deviation increase in tariff cuts increases the dropout rates by approximately 1% in the medium run and 3% in the long run.

In conclusion, while increased education expenditures and improved educational outcomes are often seen as catalysts for better health outcomes, our analysis suggests a contrary trend in regions most affected by tariff cuts. The evidence points to a lasting decline in educational investments and attainment, which may hinder efforts to reduce mortality rates in these regions.

M Robustness of the Mechanism Exploration

The reason for not including placebo tests (or parallel trends tests) and estimations with pre-trends for the dependent variables presented in Table 6 of the article is, unfortunately, due to data constraints for periods before trade liberalization in Brazil. Below, we provide a detailed explanation of the data limitations for each category analyzed in the panels of

Table 6.

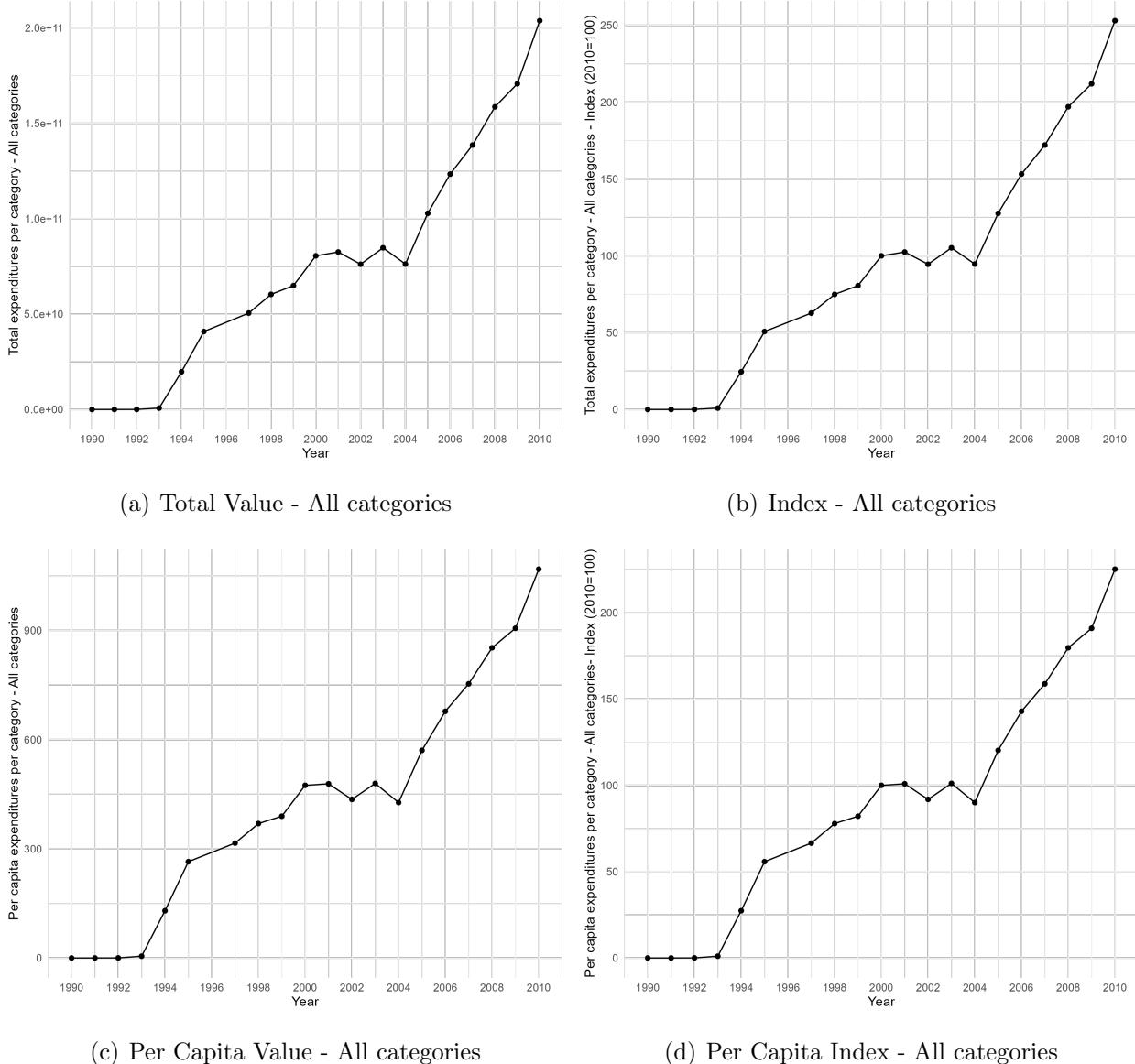
M.1 Panel A

First, in Panel A of Table 6, referring to local government expenditures, note that the start period of the three variables of interest - total expenditures per function, expenditures in H&S, and, the share of H&S over total expenditures per category – is 1995, the last year of the trade reform. For clarification, we constructed these variables by computing annual government spending per category at the municipality level with data from the Ministry of Finance (*Ministério da Fazenda - Secretaria do Tesouro Nacional*). Although the Ministry of Finance provides data on these expenditures at the municipality level (used to construct the measures per micro-region) dating back to 1990, this timeframe is insufficient to construct placebo tests. Our approach, as