DISSERTATION

ESSAYS ON TRADE AND DEVELOPMENT

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ABSTRACT

ESSAYS ON TRADE AND DEVELOPMENT

This dissertation examines the multifaceted impacts of trade-induced economic shocks on the economic, social, and political landscapes of developing economies, with an in-depth focus on Brazil. While intensifying global demand for commodities can drive export growth and bolster local labor markets, it also presents significant challenges: reshaping productive structures, raising regulatory and property-rights concerns, and influencing local political dynamics. By integrating insights from three interrelated essays, this work enhances our understanding of how exogenous shocks reverberate across multiple dimensions of development in resource-rich contexts.

The first essay investigates the impact of resource booms on export performance and structural transformation. Using a shift-share instrument that leverages heterogeneous exposure to Chinese demand following China's 2001 WTO accession, I examine how Brazil's recent commodity boom affected local economies. The findings reveal that more exposed regions experienced higher export values and greater export concentration, driven by a shift from exporting resource-based manufactures to primary products, leading to a decline in export sophistication. This movement toward raw materials — rather than processed, higher-value-added goods — suggests a reduction in average value added and productivity within the resource or booming sector, highlighting an additional channel through which resource booms can undermine overall productivity. Although wage growth occurred in both the primary and service sectors, primary-sector employment remained stable while manufacturing jobs contracted, aligning with a Dutch disease dynamic. There is also suggestive evidence that these manufacturing job losses coincided with a rise in likely informal activities. Overall, the results underscore the trade-offs inherent in resource booms: short-term gains in export value and sectoral wages may come at the expense of long-term development challenges.

Given Brazil's similarities to other commodity exporters, these findings may reflect broader trends across developing economies.

The second essay examines the relationship between the institutionalization of a renewable resource market with poorly defined property rights and local crime rates. We focus on the regulation of the donkey hide trade in Brazil, driven by foreign demand for *ejiao*, a Traditional Chinese Medicine product. Using a quasi-experimental research design, we leverage the timing of regulatory measures alongside spatial variations in donkey populations across Brazilian municipalities to provide causal evidence that the slaughtering of free-roaming donkeys led to an increase in crime and violence. We further explore the role of market illegality, finding that the impact on crime was twice as large during periods when the trade was illegal in Brazil. Our results carry important policy implications for developing countries grappling with resource booms and weak property rights. These findings emphasize the need for effective regulation, robust monitoring, and enforcement mechanisms to mitigate the social costs associated with natural resource exploitation.

The third essay explores how trade-induced economic shocks distort voter perceptions of incumbents, weakening electoral accountability. While economic voting allows citizens to reward or punish incumbents based on local economic conditions, external factors can obscure the true sources of these fluctuations. Using a game-theoretic framework and a difference-in-differences approach, I analyze the reelection prospects of Brazilian mayors from 2000 to 2020 in response to two external shocks: the legalization of genetically engineered (GE) soy seeds in 2003 and the global commodity price boom of the 2000s. The theoretical framework shows how exogenous regional economic shocks influence incumbents' electoral prospects by shaping voter perceptions of competence. Empirically, a one standard deviation increase in potential soy yield led to a five percent rise in incumbents' reelection probabilities, as voters credited local leaders for economic upturns driven by global market forces. This dynamic allows short-sighted incumbents to remain in office during booms while penalizing capable mayors during downturns, distorting the accountability mechanism. By highlighting these distortions, this essay contributes to the literature on

electoral accountability in commodity-exporting countries and underscores the need to strengthen voter information and institutional resilience in emerging democracies.

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Finally, I sincerely appreciate the dedicated staff of the College of Liberal Arts and, in particular, the Department of Economics at Colorado State University, whose assistance has been fundamental along the way.

EPIGRAPH

Eu sou da América do Sul	*I am from South America*
Eu sei, vocês não vão saber	*I know, you will not know about it*
Mas agora sou cowboy	*But now I'm a cowboy*
Sou do ouro, eu sou vocês	*I'm from the gold, I'm you*
Sou o mundo, sou Minas Gerais.	*I'm the world, I'm Minas Gerais.*

"Para Lennon e McCartney"*, by Fernando Brant, Márcio Borges, Lô Borges*

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Introduction

This dissertation examines the multifaceted impacts of trade-induced economic shocks on the economic, social, and political landscapes of developing economies, with an in-depth focus on Brazil. While intensifying global demand for commodities can drive export growth and bolster local labor markets, it also presents significant challenges: reshaping productive structures, raising regulatory and property-rights concerns, and influencing local political dynamics. The effects of these shocks are often unevenly distributed, disproportionately affecting regions that are more reliant on commodity exports and altering long-term development trajectories. In many cases, resource booms lead to structural transformations that favor primary sectors at the expense of industrial diversification, contribute to governance challenges related to the management of newly generated wealth, and create distortions in voter accountability mechanisms. By integrating insights from three interrelated essays, this work enhances our understanding of how exogenous shocks reverberate across multiple dimensions of development in resource-rich contexts, shedding light on the economic reallocation effects, institutional weaknesses, and political consequences that arise in the wake of commodity-driven economic shifts.

The first essay investigates the causal impact of resource booms on export performance and structural transformation in resource-rich regions. Using a shift-share instrument that leverages heterogeneous exposure to Chinese demand following China's 2001 WTO accession, along with the *ex-ante* composition of regional export baskets, I examine the effects of the recent commodity boom on Brazilian local economies. The findings reveal that more exposed regions experienced higher export values and greater export concentration, driven by a shift from exporting resource-based manufactures to primary products, leading to a decline in export sophistication. This movement toward raw materials — rather than processed, higher-value-added goods — suggests a reduction in average value added and productivity within the resource or booming sector, highlighting an additional channel through which resource booms can undermine overall productivity. Although wage growth occurred in both the primary and service sectors, primary-sector employment remained stable while manufacturing jobs contracted, aligning with a Dutch disease dynamic. There is also suggestive evidence that these manufacturing job losses coincided with a rise in likely informal activities. Overall, the results underscore the trade-offs inherent in resource booms: short-term gains in export value and sectoral wages may come at the expense of long-term development challenges. Given Brazil's similarities to other commodity exporters, these findings may reflect broader trends across developing economies.

The second essay examines the relationship between the institutionalization of a renewable resource market with poorly defined property rights and local crime rates. Focusing on the regulation of the donkey hide trade in Brazil, driven by foreign demand for *ejiao*, a Traditional Chinese Medicine product, we use a quasi-experimental research design to analyze its effects. By leveraging the timing of regulatory measures alongside spatial variations in donkey populations across Brazilian municipalities, we provide causal evidence that the slaughtering of free-roaming donkeys led to an increase in crime and violence. We further explore the role of market illegality, finding that the impact on crime was twice as large during periods when the trade was illegal in Brazil. These findings emphasize the need for effective regulation, robust monitoring, and enforcement mechanisms to mitigate the social costs associated with natural resource exploitation and carry important policy implications for developing countries grappling with resource booms and weak property rights.

The third essay examines how trade-induced economic shocks distort voters' perceptions of incumbent performance, weakening electoral accountability. While economic voting allows citizens to reward or punish incumbents based on economic conditions, external factors can obscure the true origins of economic fluctuations. This study develops a game-theoretic framework and employs a difference-in-differences approach to analyze the reelection prospects of Brazilian mayors from 2000 to 2020 following two major external shocks: the legalization of genetically engineered (GE) soy seeds in 2003 and the global commodity price boom of the 2000s. The theoretical framework shows how exogenous regional economic shocks influence incumbents' electoral prospects by shaping voter perceptions of competence. Empirically, a one standard deviation increase in potential soy yield led to a five percent rise in incumbents' reelection probabilities, as voters credited local leaders for economic upturns driven by global market forces. This dynamic enables short-sighted or less competent incumbents to remain in office during booms while penalizing capable mayors during downturns, distorting the accountability mechanism. By highlighting these distortions, this essay contributes to the literature on electoral accountability in commodity-exporting countries and underscores the need to strengthen voter information and institutional resilience in emerging democracies.

Taken together, these essays provide new insights into the economic and institutional vulnerabilities that arise in resource-rich developing economies. The findings illustrate how trade-induced shocks can drive economic specialization that hampers long-term growth, exacerbate governance challenges in managing newly created wealth, and distort the democratic process by shaping voter behavior in unintended ways. More broadly, this dissertation contributes to ongoing discussions on how external economic shocks interact with domestic institutions—both intentionally and unintentionally—shaping development outcomes beyond the immediate economic benefits of rising exports. Understanding these dynamics is critical for policymakers seeking to mitigate the adverse effects of resource dependency, strengthen regulatory frameworks, and enhance mechanisms of democratic accountability in commodity-exporting nations.

By way of conclusion, this dissertation underscores the importance of systematically addressing the structural specificities of less developed economies in both theoretical and empirical work. A deeper understanding of these structural constraints is essential for assessing how developing economies integrate into the global system and how their interactions with advanced economies shape their development trajectories. Looking inward is not merely an intellectual exercise but a necessary step in overcoming the persistent barriers to economic development. Structural obstacles cannot be dismissed as secondary concerns; rather, they may hold the key to fostering more sustainable and equitable growth in the long run.

Chapter 1

Resource boom, export composition, concentration, and sophistication: evidence from Brazilian local economies

[1]

1.1 Introduction

The economic phenomenon known as Dutch disease continues to pose a persistent challenge for resource-abundant countries, especially amid rising global commodity prices in recent decades. It occurs when a surge in revenue from booming natural resources undermines other economic sectors, particularly manufacturing. This raises key questions about how these booms impact export structures and broader economic development.

Extensive literature highlights the negative effects of natural resource booms on non-resource tradable sectors (Corden, 1984; van der Ploeg & Venables, 2011; Van Wijnbergen, 1984a). These booms often redirect labor and capital from manufacturing to resource industries, leading to deindustrialization (Corden & Neary, 1982; Matsuyama, 1992). Simultaneously, resource windfalls frequently boost demand for non-tradable goods and services, intensifying manufacturing contraction (Alberola & Benigno, 2017; Corden, 1984; Van Wijnbergen, 1984b; Venables, 2016). Beyond intersectoral adjustments, recent evidence suggests that within manufacturing, commodity booms reallocate market share away from exporters and capital-intensive firms, potentially reducing av-

[1] A working paper version of this essay can be found here. I am grateful for insightful discussions, comments, and suggestions from Edward Barbier, Elissa Braunstein, Firat Demir, Rafael Dix-Carneiro, Iasmin Goes, Timothy Komarek, Ray Miller, Fernando Rugitsky, Gilberto Tadeu Lima, Daniele Tavani, and Ramaa Vasudevan. I also appreciate the feedback from seminar participants at City College of New York, Colorado State University, Denison University, Governors State University, New Mexico State University, Smith College, Universidad de Los Andes, University of Massachusetts Amherst, University of Utah, the 50th Eastern Economic Association Annual Conference, and the 5th International Growth Workshop in Rio de Janeiro. Any remaining errors are my own.

erage productivity (Heresi, 2023). Yet, little attention has been paid to how export baskets evolve after such booms, leaving a gap in understanding the broader economic consequences of resource dependence. This is particularly important as export baskets reflect the productive structure of local economies, especially their most dynamic firms (Bernard & Jensen, 2004).

This study focuses on Brazil, a resource-rich developing country and major global exporter. Leveraging a novel shift-share instrument, I examine how exogenous trade shocks — specifically those linked to commodity production — reshape the composition, concentration, and sophistication of regional export baskets over time. Understanding these shifts is key to identifying localized structural changes caused by resource booms. Moreover, these findings could have broader relevance for other commodity-exporting countries with similar economic structures.

Policy concerns regarding export concentration in primary products — and its potential negative effects on terms of trade, income volatility, and long-term economic growth — have long been recognized. Foundational works by Prebisch (1949) and Singer (1950) underscore the developmental challenges tied to dependence on commodity production. Research on the "natural resource curse" further explores how resource dependence negatively impacts economic growth (Barbier, 2019; Isham et al., 2005; Sachs & Warner, 1995, 2001; van der Ploeg & Venables, 2011, 2013). Although these macroeconomic effects are well-documented, less attention has been given to how resource booms and external demand shocks reshape the composition and concentration of both resource and non-resource sectors. Bahar and Santos (2018) provide a notable exception, finding that countries with larger shares of natural resource exports tend to have more concentrated non-resource export baskets. However, their national-level analysis overlooks regional dynamics and also does not fully consider shifts in export structures. Addressing this gap is critical since export diversification and sophistication are widely recognized as key drivers of growth (Cadot et al., 2011; Hausmann et al., 2007; Hidalgo et al., 2007; Imbs & Wacziarg, 2003; Klinger & Lederman, 2006). Furthermore, regional dynamics often diverge from national trends following resource booms (Allcott & Keniston, 2018; Cust & Poelhekke, 2015; Marchand & Weber, 2018; Pelzl &

Poelhekke, 2021), underscoring the importance of accounting for these variations to comprehend their long-term developmental consequences fully.

This paper addresses these gaps by leveraging a quasi-natural experiment that generated varied export demand across Brazilian regions. The analysis focuses on the surge in Chinese export demand following China's accession to the World Trade Organization (WTO) in 2001. Using a novel identification strategy, I examine the causal effects of resource booms on export performance and sectoral employment in Brazil from 2000 to 2019. Brazil presents a particularly relevant and timely case, as its share of resource exports rose from under 50% in 2000 to nearly 80% by 2019. This approach allows for precise identification of the causal effects of resource booms, overcoming challenges such as reverse causality and omitted variable bias. It also aligns with recent studies exploring the impact of resource booms on US regional economies (Allcott & Keniston, 2018; Feyrer et al., 2017; James & Smith, 2020).

Although some studies have explored the effects of China's rise as a dominant global trade player on Brazil, this literature remains relatively limited. Costa et al. (2016) use a quasi-experimental approach to examine how China's ascent impacted Brazilian local labor markets. They find that regions exposed to Chinese import competition experienced slower growth in manufacturing wages between 2000 and 2010. Conversely, regions benefiting from heightened Chinese export demand saw faster wage growth. Carreira et al. (2024) employ a similar empirical strategy to assess the impacts of trade shocks on deforestation in Brazil, observing significant impacts on land use but no direct link between exposure to Chinese demand and deforestation. While these studies offer valuable insights into labor markets and land use, the effects of such external shocks on regional export dynamics and productive structures remain underexplored.

In this paper, I construct a shift-share instrument that leverages the heterogeneous exposure of regions to China's export demand shock, based on the *ex-ante* composition of regional export baskets. The results show a notable increase in total export value and heightened concentration in the most exposed regions. For example, moving a region from the 25th to the 75th percentile of exposure leads to a $270 million increase in export value (14% growth) and a 0.008 rise in the

Herfindahl-Hirschman Index (HHI). This concentration is driven by a focus on already-exported products, with minimal changes in export variety. Interestingly, the share of non-resource (i.e., manufacturing) exports remains stable, showing no relative decline in the most affected regions compared to less exposed areas. However, I find suggestive evidence of adjustment within the resource basket, indicating a shift toward primary products at the expense of resource-based manufactures.

Building on the extensive literature on export basket sophistication (Hausmann et al., 2007; Hidalgo & Hausmann, 2009; Hidalgo et al., 2007; Jarreau & Poncet, 2012), I investigate the "primarization" effect — where regions shift toward exporting lower value-added resource goods — triggered by the surge in Chinese demand. The analysis reveals a decline in the average complexity of export baskets in more exposed regions relative to less exposed ones. This shift toward raw materials and basic commodities, rather than processed, higher-value-added goods, may indicate a decline in average value added and productivity *within* the resource or booming sector, highlighting an additional channel through which resource booms can contribute to overall productivity reductions. This dynamic raises concerns about the long-term development of these local economies, particularly in terms of technological progress and upgrading.

I also explore how these structural changes affect labor market dynamics. Despite a shift toward more primary exports, this change does not increase employment in the primary sector. Instead, I find wage increases in both the primary and service sectors. Concurrently, there is a notable decline in manufacturing employment in the most affected regions, consistent with the Dutch disease phenomenon. I also find suggestive evidence that this contraction in manufacturing employment is accompanied by a rise in likely informal activities.

These findings carry important policy implications for Brazil and other resource-rich developing economies. While the resource boom driven by Chinese demand has significantly boosted regional export values and sector-specific wages, it has also led to a shift toward simpler, lower-value-added exports, with consequences for sectoral employment, particularly in manufacturing and potentially the informal sector. These results challenge the predominantly positive view of the

China-led export demand shock in Brazilian labor markets presented by Costa et al. (2016), revealing structural changes that could undermine long-term economic development, in part by reducing total factor productivity in the most exposed regions. Furthermore, they align with recent evidence from Branstetter and Laverde-Cubillos (2024), who find that Colombia's recent resource boom, while associated with income growth, negatively impacted technological development through persistent declines in R&D spending and investment in technological upgrading within the manufacturing sector. These dynamics may be unfolding in other resource-dependent economies, highlighting the importance of policies that not only capture short-term benefits from resource booms but also promote sustained economic diversification to support long-term development.

This study also contributes to the growing literature on the global repercussions of China's rise as a dominant force in international trade. Most prior research has focused on the effects of Chinese import competition on manufacturing employment and wages (Autor et al., 2013, 2014; Dix-Carneiro et al., 2023; Pierce & Schott, 2016). However, fewer studies have examined the demand-side effects of China's economic expansion. For example, Dauth et al. (2014) use a reduced-form approach to explore how rising imports from and exports to China and Eastern Europe affect labor markets in Germany. Yet, research specifically addressing the demand-side effects of China's expansion in developing economies remains limited.

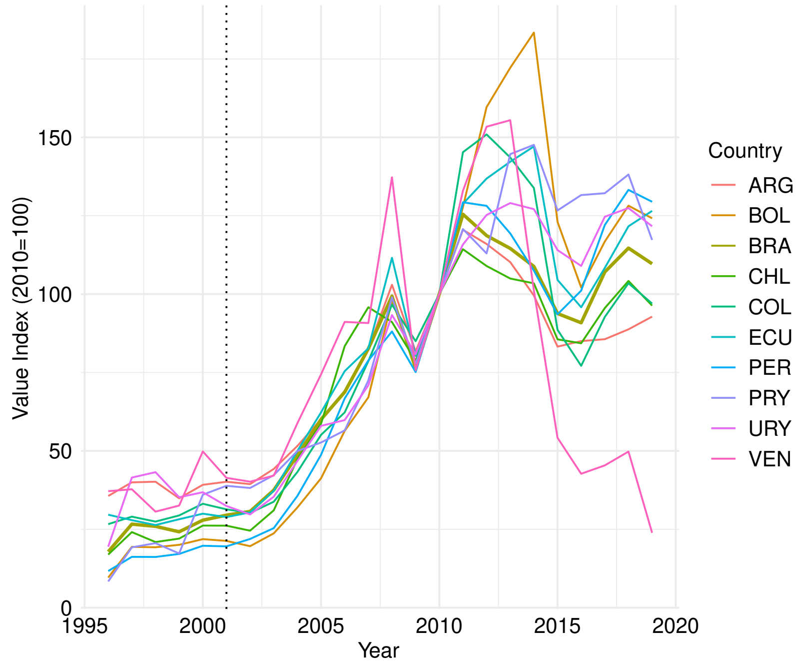
The remainder of the paper is organized as follows. Section 1.2 provides background on the effects of China's WTO accession on resource-rich economies, with a focus on South America, and outlines recent trends in Brazil's resource boom. Section 1.3 describes the data used in the empirical analysis, while Section 1.4 details the empirical strategy and identification approach. Section 1.5 presents the main results, and Section 1.6 concludes with a discussion of broader implications.

1.2 Background: Transformative Trade Dynamics

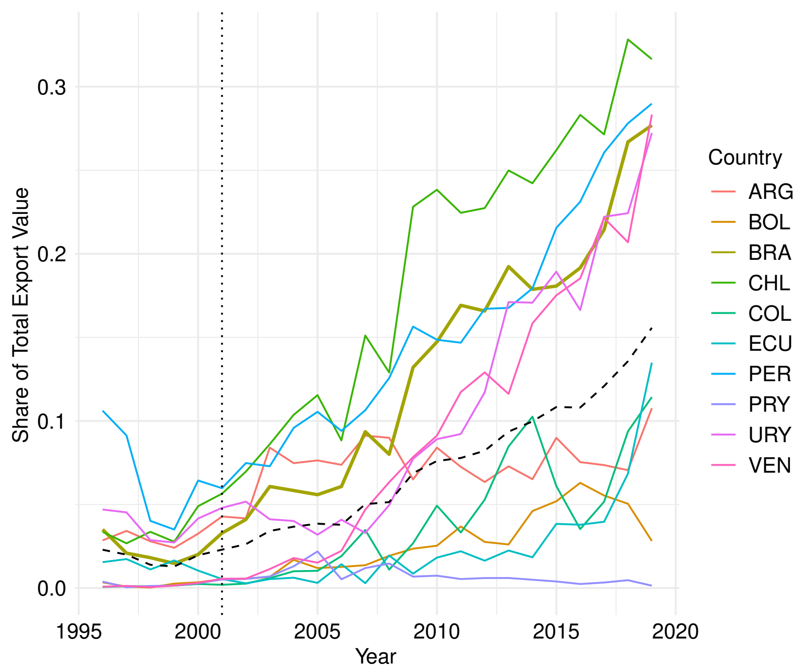
1.2.1 China's ascent to the WTO: transformation in trade patterns

China's rapid emergence as a global economic powerhouse has profoundly altered the demand for primary goods over recent decades. Its remarkable economic growth, abundant reserves of labor, land, and capital, and deepening integration into the global economy have driven substantial shifts in global economic dynamics. China's accession to the WTO in 2001 marked a pivotal moment in international trade, transforming trade patterns worldwide (Autor et al., 2013). For developing countries, China quickly became a dominant exporter of manufactured goods and a major importer of raw materials (Costa et al., 2016).

The surge in demand for primary goods has been particularly pronounced in countries historically reliant on such exports, particularly across South America. Figure 1.1 shows that the total value of exports from South American countries has significantly increased over the past three decades, with a large portion of this growth attributable to trade with China. Panel (a) illustrates a clear upward trajectory in total export values for each country since 2001, coinciding with China's WTO accession. Panel (b) highlights the sharp rise in the share of total exports destined for China, with the average share across the sample increasing from approximately 3% in 2001 to over 15% by 2019. Resource-dependent economies such as Chile, Peru, and Brazil saw shares exceeding 25% in 2019.



(a) Total Value — Index

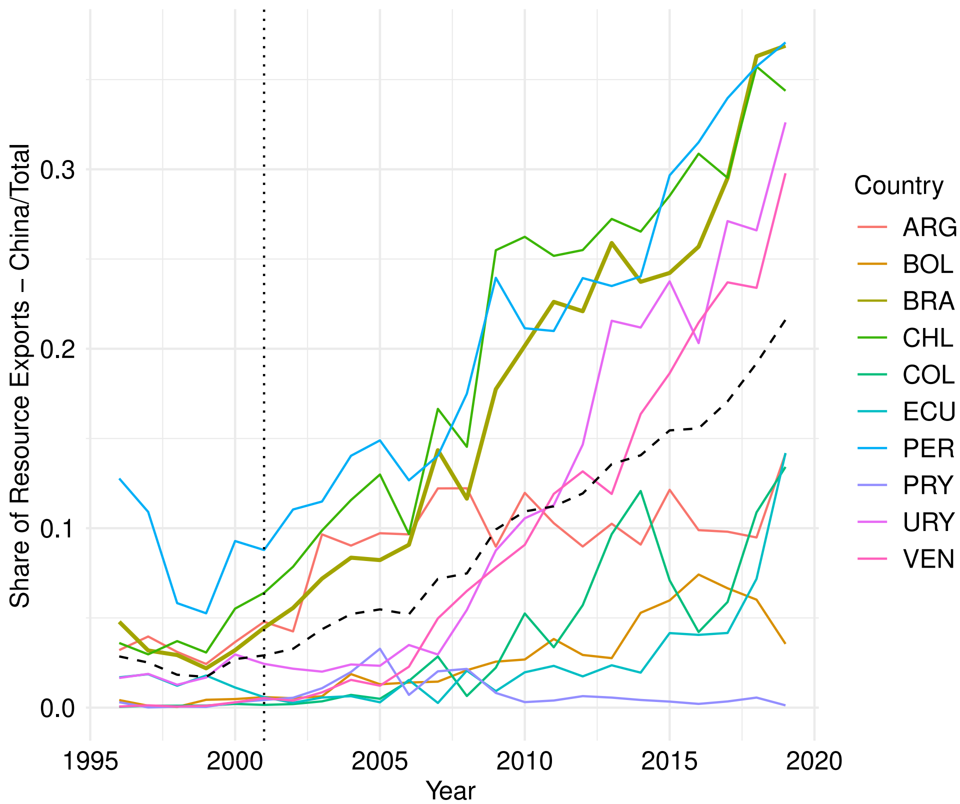


(b) Share of Total Exports Value — China/Total

Source: Bilateral trade flow data from the BACI database, developed by the *Centre d'Études Prospectives et d'Informations Internationales* (CEPII). Dotted lines indicate 2001, and the dashed line in panel (b) represents the yearly average. Thicker line highlights Brazil.

Figure 1.1: Total exports value — South American countries

This trend is even more pronounced when we focus on the share of total export value represented by resource goods, including primary products and resource-based manufactures, as shown in Figure 1.2. The largest South American economies experienced a significant increase in these shares, rising from around 3% in 2001 to over 20% by 2019. This pattern underscores China's growing role as a major importer of resource-based products from these countries. Notably, Chile, Peru, and Brazil saw shares exceed 35%. Among these, Brazil stands out not only for its significant export share but also for its broader economic implications, making it a compelling case study for understanding the full impact of China's ascent.



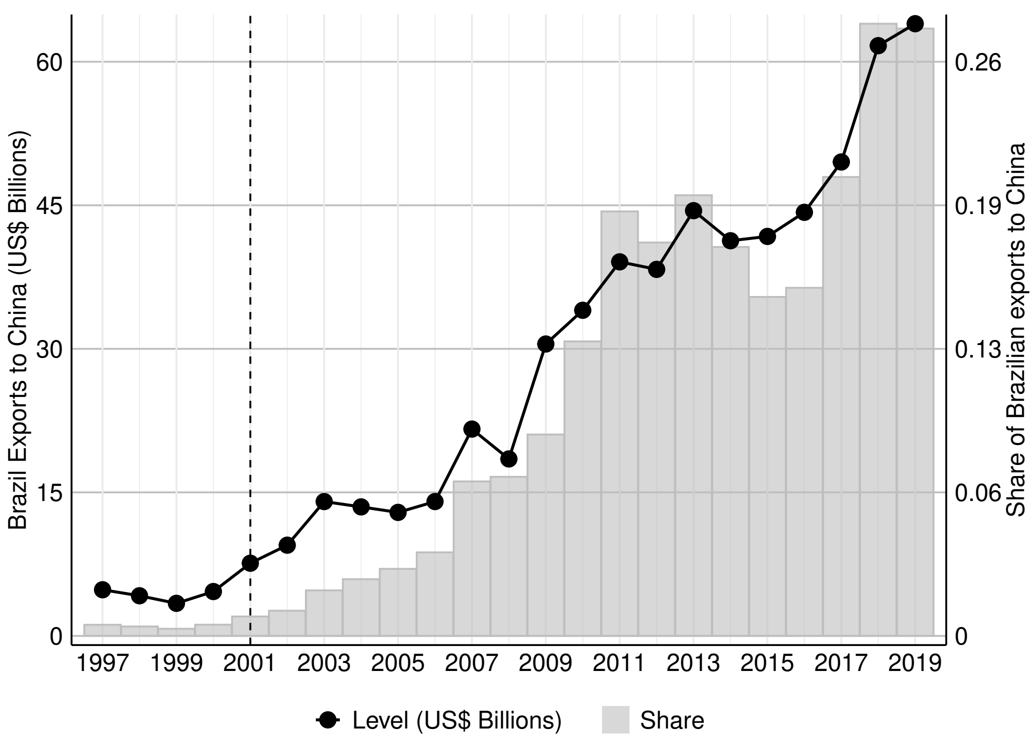
Source: Bilateral trade flow data from the BACI database, developed by the *Centre d'Études Prospectives et d'Informations Internationales* (CEPII). Product classification follows Lall (2000). Dotted line indicates 2001, and the dashed line represents the yearly average. Thicker line highlights Brazil.

Figure 1.2: Share of resource exports — China/Total

1.2.2 The China-driven resource boom in Brazil

Building on broader South American trends and given its economic significance for the region, Brazil's experience offers a detailed case study of how China's rise in international trade has fueled a commodities boom, particularly in soybeans and iron ore (Carreira et al., 2024). Brazil provides a compelling context for examining China's impact on the export composition of developing countries for several reasons.

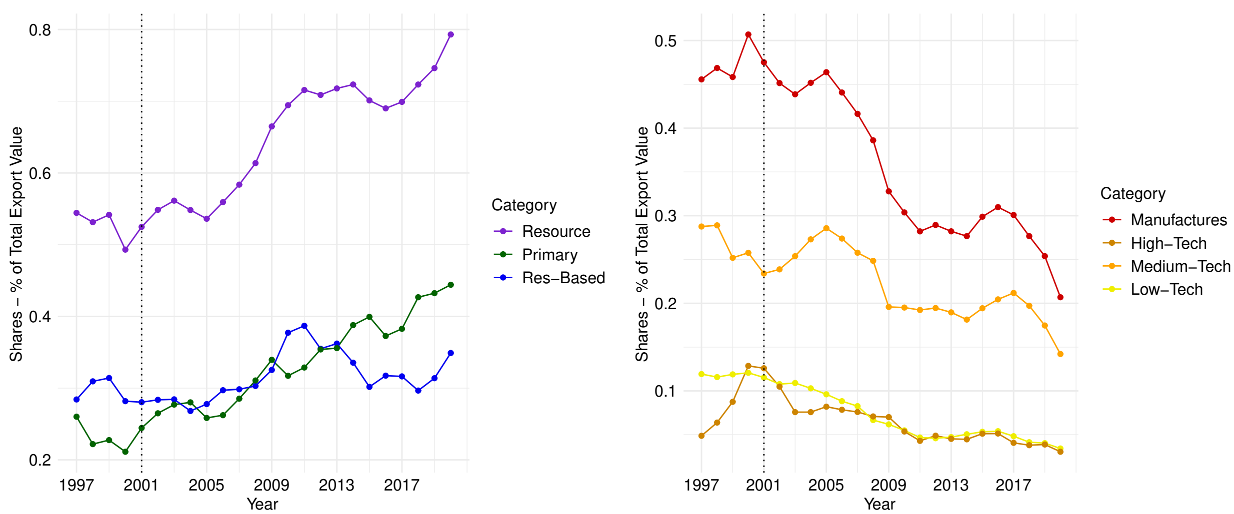
First, China's significance as an export destination for Brazil surged in recent years, surpassing its importance to other major South American economies, as shown in Figure 1.1. Figure 1.3 illustrates China's growing relevance as a key export market for Brazilian products over the past decades, both in terms of total export value and the share of exports destined for China. By the late 2000s, China had already become Brazil's largest trade partner.



Source: Export value data is based on declarations by exporters in Brazil (SISCOMEX, Ministry of Industry, Foreign Trade, and Services). Dashed line indicates 2001.

Figure 1.3: Brazil's exports to China - total value and share

Second, the trade pattern between Brazil and China aligns with broader South American trends, as Brazilian exports increasingly shift towards agricultural and extractive products. Figure 1.4 tracks the evolution of export shares across different product classifications, based on the definitions by Lall (2000), which categorize goods by their technological content. Panel (a) reveals a substantial increase in resource exports relative to total exports from 2001 onward, with a shift from resource-based manufactures to primary products. In panel (b), a significant decline appears in the share of manufacturing exports at the aggregate level, with decreases observed across all subcategories.



(a) Resource

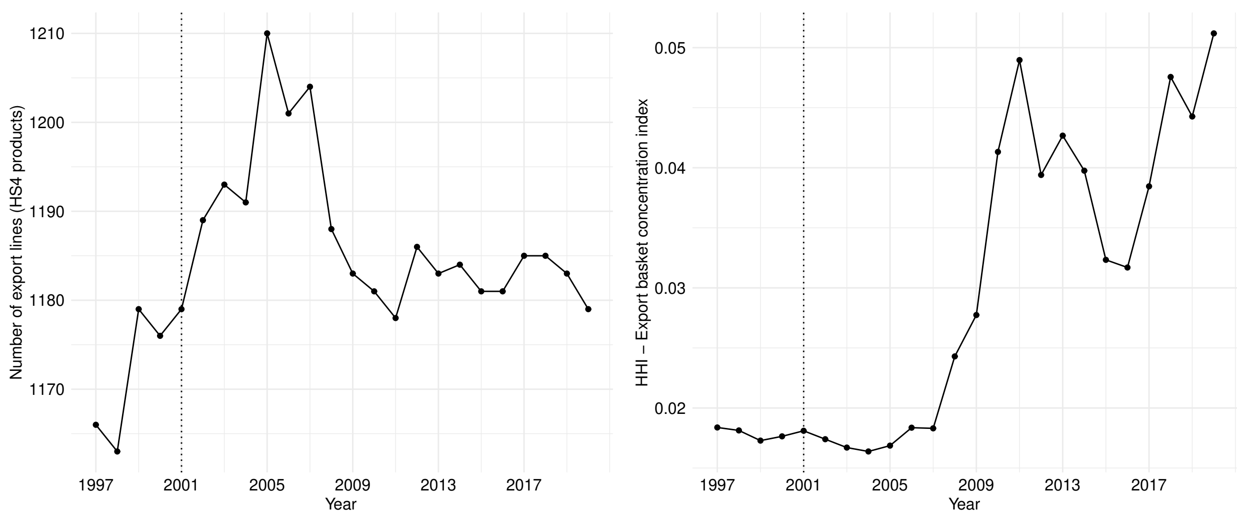
(b) Manufactures

Source: Export value and composition data are based on declarations by exporters in Brazil (SISCOMEX, Ministry of Industry, Foreign Trade, and Services). Product classification follows Lall (2000). Dotted lines indicate 2001.

Figure 1.4: Brazil's export shares, 1997-2019

Third, Brazil's vast size and diverse geography give rise to regional economies with varied comparative advantages, enabling the identification of heterogeneous trade effects without relying on cross-country regressions. This approach allows a closer examination of the causal impacts of a resource boom on the export baskets of local economies that resemble typical small open economies.

Beyond assessing the impacts on the value and composition of the export basket, this paper focuses on its concentration following the resource boom. Figure 1.5 provides high-level evidence of stability in the number of products exported by Brazil from 2000 to 2019, while also revealing a clear increase in the Herfindahl-Hirschman Index (HHI), indicating growing concentration in Brazil's export basket over time along the intensive margin.[2] This rise in export basket concentration temporally coincides with China's growing global influence.



(a) Number of export lines (b) HHI

Source: Export value and composition data are based on declarations by exporters in Brazil (SISCOMEX, Ministry of Industry, Foreign Trade, and Services). Dotted lines indicate 2001.

Figure 1.5: Export basket concentration: Brazil, 1997-2019

1.3 Data Description

1.3.1 Regional export data

To investigate regional export dynamics in Brazil, I use the SISCOMEX dataset, an administrative source maintained by Brazil's Ministry of Industry, Foreign Trade, and Services. This dataset captures monthly export data from 1997 to 2024, documenting the tax jurisdiction or fiscal location

[2]The construction of this concentration measure is discussed in detail in Section 1.3.

of the exporting firm.[3] For the analysis, I aggregate this data to the municipal and yearly levels from 1997 to 2019, focusing on periods before and after the quasi-natural experiment triggered by the Chinese export demand shock, excluding the period affected by COVID-19.

Given the data structure and the relatively low share of net exports as a percentage of GDP in Brazil during the 1990s, I focus on municipalities with consistent export activity during the years studied. This results in a smaller sample than the total number of municipalities recorded during the same period. To reduce potential distortions and gain clearer insights into the resource boom's impact on Brazil's local economies, I aggregate municipalities into micro-regions. These micro-regions, encompassing groupings of economically integrated municipalities with similar geographic and productive characteristics, are delineated by the Brazilian Institute of Geography and Statistics (IBGE) and are widely recognized in economic literature for characterizing regional economies in Brazil (e.g. Costa et al. (2016), Dix-Carneiro and Kovak (2017), Dix-Carneiro et al. (2018), Hirata and Soares (2020), Ogeda et al. (2024), and Ponczek and Ulyssea (2022)). As noted earlier, these local economies closely approximate the conditions of small open economies. The aggregation results in a dataset of 424 consistently observed exporting micro-regions.

The export data at the municipal level disaggregates products based on the Harmonized System (HS) classification at the four-digit level, corresponding to headings rather than subcategories. This classification encompasses over 1,200 product lines. To classify products as resource-based or non-resource-based, I use the technological definitions provided by Lall (2000), which categorize goods based on their technological content using the Standard Industry Trade Classification (SITC 3-digit, revision 2). I then cross-reference this classification with the HS to categorize the products exported by Brazilian local economies.

Figure 1.6 contextualizes the observed expansion of resource product exports in Brazil over recent decades and their geographic distribution across the country. It presents the change in the

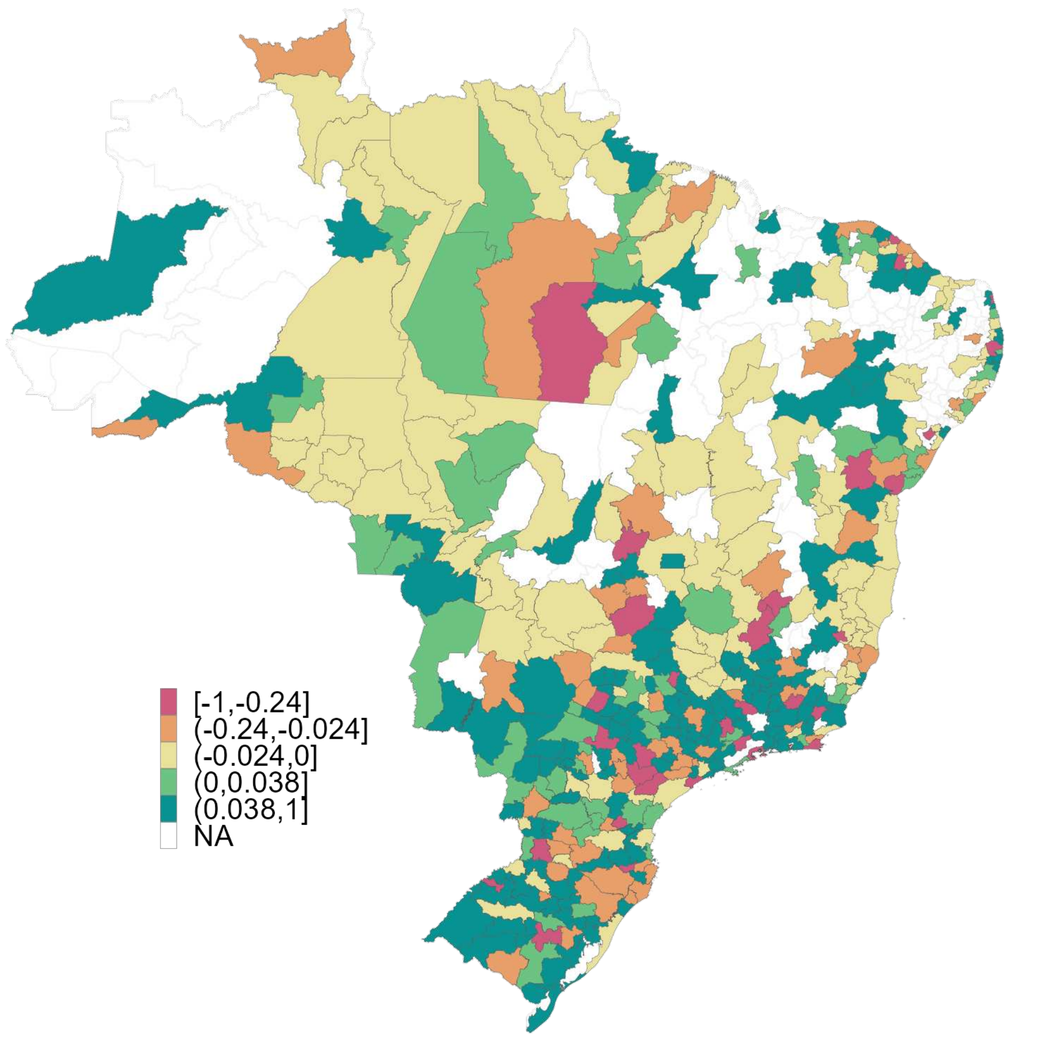
[3]While the primary analysis uses municipal-level data, the Appendix A presents state-level results, measured at the locality of production, confirming that findings are consistent across different levels of aggregation.

share of resource exports, which includes both primary products and resource-based manufactures, relative to the total export value across local economies in Brazil.

Further disaggregation in Figure 1.7 shows the variations in export shares of each product classification across Brazilian local economies from 2000 to 2019, excluding high-tech manufactures due to their limited relevance during this period.[4]

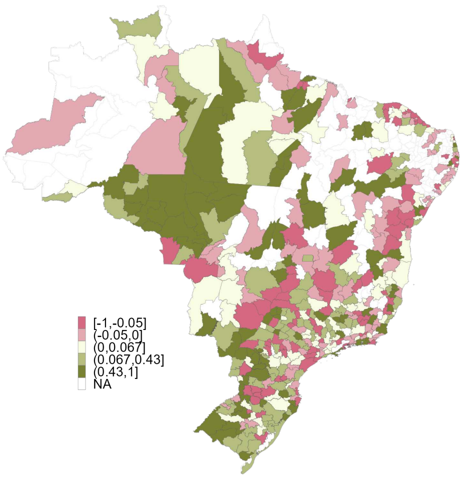
Comparing panels (a) and (b), regions in the North, Central-West, Southeast, and South — particularly those historically focused on agrarian production — show an expansion in primary product exports from 2000 to 2019. Concurrently, these regions have reduced the share of resource-based manufactures in their export baskets, with some exceptions. Regarding manufacturing, as depicted in panels (c) and (d), significant heterogeneity exists in the evolution of export patterns across Brazilian regions during this period. Notably, the Southeast region, particularly its metropolitan areas, which are historically more economically developed and possess a more diversified productive structure, appears to have increased its exports of medium-tech manufactures. Many locations experiencing such shifts also witnessed relative reductions in exports of low-tech manufactures.

[4]The Appendix A presents the spatial distribution of the share of total export value for each category across Brazilian local economies in 2000 and 2019.

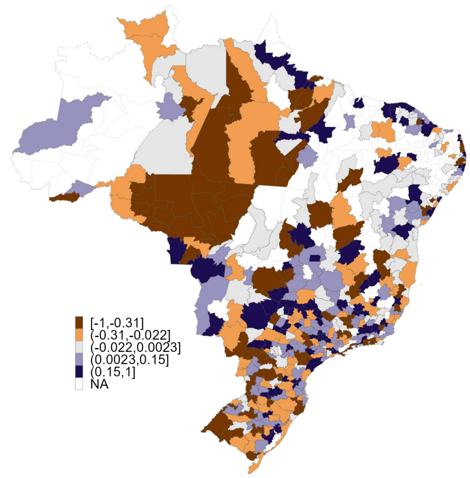


Source: Export value and composition data are based on declarations by exporters in Brazil (SISCOMEX, Ministry of Industry, Foreign Trade, and Services). Product classification follows Lall (2000).

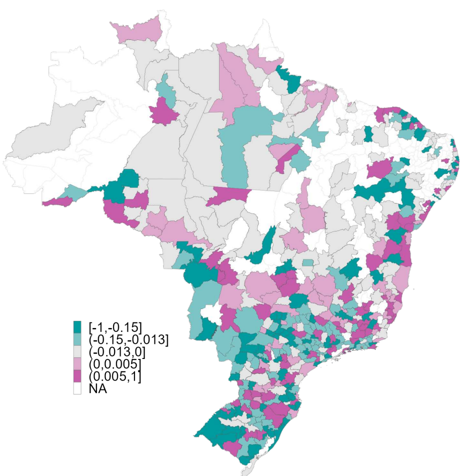
Figure 1.6: Difference in the share of resource exports: 2019 - 2000



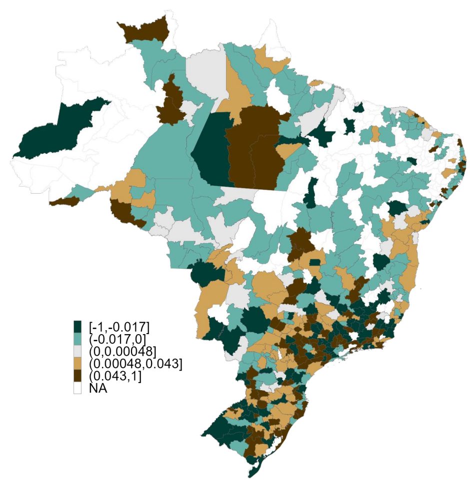
(a) Primary



(b) Resource Based



(c) Low Tech



(d) Medium Tech

Source: Export value and composition data are based on declarations by exporters in Brazil (SISCOMEX, Ministry of Industry, Foreign Trade, and Services). Product classification follows Lall (2000).

Figure 1.7: Difference in share of exports per category: 2019 - 2000

1.3.2 Export basket concentration

To analyze the relationship between the resource boom and export concentration, I employ two widely recognized measures of concentration for each local economy and year: the number of export lines (products exported with a value above zero) and the Herfindahl-Hirschman Index

(HHI) (as used in Bahar and Santos (2018), Cadot et al. (2011), Imbs and Wacziarg (2003), and Koren and Tenreyro (2007)).

The HHI quantifies the concentration of export activity within a region, normalized to range between 0 (indicating no concentration) and 1 (indicating maximum concentration). It is calculated using the following formula:

$$HHI_r = \frac{\sum_k s_{r,k}^2 - \frac{1}{N_r}}{1 - \frac{1}{N_r}} \tag{1.1}$$

where $s_{r,k} = \frac{X_{r,k}}{\sum_{k=1}^{N} X_{r,k}}$ represents the share of export line k (with an export value of $X_{r,k}$) in the total exports of micro-region r, and N_r is the number of export lines in that region.

1.3.3 Export basket sophistication

In addition to examining total value and concentration, I explore the impacts of the resource boom on export sophistication across local economies. To measure export basket sophistication, I adopt the methodology established by Hausmann et al. (2007), which evaluates the complexity of a region's exports by comparing them to the income levels of countries with similar export structures. Specifically, I construct an annual measure of sophistication for each Brazilian micro-region's export basket.

The process begins with the calculation of a country's revealed comparative advantage (RCA) in exporting a specific good, following the approach of Balassa (1965). The RCA index $RCA_{k,j}$ is defined as:

$$RCA_{j,k} = \frac{\frac{x_{j,k}}{X_j}}{\sum_j \frac{x_{j,k}}{X_j}} \tag{1.2}$$

where $x_{j,k}$ is the value of exports of good k by country j and $X_j = \sum_k x_{j,k}$ is the total value of country j's exports.

Next, I calculate an intrinsic sophistication level P_k for each good k. This level is determined as the weighted average of the income levels of countries exporting good k, with weights corresponding to the RCA of each country:

$$P_k = \sum_j RCA_{j,k} \times Y_j \tag{1.3}$$

where Y_j is the per capita income of country j, measured as the real GDP per capita in PPP.

This measure, commonly referred to as "PRODY" in the literature (e.g., Hausmann et al. (2007), Hidalgo and Hausmann (2009), and Hidalgo et al. (2007)), reflects the average income level associated with the production and export of good k, weighted by each exporter's comparative advantage. Essentially, this measure infers from observed trade patterns which products require higher levels of economic development for their export, rather than directly determining intrinsic product features such as embedded technology (Jarreau & Poncet, 2012).

The data for these calculations come from the CEPII-BACI database, which consolidates information from the United Nations Statistical Division's COMTRADE database. This dataset contains annual bilateral trade values at the 6-digit level of the HS classification for over 200 countries, starting from 1995. For this analysis, I aggregate products to the 4-digit level and use data from 1997 to 2000 to establish average RCA measures for each product and country prior to China's accession to the WTO. Additionally, I use data from the World Development Indicators (WDI) and the Penn World Table (PWT) to compute the average real per capita income for each country during the same period.

Following the construction of the product-level sophistication index, I compute a regional export sophistication level, denoted S_r, for each local economy's export basket following Jarreau and Poncet (2012). This index is calculated as the weighted sum of the sophistication levels P_k of each exported good k, with weights representing the share of each good in the micro-region's total exports:

$$S_r = \sum_k s_{r,k} P_k \tag{1.4}$$

where $s_{r,k}$ is defined as in Equation (1.1). Although the time subscript is omitted for simplicity, this measure of regional export basket sophistication is constructed annually, based on the export data discussed earlier.

1.3.4 Local exposure to the resource boom

To quantify the impact of increased export demand from China at the local level, I begin by calculating a simple measure of local exposure. This measure is constructed by classifying exports by product category for each region in 2000 and 2019. Using data from SISCOMEX, I compile export values for each product at the micro-region level for the year 2000. The export share of each product in each locality is then calculated by dividing the export value for each product by the total export value in the micro-region. Additionally, I incorporate international trade data from the CEPII-BACI database, focusing on the years 2000 and 2019. The values for 2000 are adjusted to 2019 US dollars using the US GDP deflator provided by the US Bureau of Economic Analysis, ensuring consistent comparison and enabling the construction of measures of increased Chinese demand for products.