outlined for instance in column 6 of Table 1 of the manuscript, involves utilizing the (log) difference of the dependent variable — overall mortality rate — between 1985 and 1991 to establish pre-trends. This enables us to demonstrate the lack of correlation between regional tariff changes and pre-liberalization changes in mortality rates.

Moreover, using local government expenditure data from the late 1980s and early 1990s in Brazil presents additional challenges. As highlighted in Dix-Carneiro et al. (2018, p. 169), local government expenditure data dating back to 1990 is often unreliable, partly because of i) measurement error due to hyperinflation, and, ii) frequent missing information. We illustrate these challenges with the below visualizations.

First, it is worth highlighting the large variation in the total value of disaggregated government expenditures at the municipality level in the first years of the series. This is arguably due, as discussed above, to measurement error in the hyperinflation years. In Figure 25, we show the evolution of total expenditures per function in all categories for municipalities in Brazil, aggregated to the national level. In panels (a) and (b) we show, respectively, the total value of expenditures over time and an index, based on 2000, for better comparison. In panels (c) and (d), we present the same visualizations but for the total government expenditure per capita and its index, also based on 2000. Note that it is from 1994 (the year that marks monetary stabilization in the country) and 1995 that we have comparable observations throughout the extensive period of our analysis.

Figure 25: Total Expenditures per Category - All categories - Brazil aggregated from municipalities

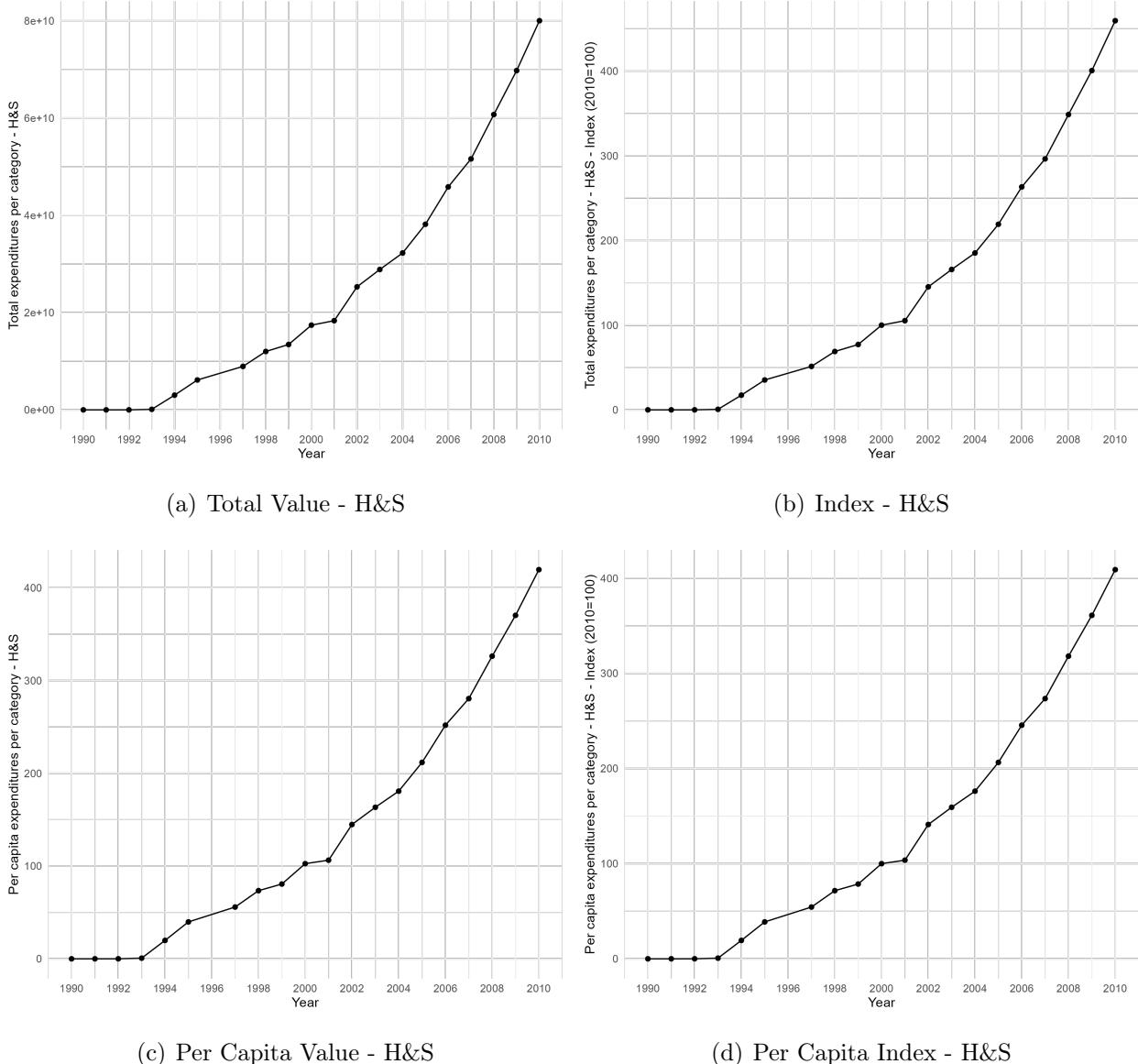


Source: Annual government spending per category at the municipality level is from the Ministry of Finance (*STN - Secretaria do Tesouro Nacional*).

This same trend is observed when focusing on municipal spending on health and sanitation. In Figure 26, we show the evolution of municipality-level government expenditures in the H&S category, aggregated at the national level, over time. Again, panels (a) and (b)

present, respectively, the total value of government expenditure on H&S and the value index based on 2000. Panels (c) and (d), in turn, present the value and index of this government expenditure category in per capita terms.

Figure 26: Total Expenditures per Category - H&S- Brazil aggregated from municipalities



Source: Annual government spending per category at the municipality level is from the Ministry of Finance (*STN - Secretaria do Tesouro Nacional*).

Second, regarding frequent missing information, Figure 27 illustrates the frequency of

missing municipalities per year in the municipality-level disaggregated government expenditure data.⁶ The dashed lines in red describe the beginning and end of the trade reform. Notably, 1995 has the least missing data between the treatment years. With that in mind, we focus on data after Brazil stabilized its currency, that is, from 1994 onwards (the year of the implementation of *Plano Real*), and with a more reliable number of observations.