This initial measure provides a raw estimate of local exposure to heightened export demand from China by multiplying the increase in demand for each product exported by Brazil to China between 2000 and 2019 by the relative significance of that product in the export basket of each region. However, this measure may still be endogenous, as local factors influencing export performance could affect the outcomes.

To address this potential endogeneity, I construct an exogenous measure of local exposure using a shift-share, or "Bartik", instrument. This method, drawing on the approaches of Costa et al. (2016) and Carreira et al. (2024), isolates the effect of global and Brazil-specific shocks on trade patterns It parallels the methodology used to identify the "China shock" in the US economy (e.g., Autor et al. (2013, 2014, 2019, 2020)). The shift-share instrument addresses the endogeneity problem by leveraging variations in global demand shifts that are independent of local conditions in Brazil.

The first step in constructing the instrument involves conducting auxiliary regressions for all countries except Brazil, weighted by initial import values, to isolate China-specific demand shocks:

$$\frac{\Delta \tilde{I}_{j,k,00/19}}{\tilde{I}_{j,k,00}} = \beta_k + \psi_{China,k} + v_{j,k} \tag{1.5}$$

where $\frac{\Delta \tilde{I}_{j,k,00/19}}{\tilde{I}_{j,k,00}}$ is the growth rate in imports of product k by country j from all countries other than Brazil between 2000 and 2019; β_k is the product fixed effect that captures the world average growth of net-of-Brazil imports of product k; $\psi_{China,k}$ is a China-product specific dummy that measures the deviation of the China import growth rate of product k in comparison to the one from the rest of the world. The estimated $\hat{\psi}_{China,k}$ represents the predicted change in global exports to China (excluding Brazil) induced by China-specific factors between 2000 and 2019.

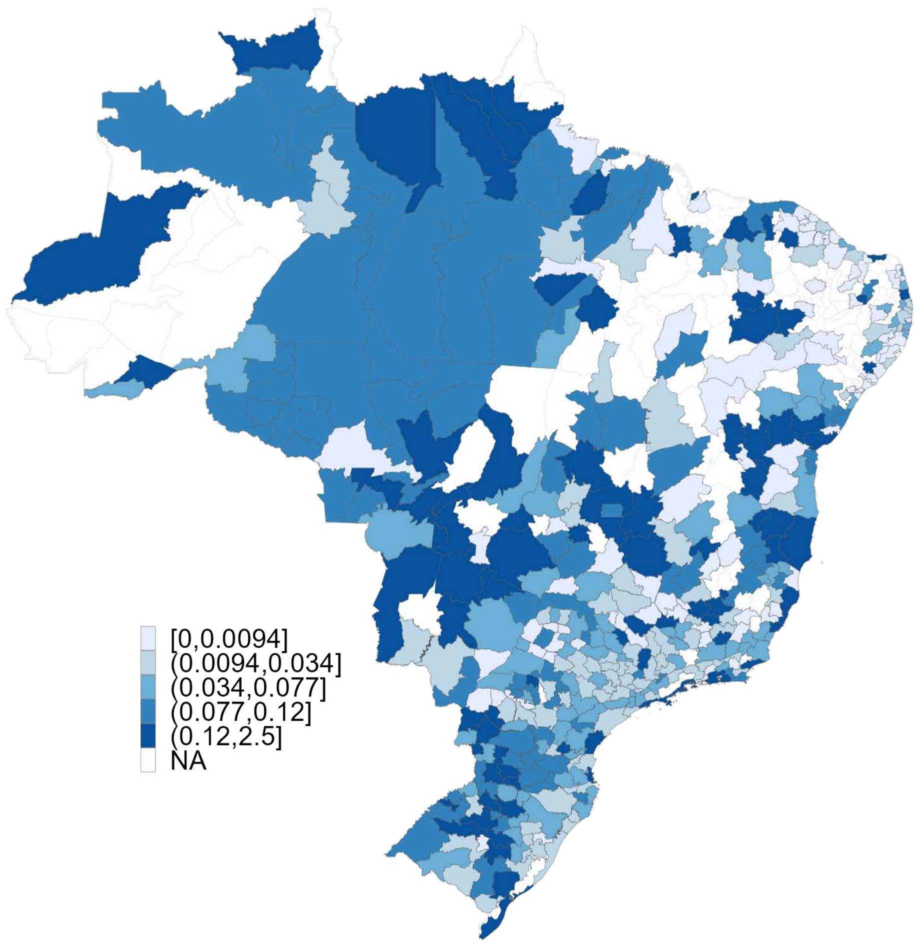
Using the estimated $\hat{\psi}_{China,k}$ and the share of product exports per micro-region in Brazil, I then construct the "Bartik" instrument that quantifies local exposure to China-induced export demand:

$$\Delta \tilde{X}_r = \frac{1}{X_{r,_{00}}} \sum_k \frac{X_{r,k,_{00}}}{X_{B,k,00}} \times X_{BC,k,_{00}} \hat{\psi}_{China,k} \tag{1.6}$$

where $X_{r,00}$ is the total value of exports in region r in 2000; $X_{r,k,_{00}}$ is the total value of exports of product k within region r in 2000; $X_{B,k,_{00}}$ is the Brazil-wide total export value for product k in 2000 and, $X_{BC,k,_{00}}$ is the Brazilian exports of product k to China in 2000.[5] To ensure that the results are not skewed by outliers, I winsorize $\Delta \tilde{X} r$ at the 1st and 99th percentiles.

Finally, Figure 1.8 maps the local exposure to Chinese export demand using the shift-share instrument calculated from Equation (1.6). Notably, local economies in the Central-West, parts of the North, and the South of Brazil exhibit significant exposure to China-induced growth in export demand, highlighting the heterogeneous regional impacts of the resource boom.

[5] In the main results, I use HS2 products to compute the shift-share, resulting in 92 categories. However, the results are virtually identical when using the HS4 classification with approximately 1,200 categories.



Source: Regional exposures to China's export demand, $\Delta\tilde{X}_r$, are computed according to Equation (1.6). Data from CEPII-BACI and SISCOMEX are used to compute the shift-share instrument.

Figure 1.8: Exposure to China's Export Demand - $\Delta\tilde{X}_r$

1.3.5 Additional variables

In addition to trade data, I incorporate local labor market variables to further investigate the channels through which the resource boom may influence regional patterns of structural change.

For this analysis, I use individual-level labor market and socioeconomic data from the Brazilian Demographic Census for the years 2000 and 2010, provided by the Brazilian Institute of Geography and Statistics (IBGE). Following Costa et al. (2016), I restrict the sample to individuals aged 18 to 60, who are most likely to be active in the labor market. Within this cohort, I calculate sectoral employment shares and average hourly wages for employed individuals. Wages are adjusted for inflation using the Brazilian Consumer Price Index (IPCA) and are expressed in 2010 Brazilian reais.

1.4 Empirical Strategy and Identification

My empirical objective is to analyze the effects of the surge in demand for resource exports, driven by China's accession to the WTO in 2001, on the export baskets of regional economies across Brazil. As outlined in Section 1.1, this analysis focuses specifically on changes in export value, composition, concentration, and sophistication over time.

Concerning the empirical strategy, recent research has established a formal framework for identifying assumptions in shift-share regression designs (Borusyak et al., 2022; Goldsmith-Pinkham et al., 2020). Building on the work of Goldsmith-Pinkham et al. (2020) and Borusyak et al. (2022), my identification assumption relies on the notion that the trade shock induced by China, denoted as $\Delta\tilde{X}_r$, is orthogonal to local political and institutional dynamics across micro-regions in Brazil. This independence is largely assured by focusing on the relative effects of increased demand from China compared to all other countries worldwide, excluding Brazil. Therefore, the "Bartik" instrument can be considered a quasi-exogenous shock to local political and institutional dynamics in Brazil.

To achieve this empirical objective, I employ a long-difference or first-difference specification, similar to the methodologies used by Autor et al. (2013), Costa et al. (2016), and Carreira et al. (2024). Specifically, I analyze changes over time in the variables of interest with the following specification, using the micro-region as the unit of analysis:

$$\Delta y_{r,t} = y_{r,t} - y_{r,2000} = c + \beta\Delta\tilde{X}_r + \alpha_{s,t} + \varepsilon_{r,t} \tag{1.7}$$

where $\Delta y_{r,t}$ represents the change in the outcome variable in region r from 2000 to t (2019 in the primary results), $\Delta\tilde{X}r$ denotes the measure of local exposure for region r to the China-induced export demand shock (as detailed in Equation (1.6)), and $\alpha_{s,t}$ are state-time fixed effects. This long-difference specification captures variation in $\Delta\tilde{X}_r$ across micro-regions within states, allowing for clear treatment-control comparisons. In all estimations, I cluster the standard errors at the meso-region level — a larger grouping of micro-regions defined by IBGE — to account for potential spatial correlation in outcomes.[6]

The model specified in Equation (1.7) serves as the baseline for the main results, utilizing a first-difference specification. Alternatively, I implement a dynamic difference-in-differences (DiD) model. In this approach, the measure of exposure to the China-induced regional export demand shock is interacted with year indicators, and I analyze the variables of interest in levels rather than relative changes. This event-study design aligns with recent advancements in the literature (Borusyak et al., 2024; De Chaisemartin & d'Haultfoeuille, 2020; Roth et al., 2023), allowing the assessment of whether treatment and control micro-regions exhibited similar trends in export basket dynamics before the exogenous resource boom. The equivalent dynamic DiD specification to Equation (1.7) is expressed as follows:

$$y_{r,t} = c + \sum_{t=1997}^{2019} \beta_t \mathbb{1}\left\{\tau = t\right\} \Delta\tilde{X}_r + \mu_r + \alpha_{s,t} + \varepsilon_{r,t} \tag{1.8}$$

In Equation (1.8), the year 2001 is set as the baseline treatment year, with $\Delta\tilde{X}r$ serving as the treatment variable. Additionally, μr represents micro-region fixed effects, and $\mathbb{1}$ denotes year indicators. Since all micro-regions were affected simultaneously by China's accession to the WTO in 2001, this empirical approach is not subject to the recent methodological criticisms of the DiD literature (Callaway & Sant'Anna, 2021; De Chaisemartin & d'Haultfoeuille, 2022; Goodman-Bacon, 2021).

As outlined, the dynamic difference-in-differences specification offers a more flexible version of the baseline model, enabling a rigorous empirical evaluation of the parallel trends assumption. Under this assumption, the coefficients β_t for years preceding 2001 should not exhibit significant

[6] In the Appendix A, I demonstrate the robustness of the main results to the inference procedures recommended by Borusyak et al. (2022) to address cross-region residual correlation in shift-share designs.

deviations from zero, either individually or collectively, across all outcome variables of interest. This assessment is crucial for validating that before the treatment — China's accession to the WTO in 2001 — the treatment and control groups experienced similar trends in the variables under study, thus supporting the credibility of the causal inferences.

Lastly, evaluating the validity of the shift-share instrument is crucial for accurately measuring the regional impacts of the significant increase in Brazilian exports to China following 2001. The credibility of this instrument is central to the identification strategy. Figure 1.9 illustrates the relationship between the endogenous measure of exposure and the instrument, revealing a significant correlation. This correlation underscores that the measure of the impact of increased Chinese export demand on Brazilian micro-regions and the instrumental variable ($\Delta\tilde{X}_r$), which captures estimated changes in Chinese demand, are closely aligned.

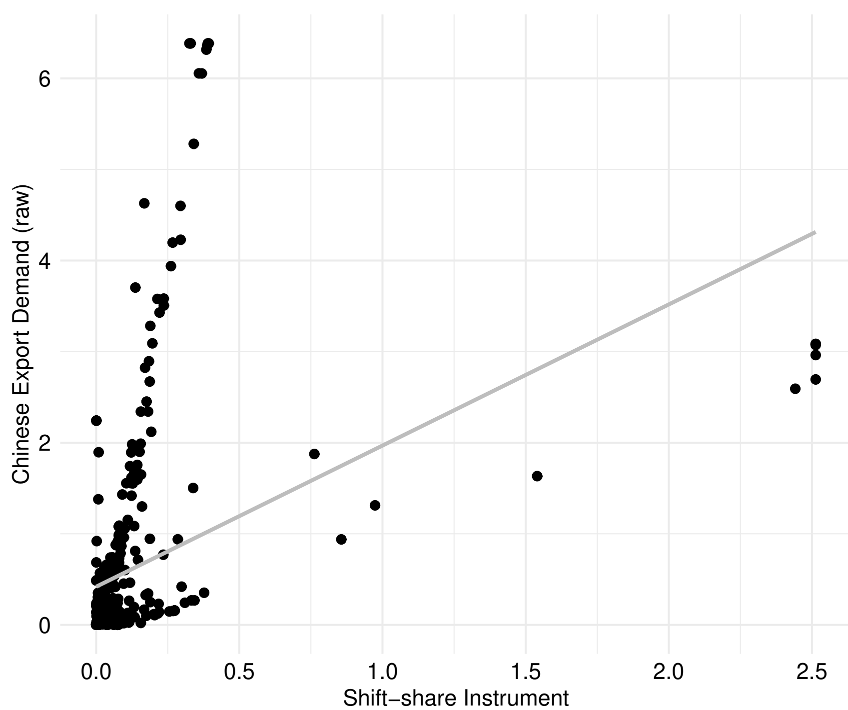


Figure 1.9: Correlation: Chinese export demand and shift-share instrumental variable

To further validate the shift-share instrument, I conduct first-stage regression analyses, detailed in the Appendix A, incorporating state-year fixed effects to control for potential unobserved heterogeneity. These analyses support the instrument's validity, reinforcing the robustness of the empirical findings.

1.5 Impacts of China-Induced Resource Boom

1.5.1 Export value and concentration

I begin by estimating Equation (1.7), using changes in the total export value per micro-region as the dependent variable, analyzed both in levels and growth rates. The expectation is that local economies most affected by the resource boom would see a significant increase in their total export value compared to less impacted regions. The results, presented in Table 1.1, confirm this hypothesis, showing a notable relative increase in both export value and growth rate in micro-regions most influenced by the surge in Chinese demand.

The analysis starts with a simple specification without controls, where observations are weighted by the total export value of each micro-region in 2000, the base year. This approach is crucial for addressing the correlation between the variance in export basket values and the economic size of the regions, as shown in column 1 of Table 1.1. To refine the model and control for potential confounding factors, I introduce state-year fixed effects in column 2, accounting for time-varying regional characteristics that may influence export dynamics. The inclusion of these fixed effects not only maintains the qualitative nature of the initial results but also enhances their statistical robustness. Finally, in column 3, I employ a 2SLS regression model using the shift-share as an instrumental variable for the observed local export growth to China. This specification further reinforces the findings, demonstrating a strong correlation between increased exposure to Chinese demand and export growth.

From the preferred specification in column 2, a micro-region at the 75th percentile of exposure to Chinese demand ($\Delta\tilde{X}r$ = 0.108) experienced, on average, an increase of approximately \$270 million in total export value, compared to a micro-region at the 25th percentile ($\Delta\tilde{X}r$ = 0.015). This relative change corresponds to a growth of about 14% in export value. These results highlight the substantial economic impact of Chinese demand on the most affected Brazilian micro-regions.

Table 1.1: Commodity boom and export value

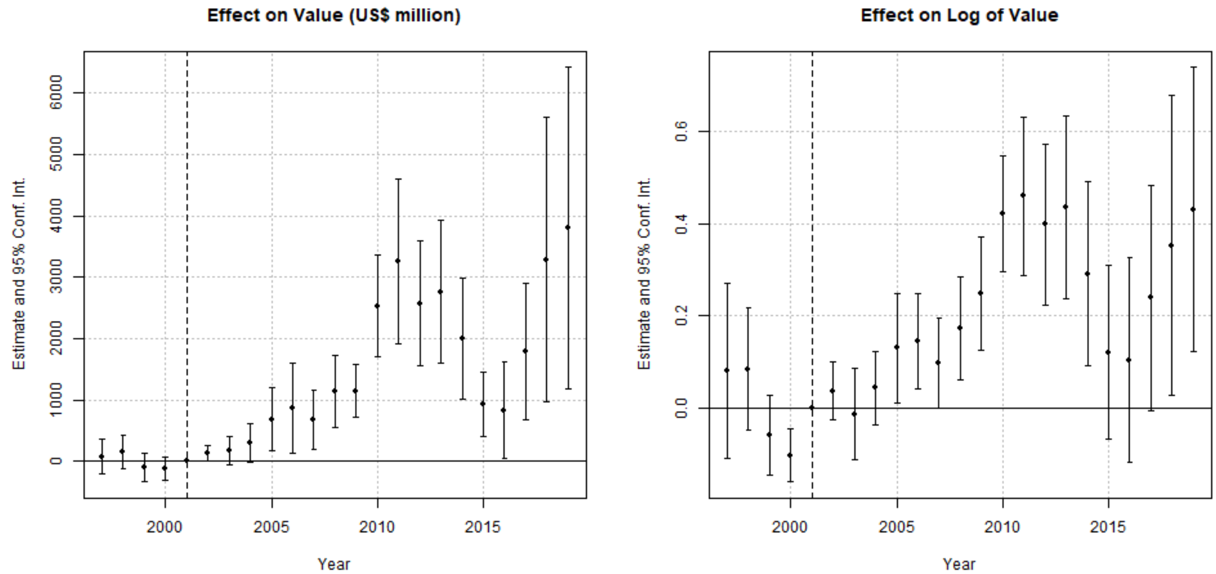
Dependent variable:	Δ **Value of exports (US$ millions)**			$\%\Delta$ **Value of exports**		
	OLS	OLS	2SLS	OLS	OLS	2SLS
	(1)	(2)	(3)	(1)	(2)	(3)
$\Delta\tilde{X}_r$	3,312.141**	2,902.794***		1.730	1.486*	
	(1,628.508)	(1,081.950)		(1.169)	(0.795)	
ΔX_r			2,537.285***			1.299*
			(886.315)			(0.669)
Weighted	X	X	X	X	X	X
State-year fixed effects		X	X		X	X
Observations	424	424	424	424	424	424
Adjusted R^2	0.182	0.625	0.296	0.002	0.066	0.066
KP F-stat			99.7			99.7

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

To assess the causal effects of the China-induced export demand shock on regional export values, I test the parallel trends assumption underlying the identification strategy. This assumption posits that, prior to China's WTO accession, regions with *ex-post* varying levels of exposure to the trade shock would exhibit *ex-ante* similar trends in their export values. To evaluate this, I employ the event-study specification outlined in Equation (1.8), focusing on the coefficients resulting from the interaction of the treatment indicator, $\Delta\tilde{X}_r$, with year dummies. Figure 1.10 illustrates the dynamic effects of the China-induced demand shock on regional export values, along with 95% confidence intervals.

The results in Figure 1.10 visually corroborate the findings in Table 1.1. The overall insignificance of pre-treatment coefficients supports the assumption of parallel trends, thereby reinforcing the robustness of the research design.

Next, I examine the implications of the resource boom on the concentration of regional export baskets. While the increase in export value for regions most affected by the shock is evident, its effect on export concentration — measured by the number of exported products and the HHI — is more complex. One might expect that the resource windfall could lead to a heightened con-



(a) Dynamic effects on export value

(b) Dynamic effects on growth of export value

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (1.8), where the dependent variables are the regional export value in level (US$ millions) and log, respectively, in year $t = 1997, ..., 2019$. The regressions include micro-regions fixed effects and state-year fixed effects. Standard errors are adjusted for 129 meso-region clusters and the observations are weighted by total exports in 2000.

Figure 1.10: Dynamic effects of the resource boom on export value and growth

centration of exports in products experiencing surging demand following China's WTO accession. Alternatively, increased demand for certain products could drive greater diversification of the export basket, potentially facilitated by backward linkages in the production structure.

Table 1.2 presents the results from the estimation of Equation (1.7), focusing on changes in the number of export lines and the HHI between 2000 and 2019 as dependent variables. Across the two main specifications, the relationship between exposure to the resource boom and the number of export lines is negative but does not reach statistical significance. This suggests that the regions most affected by the export demand shock did not alter the number of products they exported between 2000 and 2019 compared to less impacted regions.

However, the analysis of the regional export basket concentration index reveals significant effects of the resource boom. Regions most impacted by the shock experienced an increase in the HHI of their export baskets relative to less affected regions. The estimates in column 2 indicate that a micro-region at the 75th percentile of exposure to Chinese demand ($\Delta\tilde{X}r$ = 0.108) saw, on average, an increase of 0.008 in the HHI associated with its export basket compared to a micro-region at the 25th percentile ($\Delta\tilde{X}r$ = 0.015).

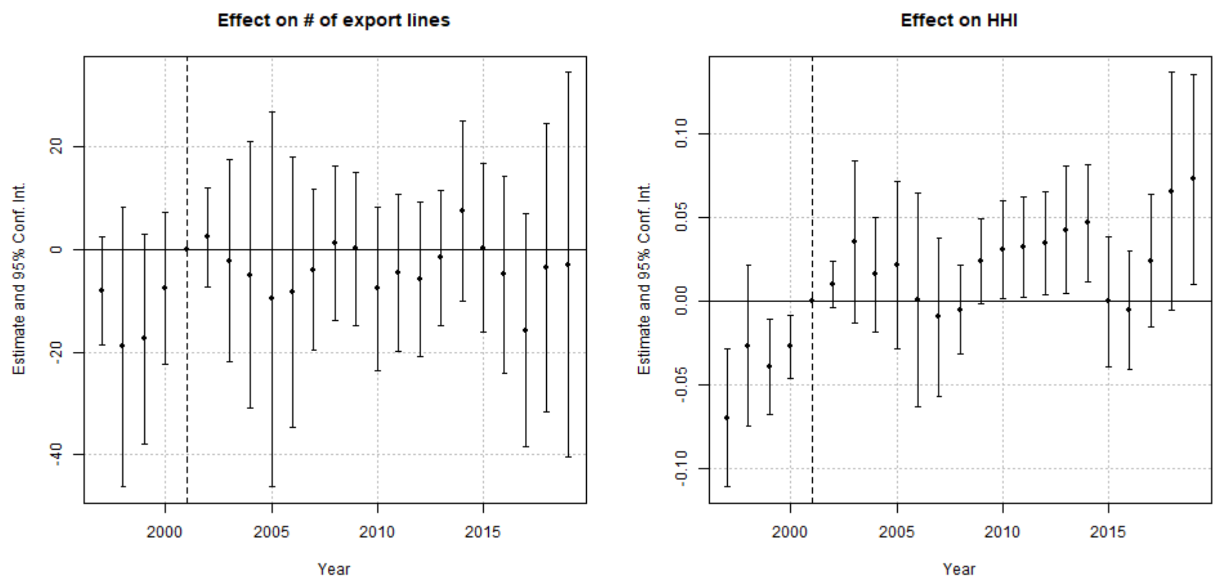
These findings suggest that the concentration of the export basket in regions most affected by the resource boom occurred primarily on the intensive margin rather than the extensive margin. In other words, while the number of products exported by these regions remained relatively stable, the export basket itself became more concentrated. Consequently, the increase in total export value in these regions was concentrated in a few products that were already part of the export repertoire, significantly increasing their share in the total export basket.

Table 1.2: Commodity boom and export concentration: number of lines and HHI

Dependent variable:	**Δ Lines**			**Δ HHI**		
	OLS	OLS	2SLS	OLS	OLS	2SLS
	(1)	(2)	(3)	(1)	(2)	(3)
$\Delta\tilde{X}_r$	3.658	-16.494		0.085	0.083***	
	(30.516)	(29.669)		(0.052)	(0.029)	
ΔX_r			-14.417			0.073***
			(25.858)			(0.026)
Weighted	X	X	X	X	X	X
State-year fixed effects		X	X		X	X
Observations	424	424	424	424	424	424
Adjusted R^2	-0.002	0.191	0.117	0.053	0.324	0.249
KP F-stat			99.7			99.7

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

To further validate the causal interpretation of these results, I re-estimate the event-study specification provided by Equation (1.8) for the measures of export concentration. Figure 1.11 illustrates the dynamic effects of the China-induced demand shock on export basket concentration, complete with 95% confidence intervals.



(a) Dynamic effects on export lines

(b) Dynamic effects on HHI

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (1.8), where the dependent variables are the number of exported lines and the HHI associated with regional export baskets, respectively, in year $t = 1997, ..., 2019$. The regressions include micro-regions fixed effects and state-year fixed effects. Standard errors are adjusted for 129 meso-region clusters and the observations are weighted by total exports in 2000.

Figure 1.11: Dynamic effects of the resource boom on export concentration

Similar to the findings on export value, the results depicted in Figure 1.11 are consistent with those presented in Table 1.2. The overall insignificance of pre-treatment coefficients for the years leading up to the trade shock further supports the validity of the parallel trends assumption.

1.5.2 Heterogeneity analysis

To address this, I conduct a heterogeneity analysis by categorizing local export baskets into resource-based and non-resource-based segments, following the classifications by Lall (2000). I then re-estimate the previous models for each category to discern the differential impacts on both the relative value of total exports and the concentration of regional export baskets. This approach allows me to determine whether the surge in export values and changes in export basket concentration are predominantly associated with resource-intensive products or a broader array of goods.

Table 1.3 presents the disaggregated results by broad type of export basket, focusing on variations in the total value of exports and the growth rate of these values from 2000 to 2019. The findings indicate that the increases documented in Table 1.1 are predominantly driven by the expansion of resource exports. Notably, there is no significant impact on the non-resource export basket in the regions most affected by the trade shock compared to less affected regions.

Table 1.3: Commodity boom and export value - Resource and non-resource baskets

Dependent variable:	Δ **Value of exports (US$ millions)**			$\%\Delta$ **Value of exports**		
	OLS	OLS	2SLS	OLS	OLS	2SLS
	(1)	(2)	(3)	(1)	(2)	(3)
Panel A: Resource basket						
$\Delta\tilde{X}_r$	2,747.192**	2,366.870**		1.405	1.449*	
	(1,302.870)	(940.415)		(1.219)	(0.819)	
ΔX_r			2,228.851**			1.365*
			(870.496)			(0.750)
Observations	406	406	406	406	406	406
Adjusted R^2	0.1883	0.710294	0.290033	-0.000	0.071	0.071
KP F-stat			94.1			94.1
Panel B: Non-resource basket						
$\Delta\tilde{X}_r$	504.032	70.895		0.203	-0.730	
	(437.353)	(364.996)		(0.364)	(1.122)	
ΔX_r			55.404			-0.570
			(277.034)			(0.852)
Observations	314	314	314	314	314	314
Adjusted R^2	0.034	0.265	0.264	-0.001	0.071	0.075
KP F-stat			82.5			82.5
Weighted	X	X	X	X	X	X
State-year fixed effects		X	X		X	X

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 1.4 presents results for export basket concentration. The impact on the HHI occurs primarily within the resource basket of the most affected regions, though with some variability compared to Table 1.2.

Next, I explore how the composition of the export basket in the regions most affected by the China-induced trade shock evolved between 2000 and 2019. To this end, I estimate Equation (1.7) using the share of total export value for each product sub-category, as defined by Lall (2000), as the

Table 1.4: Commodity boom and export concentration - Resource and non-resource baskets

Dependent variable:	Δ **Lines**			Δ **HHI**		
	OLS	OLS	2SLS	OLS	OLS	2SLS
	(1)	(2)	(3)	(1)	(2)	(3)
Panel A: Resource basket						
$\Delta\tilde{X}_r$	-10.209	-15.654		0.065	0.058***	
	(10.562)	(15.032)		(0.045)	(0.022)	
ΔX_r			-14.741			0.054**
			(13.923)			(0.023)
Observations	406	406	406	402	402	402
Adjusted R^2	0.004	0.165	0.060	0.030	0.429	0.353
KP F-stat			94.1			93.1
Panel B: Non-resource basket						
$\Delta\tilde{X}_r$	44.253	-11.653		-0.029	0.103	
	(50.747)	(21.497)		(0.093)	(0.085)	
ΔX_r			-9.107			0.081
			(14.816)			(0.080)
Observations	314	314	314	292	292	292
Adjusted R^2	0.035	0.348	0.346	-0.000	0.301	0.238
KP F-stat			82.5			76.2
Weighted	X	X	X	X	X	X
State-year fixed effects		X	X		X	X

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

dependent variable. The outcomes, summarized in Table 1.5, suggest that there was no substantial aggregate structural change in the relative composition of the export basket between the regions most and least affected by the shock. This stability in export composition is expected, considering that at the highly disaggregated level of analysis, regions with a comparative advantage *ex-ante* generally continued to focus on exporting similar products. This is a distinctive aspect of the quasi-natural experiment used here: unlike previous studies evaluating resource windfalls characterized by the discovery of substantial natural reserves, this shock amplified the pre-existing advantages of historically resource-exporting regions rather than altering their comparative advantages.

While the overall share of resource exports remained stable, the composition within these categories appears to have shifted relatively in the most impacted regions. Despite modest statistical significance, there is a relative reduction in the share of resource-based manufactures, coupled with an increase in the share of primary product exports in the localities more affected by the

Table 1.5: Commodity boom and export shares

Classification	Share of export value	Estimated coefficients	
	Average 2000-2019	OLS 2000-2019	2SLS 2000-2019
Resource	**0.639**	**-0.005**	**-0.004**
		(0.018)	**(0.016)**
Primary	0.326	0.070	0.061
		(0.048)	(0.039)
Resource-based	0.313	-0.075*	-0.066*
		(0.043)	(0.035)
Manufactures	**0.361**	**0.005**	**0.004**
		(0.018)	**(0.016)**
Low-Tech	0.072	0.011	0.009
		(0.013)	(0.011)
Medium-Tech	0.227	-0.005	-0.004
		(0.024)	(0.021)
High-Tech	0.067	-0.001	-0.001
		(0.003)	(0.002)
Observations		424	424
KP F-stat			99.7

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In all regressions, observations are weighted and state-year fixed effects are added. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

shock compared to less impacted economies. Essentially, while the overall balance between resources and manufactures stayed largely stable in the regions most affected by the shock, there is evidence of a "primarization" within the resource basket. This trend reflects a shift from exporting more complex, resource-based manufactures to simpler, primary products. For instance, regions that exported soybean oil in the early 2000s may have transitioned to exporting raw soybeans, indicating a move toward less complex items within the same production chain.

This shift towards "primarization" suggests a potential decline in the sophistication or complexity of the products exported by these regions over the past two decades. This trend could significantly impact long-term economic diversification and value addition in the affected regions. I explore these implications further in the following subsection.

1.5.3 Export sophistication

To examine whether the export demand shock led to changes in the complexity of export baskets, I use the established export sophistication measure outlined in Equation (1.4) (Hausmann et al., 2007; Jarreau & Poncet, 2012). This measure helps determine if shifts in the concentration and composition of local Brazilian export baskets, induced by the export demand shock, are associated with a decline in the average complexity of exported goods.

I estimate Equation (1.7), with the change in the export sophistication index, $\Delta S_{r,t}$, for each micro-region as the dependent variable. The analysis is conducted using two distinct real income per capita measures from the World Development Indicators (WDI) and the Penn World Table (PWT). The findings, presented in Table 1.6, reveal a relative decrease in the sophistication index of the export baskets in micro-regions most affected by the increased Chinese demand compared to less impacted areas.

The estimates in column 2 indicate that a micro-region at the 75th percentile of exposure to Chinese demand experienced, on average, a decline of just over 30 points in the sophistication index of its export basket compared to a micro-region at the 25th percentile of the shock distribution. This decline reflects a shift toward less complex, lower-value-added activities. Similar to Heresi (2023), this shift can be linked to a reduction in average value added and productivity within the sector following a resource boom. However, in this case, the effect is observed within the resource or booming sector itself, highlighting an additional channel through which resource booms can contribute to overall productivity declines. Figure A.1 illustrates the combined direct and indirect effects of the resource boom, showing how these dynamics may lead to reductions in total factor productivity in the most exposed localities. Given the central role of average productivity in sustaining long-run growth, this additional mechanism suggests that the commodity boom may, in fact, hinder long-term economic development in these regions.

As with the analyses of export value and concentration, I further evaluate the causal implications of these findings on export sophistication using the event-study approach specified in Equation (1.8). Figure 1.12 illustrates these dynamic effects, reinforcing the findings in Table 1.6. The

Table 1.6: Commodity boom and export sophistication

Dependent variable:	Δ **Export sophistication (WDI)**			Δ **Export sophistication (PWT)**		
	OLS	OLS	2SLS	OLS	OLS	2SLS
	(1)	(2)	(3)	(1)	(2)	(3)
$\Delta \tilde{X}_r$	35.522	-327.907*		6.714	-371.524*	
	(89.544)	(179.980)		(83.741)	(190.815)	
ΔX_r			-286.619*			-324.743**
			(156.630)			(164.629)
Weighted	X	X	X	X	X	X
State-year fixed effects		X	X		X	X
Observations	424	424	424	424	424	424
Adjusted R^2	-0.001	0.248	0.238	-0.002	0.146	0.165
KP F-stat			99.7			99.7

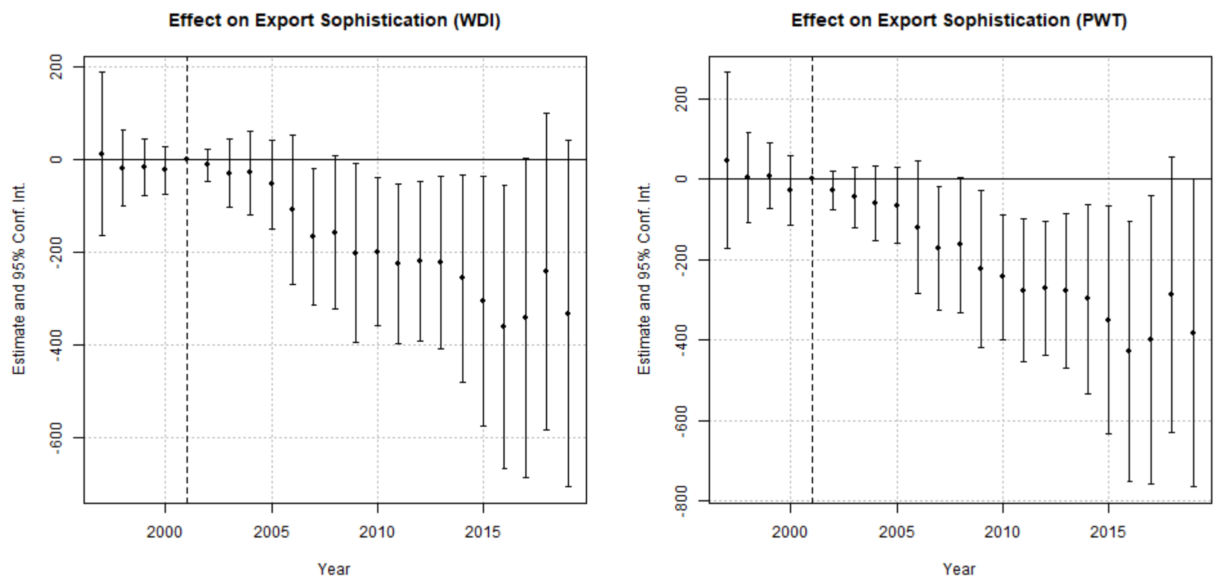
Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta \tilde{X}_r$ as the IV for ΔX_r. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

figures show a clear post-treatment decline in export sophistication for the most affected regions, with pre-treatment coefficients remaining largely insignificant. This pattern supports the validity of the parallel trends assumption and strengthens the causal link between the export demand shock and reduced export complexity.

1.5.4 Employment variables

The findings presented thus far suggest a degree of structural change, particularly in the export baskets of regions most affected by the resource boom compared to those less impacted. This subsection extends the analysis to examine whether these changes in export dynamics are mirrored in local labor markets, with a focus on sectoral employment composition and wage variations.

To explore these labor market effects, I adapt the shift-share instrument from Equation (1.6) to capture export demand growth from 2000 to 2010. This period represents the most recent year for which comprehensive micro-data from the Brazilian Demographic Census are available. This adaptation allows for an evaluation of whether the China-induced export demand shocks, as captured by $\Delta \tilde{X}_r$, led to sectoral employment shifts or wage adjustments, as economic theory would predict.



(a) Dynamic effects on sophistication (WDI)

(b) Dynamic effects on sophistication (PWT)

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (1.8), where the dependent variables are the sophistication indexes associated with regional export baskets as described in Equation (1.4) in year $t = 1997, ..., 2019$. The regressions include micro-regions fixed effects and state-year fixed effects. Standard errors are adjusted for 129 meso-region clusters and the observations are weighted by total exports in 2000.

Figure 1.12: Dynamic effects of the resource boom on export sophistication

Building on Costa et al. (2016), I use the modified shift-share instrument to investigate the impacts of Chinese demand shocks on local labor market outcomes in Brazil. Table 1.7 presents regression results at the micro-region level, focusing on changes in log average hourly wages and private sector employment rates between 2000 and 2010. Notably, regions most affected by the resource boom show no significant variation in log average hourly wages compared to less impacted localities. The effect of the Chinese demand shock on the aggregate employment rate is positive but statistically significant only at the 10% level in specifications that include state-year fixed effects.

Next, I assess changes in local employment composition across sectors by estimating the specification from Equation (1.7), using sectoral employment shares as the dependent variables. The literature on the resource curse, particularly studies on the Dutch disease, suggests that trade shocks, such as this commodity boom, might lead to a labor shift toward booming sectors at the expense

Table 1.7: Commodity boom, employment and remuneration - aggregate results

Dependent variable:	Δ Log average hourly wages			Δ Employment share		
	OLS	OLS	2SLS	OLS	OLS	2SLS
	(1)	(2)	(3)	(1)	(2)	(3)
$\Delta\tilde{X}_r$	0.058*	0.024		0.023	0.026*	
	(0.034)	(0.019)		(0.016)	(0.013)	
ΔX_r			0.011			0.012*
			(0.009)			(0.006)
Weighted	X	X	X	X	X	X
State-year fixed effects		X	X		X	X
Observations	424	424	424	424	424	424
Adjusted R^2	0.008	0.625	0.623	0.007	0.335	0.342
KP F-stat			751.2			751.2

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

of manufacturing employment. Table 1.8 presents the results for changes in sectoral employment shares between 2000 and 2010 across Brazilian micro-regions.

In Panels A and C of Table 1.8, no statistically significant impact is observed on the share of workers in the primary and services sectors, respectively, in regions experiencing relatively larger increases in Chinese demand. Although the estimated coefficients point in the expected direction, the absence of significant changes in the employment share of the booming sector could be attributed to the labor-saving technologies increasingly adopted in commodity production, such as in the soybean industry (Bustos et al., 2016, 2020; Farrokhi & Pellegrina, 2023; Pellegrina, 2022). In contrast, Panel B reveals a statistically significant decline in the share of manufacturing employment in regions more affected by the trade shock compared to less impacted counterparts.

These results partially align with the Dutch disease literature. While there is no significant increase in the share of workers in the booming sector in regions most affected by the export demand shock, there is a notable decline in manufacturing employment. Importantly, Panel D indicates that the adjustment appears to have occurred within the residual category of our sector classification — individuals reporting income from employment but without specifying occupations. These posi-

Table 1.8: Commodity boom and sectoral employment patterns

Dependent variable: Δ **Employment share**			
	OLS	OLS	2SLS
	(1)	(2)	(3)
Panel A: Primary sector			
$\Delta\tilde{X}_r$	0.048	0.032	
	(0.029)	(0.020)	
ΔX_r			0.014
			(0.009)
Adjusted R^2	0.0208	0.283	0.287
F Statistic			751.2
Panel B: Manufacturing sector			
$\Delta\tilde{X}_r$	-0.014	-0.020**	
	(0.013)	(0.008)	
ΔX_r			-0.009**
			(0.004)
Adjusted R^2	0.005	0.172	0.152
F Statistic			751.2
Panel C: Services (nontraded) sector			
$\Delta\tilde{X}_r$	-0.038*	-0.022	
	(0.020)	(0.018)	
ΔX_r			-0.010
			(0.008)
Adjusted R^2	0.013	0.326	0.335
F Statistic			751.2
Panel D: Residual sector			
$\Delta\tilde{X}_r$	0.005	0.010**	
	(0.007)	(0.004)	
ΔX_r			0.005**
			(0.002)
Adjusted R^2	0.001	0.353	0.353
F Statistic			751.2
Observations	424	424	424
Weighted	X	X	X
State-year fixed effects		X	X

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

tions are likely more closely linked to the informal labor market than those in other sectors. Thus, the reduction in manufacturing employment is accompanied by a modest increase in the share of individuals in these unspecified positions in the regions most impacted by the shock.

Table 1.9 further investigates changes in average wages across various sectors by applying long differences to log average wages in the primary, manufacturing, service , and the residual sectors. The results reveal a significant wage impact of the Chinese demand shock, particularly in the primary and service sectors. This suggests that while the resource boom predominantly raises wages in directly affected sectors, such as agriculture and extractive industries, it also has positive spillover effects on wages in indirectly linked sectors.

Overall, the findings on employment dynamics following the China-induced export demand shock suggest that the significant increase in the value and concentration of primary product exports in the most affected regions does not correspond with a proportional rise in employment shares. Instead, the primary impact manifests in wage dynamics. Notably, there is a marked increase in average hourly wages within the primary sector, reflecting wage pressures driven by heightened demand. Additionally, consistent with Corden and Neary (1982) and Corden (1984), there is a significant rise in service sector wages in regions most impacted by the export demand surge, indicating positive income spillovers from the primary to non-traded sectors.

1.6 Concluding remarks

This paper examines the impact of resource booms on local export baskets in Brazil, a key example of a resource-rich developing country, focusing on export value, concentration, and composition. By leveraging a shift-share instrument that captures heterogeneous exposure to Chinese export demand following China's WTO accession in 2001, I analyze how this resource boom affected regional export dynamics across Brazilian local economies. The findings reveal significant increases in total export value and heightened concentration in export baskets within the most impacted regions. Contrary to initial expectations, the share of manufacturing exports in these regions

Table 1.9: Commodity boom and sectoral remuneration patterns

Dependent variable: Δ **Log average hourly wages**	OLS (1)	OLS (2)	2SLS (3)
Panel A: Primary sector			
$\Delta\tilde{X}_r$	0.489*	0.322***	
	(0.259)	(0.079)	
ΔX_r			0.146***
			(0.037)
Adjusted R^2	0.021	0.283	0.287
F Statistic			751.2
Panel B: Manufacturing sector			
$\Delta\tilde{X}_r$	0.036	-0.018	
	(0.036)	(0.030)	
ΔX_r			-0.008
			(0.013)
Adjusted R^2	0.000	0.354	0.355
F Statistic			751.2
Panel C: Services (nontraded) sector			
$\Delta\tilde{X}_r$	0.085***	0.052***	
	(0.030)	(0.016)	
ΔX_r			0.023***
			(0.007)
Adjusted R^2	0.023	0.682	0.718
F Statistic			751.2
Panel D: Residual sector			
$\Delta\tilde{X}_r$	0.021	-0.020	
	(0.086)	(0.054)	
ΔX_r			-0.009
			(0.024)
Adjusted R^2	-0.001	0.176	0.176
F Statistic			751.2
Observations	424	424	424
Weighted	X	X	X
State-year fixed effects		X	X

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

did not significantly decline; instead, the increased concentration is driven by a focus on a narrower range of previously exported products, with minimal changes in the overall variety of goods.

The heterogeneity analysis indicates that these aggregate changes are primarily driven by shifts within the resource basket rather than the non-resource basket, highlighting a nuanced pattern of specialization across regions. Specifically, there is a shift toward exporting primary products at the expense of resource-based manufactures in the regions most affected by the surge in Chinese demand.

To further explore this "primarization" effect, I constructed an index of export basket sophistication, drawing upon the methodologies of Hausmann et al. (2007) and Jarreau and Poncet (2012). The analysis shows a decline in the average complexity of export baskets in regions more impacted by the shock compared to less affected areas. The shift toward simpler, lower-value-added goods — favoring raw materials over processed products — reveals another channel through which commodity booms reshape local productive structures. Similar to Heresi (2023), I document a within-sector adjustment, where a decline in value-added production within the resource or booming sector may be linked to a reduction in average productivity. This mechanism, if it indeed affects total factor productivity, could have lasting developmental consequences for these local economies and the broader Brazilian economy. Future research could focus on developing a theoretical framework to more comprehensively analyze the implications of such specialization patterns on regional growth trajectories.

Beyond export dynamics, this study also examines the broader implications of this trade shock for structural change, particularly in local labor markets. Despite the shift to a more "primary" export orientation, there was no corresponding increase in employment within the primary sector. Instead, this sector experienced substantial wage increases, suggesting that labor market adjustments occurred primarily along the intensive margin rather than through large-scale reallocation of workers. Additionally, relatively higher wage growth was observed in the service sector in regions more exposed to the China-induced trade shock. Meanwhile, manufacturing employment contracted significantly in the most affected regions, consistent with the Dutch disease phenomenon discussed in the resource curse literature. This contraction in manufacturing employment was partially offset by an increase in employment within a residual category — likely linked to the informal labor market — suggesting that resource booms may not only induce structural shifts in formal employment but also contribute to the expansion of less stable, informal economic activities.

In sum, the resource boom driven by Chinese demand has considerably reshaped regional economic dynamics in Brazil. While it substantially increased export values and boosted wages in certain sectors, it also led to a structural shift toward simpler, lower-value-added exports, with adverse effects on manufacturing employment. These findings offer a more nuanced perspective on the impacts of the China-led export demand shock in Brazilian labor markets, expanding on the evidence provided by Costa et al. (2016) and uncovering deeper structural changes that may hinder long-term economic diversification. These insights are crucial for policymakers in resource-rich developing countries, underscoring the need for strategies that not only capitalize on the immediate benefits of resource booms but also address the sustained pressures of maintaining diversified productive structures at the regional level.

Furthermore, the implications of this study extend beyond Brazil, offering broader lessons for other resource-rich developing economies facing similar external shocks. As recent evidence from Colombia suggests (Branstetter & Laverde-Cubillos, 2024), resource booms often come with a decline in technological development and long-term competitiveness. The patterns observed in Brazil may be part of a larger global trend, underscoring the urgent need for policies that foster economic diversification and technological upgrading.

By way of conclusion, while this study provides new causal evidence on resource booms and regional export dynamics, further research is needed to fully understand their long-term growth and development implications. Although this analysis captures additional partial equilibrium effects of China's WTO accession on the Brazilian economy — extending the limited literature, including works by Costa et al. (2016) and Carreira et al. (2024) — the findings presented here broaden the scope of existing research on developing economies and suggest new avenues for future investigation.

Chapter 2

Donkey business: trade, resource exploitation, and crime

[7]

2.1 Introduction

Natural resources account for over 20% of global goods exchange, and their sustainable use is becoming increasingly critical as countries around the world grapple with resource depletion and overexploitation (Ruta & Venables, 2012). In many renewable resource markets, unclear property rights and weak regulatory enforcement lead to overuse, a pressure often exacerbated by international trade (Copeland & Taylor, 2009). This dynamic raises concerns about the long-term sustainability of natural resource use, particularly in regions where regulation is insufficient or inconsistently enforced, leading to both economic and ecological consequences.

One timely and notable example of these pressures is the rising demand for *ejiao*, a Traditional Chinese Medicine product made from donkey hides, which has sparked a global trade that significantly impacts donkey populations and local economies, particularly in low- and middle-income countries (Lesté-Lasserre, 2019). Once a luxury item reserved for China's elite, *ejiao* has gained popularity among the broader middle class due to rising incomes, creating a shortage of donkeys within China (Sheikh & Lohre, 2023; Skippen et al., 2021; Waters, 2019). In this context, Brazil has emerged as a key supplier of donkey hides, especially after regulating this trade in 2017 and establishing export-oriented slaughterhouses. However, Brazil faces a pressing challenge: the ab-

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[7] This chapter is co-authored with Lucas Corrêa-Dias. A working paper version of this essay can be found here. We benefited from fruitful discussion with and useful comments and impressions from Ariaster Chimeli, Daniel da Mata, Daniele Tavani, Elissa Braunstein, Iasmin Goes, João Paulo Pessoa, Pedro Amoni, Rafael Araujo, Rafael Dix-Carneiro, Ramaa Vasudevan, Timothy Komarek, Sammy Zahran and Vítor Possebom. We are thankful for seminar participants at the NEREUS-USP, São Paulo School of Economics, University of Wyoming, and Colorado State University. We are grateful to Daniel Wietzel for comments and code suggestions. Any remaining errors are our own.

sence of clear property rights for donkeys, many of which roam freely in semi-arid regions with high poverty levels. Once vital to local communities, modernization and economic changes have diminished the donkeys' value, leading to increased abandonment. The potential for quick profits through their sale to slaughterhouses has complicated ownership claims, leading to conflicts over these donkeys and raising concerns about violence and the sustainability of this trade.

We examine the relationship between contestable markets – particularly those with poorly defined or enforced property rights – and crime, focusing on the direct consequences of institutionalizing the donkey hide trade on crime and violence dynamics at the local level in Brazil. Three main features contribute to the escalation of crime around this market: (i) the emergence of dealers and intermediaries' agents that fuel the supply of free-roaming donkeys to slaughterhouses, (ii) the legal status of slaughtering activities, which were under dispute in the Brazilian judicial system and considered illegal in some periods, and (iii) the association of donkey hide trade with other criminal activities, such as parallel trafficking[8](Sheikh & Lohre, 2023; The Donkey Sanctuary, 2022). This research addresses the question: What impact does regulating a contestable market have on local crime rates? Our inquiry seeks to understand the socio-economic implications of market regulation, particularly the negative externalities associated with the donkey hide trade.