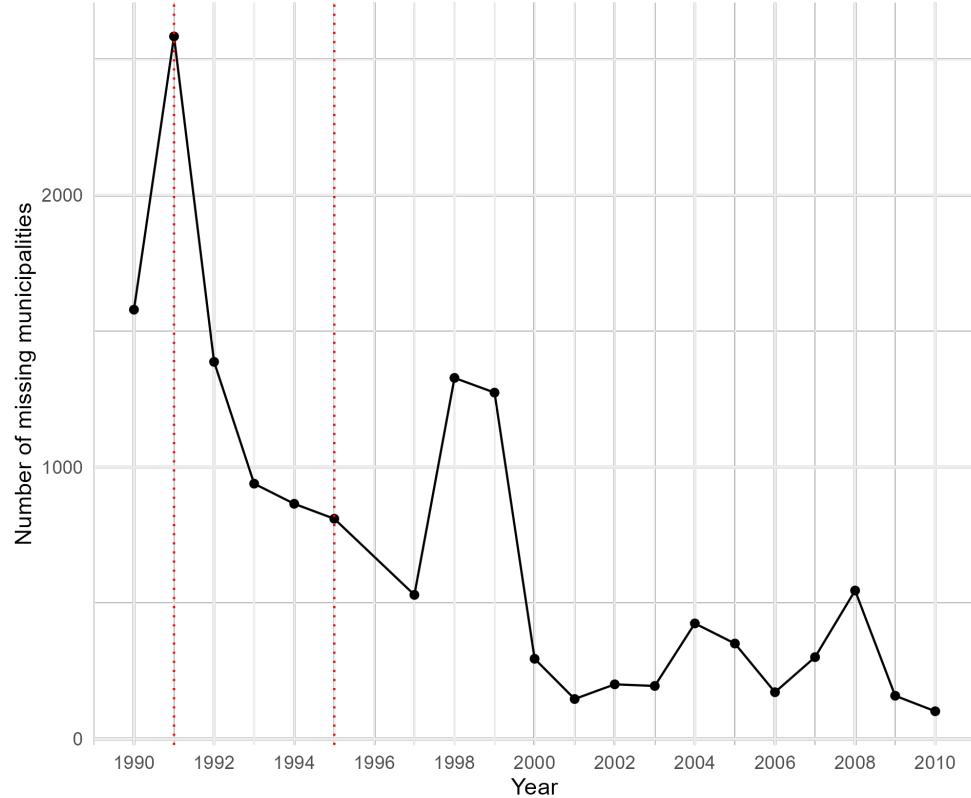


Figure 27: Number of missing municipalities per year - Total expenditures per category

Source: Annual government spending per category at the municipality level is from the Ministry of Finance (*STN - Secretaria do Tesouro Nacional*).

M.2 Panels B and C

In Panels B and C, we delve into the expenditures and production of healthcare infrastructure, specifically focusing on hospital and outpatient systems (using data from *SIH - Sistema de Informações Hospitalares do SUS* and *SIA - Sistema de Informações Ambulatoriais do SUS*, respectively) as well as the total number of medical establishments.

⁶For context, the number of municipalities in the 2010 Census in Brazil was slightly above 5500.

First, it is worth highlighting that consolidated data on hospital production by place of hospitalization (thus, at the municipality level) are only available from 1992 onwards. Therefore, for a similar reason to that discussed previously for Panel A of Table 6, it is not possible to present the parallel trends or placebo test for such dependent variables, as well as the inclusion of pre-trends, since the first available observation is already within the treatment period - i.e. the trade liberalization episode. This is why, in the first column of Panel C, we use 1992 as the base year. Furthermore, for reasons similar to those discussed concerning Panel A, in the first column of Panel B, we use the year 1995 as a base because it is a monetary variable.

The restriction is similar to the outpatient system, in which consolidated data on outpatient production by place of care (municipality level) is only available from mid-1994 onwards. This is the reason why we used 1995 as a base in the second column of Panel B and the last two columns of Panel C, and were unable to present placebo tests for such dependent variables.

The only potentially distinct case concerns the last column of Panel B, referring to the total number of medical establishments. For clarification, we computed the number of health establishments at the municipality level - and, later, aggregated to micro-regions - from two data sources: i) *Pesquisa de Assistência Médico-Sanitária* (1992, 1999, 2002), and, ii) *Cadastro Nacional de Estabelecimentos de Saúde* (2005-2010), both available at DataSUS (administrative dataset from the Ministry of Health). For reasons similar to those discussed previously, we chose to establish the year 1992 as the basis for treatment. However, it is important to highlight that the *Pesquisa de Assistência Médico-Sanitária* was carried out continuously between 1981 and 1990 in Brazil. Despite the slight methodological change in measuring medical establishments between the late 1980s and 1992, we present detailed estimates in Table 18.