To explore this question, we first document the increase in donkey slaughters following the regulation of the market. Before the rise in Chinese demand, there were almost no recorded slaughters, but they surged to over 8,000 per month at their peak, according to official data. These numbers dropped sharply after Brazilian authorities imposed a ban, followed by a significant increase in the slaughter of "other equids", such as horses and mules, exclusively in the region where the donkey market was active. Donkey slaughters resumed as soon as the ban was lifted, and trade statistics show that the export boom of equids' raw skin and hides remained largely unaffected by regulatory changes. This pattern, along with extensive anecdotal evidence, suggests that slaughterhouses may have operated illegally during the national ban.

[8]For instance, Su et al. (2022) show that at least 13 groups of CITES-listed species are involved in the densely connected donkey skin product network, facilitated through online wildlife trade.

Leveraging the timing of national regulation of the donkey market, combined with variations in the historical presence of donkeys across Brazilian municipalities, we use a difference-in-differences research design to estimate the causal impact of the donkey market on violence. Our results show a positive effect of the donkey market on crime rates, measured by homicides, in Brazilian municipalities. Regions with higher donkey populations before the regulation experienced relatively higher crime rates after the institutionalization of the market. To put this in perspective, a one-standard-deviation increase in donkeys per thousand inhabitants corresponds to a quarterly increase of approximately 1.31 homicides per 100,000 inhabitants in Northeast municipalities. Moreover, we find that this effect is stronger for municipalities closer to a slaughterhouse, and that there are null effects for other climate-resilient livestock, strengthening the causal interpretation of our results. These findings are consistent across various samples, time periods, alternative treatment variables, and remain robust when controlling for different crime dynamics in municipalities with varying initial characteristics.

Finally, we study the role of the illegality of the market in criminal dynamics. Using judicial decisions that imposed a national ban on slaughtering, which was later lifted, we apply a triple-difference strategy and find that illegality plays a major role in driving the observed effects. Our estimates indicate that the effect on crime was twice as large during the illegal period. These results, however, do not diminish the role of open access and lack of property rights over the animals in driving violence, as the effects remain positive and both economically and statistically significant during the period of legal slaughter. In light of this, we conclude that both resource booms and illegality in such contestable markets play a critical role in explaining the effects on crime and violence.

Our findings make at least three contributions to the existing literature. First, we contribute to the broad literature on trade in natural resources and economic development (Ruta & Venables, 2012). Specifically, we add to the literature on the consequences of (wildlife) renewable resource overuse (Brander & Taylor, 1998; Copeland & Taylor, 2009; Taylor, 2011) by demonstrating an unintended welfare loss associated with a pervasive feature of such markets – open access to the resources. Previous studies document externalities associated with wildlife depletion. For instance, Feir et al. (2024) show that the near extinction of the North American bison[9] had long-lasting effects on populations that relied on them, while Frank and Sudarshan (2024) argue that the collapse of vultures in India had important consequences for human health. We show that the institutionalization of a new wildlife market, combined with open access and resource overuse, may lead to increased crime in resource-abundant regions.

Second, we provide empirical evidence on the impact of regulatory interventions in a natural-resource market on local crime rates, shedding light on an underexplored aspect of contestable market operations. Prior studies have uncovered correlations between regulatory interventions in illegal drug markets and subsequent spikes in violence (Angrist & Kugler, 2008; Castillo et al., 2020; Dell, 2015; Dube et al., 2016; Mejıa & Restrepo, 2011). Another strand of the literature suggests that illegal markets often foster competition, leading to violence among competitors and within organizations (Adda et al., 2014; Dragone et al., 2019; Owens, 2014; Reuter, 2009). In the case of natural resources, Chimeli and Soares (2017) and Pereira and Pucci (2024) find that violence is associated with the illegal status of logging and mining in the Brazilian Amazon, respectively. We find a positive effect on crime following the institutionalization of the donkey market. While prohibition and illegal activity intensify this effect, it remains significant even during periods of legal market operations. This suggests that crime can result from markets with poorly defined property rights, even when these markets are not strictly illegal.

Finally, we uncover a previously underexplored consequence of China's rise in global trade, particularly in its effects on developing countries most impacted by the exploitative depletion of natural resources. While previous research has focused on the impact of Chinese import competition on variables such as manufacturing employment (Autor et al., 2013; Pierce & Schott, 2016) and workers' earnings (Autor et al., 2014; Dix-Carneiro et al., 2023), there is less work examining the demand-side effects of China's economic ascendancy, especially on natural-resource-based markets. Concerning the *ejiao* market and donkey hide trade, other studies show detrimental effects

[9] For a detailed discussion on this process, see the seminal contribution in Taylor (2011).

on rural livelihoods in African countries where donkeys are important assets (Carder et al., 2019; Goodrum et al., 2022; Maggs et al., 2023). Our findings, however, reveal an important resource-boom effect: international demand has made previously worthless animals valuable. Combined with open access, excessive hunting, and weak enforcement in Brazil, crime emerges as an unintended consequence of this boom and the operation of the market.

These findings carry significant policy implications, underscoring persistent externalities in the trade-resource depletion relationship. This pattern of unintended welfare consequences from over-exploitation recalls the historical decline of the North American buffalo over a century ago (Feir et al., 2024; Taylor, 2011) and aligns with research linking commodity price shocks to crime (Dube & Vargas, 2013; Komarek, 2018). Similar resource booms are emerging across many developing economies, notably in the donkey hide trade. As governments face challenges in implementing and enforcing effective regulations, our results suggest that global coordination may be essential to mitigating these social costs.

The remainder of this paper is organized as follows: Section 2.2 provides background information on the *ejiao* market, detailing the rise of the donkey hide market in Brazil and its regulatory framework. Section 2.3 describes the data used in our empirical analysis and outlines our empirical strategy. In Section 2.4, we present the main results and robustness checks. Section 2.5 offers a falsification exercise, while Section 2.6 explores the validity of our proposed mechanism. Finally, Section 2.7 concludes with a discussion of the implications and policy relevance of our findings.

2.2 Background

2.2.1 *Ejiao* and donkeys' hide trade

Donkey products have long been revered in Traditional Chinese Medicine for their purported medicinal and beautifying properties. Central to this is *ejiao*, a gelatin made from donkey hides, believed to offer health benefits like blood enrichment, immune enhancement, and cancer prevention. Despite its long-standing tradition and perceived efficacy, empirical evidence supporting these claims is lacking (Li et al., 2017; Wang et al., 2012).

Originally reserved for the elite in Chinese society, *ejiao* has witnessed a notable shift in its consumer base, predominantly due to a substantial increase in average income in China. As a result, the product now primarily targets the burgeoning middle class, leading to a significant surge in demand in recent years (Köhle, 2018; Lesté-Lasserre, 2019). This escalating demand has overwhelmed the domestic supply of donkeys in China (Sheikh & Lohre, 2023; Skippen et al., 2021; Waters, 2019). Data from the Food and Agricultural Organization of the United Nations (FAO) indicate a drastic decline in the Chinese donkey population from approximately 11 million in 1994 to less than 2.7 million in 2018. To sustain the *ejiao* market in China alone, an estimated 5-6 million donkeys annually are required (The Donkey Sanctuary, 2022), a figure far surpassing China's current population.

Consequently, suppliers have turned to sourcing donkey hides from developing countries, particularly in South America and Africa, where the donkey hide trade has flourished.[10] Donkeys are slaughtered in these regions, and their skins are exported to Chinese factories for *ejiao* production. This global trade has put immense pressure on donkey populations, sparking concerns over sustainability, animal welfare, and the livelihoods of communities reliant on these animals (Bennett & Pfuderer, 2020; Goodrum et al., 2022; Waters, 2019).

The protracted gestation period of donkeys, lasting up to 14 months, presents a significant obstacle to sustainable production, particularly amid the surge in demand. This slow reproductive rate, coupled with the exploitative nature of the trade, heightens the risk to donkey populations, particularly in developing regions. Beyond animal welfare concerns, the trade's impact extends to the communities that rely on donkeys for agriculture and transportation. Maggs et al. (2023) highlights the vital economic and social roles donkeys play in West African households, where their loss can severely disrupt livelihoods.

In addition to its direct effects, the donkey hide trade is often intertwined with broader wildlife trafficking. Hides are often smuggled alongside other banned wildlife products, a practice known

[10] Panel (a) of Figure B.1 illustrates the global distribution of donkey populations, while panel (b) provides examples of products derived from donkey hides. Also, see detailed coverage of the recent trend in Nature, The Economist, and The New York Times.

as parallel trafficking (Sheikh & Lohre, 2023; The Donkey Sanctuary, 2022). Scientists have also raised concerns about this trade facilitating the spread of zoonotic and other infectious diseases, further complicating the array of issues posed by the donkey hide industry (Goodrum et al., 2022; Skippen et al., 2021).

In response to growing concerns over the economic and social impacts of the trade, several governments – such as Burkina Faso, Ghana, Nigeria, and Senegal – implemented export restrictions on donkey products. However, these unilateral efforts were often undermined by illegal animal transport and hide smuggling across borders (Voice of America, 2022). Most recently, the African Union introduced a continent-wide ban on the donkey skin trade, which has been praised by animal welfare advocates for its potential to protect both donkey populations and the livelihoods of communities dependent on them.[11]

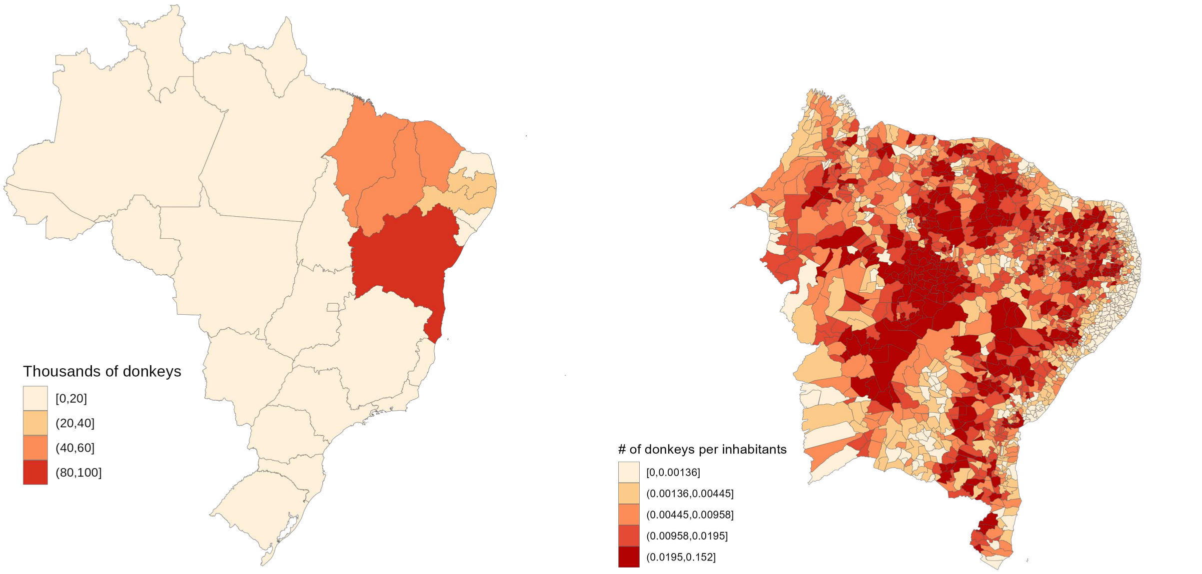
2.2.2 The donkey market in Brazil

The historical significance of donkeys in Brazil dates back to the 16th century when Portuguese colonizers introduced them to the Brazilian Northeast. Alongside other equid species, donkeys played a vital role in the region's development due to their resilience and adaptability to local environmental conditions. They became indispensable for various tasks such as packing, drafting, and transportation, particularly in the semi-arid regions of the Northeast (Carneiro et al., 2018; Salles et al., 2013).

This historical prominence has led to a highly concentrated distribution of donkey populations, with approximately 90% of all donkeys in Brazil located in the Northeast.[12] Panel (a) of Figure 2.1 shows the number of donkeys across Brazilian states, emphasizing the Northeast's concentration. Panel (b) shows the number of donkeys per inhabitant and zooms in on municipalities in the Northeast, particularly highlighting the cluster in the more arid regions.

[11]For more details, see coverage from BBC, The New York Times, Reuters, and The Atlantic.

[12]Figures B.2 and B.3 illustrate the spatial distribution of donkeys in Brazilian municipalities in 2017, highlighting this regional concentration.



(a) Number of donkeys

(b) Donkeys per capita - Northeast municipalities

Source: Data on donkey population is from the 2017 Census of Agriculture. Population data is from IBGE.

Figure 2.1: Spatial distribution of donkeys in Brazil

Notwithstanding, unlike many other developing countries, Brazil's Northeast has experienced a significant decline in the utility of donkeys as working animals over the past two decades. This decline is largely due to improvements in agricultural technology, greater access to credit, and rising living standards. As a by-product of regional development, the advent of motorized vehicles has further contributed to a decline in the utilization of donkeys, leading to instances of abandonment and uncontrolled breeding. This shift is underscored by anecdotal evidence, such as former President Luiz Inácio "Lula" da Silva's famous remark, months after leaving government with 87% of approval, that "the thing that makes us most proud is that folks are exchanging our beloved donkey for a motorcycle" (O Globo, 2011). As a result, free-roaming and feral donkey populations have proliferated in the semi-arid regions of Brazil, contributing to challenges for local governments, particularly an increase in traffic accidents caused by these animals (Carneiro et al., 2018; G1 - Globo, 2017b; Gameiro et al., 2022; Salles et al., 2013).

The existing population of donkeys, now freely roaming the arid regions of Northeast Brazil, has not escaped notice amidst the burgeoning market for donkey hide. In late 2015, then-Minister of Agriculture, Kátia Abreu, shared a noteworthy anecdote on Twitter.[13] She recounted that “at the entrepreneurs’ seminar, an investor caught our attention with an interest that seemed like a joke to us, but it wasn’t. He wants to export donkeys to China.” The Minister also expressed astonishment at the magnitude of the investor’s demand: “Unbelievable, but their demand is 1 million donkeys per year[...]” (Valor Econômico, 2015).

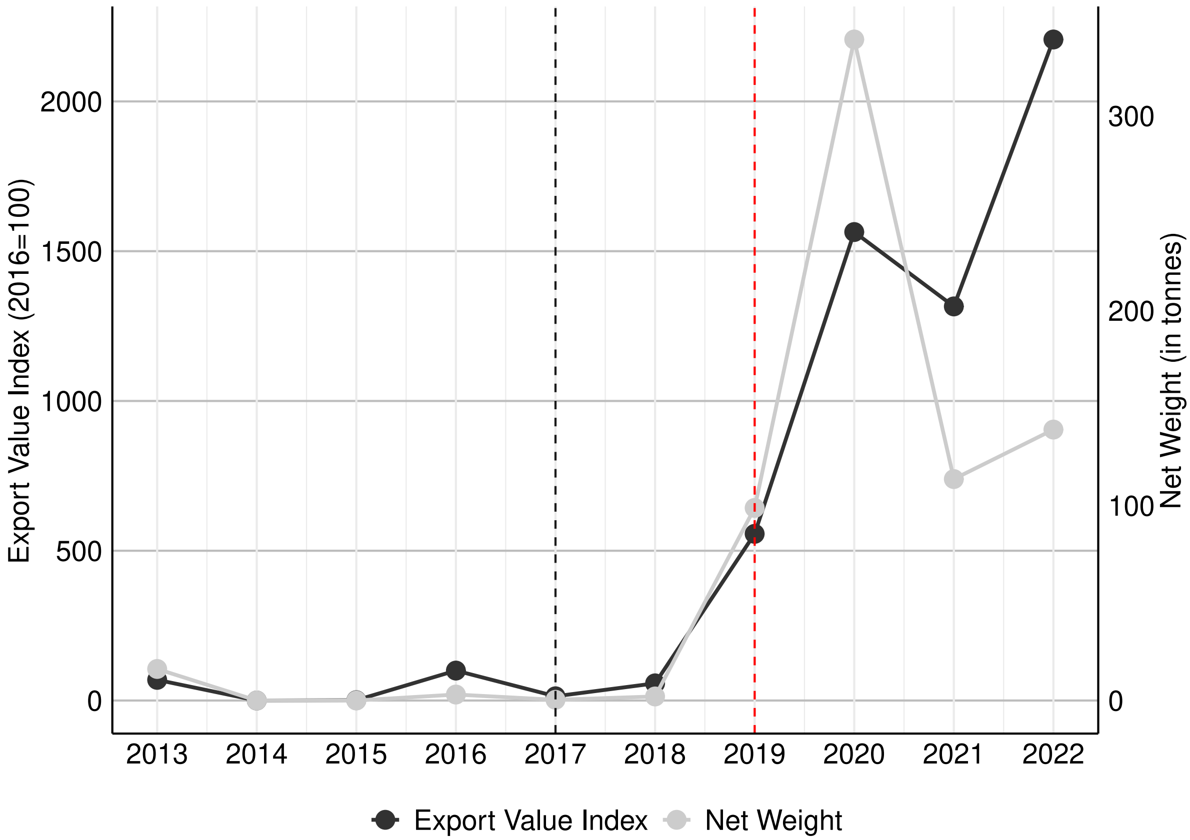
In 2016, public officials began using the presence of donkeys on roads as a key justification for legalizing donkey slaughter in Bahia for meat and hide export (Gameiro et al., 2021, 2022). Local government discourse framed this trade as a humane solution to the long-standing issue of roaming animals (Diário Oficial da Bahia, 2016). Subsequently, in 2017, the federal government established the legal framework for animal slaughter and transportation, marking the complete regulation of the activity in the country – under *Decreto 9.013, March 2017*. With local government support, three slaughterhouses linked to Chinese entrepreneurs were established in Bahia state between mid-2017 and 2018.[14]

As a result, exports of donkey products in Brazil saw considerable growth. Figure 2.2 illustrates the rise in both value and net weight of raw skins and hides from donkeys and other equids, likely materials used in *ejiao* production. Exports surged from around 3 tonnes in 2016 to approximately 340 tonnes by 2020, with an even steeper rise in export value. Concurrently, the number of slaughtered donkeys soared by 4,788% during the same period (Gameiro et al., 2021, 2022). Figure 2.3 shows the trajectory of slaughtered donkeys and other equids (horses and mules) in Brazil from 2013 to 2021. The reduction in Brazil’s donkey population over recent decades is notable. According to estimates from FAO Stat, the number of donkeys decreased by approximately 43% between 1996 (1.370 million) and 2022 (782 thousand). This decline underscores the impact of market forces and regulatory decisions on donkey demographics in Brazil.

With no established commercial production chain for donkeys in Brazil, sourcing them typically involves capturing animals from rural communities or directly from roads for slaughter

[13] The original tweet from then-Minister of Agriculture, Kátia Abreu, is accessible via this link.

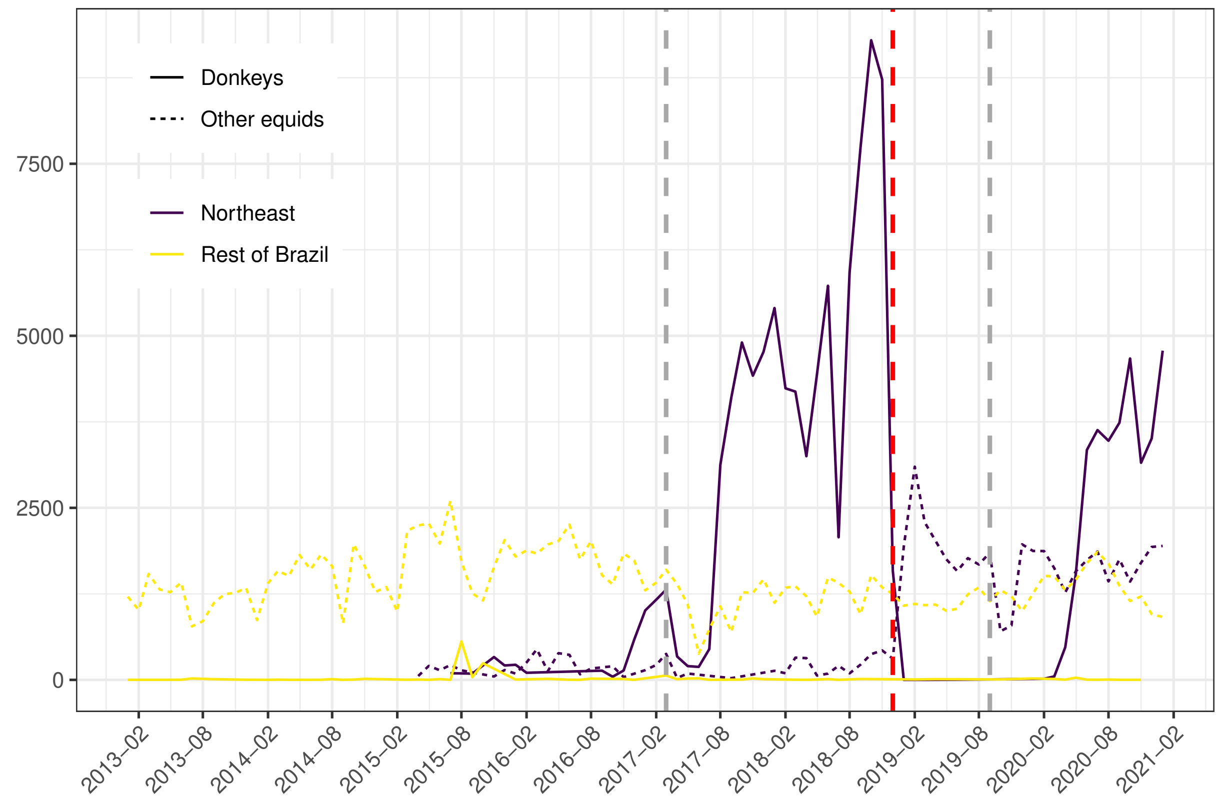
[14] Panel (c) of Figure B.1 shows free-roaming donkeys on a state road in Ceará, while panel (d) provides an example of a Chinese donkey slaughterhouse in Amargosa, Bahia.



Notes: Data on exports of raw skins and hides from donkeys and other equids are sourced from SISCOMEX, affiliated with Brazil's Ministry of Industry, Foreign Trade, and Services. The black dashed line represents the year of regulation of the donkey market, while the red dashed line marks the timing of the prohibition.

Figure 2.2: Exports of donkeys in Brazil

(Gameiro et al., 2022). This has led to the rise of a network of dealers and intermediaries who pay for the capture of donkeys, which are then sold to one of the few slaughterhouses (BBC-Brasil, 2021; UOL, 2023). Anecdotal evidence suggests that donkeys are collected from various points in the Northeast region, often caught and sold by impoverished farmers for rewards ranging from 20 to 50 Brazilian reais (roughly US$4 to $10), while their hides can fetch over US$4,000 in China (BBC-Brasil, 2021). The widespread absence of property rights over these animals, coupled with the potential for quick profits in poor areas through their sale to slaughterhouses, may contribute to rising criminal activities. Our analysis explores whether areas with high donkey populations before market institutionalization experienced a subsequent increase in crime, as disputes over ill-defined property rights often escalate into violence.



Notes: Data on slaughters of donkeys and other equids is from *Serviço de Inspeção Federal* (SIF), associated with the Brazilian Ministry of Agriculture, Livestock and Food Supply. The grey dashed lines represent the months of the market's regulation and continuation, while the red dashed line indicates the timing of the prohibition.

Figure 2.3: Slaughters of donkeys in Brazil

Importantly, following repeated incidents of animal mistreatment reported by the press and denounced by animal protection NGOs and local communities, temporary bans were imposed on donkey slaughter in Brazil (Gameiro et al., 2021; Skippen et al., 2021). The first major ban was instituted by the end of 2018 but was lifted in the second half of 2019 due to extreme pressure from business groups and local politicians. They argued that the interruption of slaughter had caused serious damage to the public economy as the activity generated jobs and income.[15] However, despite the ban, evidence suggests the practice persisted. Notably, following the ban on donkey slaughter, the number of "other equids" – horses and mules – slaughtered in the Northeast increased significantly, deviating from historical trends, as shown in Figure 2.3. This pattern supports the

[15]For instance, the then-mayor of Amargosa, home to the region's largest slaughterhouse, claimed that the facility was the city's second-largest private employer (BBC-Brasil, 2021).

abundant anecdotal evidence that illegal slaughters have surfaced in the Northeast region after the ban imposed by the federal government (BBC-Brasil, 2021; G1 - Globo, 2017a; Gameiro et al., 2021; UOL, 2023). Interestingly, even after the ban was lifted, the slaughter of "other equids" continued alongside donkeys in the Northeast, reaching levels near the peak before the prohibition. This suggests a persistent pattern of semi-licit activity, likely sustained by fixed costs already incurred and the risk of renewed prohibitions driven by public pressure.

The potential for quick profits in impoverished areas, coupled with the persistence of illegal operations and the absence of clear property rights, presents significant policy challenges. Our analysis underscores the need for a thorough evaluation of both the benefits and external costs of this activity in Brazil. These findings also offer valuable insights for developing nations grappling with the surge in demand for donkey hides, as they confront the exploitative nature of the market and the difficulties of enforcing unilateral bans on the trade.

2.3 Data Description and Empirical Strategy

2.3.1 Data

Donkeys' presence

We rely on data regarding the presence of donkeys per municipality in Brazil obtained from the Census of Agriculture, conducted every ten years by the Brazilian National Statistical Institute (*Instituto Brasileiro de Geografia e Estatística* - IBGE). These agricultural census records are compiled through direct interviews with agricultural establishment managers and are accessible online through IBGE, aggregated at the municipality level. Our primary focus is on the 2017 wave, collected in the months before the donkey market's regulation, with the 1995 wave used for robustness checks. Combining this information with population data from IBGE allows us to construct our primary measure of continuous treatment in our empirical approach: the number of donkeys per 1,000 inhabitants. Figure 2.1 illustrates the spatial distribution of donkeys across Brazil and the donkey population density in the Northeast region. We also gather information on other types of livestock from the Census of Agriculture to conduct placebo exercises.

Crime data

Our outcome variable, serving as an indicator of the incidence of crime, is the homicide rate per 100,000 inhabitants. We use mortality records from DATASUS, an administrative dataset from the Ministry of Health, available at the municipality level. We use quarterly data from 2013 to 2021. Population data from IBGE are then employed to compute mortality rates per 100,000 inhabitants. Throughout our empirical investigation, we concentrate on homicide rates per place of occurrence. However, results are consistent when using rates by place of residence.[16]

It is important to note that homicide rates typically exhibit higher reporting accuracy compared to other types of crime and violence in developing countries (Soares, 2004), and the unified system of public health maintained by the Brazilian government ensures a certain level of consistency in definitions across regions. This approach is commonly employed in the literature examining the relationship between economic shocks and crime in Brazil (Chimeli & Soares, 2017; Dix-Carneiro et al., 2018; Soares & Souza, 2023). Data on other forms of crime in Brazil are typically collected by state-level police forces. Unfortunately, states in the Northeast region of the country lack systematic data collection at the municipality level for these variables over time, precluding their use in our analysis.

Other variables

Our analysis also incorporates additional municipality-level data from the population census (IBGE) and mortality records (DATASUS). Specifically, we use control variables such as GDP per capita and the share of agriculture in GDP, both sourced from IBGE. Furthermore, we include mortality rates from suicide and transit accidents per 100,000 inhabitants, obtained from DATASUS. These variables help account for differences in urbanization and development across municipalities, allowing us to control for other potential influences on local crime rates.[17]

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[16]Results using the homicide rate per place of residence as the dependent variable in our main specifications are provided in the Appendix B.

[17]Table B.1 describes the summary statistics for the relevant variables used in our analysis.

2.3.2 Empirical Strategy

We employ a difference-in-difference (DiD) research design, capitalizing on the natural variability in donkey density per capita and the timing of policy shifts. Our identification strategy hinges on the unique timing of the intervention, enabling us to discern divergent responses across regions to the regulation of the activity. Building upon the institutional context described in Section 2.2, we use the regulation changes outlined in March 2017 as the treatment time. Notably, all treated units experience policy changes simultaneously, mitigating potential drawbacks associated with conventional DiD estimators, as elucidated in recent literature (De Chaisemartin & d'Haultfoeuille, 2020; Goodman-Bacon, 2021; Roth et al., 2023).

We begin estimating the following difference-in-difference model:

$$\text{Homicides}_{i,t} = \alpha + \beta\left(\text{Slaughter}_t \times \text{Donkeys}_i\right) + \theta_i + \mu_{s,t} + \gamma_t X_i + \varepsilon_{i,t} \tag{2.1}$$

where $\text{Homicides}_{i,t}$ indicates the homicide rate per 100,000 inhabitants for municipality i in quarter t, Slaughter_t is a dummy variable equal to 1 for the quarters after March 2017 – when the slaughtering activity was made legal and started –, Donkeys_i is our exposure variable indicating the relevance of donkeys in terms of number of animals per 1,000 inhabitants in municipality i. Additionally, θ_i is a municipality fixed effect and $\mu_{s,t}$ represents state-quarter specific dummies. We include the state-quarter fixed effects on the estimation to deal with simultaneous changes beyond the regulation of the donkey market that could confound our estimates. Notably, the market's regulation may induce economic repercussions that indirectly influence violence incidences, such as enhanced income opportunities or altered agricultural patterns. By including these terms, we control for systematic variations across states arising from policy or socioeconomic shifts. In some specifications we also control for time-varying covariates (X_i), interacting their values in the baseline with quarter dummies.[18] Finally, $\varepsilon_{i,t}$ is the idiosyncratic error term.

[18] We use GDP per capita, the share of agriculture in GDP, traffic accidents and suicides per 100,000 inhabitants to control for different trends in municipalities with different levels of urbanization and development.

Moreover, we adopt the equivalent event study specification to evaluate whether treatment and control municipalities demonstrated similar trends in homicides before the legislative change. The event study specification is described by Equation (2.2):

$$\text{Homicides}_{i,t} = \sum_{t \neq 2017Q1} \beta_t \left(\text{Slaughter}_t \times \text{Donkeys}_i\right) + \theta_i + \mu_{s,t} + \gamma_t X_i + \varepsilon_{i,t} \quad (2.2)$$

This equation presents a more adaptable version of the baseline model, facilitating an empirical assessment of the parallel trends assumption. The crucial distinction lies in the incorporation of a series of parameters, β_t, each corresponding to a specific quarter in our panel – except from the baseline, for which we use the first quarter of 2017. Assuming parallel trends, we expect the coefficients β_t for all periods before our treatment to demonstrate no significant deviation from zero.

Finally, all regressions are weighted by population. This ensures that municipalities with larger populations contribute proportionally to the estimation process, preventing biases and heteroskedasticity stemming from population size differentials in the crime rates (Solon et al., 2015). Also, to mitigate the potential underestimation of standard errors due to serial and spatial correlation in the residuals, we adopt a clustering approach for standard errors at the micro-region level – economically integrated groups of municipalities with similar geographic and productive characteristics, as defined by IBGE and widely used in the literature.[19] By doing so, we accommodate for arbitrary correlation structures over time and allow for unobserved shocks to be correlated in space.[20]

[19]Donkeys, particularly wild or feral ones, have relatively short home ranges, averaging about 32 square kilometers (Klingel, 1998; Rudman, 1998). This range indicates that clustering at the micro-region level is well-suited to capture potential spatial spillovers.

[20]We assess the statistical robustness of our results by using alternative spatial clustering methods in the Appendix B. The findings remain consistent with the main results.

2.4 Effects of Donkey Market Regulation

2.4.1 Main results

Our main findings, shown in Table 2.1, focus on municipalities in the Northeast region. These municipalities were selected for several reasons. First, they hold 90% of Brazil's donkey population.[21] Second, this region is the primary hub for donkey slaughter and export to China, matching the locations of newly established slaughterhouses. Third, anecdotal evidence highlights a well-developed ecosystem around the donkey hide market in the region, from capturing free-roaming animals to their sale to slaughterhouses for export – even in the illegal period. Each column in Table 2.1 presents estimated coefficients based on Equation (2.1), using quarterly data described in Section 2.3.[22]

Column (1) uses a standard two-way fixed effects model, controlling for both quarter and municipality fixed effects. Column (2), our primary specification, includes municipality fixed effects alongside state-quarter fixed effects, accounting for systematic changes in crime activity driven by state-level policies or socioeconomic shifts. Column (3) extends the model by incorporating baseline municipal characteristics related to income, economic development, and urbanization, such as suicide and transit accident mortality rates per 100,000 inhabitants (measured in 2012), the share of agriculture in municipal GDP, and the natural logarithm of per capita GDP (both measured in 2010). These variables, interacted with quarter-fixed effects, capture municipality-specific socioeconomic patterns, and their inclusion yields negligible changes to our estimates.

Overall, the results in Table 2.1 demonstrate a statistically significant and positive causal impact of the institutionalization of the donkey market on local crime rates, as measured by homicide rates, in municipalities with higher donkey population densities. To quantify the economic significance, column (2) of Table 2.1, our preferred specification, shows that a one-standard-deviation increase in donkeys per thousand inhabitants leads to an increase of approximately 1.31 homicides

[21]This concentration is visually evident in Figure 2.1 and further detailed in Figure B.3, which shows the donkey population share across Brazil's regions.

[22]We also provide robustness checks using yearly data in the Appendix B, where the results remain quantitatively and qualitatively consistent.

Table 2.1: Market regulation and crime rate

Dep. var: Homicides$_{i,t}$	OLS/DiD (1)	OLS/DiD (2)	OLS/DiD (3)
Slaughter × Donkeys	0.0729***	0.0743***	0.0277**
	(0.0208)	(0.0205)	(0.0124)
Observations	63,604	63,604	63,604
Adj. R^2	0.44852	0.48511	0.4743
Quarter fixed effects	✓		
Municipality fixed effects	✓	✓	✓
State-Quarter fixed effects		✓	✓
Baseline charac. × Quarter fixed effects			✓

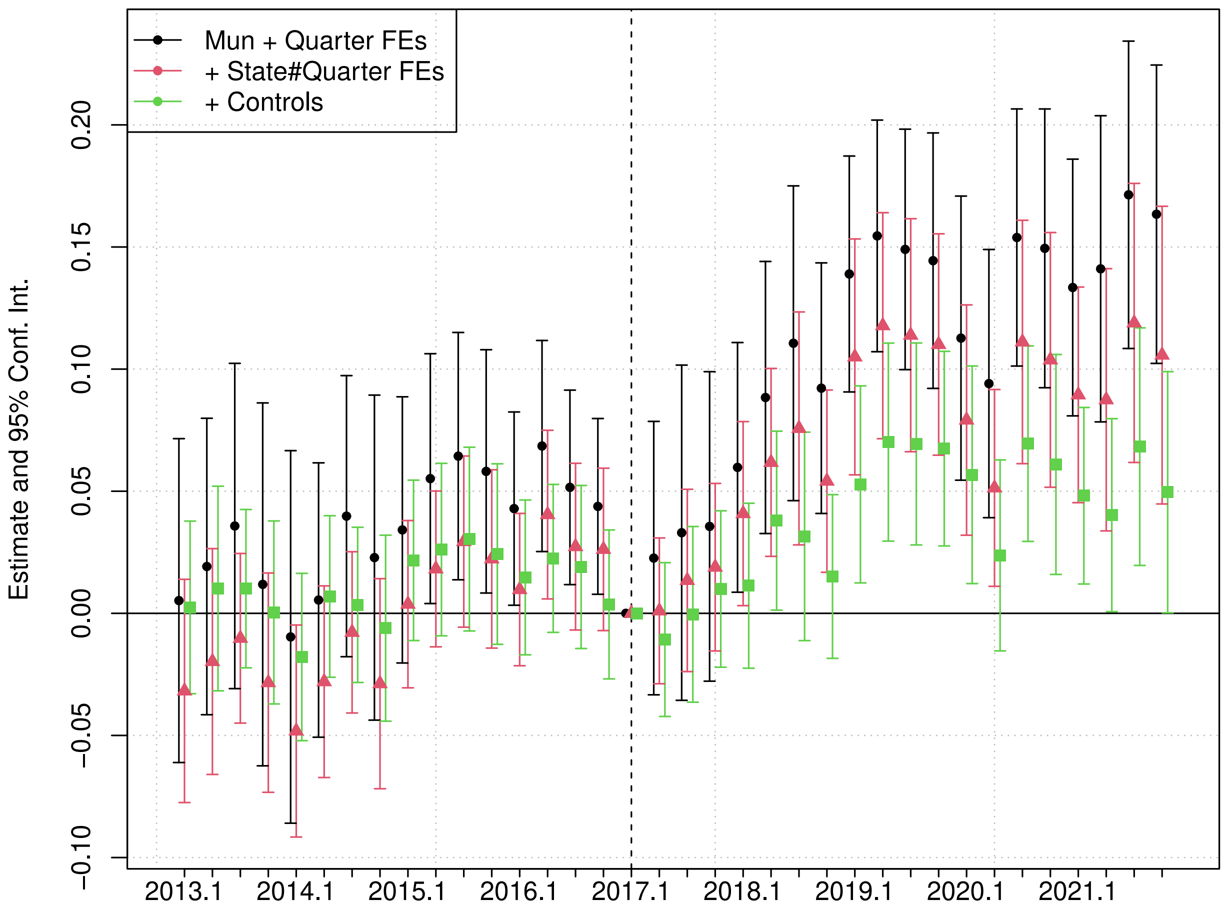
Notes: Standard errors (in parentheses) are clustered for the micro-region level. Unit of analysis is a municipality. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

per 100,000 inhabitants $[= (17.691) \times (0.0743)]$ in Northeast municipalities. Given the region's average quarterly homicide rate of about 9.95 per 100,000 inhabitants, this increase represents a rise of roughly 13% above the average rate. Additionally, moving a municipality from the 25th to the 75th percentile of donkey density corresponds to an increase of about 1.04 homicides per 100,000 inhabitants $[= (16 - 2) \times (0.0743)]$.

To establish the causal effects of the donkey market on crime rates, it is fundamental to test the validity of the parallel trends assumption. This assumption requires that municipalities with different donkey population density exhibit similar trends in homicide rates prior to the policy change. To examine this, we employ the event-study specification outlined in Equation (2.2), focusing on the interaction of the treatment indicator with the respective quarter dummies.[23] Figure 2.4 illustrates the outcomes, along with 95% confidence intervals. Notably, the pre-treatment coefficients are shown not to be significant, both individually and collectively, at conventional significance levels. These results support the validity of the parallel trends assumption and strength the robustness of our research design. Moreover, note that the addition of state-specific time trends are sufficient

[23]We conduct a similar analysis using yearly data in the Appendix B, yielding consistent results.

to establish conditional parallel trends prior the policy intervention, which justifies the choice of our preferred specification. Even so, we report the results with the inclusion of baseline controls, which does not change our main findings.



Notes: Each point reflects an individual regression coefficient, at quarter frequency, $\hat{\beta}_t$ following Equation (2.2). Standard errors are adjusted to micro-region clusters. Solid lines show 95 percent confidence intervals.

Figure 2.4: Dynamic impacts: market regulation and crime rates

This flexible specification allows us to track the evolution of treatment effects over time. As shown in Figure 2.4, post-treatment coefficients become statistically significant about four quarters after the policy change. Notably, the impact on homicide rates intensifies following the temporary ban on donkey-related activities at the end of 2018. It is important to note that the dip in point estimates during the first two quarters of 2020 likely reflects reduced activity due to COVID-19 lockdowns.

Juxtaposing these findings with Figures 2.2 and 2.3, we observe a substantial surge in homicide rates during the prohibition period in 2019. We interpret this phenomenon in light of a concurrent increase in the illegal slaughter of other equids in Bahia and the rise in exports (both value and weight) of donkey-hide-related products during this period, suggesting that the market persisted through illicit channels.[24] This persistence likely exacerbated adverse effects on crime dynamics, particularly in areas with a higher pre-treatment prevalence of donkeys. This logic suggests that if illegal market operations drive up average prices per donkey – due to a reduced supply, given the increased risk of capturing these animals and transporting them to illicit slaughterhouses, combined with relatively inelastic demand – the activity becomes more financially attractive to factions competing for control over these animals. This heightened incentive may have outweighed the expected costs of getting caught, likely increasing participation in this market – characterized by poorly defined property rights – and further intensifying criminal activities and associated negative externalities. We explore this channel in detail in Section 2.7.

2.4.2 Robustness

We conduct several exercises to asses the robustness of our main findings. First, we evaluate the robustness of our results across different samples of municipalities and regional units of analysis. While our primary analysis focuses on municipalities in the Northeast region, it is important to verify whether the results hold across alternative samples. Columns (1) to (3) of Table 2.2 present these outcomes. In column (1), we analyze municipalities within the Northeast region that reported a positive number of donkeys in 2017. Column (2) narrows the focus to municipalities in Bahia state, where donkey slaughterhouses are concentrated. In column (3), we aggregate municipalities

[24]This interpretation is supported by anecdotal evidence of the activity's continuation, now conducted illegally, in the Northeast region following the federal government's ban. See, for example, journalistic reports by BBC-Brasil (2021), G1 - Globo (2022), and UOL (2023).

into micro-regions and use this aggregation as the unit of analysis in the regression.[25] Across all subsamples, the results consistently demonstrate that regulatory interventions in the donkey market have statistically significant and positive effects on homicide rates in localities with a higher density of donkeys per inhabitant.

Next, we explore whether the base year of our donkey population density measure, a key component of the treatment variable, influences our results. One concern is that municipalities with high donkey counts in the agricultural census might not coincide with those having large numbers of free-roaming donkeys. For instance, municipalities that underwent mechanization in the early 2000s could report fewer donkeys in 2017 despite a significant free-roaming population. To account for this, we construct the donkey density measure using an earlier wave of census data. Additionally, we may be capturing different dynamics in poor rural areas not controlled by the initial characteristics included in Table 2.1. To address this, we conduct a placebo test using the number of caprine stock (goats) per 1,000 inhabitants. We use goat population density as a placebo because, like donkeys, they are historically significant livestock in the region and a valuable asset in climate-vulnerable areas of Brazil (Da Mata & Resende, 2020), but their market was not affected by mechanization or exports to China.

Columns (4) and (5) of Table 2.2 summarize the results from estimating Equation (2.1) using an alternative year for constructing the donkey population density variable and conducting the placebo test with the goat population density as the treatment variable, respectively. Specifically, we use the 1995 Agricultural Census for donkeys and the 2017 census for goats.[26] Our findings show that using alternative treatment measures does not alter the main results. Furthermore, the placebo test yields a small, statistically insignificant coefficient, reinforcing our claim that changes in crime dynamics are due to the donkey market.[27]

[25] We assess the robustness of these micro-region-level results using alternative spatial clusters for the standard errors in the Appendix B.

[26] Since the goat market was unaffected by mechanization or exports to China, and goats are valuable livestock with no free-roaming populations, endogeneity concerns are minimal when using 2017 data.

[27] We also conduct several sensitivity analyses in the Appendix B, demonstrating the robustness of our main findings to the exclusion of observations from the sample.

Table 2.2: Robustness to alternative samples and treatments

Dep. var: Homicides$_{i,t}$	Alternative sample			Alternative treatment	
	Some donkey (1)	BA only (2)	Micro-region (3)	Donkey rate 1995 (4)	Placebo (5)
Slaughter × Donkeys	0.0474**	0.0711*	0.1346***		
	(0.0190)	(0.0414)	(0.0383)		
Slaughter × Donkeys in 1995				0.0203***	
				(0.0063)	
Slaughter × Goats					0.0004
					(0.0003)
Observations	58,889	14,974	6,752	63,604	63,604
Adj. R^2	0.4572	0.4839	0.7667	0.4846	0.4828
Municipality fixed effects	✓	✓		✓	✓
Micro-region fixed effects			✓		
State-Quarter fixed effects	✓	✓	✓	✓	✓

Notes: Standard errors (in parentheses) are clustered for the micro-region level. Unit of analysis is a municipality in columns (1),(2),(3), and (5), and a micro-region in column (4). All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Taken together, the consistency of results across different samples and the insignificance of the placebo test bolster our conclusions about the causal impact of the donkey market on crime rates.

2.5 Falsification

In the previous section, we presented empirical evidence that the legalization of the donkey market in Brazil—marked by large free-roaming populations and ill-defined property rights—led to a significant increase in crime rates in areas with greater exposure to donkeys before the legislation. The underlying rationale is that the nature of the activity fosters conflict resolution through violence.

To further assess the validity of our findings, beyond the robustness and placebo tests, we conduct a falsification test. The logic is straightforward: if an omitted variable is driving the observed increase in homicide rates, we would expect the market legalization to also impact other causes of mortality, such as those related to local economic conditions or cycles. For this falsification test, we examine the effects of market legalization on mortality rates from undetermined causes, from causes typically associated with income shocks (e.g., cardiovascular and liver diseases, as well as neoplasms), and on child mortality (under 14), which should remain largely unaffected by crime-related factors, given their lack of participation in the market.

The results are presented in Table 2.3. Notably, none of the coefficients associated with market legalization are statistically significant at conventional levels. Besides, even in the case of mortality rates for children from 0-6 years old where the coefficient is significant at the 10% significance level, the magnitude of the effect is *de minimis*. This indicates that the effect of market legalization on other causes of mortality is negligible, providing no evidence of a broader impact across unrelated mortality rates.[28]

Table 2.3: Falsification: effects of market regulation on specific mortality causes

Dep. var.: Mortality causes	Undetermined (1)	Cardiovascular (2)	Neoplasm (3)	Liver (4)	0-14 years (5)	0-6 years (6)	7-14 years (7)
Slaughter × Donkeys	-0.0014	-0.0021	-0.0001	-0.0014	0.0099*	0.0093*	0.0006
	(0.0097)	(0.0124)	(0.0030)	(0.0097)	(0.0058)	(0.0050)	(0.0014)
Observations	63,604	63,604	63,604	63,604	63,604	63,604	63,604
Adj. R^2	0.4367	0.8695	0.2888	0.4367	0.8326	0.8456	0.1643
Municipality fixed effects	✓	✓	✓	✓	✓	✓	✓
State-Quarter fixed effects	✓	✓	✓	✓	✓	✓	✓

Notes: Standard errors (in parentheses) are clustered at the micro-region level. Unit of analysis is a municipality. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

These findings suggest that there are no relevant omitted factors in our previous estimates that broadly impact mortality rates unrelated to criminal activity during the period. The specific mortality rates, theoretically independent of criminality or primarily driven by income variations, remain unaffected by the donkey market legalization. Alongside the robustness and placebo tests, this falsification test bolsters the robustness of our empirical approach and further validates the credibility of our identification strategy.

[28] In the Appendix B, we present additional evidence for this falsification exercise from event study specifications demonstrating the negligible effects of donkey market regulation on these specific causes of mortality.

2.6 Mechanism Validation

Next, we evaluate the validity of our proposed mechanism underlying the effect of donkey market regularization on crime rates – that the characteristics of the market lead to increased violence as disputes over a now valuable resource with ill-defined property rights are resolved through violent means.

2.6.1 Distance to slaughterhouses

First, if our hypothesis holds we expect the crime effects following the legislative change to be more pronounced in regions closer to slaughterhouses. The reasoning is straightforward: given the presence of free-roaming donkeys, the cost of capturing and transporting the animals is lower near slaughterhouses. Furthermore, if information networks play a role in such markets, players closer to the slaughterhouses are better informed on prices and the donkey demand. Consequently, the economic incentive to participate in this activity should be greater in these areas compared to more distant municipalities, where transportation costs are higher and there is less information.

We assess this mechanism in Table 2.4. Column (1) shows the baseline result from our preferred specification, while columns (2) to (4) focus on municipalities within 250 kilometers, 500 kilometers, and 1000 kilometers of the nearest slaughterhouse—either Amargosa or Itapetinga, both in Bahia, where donkey slaughter and export activities are concentrated. The results in Table 2.4 corroborate our hypothesis, indicating that proximity to slaughterhouses significantly amplifies the impact of donkey market regulation on homicide rates. Specifically, municipalities within 250 km of a slaughterhouse experience a nearly double increase in homicide rates compared to the baseline. Importantly, the effect diminishes steadily and monotonically as the distance from the slaughterhouse increases.

2.6.2 Prohibition and enforcement

Second, we examine the temporal overlap between the 2019 ban on the donkey market, the evidence of continued illegal activity, and the observed increase in homicide rates in the most

Table 2.4: Effects of market regulation on crime rate per distance to slaughterhouse

	Distance to closest slaughterhouse			
Dep. var: Homicides$_{i,t}$	Baseline (1)	250km (2)	500km (3)	1000km (4)
Slaughter × Donkeys	0.0743*** (0.0205)	0.1450*** (0.0421)	0.0870*** (0.0326)	0.0553*** (0.0166)
Observations	63,604	9,889	18,824	53,924
Adj. R^2	0.4674	0.4210	0.4491	0.4268
Municipality fixed effects	✓	✓	✓	✓
State-Quarter fixed effects	✓	✓	✓	✓

Notes: Standard errors (in parentheses) are clustered for the micro-region level. Unit of analysis is a municipality. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

affected municipalities. If our suggested mechanism is correct, we expect the policy change to have intensified crime rates during the ban, likely due to rising donkey prices and the increased economic benefits of participating in the exploitation of resources with ill-defined property rights. However, the illegal nature of the activity may have also raised the cost of participation, especially with efficient enforcement.

To explore this, we estimate a modified version of Equation (2.1), which includes an additional treatment variable indicating the quarters when the activity was illegal. Specifically, we estimate the following triple-differences model:

$$\begin{aligned}\text{Homicides}_{i,t} = \alpha &+ \beta\left(\text{Slaughter}_t \times \text{Donkeys}_i\right) \\ &+ \delta\left(\text{Illegal}_t \times \text{Slaughter}_t \times \text{Donkeys}_i\right) \\ &+ \theta_i + \mu_{s,t} + \gamma_t X_i + \varepsilon_{i,t}\end{aligned} \tag{2.3}$$

where Illegal_t equals 1 between the fourth quarter of 2018 and the third quarter of 2019.

The results are summarized in Table 2.5. Across all specifications, from the canonical two-way fixed effects model in column (1) to the inclusion of state-quarter fixed effects and baseline characteristics in column (3), the impact of the legislative change on crime rates in more exposed municipalities intensifies during the ban. The larger coefficients during this period supports our hypothesis that the rise in crime is driven by increased economic benefits from capturing and selling donkeys to slaughterhouses, compounded by limited enforcement capabilities from local governments.

Table 2.5: Illegality and homicide rate

Dep. var: Homicides$_{i,t}$	OLS/DiD (1)	OLS/DiD (2)	OLS/DiD (3)
Slaughter × Donkeys	0.0664***	0.0681***	0.0243**
	(0.0201)	(0.0197)	(0.0121)
Illegal × Slaughter × Donkeys	0.0328***	0.0309***	0.0166**
	(0.0120)	(0.0092)	(0.0075)
Observations	63,604	63,604	63,604
Adj. R^2	0.4324	0.4675	0.4746
Municipality fixed effects	✓	✓	✓
State-Quarter fixed effects		✓	✓
Baseline charac. × Quarter fixed effects			✓

Notes: Standard errors (in parentheses) are clustered at the micro-region level. Unit of analysis is a municipality. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

These findings have significant implications for policymakers regulating markets tied to natural resource exploitation. Like Chimeli and Soares (2017) and Araujo et al. (2024), we find that prohibiting an exploratory activity, without sufficient monitoring or effective enforcement, leads to unintended consequences. These include the persistence of the activity even during its illegality, amplifying negative externalities, especially increased criminality and violence in the most exposed municipalities of the Northeast region.

2.6.3 Characterizing the victims

Lastly, if our story is correct, we expect that individuals more likely to participate in the market are disproportionately impacted by the violence associated with it. Given that we are examining a resource-based market with violent resolution of ill-defined property rights, young males – who are more prone to engage in risky economic activities like capturing and transporting donkeys to slaughterhouses – are particularly vulnerable. This effect should manifest as a relative increase in homicide rates among specific groups. Table 2.6 provides further evidence supporting this narrative. To better understand the rise in violence, we break down the dependent variable by characteristics typically associated with criminal violence, rather than domestic or interpersonal conflicts.

First, we restrict the analysis to homicides involving firearms versus non-firearm-related causes, as firearm violence is more indicative of criminal activity. Next, we examine homicides by gender, separating male and female victims. Finally, we focus on male homicides, specifically comparing young males (ages 20 to 40) with other age groups.

Table 2.6: Effects of market regulation on specific homicide rates

Dep. var: Specific Homicides$_{i,t}$	Baseline (1)	Gun-related (2)	Not Gun-related (3)	Male (4)	Female (5)	Young Male (6)	Not Young Male (7)
Slaughter × Donkeys	0.0743***	0.0717***	0.0025*	0.0713***	0.0030**	0.0416***	0.0326***
	(0.0205)	(0.0201)	(0.0014)	(0.0196)	(0.0014)	(0.0125)	(0.0083)
Observations	63,604	63,604	63,604	63,604	63,604	63,604	63,604
Adj. R^2	0.4674	0.4564	0.1298	0.4626	0.0664	0.3939	0.2968
Municipality fixed effects	✓	✓	✓	✓	✓	✓	✓
State-Quarter fixed effects	✓	✓	✓	✓	✓	✓	✓

Notes: Standard errors (in parentheses) are clustered at the micro-region level. Unit of analysis is a municipality. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.6 indicates that the increase in violence in donkey-dense areas after the market institutionalization is predominantly driven by firearm violence against prime-aged men. The point estimates in columns (2) and (3) show that firearm violence, compared to non-firearm causes, is the main driver of the homicide increase, underscoring the criminal nature of the violence. Comparing columns (4) and (5), we observe that this rise in homicide rates is almost entirely driven by male victims. When breaking down male homicides by age, columns (6) and (7) reveal that the increase is concentrated among men aged 20 to 40, though there is also a significant rise in homicide rates for men above this age threshold. These results suggest that the observed increase in violence in municipalities more exposed to the donkey market is more likely driven by criminal violence rather than domestic or family-related disputes.

In summary, the evidence across these channels – proximity to slaughterhouses, the effects of prohibition and enforcement, and the characterization of victims – collectively supports the validity of our proposed mechanism. Taken together, these findings reinforce the conclusion that the market's characteristics are directly linked to the rise in criminal violence, measured by homicides, in the more affected municipalities.

2.7 Concluding Remarks

The exploitation of resources through international trade is a long-standing issue in development economics. This paper investigates how the regulation of a new renewable-resource market – characterized by open access and poorly defined property rights – affects local violence. Focusing on the regulatory changes that established Brazil's donkey hide trade, we find that the market led to increased crime and violence in resource-rich regions. Importantly, our results show that both the foreign demand boom itself and the illegal activity in this market play important roles in explaining these results.

Our findings have important policy implications. First, they highlight the relationship between violence and property rights within a market. When property rights are poorly defined – such as with free-roaming donkeys in Brazil – the institutionalization of such markets can lead to unintended social consequences and welfare losses. Second, prohibition policies without proper enforcement capabilities may amplify such negative effects. We document that making the donkey market illegal doubles its effect on local crime rates and violence. Finally, our results point out to the global repercussions of resource exploitation. The widespread negative effects of the donkey hide trade across the world underscore the need of global regulation in markets where there is predatory exploitation of natural resources. In the case of the donkey hide market, the inefficiency of national-level bans is evident, and broader efforts, such as the recent African Union's continent-wide ban on donkey skin trade, represent a more coordinated and potentially positive approach. Nonetheless, our results suggest that global-level coordination may still be necessary in such cases, particularly given the extent of the social costs involved.

By way of conclusion, while this study offers valuable insights into the causal links between the donkey hide market and crime in Brazil, much remains to be explored regarding the broader implications of this trade in other developing regions. These findings resonate with a long-standing literature on trade and resource depletion, highlighting that issues from past centuries – such as the exploitation and eventual depletion of resources like the North American buffalo (Feir et al., 2024; Taylor, 2011) – persist today and continue to emerge across the developing world, with significant welfare implications.

Chapter 3

Exogenous economic shocks and incumbents' electoral fortunes: evidence from Brazilian mayoral elections

[29]

3.1 Introduction and Motivation

In the realm of political science, research dating back to at least Kramer (1971) has consistently established a positive correlation between economic outcomes and support for incumbent candidates, particularly in the context of general elections (Lewis-Beck & Stegmaier, 2000; Lewis-Beck et al., 2008; Nadeau et al., 2013). However, the straightforward nature of this economic voting hypothesis becomes more complex when examining smaller political units, such as states, counties, and municipalities (Lewis-Beck & Stegmaier, 2007, 2019).

This correlation between economic outcomes and political support can be attributed to two key aspects of voter behavior. Firstly, there is substantial international evidence supporting retrospective voting, indicating that citizens tend to weigh the performance of incumbents heavily when making their electoral decisions (Carlin et al., 2015; Duch & Stevenson, 2008). Secondly, various studies suggest that voters also consider external factors beyond the control of incumbents when evaluating their performance (Achen & Bartels, 2017; Healy et al., 2010). This second aspect holds particular importance for developing countries, which historically face a higher susceptibil-

[29]This chapter is a preliminary version of a co-authored paper with Iasmin Goes. I benefited from fruitful discussion with and useful comments and impressions from Elissa Braunstein, Daniele Tavani, Ramaa Vasudevan and Ray Miller. I am also thankful for the participants of graduate seminars at the Western States Graduate Workshop. Any remaining errors are my own.

ity to exogenous shocks. Despite the wealth of research in this area, the majority of evidence is concentrated within advanced economies, creating a notable gap in the literature.

This paper seeks to address this gap by investigating the ramifications of trade-related economic shocks on local-level voting behavior within the framework of a developing country. Specifically, our focus centers on understanding how exogenous shocks linked to commodity production can influence the voting choices of individuals residing in municipalities most affected by these shocks. We aim to discern whether these regional economic shocks have a tangible impact on the electoral fortunes of local incumbents. Importantly, we exclusively focus, in both theoretical-formal and empirical-econometric fronts of this paper, on executive elections rather than legislative ones. This emphasis is informed by a robust body of literature on the "clarity of government responsibility", which underscores the executive's pivotal role as a focal point for voters, particularly in countries characterized by presidential systems. Consequently, voters often disproportionately reward or penalize executive incumbents for factors beyond their direct control (Ang et al., 2022; Hobolt et al., 2013; Powell & Whitten, 1993; Samuels, 2004).

To gain additional insight into these dynamics, we initially develop a theoretical framework that models the impact of exogenous regional economic shocks on the reelection prospects of incumbents in local elections. Our framework posits that the interplay between the intensity of these shocks and the competence of politicians shapes electoral outcomes. Drawing upon existing literature (Ashworth et al., 2018; Wolfers, 2007), we scrutinize how varying degrees of exposure to these exogenous economic shocks can either magnify or attenuate the influence of incumbent competence on electoral fortunes, with a particular emphasis on the pivotal role of voter learning in assimilating additional information.

Importantly, the existing literature related to our research question is growing but still limited. In a recent effort, Campello and Zucco (2020) examined the impact of aggregate economic performance on support for chief executives in Latin American economies characterized by exposure to exogenously induced economic volatility. They argue that such volatility may restrict the ability of voters to hold leaders accountable through elections. However, their findings indicate that presidential elections in these countries do respond to international commodity prices and interest rates.

Focusing on the Brazilian case, Novaes and Schiumerini (2022) demonstrate that increases in the price of agricultural commodities significantly bolster the prospects of incumbents in local elections, while negative shocks exacerbate their incumbency disadvantage, particularly in rural municipalities.

Yet, irrespective of the unit of analysis, this literature faces a notable limitation — namely, the challenge of accounting for various factors that can introduce bias into the analysis, including omitted variables and reverse causality.

We overcome this limitation by leveraging two "quasi-natural" experiments that generated heterogeneous effects on agriculture productivity and export dynamics across municipalities in Brazil. This paper endeavors to measure the distinct impacts of emerging agricultural technology and heightened trade demand, both of which generated heterogeneous effects on regional economic conditions, on voting behavior at the local level in Brazil. Specifically, we dissect the effects of the introduction of genetically engineered soy seeds, as explored by Bustos et al. (2016) and Bustos et al. (2020), and the surge in demand from China, as examined by Costa et al. (2016) and Carreira et al. (2024), on the electoral performance of incumbents in municipal elections spanning from 2000 to 2020.

Bustos et al. (2016) exploit the introduction of genetically engineered soybean seeds in Brazil as a "quasi-natural" experiment that generated heterogeneous effects on agricultural productivity across areas with different soil and weather characteristics. Their findings suggest that this technical change in soy production led to increased output per worker, a reduction in agricultural labor intensity, and industrial growth. Using a similar empirical strategy, Bustos et al. (2020) show that agricultural productivity growth generated an increase in savings that, through the credit system, had positive spillover effects in terms of capital accumulation in municipalities linked to the soy-producing regions.

Costa et al. (2016) investigate China's emergence in the world market after joining the World Trade Organization in 2001 as a "quasi-natural" experiment that generated export and import demand shocks for Brazilian local labor markets. Their research reveals that local labor markets more affected by Chinese import competition experienced slower growth in manufacturing wages between 2000 and 2010, but they also documented faster wage growth in locations benefiting from rising Chinese commodity demand during the same period. Furthermore, in a recent paper, Carreira et al. (2024) follow a similar empirical strategy to evaluate the impacts of trade on deforestation in Brazil and find no association between exposure to demand from China and deforestation, though they note that trade does induce the conversion of cropland to pastureland.

Our preliminary findings reveal a significant and economically meaningful effect of exogenous regional economic shocks on incumbency effects in municipal elections. Specifically, we observe that neither the shock to agricultural productivity nor the trade shock had any discernible impact on incumbency effects in the 2004 elections, which aligns with expectations given the limited time for these shocks to propagate across regions. However, in the subsequent election, in 2008, we identify an amplification of the incumbency effect in municipalities more exposed to the shock of agricultural technical change compared to less exposed localities. This suggests that incumbents were disproportionately favored in areas where the adoption of genetically engineered soybean seeds led to higher potential yields in production. Additionally, we find a positive impact of the trade shock on incumbency when complete interaction terms are included. Although our evidence is preliminary and limited, it offers promise and indicates a causal relationship between improvements in regional economic conditions and an increased likelihood of an incumbent's victory, validating the hypothesis that incumbents benefit from positive economic volatility even when they are not responsible for or capable of affecting such fluctuations. Another noteworthy finding from our initial exploration is that, while exposure to exogenous regional economic shocks appears to influence the electoral fortunes of incumbents in local elections, this impact is transient. Notably, in subsequent elections (from 2012 to 2020) we do not observe significant impacts of the regional economic shocks on the reelection probability of incumbent mayors.

This paper makes both theoretical and empirical contributions to the study of incumbency effects and democratic accountability. While prior research on incumbency effects has predominantly focused on advantages or disadvantages within a single national context, our work provides a theoretical framework and supporting evidence indicating that incumbency effects can vary at the sub-national level, as illuminated by Novaes and Schiumerini (2022). Furthermore, our study underscores the informational challenges that voters face when making candidate choices, contributing to the ongoing debate concerning voter competence in engaging in electoral accountability. While a substantial body of literature interprets the impact of external shocks on elections as evidence of voters irrationally reacting to irrelevant outcomes (Achen & Bartels, 2017), we argue that voters can gain insights into competence by observing how incumbents have prepared for or managed such shocks (Ashworth et al., 2018).

Finally, this paper also relates to and contributes to the growing literature on the global effects of China's rise as a major player in international trade. This includes research that has examined the impact of Chinese import competition on economic variables such as manufacturing employment (Autor et al., 2013; Pierce & Schott, 2016) and workers' earnings (Autor et al., 2014; Dix-Carneiro et al., 2023). However, there is a smaller body of work that considers the demand-side effects of China's ascent.

The remainder of the paper is structured as follows. Section 3.2 provides background information on the new soy agriculture technology and the local impacts of recent trade shocks in Brazil. In Section 3.3, we develop a simple theoretical-formal framework to examine how exogenous regional economic shocks might impact the electoral fortunes of incumbents in local elections, deriving testable hypotheses from the main predictions. Section 3.4 describes the data used in our empirical exercise and discusses our empirical strategy to examine the main hypotheses derived from the model. Section 3.5 presents the preliminary results of the paper. Lastly, in Section 3.6, we discuss the main findings of the paper and their implications, and we briefly outline the next steps in this research project.

3.2 Technological Change and Commodity Boom: General Context

3.2.1 The introduction of GE soybean seeds and the technological change in production

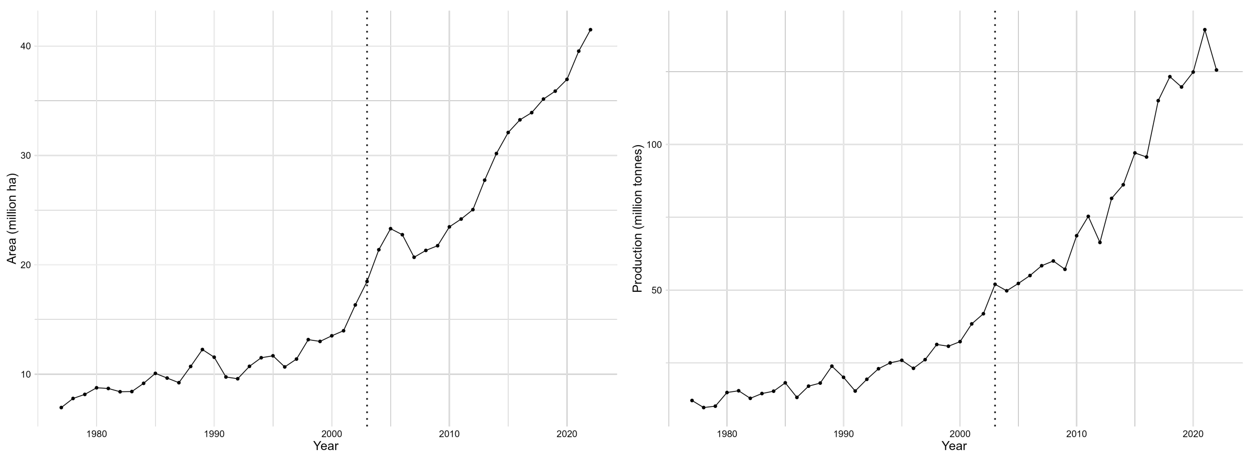
The remarkable surge in crop production within Brazil in recent decades can be chiefly attributed to the introduction of genetically engineered (GE) soybean seeds. These GE soy seeds possess superior herbicide resistance compared to their conventional counterparts. By employing GE seeds, farmers have substantially reduced the need for extensive soil preparation to weed out unwanted plants. Instead, they can directly apply herbicides to the soil, effectively eliminating weeds while safeguarding the soy plants. This innovative approach not only conserves valuable time and labor but also augments potential crop yields.

Crucially, the adoption of this new technology has enabled the expansion of soy production into areas that were previously considered economically nonviable using traditional soy seeds. Our empirical approach to studying the local effects of soy technological change builds upon the work of Bustos et al. (2016) and Bustos et al. (2020). We leverage the legalization of GE soy seeds in Brazil as a source of temporal variation and the differential impact of this technology across regions as a source of cross-sectional variation. This methodology allows us to disentangle the effects of soybean technological change from other confounding factors.

Specifically, the first generation of GE soy seeds was introduced in the United States in 1996, but it was not until 2003 that the Brazilian government legalized their commercialization and planting across the entire country. This temporal gap provides the necessary variation for our analysis, with the year 2003 marking the legalization of GE soy in Brazil. Furthermore, the adoption of GE soy seeds had varying effects on potential yields, depending on local soil and weather conditions, creating the required cross-sectional variation for our study.

Before delving deeper into our empirical strategy and the analysis of voting behavior, it is crucial to examine some key trends in soybean production in Brazil over the past decades. Firstly,

Figure 3.1 illustrates changes in the area planted with soybeans (in millions of hectares) and total production (in millions of tonnes) from the mid-1970s onward across the country. Notably, there is a pronounced increase in the area devoted to soy production, which had a substantial impact on total production, particularly after 2003, coinciding with the adoption of genetically engineered seeds.



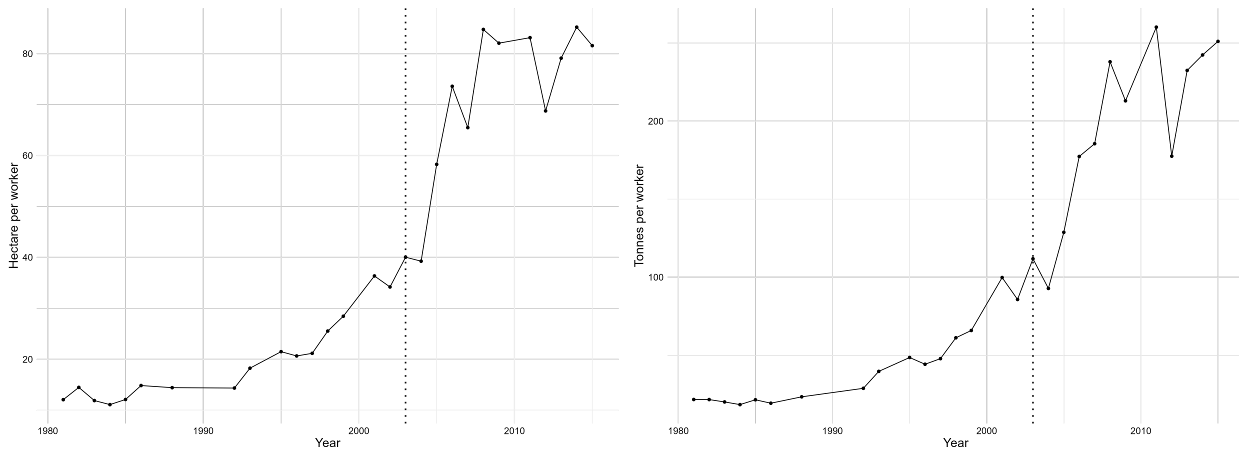
(a) Area planted (b) Total production

Source: Data on soybean planted area and total production is from the Brazilian Ministry of Agriculture, computed by *Companhia Nacional de Abastecimento* (CONAB).

Figure 3.1: Area planted and total production of soybean: 1976-2022

Next, we analyze the behavior of these trends concerning their impact on the number of workers in soybean-related activities. Figure 3.2 illustrates the evolution of area planted per worker (in hectares per worker) and output per worker (or labor productivity, in tonnes per worker) over time. While both variables show consistent increases from the 1990s, there is a noticeable change in slope and level for both graphs after 2003.

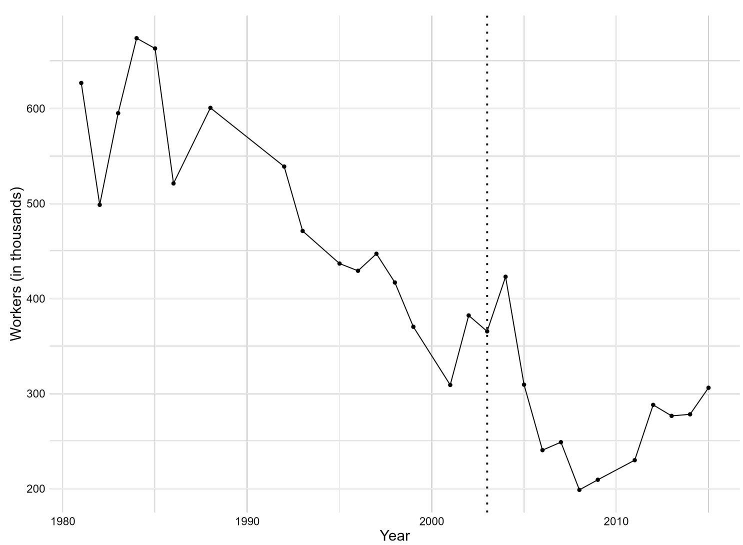
Lastly, we track the evolution of employment in soybean production in terms of thousands of workers from 1981 to 2015. Importantly, these initial trends suggest that the technical change in soy production brought about by the introduction of GE seeds has been labor-saving over the past few decades in Brazil.



(a) Area planted per worker (b) Output per worker

Source: Data on soybean planted area and total production is from the Brazilian Ministry of Agriculture, computed by *Companhia Nacional de Abastecimento* (CONAB). We construct measures of the number of workers in soy production at the state level in Brazil using *Pesquisa Nacional por Amostra de Domicílios* (PNAD), carried out by the Brazilian Institute of Geography and Statistics.

Figure 3.2: Area planted per worker and output per worker in soybean production: 1980-2015



Source: We construct measures of the number of workers in soy production at the state level in Brazil using *Pesquisa Nacional por Amostra de Domicílios* (PNAD), carried out by the Brazilian Institute of Geography and Statistics.

Figure 3.3: Employment in soybean production: 1980-2015

In summary, these visualizations indicate that the adoption of GE soy seeds coincides with an increase in labor productivity and a significant expansion in the area planted with soybeans in

Brazil. To provide a broader perspective, Table 3.1 compares land use and labor intensity in various agricultural activities in Brazil during agricultural census years.

Table 3.1: Land Use and Labor Intensity by Agricultural Activity

	Land use (million ha)						**Labor intensity (worker per 1,000 ha)**		
Principal Activity	1996		2006		2017		1996	2006	2017
	Area	**Share**	**Area**	**Share**	**Area**	**Share**			
Permanent crops	7.5	2.1%	11.7	3.5%	7.8	2.2%	127.2	126.7	139.5
Seasonal crops	34.3	9.7%	44.6	13.4%	55.6	15.8%	110.8	83.7	55.7
Soy	9.2	2.6%	17.9	5.4%	30.7	8.8%	28.6	17.1	13.1
Cattle ranching	177.7	50.3%	160.0	47.9%	159.5	45.4%	25.9	30.6	30.5
Forestry	94.3	26.7%	91.7	27.5%	101.4	28.8%	34.4	42.6	32.0
Not usable and other uses	39.8	11.2%	25.6	7.7%	27.0	7.7%	-	-	-
Total	353.6	100%	333.7	100%	351.3	100%	55.6	49.7	43.0

Notes: The data on land use and labor intensity in agriculture is from the Brazilian Agricultural Censuses of 1996, 2006 and 2017.

Notably, seasonal crops, driven predominantly by the expansion of soybean production, experienced substantial growth during the analyzed period. Alongside this expansion, there was a consistent reduction in labor intensity in these activities compared to others. Overall, the reduction in labor intensity across all agricultural activities in the country in recent decades can be attributed, in part, to the adoption of GE seeds, as evidenced by the significant reduction in labor intensity observed in soybean production between 1996 and 2006. The introduction of GE soybean seeds and the associated technological changes have thus played a pivotal role in reshaping the agricultural landscape and labor dynamics in Brazil.

3.2.2 China's export demand: the resource boom

Another pivotal economic event that exerted a profound influence on the Brazilian agriculture sector in recent decades was the emergence of China on the global stage. After a decade of remarkable domestic growth rates, China's accession to the World Trade Organization (WTO) in 2001 heralded a transformative shift in international trade dynamics (Autor et al., 2013). For developing countries in general, China swiftly transformed into a major exporter of manufactured goods and a significant importer of commodities.

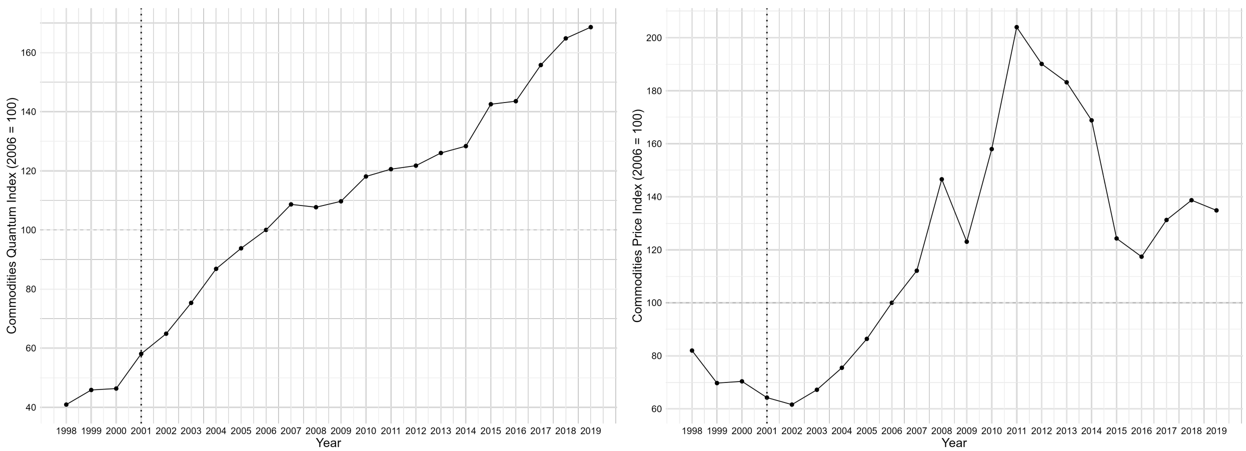
The Brazilian experience closely mirrored this pattern, with China's ascent in international trade creating a resource boom for commodities such as soybeans and iron ore (Carreira et al., 2024). The confluence of the introduction of genetically engineered (GE) soy seeds and the surging demand for soybeans fostered a favorable environment for Brazilian agriculture. Consequently, Brazil claimed the title of the world's leading soybean exporter, with China emerging as the primary destination for Brazilian soy production.

To grasp the magnitude of this expansion in Chinese commodity demand, Costa et al. (2016) point out that the share of agricultural and extractive sectors in the exports of non-high-income countries to China underwent a dramatic transformation. This share rose from less than 20% in 1995 to nearly 70% in 2010. Between 1995 and 2006, Brazilian exports of soybeans to China experienced a substantial increase of $2.8 billion, and this growth accelerated even further, surging by an impressive $17.4 billion between 2006 and 2017 (Carreira et al., 2024). This period also witnessed a simultaneous expansion in Brazilian exports of various other commodities and minerals to China.

Figure 3.4 visually demonstrates the expansion of commodity exports from Brazil from the late 1990s up to 2019 (pre-COVID shock). Notably, there is a pronounced upward trend in the volume of commodity exports starting in the early 2000s. This trend appears to align temporally with an increase in the prices of these commodities in the global market, as indicated in panel (b). Importantly, these increasing trends seem to coincide with China's accession to the WTO.

To explore this possibility further, Figure 3.5 dissects the expansion of Brazilian total exports, dividing them into quantum and prices, directed towards China. As previously suggested, there is a striking similarity between the trends presented in Figure 3.4 for all commodity exports from Brazil and the total exports to China.

Lastly, Figure 3.6 assesses the value of total exports to China and other destinations. Panel (a) underscores the substantial increase in the value of exports from Brazil to China, largely propelled



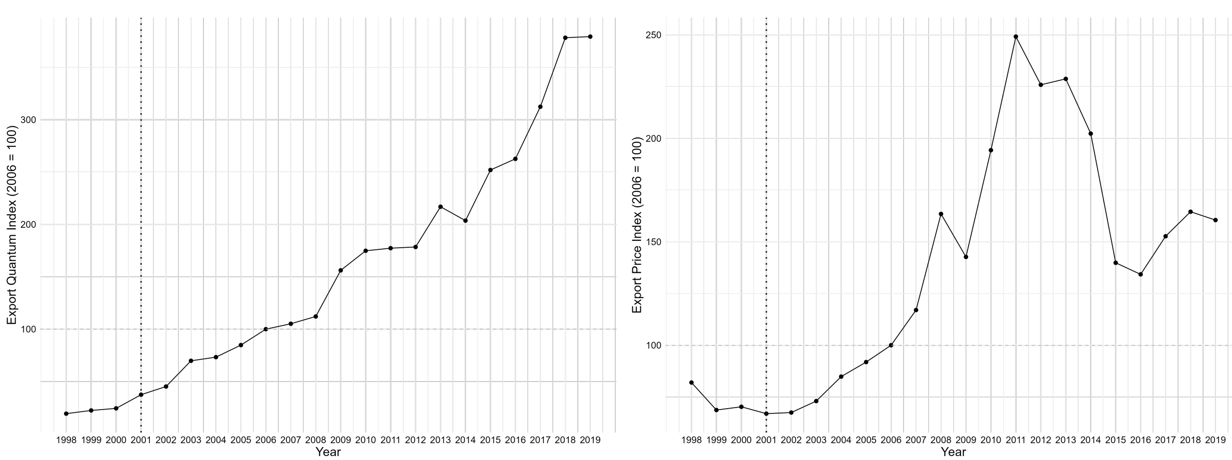
(a) Commodity export quantum index

(b) Commodity export price index

Source: Data on commodity export quantum and price index is from the Brazilian Ministry of Development, Industry, Foreign Trade, and Services, computed by *Secretaria de Comércio Exterior* (SECEX).

Figure 3.4: Total Commodity Exports - Quantum and Price

by the boom in the value of commodity exports. In Panel (b), we unravel the share of total exports from Brazil to China and all other destinations comprising goods from agriculture and extractive sectors. While there is a general increase in the share of these goods in the Brazilian export mix for all countries, this trend is strikingly pronounced concerning China.

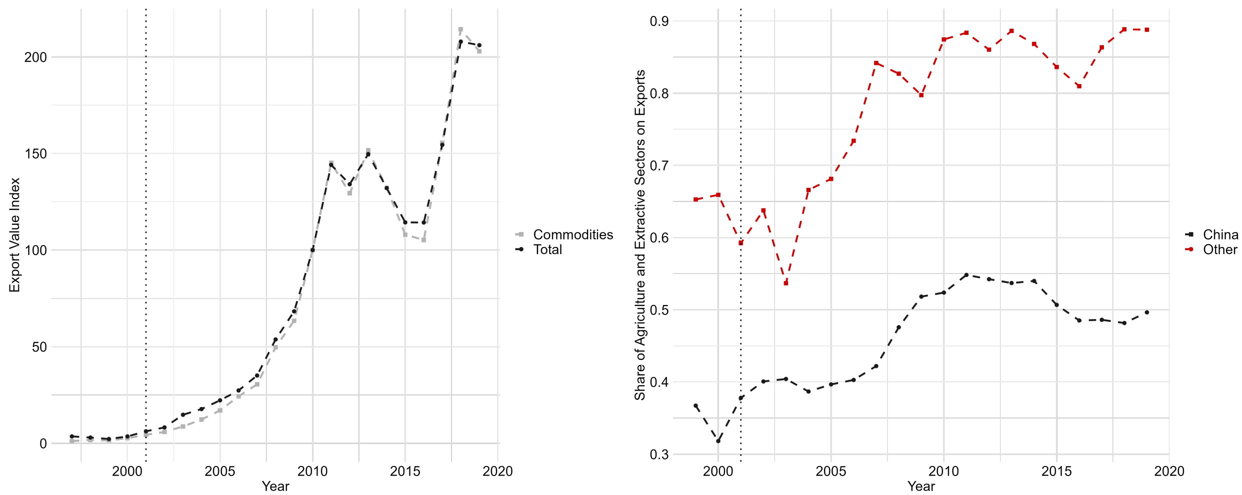


(a) Export quantum index

(b) Export price index

Source: Data on export quantum and price index for China is from the Brazilian Ministry of Development, Industry, Foreign Trade, and Services, computed by *Secretaria de Comércio Exterior* (SECEX).

Figure 3.5: Exports to China - Quantum and Price



(a) Export value index - China (b) Share of agriculture and extractive exports

Source: Data on the value of total exports and commodity exports to China is based on the declaration of exporters in Brazil (SISCOMEX from the Ministry of Industry, Foreign Trade and Services).

Figure 3.6: Exports to China and Other Destinations - Total and Commodities

In summary, the emergence of China as a major global player in trade triggered a resource boom for Brazil, propelling the country's commodity exports, particularly in the agricultural and extractive sectors. This transformation in trade dynamics has had far-reaching implications for the Brazilian economy, with significant consequences for the agricultural sector.

3.3 Theoretical Framework

Elections serve as a cornerstone of democratic representation, offering voters the dual functions of sanctioning ineffective governments and selecting capable representatives (Campello & Zucco, 2020; Fearon, 1999; Przeworski et al., 1999). This distinction is embodied in two primary models: i) sanctioning and ii) selection. In sanctioning models, politicians are viewed as uniform actors primarily concerned with retaining office, with elections serving as contracts where voters set performance thresholds based on retrospective evaluations. Conversely, selection models recognize the diversity among politicians in terms of policy preferences, integrity, and competence. Voters, in this framework, aim to identify and elect representatives who align with their policy preferences and demonstrate competency and integrity, incorporating both past performance and future expectations into their decision-making process (Ashworth, 2012).

However, the reality often presents a blend of both mechanisms, as voters assess past performance to anticipate future behavior while retaining the capacity to impose sanctions if necessary. The existing literature largely explores the circumstances favoring one mechanism over the other, suggesting selection's effectiveness in environments where agents are intrinsically motivated to align with objectives. Selection mechanisms thrive in less corrupt political systems with stable party programs while sanctioning mechanisms dominate in less institutionalized settings with limited voter information (Fearon, 1999; Mansbridge, 2009).

Fearon (1999) developed a canonical theoretical model emphasizing a fundamental aspect of agency models in electoral dynamics: the rational voter's primary aim is to elect the politician promising the highest expected future performance. Thus, despite the potential for politicians to take costly actions, voter behavior during elections revolves around assessing incumbent quality, emphasizing voter learning. However, the selection mechanism within this model leaves little room for agent updating regarding politician types.

We delve into this concept and examine how exogenous economic shocks influence voter behavior and electoral outcomes, building upon the canonical Bayesian learning framework (Achen, 1992; Achen & Bartels, 2017; Bartels, 1993; Wolfers, 2007). Adapting theoretical frameworks from Ashworth et al. (2018) and Wolfers (2007), we directly relate observable exogenous regional economic shocks to the electoral fortunes of incumbents in local elections. This analysis initially aligns with the conventional wisdom that electoral outcomes are unaffected by observable shocks if voters are rational. However, we demonstrate that this conclusion is not robust to minor modifications, prompting hypotheses to be addressed in subsequent analysis.

Crucially, our model is classified as a "'pure selection" theoretical framework, where governance outcomes are solely determined by politician type, observable economic shocks, and idiosyncratic factors. While this assumption simplifies the model, it does not alter its core message. Despite the potential for politicians to undertake actions, voter behavior remains rooted in assessing incumbent quality during elections. While actions by politicians may provide additional information about the incumbent's type, the fundamental logic of voter behavior remains unchanged (Ashworth et al., 2017). Thus, our main insights remain unaffected. In our empirical analysis, we predominantly focus on positive exogenous economic shocks, making the additional information mechanism about incumbent types less relevant compared to cases involving natural disasters. Nonetheless, we evaluate scenarios where the model operates *as if* the sanction mechanism outweighs selection dynamics, which does not impact the main testable hypotheses derived from it.

3.3.1 Baseline model

In every election within a municipality, two politicians compete: an incumbent (I) and a challenger (C). Each politician can be categorized into two possible "types"[30]: either a good (with productivity $\bar{\gamma}$) or a bad ($\underline{\gamma}$) type, directly linked to the achievable governance outcome. The probability of an incumbent being a good type is denoted by $\pi_I \in (0,1)$, while the challenger has a probability $\pi_C \in (0,1)$ of being a good type.

In each period, a representative voter observes a governance outcome. To simplify our analysis, consider that this outcome is directly related to the regional economic conditions observed by the voter. However, this outcome is a noisy indicator of the incumbent's ability, as it reflects both the mayor's efforts and exogenous shocks. Specifically, the outcome is determined by a production function incorporating the incumbent's type and two orthogonal shocks: i) an observable exogenous regional economic shock (x) and, ii) an unobservable idiosyncratic shock ($\varepsilon \sim N(0,1)$).

Following (Ashworth et al., 2018), we begin with a simple additive function for the regional governance outcome in period t, denoted as $g_{r,t}$:

$$g_{r,t} = \gamma_t + x_{r,t} + \varepsilon_t \tag{3.1}$$

[30]This assumption is a simplification from the canonical models. For instance, Wolfers (2007) consider normally distributed types.

where γ_t is the type of the incumbent in period t, $x_{r,t}$ is the *exogenous regional economic shock intensity* (in region or municipality r), and ε_t is the idiosyncratic shock.

Between governance periods, an electoral process occurs. During elections, voters reelect the incumbent if their belief that the incumbent is a good type, based on observed governance outcomes, exceeds their prior belief that the challenger is a good type. Employing Bayes' rule, a voter will reelect an incumbent if and only if:

$$\pi_C \leq \frac{\pi_I \phi\left(g_{r,1} - x_{r,1} - \bar{\gamma}\right)}{\pi_I \phi\left(g_{r,1} - x_{r,1} - \bar{\gamma}\right) + (1 - \pi_I)\phi\left(g_{r,1} - x_{r,1} - \underline{\gamma}\right)} \tag{3.2}$$

where ϕ is the probability density function of the standard normal distribution.

The voter's objective is to select the candidate with the highest type, using observed regional economic shock intensity ($x_{r,1}$) and governance outcomes ($g_{r,1}$) to infer the incumbent's type. The regional governance outcome serves as a biased signal of the incumbent's type due to observed shock intensity. To mitigate this bias, we follow Ashworth et al. (2018) and interpret that the voter "filters out" the shocks' effect from the governance outcome to form posterior beliefs.

Rearranging terms in Equation (3.2), the voter reelects the incumbent if and only if:

$$\frac{\phi\left(g_{r,1} - x_{r,1} - \bar{\gamma}\right)}{\phi\left(g_{r,1} - x_{r,1} - \underline{\gamma}\right)} \geq \left(\frac{\pi_C}{1 - \pi_C}\right)\left(\frac{1 - \pi_I}{\pi_I}\right) \tag{3.3}$$

Similar to Ashworth et al. (2018), note that the left-hand side of this inequality is strictly increasing and continuous in $g_{r,1} - x_{r,1}$. So there is a unique number, κ, such that Equation (3.3) holds with equality when $g_{r,1} - x_{r,1} = \kappa$.

The voter sets a reelection threshold that governance outcomes must meet, adjusting it to offset the economic shock intensity's effect. Specifically, for a given exogenous regional economic shock $x_{r,1}$, the voter's reelection threshold is given by $\hat{g}_r(x_{r,1}) = \kappa + x_{r,1}$. Therefore, the voter reelects the incumbent if:

$$g_{r,1} \geq \hat{g}_r(x_{r,1}) = \kappa + x_{r,1} \tag{3.4}$$

Consequently, if an exogenous shock increases governance outcomes by 1 unit, the voter increases the reelection threshold by exactly 1 unit. As a result, the incumbent's electoral prospects remain unaffected by the presence or magnitude of exogenous economic shocks. Specifically, the probability of reelection is constant regardless of the intensity of the exogenous economic shock, as represented by Equation (3.5). This aligns with the conventional expectation of voter rationality that incumbent electoral fortunes should not be influenced by factors outside their control.

$$Pr\left(\gamma_I + a_1 + \varepsilon_1 \geq \hat{g}_r(a_1)\right) = Pr\left(\gamma_I + \varepsilon_1 \geq \kappa\right) \tag{3.5}$$

3.3.2 Exogenous shocks and politicians' types

So far, our model has not incorporated the interaction between the incumbent's type and the intensity of exogenous economic shocks, which contradicts extensive literature documenting incumbents' responsiveness to phenomena like natural disasters.

To address this omission, we modify the regional governance outcome in period t to include this interaction:

$$g_{r,t} = \left(\gamma_t + x_{r,t}\right)^{\sigma} + \varepsilon_t \tag{3.6}$$

where $\sigma > 0$, and each possible exogenous economic shock is bounded below by $-\underline{\gamma}$ to ensure governance outcomes increase with shock intensity. Setting $\sigma = 1$ reverts us to the special case of additively separable outcomes as in Equation (3.1).

Similar to the analysis in the baseline version of the model, the voter reelects the incumbent in region r if and only if:

$$\frac{\phi\left(g_{r,1} - (\bar{\gamma} + x_{r,1})^{\sigma}\right)}{\phi\left(g_{r,1} - (\underline{\gamma} + x_{r,1})^{\sigma}\right)} \geq \left(\frac{\pi_C}{1-\pi_C}\right)\left(\frac{1-\pi_I}{\pi_I}\right) \tag{3.7}$$

Rearranging terms and remembering that $\varepsilon \sim N(0,1)$, the voter reelects the mayor in region r if and only if:

$$\frac{e^{-\left(g_{r,1} - (\bar{\gamma} + x_{r,1})^{\sigma}\right)^2}}{e^{-\left(g_{r,1} - (\underline{\gamma} + x_{r,1})^{\sigma}\right)^2}} \geq \left(\frac{\pi_C}{1-\pi_C}\right)\left(\frac{1-\pi_I}{\pi_I}\right) \tag{3.8}$$

Taking logs on both sides, we can again calculate the voter's reelection threshold in a simplified manner as follows:

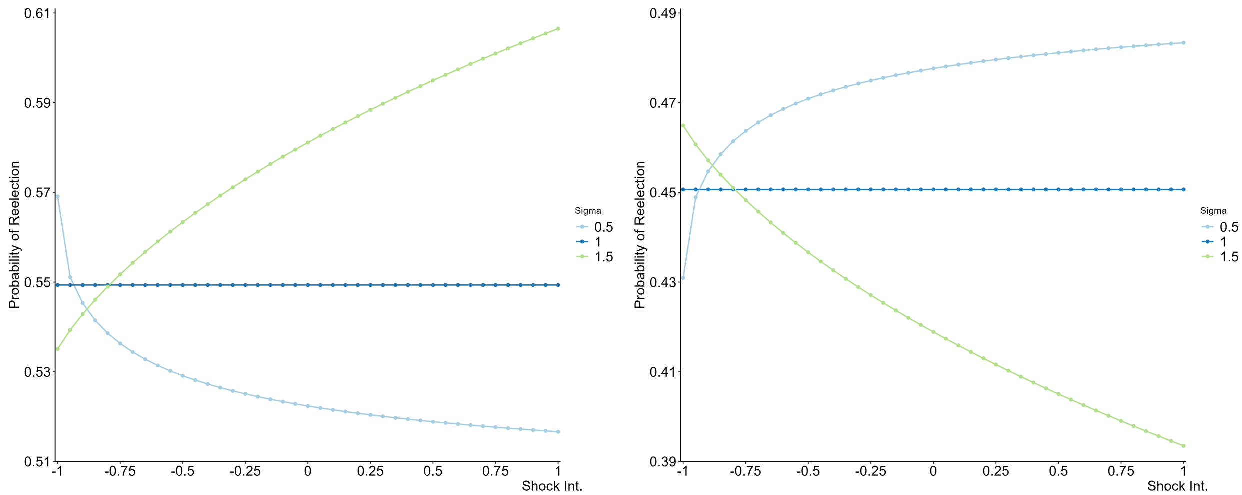
$$\hat{g}_r(x_{r,1}) = \frac{\ln\left(\frac{\pi_C}{1-\pi_C}\right)\left(\frac{1-\pi_I}{\pi_I}\right)}{2\left[(\bar{\gamma}+x_{r,1})^{\sigma}-(\underline{\gamma}+x_{r,1})^{\sigma}\right]} + \frac{(\bar{\gamma}+x_{r,1})^{\sigma}+(\underline{\gamma}+x_{r,1})^{\sigma}}{2} \tag{3.9}$$

The probability that the incumbent in the local election is reelected is, again, the probability that $g_1 \geq \hat{g}_r(x_{r,1})$ or, equivalently, the probability that $(\gamma_I + x_{r,1})^{\sigma} + \varepsilon_1 \geq \hat{g}_r(x_{r,1})$. We can write this as follows:

$$Pr\left(g_{r,1} \geq \hat{g}_r(x_{r,1})\right) = 1 - \left[\pi_I \Phi\left(\hat{g}_r(x_{r,1}) - (\bar{\gamma}+x_{r,1})^{\sigma}\right) + (1-\pi_I)\Phi\left(\hat{g}_r(x_{r,1}) - (\bar{\gamma}+x_{r,1})^{\sigma}\right)\right] \tag{3.10}$$

where Φ represents the cumulative density function of the standard normal distribution.

Analyzing Equations (3.9) and (3.10), we can compute the probability of reelection as a function of the exposure to the exogenous regional economic shock. Notably, if voters perceive an equal probability for any candidate to be a good type of politician *ex-ante*, the probability of reelection varies based on shock intensity. This implies a direct impact of regional economic shock exposure on the electoral fortunes of an incumbent, notwithstanding the presence of a proxy for the sanctioning mechanism potentially outweighing the pure selection mechanism as discussed earlier. For instance, Figure 3.7 illustrates this variation for different assumptions of candidate probabilities, highlighting implications for voter rationality. In Figure 3.7 we present the results considering that $\bar{\gamma} = \frac{3}{2}$ and $\underline{\gamma} = \frac{1}{2}$. In Panel (a), we plot the optimistic case, in which any candidate has a probability of $\frac{3}{4}$ of being a good type. Panel (b), shows the pessimistic case, in which any candidate has a probability of $\frac{1}{4}$ of being a good type.



(a) Optimistic case – $\pi_I = \pi_C = \frac{3}{4}$

(b) Pessimistic case – $\pi_I = \pi_C = \frac{1}{4}$

Figure 3.7: Reelection probability and shock intensity: the sanctioning variation

In our analysis of these simple examples, we draw preliminary insights into the rationality of voter behavior. A crucial aspect lies in the parameter σ, which determines how electoral outcomes respond to observable shocks. When $\sigma = 1$, our model predicts electoral outcomes to remain unresponsive to such shocks. However, this contradicts empirical evidence suggesting that voters do react to observable shocks (Achen & Bartels, 2004, 2017; Healy et al., 2010; Wolfers, 2007). This inconsistency implies potential voter irrationality. Conversely, for σ values such as $\frac{1}{2}$ or $\frac{3}{2}$, indicative of rational voter behavior, we expect electoral fortunes to react sensitively to observable shocks. Specifically, when $\sigma = \frac{1}{2}$, we observe a decrease in the probability of re-election with heightened exposure to shocks in one scenario and an increase in the other, illustrating a reversed pattern when $\sigma = \frac{3}{2}$.