We adopt a comprehensive approach, incrementally saturating the model to ensure robustness. The magnitude of the coefficient decreases marginally but remains statistically non-significant for the medium-run and significant for the long run with the incremental saturation of the model, involving the weighting of the observations by the average population between 1991 and 2000 - for the medium-run - and, 1991 and 2010, for the long-run (column 2), the inclusion of state fixed effects (column 3) and a variable capturing the pre-period trend in total medical establishment rates (columns 4 and 5).

Furthermore, Column 6 provides evidence that regional tariff changes are uncorrelated with pre-trends by regressing pre-liberalization changes in medical establishments directly

against future trade shocks. The non-significance of the coefficients in our placebo test corroborates the previous evidence obtained with the inclusion of trend variables in the estimations.

Table 18: Regional tariff changes and log changes total medical establishment rates

	OLS (1)		OLS (2)		OLS (3)		OLS (4)		2SLS (5)		Placebo (6)
<i>Dep. var:</i> $\Delta \log(\text{TotalMedEst}_r)$	1992-1999	1992-2010	1992-1999	1992-2010	1992-1999	1992-2010	1992-1999	1992-2010	1992-1999	1992-2010	1985-1992
RTC_r	1.775*** (0.415)	-5.188*** (0.754)	0.954** (0.467)	-3.497*** (0.989)	-0.157 (0.686)	-3.262*** (0.583)	-0.316 (0.619)	-3.326*** (0.594)	-0.223 (0.618)	-3.326*** (0.545)	-0.619 (0.563)
$\Delta_{85-92} \log(\text{TotalMedEst}_r)$							-0.259** (0.108)	-0.102 (0.125)	-0.107 (0.0679)	-0.102 (0.130)	

Notes: There are 411 micro-region observations. Standard errors (in parentheses) are adjusted for 91 meso-region clusters. Unit of analysis r is a micro-region. In column 1, observations are not weighted; in column 2, observations are weighted by population; column 3 adds state fixed effects to column 2; column 4 adds pre-trends to column 3; column 5 shows two-stage least squares, with an instrument for $\Delta_{85-91} \log(\text{TotalMedEst}_r)$. Column 6 presents a placebo test, with observations weighted by population and considering state fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

References

- Arceo, E., Hanna, R., & Oliva, P. (2016). Does the effect of pollution on infant mortality differ between developing and developed countries? evidence from mexico city. *The Economic Journal*, 126(591), 257–280.
- Borusyak, K., Hull, P., & Jaravel, X. (2022). Quasi-experimental shift-share research designs. *The Review of Economic Studies*, 89(1), 181–213.
- Braga, B. (2018). The effects of unilateral trade liberalization on marriage and fertility choices: Evidence from brazil. Available at SSRN 3161625.
- Carrillo, B., & Feres, J. (2019). Provider supply, utilization, and infant health: Evidence from a physician distribution policy. *American Economic Journal: Economic Policy*, 11(3), 156–96.
- Charris, C., Branco, D., & Carrillo, B. (2023). Economic shocks and infant health: Evidence from a trade reform in brazil. *Journal of Development Economics*, 166.
- Chay, K. Y., & Greenstone, M. (2003). The impact of air pollution on infant mortality: Evidence from geographic variation in pollution shocks induced by a recession. *The Quarterly Journal of Economics*, 118(3), 1121–1167.
- Clancy, L., Goodman, P., Sinclair, H., & Dockery, D. W. (2002). Effect of air-pollution control on death rates in Dublin, Ireland: An intervention study. *The Lancet*, 360(9341), 1210–1214.