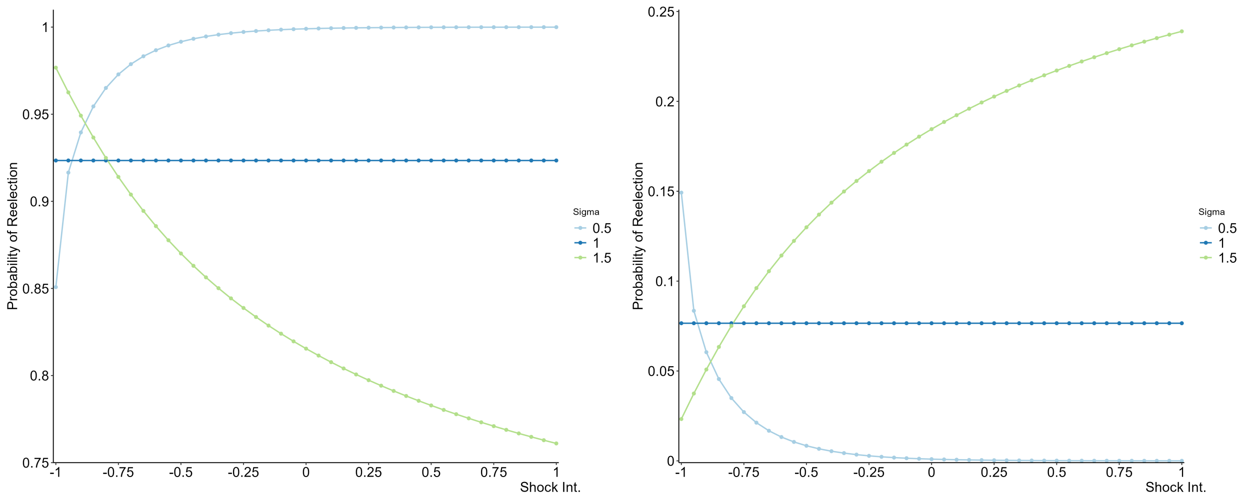
Ashworth et al. (2018) aptly highlight the absence of a principled basis for determining a specific value for σ from the outset, presenting a significant challenge in interpreting the model's implications for empirical findings and drawing inferences about voter rationality. Justifying the choice of σ necessitates a substantive argument aligning with the reality depicted by the model. However, the current selection of σ lacks such substantiation.

To illustrate this challenge, consider a straightforward scenario based on our previous discussion, where positive exogenous regional economic shocks result from the introduction of new agricultural technology or an export demand shock across Brazilian municipalities. Here, consider that these shocks significantly shape how voters perceive the competence (or type) of incumbent politicians. In the absence of a shock ($x_{r,1} = 0$), voters remain uninformed regarding type, but when regions experience positive economic shocks ($x_{r,1} > 0$), voters gain insights. Consequently, economic shock intensity could significantly influence electoral outcomes. It becomes apparent that how shock exposure influences voter information acquisition is pivotal in comprehending the impact of economic shocks on voter behavior and the electoral fortunes of incumbents. This discussion underscores that determining the value of σ in the model hinges on a substantive assumption directly linked to qualitative features of the political environment. Specifically, it dictates how exogenous shocks shape voter information acquisition: if $\sigma < 1$, heightened exposure to regional economic shocks implies increased voter information acquisition about the incumbent's type, while $\sigma > 1$ suggests reduced voter information acquisition. Notably, $\sigma = 1$ represents the sole scenario where shock exposure fails to impact voter information acquisition.

This insight is directly observable in Figure 3.7. In both cases, where politicians have equal probabilities of being the good type, complete information would render voters indifferent between reelecting the incumbent or electing the competitor, resulting in a reelection probability of $\frac{1}{2}$. However, in Panel (a), where voters hold optimistic views regarding the average quality of politicians, greater exposure to exogenous regional economic shocks increases voter information acquisition regarding incumbents' types, leading to a re-election probability approaching 50%. Conversely, in Panel (b), we observe a similar pattern for the pessimistic case, with the probability of re-election still tending towards 50% with greater exposure to shocks when voters derive enhanced information regarding types.

We proceed to assess more realistic scenarios wherein voters initially perceive a discrepancy in the *ex-ante* probabilities of candidates being of the "good" type. Figure 3.8 illustrates these scenarios, with Panel (a) presenting a scenario where the incumbent holds a probabilistic advantage

($\pi_I = \frac{2}{3}$) over the competitor ($\pi_C = \frac{1}{3}$). Conversely, Panel (b) depicts the inverse scenario, where the competitor is assumed, on average, to be of the good type ($\pi_I = \frac{1}{3}$) compared to the incumbent ($\pi_C = \frac{2}{3}$).



(a) Type advantage – $\pi_I = \frac{2}{3} \wedge \pi_C = \frac{1}{3}$

(b) Type disadvantage – $\pi_I = \frac{1}{3} \wedge \pi_C = \frac{2}{3}$

Figure 3.8: Reelection probability as a function of shock intensity - different probabilities

The interpretation of these scenarios closely mirrors the previous one. In Panel (a) of Figure 3.8, where the incumbent is more likely to be of the "good" type compared to the competitor, the probability of reelection tends towards one with greater exposure to the regional economic shock when such intensity increases the informational level of voters. Conversely, in Panel (b), when the competitor has the probabilistic advantage, the probability of reelection tends to zero with greater exposure to shocks, in the case of heightened information acquisition by voters.

3.3.3 Predictions and testable hypotheses

Our primary objective in this paper is to examine how exogenous economic shocks, beyond the control of incumbents, impact the likelihood of their reelection in municipal elections. Drawing upon our theoretical framework, we posit that the interaction between the intensity of these shocks and the inherent competence associated with a politician's type significantly influences reelection probabilities.

In the simplified scenario of additively separable governance outcomes, exposure to shocks does not sway reelection probabilities. This is because voters tend to discount this exposure when assessing incumbent competence amidst noisy governance outcomes. However, moving beyond this simplified framework, our theoretical-formal model explores how shock exposure can either amplify or diminish the effect of incumbent competence on governance outcomes. Amplification occurs when heightened exposure to exogenous economic shocks enhances the impact of incumbent competence, leading to more favorable outcomes in terms of selection. For instance, skilled incumbents may demonstrate superior political ability and budget management during economic shocks, resulting in increased reelection prospects. Conversely, muting occurs when shocks overshadow incumbent competence, yielding similar outcomes regardless of incumbent quality. Understanding these dynamics is essential for comprehending the repercussions of economic shocks on governance outcomes and deciphering the role of incumbent competency within such contexts.

With these insights in mind, our first testable hypothesis stemming from the theoretical-formal model is whether exogenous regional economic shocks influence the reelection probability of mayoral incumbents, and if so, in what manner?

Furthermore, while we can speculate on how economic shocks might magnify or diminish the effect of politician type on reelection probabilities based on our theoretical predictions, empirical evaluation is crucial. By controlling for factors correlated with competent incumbents (e.g., provision of public goods, enhancements in human development indicators), we are potentially able to provide illuminating evidence in that regard. Thus, the second testable hypothesis that arises from the theoretical framework is whether the intensity of exogenous regional economic shocks amplifies or mutes the effect of type on voters' decisions in local elections.

3.4 Data Description and Empirical Strategy

The theoretical framework outlined in Section 3.3 lays the groundwork for two key propositions that we aim to empirically investigate. In this section, we detail the data sources and the methodology employed to empirically evaluate these predictions.

3.4.1 Data description

In our analysis, we rely on a diverse range of data sources to explore both electoral and economic dynamics within Brazil's municipalities. These data sources provide insights into the outcomes of municipal elections, the impact of technological changes in agriculture, and the effects of China's growing demand for commodities on local economies.

Electoral data

We commence our analysis by accessing electoral data from Brazil's Superior Electoral Court, known as the *Tribunal Superior Eleitoral* or TSE. Our primary focus centers on municipal elections, which occur every four years, and we analyze these events starting from 2000 (pre-treatments) onwards.

Within this electoral dataset, we pay particular attention to key variables at the candidate level associated with the election outcomes. We focus on the share of votes for each mayoral candidate in each municipality, their party affiliations, and their characteristics. Particularly, we construct a measure of incumbency at the candidate level that considers if the candidate won the previous mayoral election in a given municipality. This binary variable, along with the electoral outcome, are the main interest variables for our preliminary empirical strategy.

However, it is essential to underscore our intention to expand incumbency measures in the future to enhance the robustness of our results and consider potential mechanisms underlying the amplification of incumbency effects. For instance, we plan to broaden our measure of incumbency to include not only the winner of the previous election as the incumbent but also potential political successors, as re-election in Brazil is limited to one term. Constructing a more comprehensive measure of incumbency might involve considering candidates from the previous winning coalition as incumbents, thereby providing a more nuanced perspective.

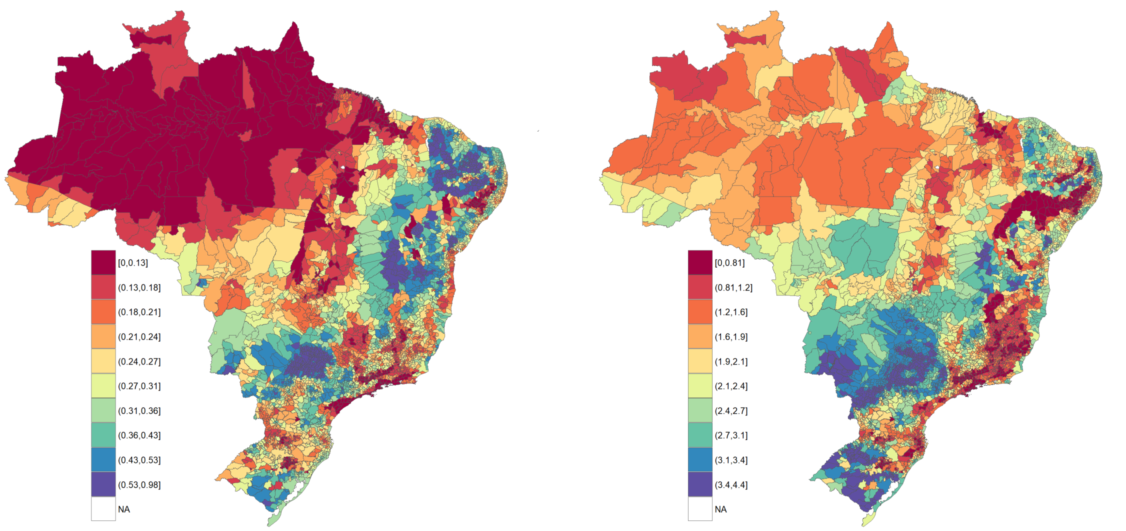
Productivity shock

Although the adoption of the GE soybean seed can be seen as exogenous to previous production conditions, it is possible to argue that the expansion of planted area may have some correlation with specific factors of Brazilian municipalities or even individual characteristics of decision makers. Consequently, to study the impacts of the technological change shock on political outcomes at the municipality level, we need to construct an exogenous measure of this technological change.

We obtain a measure of potential soy yields in different Brazilian regions from the FAO Global Agro-Ecological Zones (FAO-GAEZ) database. These yields are calculated by incorporating local soil and weather characteristics into an agronomic model that predicts the maximum attainable yield for each crop in a given area. Crucially for our analysis, the FAO-GAEZ database reports potential yields under different technologies or input combinations. The low input potential yields consider traditional production technology with minimal use of modern inputs such as fertilizers and herbicides. The high input potential yields consider the use of modern inputs such as GE seeds and fertilizers.

Figure 3.9 presents our measure of the potential yield of soy production, in terms of tonnes per hectare, under the low and high agricultural technology at the municipality level in Brazil. For clarity of exposition, we aggregate the potential yields into deciles. It is important to note the large variation in production capacity for the municipalities in the top deciles of the distribution: even for the very productive regions ex-ante, the potential yield that could be achieved in terms of tonnes of soy per hectare with the new technology is 3 to 5 times larger.

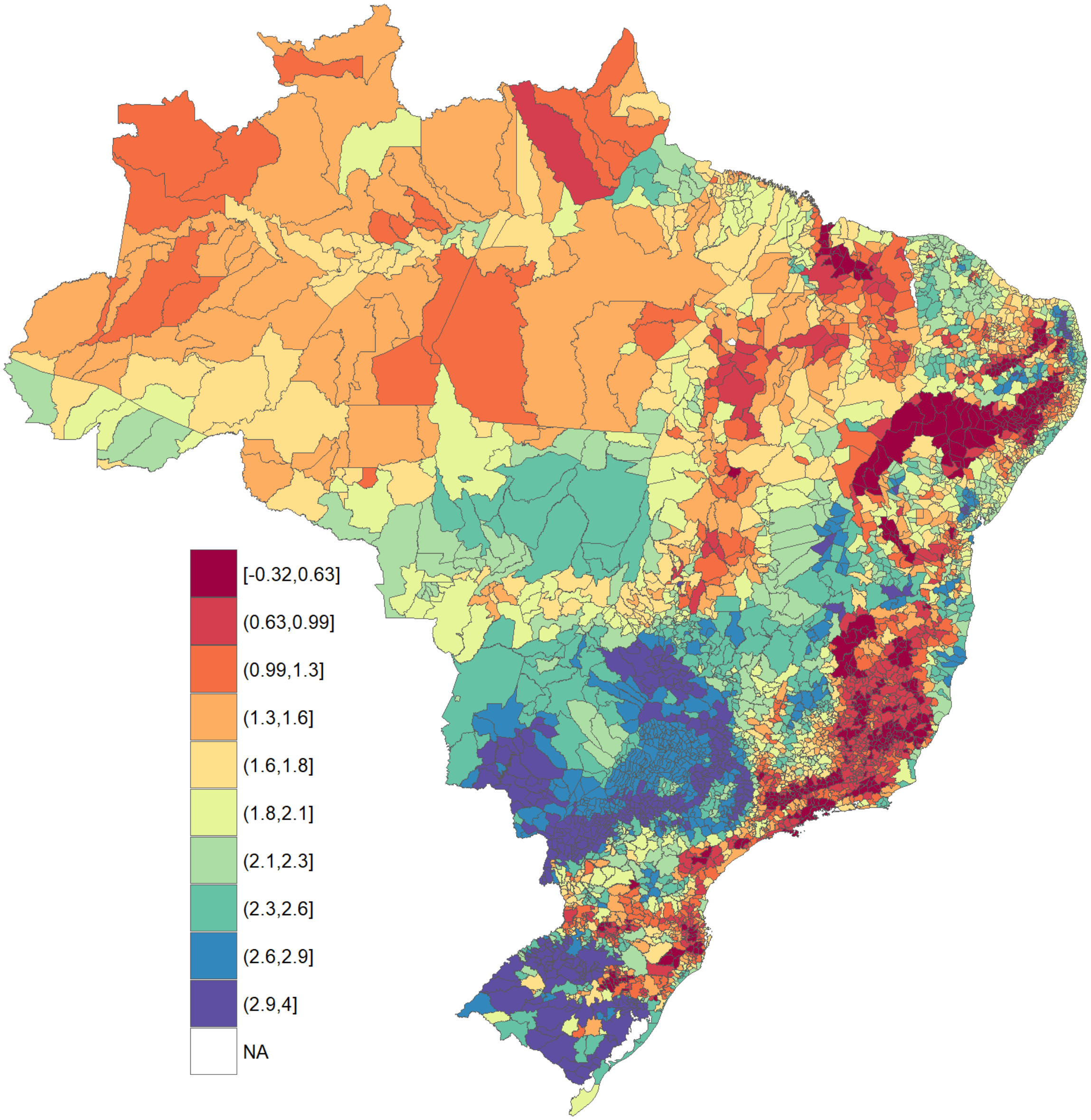
We construct a measure of local exposure to productivity gains from adopting GE soy seeds in municipality r as the difference in potential yield in the high and low input scenarios aggregated at the municipality level. This difference in yields between the high and low technology captures the effect of moving from traditional agriculture to a technology that uses improved seeds and optimum weed control, among other characteristics. Figure 3.10 illustrates the resulting measure of technical change in soy at the municipality in Brazil.



(a) Low (b) High

Source: Potential soy yields, in terms of tons per hectare, using low and high agricultural technology at the municipality level are constructed using FAO-GAEZ data.

Figure 3.9: Potential Soy Yield under Low and High Agricultural Technology



Source: Potential soy yields, in terms of tons per hectare, using low and high agricultural technology at the municipality level are constructed using FAO-GAEZ data.

Figure 3.10: Technological Change in Soy: Municipalities

Commodity boom

Our analysis also examines the effects of China's increasing demand for commodities, often referred to as the "commodity boom". To investigate this phenomenon, we gather data on international trade in goods from the BACI database developed by *Centre d'Etudes Prospectives et d'Informations Internationales* (CEPII), which reconciles the data separately reported by importers and exporters in the United Nations Statistical Division's COMTRADE database. This data is largely used in the international trade literature and contains the total annual value of bilateral trade at the 6-digit level of the Harmonized System classification for more than 200 countries from 1995 to 2022. We use data for 2000 and 2019 in the analysis below. Given that the CEPII data is denominated in thousands of current US dollars, we adjust 2000 values to 2019 US dollars employing the US GDP deflator from the US Bureau of Economic Analysis.

To craft a local-level measure gauging the impact of heightened export demand from China, our empirical strategy necessitates categorizing employed individuals in the 2000 census data and products in the 2000 and 2019 trade data into sectors. For this purpose, we leverage individual-level labor market and socioeconomic data from the Brazilian Demographic Census for 2000 and 2010, sourced from the Brazilian National Statistical Institute (IBGE). We narrow down our sample to the sub-population most likely to partake in the labor market, defining the workforce as individuals aged between 18 and 60 years. Subsequently, we calculate the sector employment share by dividing the workforce in each sector by the total workforce in each municipality. The 2000 Brazilian census captures individuals' sector of activity according to the 5-digit *"CNAE Domicílio"* classification. We thus construct a concordance, following Costa et al. (2016), assigning products in the trade data to *"CNAE Domicílio"* sectors, yielding a total of 82 traded goods sectors, including 24 agricultural sectors, 8 extractive sectors, and 50 manufacturing sectors.

These steps allow us to generate a measure of the exposure of local economies (municipalities) in Brazil to the heightened export demand for China. Nonetheless, the challenge of the empirical strategy is to obtain an exogenous measure of the increase in Chinese demand for commodities at the local labor market level. Nonetheless, drawing from the existing literature (Autor et al., 2013, 2014; Carreira et al., 2024; Costa et al., 2016), we can construct a plausible exogenous measure of local exposure to China's growing demand for commodities - the "commodity boom" - for each municipality in Brazil as follows.

We start by first removing the influence of world- and Brazil-specific shocks in the observed changes in trade between Brazil and China.[31] We do so by running the following auxiliary regressions for all countries except Brazil, weighted by the initial import values:

$$\frac{\Delta \tilde{I}_{ij00/19}}{\tilde{I}_{ij00}} = \beta_j + \psi_{China,j} + v_{ij} \tag{3.11}$$

where $\frac{\Delta \tilde{I}_{ij00/19}}{\tilde{I}_{ij00}}$ is the growth rate in imports of product j by country i from all countries other than Brazil between 2000 and 2019; β_j is the product fixed effect that captures the world average growth of net-of-Brazil imports of product j; $\psi_{China,j}$ is a China-product specific dummy that captures the deviation of the China import growth rate of product j in comparison to the one from the rest of the world. The estimated $\hat{\psi}_{China,j}$ captures the predicted change in global exports to China (excluding Brazil) induced by China-specific factors between 2000 and 2019.

Next, we can construct a measure of local exposure to China-induced export demand using the estimated $\hat{\psi}_{China,j}$:

$$\Delta \tilde{X}_r = \frac{1}{L_{r00}} \sum_j \frac{L_{rj_{00}}}{L_{Bj00}} \times \frac{X_{BCj00} \hat{\psi}_{China,j}}{100} \tag{3.12}$$

where L_{i00} is the size of the workforce in region r in 2000; L_{ij00} is the size of the workforce in sector j within region r in 2000; L_{Bj00} is the Brazil-wide workforce in sector j in 2000 and X_{BCj00} is the Brazilian exports of product j to China in 2000. We follow the literature and divide the Bartik-instrument or shock by 100 to better visualize the coefficients in our main results (Autor et al., 2013; Goldsmith-Pinkham et al., 2020).

[31] This process is similar to the construction of the Chinese import competition measure for the US economy - or the "China shock"-, for instance in Autor et al. (2013, 2014).

Figure 3.11 plots the map of local exposure to the Chinese demand computed from Equation (3.12). We see that municipalities in the South, Center-West, and North regions of Brazil were among the most exposed to China-induced growth of export demand.

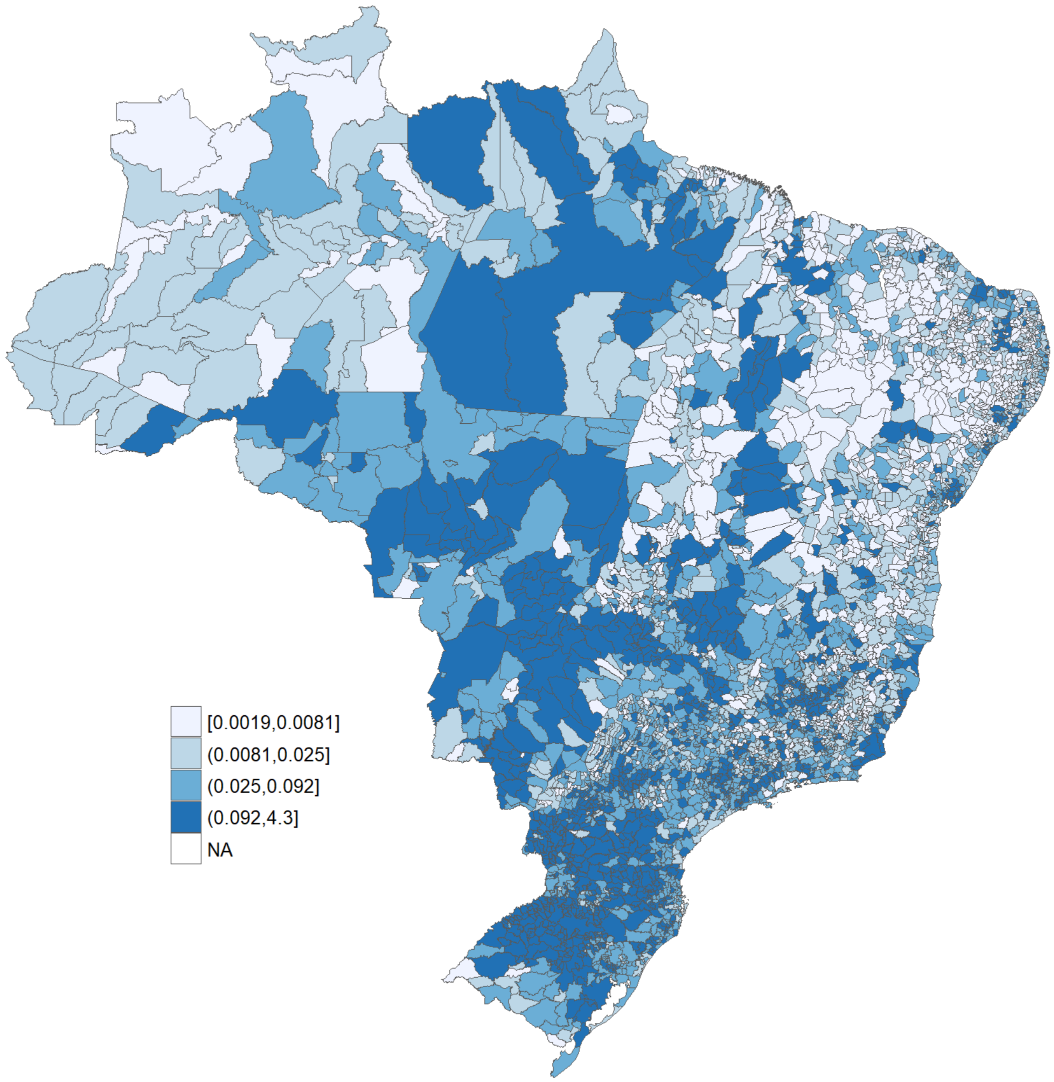


Figure 3.11: Exposure to China's Export Demand - $\Delta \tilde{X}_r$

Additional variables

In addition to electoral and economic data, we incorporate information from other sources to create a comprehensive dataset for analysis.

We use data from the agricultural census from IBGE, collected every ten years (1996, 2006, and 2017). The agricultural census data are collected through direct interviews with the managers of each agricultural establishment and are made available online by the IBGE aggregated at the municipality level. The agricultural variables of interest are the share of land planted with soy, the value of production per worker, and labor intensity. The last two variables are aggregated across all agricultural activities. This is because the unit of observation in the census is the agricultural establishment, and these tend to perform several activities. As a result, it is not possible to obtain a measure of employment by crop.

Additionally, we use the Brazilian population census to construct baseline control variables for each municipality. As for the shift-share instrument, we use data from the last two rounds of the census (2000 and 2010). This allows us to observe the variables of interest before and after the legalization of GE soy seeds and the outset of the commodity boom.

3.4.2 Empirical strategy and identification

To empirically evaluate the main theoretical predictions of the formal-theoretical model developed earlier using the data described above, our first objective is to disentangle the effects of two distinct factors on incumbency effects in Brazil: exposure to agriculture productivity shocks and increased international demand for commodities.

First, regarding the exposure to the new GE soy seed, our approach seeks to identify the causal effect of new agricultural technologies on electoral outcomes by strategically utilizing the timing of their adoption and their varied impact on potential yields across geographical regions. To establish the exogeneity of the timing of adoption concerning Brazilian economic developments, it is important to note that genetically engineered (GE) soy seeds were first commercially released in the United States in 1996. Given that these seeds originated in the United States, their approval for commercialization in 1996 can be viewed as exogenous to developments in the Brazilian economy, as it was driven by factors external to Brazil's economic circumstances.

Notwithstanding, the legalization of GE soy seeds in Brazil occurred in 2003, influenced in part by pressure from Brazilian farmers. However, it is crucial to acknowledge that reports of GE soy seed smuggling from neighboring Argentina date back to 2001. In our empirical analysis, our ideal approach is to compare outcomes before and after 2001. This choice is significant because the year 2000, our first mayoral election on the sample, precedes both the legalization of GE soy seeds and the initial reports of smuggling, ensuring the exogeneity of the timing in our analysis.

Furthermore, as highlighted earlier, the impact of this new technology on potential yields varied based on soil and weather characteristics. We intend to capitalize on these exogenous differences in potential yields across geographical areas as a source of cross-sectional variation in treatment intensity, utilizing the previously constructed measure ΔA_r^{soy}.

Concerning the trade shock, recent research has developed a formal framework to establish identifying assumptions for shift-share regression designs (Borusyak et al., 2022; Goldsmith-Pinkham et al., 2020). Building on the work of Goldsmith-Pinkham et al. (2020) and Borusyak et al. (2022), the identifying assumption is that the China-induced trade shock, denoted as $\Delta\tilde{X}_r$, is independent of local political and institutional dynamics across municipalities in Brazil. This independence is largely guaranteed by our consideration of the relative effects of increased demand from China to all other countries in the world, excluding Brazil. Therefore, our "Bartik" instrument can be regarded as good as an exogenous shock to local political dynamics and choices.

Furthermore, it is crucial to examine the relationship between these two exogenous shocks. Our identification strategy hinges on the fact that the two analyzed exogenous economic shocks exhibit a low positive correlation (slightly below 0.05), ensuring they are not highly colinear. This variation in differential exposure to both shocks is essential, as it enables us to dissect and study the distinct effects of each on incumbency effects. Without this essential variation in differential exposure to both shocks, it would be impossible to discern the unique contributions of each to changes in incumbency effects across municipalities. Notably, there exists significant spatial diversity in the degree of exposure to these two shocks, as illustrated in Figures 3.10 and 3.11.

To pursue our empirical objective, initially evaluating the first theoretical prediction derived from our theoretical model, and with the measures constructed for local exposure to these arguably exogenous economic shocks in mind, we utilize the following specification with mayoral candidates as the unit of analysis:[32]

$$\begin{aligned} y_{j,r,t} = \beta_0 + \beta_1 Incumbency_{j,r,t-1} + \beta_2 \Delta A_r{}^{soy} + \beta_3 \Delta \tilde{X}_r + \beta_4 \left(\Delta A_r{}^{soy} \times \Delta \tilde{X}_r \right) + \\ \beta_5 \left(Incumbency_{j,r,t-1} \times \Delta A_r{}^{soy} \right) + \beta_6 \left(Incumbency_{j,r,t-1} \times \Delta \tilde{X}_r \right) + \\ \beta_7 \left(Incumbency_{j,r,t-1} \times \Delta A_r{}^{soy} \times \Delta \tilde{X}_r \right) + \alpha_{r,t} + \varepsilon_{j,r,t} \quad (3.13) \end{aligned}$$

where $y_{j,r,t}$ is the electoral outcome variable for candidate j in municipality r at election year $t \in \{2004, 2008, 2012, 2016\}$, $\Delta A_r{}^{soy}$ is the potential yield of soy under the high technology minus the potential yield of soy under low technology, $\Delta \tilde{X}_r$ is the measure of local exposure, for municipality r, to the China-induced export demand shock and $\alpha_{r,t}$ are region-time fixed effects (we explore different possibilities in the preliminary results). This comprehensive framework allows us to rigorously analyze the intricate interplay between these exogenous economic factors and incumbency effects in Brazilian mayoral elections.

To relate such a specification to our first theoretical prediction, note that we are directly interested in the statistical significance of the interaction terms between economic shocks and the binary incumbency variable. That is, our main interest lies on the coefficients β_5, β_6 and β_7. In either case or jointly, we are interested in evaluating whether exposure to exogenous regional economic shocks has a statistically significant impact on the electoral fortunes of incumbents in localities more impacted by these shocks compared to less impacted regions. Understanding the

[32] We follow the recent literature on shift-share analyses to construct the Bartik instrument and use it directly as an independent variable, instead of generating an instrumental variable for the change in exports demand from China. For a detailed analysis of the methods see, for instance, Adao et al. (2019), Borusyak et al. (2022), and Goldsmith-Pinkham et al. (2020).

direction of this impact is equally crucial. By assessing the statistical significance and sign of these coefficients, we can derive initial insights into how exogenous economic shocks shape voter perceptions of incumbent mayors. This analytical approach enables us to interpret how voters absorb information about different types of politicians, as predicted by our theoretical framework.

In addition to our primary specification, we explore the relative effects of shocks over time to enhance the robustness of our analysis and identify avenues for future research. To achieve this, we modify our regression model in (3.13) to pool data across all years, using 2004 as the base year.[33] This dynamic specification can be described as follows:

$$y_{j,r,t} = \sum_{t=2004}^{2020} \mathbb{1}\left\{\tau = t\right\} \Big[\alpha Incumbency_{j,r,t-1} + \delta_1 \left(Incumbency_{j,r,t-1} \times \Delta A_r{}^{soy}\right) \\ + \delta_2 \left(Incumbency_{j,r,t-1} \times \Delta \tilde{X}_r\right) + \delta_3 \left(Incumbency_{j,r,t-1} \times \Delta A_r{}^{soy} \times \Delta \tilde{X}_r\right)\Big] \\ + \mu_{r,t} + \varepsilon_{j,r,t} \quad (3.14)$$

where $\mu_{r,t}$ represent the municipality-election or municipality-year fixed effects. The term $\mathbb{1}$ is a year indicator.

Lastly, it is important to highlight that, to empirically investigate the second hypothesis derived from our theoretical framework, further data collection is imperative. Specifically, we need to gather data on public goods provision, economic indicators, and social development measures at the municipal level throughout the analysis period. Additionally, incorporating control variables related to candidates can help mitigate potential confounding factors associated with politicians' competence and refine our empirical results *vis-à-vis* the theoretical model. This comprehensive approach, to be followed in the coming months, will strengthen our understanding of the relation-

[33] Unfortunately, we have not been able to construct measures of incumbency for previous mayoral elections so far. Hence, the canonical difference-in-differences specification here must be taken with "a grain of salt" due to the impossibility of testing the parallel trend assumption up to this point. This is an important point to focus on in the coming months.

ship between economic shocks, political incumbency, and electoral outcomes, paving the way for a more conclusive research effort.

3.5 Preliminary Results

3.5.1 Incumbency effects

In this section, we delve into the ramifications of exogenous regional economic shocks on electoral outcomes, particularly focusing on incumbents versus challengers in mayoral elections across Brazil.[34] To assess this relationship, we employ Equation (3.13), which includes three pivotal interaction terms: i) the interaction between incumbency and soy technical change, ii) the interaction between incumbency and the trade shock, and, iii) the interaction of both exogenous shocks with incumbency. We also introduce locality fixed effects at two levels of aggregation, including micro-region[35] and municipality, to control for specific local factors influencing our analysis. Besides, in all specifications, we cluster the standard errors at the micro-region level to account for potential spatial correlation in outcomes.

First, we evaluate the impact of the exogenous economic shocks on the incumbent's probability of reelection in each electoral process - that is, in each year of the sample separately. Tables 3.2 and 3.3 present preliminary findings from our estimation of Equation (3.13) for mayoral elections in Brazil between 2004 and 2020. It is essential to note that the incumbency variable is defined at $t-1$. In other words, for instance for the 2004 election, a candidate is considered an incumbent if they won the mayoral election in that municipality in 2000. Importantly, in Table 3.2 we only consider micro-region fixed effects to be able to assess the effects of the exogenous shocks on the overall probability of victory of any candidate. Conversely, Table 3.3 presents the preferred specification with municipality-year fixed effects.

[34]It is important to highlight that we present extensive evidence regarding the validity of the instruments in the Appendix C.

[35]These micro-regions involve groupings of economically integrated municipalities with similar geographic and productive characteristics. They are delineated by the Brazilian Institute of Geography and Statistics (IBGE) and are commonly used in economic literature to characterize regional economies in Brazil (Costa et al., 2016; Dix-Carneiro & Kovak, 2017; Dix-Carneiro et al., 2018).

Table 3.2: Exogenous shocks and incumbency effects: micro-region fixed effects

Dep. var.: *Probability of victory*																				
Election year	2004				2008				2012				2016				2020			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)	(17)	(18)	(19)	(20)
$Incumbency_{j,r,t-1}$	0.224***	0.257***	0.227***	0.232***	0.305***	0.380***	0.303***	0.278***	0.234***	0.211***	0.232***	0.232***	0.151***	0.143***	0.153***	0.146***	0.377***	0.408***	0.377***	0.381***
	(0.0297)	(0.0130)	(0.0298)	(0.0310)	(0.0236)	(0.0114)	(0.0239)	(0.0253)	(0.0294)	(0.0134)	(0.0295)	(0.0317)	(0.0325)	(0.0136)	(0.0326)	(0.0335)	(0.0261)	(0.0117)	(0.0262)	(0.0276)
ΔA_r^{soy}	-0.0114**		-0.0117***	-0.0145***	-0.0289***		-0.0287***	-0.0348***	-0.00862*		-0.00847*	-0.0102*	-0.0121**		-0.0120**	-0.0159***	-0.0220***		-0.0220***	-0.0240***
	(0.00450)		(0.00449)	(0.00477)	(0.00549)		(0.00548)	(0.00568)	(0.00496)		(0.00496)	(0.00519)	(0.00548)		(0.00547)	(0.00574)	(0.00482)		(0.00482)	(0.00509)
$\Delta \tilde{X}_r$		0.00633	0.00683*	-0.00913		-0.0126**	-0.0120**	-0.0512***		-0.00723	-0.00712	-0.0167		0.000704	0.000694	-0.0218**		-0.00334	-0.00300	-0.0140
		(0.00396)	(0.00400)	(0.00781)		(0.00541)	(0.00547)	(0.0115)		(0.00503)	(0.00502)	(0.0113)		(0.00530)	(0.00525)	(0.0108)		(0.00503)	(0.00499)	(0.0112)
$\Delta A_r^{soy} \times \Delta \tilde{X}_r$				0.0104**				0.0242***				0.00619				0.0145**				0.00706
				(0.00500)				(0.00619)				(0.00532)				(0.00632)				(0.00579)
Interactions																				
Incumbency × GE soy	0.0159		0.0172	0.0140	0.0430***		0.0427***	0.0586***	-0.0116		-0.0124	-0.0121	-0.00664		-0.00534	-0.000927	0.0169		0.0171	0.0146
	(0.0154)		(0.0153)	(0.0161)	(0.0123)		(0.0123)	(0.0135)	(0.0150)		(0.0151)	(0.0166)	(0.0158)		(0.0156)	(0.0164)	(0.0131)		(0.0131)	(0.0139)
Incumbency × Chinese demand		-0.0221	-0.0242	-0.0486		0.0154	0.0135	0.121***		0.0122	0.0143	0.0152		-0.0218	-0.0207	0.00752		-0.00213	-0.00373	-0.0203
		(0.0219)	(0.0218)	(0.0478)		(0.0153)	(0.0159)	(0.0282)		(0.0228)	(0.0228)	(0.0472)		(0.0199)	(0.0197)	(0.0401)		(0.0166)	(0.0166)	(0.0343)
Incumbency × Both				0.0133				-0.0658***				-0.00103				-0.0170				0.00983
				(0.0246)				(0.0189)				(0.0224)				(0.0232)				(0.0171)
Micro-region FE	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X
Municipality FE																				
Observations	15,268	15,268	15,268	15,268	14,401	14,401	14,401	14,401	14,714	14,714	14,714	14,714	14,964	14,964	14,964	14,964	17,279	17,279	17,279	17,279
R^2	0.053	0.053	0.053	0.054	0.122	0.121	0.122	0.123	0.047	0.047	0.047	0.047	0.035	0.035	0.035	0.036	0.139	0.138	0.139	0.139

Notes: Standard errors (in parentheses) are adjusted for 556 micro-region clusters. Unit of analysis is a mayoral candidate. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 3.3: Exogenous shocks and incumbency effects: municipality-year fixed effects

Dep. var.: *Probability of victory*																				
Election year	2004				2008				2012				2016				2020			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)	(17)	(18)	(19)	(20)
$Incumbency_{j,r,t-1}$	0.249***	0.291***	0.251***	0.260***	0.337***	0.421***	0.335***	0.305***	0.258***	0.226***	0.256***	0.257***	0.164***	0.138***	0.166***	0.159***	0.403***	0.435***	0.405***	0.413***
	(0.0438)	(0.0190)	(0.0440)	(0.0456)	(0.0349)	(0.0165)	(0.0354)	(0.0376)	(0.0448)	(0.0203)	(0.0450)	(0.0482)	(0.0481)	(0.0200)	(0.0482)	(0.0495)	(0.0367)	(0.0160)	(0.0370)	(0.0389)
Interactions																				
Incumbency × GE soy	0.0205		0.0220	0.0168	0.0477***		0.0475***	0.0670***	-0.0157		-0.0172	-0.0177	-0.0172		-0.0158	-0.0112	0.0161		0.0166	0.0114
	(0.0226)		(0.0225)	(0.0235)	(0.0182)		(0.0182)	(0.0201)	(0.0229)		(0.0229)	(0.0252)	(0.0234)		(0.0232)	(0.0244)	(0.0185)		(0.0184)	(0.0196)
Incumbency × Chinese demand		-0.0249	-0.0276	-0.0642		0.0127	0.0103	0.145***		0.0184	0.0211	0.0174		-0.0254	-0.0237	0.00916		-0.0101	-0.0113	-0.0451
		(0.0338)	(0.0336)	(0.0723)		(0.0224)	(0.0230)	(0.0459)		(0.0339)	(0.0339)	(0.0672)		(0.0286)	(0.0283)	(0.0590)		(0.0227)	(0.0227)	(0.0481)
Incumbency × Both				0.0211				-0.0836***				0.00206				-0.0186				0.0203
				(0.0377)				(0.0317)				(0.0332)				(0.0360)				(0.0244)
Micro-region FE																				
Municipality FE	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X	X
Observations	15,201	15,201	15,201	15,201	14,260	14,260	14,260	14,260	14,651	14,651	14,651	14,651	14,843	14,843	14,843	14,843	17,173	17,173	17,173	17,173
R^2	0.086	0.086	0.086	0.086	0.153	0.152	0.153	0.154	0.074	0.074	0.075	0.075	0.065	0.065	0.065	0.066	0.173	0.173	0.173	0.173

Notes: Robust standard errors in parentheses. Unit of analysis is a mayoral candidate. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Our first results offer intriguing insights into the dynamics of incumbency effects amidst exogenous economic shocks. Initially, incumbents generally exhibit a substantial advantage over challengers, as anticipated. Surprisingly, exposure to regional economic shocks does not appear to directly impact the likelihood of victory for an average candidate.

However, it is crucial to consider an important caveat related to the statistical significance of the coefficient associated with the agricultural productivity shock (no interaction) before accounting for municipality fixed effects. In essence, this variable, due to its high spatial correlation with the significance of agricultural activity across regions, may primarily capture the inherent characteristics of municipalities throughout Brazil. This result aligns, to some extent, with the argument put forth by Novaes and Schiumerini (2022), who suggest that rural municipalities exhibit distinct voter performance evaluations compared to urban areas. This inherent characteristic, along with several other covariates related to the effect of the shocks in level, is adequately controlled with the inclusion of municipality-year fixed effects.

The most notable and promising preliminary finding concerns the influence of exogenous economic shocks on the incumbency effect. As the results are qualitatively (and quantitatively) similar in both specifications, we base our main preliminary findings on the preferred specification presented in Table 3.3.

Notably, in the initial election under analysis - in 2004 - regional economic shocks did not significantly impact the effect of incumbency in municipal elections across the country. This observation is expected since the shocks commenced in 2001 and 2003 and would reasonably have delayed impacts on economic activity. Nevertheless, our findings indicate an interesting pattern for the 2008 municipal elections. While we do not observe a significant effect of the trade shock on the incumbent's victory probability without the introduction of the complete interaction term, it is essential to emphasize that the variation in soybean production's potential yield, our metric for technical change in agriculture, significantly enhances the incumbent's advantage in mayoral elections in municipalities more exposed by the shock compared to less exposed localities, both statistically and economically.

To contextualize this effect, our results suggest that a one standard deviation positive shock in soy potential yield increased the probability of an incumbent mayor's reelection in 2008 by approximately 5% in columns (5) and (7). While these findings are preliminary, they contribute to the literature by identifying the amplification of the incumbency effect in local economies profoundly affected by positive exogenous economic shocks in a resource-rich country, where the agricultural sector plays a pivotal economic role.

We consider a specification that includes an interaction between the exposures to GE soy seed and Chinese demand in column (8). The addition of this interaction does not alter the sign of the coefficients related to GE soy technology although it increases its magnitude. However, we see a meaningful change in the coefficients associated with Chinese demand, with gains in terms of statistical and economic significance. Regarding the impact of the trade shock, the coefficients associated with it can be (approximately) interpreted as the effect of a US$1000 increase in exports per worker on changes in the dependent variable. Thus, an increase of US$1000 in exports per worker in a given municipality indicates a 14.5% increase in the probability of re-election of the incumbent in the mayoral elections. To provide greater context, an increase of one standard deviation in a municipality's exposure to the trade shock is associated with an approximate increase in reelection probability of 8.4% for the incumbent mayor.

Nevertheless, the point estimates of the full interaction terms suggest that areas extensively exposed to both trade and technology shocks experienced less amplification of the incumbency effect compared to regions solely affected by GE soy or the trade shock. This relative attenuation of the increase in the incumbency effect compared to regions more exposed to GE soy or the trade shock alone might be related to the information impact of these exogenous shocks on voters' behavior. To comprehensively understand this result, further exploration of the potential mechanisms behind the impacts of shocks on the probability of reelection of incumbents is necessary. This directly relates to the testing of our second theoretical prediction, which will be the focal point of our research in the forthcoming months.

Another important result found in our initial empirical exploration is that, although exposure to the exogenous regional economic shocks appears to affect the electoral fortunes of incumbents in local elections, this impact is transient. Note that, in the subsequent analyzed elections, from 2012 to 2020, we do not observe any statistical significance (at conventional levels) for the coefficients associated with the interaction terms of the economic shocks with the incumbency variable. Although a more detailed exploration needs to be done to better understand such dynamics, it is possible to argue that voters reward incumbents only in elections immediately associated (in temporal terms) with the peak of such regional economic shocks, and in subsequent electoral processes there is complete absorption of such exogenous effects on the evaluation of candidates in competition. This interesting pattern will also be the focus of analysis in the next steps of this research.

To assess the robustness of our preliminary results, we estimate some alternative specifications. Initially, we pooled all elections together and ran a simple two-way fixed effect model, controlling for region fixed-effects and year or election fixed effects. The results are presented in Table 3.4.

Table 3.4: Exogenous shocks and incumbency effects: two-way fixed effects with pooled sample

Dep. var.: *Probability of victory*								
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$Incumbency_{j,r,t-1}$	0.264***	0.288***	0.264***	0.256***	0.261***	0.284***	0.261***	0.254***
	(0.0128)	(0.00600)	(0.0127)	(0.0137)	(0.0133)	(0.00626)	(0.0133)	(0.0144)
ΔA_r^{soy}	-0.0167***		-0.0168***	-0.0215***				
	(0.00273)		(0.00274)	(0.00290)				
$\Delta \tilde{X}_r$		0.00109	0.00200	-0.0251***				
		(0.00346)	(0.00352)	(0.00699)				
$\Delta A_r^{soy} \times \Delta \tilde{X}_r$				0.0151***				
				(0.00302)				
Interactions								
Incumbency × GE soy	0.0127**		0.0138**	0.0180***	0.0117*		0.0127*	0.0167**
	(0.00634)		(0.00629)	(0.00691)	(0.00664)		(0.00659)	(0.00724)
Incumbency × Chinese demand		-0.00700	-0.00992	0.0206		-0.00689	-0.00976	0.0199
		(0.00835)	(0.00858)	(0.0160)		(0.00886)	(0.00909)	(0.0173)
Incumbency × Both				-0.0152**				-0.0142*
				(0.00755)				(0.00811)
Micro-region FE	X	X	X	X				
Municipality FE					X	X	X	X
Election (Year) FE	X	X	X	X	X	X	X	X
Observations	76,626	76,626	76,626	76,626	76,626	76,626	76,626	76,626
R^2	0.071	0.071	0.071	0.071	0.085	0.085	0.085	0.085

Notes: Standard errors (in parentheses) are adjusted for 556 micro-region clusters. Unit of analysis is a mayoral candidate. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

The results are, in general, qualitatively similar to those presented previously, indicating the robustness of our preliminary estimates. In short, we find that, in the aggregate or average of years, the level of exposure to the technological shock in agriculture positively impacted the electoral fortunes of incumbents in elections in these municipalities compared to regions less impacted by such a shock. However, an important difference concerns the non-significance of the trade shock in any of the specifications when we consider the pooled sample of elections.

Finally, we estimate the classic dynamic difference-in-difference model presented in Equation (3.14). Table 3.5 presents the estimation results of such dynamic specification.[36] Even with the caveats highlighted previously regarding such an empirical strategy in our case, the results are also qualitatively quite similar to those discussed above.

[36]In the Appendix C we present a visual representation of the coefficients of interest.

3.5.2 Amplification, muting and other mechanisms

As highlighted previously, our analysis has yet to empirically evaluate the second theoretical prediction of our model: whether the intensity of exogenous regional economic shocks amplifies or mutes the effect of incumbency on voters' decisions in local elections. This avenue of inquiry holds significant importance and will be a central focus in the forthcoming months of this research. It will serve as a fundamental approach to elucidate the mechanisms through which exogenous regional economic shocks influenced the electoral outcomes of incumbents in local elections in Brazil in the short run.

Table 3.5: Exogenous shocks and incumbency effects: difference-in-differences model

Dep. var.: *Probability of victory*	
Baseline - 2004 elections	Coefficients
Incumbency ×2008	0.0461
	(0.0489)
Incumbency ×2012	0.0101
	(0.0520)
Incumbency ×2016	-0.102*
	(0.0527)
Incumbency ×2016	0.165***
	(0.0483)
Inc × GE soy ×2008	0.0500**
	(0.0250)
Inc × GE soy ×2012	-0.0455*
	(0.0270)
Inc × GE soy ×2016	-0.0274
	(0.0273)
Inc × GE soy ×2020	-0.0128
	(0.0247)
Inc × China shock ×2008	0.193***
	(0.0738)
Inc × China shock ×2012	0.0608
	(0.0842)
Inc × China shock ×2016	0.0660
	(0.0834)
Inc × China shock ×2020	-0.0167
	(0.0768)
Inc × Both ×2008	-0.0910***
	(0.0339)
Inc × Both ×2012	0.00352
	(0.0362)
Inc × Both ×2016	-0.0343
	(0.0370)
Inc × Both ×2020	0.0193
	(0.0331)
Municipality FE	X
Election (Year) FE	X
Municipality × Year FE	X
Observations	76,128
R^2	0.114

Notes: Standard errors (in parentheses) are adjusted for 556 micro-region clusters. Unit of analysis is a mayoral candidate. $^{*}\ p < 0.1$, $^{**}\ p < 0.05$, $^{***}\ p < 0.01$

3.6 Discussion and Next Steps

In this paper, we examine the effects of exogenous economic shocks on the electoral fortunes of incumbents in local elections. First, we developed a theoretical framework that highlights how the interplay between the intensity of these exogenous economic shocks and the competence of politicians shapes electoral outcomes. Drawing upon existing literature, we scrutinize how varying degrees of exposure to these shocks can either magnify or attenuate the influence of incumbent competence on electoral fortunes, with a particular emphasis on the pivotal role of voter learning in assimilating additional information.

We evaluate one of the main predictions of the model estimating the effects of two prominent exogenous economic shocks – the introduction of new agricultural technology and increased trade demand – on local-level voting behavior in Brazilian mayoral elections spanning the period from 2000 to 2020. Methodologically, we assess local exposure to an agricultural technological shock, quantified by the productivity gains from the introduction of genetically engineered soy seeds in Brazil in 2003. Additionally, we analyze exposure to heightened export demand from China, employing a shift-share strategy that considers the growth of demand from China for specific products and the sectoral composition of employment across Brazilian municipalities. Through this rigorous analysis, we aim to disentangle the distinct impacts of local exposure to these economic shocks on incumbency effects, leveraging electoral data sourced from Brazil's Superior Electoral Court.

Our preliminary results regarding mayoral elections across Brazil from 2004 to 2020 reveal intriguing patterns regarding the impact of regional economic shocks on the incumbency effect. In the 2008 elections, while the trade shock alone does not significantly influence incumbents' victory probability (without the inclusion of the complete interaction term), the variation in soybean production's potential yield significantly enhances their advantage, indicating an amplification of the incumbency effect in municipalities more exposed to this shock. Specifically, a one standard deviation positive shock in soy potential yield increases the probability of an incumbent mayor's reelection by approximately 5%. This suggests that the primary driver of economic voting behavior in local elections in Brazil appears to be the productivity gains that altered the country's comparative advantage in global markets.

Additionally, the inclusion of complete interaction terms suggests that areas extensively exposed to both trade and technology shocks experienced less amplification of the incumbency effect compared to regions solely affected by individual shocks. However, the impact of these economic shocks appears transient, with subsequent elections from 2012 to 2020 showing no statistical significance in the interaction terms, implying a short-term effect of the exogenous economic shocks on incumbents' electoral fortunes. Further exploration is needed to understand the underlying mechanisms behind these dynamics, which will be the focus of future research.

By way of conclusion, it is worth considering potential paths for advancing the research initiated in this paper. Firstly, it is imperative to acknowledge the preliminary nature of our results, underscoring the importance of conducting robustness checks with varied incumbency measures throughout the years to enhance the reliability of our findings. Besides, a key focus of our research moving forward will be to empirically evaluate whether the intensity of exogenous regional economic shocks amplifies or mutes the effect of incumbency perceived competence on voters' decisions in local elections. This inquiry is essential for elucidating the mechanisms through which these shocks influence the electoral outcomes of incumbents in Brazil in the short run. By conducting further analyses and exploring additional data, we aim to provide deeper insights into the dynamics of incumbency effects amidst exogenous economic shocks, contributing to the existing literature on voter rationality and economic voting. Additionally, robustness checks through alternative specifications will help validate our preliminary findings and ensure the reliability of our conclusions.

Furthermore, one of our research project's primary objectives in the near future is to delve deeper into the mechanisms underlying the increased incumbency impact in regions more profoundly affected by the exogenous economic shocks in Brazil. For instance, we intend to explore whether incumbents perform better in government due to increased revenue and how the composition of government expenditures evolves following these exogenous shocks. We are particularly interested in investigating the impacts of these episodes on the provision of public goods, with a specific focus on areas such as health and education.

Importantly, these shocks, particularly the agricultural productivity, seem to have disproportionately benefited large rural producers in Brazil. Given the historical concentration of land in the country among the wealthiest segments of the population, it is likely that the economic benefits associated with such an exogenous shock would also have been concentrated among the upper echelons of the income and wealth distribution. Consequently, we aim to explore whether incumbents affiliated with certain political parties were more favored in local elections and what role electoral donations might have played in this dynamic. Lastly, it is also worth highlighting that the nature of these economic shocks may be related to a relative polarization of the political spectrum in the most affected regions compared to those least affected. These are just a few of the questions we aspire to address shortly as we continue to unravel the intricate relationship between economic shocks, incumbency effects, and local-level voting behavior in Brazil.

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A.1 Appendix A

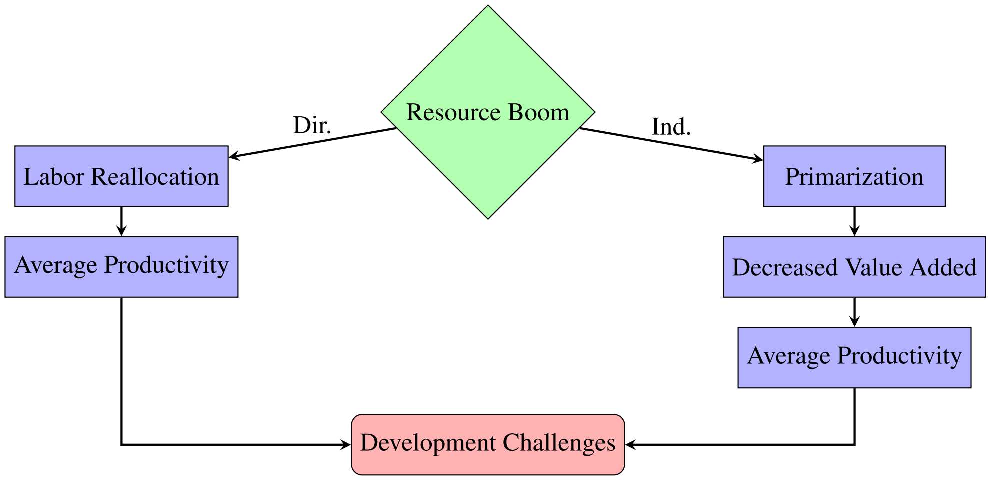
Table A.1: Descriptive statistics at the micro-region level - Long-differences (2000-2019)

Variables	*Source*	*Mean*	*Standard Deviation*	*Min*	*Max*	*Observations*
Chinese Demand Shock						
ΔX_r	COMEXSTAT	0.62	1.21	0.00	6.38	424
$\Delta \tilde{X}_r$ (instrument)	COMEXSTAT and BACI-CEPII	0.12	0.31	0.00	2.51	424
Regional Export Basket Value						
Δ Export Value	COMEXSTAT	333401105.96	1247641603.99	-1383672724.60	20543485713.80	424
Δ Log Export Value	COMEXSTAT	1.59	2.27	-6.16	12.53	424
Regional Export Basket Concentration						
Δ Lines	COMEXSTAT	47.77	83.84	-178.00	558.00	424
Δ HHI	COMEXSTAT	-0.04	0.33	-1.00	0.98	424
Regional Export Basket Sophistication						
Δ $S_{r,t}$ (WDI)	WDI, BACI-CEPII, and COMEXSTAT	225.36	1757.51	-4735.18	24695.02	424
Δ $S_{r,t}$ (PWT)	PWT, BACI-CEPII, and COMEXSTAT	287.96	2924.96	-8998.35	44524.66	424
Regional Export Basket Composition						
Δ Share of Resources (Prim. + RB)	COMEXSTAT and Lall (2000)	0.07	0.32	-1.00	1.00	424
Δ Share of Primary Products	COMEXSTAT and Lall (2000)	0.13	0.39	-1.00	1.00	424
Δ Share of Resource-Based Man.	COMEXSTAT and Lall (2000)	-0.06	0.40	-1.00	1.00	424
Δ Share of Low-Tech Man.	COMEXSTAT and Lall (2000)	-0.08	0.25	-1.00	1.00	424
Δ Share of Medium-Tech Man.	COMEXSTAT and Lall (2000)	0.01	0.24	-1.00	1.00	424
Δ Share of High-Tech Man.	COMEXSTAT and Lall (2000)	-0.01	0.11	-1.00	0.54	424

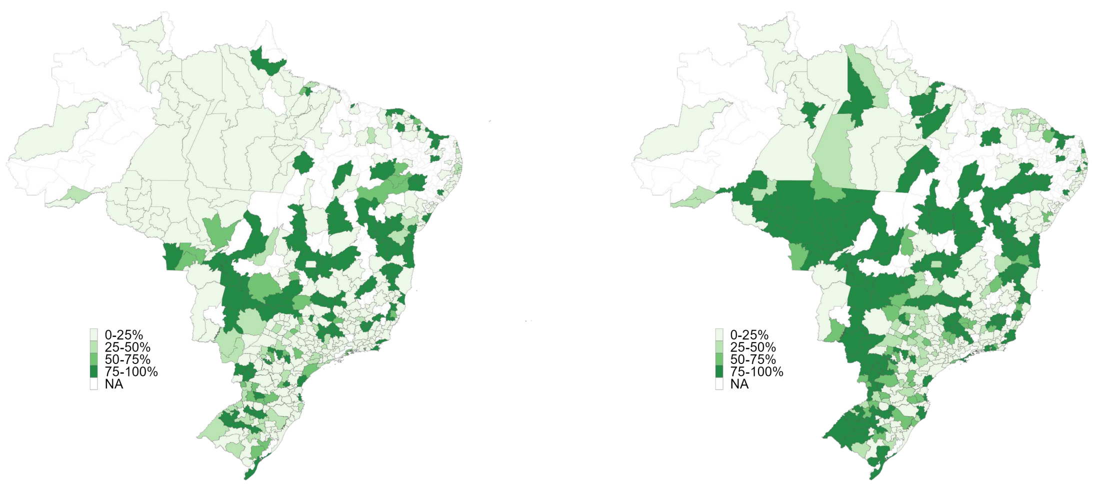
Table A.2: Descriptive statistics - Yearly dataset

Variables		*Mean*	*Standard Deviation*	*Min*	*Max*	*Observations*
Regional Export Basket Value						
Export Value	overall	392564333.77	1221829523.673	1.017	22706823321	N = 9752
	between		1055119999.552	1026.443	13700143738.131	n = 424
	within		577278271.726	-6180431422.417	16072092164.325	T = 23
Log Export Value	overall	17.383	2.794	0.016	23.846	N = 9752
	between		2.714	6.093	23.294	n = 424
	within		1.305	2.676	22.666	T = 23
Regional Export Basket Concentration						
Number of Export Lines	overall	69.482	136.295	1	1079	N = 9752
	between		129.705	1	1011.565	n = 424
	within		34.059	-293.736	502.308	T = 23
HHI	overall	0.457	0.299	0	1	N = 9752
	between		0.237	0.04	1	n = 424
	within		0.194	-0.437	1.22	T = 23
Regional Export Basket Sophistication						
$S_{r,t}$ (WDI)	overall	2551.369	1856.562	444.085	49477.073	N = 9752
	between		1551.084	719.364	21413.945	n = 424
	within		1060.919	-15435.716	30614.497	T = 23
$S_{r,t}$ (PWT)	overall	3564.521	3073.459	676.219	87262.215	N = 9752
	between		2448.242	1080.277	37069.303	n = 424
	within		1875.831	-28969.087	59412.882	T = 23

Figure A.1: Potential mechanism — Direct and indirect effects



A.1.1 Composition of regional export baskets



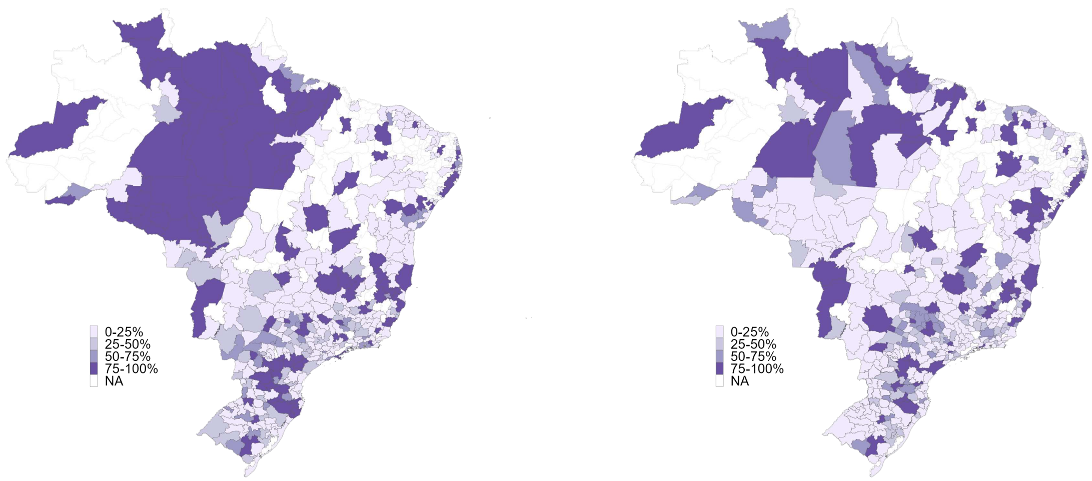
(a) 2000

(b) 2019

Source: Data on the value of exports is based on the declaration of exporters in Brazil (SISCOMEX from the Ministry of Industry, Foreign Trade and Services).

Figure A.2: Share of Exports — Primary Products

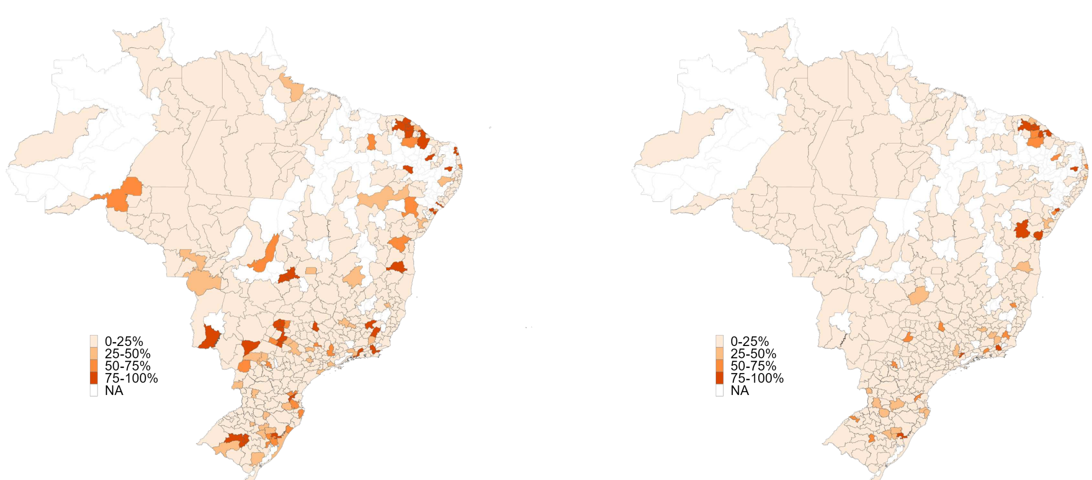
Figure A.3: Share of Exports — Resource Based Products



(a) 2000 (b) 2019

Source: Data on the value of exports is based on the declaration of exporters in Brazil (SISCOMEX from the Ministry of Industry, Foreign Trade and Services).

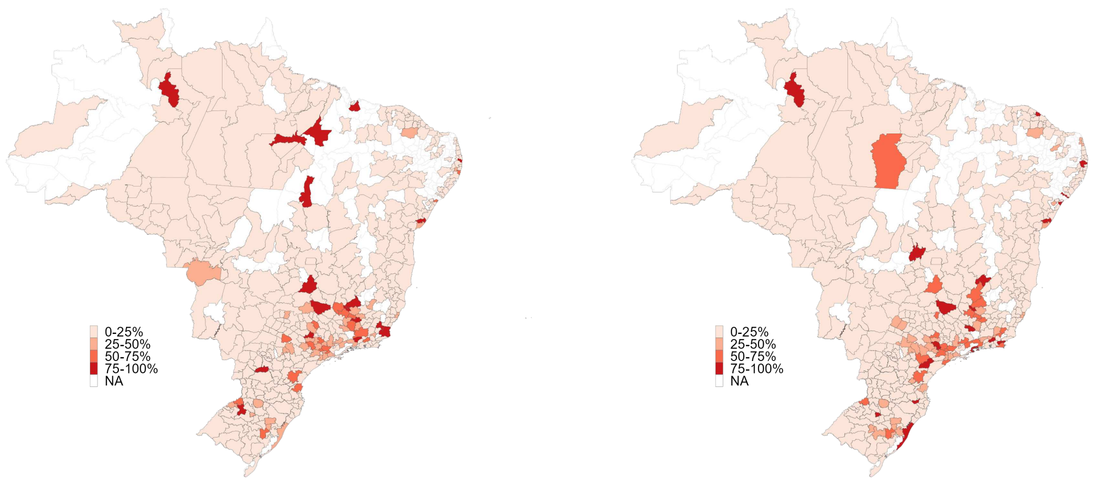
Figure A.4: Share of Exports — Resource Based Products



(a) 2000 (b) 2019

Source: Data on the value of exports is based on the declaration of exporters in Brazil (SISCOMEX from the Ministry of Industry, Foreign Trade and Services).

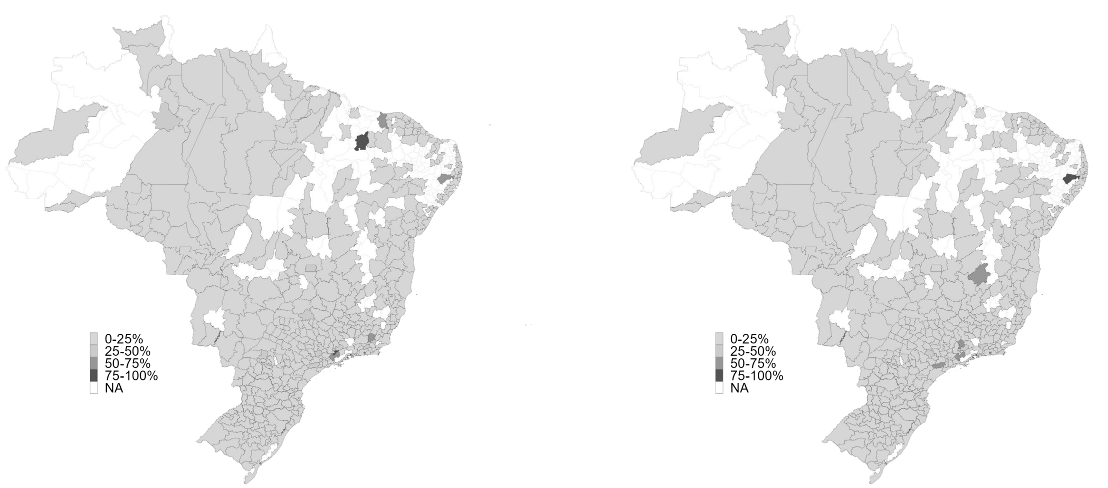
Figure A.5: Share of Exports — Low-Tech Manufactures



(a) 2000 (b) 2019

Source: Data on the value of exports is based on the declaration of exporters in Brazil (SISCOMEX from the Ministry of Industry, Foreign Trade and Services).

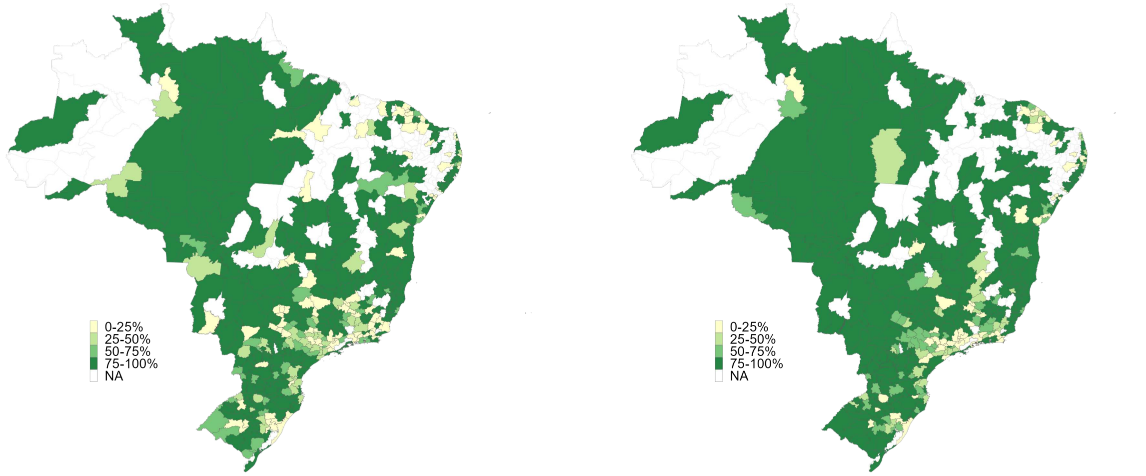
Figure A.6: Share of Exports — Medium-Tech Manufactures



(a) 2000 (b) 2019

Source: Data on the value of exports is based on the declaration of exporters in Brazil (SISCOMEX from the Ministry of Industry, Foreign Trade and Services).

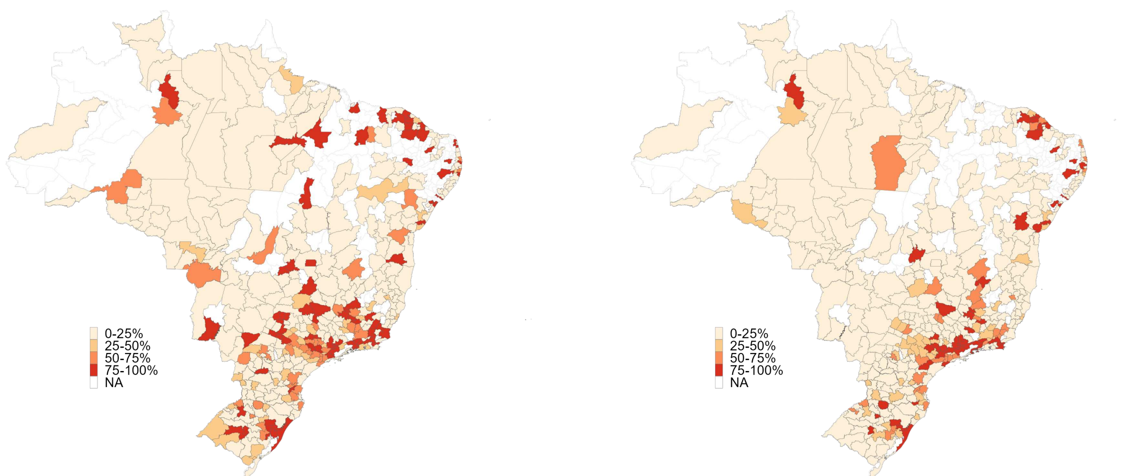
Figure A.7: Share of Exports — High-Tech Manufactures



(a) 2000 (b) 2019

Source: Data on the value of exports is based on the declaration of exporters in Brazil (SISCOMEX from the Ministry of Industry, Foreign Trade and Services).

Figure A.8: Share of Exports — Resources



(a) 2000 (b) 2019

Source: Data on the value of exports is based on the declaration of exporters in Brazil (SISCOMEX from the Ministry of Industry, Foreign Trade and Services).

Figure A.9: Share of Exports — Manufactures

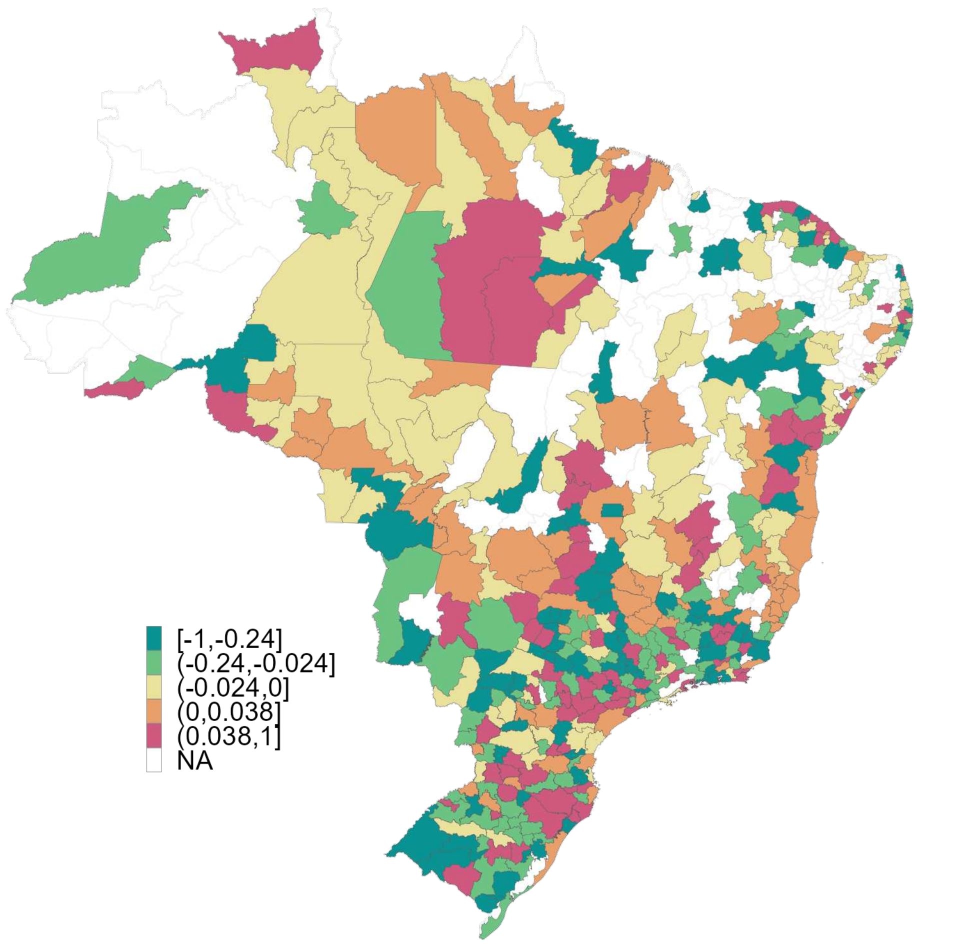


Figure A.10: Difference in the share of manufactures exports: 2019 - 2000

Table A.3: First Stage Regressions - Chinese Export Demand and Shift-Share Instrument

Dependent variable:	Chinese Export Demand ΔXX_r HS4 - 1217 categories 2019 - 2000	Chinese Export Demand ΔXX_r HS2 - 92 categories 2019 - 2000	Chinese Export Demand ΔXX_r HS2 - 92 categories 2010 - 2000
$\Delta \tilde{XX}_r$	1.763***	1.550***	2.282***
	(0.299)	(0.170)	(0.077)
Constant	0.570***	0.419***	0.085***
	(0.120)	(0.056)	(0.017)
Observations	439	439	439
Adjusted R^2	0.074	0.160	0.666
KP F-Stat	34.873	83.185	874.300

Notes: Robust standard errors in parentheses. Unit of analysis is a micro-region r. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

A.1.2 First-stage regression and additional visual evidence

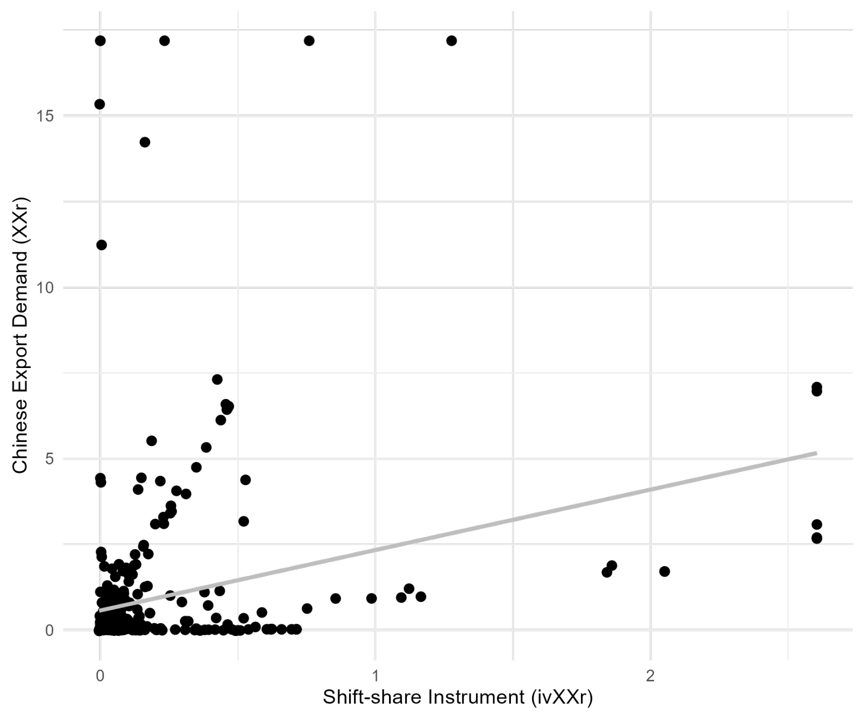


Figure A.11: Correlation: Chinese export demand and shift-share instrumental variable - HS4

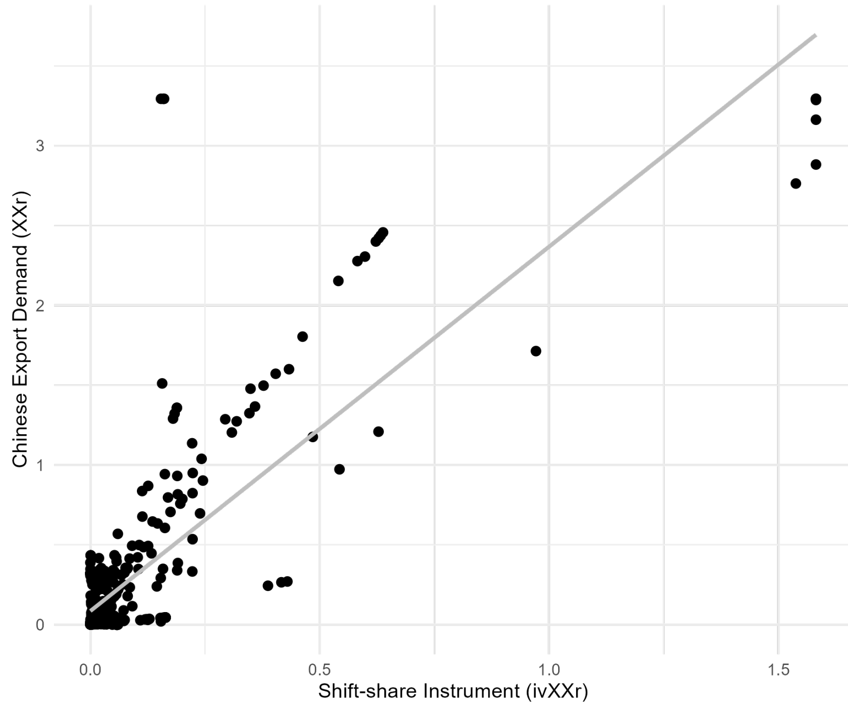


Figure A.12: Correlation: Chinese export demand and shift-share instrumental variable - HS2 and 2010-2000 (Census variables)

A.1.3 Robustness to alternative inference procedures

In this section, I show that the baseline results are very similar using the inference procedures proposed by Borusyak et al. (2022), which address cross-region correlation in residuals in shift-share designs. Tables A.4, A.5, and A.6 indicates that the baseline results for export basket value, concentration and sophistication, presented in Tables 1,2, and 6 of the manuscript, are not qualitatively altered when following alternative inference procedures.

A.1.4 Robustness to different temporalities and parallel trends

Finally, I present additional estimations aimed at assessing the robustness of the primary findings. Tables A.7, A.8, A.9, and A.10 delve into the stability of the coefficients associated with the variation in the total value of exports, the growth rate of such values, the number of export lines, and the HHI over different periods, respectively. Particularly, I scrutinize a shorter-term impact by estimating the disparities in the variables of interest between 2000 and 2010, juxtaposed with

Table A.4: Commodity boom and export value (Borusyak et al. (2022) robust standard errors)

Dependent variable:	Δ **Value of exports**			$\%\Delta$ **Value of exports**		
	OLS	OLS	2SLS	OLS	OLS	2SLS
	(1)	(2)	(3)	(1)	(2)	(3)
$\Delta \tilde{X}_r$	3,312,141,960*	2,902,794,084***		1.730**	1.486***	
	[801,244,029]	[710,144,339]		[0.362]	[0.256]	
ΔX_r			2,537,285,580***			1.299**
			[1,018,991,070]			[0.767]
Weighted	X	X	X	X	X	X
State-year fixed effects		X	X		X	X

Notes: This table presents an alternative approach to inference on the baseline results in Table 1 of the manuscript. There are 96 industry observations in each regression (industry-level regressions). Borusyak et al. (2022) robust standard errors are reported in brackets. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A.5: Commodity boom and export concentration (Borusyak et al. (2022) robust standard errors)

Dependent variable:	Δ **Lines**			Δ **HHI**		
	OLS	OLS	2SLS	OLS	OLS	2SLS
	(1)	(2)	(3)	(1)	(2)	(3)
$\Delta \tilde{X}_r$	3.658	-16.494		0.085***	0.083***	
	[40.527]	[58.779]		[0.026]	[0.025]	
ΔX_r			-14.417			0.073**
			[80.825]			[0.031]
Weighted	X	X	X	X	X	X
State-year fixed effects		X	X		X	X

Notes: This table presents an alternative approach to inference on the baseline results in Table 2 of the manuscript. There are 96 industry observations in each regression (industry-level regressions). Borusyak et al. (2022) robust standard errors are reported in brackets. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

the coefficients already presented for the long differences between 2000 and 2019. Importantly, I use the variation of the shift-share instrument with the final year of 2010 in the shorter period window. Across virtually all cases, the results exhibit qualitative similarity in both temporalities, underscoring the robustness of the findings to different temporal specifications.

Additionally, I incorporate placebo tests or parallel trend assessments in Column (4) of all tables. Here, I regress the shock exposure variable on a preceding difference in the dependent variable (between 1997 and 2000 — before the occurrence of the resource boom), with observations

Table A.6: Commodity boom and export sophistication (Borusyak et al. (2022) robust standard errors)

Dependent variable:	Δ $S_{r,t}$ **(WDI)**			Δ $S_{r,t}$ **(PWT)**		
	OLS	OLS	2SLS	OLS	OLS	2SLS
	(1)	(2)	(3)	(1)	(2)	(3)
$\Delta\tilde{X}_r$	35.52	-327.91***		6.7138	-371.52***	
	[119.723]	[109.899]		[130.047]	[131.002]	
ΔX_r			-286.62**			-324.74**
			[122.296]			[149.761]
Weighted	X	X	X	X	X	X
State-year fixed effects		X	X		X	X

Notes: This table presents an alternative approach to inference on the baseline results in Table 6 of the manuscript. There are 96 industry observations in each regression (industry-level regressions). Borusyak et al. (2022) robust standard errors are reported in brackets. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

weighted and state-year fixed effects considered. In Tables A.7 and A.8, the placebo test reveals a contrasting trend prior to treatment. In other words, there seems to be a reversal of the trend with the treatment - the regions most affected subsequently by the Chinese demand shock displayed, before the treatment, negative changes in the value of total exports. However, concerning HHI (Table A.10), the placebo test suggests that the hypothesis of parallel trends before treatment appears to be observed.

Table A.7: Commodity boom and change in export value: short and medium run

Dependent variable: Δ **Value of exports**								
	OLS		OLS		2SLS		Placebo	
	(1)		(2)		(3)		(4)	
	2000-2010	2000-2019	2000-2010	2000-2019	2000-2010	2000-2019	1997-2000	1997-2000
$\Delta\tilde{X}_r$	2,423,790,438***	3,312,141,960**	3,342,149,578***	2,902,794,084***			-409,607,604***	-246,990,431***
	(756,532,864)	(1,628,508,368)	(593,872,297)	(1,081,950,192)			(121,751,038)	(76,174,580)
ΔX_r					1,759,577,719***	2,537,285,580***		
					(335,237,141)	(886,315,027)		
Weighted	X	X	X	X	X	X	X	X
State-year fixed effects			X	X	X	X	X	X
Observations	417	424	417	424	417	424	406	406
Adjusted R^2	0.116	0.182	0.305	0.625	0.286	0.296	0.314	0.309
KP F-stat					952.8	99.7		

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. Column 4 presents a placebo test, with observations weighted by population and considering state fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A.8: Commodity boom and percent change in export value: short and medium run

Dependent variable: **% Δ Value**								
	OLS		OLS		2SLS		Placebo	
	(1)		(2)		(3)		(4)	
	2000-2010	2000-2019	2000-2010	2000-2019	2000-2010	2000-2019	1997-2000	1997-2000
$\Delta\tilde{X}_r$	3.156***	1.730	2.467***	1.486*			-0.308***	-0.182***
	(0.755)	(1.169)	(0.911)	(0.795)			(0.085)	(0.054)
ΔX_r					1.299***	1.299*		
					(0.493)	(0.669)		
Weighted	X	X	X	X	X	X	X	X
State-year fixed effects			X	X	X	X	X	X
Observations	417	424	417	424	417	424	406	406
Adjusted R^2	-0.002	0.002	0.068	0.066	0.067	0.066	-0.017	-0.018
KP F-stat					952.8	99.7		

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. Column 4 presents a placebo test, with observations weighted by population and considering state fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A.9: Commodity boom and number of exports lines: short and medium run

Dependent variable: **Δ Lines**								
	OLS		OLS		2SLS		Placebo	
	(1)		(2)		(3)		(4)	
	2000-2010	2000-2019	2000-2010	2000-2019	2000-2010	2000-2019	1997-2000	1997-2000
$\Delta\tilde{X}_r$	2.097	3.658	-4.929	-16.494			6.901	6.198
	(21.455)	(30.516)	(21.539)	(29.669)			(10.319)	(6.902)
ΔX_r					-2.595	-14.417		
					(11.293)	(25.858)		
Weighted	X	X	X	X	X	X	X	X
State-year fixed effects			X	X	X	X	X	X
Observations	417	424	417	424	417	424	406	406
Adjusted R^2	-0.001	-0.002	0.01412	0.191	0.053	0.117	0.297	0.301
KP F-stat					952.8	99.7		

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. Column 4 presents a placebo test, with observations weighted by population and considering state fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

A.1.5 Robustness to alternative spatial aggregations

First, Figure A.13 maps the local exposure to Chinese export demand — the shift-share instrument — computed at the state level. The similarities with the regional exposure shown in Figure 1.8 are quite clear.

Additionally, I present the event-study specification outlined in Equation (1.8) for state-level observations. Figures A.14, A.15, and A.16 depict the dynamic effects of the China-induced de-

Table A.10: Commodity boom and HHI: short and medium run

Dependent variable: Δ **HHI**	OLS		OLS		2SLS		Placebo	
	(1)		(2)		(3)		(4)	
	2000-2010	2000-2019	2000-2010	2000-2019	2000-2010	2000-2019	1997-2000	1997-2000
$\Delta\tilde{X}_r$	0.085***	0.085	0.079***	0.083***			-0.001	-0.003
	(0.025)	(0.052)	(0.026)	(0.029)			(0.029)	(0.018)
ΔX_r					0.042***	0.073***		
					(0.015)	(0.026)		
Weighted	X	X	X	X	X	X	X	X
State-year fixed effects			X	X	X	X	X	X
Observations	417	424	417	424	417	424	406	406
Adjusted R^2	0.041	0.053	0.165	0.324	0.147	0.249	0.057	0.057
KP F-stat					952.8	99.7		

Notes: Unit of analysis r is a micro-region. Standard errors (in parentheses) are adjusted for 129 meso-region clusters. In column 1, observations are weighted by the total exports in 2000; column 2 adds state-year fixed effects to column 1; column 3 presents the 2SLS using $\Delta\tilde{X}_r$ as the IV for ΔX_r. Column 4 presents a placebo test, with observations weighted by population and considering state fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

mand shock on the value, concentration, and sophistication of state-level export baskets, respectively, along with 95% confidence intervals.

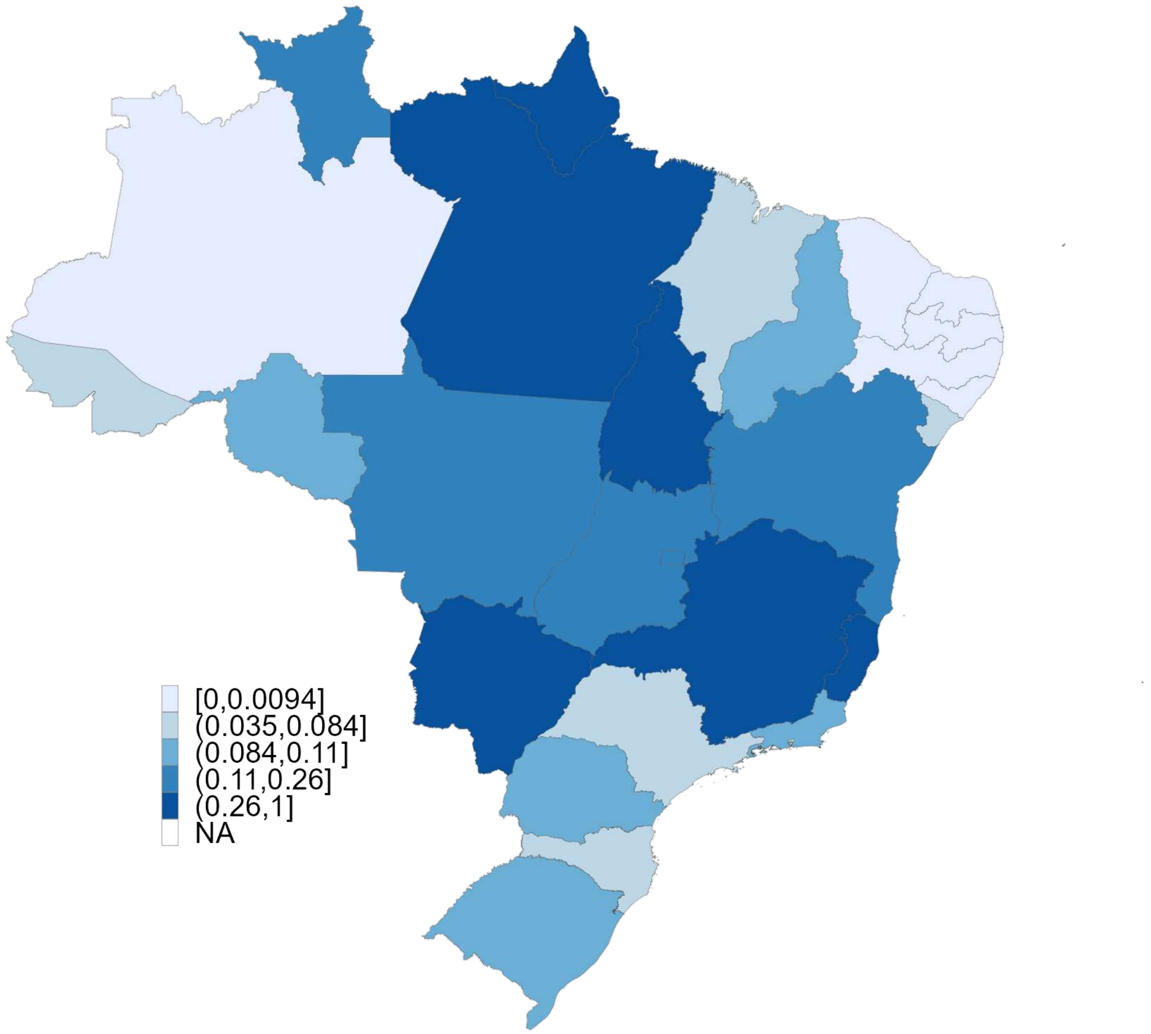
Overall, the results indicate the robustness of the causal findings to alternative spatial aggregations, which, unlike the municipality-level data, have export data computed at the production locality.

A.1.6 Sensitivity Analysis

To further assess the robustness of the baseline results presented in Tables 1, 2, and 6 of the manuscript, I conducted a sensitivity analysis focusing on the coefficients associated with the impacts of regional exposure to Chinese demand on key dependent variables.

For this analysis, I re-estimated the preferred specification — using weighted observations and including state-year fixed effects — by sequentially excluding one of the top or bottom 40 micro-regions based on: i) the magnitude of the China-induced export demand shock, and ii) the magnitude of the dependent variables of interest.

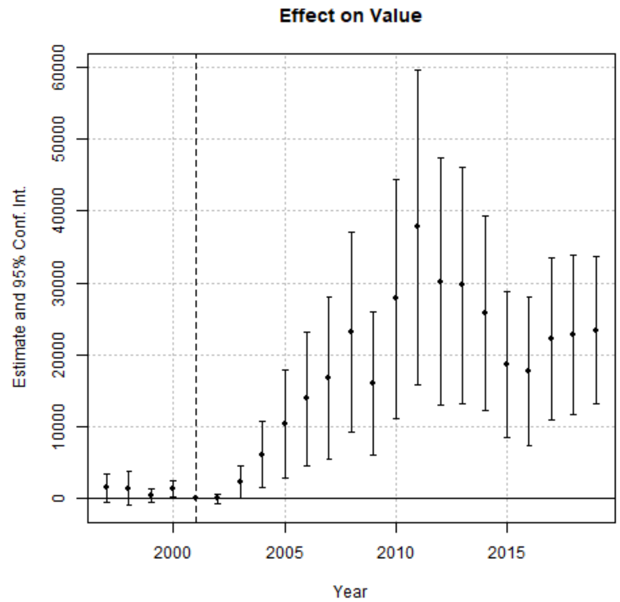
Figure A.13: Exposure to China's Export Demand - $\Delta\tilde{X}_r$ - State Level



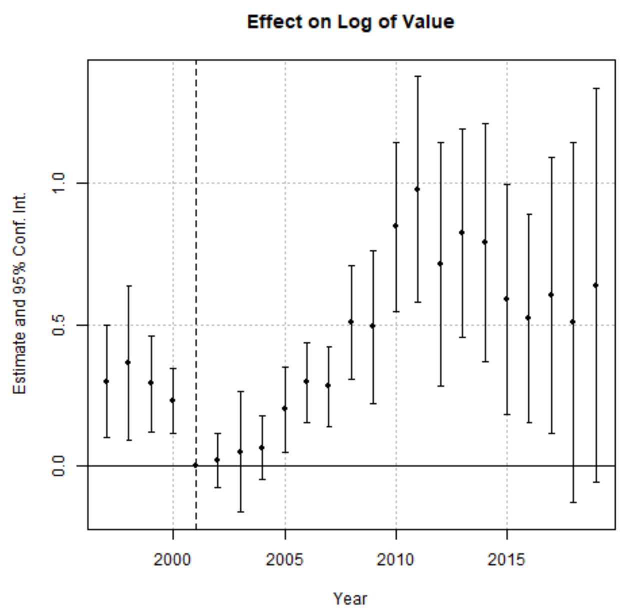
Source: Regional exposures to China's export demand, $\Delta\tilde{X}_r$, are computed according to Equation (1.6) for the state level. Data from CEPII-BACI and SISCOMEX are used for computing the shift-share instrument.

Exclusion of Observations Ranked by Magnitude of Shift-Share Instrument

First, Figures A.17 and A.18 graphically present, respectively, the results of excluding the top and bottom-ranked micro-regions in terms of $\Delta\tilde{X}r$, highlighting the coefficient of interest and the 95% confidence interval for each regression where the dependent variable is either the difference in regional export value or the growth rate of export value. The analysis begins with the highest or



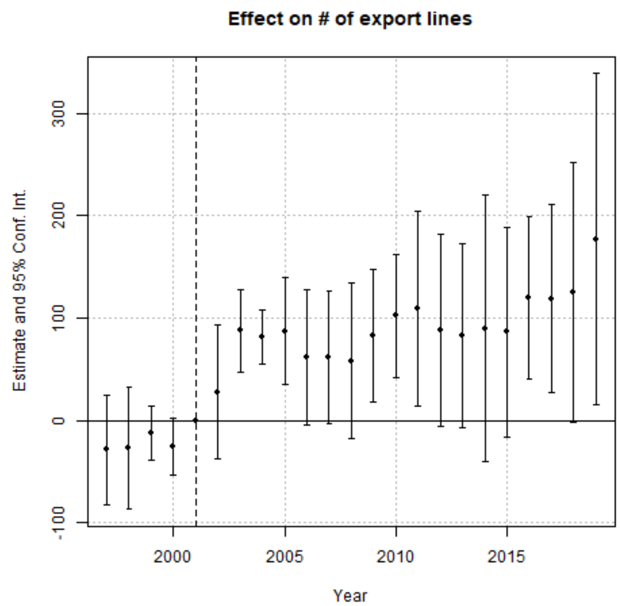
(a) Dynamic effects on export value



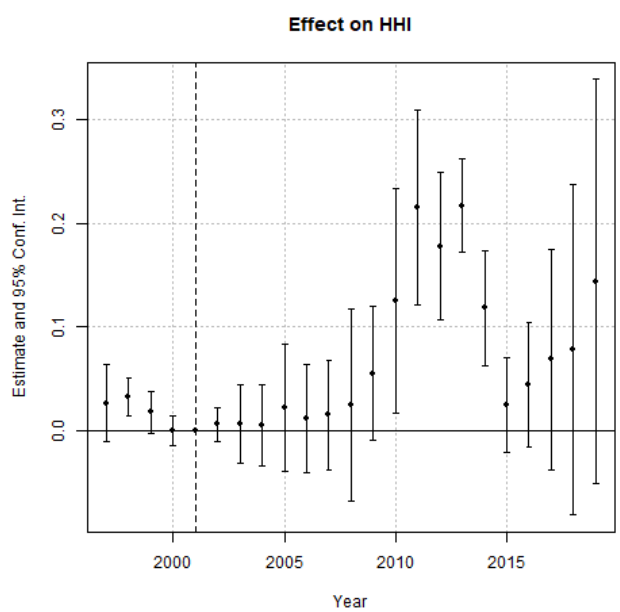
(b) Dynamic effects on log export value

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (1.8), where the dependent variables are the state export value in level and log, respectively, in year $t = 1997, ..., 2019$. The regressions include state and year fixed effects. Standard errors are adjusted for 27 state clusters and the observations are weighted by total exports in 2000.