- Clark, D., & Royer, H. (2013). The effect of education on adult mortality and health: Evidence from britain. *American Economic Review*, 103(6), 2087–2120.
- Conti, G., Heckman, J., & Urzua, S. (2010). The education-health gradient. *American economic review*, 100(2), 234–238.
- Currie, J., & Neidell, M. (2005). Air pollution and infant health: What can we learn from california's recent experience? *The Quarterly Journal of Economics*, 120(3), 1003–1030.
- Currie, J., Neidell, M., & Schmieder, J. F. (2009). Air pollution and infant health: Lessons from new jersey. *Journal of Health Economics*, 28(3), 688–703.
- Currie, J., Zivin, J. G., Mullins, J., & Neidell, M. (2014). What do we know about short-and long-term effects of early-life exposure to pollution? *Annu. Rev. Resour. Econ.*, 6(1), 217–247.
- Cutler, D. M., & Lleras-Muney, A. (2010). Understanding differences in health behaviors by education. *Journal of health economics*, 29(1), 1–28.
- Davies, N. M., Dickson, M., Davey Smith, G., Van Den Berg, G. J., & Windmeijer, F. (2018). The causal effects of education on health outcomes in the uk biobank. *Nature human behaviour*, 2(2), 117–125.
- Dix-Carneiro, R., & Kovak, B. K. (2017). Trade liberalization and regional dynamics. *American Economic Review*, 107(10), 2908–46.
- Dix-Carneiro, R., Soares, R. R., & Ulyssea, G. (2018). Economic shocks and crime: Evidence from the brazilian trade liberalization. *American Economic Journal: Applied Economics*, 10(4), 158–95.
- Gaddis, I., & Pieters, J. (2017). The gendered labor market impacts of trade liberalization evidence from brazil. *Journal of Human Resources*, 52(2), 457–490.
- Galama, T. J., Lleras-Muney, A., & Van Kippersluis, H. (2018). *The effect of education on health and mortality: A review of experimental and quasi-experimental evidence*. (tech. rep.). National Bureau of Economic Research.
- Graff Zivin, J., & Neidell, M. (2013). Environment, health, and human capital. *Journal of Economic Literature*, 51(3), 689–730.
- Heutel, G., & Ruhm, C. (2016). Air pollution and procyclical mortality. *Journal of the Association of Environmental and Resource Economists*, 3(3), 667–706.
- Knittel, C. R., Miller, D. L., & Sanders, N. J. (2016). Caution, drivers! children present: Traffic, pollution, and infant health. *Review of Economics and Statistics*, 98(2), 350–366.

- Kume, H., Piani, G., & Souza, C. F. B. (2003). A política brasileira de importação no período 1987-98: Descrição e avaliação. In K. Corseuil & H. Kume (Eds.), *A abertura comercial brasileira nos anos 1990: Impactos sobre emprego e salário*.
- Lager, A. C. J., & Torssander, J. (2012). Causal effect of education on mortality in a quasi-experiment on 1.2 million swedes. *Proceedings of the National Academy of Sciences*, 109(22), 8461–8466.
- Lang, M., McManus, T. C., & Schaur, G. (2019). The effects of import competition on health in the local economy. *Health Economics*, 28(1), 44–56.
- Lleras-Muney, A. (2005). The relationship between education and adult mortality in the united states. *The Review of Economic Studies*, 72(1), 189–221.
- Manuscript. (2023). *Trade liberalization and mortality rates: Evidence of pro-cyclical mortality from brazil* (tech. rep.).
- Mariano, L. (2023). *Dropping out of higher education after trade liberalization* (Working Paper). FGV EESP.
- Mustafić, H., Jabre, P., Caussin, C., Murad, M. H., Escolano, S., Tafflet, M., Périer, M.-C., Marijon, E., Vernerey, D., Empana, J.-P., et al. (2012). Main air pollutants and myocardial infarction: A systematic review and meta-analysis. *The Journal of the American Medical Association*, 307(7), 713–721.
- Ruhm, C. (2000). Are recessions good for your health? *The Quarterly Journal of Economics*, 115(2), 617–650.
- Ruhm, C. (2003). Good times make you sick. *Journal of Health Economics*, 22(4), 637–658.
- Ruhm, C. (2005). Healthy living in hard times. *Journal of Health Economics*, 24(2), 341–363.