Figure A.14: Dynamic Effects of the Resource Boom on Export Value and Growth - State Level



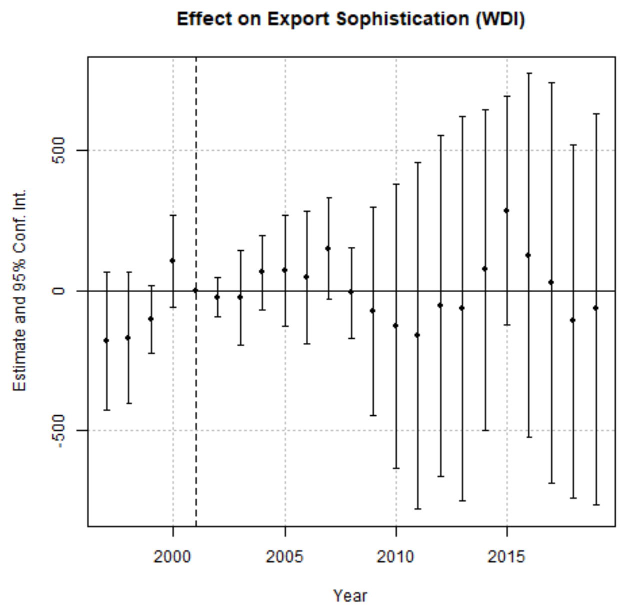
(a) Dynamic effects on export lines



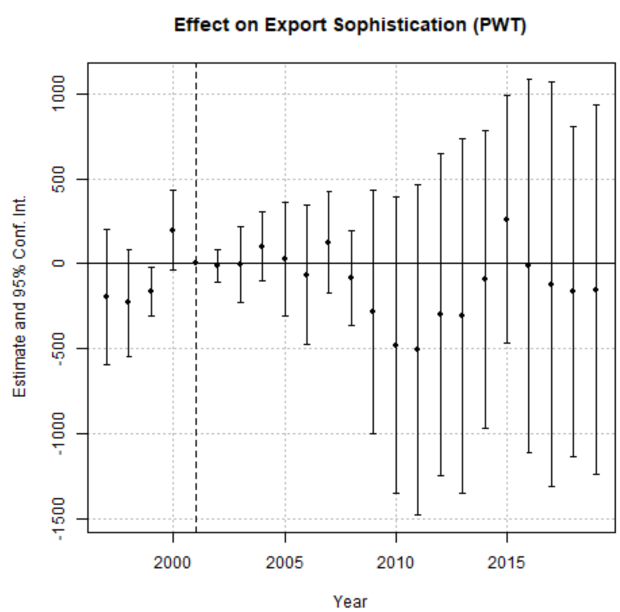
(b) Dynamic effects on HHI

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (1.8), where the dependent variables are the number of export lines and the HHI associated with state export baskets, respectively, in year $t = 1997, ..., 2019$. The regressions include state and year fixed effects. Standard errors are adjusted for 27 state clusters and the observations are weighted by total exports in 2000.

Figure A.15: Dynamic Effects of the Resource Boom on Export Concentration - State Level



(a) Dynamic effects on sophistication (WDI)



(b) Dynamic effects on sophistication (PWT)

Notes: Each point reflects an individual regression coefficient $\hat{\beta}$ following Equation (1.8), where the dependent variables are the sophistication indexes associated with state export baskets as described in Equation (1.4) in year $t = 1997, ..., 2019$. The regressions include state and year fixed effects. Standard errors are adjusted for 27 state clusters and the observations are weighted by total exports in 2000.

Figure A.16: Dynamic Effects of the Resource Boom on Export Sophistication - State Level lowest $\Delta\tilde{X}r$ on the left. The dotted black line represents the coefficient obtained in the baseline regression.

Similarly, Figures A.19 and A.20, respectively, present the results of excluding the top and bottom-ranked micro-regions based on $\Delta\tilde{X}_r$, with the difference in export lines and HHI as the dependent variables.

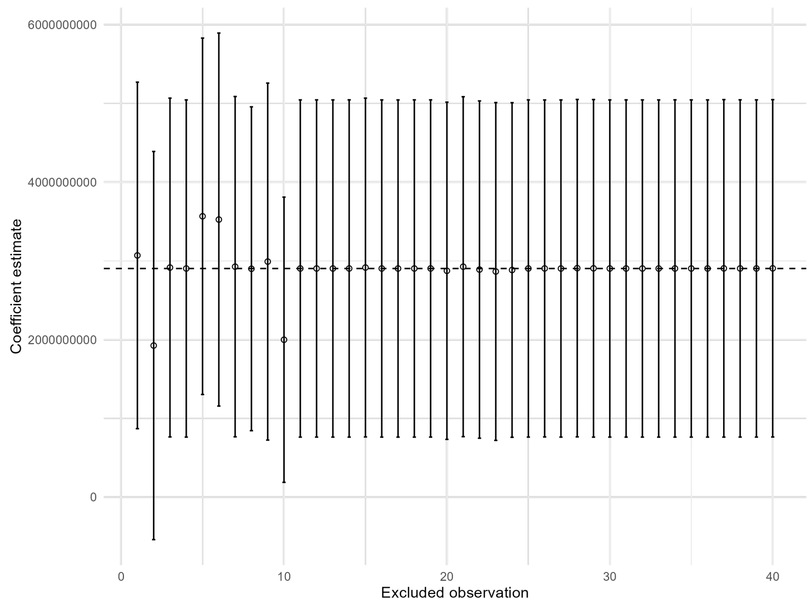
Lastly, I perform a similar sensitivity analysis for export sophistication measures. Figures A.21 and A.22 present the results of excluding the top and bottom-ranked micro-regions based on $\Delta\tilde{X}_r$, with the difference in export sophistication as defined in Equation (1.4) for both the WDI and PWT data sources.

Taken together, the evidence from this subsection further support the causal relations presented in the main results of the manuscript. The results show that outliers of the shift-share instrument indeed do not drive the main results of the manuscript, in addition to the winsorizing process already described earlier.

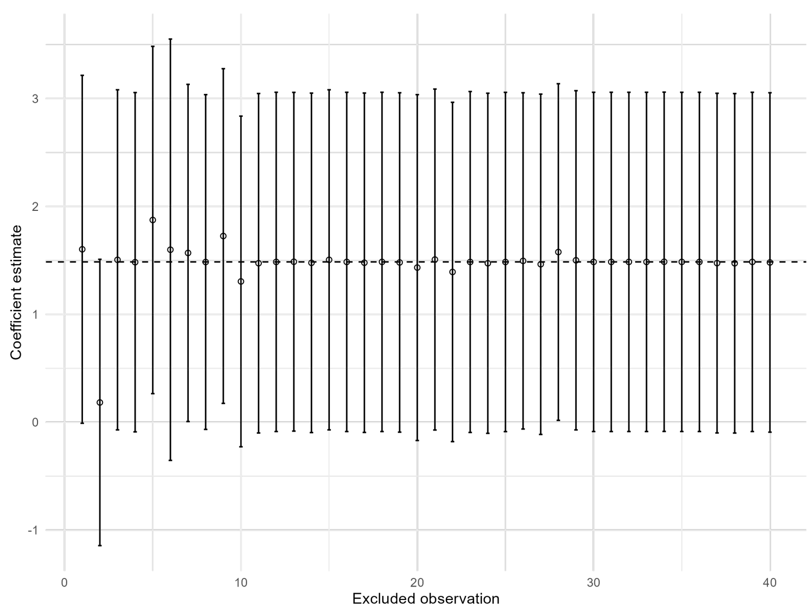
Exclusion of Observations Ranked by Value of Dependent Variables

Next, I conduct a similar sensitivity analysis by excluding observations with the highest and lowest magnitudes of the dependent variables of interest. Figures A.23, A.24, A.25, A.26, A.27, and A.28 present the estimated coefficients and 95% confidence intervals for the difference in regional export basket value, growth rate of the value, number of export lines, HHI, and export sophistication using both WDI and PWT data.

Overall, the results from this subsection indicate the stability of the estimated coefficients to the exclusion of observations at the top and bottom of the distribution of the magnitude of the dependent variables of interest.



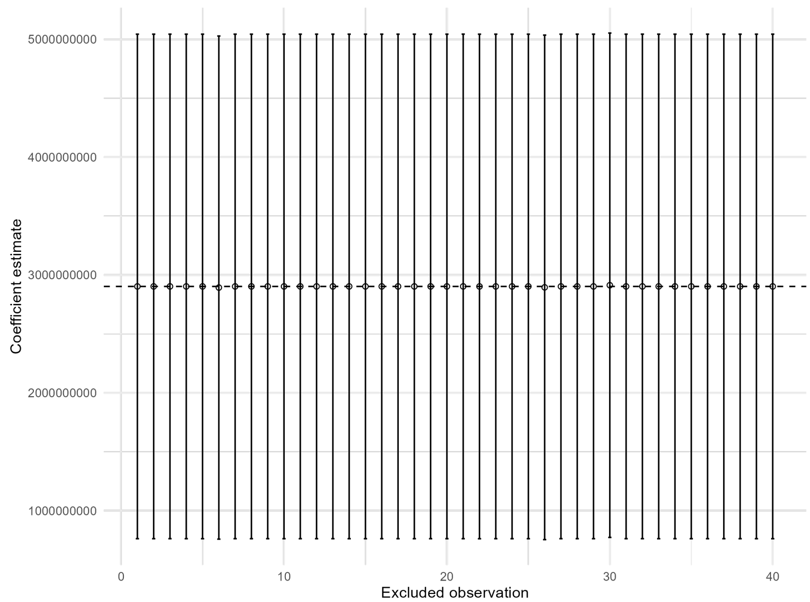
(a) Δ Value of Exports



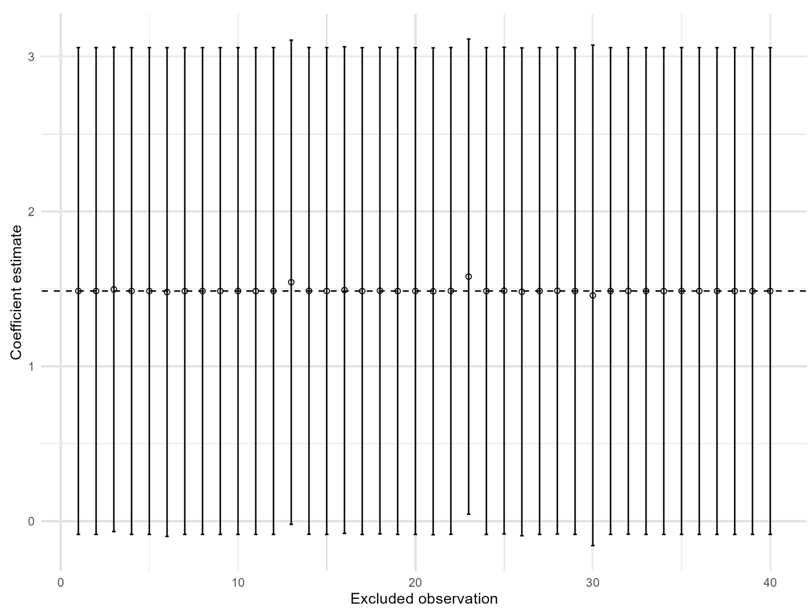
(b) $\%\Delta$ Value of Exports

Figure A.17: Commodity Boom and Export Value (Exclusion of Micro-Regions - Top 40 Based on $\Delta\tilde{X}r$ Value)

Notes: The figure shows the robustness of the results to excluding, one by one, the top 40 micro-regions with the highest $\Delta\tilde{X}r$. The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—highest $\Delta\tilde{X}_r$ exposure is the first observation.



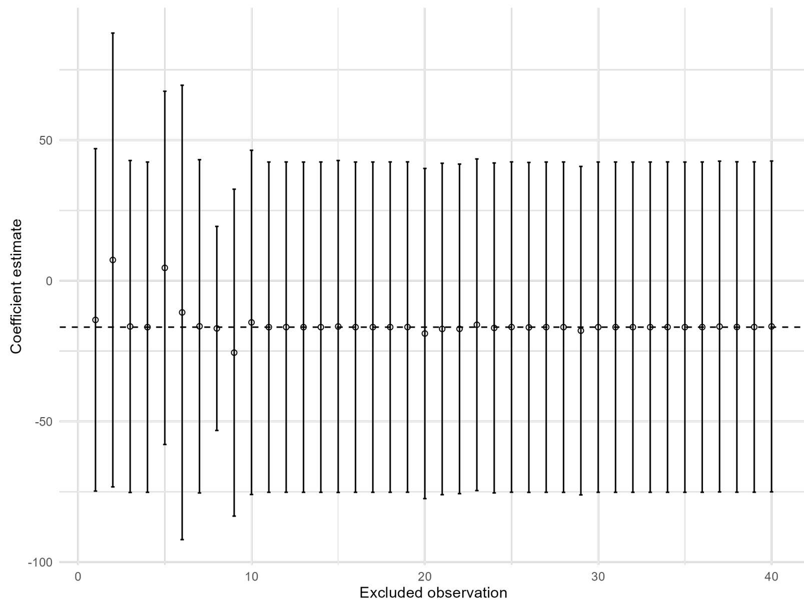
(a) Δ Value of Exports



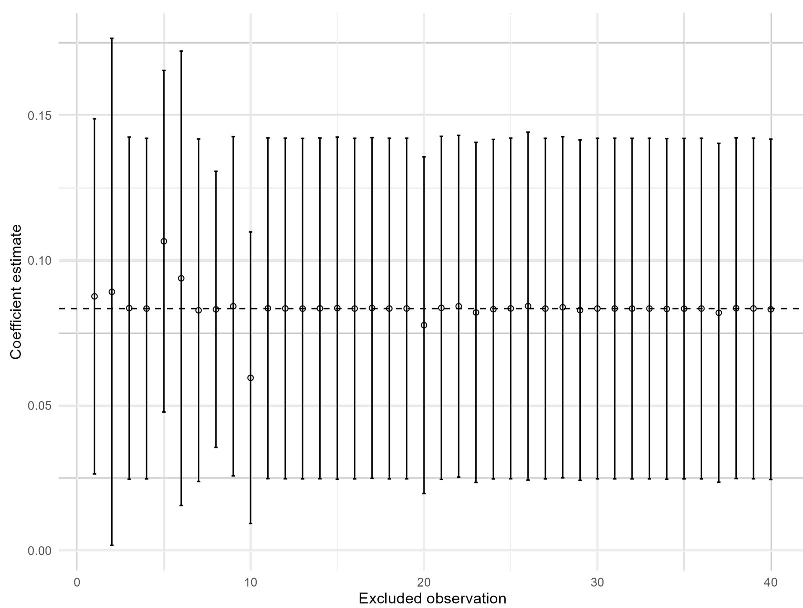
(b) $\%\Delta$ Value of Exports

Figure A.18: Commodity Boom and Export Value (Exclusion of Micro-Regions - Low 40 Based on $\Delta\tilde{X}r$ Value)

Notes: The figure shows the robustness of the results to excluding, one by one, the last 40 micro-regions with the lowest $\Delta\tilde{X}r$. The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—lowest $\Delta\tilde{X}_r$ exposure is the first observation.



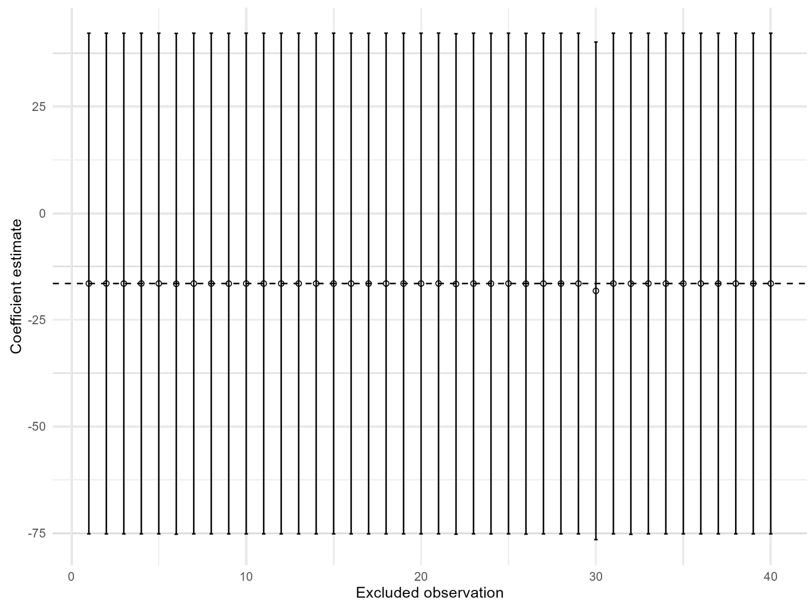
(a) Δ Lines



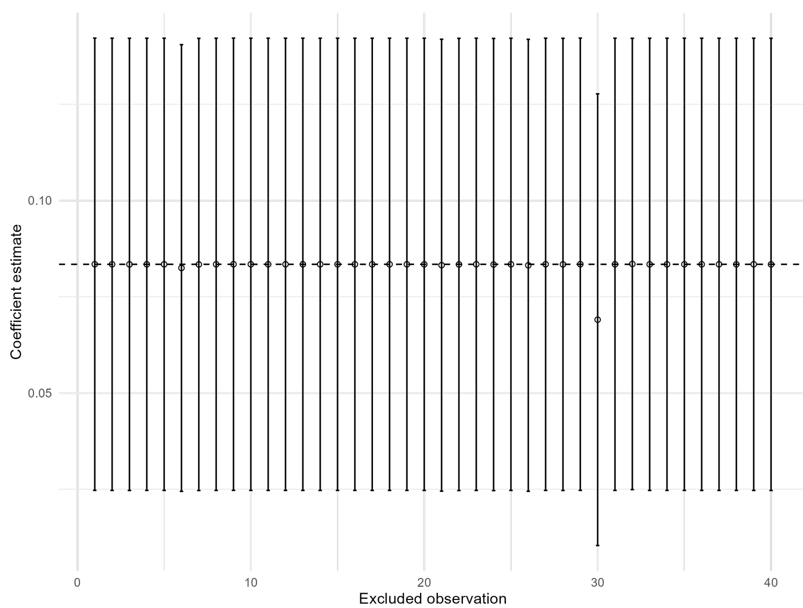
(b) Δ HHI

Figure A.19: Commodity Boom and Export Concentration (Exclusion of Micro-Regions - Top 40 Based on $\Delta\tilde{X}r$ Value)

Notes: The figure shows the robustness of the results to excluding, one by one, the top 40 micro-regions with the highest $\Delta\tilde{X}r$. The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—highest $\Delta\tilde{X}_r$ exposure is the first observation.



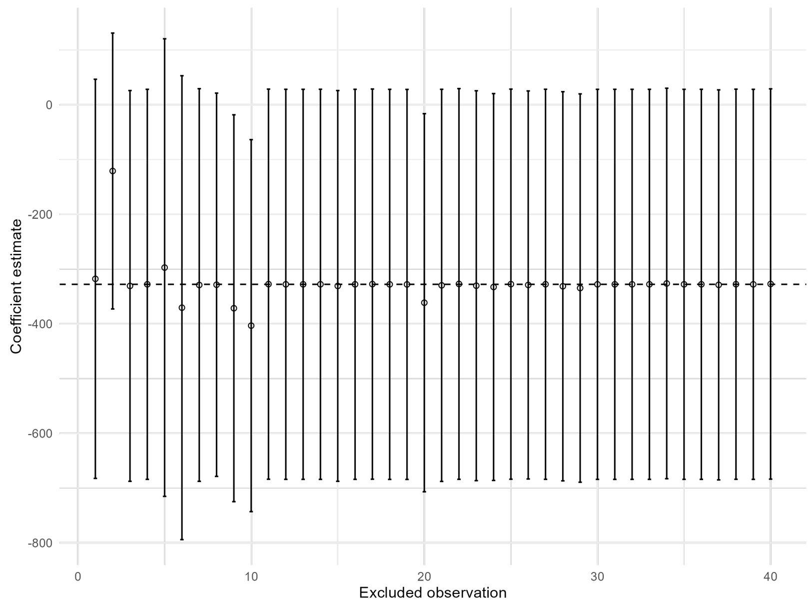
(a) Δ Lines



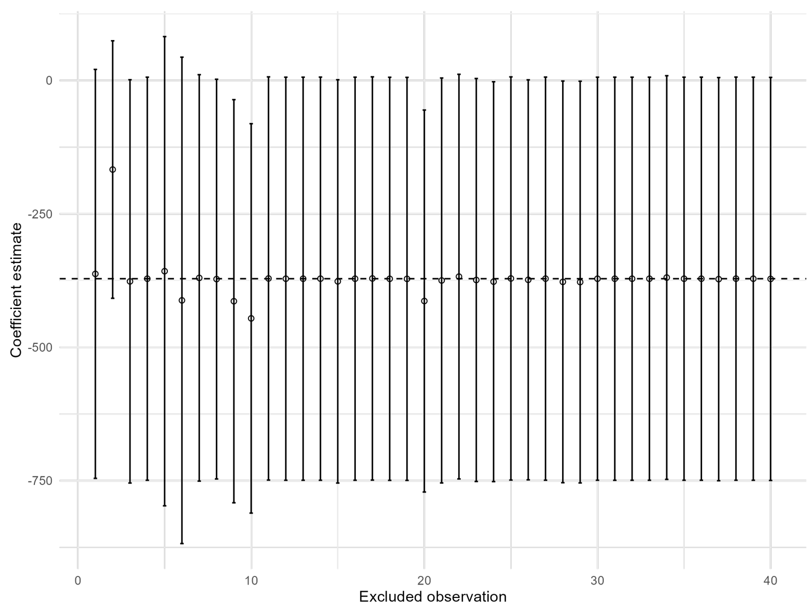
(b) Δ HHI

Figure A.20: Commodity Boom and Export Concentration (Exclusion of Micro-Regions - Low 40 Based on $\Delta\tilde{X}r$ Value)

Notes: The figure shows the robustness of the results to excluding, one by one, the last 40 micro-regions with the lowest $\Delta\tilde{X}r$. The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—lowest $\Delta\tilde{X}_r$ exposure is the first observation.



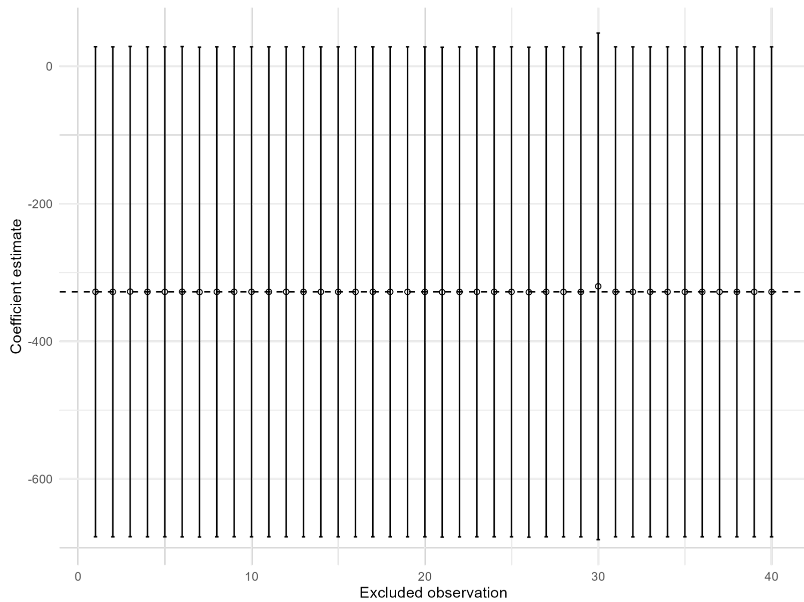
(a) $\Delta\ S_{r,t}$ (WDI)



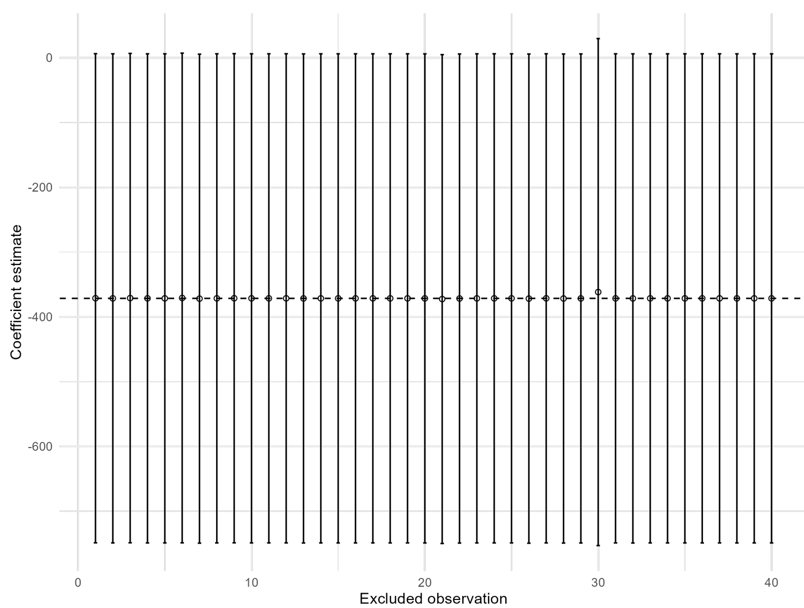
(b) $\Delta\ S_{r,t}$ (PWT)

Figure A.21: Commodity Boom and Export Sophistication (Exclusion of Micro-Regions - Top 40 Based on $\Delta\tilde{X}r$ Value)

Notes: The figure shows the robustness of the results to excluding, one by one, the top 40 micro-regions with the highest $\Delta\tilde{X}r$. The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—highest $\Delta\tilde{X}_r$ exposure is the first observation.



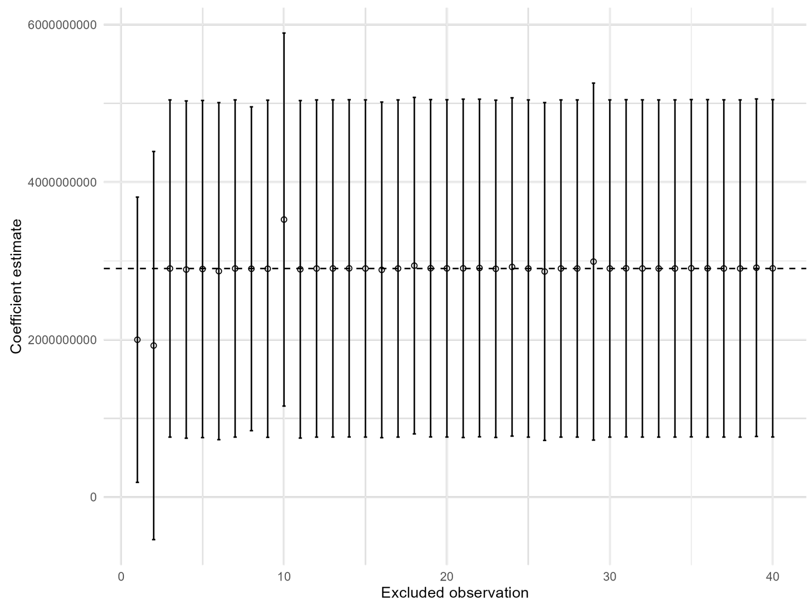
(a) $\Delta\ S_{r,t}$ (WDI)



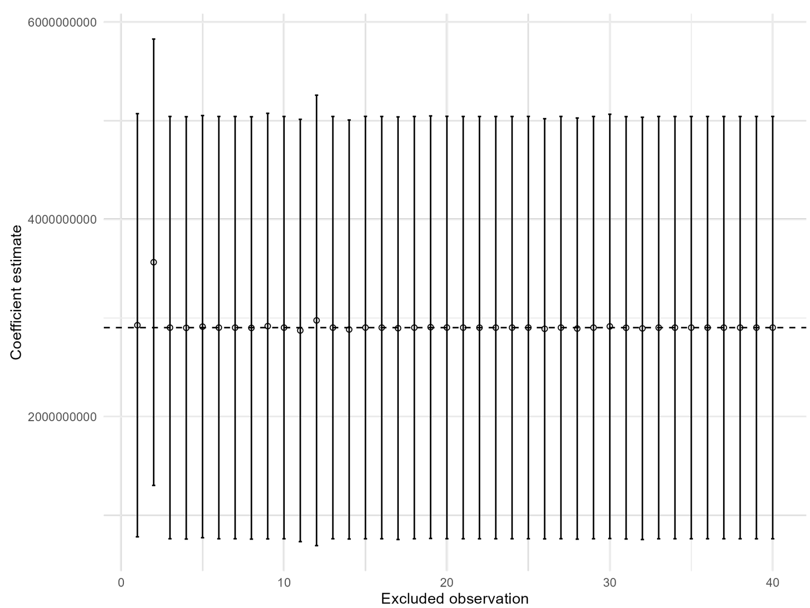
(b) $\Delta\ S_{r,t}$ (PWT)

Figure A.22: Commodity Boom and Export Sophistication (Exclusion of Micro-Regions - Low 40 Based on $\Delta\tilde{X}r$ Value)

Notes: The figure shows the robustness of the results to excluding, one by one, the last 40 micro-regions with the lowest $\Delta\tilde{X}r$. The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—lowest $\Delta\tilde{X}_r$ exposure is the first observation.



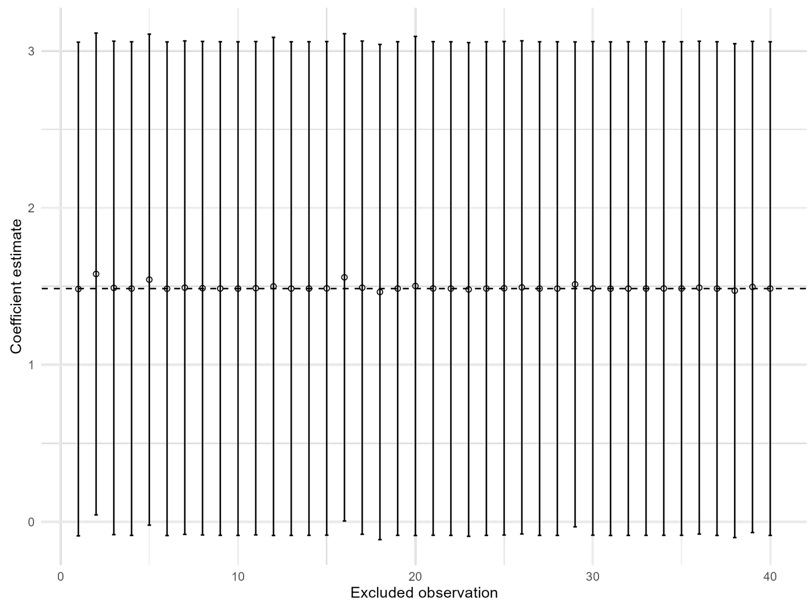
(a) Δ Value of Exports — Top



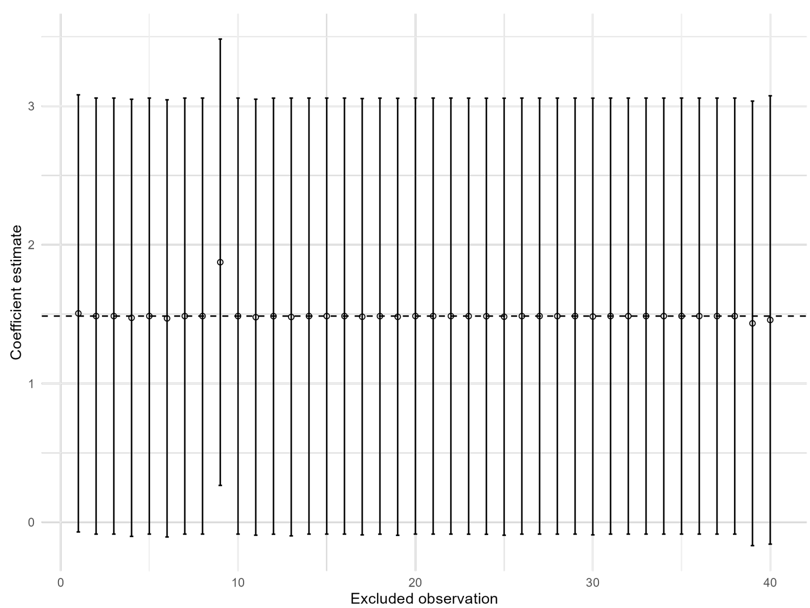
(b) Δ Value of Exports — Low

Figure A.23: Commodity Boom and Export Value (Exclusion of Micro-Regions - Top and Low 40 Based on Δ Value of Exports)

Notes: The figures show the robustness of the results to excluding, one by one, the top and bottom 40 micro-regions with the highest and lowest Δ Value of Exports. The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—highest or lowest Δ Value of Exports is the first observation.



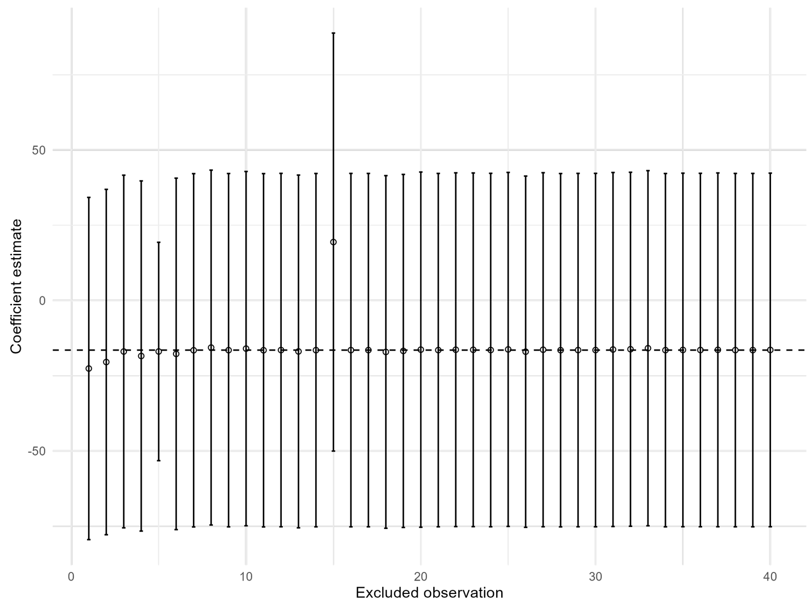
(a) %Δ Value of Exports — Top



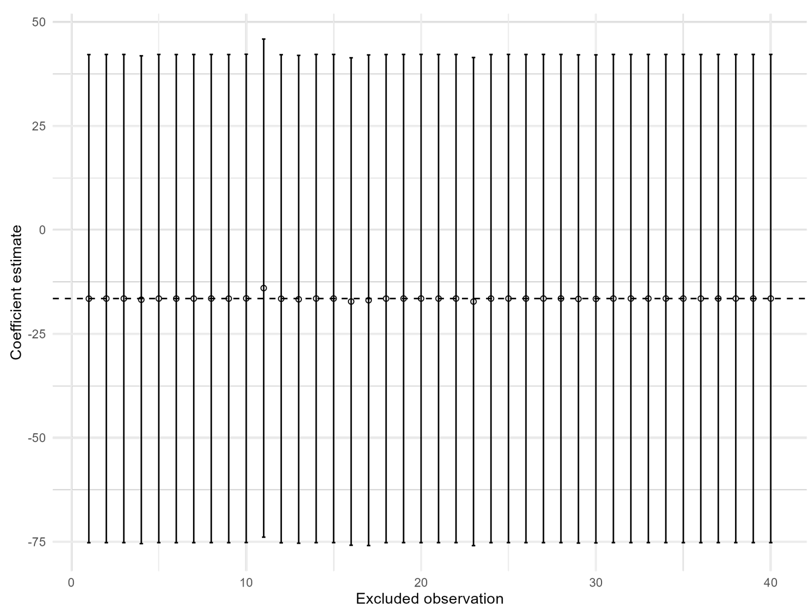
(b) %Δ Value of Exports — Low

Figure A.24: Commodity Boom and Growth Rate of Export Value (Exclusion of Micro-Regions - Top and Low 40 Based on %Δ Value of Exports)

Notes: The figures show the robustness of the results to excluding, one by one, the top and bottom 40 micro-regions with the highest and lowest %Δ Value of Exports. The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—highest or lowest %Δ Value of Exports is the first observation.



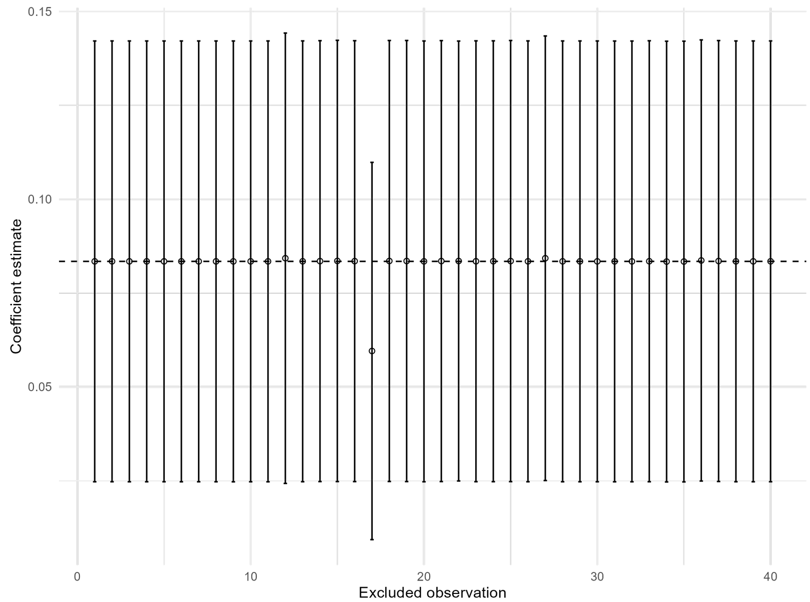
(a) Δ Lines — Top



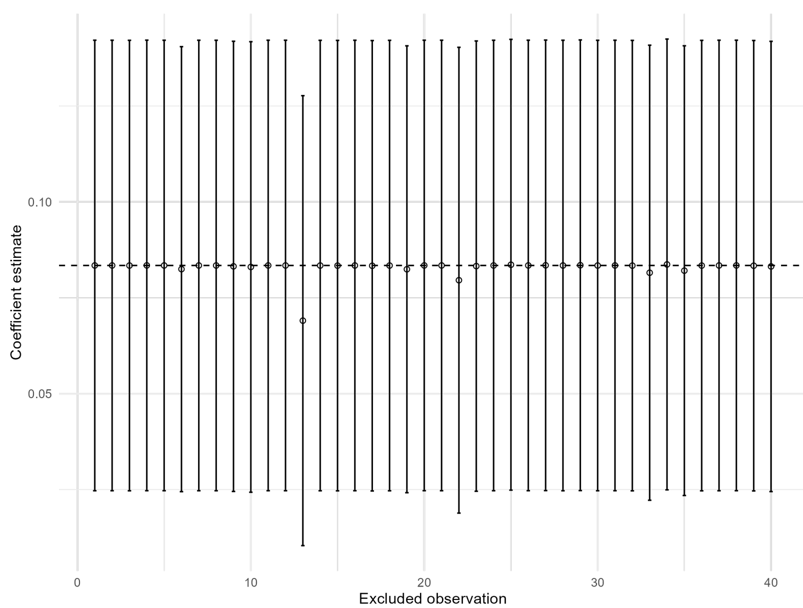
(b) Δ Lines — Low

Figure A.25: Commodity Boom and Number of Export Lines (Exclusion of Micro-Regions - Top and Low 40 Based on Δ Lines)

Notes: The figures show the robustness of the results to excluding, one by one, the top and bottom 40 micro-regions with the highest and lowest Δ Lines. The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—highest or lowest Δ Lines is the first observation.



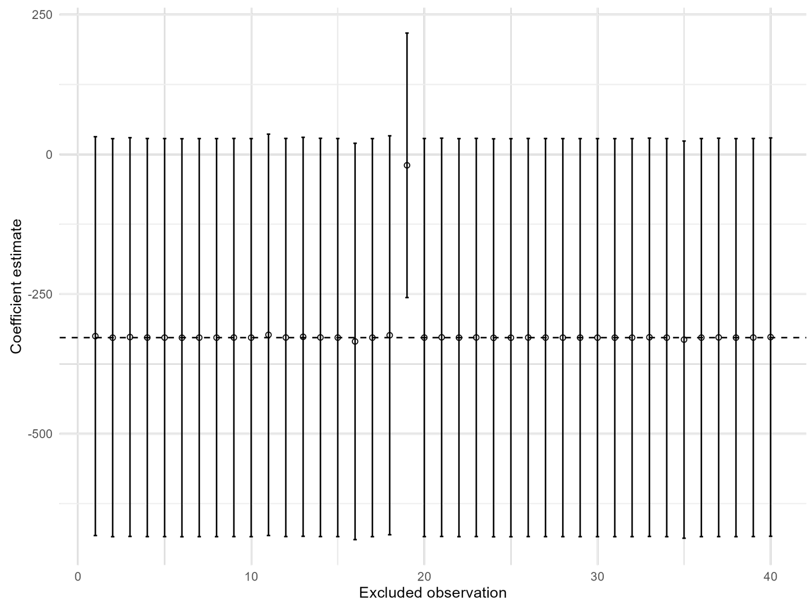
(a) Δ HHI — Top



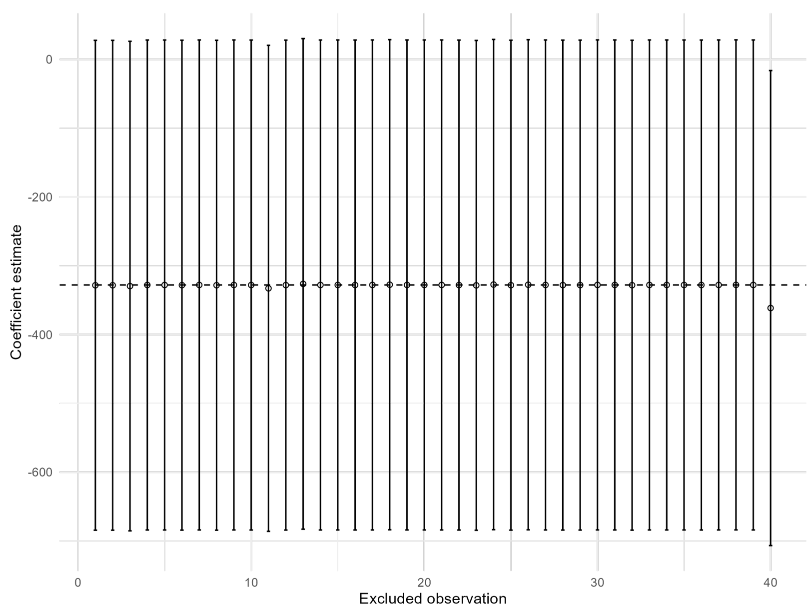
(b) Δ HHI — Low

Figure A.26: Commodity Boom and Export Concentration (HHI) (Exclusion of Micro-Regions - Top and Low 40 Based on Δ HHI)

Notes: The figures show the robustness of the results to excluding, one by one, the top and bottom 40 micro-regions with the highest and lowest Δ HHI. The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—highest or lowest Δ HHI is the first observation.



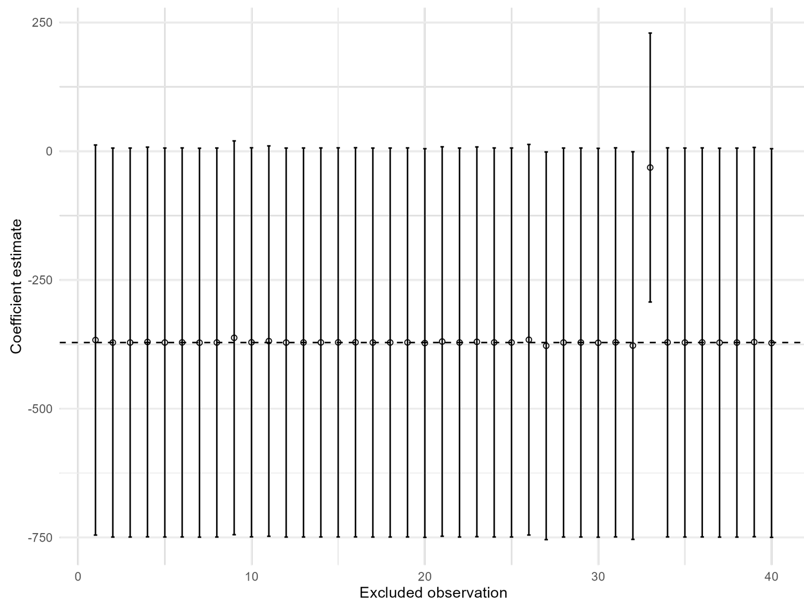
(a) $\Delta\ S_{r,t}$ (WDI) — Top



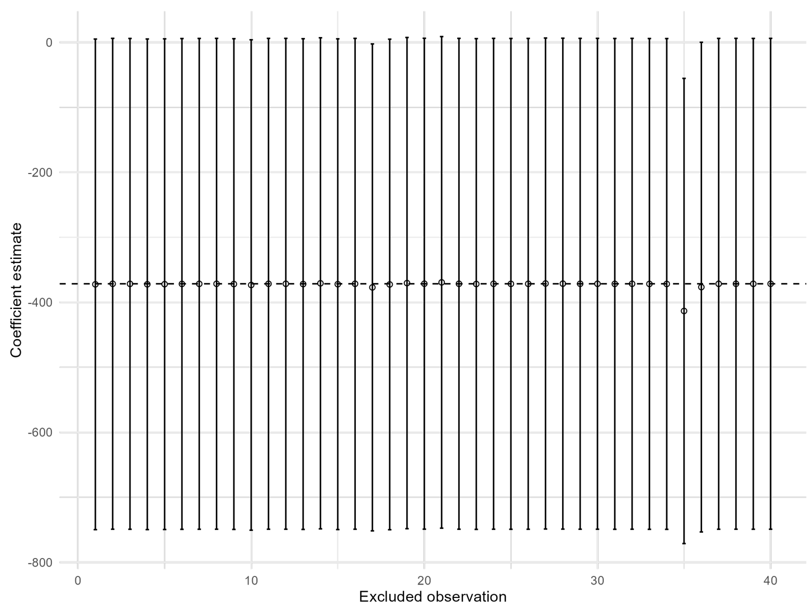
(b) $\Delta\ S_{r,t}$ (WDI) — Low

Figure A.27: Commodity Boom and Export Sophistication (WDI) (Exclusion of Micro-Regions - Top and Low 40 Based on $\Delta\ S_{r,t}$ (WDI))

Notes: The figures show the robustness of the results to excluding, one by one, the top and bottom 40 micro-regions with the highest and lowest $\Delta\ S_{r,t}$ (WDI). The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—highest or lowest $\Delta\ S_{r,t}$ (WDI) is the first observation.



(a) $\Delta\, S_{r,t}$ (PWT) — Top

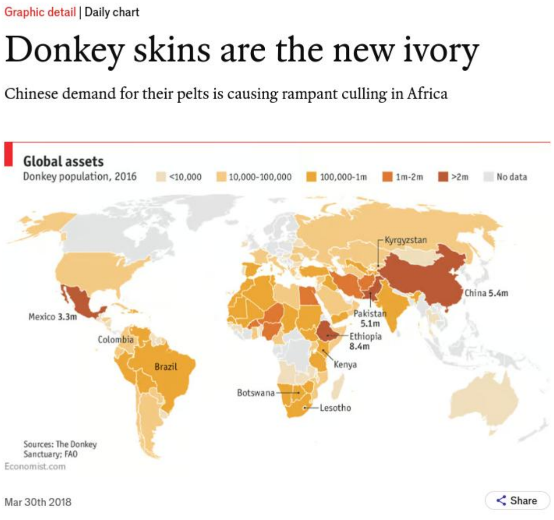


(b) $\Delta\, S_{r,t}$ (PWT) — Low

Figure A.28: Commodity Boom and Export Sophistication (PWT) (Exclusion of Micro-Regions - Low 40 Based on $\Delta\, S_{r,t}$ (PWT))

Notes: The figure shows the robustness of the results to excluding, one by one, the last 40 micro-regions with the highest and lowest $\Delta\, S_{r,t}$ (PWT). The estimated coefficients and confidence intervals at 95% are reported. Each coefficient and confidence interval emanate from a single estimation. Micro-regions are ranked from left to right—highest or lowest $\Delta\, S_{r,t}$ (PWT) is the first observation.

B.1 Appendix B



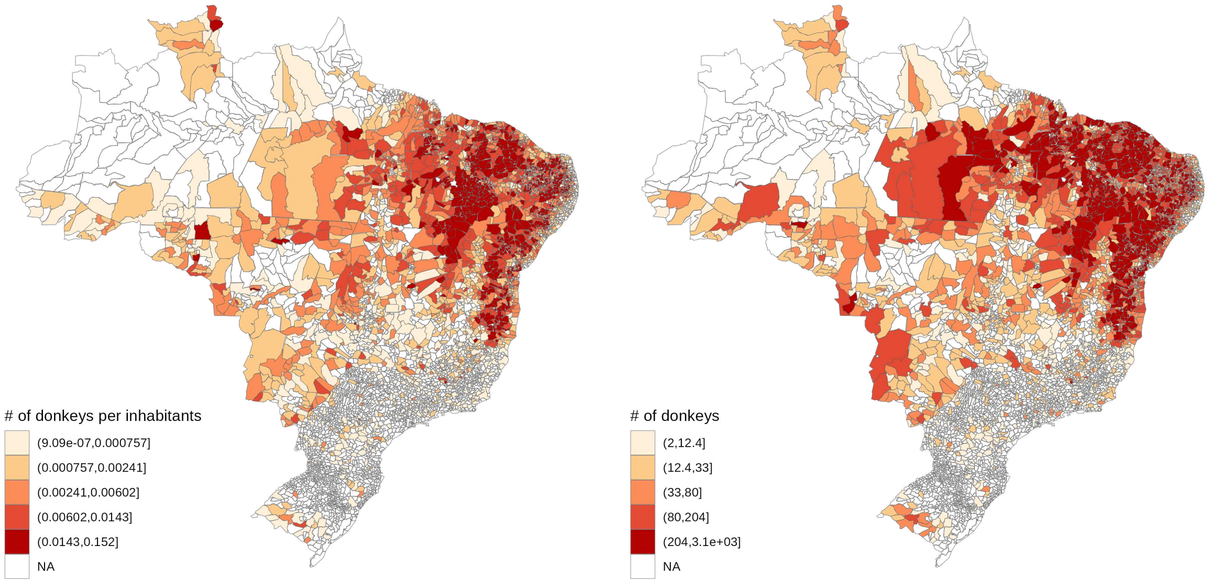
(a) The Economist (2018)

(b) *Ejiao* products. Source: The Donkey Sanctuary.

(c) Free-roaming donkeys in Northeast of Brazil. Source: G1.

(d) Chinese slaughterhouse in Amargosa - BA. Source: G1.

Figure B.1: Institutional background



(a) Number of donkeys per capita

(b) Number of donkeys

Source: Data on donkey population is from the Agricultural Census.

Figure B.2: Spatial distribution of donkeys in Brazilian municipalities (2017)

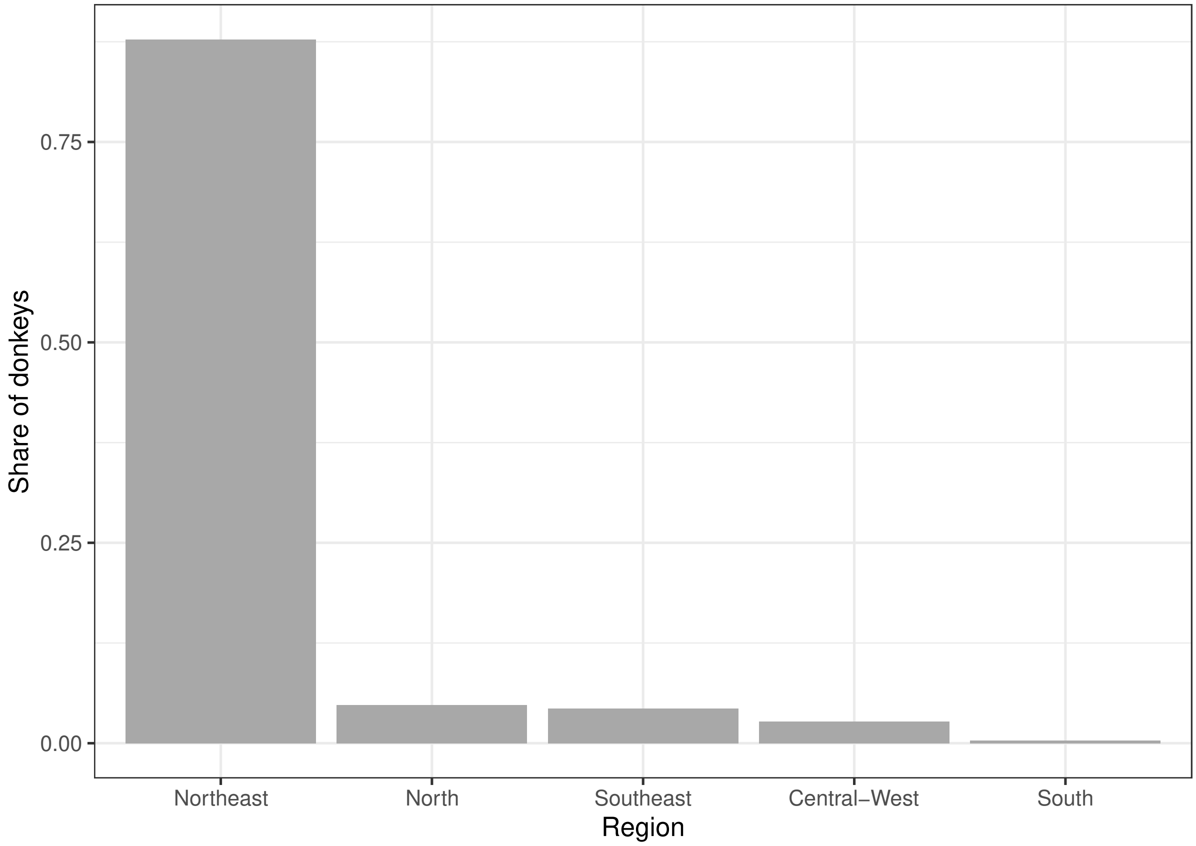


Figure B.3: Distribution of donkeys per region in Brazil (2017)

Table B.1: Descriptive statistics – quarterly data

Variables	*Source*	*Mean*	*Standard Deviation*	*Min*	*Max*	*Observations*
Homicide Rate	DATASUS	6.45	9.29	0.00	215.98	63,604
Donkey Rate	IBGE	12.59	17.69	0.00	152.49	63,604
Population	IBGE	32104.47	119368.73	1228.00	2953986.00	63,604
GDP per Capita (2010)	IBGE	6038.38	8816.70	2404.20	296884.69	63,604
Agr. Share in GDP (2010)	IBGE	0.15	0.12	0.00	0.67	63,604
Goat Rate	IBGE	314.12	774.11	0.00	8674.18	63,604
Traffic Mortality Rate (2012)	DATASUS	29.21	22.03	0.00	160.32	63,604
Suicide Mortality Rate (2012)	DATASUS	4.82	7.90	0.00	77.97	63,604

B.1.1 Yearly outcomes

In this section we present the same sequence of results from Tables 2.1, 2.2, and 2.4 using yearly data in our estimations. The results are notably similar to those presented in Section 1.3.

Table B.2: Descriptive statistics – yearly data

Variables	*Sources*	*Mean*	*Standard Deviation*	*Min*	*Max*	*Observations*
Homicide Rate	DATASUS	25.43	25.87	0.00	304.35	16,146
Donkey Rate	IBGE	12.73	17.90	0.00	152.49	16,146
Population	IBGE	31684.94	118511.17	1228.00	2953986.00	16,146
GDP per Capita (2010)	IBGE	6031.45	8757.80	2404.20	296884.69	16,146
Agr. Share in GDP (2010)	IBGE	0.15	0.11	0.00	0.67	16,146
Goat Rate	IBGE	320.04	784.54	0.00	8674.18	16,146
Traffic Mortality Rate (2012)	DATASUS	29.22	22.28	0.00	160.32	16,146
Suicide Mortality Rate (2012)	DATASUS	4.82	7.97	0.00	77.97	16,146

B.1.2 Homicides per place of residence

Next, we perform a similar robustness exercise using yearly data in our estimations but now considering the homicide rate per place of residence. Again, the results are quite similar to those presented in Section 1.3.

Table B.3: Market regulation and crime rate – yearly outcome

Dep. var: Homicides$_{i,t}$	OLS/DiD (1)	OLS/DiD (2)	OLS/DiD (3)
Slaughter × Donkeys	0.2723***	0.2757***	0.0953**
	(0.0830)	(0.0800)	(0.0486)
Observations	16,146	16,146	16,146
Adj. R^2	0.6732	0.7175	0.6908
Municipality fixed effects	✓	✓	✓
State-Year fixed effects		✓	✓

Notes: Standard errors (in parentheses) are clustered for the micro-region level. Unit of analysis is a municipality. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

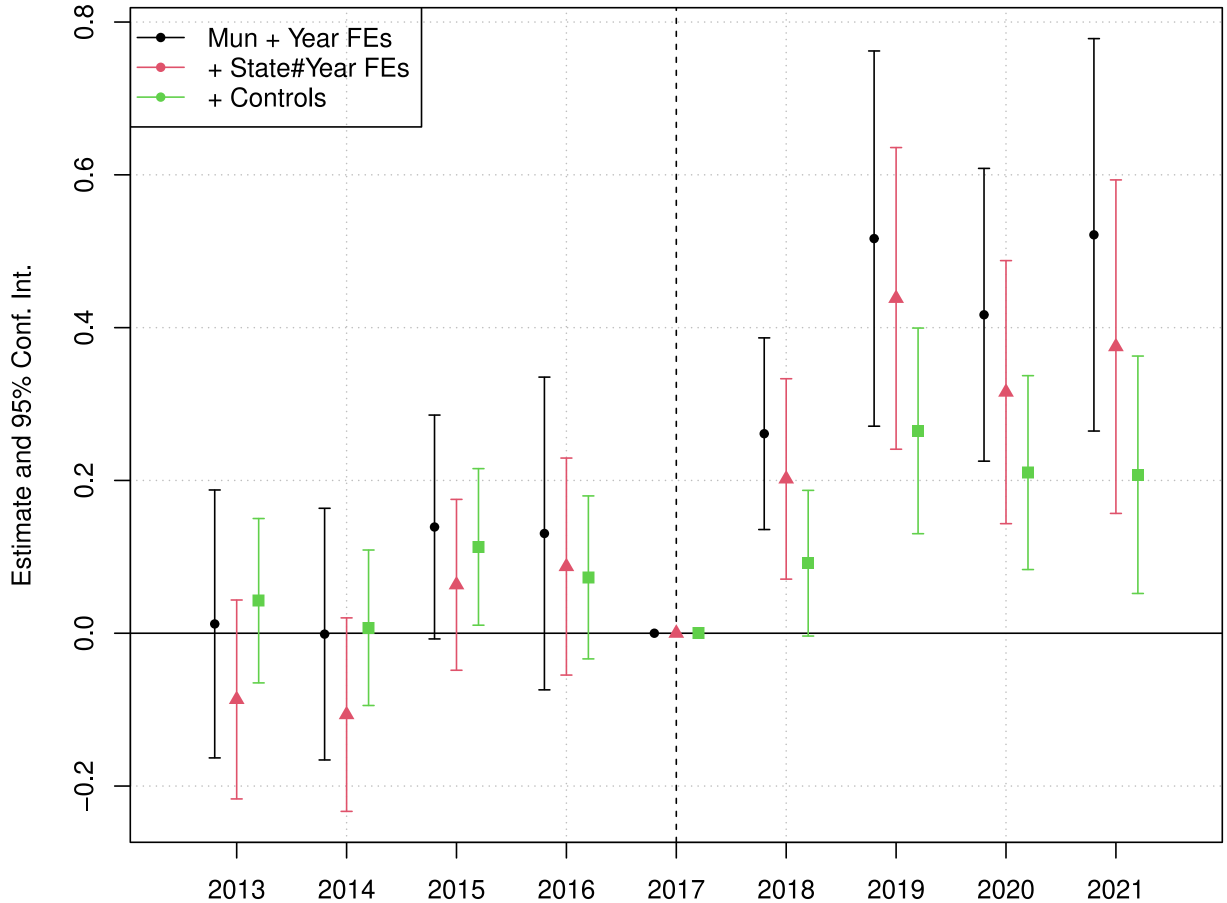
Table B.4: Robustness to alternative samples and treatments – yearly outcomes

	Alternative sample			Alternative treatment	
Dep. var: Homicides$_{i,t}$	Some donkey (1)	BA only (2)	Microregion (3)	Donkey in 1995 (4)	Placebo (5)
Slaughter × Donkeys	0.1778**	0.2260	0.5088***		
	(0.0758)	(0.1579)	(0.1508)		
Slaughter × Donkeys in 1995				0.0739***	
				(0.0247)	
Slaughter × Goats					0.0014
					(0.0011)
Observations	14,931	3,753	1,692	16,146	16,146
Adj. R^2	0.6997	0.7346	0.8529	0.7168	0.7145
Municipality fixed effects	✓	✓		✓	✓
Micro-region fixed effects			✓		
State-Year fixed effects		✓	✓	✓	✓

Notes: Standard errors (in parentheses) are clustered for the micro-region level. Unit of analysis is a municipality in columns (1),(2),(3), and (5), and a micro-region in column (4). All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

B.1.3 Robustness to alternative clusters

In this section, we evaluate the robustness of our main results to alternative spatial clusters for the standard errors. Table B.9 shows the comparison between our baseline estimates in column (2) and alternative spatial structures.



Notes: Each point reflects an individual regression coefficient, at year frequency, $\hat{\beta}_t$ following Equation (2.2). The regression includes municipality and state-time fixed effects. Standard errors are adjusted municipality clusters. Dashed lines show 95 percent confidence intervals.

Figure B.4: Event study (yearly) – Northeast region municipalities

We also perform a similar robustness exercise for the specification using micro-regions as the unit of analysis. The results are summarized in B.10.

B.1.4 Sensitivity analysis

To further evaluate the robustness of our main results, presented in Table 2.1, we perform a simple sensitivity analysis for the coefficients associated with the impacts of the donkey market regulation on local crime rates.

Table B.5: Effects of market regulation on crime rate per distance to slaughterhouse – yearly outcome

	Distance to closest slaughterhouse			
Dep. var: Homicides$_{i,t}$	Baseline (1)	250km (2)	500km (3)	1000km (4)
Slaughter × Donkeys	0.2757***	0.5357***	0.2994**	0.1981***
	(0.0800)	(0.1584)	(0.1236)	(0.0637)
Observations	16,146	2,475	4,725	13,707
Adj. R^2	0.6804	0.6576	0.6910	0.6676
Municipality fixed effects	✓	✓	✓	✓
State-Year fixed effects	✓	✓	✓	✓

Notes: Standard errors (in parentheses) are clustered for the micro-region level. Unit of analysis is a municipality. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

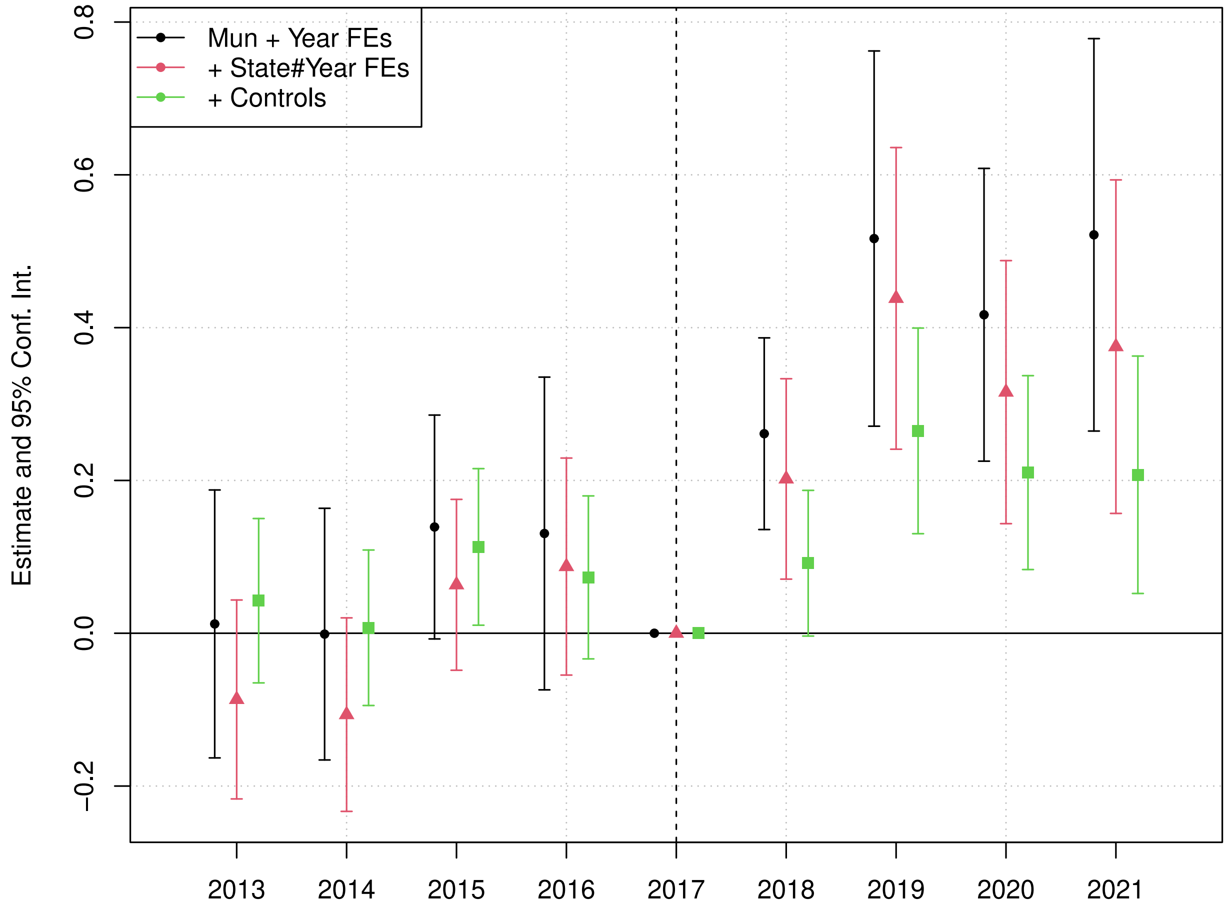
Table B.6: Market regulation and crime rate – yearly outcome per place of residence

Dep. var: Homicides$_{i,t}$	OLS/DiD (1)	OLS/DiD (2)	OLS/DiD (3)
Slaughter × Donkeys	0.2628***	0.2669***	0.1108**
	(0.0777)	(0.0729)	(0.0442)
Observations	16,146	16,146	16,146
Adj. R^2	0.6391	0.6884	0.6964
Municipality fixed effects	✓	✓	✓
State-Year fixed effects		✓	✓

Notes: Standard errors (in parentheses) are clustered for the micro-region level. Unit of analysis is a municipality. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

First, we estimated the preferred specification by removing, in each regression, one of the top and bottom 40 municipalities, respectively, in terms of i) the donkey population density, and ii) the average homicide rate for the period of analysis.

Figures B.6 and B.7 graphically present the results of the estimations removing the top- and bottom-ranked municipalities in terms of donkey population density and average homicide rates



Notes: Each point reflects an individual regression coefficient, at year frequency, $\hat{\beta}_t$ following Equation (2.2). The regression includes municipality and state-time fixed effects. Standard errors are adjusted municipality clusters. Dashed lines show 95 percent confidence intervals.

Figure B.5: Event study (yearly) – Northeast region municipalities per place of residence

respectively, emphasizing the coefficient of interest and the 95% confidence interval for each one of the regressions. The dotted black line describes the coefficient obtained in our baseline regression.

Next, we perform a similar sensitivity analysis, but now we drop the top- and bottom-ranked municipalities following the same criteria one to many. That is, we sequentially drop the top- and bottom-ranked local economies from the estimation up to 40 municipalities. Figures B.8 and B.9 present, respectively, the results for excluding regions classified by donkey population density and average homicide rates. Again, we emphasize the coefficient of interest and the 95% confidence interval for each one of the regressions. The dotted black line describes the coefficient obtained in our baseline regression.

Table B.7: Robustness to alternative samples and treatments – yearly outcomes per place of residence

	Alternative sample			Alternative treatment	
Dep. var: Homicides$_{i,t}$	Some donkey (1)	BA only (2)	Microregion (3)	Donkey in 1995 (4)	Placebo (5)
Slaughter × Donkeys	0.1905***	0.2189	0.4621***		
	(0.0705)	(0.1354)	(0.1397)		
Slaughter × Donkeys in 1995				0.4621***	
				(0.0230)	
Slaughter × Goats					0.0014
					(0.0010)
Observations	14,931	3,753	1,692	16,146	16,146
Adj. R^2	0.6693	0.7104	0.8484	0.6876	0.6852
Municipality fixed effects	✓	✓		✓	✓
Micro-region fixed effects			✓		
State-Year fixed effects		✓	✓	✓	✓

Notes: Standard errors (in parentheses) are clustered for the micro-region level. Unit of analysis is a municipality in columns (1),(2),(3), and (5), and a micro-region in column (4). All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table B.8: Effects of market regulation on crime rate per distance to slaughterhouse – yearly outcome per place of residence

	Distance to closest slaughterhouse			
Dep. var: Homicides$_{i,t}$	Baseline (1)	250km (2)	500km (3)	1000km (4)
Slaughter × Donkeys	0.2669***	0.4756***	0.2792**	0.1941***
	(0.0729)	(0.1421)	(0.1082)	(0.0557)
Observations	16,146	2,475	4,725	13,707
Adj. R^2	0.6474	0.6128	0.6516	0.6373
Municipality fixed effects	✓	✓	✓	✓
State-Year fixed effects	✓	✓	✓	✓

Notes: Standard errors (in parentheses) are clustered for the micro-region level. Unit of analysis is a municipality. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

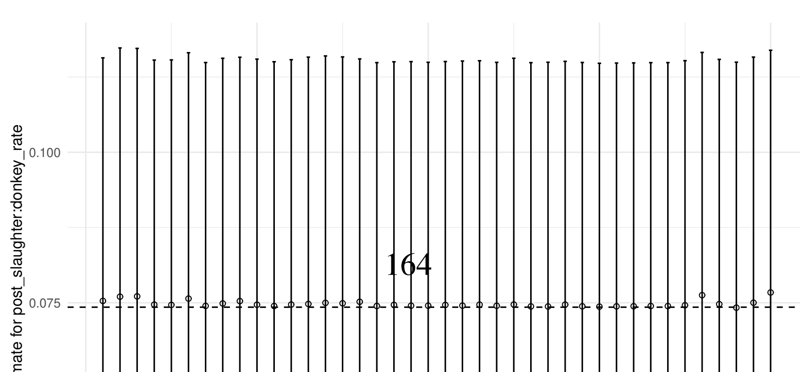


Table B.9: Market regulation and crime rates – alternative clustering

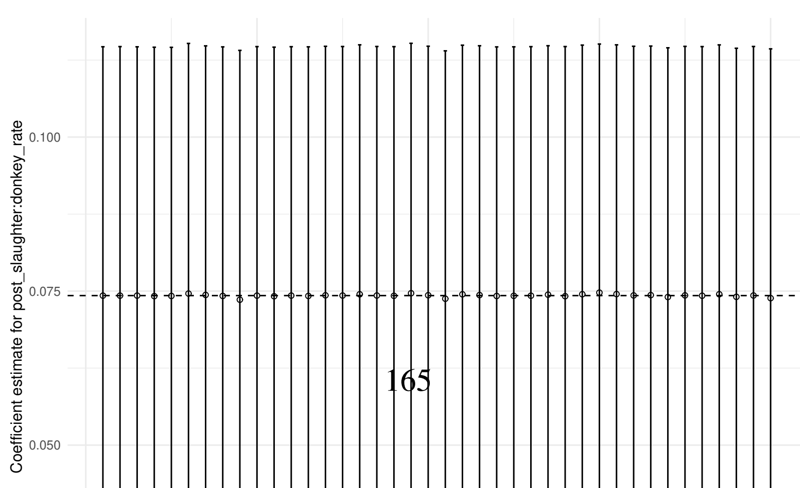
	Alternative clusters			
Dep. var: Homicides$_{i,t}$	Municipality (1)	Micro-region (2)	Meso-region (3)	State (4)
Slaughter × Donkeys	0.0743***	0.0743***	0.0743***	0.0743**
	(0.0174)	(0.0205)	(0.0207)	(0.0236)
Standard-Errors	Municipality	Micro-region	Meso-region	State
Observations	63,604	63,604	63,604	63,604
Adj. R^2	0.4851	0.4851	0.4851	0.4851
Municipality fixed effects	✓	✓	✓	✓
State-Quarter fixed effects	✓	✓	✓	✓

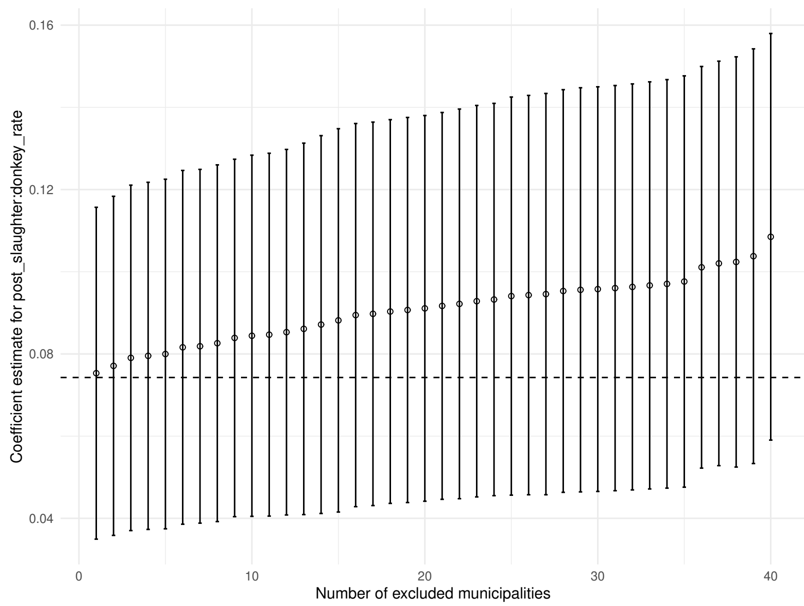
Notes: Unit of analysis is a municipality. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table B.10: Market regulation and crime rates – Micro-region aggregation and alternative clustering

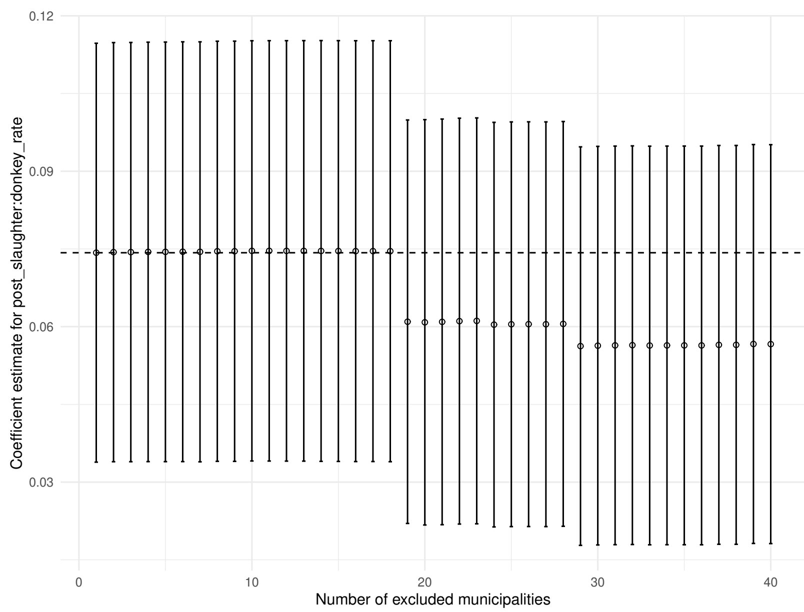
	Alternative clusters		
Dep. var: Homicides$_{i,t}$	Micro-region (1)	Meso-region (2)	State (3)
Slaughter × Donkeys	0.1346***	0.1346***	0.1346***
	(0.0383)	(0.0401)	(0.0390)
Standard-Errors	Microregion	Mesoregion	State
Observations	6,752	6,752	6,752
Adj. R^2	0.7667	0.7667	0.7667
Municipality fixed effects	✓	✓	✓
State-Quarter fixed effects	✓	✓	✓

Notes: Unit of analysis is a micro-region. All specifications are weighted by population. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.





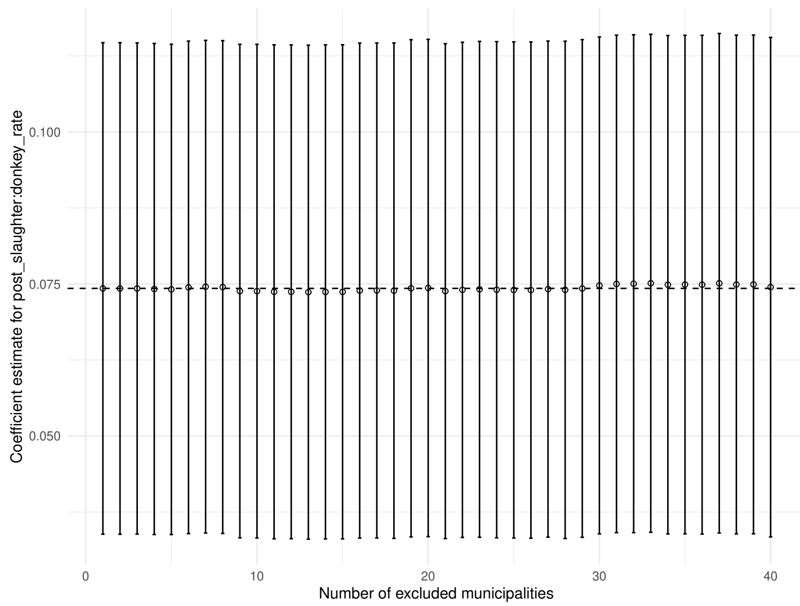
(a) Top municipalities



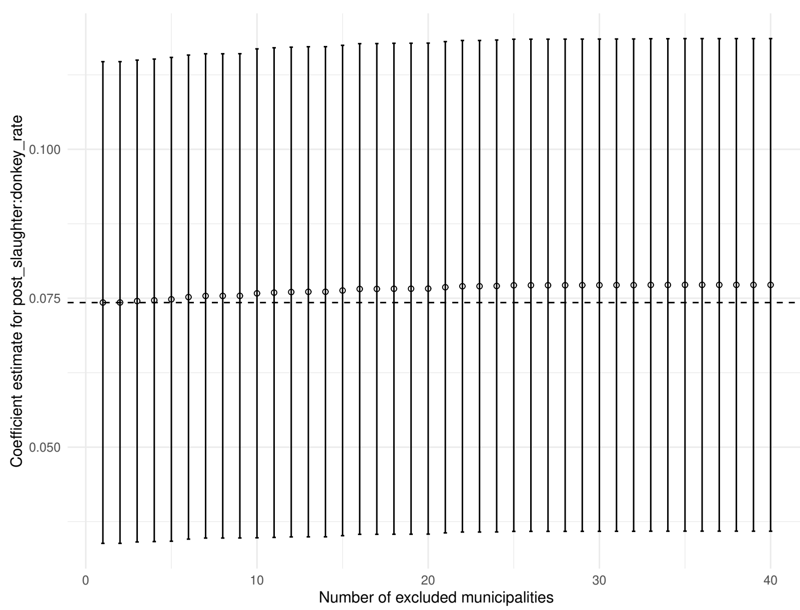
(b) Bottom municipalities

Figure B.8: Market regulation and crime rates
(Exclusion of municipalities - First and last 40 based on donkey population density)

Notes: The figure shows the robustness of the results to excluding the top and bottom 40 municipalities in terms of donkey population density. The estimated coefficients and confidence intervals at 95 percent are reported. Each estimated coefficient and confidence interval emanate from a single estimation.



(a) Top municipalities



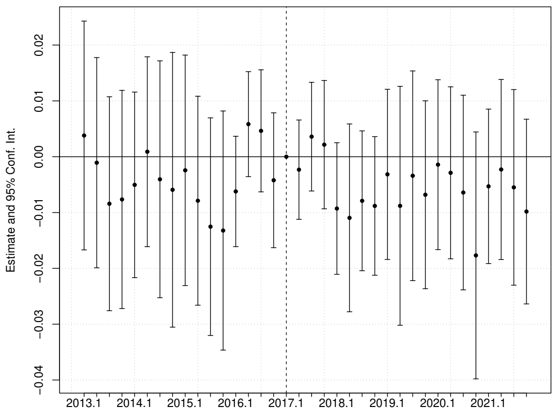
(b) Bottom municipalities

Figure B.9: Market regulation and crime rates
(Exclusion of municipalities - First and last 40 based on homicide rate)

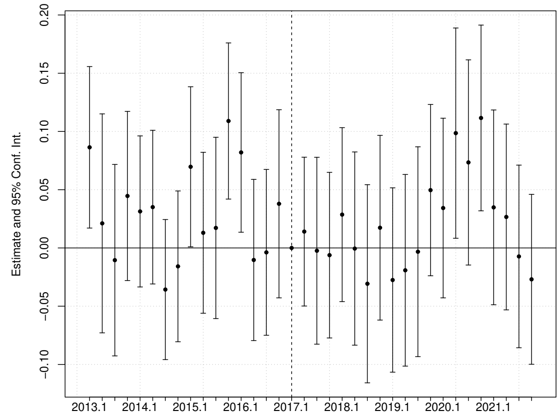
Notes: The figure shows the robustness of the results to excluding the top and bottom 40 municipalities in terms of average homicide rates. The estimated coefficients and confidence intervals at 95 percent are reported. Each estimated coefficient and confidence interval emanate from a single estimation.

B.1.5 Falsification exercise dynamic effects

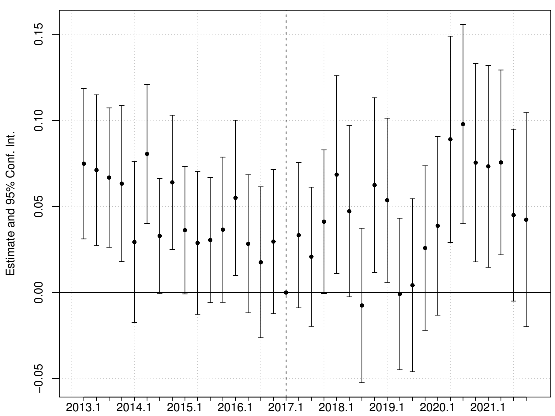
Lastly, Figure B.10 presents the dynamic effects of donkey market regulation on specific causes of mortality used in our falsification exercise, as examined in Table 2.3. Each point in the figure represents an individual regression coefficient, $\hat{\beta}_t$, estimated at the quarterly level following Equation (2.2), with the respective cause of mortality as the dependent variable. Standard errors are clustered at the micro-region level, and the solid lines depict the 95 percent confidence intervals.



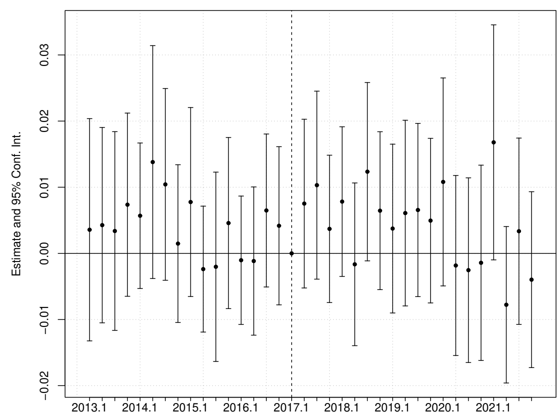
(a) Undetermined



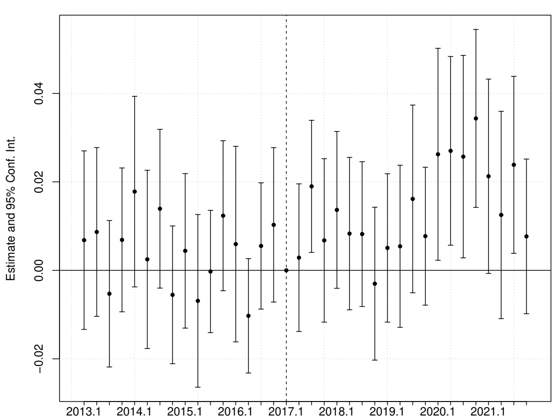
(b) Cardiovascular



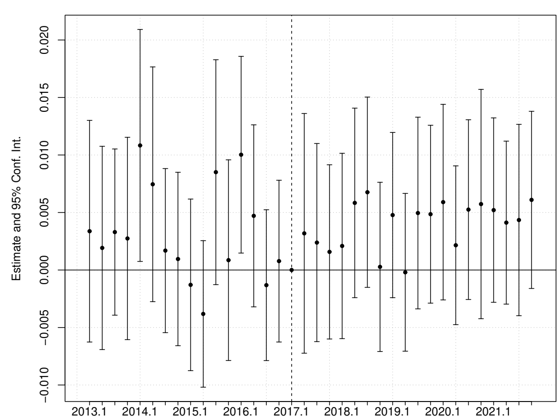
(c) Neoplasm



(d) Liver



(e) 0-6 years



(f) 7-14 years

Figure B.10: Dynamic effects for falsification

C.1 Appendix C

C.1.1 Potential yield in soy production

Before delving into our analysis of electoral data, it is essential to rigorously assess the underlying data correlations. Specifically, we investigate the interconnections between soy production expansion at the municipal level, potential yield (our instrumental variable denoting technological advancement), labor productivity, and land use intensity. To address this, we employ the following econometric specification:

$$\Delta_{96-06} y_j = \beta_0 + \beta_1 \Delta A_j{}^{soy} + \beta_2 X_{j,1991} + \delta_{s,t} + \varepsilon_{j,t} \qquad (15)$$

where Δy_j is the change in outcome variable in municipality j between 1996 and 2006 (agricultural census years), $\Delta A_j{}^{soy}$ is the potential yield of soy under the high technology minus the potential yield of soy under low technology, $X_{j,1991}$ is a set of characteristic of municipality j in 1991 and $\delta_{s,t}$ are state-period fixed effects.

Table C.1 scrutinizes the impact of technological evolution in soy production, as measured by potential yield changes, on soy cultivation area. This analysis serves to establish the validity of our technical change metric as an instrumental variable for actual soy production expansion. Our findings affirm that shifts towards high-technology inputs, specifically genetically engineered (GE) seeds, are positively correlated with soy technical advancements during the study period.

Furthermore, our estimates remain robust even when accounting for municipality-specific characteristics. This observation implies that divergent crop expansion rates among municipalities are not a consequence of initial development disparities.

Table C.2 delves into the relationship between increased soy potential yield and the expansion of genetically engineered (GE) soy's share in agricultural land from 1996 to 2006. We also conduct a falsification test by examining whether our measure of technical change in soy explains the expansion in the area planted with non-GE soy. In this case, the coefficients are negative and

Table C.1: Potential Yield and Area Planted

	Dependent variable:	
	Δ Soy area share	
	(1)	(2)
ΔA^{soy}	0.006***	0.006***
	(0.001)	(0.001)
Share rural population	0.015***	0.014***
	(0.003)	(0.004)
Log income per capita		0.007***
		(0.002)
Log pop. density		-0.003***
		(0.001)
Observations	3,652	3,652
R^2	0.273	0.279

Notes: Robust standard errors in parentheses. Unit of analysis is a municipality or minimum comparable area. Changes in dependent variables are calculated over the years 1996 and 2006. All municipality controls are from the population census of 1991. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

statistically significant at conventional levels. This finding supports our claim that the change in potential soy yield captures the benefits of adopting GE soy *vis-à-vis* traditional soy seeds.

Lastly, in Table C.3, we analyze the impact of agricultural technical change, as indicated by the change in potential yield, on agricultural output and employment. The results indicate that regions experiencing greater increases in potential soy yields witnessed larger increases in agricultural production per worker during the 1996-2006 period. However, the decrease in labor intensity lacks statistical significance.

In short, our measure of technical change in soy production emerges as a robust instrument for explaining the expansion of production and the area dedicated to GE soy cultivation. It effectively captures an exogenous increase in labor productivity in municipalities more profoundly affected by this economic transformation.

C.1.2 Chinese export demand

Next, we evaluate the validity of our shift-share or "Bartik" instrument, which leverages differential local exposure to Chinese export demand, as a measure of the regional impacts of heightened

Table C.2: Potential Yield and GE and Non-GE Soy Area Planted

	Dependent variable:			
	Δ GE soy area share		Δ Non-GE soy area share	
	(1)	(2)	(3)	(4)
ΔA^{soy}	0.013***	0.013***	-0.008***	-0.008***
	(0.002)	(0.002)	(0.001)	(0.001)
Share rural population	0.029***	0.030***	-0.013***	-0.014**
	(0.004)	(0.006)	(0.004)	(0.006)
Log income per capita		0.002		0.005
		(0.003)		(0.003)
Log pop. density		-0.001		-0.002***
		(0.001)		(0.001)
Observations	3,652	3,652	3,652	3,652
R^2	0.420	0.420	0.239	0.241

Notes: Robust standard errors in parentheses. Unit of analysis is a municipality or minimum comparable area. Changes in dependent variables are calculated over the years 1996 and 2006. All municipality controls are from the population census of 1991. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Brazilian exports to China post-2001. The credibility of this instrument, akin to the potential yield case, lies at the core of our identification strategy. Notably, our measure of the impact of increased Chinese export demand for Brazilian municipalities and the instrumental variable ($\Delta\tilde{X}_r$) reflecting trends in Chinese trade exhibit a high degree of correlation. The correlation coefficient between the observed impact and the instrumental variable for export demand shock is approximately 0.7.

Figure C.1 visually depicts the relationship between the endogenous measure of exposure and the instrument. Remarkably, our measure of the impact of increased Chinese export demand for Brazilian municipalities and the instrumental variable ($\Delta\tilde{X}_r$) reflecting trends in Chinese trade exhibit a strong degree of correlation.

Besides, to further assess the validity of our shift-share instrument, we estimate the following "first-stage" specification:

$$\Delta_{00-19}X_r = \beta_0 + \beta_1\Delta\tilde{X}_r + \delta_{s,t} + \varepsilon_{r,t} \tag{16}$$

Table C.3: Potential Yield, Productivity and Intensity

	Dependent variable:			
	Δ log output per worker		Δ log labor intensity	
	(1)	(2)	(3)	(4)
ΔA^{soy}	0.051***	0.048**	0.021	0.013
	(0.019)	(0.019)	(0.013)	(0.013)
Share rural population	0.168***	0.077	-0.051	-0.101*
	(0.065)	(0.088)	(0.048)	(0.059)
Log income per capita		-0.029		0.076**
		(0.048)		(0.034)
Log pop. density		-0.023		-0.049***
		(0.016)		(0.013)
Observations	4,149	4,149	4,149	4,149
R^2	0.037	0.038	0.057	0.064

Notes: Robust standard errors in parentheses. Unit of analysis is a municipality or minimum comparable area. Changes in dependent variables are calculated over the years 1996 and 2006. All municipality controls are from the population census of 1991. All specifications control for state-period fixed effects. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

where $\Delta_{00-19}X_r$ is the observed impact of the variation in exports to China between 2000 and 2019 in the labor market of municipality r and $\delta_{s,t}$ is a time-state or time-micro-region fixed effect.

Table C.4 presents the results of "first-stage" regressions of our shift-share instrument against the exposure of Brazilian municipalities to Chinese export demand, as described in Equation (16). Columns (2) and (3) include, respectively, state and micro-region fixed effects.

Overall, our results provide compelling evidence that the robust positive correlation observed between the variables indeed translates into a sound instrumental variable.

C.1.3 Additional results

Figure C.1: Correlation: Chinese export demand and shift-share instrumental variable

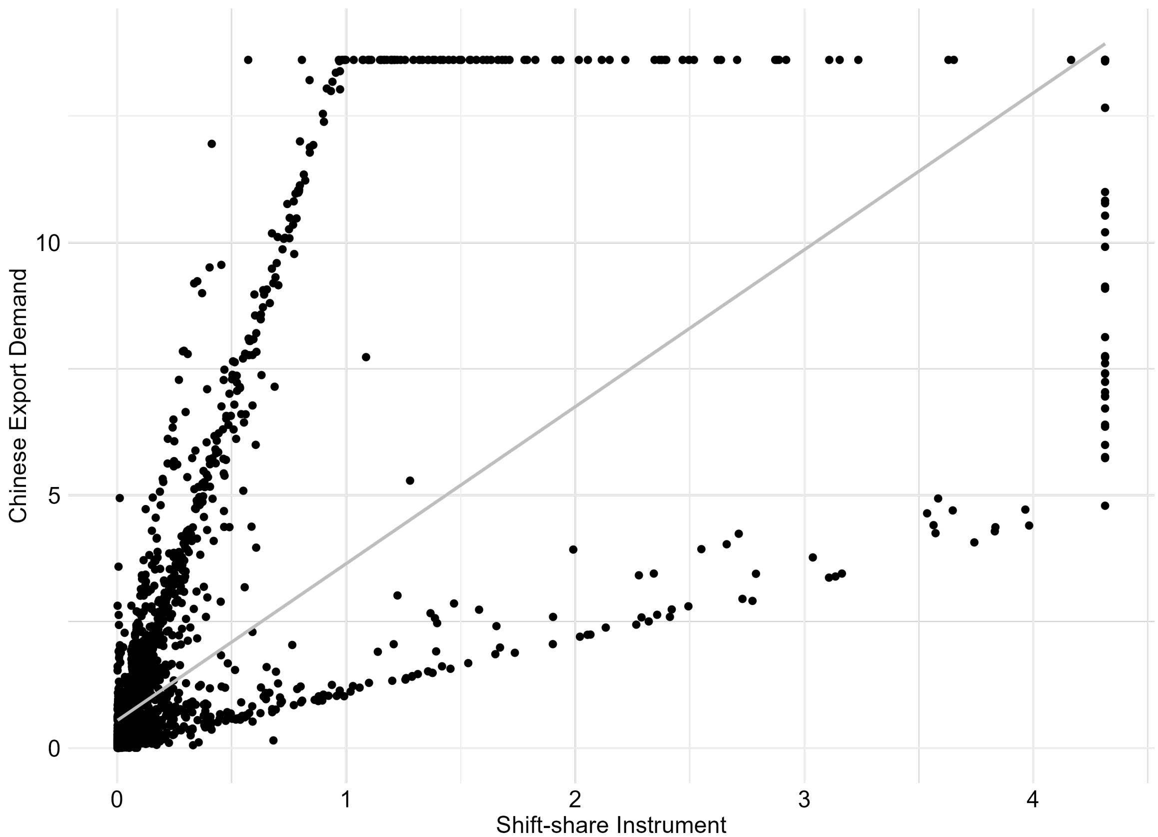
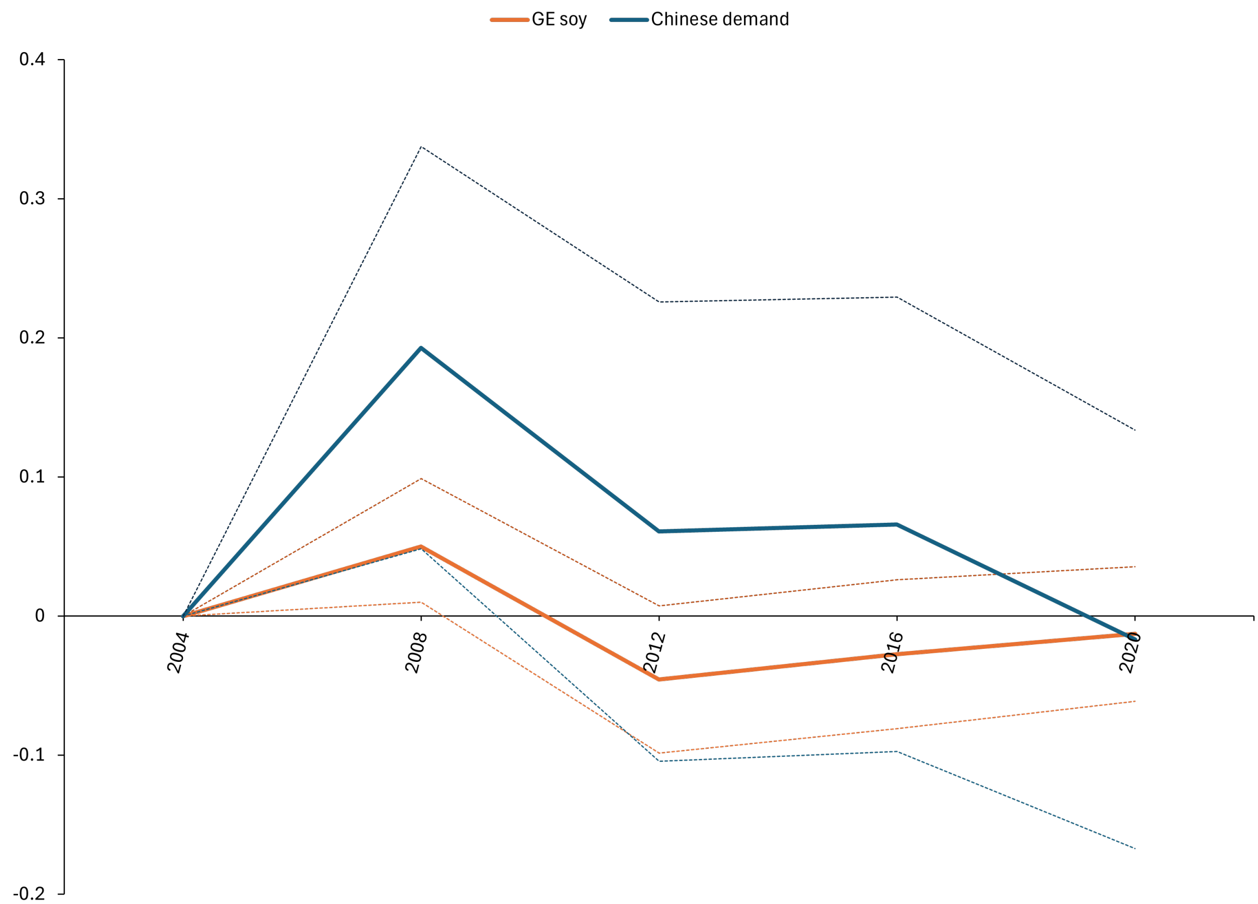


Table C.4: Exposure to China's Export Demand and Shift-Share Instrument

	Dependent variable:		
	China Export Demand	China Export Demand	China Export Demand
	(1)	(2)	(3)
$\Delta\tilde{X}_r$	3.103***	2.944***	2.539***
	(0.131)	(0.110)	(0.092)
State FE		X	
Microregion FE			X
Observations	4,258	4,258	4,258
R^2	0.707	0.750	0.856
KP F-Stat	4,044.137	237.063	25.026

Notes: Robust standard errors in parentheses. Unit of analysis is a municipality or minimum comparable area. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$



Notes: Each point reflects an individual regression coefficient $\hat{\delta_1}$ and $\hat{\delta_2}$ following Equation (3.14), where the dependent variable is the probability of winning a mayoral election in years $t = 2004, ..., 2020$. The regression includes municipality, election and municipality-year fixed effects. Standard errors are adjusted for 556 micro-region clusters. Dashed lines show 95 percent confidence intervals.

Figure C.2: Dynamic effects of exogenous regional economic shocks on probability of electoral victory - Difference-in-Difference